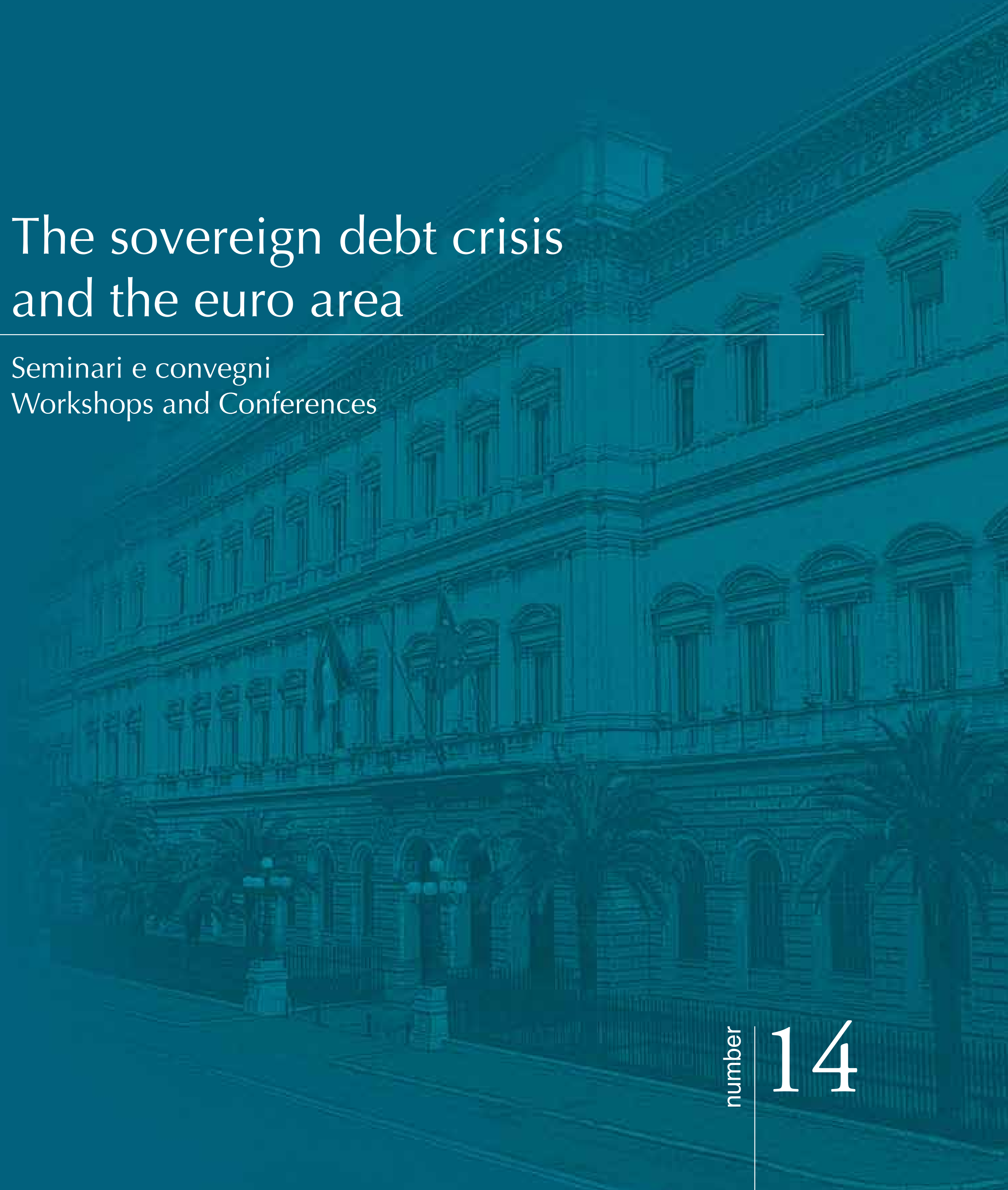
BANCA D'ITALIA
EUROSISTEMA

The sovereign debt crisis and the euro area

Seminari e convegni
Workshops and Conferences



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This volume contains papers presented at the research workshop entitled “The Sovereign Debt Crisis and the Euro Area”, held in Rome on 15 February 2013. The workshop and the preparation of this volume of papers were organized by Giuseppe Grande, Stefano Neri and Stefano Siviero, with the collaboration of Marco Romani and, for the editorial aspects, Valentina Schirosi.

The papers are available, in English only, on the Bank of Italy website at the following link:

http://www.bancaditalia.it/studiricerche/convegni/atti;internal&action=_setlanguage.action?LANGUAGE=en

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Foreword

Daniele Franco[1]

The sovereign debt crisis in the euro area has been at the centre of the economic policy debate and research analysis for some years now. The development of the crisis conditioned the European economic recovery and subjected the area to serious financial and political tensions. To address the crisis some institutional innovations were introduced to support the process of integration within the area. The success of these policy initiatives will change the role of the European Union. From the point of view of the analysis, the debt crisis provides a starting point for research which will influence academic research in the years to come: the euro area, in which sovereign nations share monetary policy while retaining responsibility for others, represents a new kind of experiment which has implications that were not well understood until a few years ago.

The Bank of Italy has invested important resources in policy analysis and research. The Economic Research and International Relations Area has carried out various research projects which analyse the impact of the sovereign debt crisis on the financial system and the economies of Italy and the other euro-area countries. The workshop held on 15 February 2013 was the first opportunity to compare and contrast the research with representatives of the academic world.[2]

The economy of Italy provides an excellent laboratory for us to study the causes and effects of the sovereign debt crisis. Italy did not suffer directly from the consequences of the 2008 financial crisis and its banking system successfully withstood the impact of the crisis. However our country was seriously affected, from the summer of 2011 on, by investors' concerns over the sustainability of the public finances in the euro area because of its high public debt and its poor growth prospects in the medium term.

The European Central Bank's exceptional monetary policy measures, together with the interventions carried out in Italy and the other countries most affected by the turmoil and with the progress made at community level following actions to reinforce the European construction, have allowed a gradual improvement in investors' confidence in the sustainability of the public finances in the euro area and in the soundness of the single currency, which was reflected in a fall in sovereign risk premiums.

However, the situation continues to be fragile. For Italy and other euro-area countries interest rate levels are still well above those before the most acute phase of the sovereign debt crisis. Investors' concerns over the future of monetary union have subsided today in comparison with the time of the greatest tensions but they have not disappeared.

As the Governor of the Bank of Italy has underlined on several occasions, monetary policy can do much, but alone it cannot lead us out of the crisis. National economic policies, which must combine the recovery of competitiveness with balanced public finances and the completion of the reform of European economic governance, will be decisive.

[1]Opening remarks by Daniele Franco, former Managing Director of the Economic Research and International Relations Area of the Bank of Italy. In May 2013 Mr. Franco was appointed State Accountant General at the Ministry for the Economy and Finance.

[2] The following participated as discussants: Nicola Borri (LUISS) and Carlo A. Favero (IGIER, Bocconi) for the first session, Giovanni Ferri (LUMSA) and Alberto F. Pozzolo (Università del Molise) for the second session and Fabio Canova (EUI) and Francesco Nucci (Università di Roma La Sapienza) for the third session.

Introduction

Giuseppe Grande, Stefano Neri and Stefano Siviero[3]

This volume is a collection of the proceedings of the research workshop on the sovereign debt crisis and the euro area held at the Bank of Italy in February 2013. The ten papers trace the structure of the workshop and are divided into three sessions, which examine the main mechanisms that allowed the sovereign debt crisis to have an impact on the economic and financial system of the euro area, with onward transmission from the government bond markets to bank intermediaries and therefore to the real economy.

The first session was dedicated to the impact of the crisis on government bond yields. The main question was to what degree recent trends could be explained on the basis of the economic fundamentals of the single countries, or whether these trends reflect factors of a systemic nature, such as risks related to fears about the reversibility of the euro or contagion effects. Another question is if the increase in public debt observed in recent years could also cause interest rates to rise in countries to date considered safer as regards the sustainability of the public finances.

The second session concentrated on the impact of sovereign debt tensions on the cost and availability of bank credit, both central to the transmission of the financial turmoil to the real economy. Although the ECB’s extraordinary measures and the other economic policy measures adopted at the national or European level have averted potentially disruptive developments, monetary and lending conditions in the euro area countries still differ greatly. One of the main aims of the papers in this session was to evaluate asymmetries in the transmission of monetary policy between the countries most affected by the tensions and those considered to be sounder, with a focus on the demand and supply factors influencing the distribution of credit to the economy.

The final session considered the macroeconomic effects of the sovereign debt crisis; it grouped some papers which aimed to quantify the impact of the tensions not only on credit and bank rates but also on the real economy, at the level of the euro area as a whole and in its main countries. The session also addressed the question of the influence of the cyclical position of the economy, the size of the public debt and financial conditions on the effectiveness of fiscal policy interventions in macroeconomic terms.

Taking the papers together, the following findings emerged. The sharp increase in the dispersion of government bond yields in the euro area observed in the summer of 2010, and the even bigger increase since July 2011, can only be partly explained by the trends of the fiscal and economic fundamentals of some individual countries. The gap between market yields and those consistent with economies’ fundamentals was, in part, caused by fears of the euro area breaking up. The marked increase in public debt observed in all the leading advanced economies in recent years is able, over the long term, to put significant upward pressure on interest rates, if we also consider the effects coming from the components of demand for government bonds which are not much influenced by interest rates (such as, for example, the management of official reserves).

The crisis decisively conditioned banking activity in terms of the cost of credit, availability of loans to households and firms, and profitability. The findings presented in the

[3] Bank of Italy, Economic Research and International Relations Area, Economic Outlook and Monetary Policy Department.

second session showed that the tensions distorted monetary conditions in the leading countries and in the euro area as a whole, in particular causing a significant increase in the cost of credit in the countries most exposed to investors' concerns over the soundness of the public finances. The replies given in the Eurosystem's bank lending survey and the data from the Central Credit Register have made it possible to quantify the impact of the crisis on the behaviour of the Italian banks, highlighting in particular the role of the greater riskiness of borrowers and that of liquidity conditions and bank capital.

Evidence of the real consequences of the sovereign debt crisis is still limited. The findings presented in the third session showed that the impact seems to have been significant not only in the countries directly affected by the tensions but also in the euro area as a whole, causing a fall in industrial production, an increase in unemployment and a depreciation of the euro. The slowdown in economic activity in 2012, even in those euro-area countries with sounder public finances, is consistent with the results of the econometric estimates.

The analyses made of the Italian economy show that fiscal policies would not be very effective in buoying economic activity in a context of high debt-servicing costs. In normal conditions, these policies would be more effective during a period of recession rather than one of expansion.

Conclusions

Eugenio Gaiotti[4]

The research workshop held at the Bank of Italy in February 2013 shows, once again, that it is a vitally important step for central banks to share the findings of their individual analysis and research work with the scientific community. I would therefore like to thank in particular the discussants, who made it possible to put under the scrutiny of the Italian academic world the findings of our research, which was closely linked to economic policy themes. Naturally, I also thank all the authors of the papers and the conference participants.

The crisis of investor confidence in the sustainability of public debt in the euro area had dire consequences for the government bond market, banking activities and macroeconomic trends. The studies presented in this volume focus on these consequences, with different nuances and points of view. At the centre of the Bank's institutional activity there are also other important issues that, for time reasons, could not be addressed during the conference, such as the effects of the crisis on the structural fragilities of Italy and the euro area and the actions taken to advance the European construction.

The research workshop documented in this volume provided useful pointers to understanding the different aspects of the crisis under way. I would like to highlight three general findings, also emphasizing some questions which they provoke and which constitute new challenges for economic analysis.

In the first place, our research work shows that the origins of the sovereign debt crisis are to be found in systemic factors besides country-specific problems and vulnerabilities. As a result – as explained some time ago by the Governor of the Bank of Italy and further repeated in the Bank's Economic Bulletin of last January - the progress made towards resolving the crisis was made possible by the combined effects of all the economic policies, at the national as well as the European level. Although it is now agreed that the tensions became systemic because doubts began to spread about the very survival of monetary union, it is however more difficult to establish precisely the influence of national factors compared with systemic ones in transmitting the crisis. Evaluating the relative contributions of the various factors underlying the crisis – the aim of some of the papers presented here – is indispensable for identifying the best ways of responding to the crisis.

In the second place, the crisis has highlighted the key role of credit supply in today's market economies. As we have often argued in the Economic Bulletin, although monetary policy managed to defuse the extremely insidious tensions on the banks' wholesale funding markets, in the current cyclical situation the supply of funding is still held back by the high risk perceived by the intermediaries as regards the effects of the recession on company balance sheets. The crisis has revealed the need to rethink the way in which, in our macroeconomic models, we interpret and represent the role of the availability of bank credit. It is necessary to investigate in more depth the factors that determine credit conditions and how the latter may influence the effectiveness of monetary policy. In a phase marked by a sharp rise in the riskiness of firms, to what point can monetary policy protect against the effects? What are the most suitable instruments to address this problem?

[4] Bank of Italy, Economic Research and International Relations Area, Head of the Economic Outlook and Monetary Policy Department. In June 2013 Mr. Gaiotti was appointed Managing Director of the Economic Research and International Relations Area.

Our analyses also show that the macroeconomic impact of the sovereign debt crisis was significant not only for the countries directly affected by the tensions, but also those considered to be fiscally sound. This evidence supports what we recently maintained in the Economic Bulletin of January 2013, where we argued that the consequences of the financial tensions which affected, during the year, some euro-area countries and the effects of the necessary consolidation of the public finances were also transmitted to economies up to then considered sound. We must now move forward in our understanding of the channels through which the tensions of the debt crisis were transmitted to economic activity in the whole area and of the close interdependence of the area's economies.

But the most important question that researchers and public authorities are called to answer relates to what economic policies can actually do. The debate on fiscal multipliers and the optimal path of fiscal adjustment is still open. There is some evidence that the multipliers depend on the cyclical conditions; the idea that the multipliers are much higher than in the past is still controversial however. Questions about the role of monetary policy are no less challenging. As the Governor has said on various occasions, monetary policy can do much, but it cannot do everything. The challenge for researchers is to understand exactly how much is "*much*" and how it can be achieved while minimizing any unwanted repercussions. On this subject, which the conference did not touch on, the Bank of Italy's researchers are hard at work.

It should also be noted that many of the assessments that emerge from the papers contained in this volume could perhaps seem fairly obvious today but this was not the case when, only a few months ago, the first research projects on these topics started up and no positions of consensus had yet been reached. These analyses carried out at the Bank of Italy contributed in a fundamental way to the policy debate on the systemic origins of the crisis, its transmission to the euro-area economies, the opportuneness of monetary policy interventions and their design. Together with the contributions from the other Eurosystem national central banks, they paved the way to the conclusions.

There is a fairly urgent need to expand and consolidate these findings and address with determination the new questions mentioned above. This will require collaboration and continuous exchanges between the central bank and the academic world. To quote John Maynard Keynes, "*it is astonishing what foolish things one can temporarily believe if one thinks too long alone*".[5] So we will be organizing an international scientific conference on these topics in September 2013, with the participation of a larger number of researchers from the academic world and other institutions to carry forward the debate and analysis we have been conducting.

[5] Keynes J. M. (1936), *The General Theory of Employment, Interest and Money*, page xxiii, reprinted and published in 1973 for the Royal Economic Society by Macmillan (London, UK) and Cambridge University Press (Cambridge, MA).

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SECTION 1

THE IMPACT OF THE SOVEREIGN DEBT CRISIS ON THE COST OF THE DEBT

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Recent estimates of sovereign risk premia for euro-area countries

Antonio Di Cesare*, Giuseppe Grande*, Michele Manna* and Marco Taboga *

January 2013

Abstract

This paper examines the recent behavior of sovereign interest rates in the euro area, focusing on the 10 year yield spreads relative to Germany for Italy and other euro area countries. Both previous analyses and the new evidence presented in the paper suggest that, in recent months, for several countries the spread has increased to levels that are well above those that could be justified on the basis of fiscal and macroeconomic fundamentals. Among the possible reasons for this gap, the analysis focuses on the perceived risk of a break up of the euro area.

JEL Classification: G12, E43, E62, H63.

Keywords: interest rates, government yield spreads, sovereign risk premia, government debt, financial crisis, sovereign debt crisis, financial contagion, euro break up, convertibility risk.

Paper presented at the Workshop "The Sovereign Debt Crisis and the Euro Area" organized by the Bank of Italy and held in Rome on February 15, 2013. The proceedings are available at: http://www.bancaditalia.it/studiricerche/convegni/atti.

A previous version of the paper was published as Bank of Italy Occasional Paper No. 128

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1. Introduction and executive summary[1]

This paper examines the recent behaviour of sovereign risk premia in a number of euro-area countries, with a particular focus on the 10-year yield spreads relative to Germany.

Using different estimation techniques and explanatory variables, the previous literature finds a statistically and economically significant relationship between sovereign risk premia and country-specific fundamentals such as the debt-to-GDP ratio, the government budget deficit and GDP growth. However, studies on the most recent period – i.e. since the onset of the Greek sovereign debt crisis at the end of 2009 – generally find that the surge in sovereign spreads experienced in several euro-area countries cannot be fully explained by changes in macroeconomic fundamentals.

The analyses presented in this paper – which in some cases are obtained building on previous studies – are broadly consistent with those of the extant literature. Our results suggest that in recent months the spectacular reduction of long-term German sovereign yields (standing at around 1.3 per cent as of end-August 2012) is to a large extent due to safe-haven flows (see Section 4.1). Moreover, for several countries we find that in the most recent period the sovereign spread vis-à-vis the German Bund has risen well above the value consistent with country-specific fiscal and macroeconomic fundamentals (see Sections 4.2-4.5).[2] For Italian government bonds, most estimates of the 10-year spread fall around 200 basis points, as opposed to a market value of almost 450 points (at end-August 2012). Furthermore, large differences between the market spreads and those warranted by fundamentals are also found on shorter maturities (2 and 5 years – see a summary of the estimates in Table 1).

These results are likely due to the fact that the models used so far do not take into account the new risks which have recently emerged in euro-area sovereign debt markets. In fact, several reasons suggest that euro-area sovereign spreads are increasingly affected by investors' concerns of a break-up of the Economic and Monetary Union (EMU – see Section 5). First, the fact that the deviation of sovereign yields from their model-based value is negative for some "core" countries and positive for "non-core" countries likely reflects the expectation that a break-up of the euro would entail an appreciation of the new national currencies for the former countries and a depreciation for the latter (compared with the parities enshrined in the single currency). Second, the divergence between sovereign spreads and their model-based values has emerged in a phase of

[1] The first version of this paper was published in September 2012, as Bank of Italy Occasional Paper No. 128. Most of the analyses presented in the paper refer to the data available at the date of its first publication.

[2] For the sake of conciseness, in this paper the value of the yield spread consistent with fundamentals is in some cases referred to as the 'fair value', as it is sometimes called in the literature.

exceptionally high volatility in financial markets, when the risk of a break-up of the euro is mentioned more and more frequently by market participants.

Other explanations are possible. These include: concerns of a further, significant deterioration of the medium-term fiscal outlook of the weaker sovereigns not captured by the available indicators; a re-pricing of sovereign risk that increases the compensation required by investors for bearing it; difficulties in assessing sovereign risk that may induce investors to make oversimplifying assumptions and take into consideration only pessimistic or worst-case scenarios. More generally, spreads may reflect the interaction between these different factors, with the possible emergence of a negative spiral between rising risk premia, deteriorating public finances, problems with banking systems, and low growth.

In future work we will assess the contribution of these alternative factors. Nonetheless, as already mentioned, the timing of the increase of sovereign yields in fiscally weak countries and the concurrent, spectacular fall of sovereign yields in fiscally sound countries seems to suggest that recent developments in sovereign euro-area debt markets can be largely traced back to concerns of a break-up of the EMU.

Table 1

Estimates of the Italian yield spreads vis-à-vis Germany consistent with fundamentals: Summary of the results (1)

(basis points)

Main determinants of the spread	Frequency of the data	Time horizon		
		2 years	5 years	10 years
Debt-to-GDP ratio	Daily	91	109	122
Debt-to-GDP ratio (nonlinear)	Quarterly	164	203	212
Fiscal/macro indicators (CDS model)	Daily	124	143	155
Fiscal/macro consensus expectations	Monthly	116	215	260
Fiscal/macro indicators (“wake-up call” model)	Monthly	–	–	80-270 (2)
Financial indicators (average value)	Daily	168	193	215
Fiscal/macro consensus expectations and financial indicators	Monthly	182	272	272
Fiscal/macro indicators and financial accounts	Yearly	–	–	112-301 (3)
Fiscal/macro indicators and contagion	Annual	–	–	80-408 (4)
Memo:				
Actual BTP-Bund spread (21 August 2012)	Daily	300	413	410
Actual BTP-Bund spread (June 2012)	Monthly	414	493	449
Actual BTP-Bund spread (2012 Q1)	Quarterly	289	371	382

(1) Unless otherwise stated, daily estimates refer to the value of the spread on 21 August 2012, monthly estimates refer to its average value in June 2012, and quarterly estimates refer to its average value in the first quarter of 2012. – (2) Data as of December 2011. – (3) Average value for 2012, as of early July. – (4) As of mid-July 2012. The lower value refers to a specification based only on fundamentals, the upper value to one including also a proxy for euro-area systemic risks; the difference between the two values cannot be ascribed to country-specific factors.

This paper is organised as follows. The second section provides an overview of recent trends in the long-term interest rates of the euro area. The third section briefly reviews recent studies. The fourth section shows alternative estimates of the values of the yield spreads vis-à-vis Germany consistent with fundamentals for a number of euro-area countries, with a focus on the 10-year maturity. The fifth section presents evidence on the ongoing concerns of a break-up of the euro area and their role in widening the dispersion of interest rates across euro-area countries. The sixth section concludes and highlights some topics for future research.

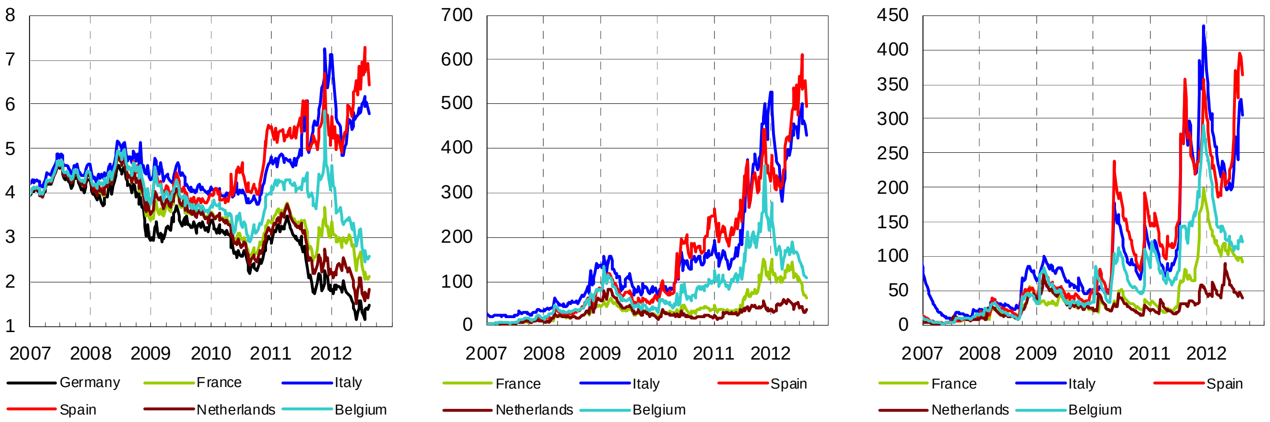
2. The rising dispersion of long-term interest rates within the euro area

Since the onset of the global financial crisis in the summer of 2007 the dispersion of the long-term government bond yields of the main euro-area countries has risen significantly (Figure 1, left-hand panel). In particular, long-term rates have considerably increased in Italy and Spain, while they have declined in Belgium, France and, above all, the Netherlands and Germany.

Figure 1

Ten-year government bonds

Yield *(end-of-week data; percentages)* | Spread vis-à-vis the German Bund *(end-of-week data; basis points)* | Yield volatility *(end-of-week annualized data; basis points)*



Source: based on Bloomberg data.

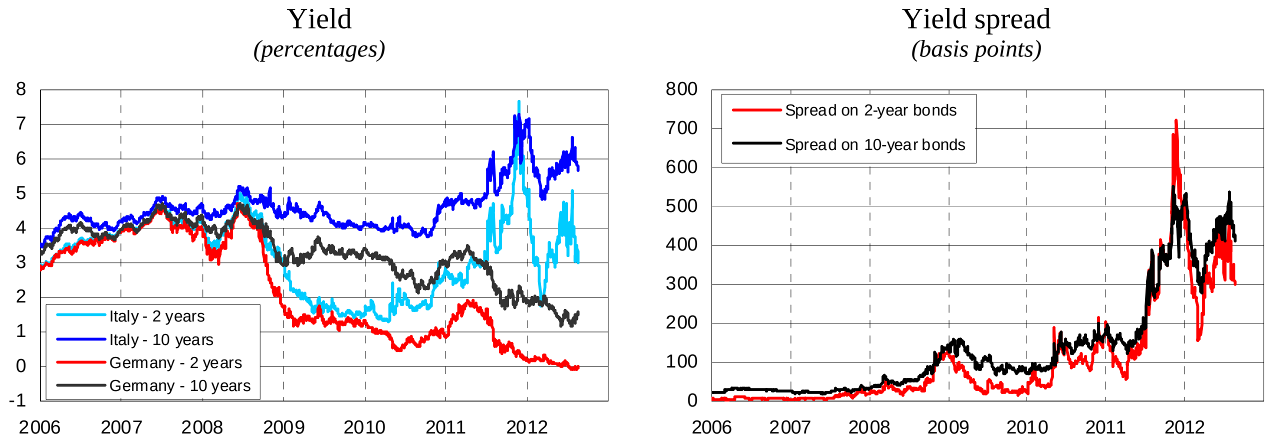
Similarly, yield spreads relative to the German Bund have recorded a significant increase in Spain and Italy, while they have risen much less in the other main euro-area countries (Figure 1, middle panel). Yield volatility has soared across the board, reaching particularly high levels for Spanish and Italian government bonds (Figure 1, right-hand panel).

The dynamics of the spread between Italian and German 10-year sovereign rates has been characterized by three different periods (Figure 2). Between mid-2009 and April 2010, the spread hovered around 85 basis points.

Figure 2

Italian and German 2- and 10-year government bonds

(daily data)



Source: based on Bloomberg data.

Subsequently, after the start of the first wave of sovereign debt tensions in May 2010 and up to June 2011, the spread was still relatively stable, although at a higher level (about 150 basis points on average).

In the third period, starting in July 2011 (after the announcement of the so-called private sector involvement in the second assistance package for Greece), the Italian 10-year yield spread has increased substantially and has become much more volatile. The tensions have involved the entire euro area, leading to a widespread increase in market volatility and to a sharp depreciation of the euro. During this period, the sources of the tensions changed. Until November 2011, the turbulence was concentrated on Italy, as shown by the widening spread of Italian sovereign bonds vis-à-vis other non-core countries, such as Spain. It was fuelled by the deterioration of macroeconomic conditions and political instability in Italy; in November 2011, the Italian 10-year sovereign rate and spread vis-à-vis Germany reached record highs of 7.3 per cent and 5.5 percentage points, respectively. In contrast, in the first half of 2012 the instability was largely driven by the deterioration of macroeconomic conditions in Greece and the difficulties of the Spanish banking sector. In this phase, Italian sovereign rates remained well below the previous peaks, hovering below 6 per cent; in contrast, Spanish yields increased significantly, with the spread between Spain and Italy turning positive (up to 1.2 percentage points in the second half of July 2012).

3. Sovereign risk premia for euro-area countries: recent literature

Ardagna, Caselli and Lane (2007) describe the three main channels through which a worsening of the public finances can affect medium- and long-term yields.[3] First, if the supply of savings is not perfectly elastic, financing the public deficit has to compete for resources with the private sector, causing real interest rates to rise.[4] Second, increases in public debt may cause fears that even sovereign borrowers may default, leading to an increase in the credit risk premia on government bonds. Third, larger public deficits may fuel expectations of inflation or exchange-rate depreciation, with repercussions on interest rates.

Most of the extensive empirical literature on the effects of fiscal imbalances on long-term interest rates does not distinguish among the three aforementioned channels and resorts to reduced-form regressions. Estimates vary greatly from country to country and depending on the method used (see the table in Annex 1, reproduced from Haugh, Ollivaud and Turner, 2009). It is widely agreed that the effects are generally small (see, among others, Balassone, Franco and Giordano, 2004), despite their being larger where the deterioration in the budget balance persists over time. Estimates for the United States indicate that a permanent increase in the debt-to-GDP ratio of 1 percentage point would raise real long-term interest rates by 3 to 5 basis points, while a permanent increase in the budget deficit would produce far larger results. Estimates for European countries, although not uniform, tend to show larger effects.

In recent years, the global financial crisis of 2007-9 and the ensuing sovereign debt crisis in the euro area have spurred a new wave of studies on the relationship between fiscal conditions and long-term interest rates. Unlike previous studies, most of these analyses relax the assumption that public debt is always honoured and allow for the possibility that interest rates on government bonds contain a default risk premium.[5] Attinasi, Checherita and Nickel (2009) estimate a dynamic panel for the 10-year spreads vis-à-vis Germany of ten euro-area countries and find that they are mainly driven by expected public debt and market liquidity, while risk aversion is not significant. Barrios,

[3] See also the box "The effects of the public debt on long-term interest rates" in Banca d'Italia (2010).

[4] As pointed out by Ardagna, Caselli and Lane (2007), it is useful to distinguish between short- and long-run effects. In an economy with a certain degree of short-run nominal stickiness, a weakening in the primary fiscal balance adds to aggregate demand and leads to an increase in nominal and real short-term rates. Insofar as price adjustment is gradual and the weakening in the primary balance is perceived to be persistent, long-term interest rates are also affected. In the longer run, to the extent that fiscal expansion crowds out private investment and results in a lower steady-state capital stock, it will be associated with a higher marginal product of capital and thus a higher real interest rate. For an analysis of the long-run implications of rising public debt for interest rates, see Engen and Hubbard (2005). An important point is made by Krugman (2012), who argues that, in a depressed economy, budget deficits do not compete with the private sector for funds, and hence do not lead to soaring interest rates.

[5] For an earlier analysis of yield spreads in the euro area, see Codogno, Favero and Missale (2005).

Iversen, Lewandowska and Setzer (2009) find a limited impact of deteriorated fiscal balances: on average an increase of 1 percentage point in the budget deficit (vis-à-vis Germany) implies a rise of only 2.4 basis points in the government bond yield spread (vis-à-vis Germany). Bernoth, von Hagen and Schuknecht (2012) show that yield spreads responded significantly to measures of government indebtedness both before and after the start of the EMU. They also find that, since the start of the EMU, markets have paid less attention to government debt levels than they did before; on the contrary, deficits and debt service ratios have been more closely monitored. Bernoth and Erdogan (2012) detect some instability in the pricing of risk between 1999 and the first quarter of 2010 and advocate for the need of time-varying coefficient models in this context.

An increasing number of papers specifically deal with the euro-area sovereign debt crisis and try to analyse its determinants. Borgy, Laubach, Mésonnier and Renne (2011) develop an arbitrage-free affine term structure model to price defaultable sovereign bonds and apply it to a panel of eight euro-area government bond yield curves. They use expected changes in debt-to-GDP ratios as a proxy of fiscal sustainability. According to their estimates (which only include the first period of the sovereign debt crisis), the conditions of the public finances were the major drivers of the increase in spreads that occurred between 2008 and mid-2011.

Other papers find that fundamentals cannot explain a significant portion of the movements of sovereign risk premia registered since the spring of 2010. Aizenman, Hutchison and Jinjarak (2011) estimate a panel model of the premia on 5-year sovereign CDSs. Their sample covers 60 countries (advanced and emerging) from 2005 to 2010 and their explanatory variables include two measures of fiscal laxity (the ratio of government debt to tax revenue and the ratio of the fiscal deficit to tax revenue) and other economic fundamentals. For the euro-area countries most exposed to sovereign tensions (Greece, Ireland, Italy, Portugal and Spain), they find that sovereign credit risk was somewhat underpriced relative to international norms in the period prior to the global financial crisis and substantially overpriced during and after the crisis. According to the authors, this could be due either to mispricing or to pricing based on future fundamentals, incorporating expectations that the fiscal outlook will deteriorate markedly in the euro-area periphery and will pose a high risk of debt restructuring.

Ardagna, Burgi, Cole and Garzarelli (2012) model the 10-year asset swap spreads relative to Germany of France, Italy and Spain as a function of fundamentals (public debt, primary deficit, expected nominal GDP growth and expected 3-month rates) and time dummies. The basic specification of the model uses only the macro fundamental variables and predicts a value of the spreads of about 40, 130 and 200 basis points for France, Italy and Spain, respectively (Table 2).

The higher spreads prevailing in recent years are accounted for by augmenting the model with time dummies that capture changes in the spreads that took place in specific periods and unrelated to country fundamentals. In particular, the very high values reached by sovereign spreads after July 2011 can only be captured by introducing a dummy for that period.

Table 2

Ten-year sovereign spreads vis-à-vis Germany:
A fundamental model augmented by time dummies (1)
(basis points)

	France	Italy	Spain
Actual 10-year spread with respect to Germany (2)	107	431	491
Fitted values of spreads			
1. Fundamentals and EMU dummy (post Jan-99)	43	129	202
2. As sub 1 + Financial Crisis dummy (post Sep-08)	58	180	204
3. As sub 2 + EMU sovereign crisis dummy (post May-10)	50	259	365
4. As sub 3 + PSI dummy (post Jul-11)	140	630	384

Source: Ardagna, Burgi, Cole and Garzarelli (2012).
(1) Fitted values of 10-year asset swap spreads with respect to Germany are obtained from a panel model estimated on monthly data from January 1990. The first estimate shown in the table is based only on macro fundamentals and a post-January 1999 dummy. The other estimates make the additional hypothesis that the events flagged by the time dummies indicated in the table have had a significant impact on sovereign yield spreads. – (2) Data as of 29 March 2012.

On the basis of a panel model of the 10-year interest rates of 21 advanced economies estimated over the period 1980-2010, the IMF finds that the current sovereign spreads with respect to Germany of some euro-area countries are well above what could be justified on the basis of fiscal and other long-term fundamentals (IMF, 2012). For Italy and Spain, in the first half of 2012 the model-based values of the spreads with respect to Germany were around 200 basis points, about half their market value.

Several studies argue that deviations of the spreads from the levels justified by fundamentals are partly due to contagion effects. Metiu (2012) finds that, between January 2008 and February 2012, Italy was hit by contagion from Spain and Portugal while these two countries, in turn, were “importers of risk” from Greece. Moreover, he finds that contagion from Spain to Italy is significant both statistically and economically: more than half of the unexpected increases in the Spanish spread are transmitted to the Italian spread, even if they are unrelated to Italian fundamentals. Similar contagion effects are found by De Santis (2012), who also finds evidence that common upward movements in the spreads are often due to safe haven phenomena that contribute to reducing the yield of the Bund. The results of Caceres, Guzzo and Segoviano (2010) and Beber, Brandt and Kavajecz (2009) are consistent with this finding; in particular, the latter authors argue that safe haven phenomena are often linked to increased demand for very liquid assets.

While a consensus is forming around the idea that contagion is an important determinant of the increase in sovereign risk premia in some countries, the economics profession still lacks a rigorous theoretical framework to understand contagion and identify policy actions that might prevent it. Moving from the empirical observation that contagion has been spreading, mainly but not only, within the euro zone, De Grauwe and Ji (2012) argue that contagion might come from self-fulfilling liquidity crises that propagate within the euro area (but not outside of it) because of the disconnect between monetary and fiscal authorities. The policy implication is that only a better integration of the two policies can prevent contagion.

4. Estimates of the values of the yield spreads vis-à-vis Germany consistent with fundamentals

In this section we present new estimates of the yield spreads vis-à-vis Germany consistent with domestic fundamentals for selected euro-area countries. Some of these estimates are based upon new approaches, while others build upon results of previous studies.

To streamline the exposition, only estimates referring to 10-year spreads are presented in the main text, while Table 1 and the graphs in Annex 2 summarize some of the empirical findings concerning the 2- and 5-year maturities for Italy. The results obtained with shorter maturities are qualitatively similar to those obtained with the 10-year maturity. For the sake of brevity, the coefficient estimates are not shown in the main text and are reported in Annex 3 only for some of the most representative models.

We start by pointing out that analyses of the spreads vis-à-vis Germany should start from an assessment of the level of the German yield. We then present different estimates of the values of sovereign yield spreads consistent with fundamentals, moving from simple models (whose regressors include only the debt-to-GDP ratio) to models that also include other fiscal variables, economic fundamentals and financial risks.

Most of the empirical models are estimated over sample periods that do not extend beyond the first half of 2011. This is due to the fact that since July 2011 the conditions of euro-area government bond markets have rapidly deteriorated (as discussed in Section 2) and have likely been increasingly affected by contagion effects and fears of a break-up of the euro area (as will be discussed in Section 5). In estimating the determinants of the sovereign risk premia, it thus seems preferable to drop the observations that refer to this last phase of exceptional instability, which has led the market prices of the government bonds of the euro area to move away from the levels justified by fundamentals. In any case, the robustness of our results to different sample periods is assessed in a number of ways.[6]

4.1 Safe haven flows and the level of the German 10-year rate

It is commonly argued that, in times of financial stress, safe haven phenomena tend to push German yields below the levels that are consistent with the perceived creditworthiness of the German sovereign. To examine this issue, the left-hand panel of Figure 3 reports the difference between the 10-year German government bond yield and the premium on the 10-year CDS on Germany (red line). This indicator, being broadly equivalent to the return of a credit-risk-free asset, should be comparable to the 10-year Eonia swap rate, which represents a proxy of the risk-free rate (black line). Until April 2010 the difference between the Eonia swap rate and the German government bond-CDS spread (blue line) was in fact very low, with the notable exception of the aftermath of Lehman's default, when it increased substantially reaching one percentage point at the end of 2008 and in early 2009. This spread started to widen again in May 2010, when there was a first phase of strong tensions in euro-area fixed income markets. Subsequently, it increased considerably and since August 2011 it has consistently remained way above the maximum level reached during the global financial crisis of 2007-9. This indicator signals that, over recent months, safe haven effects on 10-year German yields might have been as large as 130 basis points.[7] Similar patterns are also evident for the 2- and 5-year maturities (see Figure A.4 in Annex 2).

The right-hand panel of Figure 3 shows similar indicators for the US, Italy and Spain. For the latter two countries, the differential tends to be more erratic and since mid-March 2012 it has declined considerably (even becoming negative recently), because government bond yields have increased much more than the premia on sovereign CDSs.

[6] In particular, in Section 4.4 we carry out rolling regressions, while in Section 4.5.1 we run the model on shorter-sample periods.

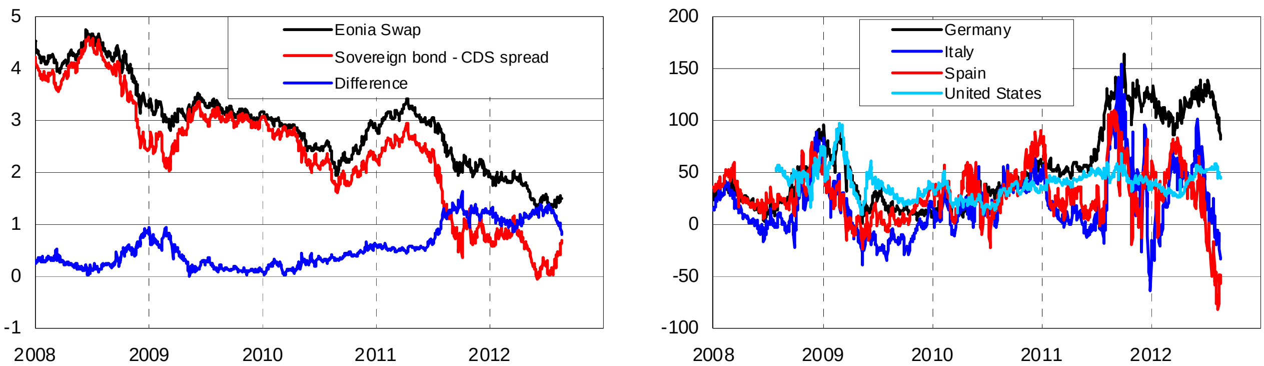
[7] It is worth noting that the CDS premium also reflects counterparty risk, which is the risk that the protection seller is not able to meet its obligation when a default event occurs. The presence of counterparty risk lowers the CDS premium because the protection buyer knows that the protection offered by the contract is not actually full. As the counterparty risk should increase during periods of stress, it seems safe to say that since mid-2011 the premium on the German CDS has actually been lower than it would otherwise have been. Thus, the presence of counterparty risk has probably increased the bond-CDS differential and lowered the difference with the Eonia swap rate. In this respect, therefore, our estimates of the safe haven effects on 10-year German yields are probably conservative.

Figure 3

Differential between the 10-year Eonia swap rate and the 10-year sovereign bond-CDS spread (1)

(daily data)

Germany *(percentages)* | Germany, Italy, Spain and United States *(basis points)*



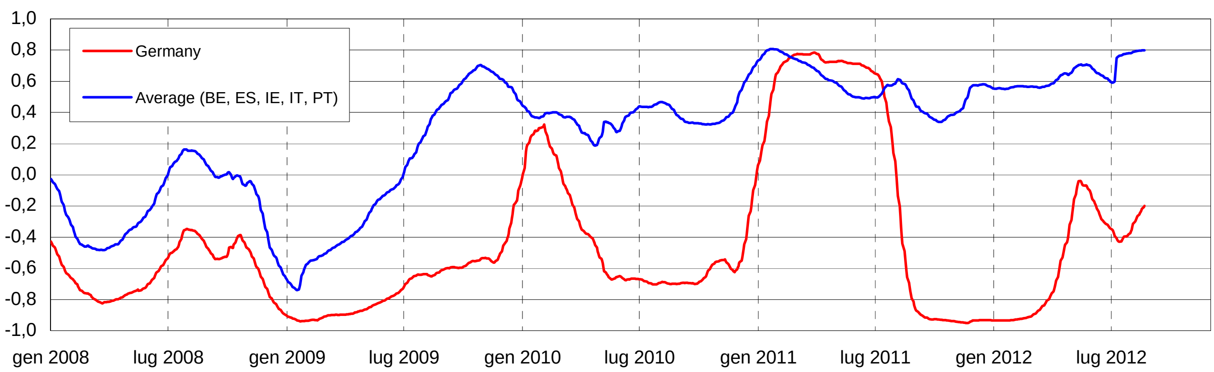
Source: based on Bloomberg data.
(1) The sovereign bond-CDS spread is the difference between the 10-year government bond yield and the premium on the 10-year sovereign CDS.

Further evidence of safe haven phenomena is provided by the co-movement between CDS premia and bond yields. In principle, there should be a positive relationship: a higher CDS spread should be associated with a higher bond yield. While this has been the case for countries with a high debt or deficit (Belgium, Ireland, Italy, Portugal, Spain; Figure 4), for Germany the correlation between the 10-year bond yield and the 10-year CDS spread has been negative both in recent months and over longer time spans. This could be interpreted as evidence of the fact that spikes in risk aversion have triggered both upward revisions of the German sovereign risk premium and safe haven phenomena, but with the effect of the latter on the Bund yield prevailing.[8]

Figure 4

Correlation between 10-year bond yield and 10-year CDS premium

(daily data; 200-day rolling correlations)



Source: based on Bloomberg and Thomson Reuters Datastream data.

[8] There could be also a role for the liquidity risk premium on German bonds that may tend to decrease significantly during periods of financial stress.

4.2 Sovereign spreads and the public debt-to-GDP ratio

A preliminary assessment of the level of the sovereign bond spread vis-à-vis the corresponding German Bund can be obtained from a simple bivariate regression model, where for each country the spread itself is regressed on a constant and the ratio of public debt to GDP (a common indicator of country risk):[9]

$$s_t = \beta_1 + \beta_2 \left(\frac{debt}{GDP} \right)_t + \varepsilon_t , \quad (1)$$

where s_t is the 10-year spread at day t and $(debt/GDP)_t$ is the debt-to-GDP ratio (kept constant within the quarter). The fitted values from this regression are used as an estimate of the fair value of the spread, while the residuals are interpreted as the portion of the spread not explained by country risk. The model is estimated using daily data from October 2000 to June 2011.

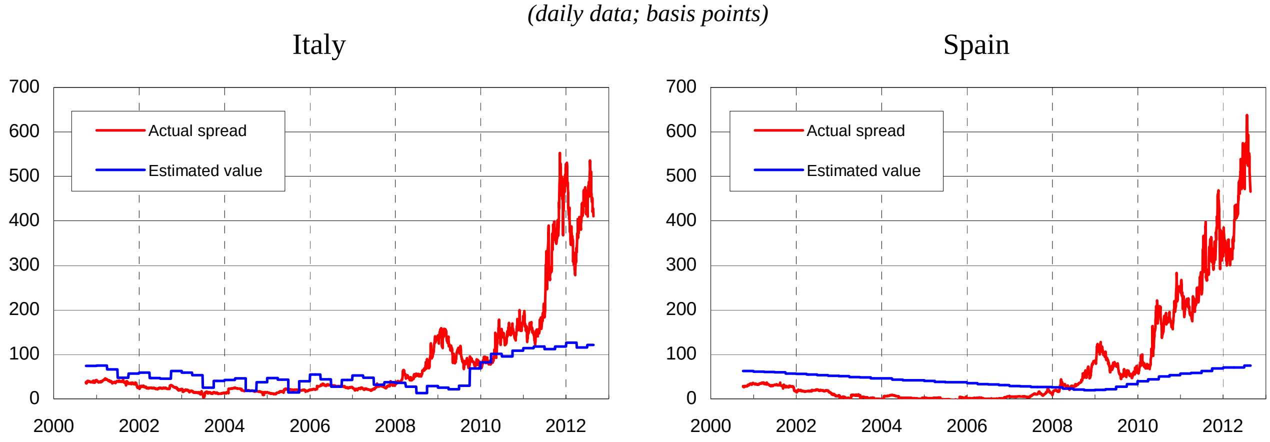
According to this simple indicator, the recent increases of both Spanish and Italian spreads with respect to Germany are much larger than would be justified by the trends in the debt-to-GDP ratios (Figure 5). In particular, the level of the Italian spread consistent with the debt-to-GDP ratio in the second half of 2012 is estimated to be around 120 basis points, against an actual value of the spread of 410 basis points (see Figure A.5 in Annex 2 for the results for 2- and 5-year maturity Italian bonds).

[9] Tests of the null hypothesis of non-stationarity for many of the time series used in this paper are not able to reject it. However, such tests are reckoned to have limited power in datasets of moderate size like ours (Nelson and Plosser, 1982, and Kwiatkowski et al., 1992). Moreover, these tests provide valid inferences only if structural breaks are absent (Perron, 1989) and if errors are reasonably homoskedastic (Kim and Schmidt, 1993). As these conditions are probably not met by our data (especially homoskedasticity), one should be very careful in interpreting the results from unit root tests. In addition, even in the presence of unit roots, OLS estimates such as those presented in this paper remain consistent (actually super-consistent, hence with smaller standard errors) if the series are also cointegrated (Phillips and Durlauf, 1986). Thus, the case remains open only if the variables are I(1) but not cointegrated. Our a priori is that this possibility is economically implausible: theory indicates that the variables under examination are strongly related in an economic equilibrium.

Figure 5

Ten-year sovereign spreads with respect to Germany as a function of debt-to-GDP ratios

(daily data; basis points)



Source: based on Bloomberg and Thomson Reuters Datastream data.

This simple regression model does not take into account possible non-linearities in the relationship between sovereign spreads and public debt-to-GDP ratios. Non-linear effects might be sizable for countries with a high public debt relative to the size of the economy (e.g., Italy, Ireland and Portugal). To account for non-linearities, we follow De Grauwe and Ji (2012). The fair value of the spread is obtained by regressing bond spreads on debt-to-GDP ratios, debt-to-GDP ratios squared, country specific dummies and interactions between these country dummies and debt-to-GDP ratios (simple and squared):

$$s_{i,t} = \beta_1 + \beta_2 \left(\frac{debt}{GDP}\right)_{i,t} + \beta_3 \left[\left(\frac{debt}{GDP}\right)_{i,t}\right]^2 + \beta_4 D_i + \beta_5 D_i \left(\frac{debt}{GDP}\right)_{i,t} + \beta_6 D_i \left[\left(\frac{debt}{GDP}\right)_{i,t}\right]^2 + \varepsilon_{i,t}\,, \quad (2)$$

where $s_{i,t}$ is the spread of country i in quarter t, $(debt/GDP)_{i,t}$ is the debt-to-GDP ratio and D_i is a country dummy.[10]

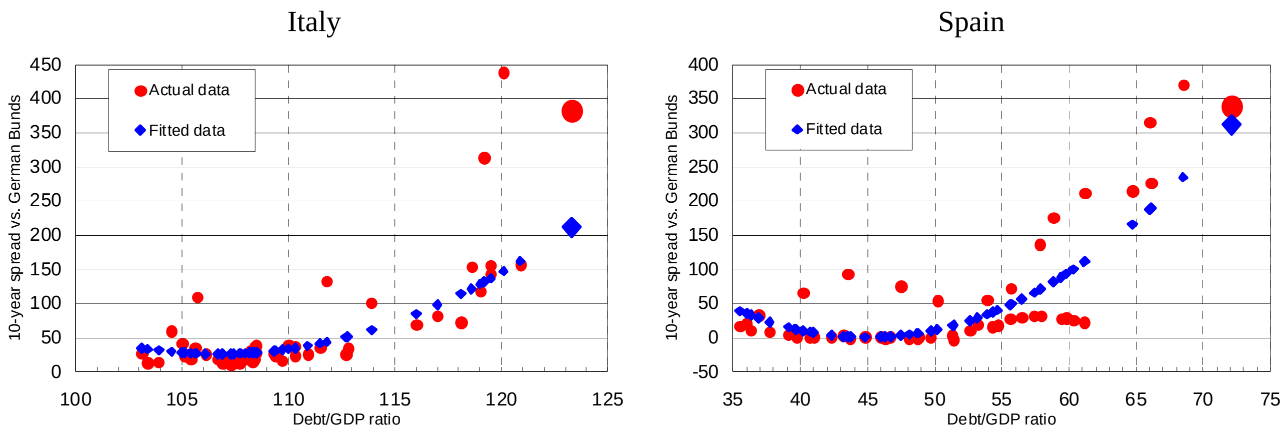
Figure 6 and Figure A.1 in Annex 2 show actual and fitted data for 10-year spreads relative to the German Bund for Belgium, France, Ireland, Italy, Portugal and Spain. The estimates are based on quarterly data of the debt-to-GDP ratio from 2000Q1 to 2011Q2. The data on spreads are quarterly averages. Fitted data from 2011Q3 to 2012Q1 are out-of-sample estimates.

[10] De Grauwe and Ji (2012) also include the ratio of the current account to GDP among the regressors, but its effect on the spread is never statistically significant. They also do not include the interaction terms between the country dummies and the debt-to-GDP ratios (simple and squared), so that the impact of the debt-to-GDP ratio is the same for all the countries in their sample.

Figure 6

Ten-year sovereign spreads with respect to Germany as a non-linear function of debt-to-GDP ratios (1)

(quarterly data; percentages and basis points)



Source: based on Bloomberg and Thomson Reuters Datastream data.
(1) The larger markers denote the latest observations (2012 Q1).

Two results stand out. First, in every country except Belgium, the relationship between the public debt-to-GDP ratio and the sovereign yield spread is non-linear and convex (the larger the debt, the higher the impact on the spread of a one percentage point increase in the debt-to-GDP ratio).[11] Second, in the first quarter of 2012 (the latest available data) the actual level of the spread is much higher than the predicted value in every country except Ireland. In Italy, the fair value of the spread is equal to about 210 basis points, as against an observed value of 380 basis points (see Figure A.6 in Annex 2 for analogous results for 2- and 5-year maturities).

4.3 Sovereign spreads, fiscal sustainability indicators and other fundamentals

Another estimate of the fair value of the sovereign spreads takes into account both fiscal sustainability and macroeconomic indicators and uses some empirical results by Aizenman, Hutchison and Jinjarak (2011). These authors estimate equilibrium sovereign CDS premia as a function of the current values of fiscal sustainability indicators (such as the ratio of public debt to GDP or the ratio of public debt to the realized tax collection) and other fundamental variables (such as inflation and the ratio of total foreign liabilities to GDP). For the euro-area countries most exposed to the tensions on government bond markets, Aizenman, Hutchison and Jinjarak (2011) calculate the ratios between the actual and the predicted values of the sovereign CDS premia for the years 2008-10. We use these ratios to get an estimate of the fair values of the 10-year yield spreads

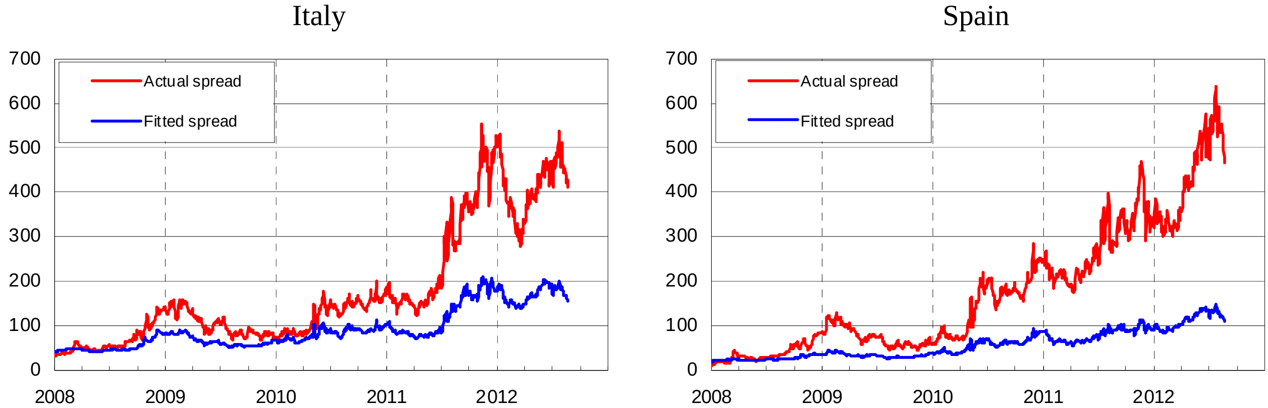
[11] In the case of Belgium, the atypical concave pattern of the fitted curve is due to the fact that in the last few quarters Belgian spreads have recorded historically high levels notwithstanding the debt-to-GDP ratio being well below its historical maxima.

with respect to Germany for Ireland, Italy, Portugal and Spain.[12] Figure 7 and Figure A.2 in Annex 2 show these estimates. Since 2012, the fitted values of the sovereign spreads with respect to Germany have hovered around 390, 180, 290 and 110 basis points for Ireland, Italy, Portugal and Spain respectively (see Figure A.7 in Annex 2 for the results for 2- and 5-year maturity Italian bonds).

Figure 7

Estimates of the 10-year sovereign spreads with respect to Germany based on the results reported by Ainzenman et al. (2011) for sovereign CDSs (1)

(daily data; basis points)



(1) Fitted values are generated on the basis of Ainzenman, Hutchison and Jinjarak (2011)'s estimates of the value of the premia on sovereign 5-year CDSs that are consistent with current fundamentals.

We then take an alternative approach, in which we estimate a model for the 10-year government bond yields of Italy and Germany and then compute the model-implied value of the spread as the difference of the fitted values of the yields. To better account for the forward-looking nature of interest rates, we use the monthly forecasts of yearly macroeconomic variables provided by Consensus Economics (based on a survey of professional forecasters) as proxies for fundamentals. For the Italian and German interest rates,[13] we estimate the following equation:

$$r_t = \alpha + \beta' \overline{EXPFUND_t} + \varepsilon_t \,, \tag{3}$$

[12] We use a three-step procedure. First, for each country the relationship between sovereign bond spreads and premia on sovereign CDSs (both calculated with respect to Germany) is derived through a linear regression estimated on daily data for the period 2008-10. Second, for each country an estimate of the level of CDS premia consistent with fundamentals ("fundamental-adjusted" CDS premia) is obtained by applying the correction terms reported in Table 4 by Aizenman, Hutchison and Jinjarak (2011) to the actual values of the CDS premia. To err on the side of caution, we use the lowest estimate of the correction term for Germany and the highest for the other countries; moreover, in order to smooth the time variation, we use the average values of the correction terms over the three-year period 2008-10. In the third and last step, the fundamental-adjusted bond spreads are computed by using the equation estimated in the first step and replacing the actual values of the CDS premia with their fundamental-adjusted values.

[13] Long time series of consensus forecasts data are only available for G7 countries.

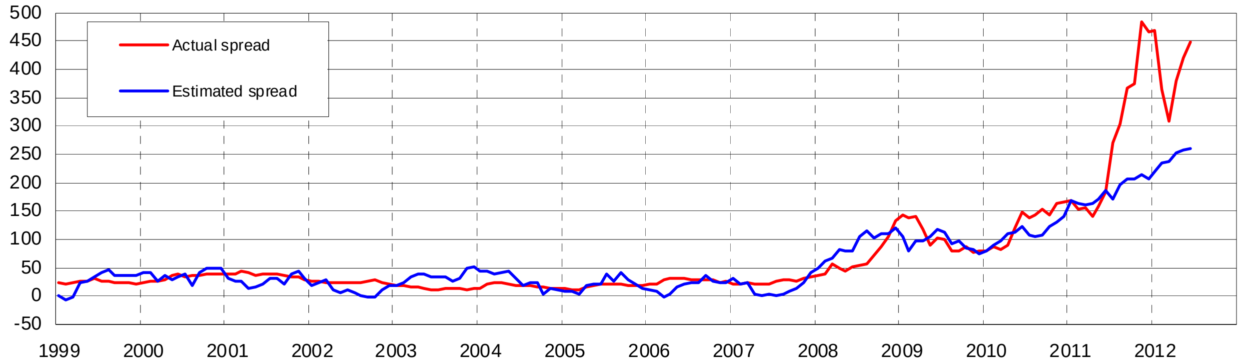
where r_t is the nominal interest rate and $\overline{EXPFUND_t}$ is a vector of variables including the 12-month-ahead forecasts at month t of one fiscal fundamental (the budget balance-to-GDP ratio) and a stream of other macroeconomic variables (three-month interest rates, GDP growth rate, consumer price inflation, unemployment rate and the current account-to-GDP ratio).[14] Regressions are estimated over the period January 1999-June 2011.

At mid-2012 the estimated value of the 10-year Italian spread with respect to Germany was equal to 260 basis points, almost 2 percentage points lower than its actual value (Figure 8). For the 2- and 5-year maturities, the gaps between the actual and estimated values of the spread were even higher (around 3 percentage points; see Figure A.8 in Annex 2).

Figure 8

Italian 10-year sovereign spread with respect to Germany as a function of consensus expectations on fiscal and other economic fundamentals (1)

(monthly data; basis points)



Source: based on Bloomberg, Thomson Reuters Datastream and Consensus Forecasts data.
(1) The estimated spread is the difference between the fitted values of the Italian and German interest rates. Interest rates are modelled as a function of the expected deficit/GDP ratio over the next 12 months and the 12-month-ahead forecasts of other macroeconomic variables (expected three-month interest rates, GDP growth rate, consumer price inflation, unemployment rate and the current account/GDP ratio). Since July 2011 the estimated spread is based on out-of-sample forecasts.

A possible weakness of our results is that the models used so far ignore the possibility that since the onset of the Greek crisis in November 2009 sovereign risk premia within the euro area may have become much more sensitive to fundamentals. This "wake-up call" hypothesis is examined by Giordano, Pericoli and Tommasino (2012), who estimate the following panel model of the 10-year spreads with respect to Germany:

$$s_{it} = \alpha_{i0} + \alpha_{i1} s_{it-1} + \beta_0 Z_{it} + \beta_1 F_t + \gamma_0 D_t + \gamma_1 D_t Z_{it} + \gamma_2 D_t F_t + \varepsilon_{it}, \tag{4}$$

[14] Rolling 12-months-ahead forecasts are computed as a weighted average of the forecasts for the current and next calendar years, in which the weights are given by the fractions of the two calendar years included in the computation window.

where D_t is a dummy variable equal to one after the outbreak of the Greek crisis in October 2009, F_t is the VIX index (regarded as a measure of global risk aversion) and $Z_{i,t}$ includes country-specific variables, such as GDP growth and the ratios of public debt, private debt and the current account to GDP (all these ratios are computed as differences with respect to those of Germany). The dataset covers nine euro-area countries (Austria, Belgium, Finland, France, Ireland, Italy, the Netherlands, Portugal and Spain) and runs from January 2000 to December 2011. Giordano, Pericoli and Tommasino (2012) find that after October 2009 financial market participants became more responsive to country-specific fundamentals than before (with countries with sounder fiscal conditions and better external positions benefiting from lower spread levels). However, even using this "wake-up-call" model, the unexplained portions of the actual yield spreads with respect to Germany remain large. In the case of Italy, the predicted value of the 10-year spread with respect to Germany ranges between 80 and 270 basis points (depending on whether investors' average sensitivity to country-specific factors is set to its pre- or post-Greek-crisis level).

4.4 Financial factors

Besides economic and fiscal fundamentals, sovereign risk premia may be affected by risks stemming from financial markets. Three factors can be singled out: 1) the surge in sovereign spread volatility has reportedly discouraged investors from holding the government bonds of some euro-area countries; 2) sovereign spreads have also been affected by strains in domestic banking systems; 3) the recent wave of sovereign debt rating downgrades might also have contributed to widen government bond spreads, due to the pervasive role of ratings in the financial industry.[15]

A preliminary assessment of the impact of these three financial factors on recent trends in euro-area sovereign yield spreads can be obtained from simple bivariate regression models, where the spread is regressed on a constant and an indicator of financial risks:

$$s_t = \beta_1 + \beta_2\, financial\_indicator_t + \varepsilon_t\,, \tag{5}$$

where s_t is the spread at time t of the country considered and $financial\_indicator_t$ is the given indicator of financial risks. As for equation (1), the fitted values from this regression are interpreted as an estimate of the fair value of the spread, while the residuals are interpreted as the portion of the spread not explained by country risk.

[15] It should be borne in mind that using financial market variables as explanatory variables of sovereign spreads may entail serious reverse causality issues. This could be particularly relevant over the last year, when developments in euro-area government bond markets have been a source of systemic risks.

For each of the six countries considered in Section 4.2, we use three different proxies of country-level financial risks, giving rise to three alternative estimates:

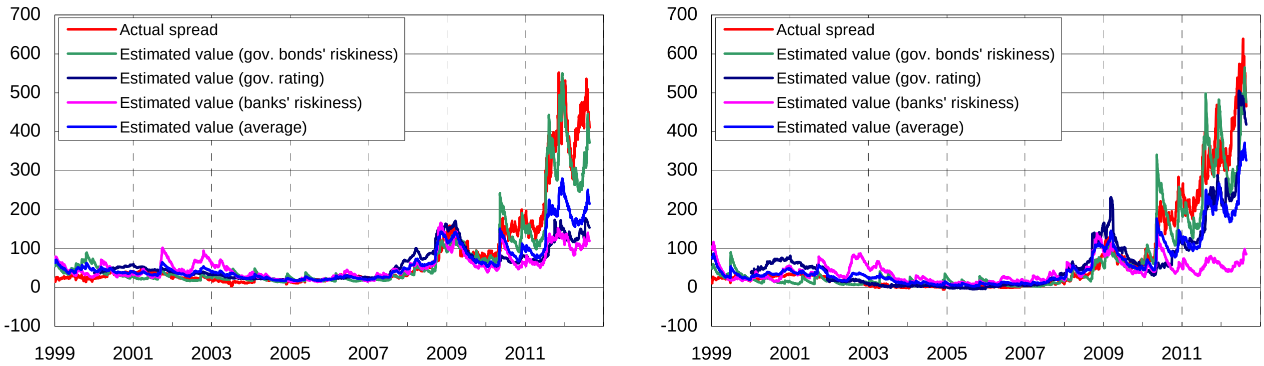
- *volatility of the sovereign spread*: this is motivated by the observation that the risk premium required to hold a given bond could be proportional to its financial riskiness, as measured by its price volatility in excess of the volatility of a safe bond. The indicator is computed as an exponentially weighted moving average (EWMA) of squared day-on-day changes in the 10-year government bond spread;[16]
- *volatility of bank stocks*: given investors' concerns about banks' conditions in the euro area, this measure takes into account the possibility that the sovereign spread of a given country might reflect the vulnerability of its banking sector, as proxied by the stock price volatility of the major banks. The indicator is calculated by applying the EWMA methodology to country indices of bank share prices;[17]
- *spread on corporate bonds having the same rating*: under the assumption that credit ratings are reliable measures of credit risk, there should be a close relationship between the spreads on sovereign and corporate bonds having the same rating. For each sovereign, this indicator is computed from the Merrill Lynch index of the corporate bonds having the same rating as the sovereign's government bonds.

Figure 9

Ten-year sovereign spreads with respect to Germany as a function of financial indicators of country risk

(daily data; basis points)

Italy | Spain



Source: based on Bloomberg and Thomson Reuters Datastream data.

We run equation (5) on daily data from January 1999 to June 2011. The fitted values from these regressions are plotted in Figure 9 for Italy and Spain and Figure A.3 in Annex 2 for the other

[16] We used the standard RiskMetrics framework for daily data, assuming null mean and a decay factor equal to 0.96.

[17] We used the Datastream indices for national banking sectors.

four countries (see Figure A.9 in Annex 2 for the results for 2- and 5- year maturity Italian bonds). The figures also show the series obtained by averaging the estimates from the three models.

The main results that emerge from the analysis can be summarized as follows:

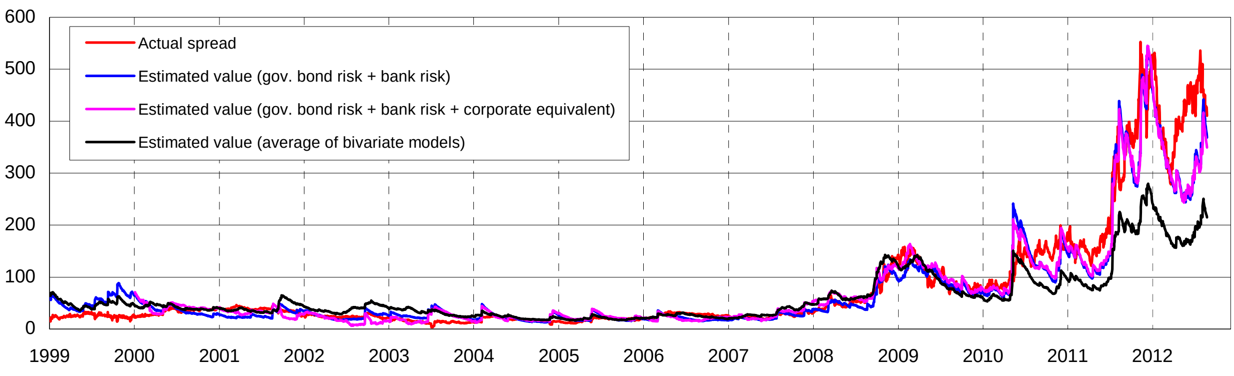
- All the proxies of country risk have significant explanatory power, particularly the volatility of the sovereign spread; the latter finding signals that financial risks stemming from short-term bond price volatility is one of the main drivers of sovereign spreads;
- Since the summer of 2011, there has been an increasing gap between the market values of sovereign spreads and their model-based values; this is true for all countries in the panel, albeit to different extents;
- Italy seems to be the most severely penalized country. On 21 August 2012 (the last day in our sample) the spread stood at 410 basis points, against an average estimated value of 215 points. On the same day, the most conservative estimate was about 370 basis points (based on the volatility of the sovereign spread), while the other estimates stood at 120 and 150 basis points (based on the volatility of bank stocks and the spread of equivalent corporate bonds, respectively).

As a robustness check, we also run multiple regressions: one including all three proxies of risk and one including only the two volatility variables. These regressions also provide evidence of a gap between the actual and the model-based value of the spreads. In particular, despite producing a remarkably good fit of the dynamics of the Italian spread until the end of March 2012 they fail to explain the surge that occurred subsequently (Figure 10).

Figure 10

Multiple regressions of the Italian 10-year sovereign spread with respect to Germany on financial indicators of country risk

(daily data; basis points)



Source: based on Bloomberg and Thomson Reuters Datastream data.

A possible concern about these estimates is that they do not take into account the possible time-variation in investors' risk aversion and the price of risk. One may interpret the explanatory variables as proxies of the quantity of risk and their regression coefficients as the price of risk (their product being the risk premium). Along these lines, it is possible to estimate how the price of risk evolved through time by running rolling regressions over shorter sub-samples. Using 2-year rolling windows, we find that in 2012 the estimated prices of risk are very close to their sample averages; even with time-varying coefficients, the estimated value of the spread is only a few basis points above the level found with the baseline model described above. Furthermore, estimates of the fair value remain virtually unchanged by adding to the regressors the level of the VIX index, which is sometimes regarded as a proxy of risk aversion.[18]

4.5 Financial factors and other fundamentals

Sovereign risk premia are likely to be a function of both financial factors and economic and fiscal fundamentals. In this section we follow two different approaches that try to take all these determinants into account.

4.5.1 Indicators of financial risks and other fundamentals

In Section 4.3 we have modelled Italian spreads with respect to Germany as a function of the consensus forecasts of macroeconomic variables. We now augment model (3) to include the three indicators of financial risks described in Section 4.4. For the Italian and German interest rates, we run the following equation:

$$r_t = \alpha + \beta' \overline{EXPFUND_t} + \gamma' \overline{FINFACT}_t + \varepsilon_t \,, \tag{6}$$

where r_t is the nominal interest rate of the country considered, $\overline{EXPFUND_t}$ is the vector of 12-month-ahead forecasts of fundamentals described in Section 4.3 and $\overline{FINFACT}_t$ is a vector including the volatility of r_t, the volatility of the share prices of the banks of the given country, and the yield on corporate bonds having the same rating as the sovereign of the given country. In an extended version of (6) the regressors also include the current level of the public debt/GDP ratio

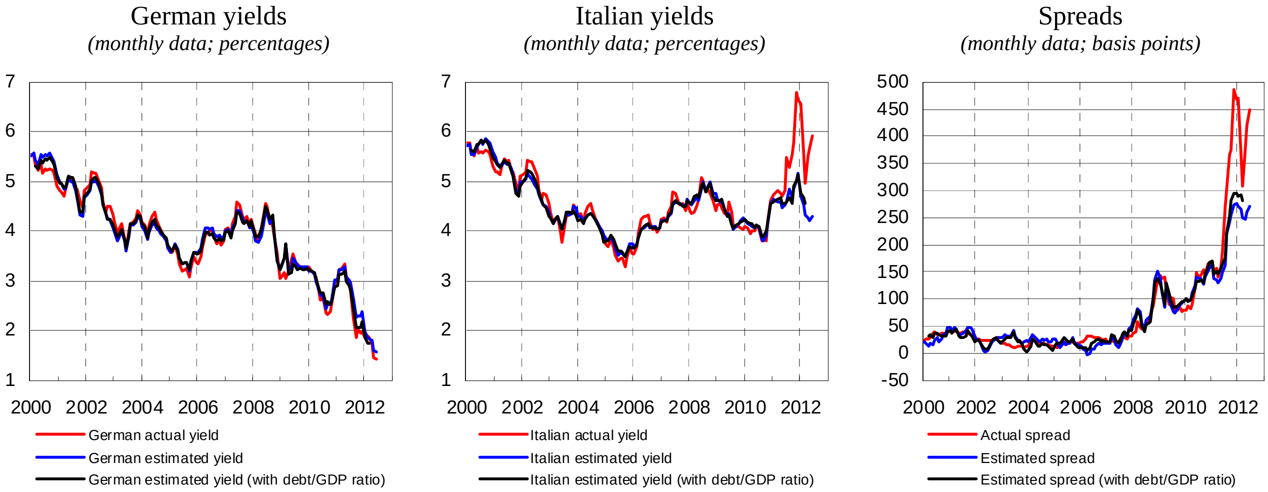
[18] These results apparently provide little support for the hypothesis that the compensation required by investors to bear sovereign risk in the euro area has significantly increased since the second quarter of 2010. However, our proxies of risk (in particular, corporate spreads) are themselves affected by the price of risk and thus may already reflect, at least in part, its possible changes over time.

(for which no consensus forecast is available), which might be an important factor for Italian sovereign risk premia. Regressions are estimated over the period January 2000-June 2011.[19]

Results for the 10-year maturity are shown in Figure 11. While the equation tracks the German 10-year yield quite well, the Italian 10-year yield turns out to be significantly higher than the fitted value (by 160 basis points at mid-2012). The fitted value of the spread at June 2012 stands at 270 basis points, almost 180 basis points lower than its actual value. For the 2- and 5-year maturities, the gaps between actual and fitted values are even larger for the Italian interest rates (about 220-230 basis points), while they are nil for the German ones (see Figure A.10 in Annex 2).

Figure 11

Ten-year government bond yields and spreads as a function of fundamentals and financial factors



Source: based on Bloomberg, Thomson Reuters Datastream and Consensus Forecast data.
(1) Yields are modelled as a function of three financial risk indicators (yield volatility, bank share price volatility and yield of corporate bonds with the same rating as the sovereign), the expected deficit/GDP ratio over the next 12 months and the 12-month-ahead forecasts of other macroeconomic variables (expected three-month interest rates, GDP growth rate, consumer price inflation, unemployment rate and the current account/GDP ratio). An extended specification also includes the current level of the debt-to-GDP ratio among the regressors. Since July 2011 fitted values are based on out-of-sample forecasts.

4.5.2 Fundamentals and the financial position of the main sectors of the economy

Grande, Masciantonio and Tiseno (2012) explain sovereign yields in terms of fundamentals and the financial position of the main sectors of the economy. For the 10-year interest rates of 18 major advanced countries, the authors estimate the following panel model:

$$r_{i,t} = \alpha_i + \beta' F_{i,t} + \gamma' B_{i,t} + \varepsilon_{i,t} , \quad (7)$$

[19] The stability of the econometric estimates is assessed by running the model on two shorter sample periods: (*i*) from 2007 to June 2011, in order to exclude the first half of the 2000s (a period of very low sovereign risk premia) from the sample period; (*ii*) from 2007 to November 2012, in order to assess how much the results are affected by the recent wave of massive instability. The results are remarkably stable across the three sample periods: the coefficients do not change sign and vary in magnitude and significance levels only for a restricted number of variables (see Annex 3).

where $F_{i,t}$ is a vector of economic and fiscal variables of country i at time t and $B_{i,t}$ is a vector of variables taken from the country's financial accounts. The latter includes the net asset holdings (defined as the balance between the stock of financial assets and that of financial liabilities) of the sectors of the economy that are the main providers or users of savings – households, non-financial corporations, the public sector, and the foreign sector. As for the fundamentals, the authors use two different specifications with and without rating dummies.[20] Rating dummies are based on end-of-year data and refer to the best rating across the three major rating agencies.[21] The model is estimated using yearly averages over the period 1995-2010 and is used to predict the average yields for 2011 and 2012.

Table 3 reports out-of-sample predictions for Italian and German interest rates for 2011 and 2012, and their actual values. Both the models with and without sovereign ratings are included. For 2011 the predicted value of the 10-year BTP-Bund spread ranges between 150 and 210 basis points, compared with an actual level of 280. The two-notch decline in Italian government bonds' best rating occurred in the last months of 2011 accounts for an increase in the fitted value of the Italian 10-year rate of more than 50 basis points.

For 2012, three different scenarios are envisaged depending on the hypotheses about the net asset holdings of households and non residents, and the other financial account variables. In the middle scenario, dubbed "Stabilization", the net asset holdings are assumed to remain broadly unchanged at the levels reached at the end of 2011. In that case, the predicted value of the 10-year BTP-Bund spread ranges between about 160 and 280 basis points, compared with an average level of the spread of nearly 400 basis points in the first half of 2012. The other two scenarios, dubbed "Recovery" and "Deterioration", assume that the changes in the net asset holdings observed in 2011 will revert or occur again in 2012, respectively. Fitted values range between about 110 and 230 basis points in the "Recovery" scenario and between about 190 and 300 basis points in the "Deterioration" scenario.

[20] With regard to fundamentals, the explanatory variables include real short-term rates, inflation, the average residual maturity of marketable public debt, and the ratio of public debt to GDP.

[21] The rating dummies are defined as follows: (*i*) for each country, the end-of-year ratings provided by the three major rating agencies are converted into a common numerical scale; (*ii*) each country is given the rating score which corresponds to the highest level of creditworthiness across the three rating agencies.

Table 3

Interest rates on Italian and German 10-year government bonds:
A model of fundamentals, credit ratings, and capital availability (1)

(average data; percentages and basis points)

Year	Scenario		Model without ratings (2)		Model with ratings (2) (3)		Observed rates (4)	
			Italy	Germany	Italy	Germany	Italy	Germany
2011	Actual data	Yield	4.19	2.72	4.81	2.68	5.42	2.61
		S.E.	(0.47)	(0.29)	(0.45)	(0.31)		
		Spread	*147*		*213*		*281*	
2012	Recovery (back at end-2010 levels)	Yield	3.54	2.42	4.71	2.38		
		S.E.	(0.54)	(0.29)	(0.58)	(0.31)		
		Spread	*112*		*233*			
	Stabilization (as at end-2011)	Yield	3.92	2.34	5.06	2.31	5.70	1.71
		S.E.	(0.50)	(0.28)	(0.54)	(0.30)		
		Spread	*158*		*275*		*399*	
	Deterioration (2011 trends continue in 2012)	Yield	4.21	2.35	5.33	2.32		
		S.E.	(0.48)	(0.28)	(0.52)	(0.30)		
		Spread	*186*		*301*			

Source: based on Grande, Masciantonio and Tiseno (2012).
(1) Yields and standard errors (S.E.) are in percentages, while spreads are in basis points. Fitted values of 10-year yields are out-of-sample predictions obtained by a panel model estimated on annual data from 1995 to 2010 for a sample of 18 major advanced countries. – (2) Fundamentals include real short-term rates, inflation, the debt-to-GDP ratio, and the average residual maturity of the public debt. – (3) Rating dummies are based on end-of-year data and refer to the best rating across the three major rating agencies. For 2012, best rating as of mid-July 2012. – (4) For 2012, average values from January to early July.

5. The perceived risk of a break-up of the euro area

The existence of large and persistent gaps between the actual levels of interest rates and what could be justified on the basis of fiscal and other macroeconomic fundamentals for several countries suggests that some common new risk factor is currently at play in the euro area.

One factor driving these gaps may be the risk of a break-up of the euro area and its systemic consequences. Doubts about the irreversibility of the euro led market participants to start guessing about the likelihood and consequences of a euro break-up and about investors' willingness to bear that risk. Fears of the reversibility of the euro can thus explain the current high dispersion of interest rates within the euro area and be a major source of uncertainty and systemic risk.

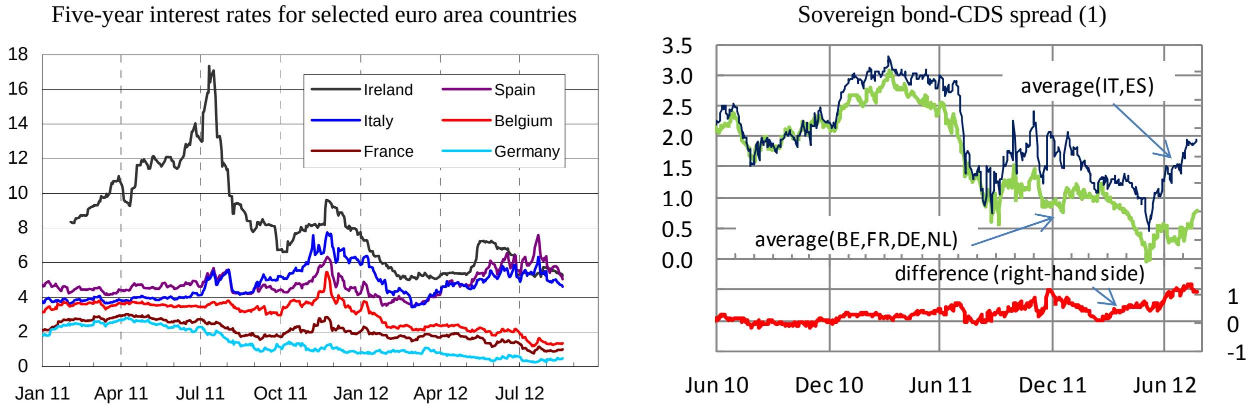
There is sound evidence that over the last year euro-area government bond markets have been increasingly affected by stories of a break-up of the euro area. Besides the abnormal levels reached in the euro area by sovereign yields and yield volatilities since the second half of 2011 (see

Section 2), some recent discontinuities in the patterns of sovereign yields call for attention. Until early March 2012, Belgian interest rates had oscillations rather similar to those of Italian and Spanish interest rates, likely due to changes in risk premia related to investors' assessments of the sustainability of the public debt in Belgium (Figure 12, left-hand panel). Subsequently, there has been a growing divergence between Belgian rates and Italian and Spanish rates, with the former becoming closer to French and German rates. This suggests a clustering of interest rates along geo-economic patterns that were discernible before the introduction of the single currency and is consistent with a progressive loss of confidence in the integrity of the euro area.

Figure 12

Interest rates within the euro area and sovereign bond-CDS spread

(daily data; percentages)



Source: based on Bloomberg and Thomson Reuters Datastream data.
(1) Average values of the sovereign bond-CDS spread for two groups of euro-area countries (Italy and Spain, on the one hand, and Belgium, France, Germany and the Netherlands, on the other). The lower panel of the graph shows the difference between the two average spreads. The sovereign bond-CDS spread is the difference between the 10-year government bond yield and the premium on the 10-year sovereign CDS.

A structural break can be observed also in the sovereign bond-CDS spread, i.e. the differential between government bond yields and the premia on sovereign CDSs. As mentioned in Section 4.1, this spread contains the risk-free rate and premia on risk factors other than sovereign default (e.g., liquidity risk). The right-hand panel of Figure 12 shows the average values of the spread for two groups of euro-area countries: the two main countries most exposed to tensions (Italy and Spain) and the other four main countries (Belgium, France, Germany and the Netherlands). The lower half of the graph also shows the difference between the two average spreads. Since July 2011 this spread has become much more volatile and dispersed across euro-area countries. More importantly, since mid-March 2012 the gap between the average spreads of the two groups of countries has consistently increased, because over the whole period bond yields have risen much more than sovereign CDS premia in Italy and Spain, and they have declined much more than sovereign CDS premia in the other main euro-area countries. The formation of such a wide gap between the average spreads of the two group of countries is consistent with the hypothesis that over recent months the huge increase in the dispersion of interest rates across euro-area countries has been due to a new common factor, namely the risk of a break-up of the euro area.

A scenario of some countries leaving the euro area has been gathering momentum for some time among financial market participants. In June 2012, the Swiss bank UBS conducted a poll of 80 central bank reserve managers who collectively control more than 8 trillion US dollars. The respondents said that a break-up of the euro area was the greatest risk to the global economy over the next 12 months (Financial Times, 2012). Nearly three quarters of them said at least one country would leave the euro area within five years. Of those, roughly a quarter said that more than one country would drop the euro.

Concerns about a possible break-up of the euro area have also become widespread in the non-financial media and the online world. The volume of searches of "euro break-up" or similar keywords using Google peaked in early December 2011 and in May and June 2012 (Figure 13, left-hand panel). As unlikely as it may be, the possibility that the interest rates of euro-area countries have been including a convertibility risk premium has recently been mentioned by the President of the ECB (Draghi, 2012).[22]

Our own quantitative analysis provides some indications that since July 2011 euro break-up risks have been a main driver of the instability of euro-area government bond markets (see also Favero, 2012, for econometric evidence of non-default components linked to break-up risks). The very fact that the deviation of sovereign yields from their estimated value has recently tended to be negative for Germany and positive for "non-core" countries likely reflects the expectation that a break-up of the euro area would entail an appreciation of the new German currency and a depreciation of the currencies of "non-core" countries (compared with the parities enshrined in the single currency). This explanation is supported by the comparison of the Google-based indicator of euro break-up risks shown in the left-hand panel of Figure 13 with the residuals from the interest rate models with macroeconomic variables and financial factors presented in Section 4.5.1 (Figure 13, right-hand panel). Model residuals are a measure of the gap between the actual level of the

[22] A straightforward way to check for the presence of a convertibility risk premium is the comparison of the yields of Italian government bonds denominated in euro and the yields of similar government bonds denominated in, say, US dollars, which are immune of the risk of redenomination. However, the two types of bonds typically do not differ only for the currency of denomination but also for a number of other factors (e.g., law to which the issuance is subject, eligibility towards central bank refinancing, liquidity of the underlying market) that may make the use of this approach extremely difficult in practice.

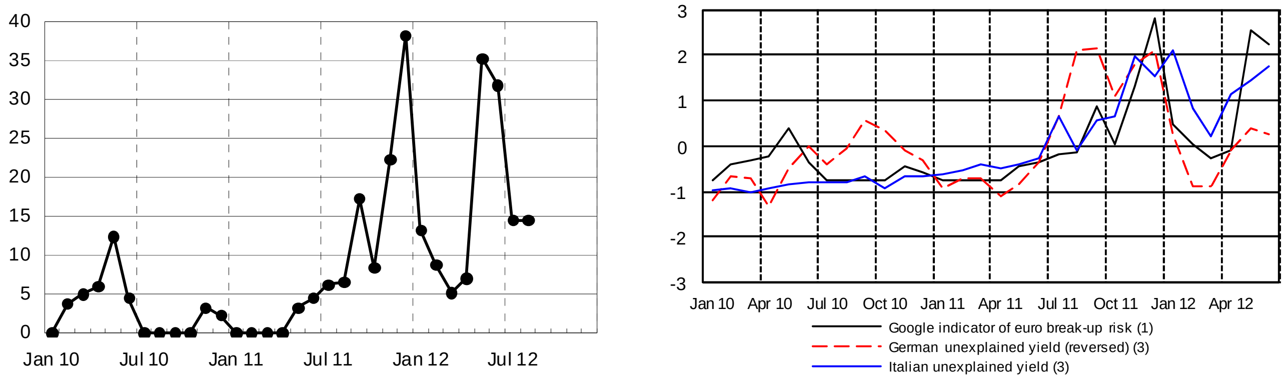
interest rate and the level that would be justified by fundamentals. Since the second half of 2011 the positive correlation between the euro break-up indicator and the portion of the Italian 10-year interest rate not justified by fundamentals is striking. For the German 10-year rate, the correlation with the euro break-up indicator is remarkable as well, although it is slightly lower than for the Italian rate (over the period January 2010-June 2012, the correlation is 0.77 and 0.56 for the Italian and the German unexplained rate, respectively).

Figure 13

Euro break-up risk and the gap between market yields and the yields consistent with fundamentals

Frequency of internet searches of euro break-up keywords (1)

Unexplained portions of sovereign German and Italian 10-year yields and euro break-up risk (2)



Source: based on Bloomberg, Thomson Reuters Datastream, Consensus Forecasts and Google data.
(1) Monthly data. Index of search volume of euro break-up keywords ("end of euro", "end of the euro", "euro break-up", "euro break up", "euro breakup" and "euro exit") typed into Google's web search engine. Data is monthly averages of weekly data: weekly data is an index which varies between 0 and 100 and is equal to 100 when the ratio "number of query X"/"total number of queries" reaches its maximum value over the period for which data are extracted. The data extraction period was January 2004-August 2012. Data downloaded on 23 August 2012. – (2) The time series are normalized to have zero mean and unit variance. – (3) Difference between the actual and fitted values of the 10-year government bond yield. Fitted values are obtained by a model that controls for both macroeconomic fundamentals and financial factors. Since July 2011 fitted values are out-of-sample forecasts.

Indicators of a generalised euro-area risk can also be computed by looking at measures of comovements of sovereign risk premia. Bufano and Manna (2012) carry out a principal component analysis of the 10-year swap spreads for the ten leading euro-area sovereign issuers.[23] They find that the first principal component explains 95% of the overall variance of sovereign swap spreads and its trend closely tracks the main phases of the sovereign debt crisis (Figure 14, left-hand panel): it is virtually unchanged until the third quarter of 2008, picks up in late 2008-early 2009, starts rising in the second quarter of 2010, reaches a maximum in the summer 2012, and sharply declines afterwards.

[23] The countries included in the sample are: Austria, Belgium, Finland, France, Germany, Ireland, Italy, the Netherlands, Portugal, Spain. The swap spread is the difference between the yield of the benchmark bond on a given maturity and the swap rate for that maturity. This measure was preferred to the perhaps more conventional yield spread with respect to the German Bund to allow the model also to provide an estimate of the fair value for the sovereign risk premium for Germany.

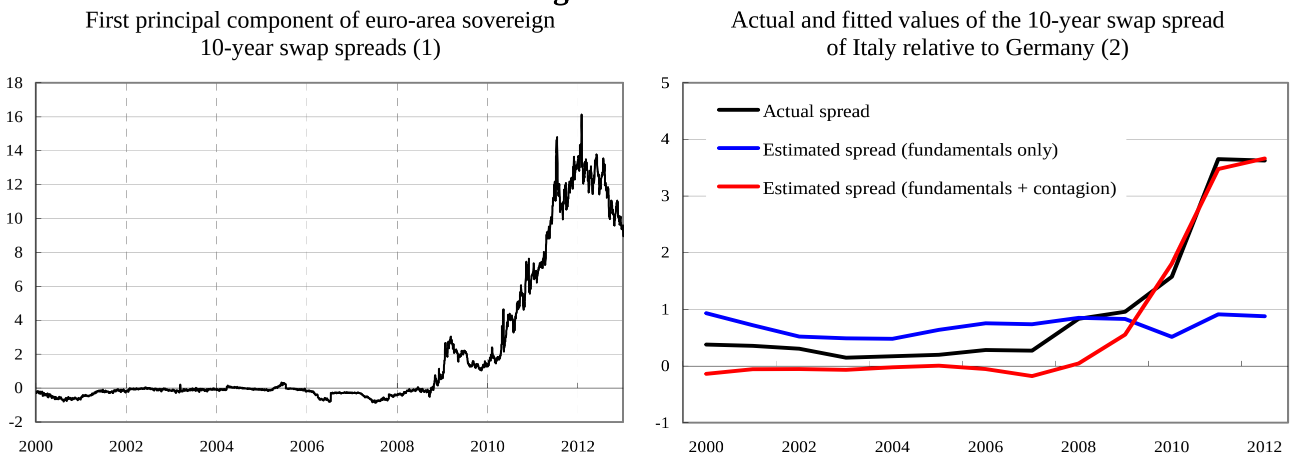
Using that indicator as a proxy of systemic euro-area risk, and building upon Bufano and Manna (2012), we estimate the following panel model for sovereign 10-year swap spreads:

$$s_{i,t} = \beta_0 + \beta_{1,i} C1_t + \beta_2 \left(\frac{debt}{GDP} \right)_{i,t} + \beta_3 \left(\frac{deficit}{GDP} \right)_{i,t} + \beta_4 E_{t,t+5} \left(\frac{\Delta GDP}{GDP} \right)_{i,t} + \varepsilon_{i,t} \, , \tag{8}$$

where *C1* is the first principal component and the expected growth in real GDP refers to a five-year horizon.[24] The right-hand panel of Figure 14 shows the fitted values of the sovereign spread of Italy relative to Germany for two different specifications. According to the specification that only includes fundamentals among the explanatory variables – model (8) without the systemic risk indicator *C1* – the predicted value of the 10-year yield spread between Italy and Germany for October 2012 is equal to 90 basis points. This number is more than 250 basis points lower than its actual value at that time and broadly unchanged from the fitted value for 2011. If one also includes the systemic risk indicator *C1* on the right-hand side of model (8), the forecast for the Italian swap spread increases considerably while that for the German swap spread declines somewhat. In this case the predicted level of the spread of Italy with respect to Germany reaches 370 basis points, close to the actual value.

Figure 14

Sovereign risk premia of euro-area countries:
A model of generalised euro-area risk



Source: based on Bufano and Manna (2012).
(1) Daily data. The principal component analysis is carried out from 2000 to2012. – (2) Annual data, in percentages. Data as of 1st October of each year. Fitted values for 2012 are out-of-sample fits.

[24] The model is estimated on data from 2000 to 2011. The estimated coefficients turn out to have the expected sign and are in line with the results found in previous studies. For details, see Bufano and Manna (2012), where a slightly different specification of model (8) is analysed, over an extended sample.

6. Conclusions

The analyses presented in this paper – which in some cases are obtained by building on the results of other studies – are broadly consistent with those of the earlier literature. In particular, financial market indicators and econometric results suggest that:

(i) In recent months the spectacular reduction of long-term German yields (standing at around 1.3 per cent as of end-August 2012) is to a large extent due to safe haven flows;

(ii) For several countries, we find robust evidence that in the most recent period the spreads vis-à-vis the German Bund have risen to levels that are significantly higher than what could be justified by fundamentals;

(iii) For Italian government bonds, most estimates of the value of the 10-year spread consistent with fundamentals are around 200 basis points, against its market value of about 450 points (at end-August 2012). The values estimated on the basis of fundamentals are markedly lower than the actual values also for the 2- and 5-year spreads.

The large gap between the market and model-based values of sovereign spreads needs to be explained. Possible alternative hypotheses are the following:

- One cannot completely rule out the possibility that financial market participants' expectations about the fiscal outlook are much more negative than one can gauge on the basis of past trends or consensus forecasts. However, given the relatively small magnitude of the estimated effects of these variables on interest rates (as explained in Section 3), it is worth observing that these pessimistic scenarios should imply a massive and persistent increase in public deficits and debts, much larger than is usually discussed anecdotally by market participants.
- Another possibility is that market participants have a biased perception of the risks associated with sovereign bonds. This might come from the difficulty of exactly measuring and quantifying these risks, which might lead investors to make oversimplifying assumptions (e.g. rule-of-thumb assessments) and take into consideration only very pessimistic or worst-case scenarios.
- Even under the hypothesis that risks are correctly measured, there may have been a surge in the price required by investors to bear these risks. Re-pricings of risk of this kind are inherently difficult to measure as they are intimately related to unobservable changes in investors' preferences and non-diversifiable risks. However, some of the regressions presented in this

paper suggest that the large discrepancies between the actual and model-based values of the spreads persist even when changes in the price of risk are controlled for by considering time-varying coefficients or commonly used proxies of investors' risk aversion.

While we plan to assess the contribution of these alternative explanations in future work, the size and persistence of the recent dynamics of interest rates that is not explained by fundamentals suggest that some common new risk factor is at play, clearly not accounted for by the models used so far.

Given the timing of the increase of sovereign yields in the countries most exposed to tensions and the concurrent, spectacular fall of sovereign yields in fiscally sounder countries, the natural and most likely candidate for the large gap between the market and model-based values of sovereign spreads is the perceived risk of a break-up of the euro area. Concerns about the fragility of the euro are increasingly and widely mentioned by a number of market observers and have apparently caught the attention of the public at large. The assumption of a prominent role of euro break-up risks is also corroborated by some new findings presented in this paper. For the bonds issued by some "core" and "non-core" countries the deviations of the yields from the values justified by fundamentals are in opposite directions. Moreover, those deviations turn out to be strongly correlated with an indicator of euro break-up risks. In conclusion, fears of the reversibility of the euro have likely played a key role in the recent huge widening of the dispersion of government bond yields across euro-area countries.

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Annex 1: The literature on the impact of fiscal variables on interest rates: A synoptic table

Summary of selected empirical works on the impact of fiscal variables on sovereign bonds, reprint from Haugh, Ollivaud and Turner (2009)

Reference	Countries	Fiscal variables (1)	Estimated effects on long-term interest rates in basis points (bps)
		Studies that focus on flow fiscal variables	
Thomas and Wu (2009)	United States	A 1% point increase in projected fiscal deficit in 5 years	30-60 bps
Bernoth et al (2006)	14 EU countries	A debt -service ratio 5% above Germany's	32 bps (spread vs. Germany, post-EMU period, some non-linear effects)
Dai and Philippon (2005)	United States	A 1% point increase in fiscal deficit lasting 3 years	20-60bp
Ardagna et al (2007)	16 OECD countries	A 1% point deterioration in primary balance	10 bps
Laubach (2003)	United States	A 1% point increase in projected fiscal deficit	25 bps
Literature review by Gale and Orzag (2003)	United States	A 1% point increase in projected fiscal deficit	40-50 bps
Literature review by Gale and Orzag (2002)	United States	A 1% point increase in projected fiscal deficit	50-100 bps (macro models) 50 bps (others)
Canzeroni, Cumby and Diba (2002)	United States	A 1% deterioration in projected fiscal balance, 5 to 10 year ahead	41-60 bps (Spread of 10-year yield over 3-month)
Linde (2001)	Sweden	A 1% deterioration in fiscal balance	25 bps after 2 years (Domestic-foreign long-term interest differential)
Reinhart and Sack (2000)	19 OECD countries G7	A 1% deterioration in fiscal balance in current and next years	9 bps (yield) 12 bps (yield)
Orr, Edey and Kennedy (1995)	17 OECD countries	A 1% point deterioration in fiscal balances	15 bps
		Studies that focus on stock fiscal variable	
Chinn and Frankel (2005)	Germany, France, Italy, UK and Spain USA	A 1% increase in current net debt A 1% increase in net public debt ratio projected 2 years ahead A 1% increase in current or projected net debt	5-8 bps 10-16 bps 5 bps over period 1998-2002, but obscured when extended to 2004
Ardagna et al (2007)	16 OECD countries	Public debt	non-linear
Engen and Hubbard (2004)	United States	A 1% point increase in debt ratio	3 bps (with ranges)
Laubach (2003)	United States	A 1% point increase in projected debt ratio	4 bps
Chinn and Frankel (2003)	Germany, France, Italy, Japan, Spain UK and USA	A 1% increase in net public debt ratio projected 2 years ahead	3-32 bps (individual country) 7-12 bps (European interest rates)
Codogno et al (2003)	9 EMU countries	Debt-to-GDP ratio	Small and significant effects on spreads for Austria, Italy and Spain
Conway and Orr (2002)	7 OECD countries	A 1% point increase in net public debt	Less than 1 bps (Real 10-year bond yields, starting from zero net debt) 1.5 bps (Real 10-year bond yields, starting from 100% net debt)
O'Donovan, Orr and Rae (1996)	7 OECD countries	A 1% point increase in net public debt	Less than 1 bps (Real 10-year bond yields, starting from zero net debt) 2 bps (Real 10-year bond yields, starting from 100% net debt)
Ford and Laxton (1995)	9 countries World	A 1% point increase in world net public debt	14 - 49 bps (Real 1-year bond yields) 15 -27 bps (Real 1-year bond yields)

Source: Haugh, Ollivaud and Turner (2009).

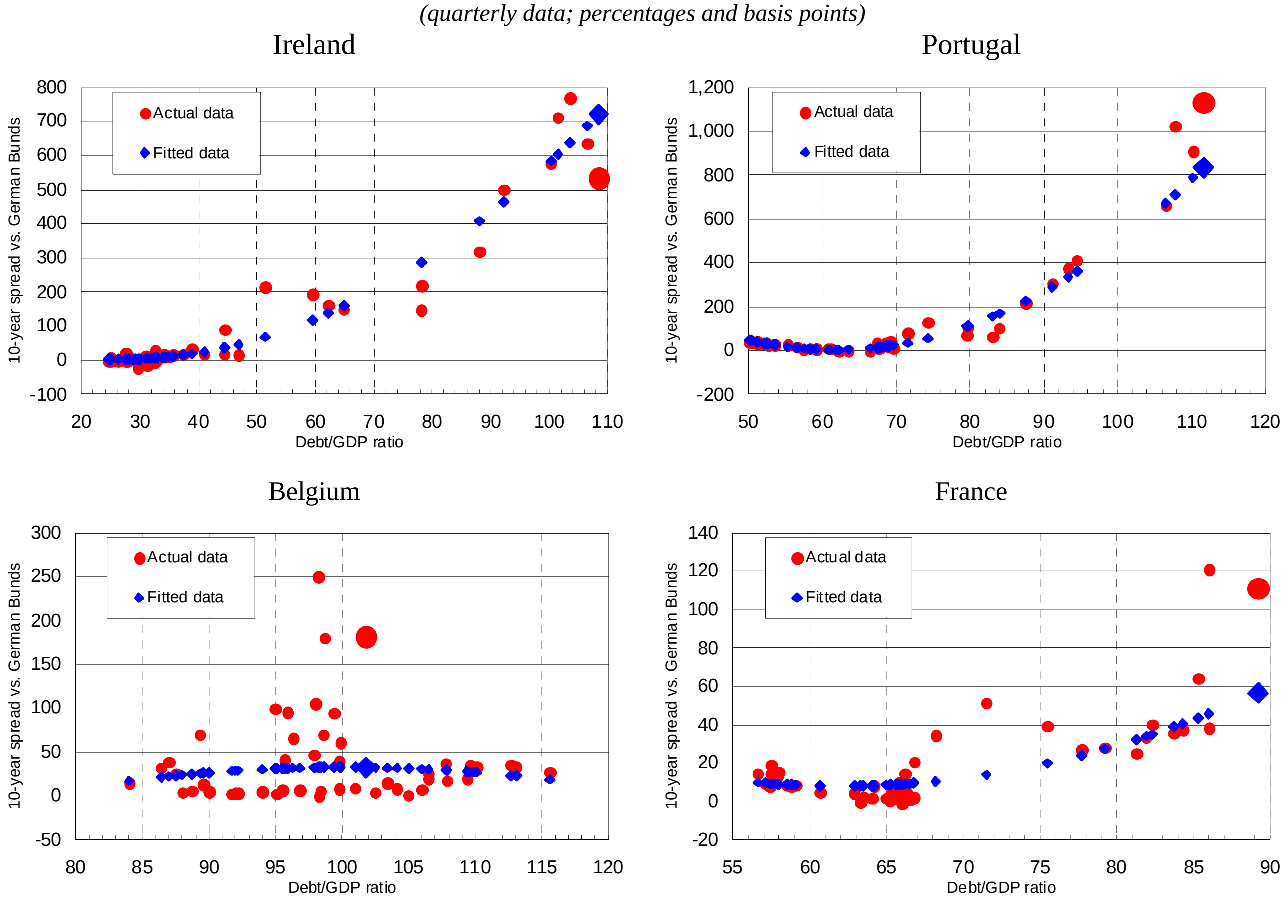
(1) All changes are expressed in relation to GDP unless otherwise specified.

Annex 2: Graphs

Figure A.1

Ten-year sovereign spreads with respect to Germany as a non-linear function of debt-to-GDP ratios (1)

(quarterly data; percentages and basis points)

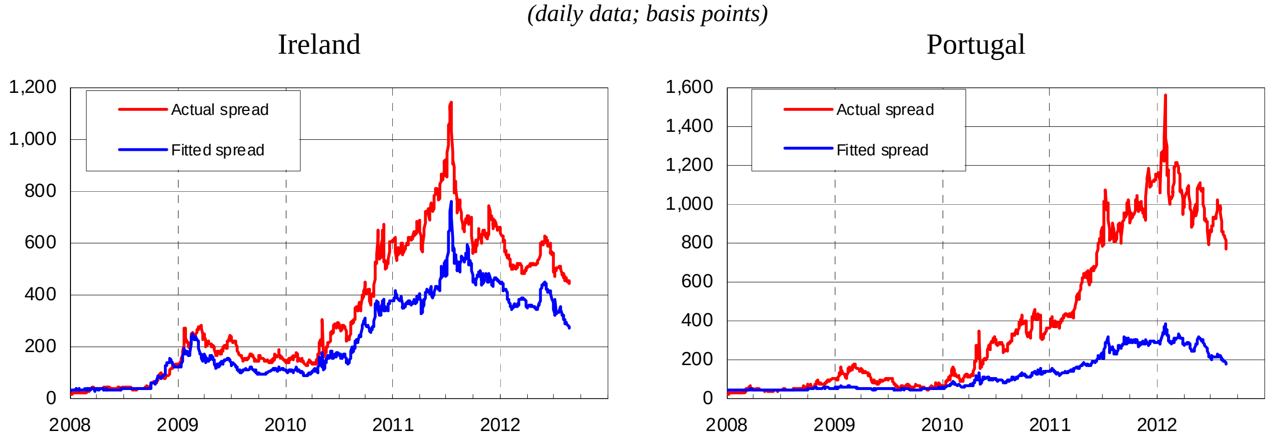


Source: based on Bloomberg and Thomson Reuters Datastream data.
(1) The larger markers denote the latest observations (2012 Q1).

Figure A.2

Ten-year sovereign spreads with respect to Germany based on the results by Ainzenman et al. (2011) for sovereign CDSs (1)

(daily data; basis points)

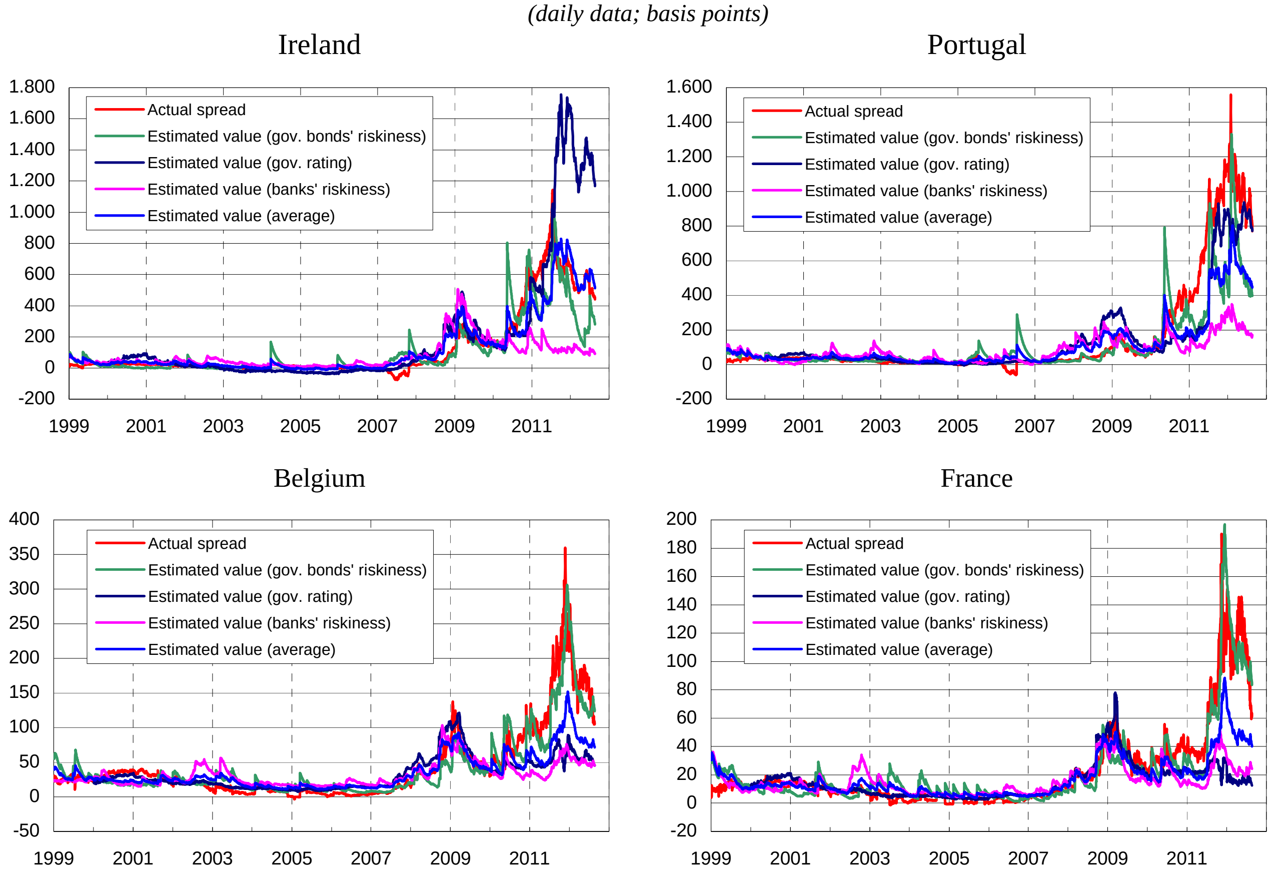


(1) Fitted values are generated by using the estimates by Ainzenman, Hutchison and Jinjarak (2011) of the value of the premia on sovereign 5-year CDSs consistent with fundamentals.

Figure A.3

Ten-year sovereign spreads with respect to Germany as a function of financial indicators of country risk

(daily data; basis points)

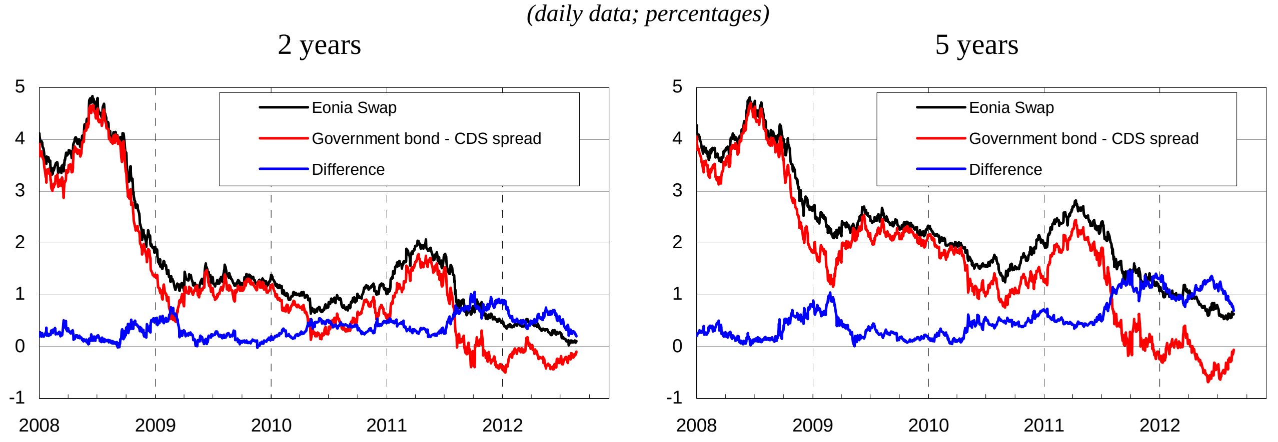


Source: based on Bloomberg and Thomson Reuters Datastream data.

Figure A.4

Differential between the Eonia swap rate and the sovereign bond-CDS spread for Germany at the 2- and 5-year maturities

(daily data; percentages)

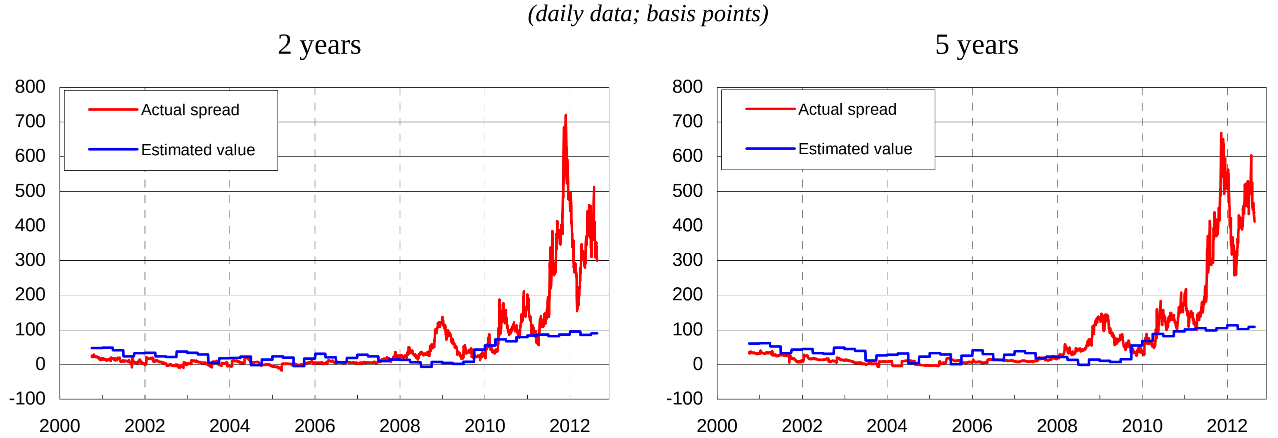


Source: based on Bloomberg data.

Figure A.5

Italian 2- and 5-year sovereign spreads with respect to Germany as a function of debt-to-GDP ratios

(daily data; basis points)



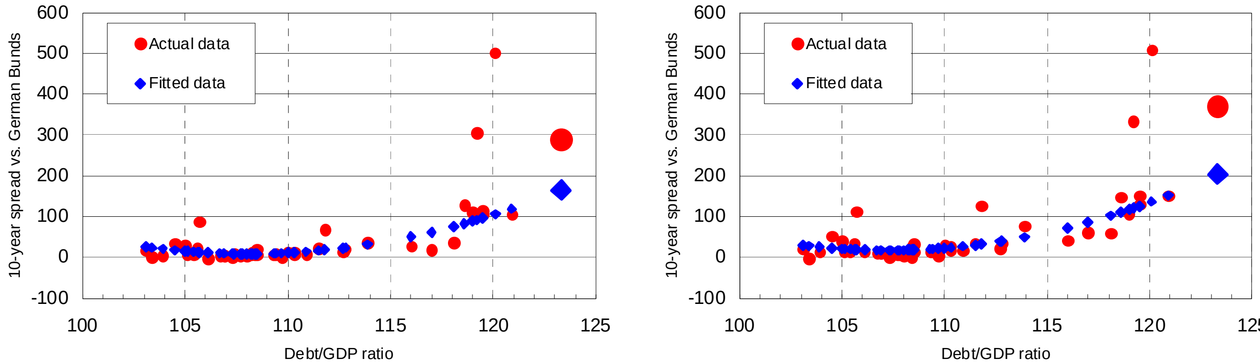
Source: based on Bloomberg and Thomson Reuters Datastream data.

Figure A.6

Italian 2- and 5-year sovereign spreads with respect to Germany as a non-linear function of debt-to-GDP ratios (1)

(quarterly data; percentages and basis points)

2 years 5 years



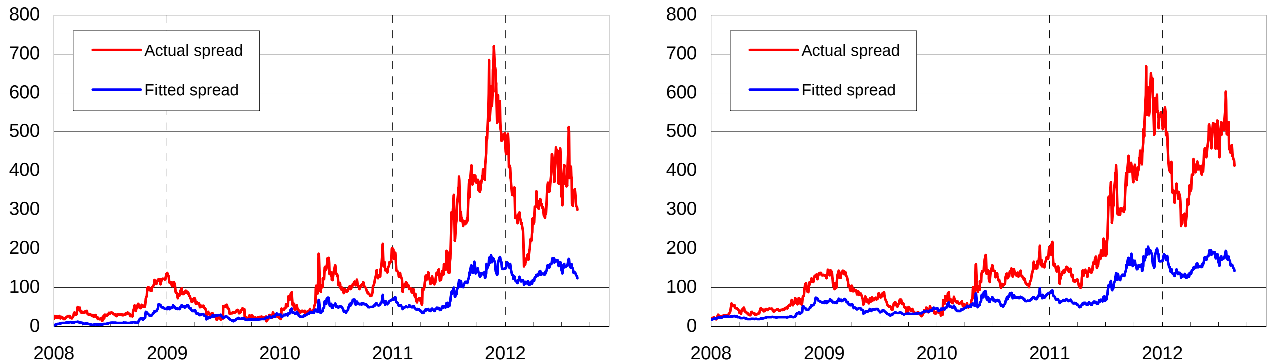
Source: based on Bloomberg and Thomson Reuters Datastream data.
(1) The larger markers denote the latest observations (2012 Q1).

Figure A.7

Estimates of the Italian 2- and 5-year sovereign spreads with respect to Germany based on the results by Ainzenman et al. (2011) for sovereign CDSs (1)

(daily data; basis points)

2 years 5 years



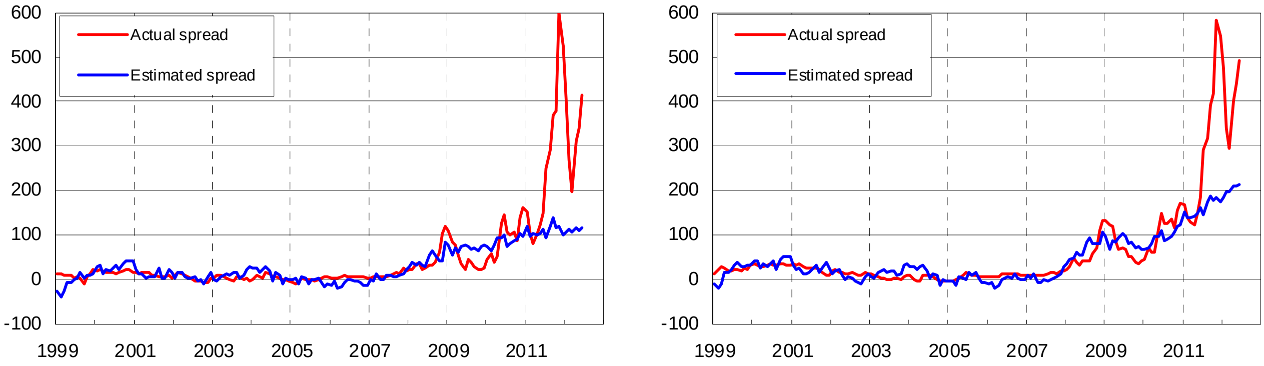
(1) Fitted values are generated on the basis of Ainzenman, Hutchison and Jinjarak (2011)'s estimates of the value of the premia on sovereign 5-year CDSs that are consistent with current fundamentals.

Figure A.8

Italian 2- and 5-year sovereign spreads with respect to Germany as a function of consensus expectations on fiscal and other economic fundamentals (1)

(monthly data; basis points)

2 years 5 years



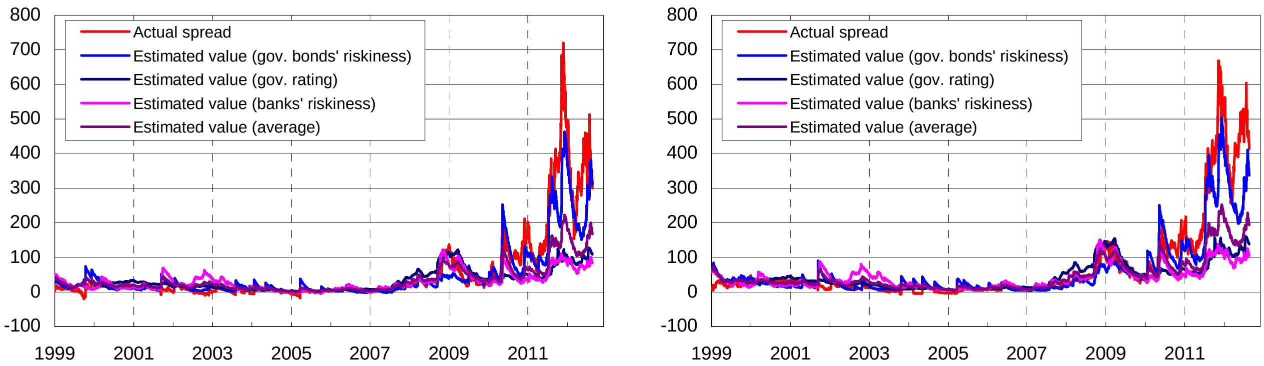
Source: based on Bloomberg, Thomson Reuters Datastream and Consensus Forecasts data.
(1) The estimated spread is the difference between the fitted values of the Italian and German interest rates. Interest rates are modelled as a function of the expected deficit/GDP ratio over the next 12 months and the 12-month-ahead forecasts of other macroeconomic variables (expected three-month interest rates, GDP growth rate, consumer price inflation, unemployment rate and the current account/GDP ratio). Since July 2011 the estimated spread is based on out-of-sample forecasts.

Figure A.9

Italian 2- and 5-year sovereign spreads with respect to Germany as a function of financial indicators of country risk

(daily data; basis points)

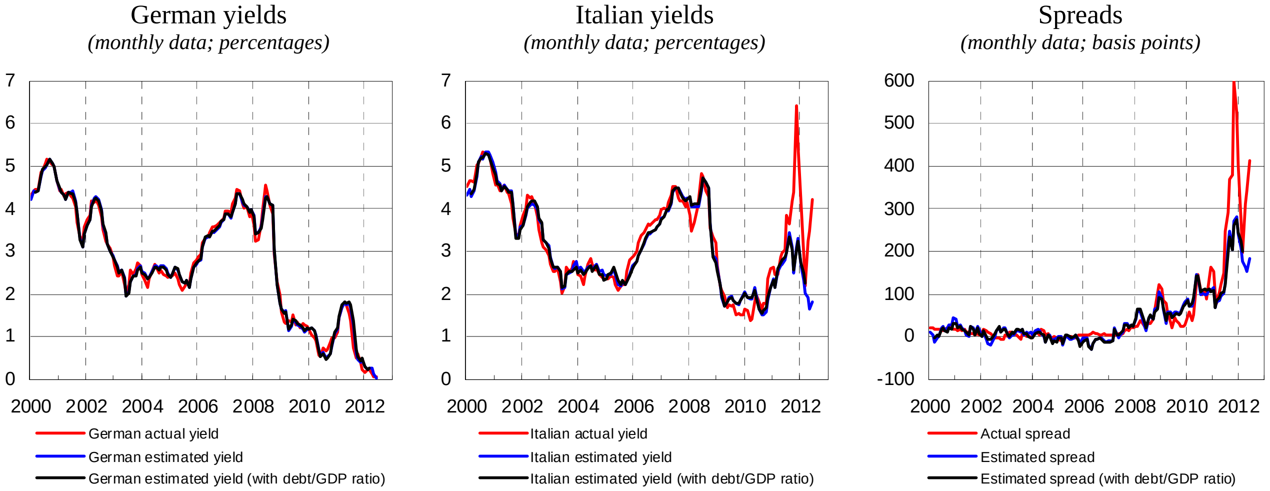
2 years 5 years



Source: based on Bloomberg, Thomson Reuters Datastream and Consensus Forecasts data.

Figure A.10a

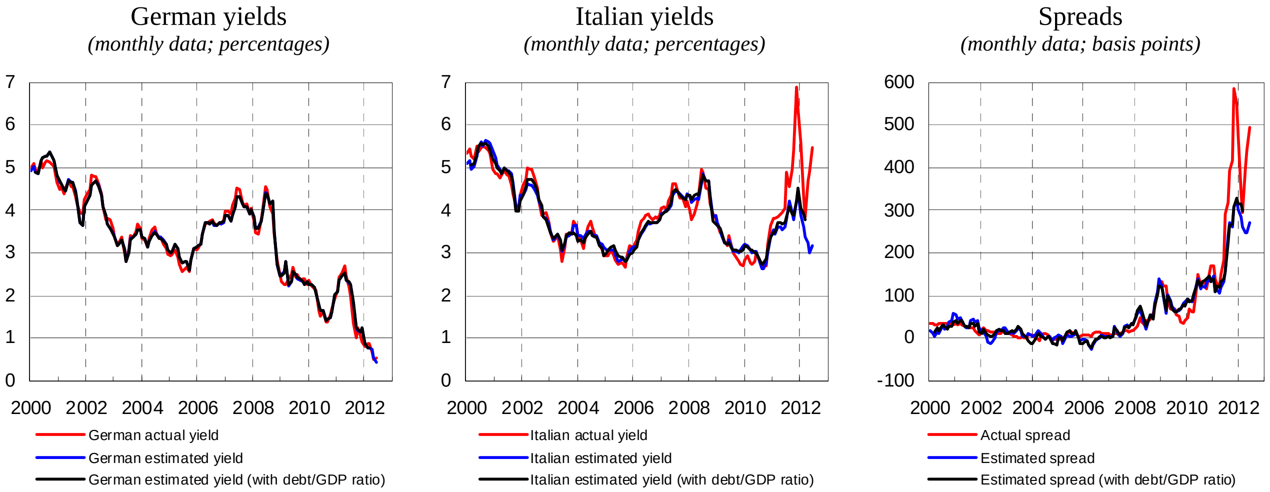
German and Italian 2-year government bond yields and spreads as a function of fundamentals and financial factors



Source: based on Bloomberg, Thomson Reuters Datastream and Consensus Forecasts data.
(1) Yields are modelled as a function of three financial risk indicators (yield volatility, bank share price volatility and yield of corporate bonds with the same rating as the sovereign), the expected deficit/GDP ratio over the next 12 months and the 12-month-ahead forecasts of other macroeconomic variables (expected three-month interest rates, GDP growth rate, consumer price inflation, unemployment rate and the current account/GDP ratio). An extended specification also includes the current level of the debt-to-GDP ratio among the regressors. Since July 2011 fitted values are based on out-of-sample forecasts.

Figure A.10b

German and Italian 5-year government bond yields and spreads as a function of fundamentals and financial factors



Source: based on Bloomberg, Thomson Reuters Datastream and Consensus Forecasts data.
(1) Yields are modelled as a function of three financial risk indicators (yield volatility, bank share price volatility and yield of corporate bonds with the same rating as the sovereign), the expected deficit/GDP ratio over the next 12 months and the 12-month-ahead forecasts of other macroeconomic variables (expected three-month interest rates, GDP growth rate, consumer price inflation, unemployment rate and the current account/GDP ratio). An extended specification also includes the current level of the debt-to-GDP ratio among the regressors. Since July 2011 fitted values are based on out-of-sample forecasts.

Annex 3: Econometric results

Table A – Econometric results of selected models

The table shows coefficient estimates and associated p-values for some of the models presented in Section 4 (the sub-section is indicated in the first row of the table). Standard errors are corrected for autocorrelation and heteroskedasticity with the Newey-West algorithm.Country codes: DE=Germany, IT=Italy.

Reference section	4.5.1		4.3		4.5.1		4.3		4.4		4.4		4.4		4.4	
Dependent variable	10-year IT yield		10-year IT yield		10-year DE yield		10-year DE yield		10-year IT-DE spread		10-year IT-DE spread		10-year IT-DE spread		10-year IT-DE spread	
	coef.	p-value	coef.	p-value	coef.	p-value	coef.	p-value	coef.	p-value	coef.	p-value	coef.	p-value	coef.	p-value
Constant	0.21	[0.65]	0.92	[0.14]	-1.16	[0.02]	3.67	[0.00]	9.21	[0.00]	-1.26	[0.78]	4.99	[0.13]	12.30	[0.00]
3-month rate	0.05	[0.43]	0.37	[0.00]	0.11	[0.24]	0.46	[0.00]								
GDP	0.19	[0.00]	-0.12	[0.00]	0.10	[0.00]	0.00	[0.95]								
CPI	-0.14	[0.02]	-0.36	[0.35]	-0.08	[0.29]	-0.39	[0.00]								
Unemployment	0.19	[0.00]	0.39	[0.00]	0.17	[0.00]	0.01	[0.88]								
Budget balance	0.06	[0.21]	0.30	[0.00]	0.01	[0.86]	0.08	[0.25]								
Current account	-0.19	[0.01]	-0.36	[0.00]	-0.03	[0.08]	-0.12	[0.00]								
Yield volatility	5.04	[0.05]			4.15	[0.28]										
Spread volatility									19.60	[0.00]					18.90	[0.00]
Banks volatility	-0.11	[0.01]			-0.10	[0.00]					33.50	[0.00]			-13.30	[0.00]
Corporate spread	0.53	[0.00]			0.74	[0.00]							0.52	[0.00]	3.90	[0.00]
R-square	0.94		0.84		0.96		0.84		0.80		0.38		0.52		0.88	
Sample period	2000:M1-2011:M6		1999:M1-2011:M6		2000:M1-2011:M6		1999:M1-2011:M6		4 Jan. 99-30 June 11		4 Jan. 99-30 June 11		4 Jan. 99-30 June 11		4 Jan. 99-30 June 11	

Table B – Robustness of the estimates of the 10-year Italian yield

The table shows the equation of the 10-year Italian yield presented in Section 4.5 estimated on three different sample periods: 2000:M1-2011:M6 (as in Section 4.5.1); 2007:M1-2011:M6; 2007:M1-2012:M11. Standard errors are corrected for autocorrelation and heteroskedasticity with the Newey-West algorithm.

Dependent variable	10-year IT yield		10-year IT yield		10-year IT yield	
	coef.	p-value	coef.	p-value	coef.	p-value
Constant	0.21	[0.65]	-1.08	[0.20]	1.06	[0.30]
3-month rate	0.05	[0.43]	0.21	[0.01]	0.05	[0.73]
GDP	0.19	[0.00]	0.12	[0.02]	0.10	[0.20]
CPI	-0.14	[0.02]	-0.23	[0.02]	-0.29	[0.02]
Unemployment	0.19	[0.00]	0.43	[0.00]	0.27	[0.00]
Budget balance	0.06	[0.21]	0.22	[0.00]	0.32	[0.00]
Current account	-0.19	[0.01]	-0.28	[0.00]	-0.37	[0.00]
Yield volatility	5.04	[0.05]	7.06	[0.00]	11.61	[0.00]
Spread volatility						
Banks volatility	-0.11	[0.01]	-0.20	[0.00]	-0.21	[0.00]
Corporate spread	0.53	[0.00]	0.46	[0.00]	0.38	[0.00]
R-square	0.94		0.90		0.89	
Sample period	2000:M1-2011:M6		2007:M1-2011:M6		2007:M1-2012:M11	

Pure or wake-up-call contagion?
Another look at the EMU sovereign debt crisis

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May 2013

Abstract

We test whether the sharp increase in sovereign spreads of euro-area countries with respect to Germany after the explosion of the Greek crisis was due to deteriorating macroeconomic and fiscal fundamentals or to some form of financial contagion. Our analysis includes indicators of domestic and external imbalances which were mostly disregarded by previous studies, and distinguishes between investors' increased attention to the variables which ultimately determine the creditworthiness of a sovereign borrower (wake-up-call contagion) and behaviour not linked to fundamentals (pure contagion). We find evidence of wake-up-call contagion but not of pure contagion.

JEL Classification: E62, H62, H63.

Keywords: sovereign bond spread; contagion; non-stationary panels.

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1 Introduction

At the beginning of 2009, ten years after the launch of the euro, many commentators viewed the single currency as a major success. In the run-up to the euro's introduction, interest rates had rapidly converged towards the low levels of the most creditworthy member states: in the period 1992-1998, the average spread of long-term government bond yields with respect to the German one had declined from about 200 to 24 basis points. From 1999 onwards spreads continued to narrow, and at the end of 2007 they were negligible (16 basis points on average). Due to the financial turmoil triggered by the Lehman Brothers bankruptcy, some tensions started to surface in September 2008, but at the end of that year the average yield spread in the euro area was still about 100 basis points. Strains on government securities markets became worrisome only towards the end of 2009 (Figure 1). The focus of concern was Greece. After a series of upward deficit revisions, the last of which equal to nearly 3 percentage points of GDP in October 2009, the Greek government estimated the deficit at 12.7 per cent of GDP in 2009, up from 7.7 per cent in 2008. The tensions spilled over from Greece to the government securities of other euro-area countries, notably Ireland, Portugal and, to a lesser extent, Spain and Italy. Three years after these events, some countries still are basically shut out of the bond market[1] and sovereign debt strains in the euro area remain worrisome and widespread, despite important progresses in fiscal adjustment by national governments.

The debate concerning the causes of the European sovereign debt crisis inflames both politics and academia. While some argue that fiscal deterioration and fundamental macroeconomic weaknesses are at the root of the crisis, others claim that spreads are well above the levels justified by fundamentals, and invoke forms of "market irrationality" and/or "contagion". The aim of the present paper is to assess the relative merits of these competing opinions through a formal econometric analysis.

Needless to say, the answer to this question has significant policy implications. Evidence of sizable and systematic mispricing of sovereign credit risk would imply that it is ill-advised to rely on markets to induce fiscal and macroeconomic discipine. Furthermore, it would strengthen the case for interventions by European Union institutions such as the European Financial Stability Facility (EFSF), the European Stability Mechanism (ESM) and the European Central Bank (ECB) in the sovereign bond markets. In fact, the Eurogroup summit of 29 June 2012 decided to use the EFSF/ESM instruments in order to stabilize the markets of member states honouring all their European commitments on schedule. Soon afterwards, the ECB decided to undertake Outright Monetary Transactions (OMT) in the secondary markets for sovereign bonds in the euro area "to address severe distortions which originate from, in particular, unfounded fears of the reversibility of the euro" (press conference following the meeting of the Governing Council on 6 September 2012).

While several other papers have studied the relationship between spreads and fiscal fundamentals in European Monetary Union (EMU), ours contributes to the discussion in three ways. First, it

[1] Greece applied for financial support in May 2010, followed by Ireland (November 2010) and Portugal (April 2011).

considers a broader set of fundamentals. One lesson of the EMU crisis is that even countries with low levels of public debt and deficits can suffer a sudden deterioration of their fiscal position, for example as an effect of financial sector bailouts (which may transform private liabilities into public debt). This risk was considered obvious for emerging markets at least since the Asian crisis of the late nineties, but it was not taken into account by the EMU rules and – as we show here – by investors. Our second contribution to the literature is to distinguish between different forms of contagion and to measure their relative importance in explaining the post-crisis behaviour of European sovereign spreads.[2] Our third contribution is methodological: for the first time we apply to sovereign spreads panel methodologies designed to detect and tackle non-stationarity and cointegration.

To give a preview of our results, we find that the explosion of the Greek crisis had a systematic impact on the other euro area countries' sovereign spreads. However, this impact differed across borrowers. In particular, investors penalized governments with weaker fiscal and macroeconomic fundamentals more heavily .

The rest of the paper proceeds as follows. In Section 2 we review the literature and clarify our definition of contagion. In Section 3 we present our dataset and in Section 4 we discuss our empirical strategies and show our results. In Section 5 we discuss several robustness checks. In Section 6 we provide numerical estimates of the long-run values of the spreads, derived from our empirical analysis. Finally, in Section 6 we draw some tentative conclusions and policy implications.

2 Literature review

Several papers assess the determinants of sovereign spreads in EMU. Starting from Codogno et al. (2003), the literature has expanded significantly in the last few years (see, among others, Favero et al. (2010), Beber et al. (2010), Schuknecht et al. (2009, 2011), Attinasi et al. (2009), Sgherri and Zoli (2009), Hallerberg and Wolff (2008)). Typically, these studies explore the role of (a) country-specific factors, namely fiscal fundamentals and market liquidity, and (b) common factors, such as the market appetite for risk. In particular, they bring to the data an empirical model such as:

$$s_{it} = \alpha_0 + \alpha_1 s_{it-1} + \beta_0 Z_{it} + \beta_1 F_t + \varepsilon_{it} \ , \ |\alpha_1| < 1 \tag{1}$$

where Z_{it} is a vector of country-specific variables and F_t is a vector of variables that are common across countries. The above-mentioned papers differ from one another in terms of data frequency (from daily to yearly), the regressors included and estimation method (in particular, some adopt a pooled cross-section/time-series approach, others provide country-specific estimates). Of course, studies using high-frequency data, such as Favero et al. (2010) and Beber et al. (2010), do not consider the role of fiscal and macro fundamentals, which are available only at lower frequencies.

[2] Of course, the two contributions are related: to understand whether spreads are excessive with respect to fundamentals, it is necessary to take a stance concerning the relevant fundamentals.

Bernoth et al. (2012) consider a slightly different dependent variable (primary instead of secondary market spreads); their sample period ends in 2009, so it does not include the post-Greek-crisis period. Although their analysis focuses on the structural break due to the introduction of EMU in 1999, it also discusses the possible effects of the Lehman bankrupcy in September 2008. Using an approach similar to ours, Bernoth et al. (2012) find that the Lehman bankrupcy increased the sensitivity of spreads to country-specific fundamentals and global factors.

Few papers consider instead the issue of contagion among sovereign securities within EMU. Some papers simply augment equation (1) with a further Z_{it} variable, which captures developments in all the other EMU countries different from i. In particular, Caceres et al. (2010) employ a measure of "distress dependence", which is built by extracting from the vector of CDS premia the unconditional marginal probability of default for each country. They then infer from those marginal distributions the joint probability of default, and build and add-up the default probability of country i conditional on the default of the other countries. Similarly, Hondroyiannis et al. (2012) add a "contagion variable", defined as a weighted combination of other countries' spreads. Neither Caceres et al. (2010) nor Hondroyiannis et al. (2012) consider the more recent years.

Our contribution borrows from a different strand of the literature, which discusses contagion concentrating on developing countries. In this literature, more precise and circumscribed definitions of contagion are used.[3] We follow, in particular, Eichengreen et al. (1996), Masson (1998) and Goldstein et al. (2000), who distinguish between three kinds of circumstances:[4]

- Wake-up-call contagion, a situation in which a crisis initially confined to one country provides new information that prompts investors to reassess the default risk of other countries (this concept is used, for example, by Goldstein, 1998, Masson, 1999, Goldstein et al., 2000). In this case, domestic fundamentals justified a flight from sovereign debt even before the crisis event, but investors did not price/perceive the risk correctly. The wake-up-call hypothesis was first put forward by Goldstein (1998) to explain contagion from Thailand (a relatively small and closed economy) to other Asian countries in the Asian crisis of the late nineties. He argues that the other countries were affected by the same structural and institutional weaknesses as Thailand (crony capitalism, weak banking system, etc.), but investors ignored those weaknesses until the Thai "wake-up call". Such behaviour is also consistent with forms of "rational inattention" (Tutino, 2011, and Wiederholt, 2010). According to rational inattention theory, given the existence of costs in aquiring and processing information, rational agents could optimally choose to ignore some information, for example concerning fundamentals.

- Shift contagion, which occurs when the normal cross-market channel intensifies after a crisis in one country. It can be seen as analogous to wake-up-call contagion except that it is due to

[3] This literature is surveyed in Pericoli and Sbracia (2003), and Dungey et al. (2005).

[4] While our contagion definitions are quite widespread in the literature, other papers use the word "contagion" differently (as discussed in the recent survey by Forbes, 2012).

increased sensitivity to common factors such as global risk aversion - the F_t term in equation (1) - instead of country-specific factors. We borrow the term and the concept from the work of Forbes and Rigobon (2002).

- Pure contagion. This residual category covers any instance of contagion that is completely unrelated not only to changes in fundamentals (as in the case of wake-up-call and shift contagion) but also to the level of fundamentals, be they country-specific (as in the case of the wake-up-call contagion) or global (as in the case of shift-contagion). Pure contagion may arise from self-fulfilling (and therefore individually rational) loss of confidence (Calvo, 1988), from irrational herding behaviour (Chari and Kehoe, 2003), or from margin calls and other wealth effects for investors, triggered by capital losses in the country which originated the crisis (Kodres and Pritsker, 2002, Kyle and Xiong, 2001, Calvo and Mendoza, 2000, Schinasi and Smith, 2000).

In distinguishing between the three types of contagion, our contribution is similar to the paper by Bekaert et al. (2011). They use an international asset pricing framework with global and local factors to predict equity returns, defining unexplained increases in factor loadings as indicative of contagion, and find evidence of systematic contagion whose severity is inversely related to the quality of countries' economic fundamentals and policies. They conclude that the wake-up-call hypothesis holds for equity markets, with markets and investors paying substantially more attention to country-specific characteristics during the crisis.

We also see the approach pioneered by Gande and Parsley (2005) as very relevant and complementary to ours. They consider a sample of emerging countries and allow rating news concerning any one of them to influence the sovereign spreads in the others. In the present paper, we likewise consider a unidirectional version of their methodology, substituting our crisis dummy with a variable summarizing Greek rating developments.[5]

Finally, let us remark that in our regressions, while taking into account the possibility that the situation of banks may have an impact on sovereign spreads, we focus on contagion across sovereign bond markets, leaving aside the issue of contagion from sovereign to other financial markets or to the banking sector (on this, see, among others, Acharya et al, 2011, Alter and Schuler, 2011, Angeloni and Wolff, 2012).

[5] Two recent papers on the EMU sovereign debt crisis use multi-equation econometric techniques and can be seen as multi-equation extensions of Gande and Parsley (2005). Arezeki et al. (2011) estimate a VAR model allowing for the mutual inter-dependence of sovereign debt markets and the stock market. De Santis (2012) allows for a long-run co-integrating relationship between spreads and other variables. Chudik and Fratzscher (2013) use the VAR methodology to study yields (not spreads) and consider stocks and foreign currencies in addition to sovereign bonds.

3 Data and descriptive statistics

Our dataset covers nine euro-area countries (Austria, Belgium, Finland, France, Ireland, Italy, Portugal, Spain and the Netherlands) using monthly data from January 2000 to December 2011. As is customary in the literature, we exclude Greece (the "ground-zero" country) from the analysis.[6]

Our dependent variable is the 10-year government bond yield spread with respect to the corresponding German Bund.[7]

In our baseline specification we consider as common factor - the F_t variable in equation (1) - the VIX, the most common indicator of the propensity of investors to bear credit risk.[8] Data on government bond yields and on the other financial market variables are taken from Thomson Financial Reuters. These data are released daily, and we compute monthly averages of them.

Like our dependent variable, country-specific fundamentals - the Z_{it} vector in equation (1) - are in differences with respect to the corresponding German variables. They include GDP growth and the ratios with respect to GDP of general government debt, private sector debt, defined as household plus non-financial corporation debt, and the current account surplus.

We also control for liquidity, measured by the difference between the country's bid-ask spread on government bonds and the German one.[9] We do not control instead for differences in debt characteristics such as inflation-indexation and currency denomination. Indeed, unlike in emerging countries, in our sample public debt is mostly in nominal terms and denominated in euros.[10]

The inclusion of private debt and the current account balance, while non-standard in the literature on advanced economies (an exception is Gourinchas and Obstfeld, 2012), is frequent in studies concerning emerging countries and has strong economic rationale inasmuch as these are indicators of the domestic and external leverage of an economy. While a current account deficit does not mean per se a higher sovereign vulnerability, it is often associated with competitiveness imbalances

[6] We have verified that our main results do not change if Greece is included in the regressions. We excluded Luxembourg, because for most of the sample period it essentially had no public debt. We had to exclude the remaining five countries because, as recent entrants to the euro, the pre-crisis period was clearly too short for us to estimate reliably our model (Estonia and Slovakia joined the union in 2011 and 2009 respectively, Cyprus and Malta in 2008, Slovenia in 2007). Moreover, private debt data are missing for the late-accession countries.

[7] An often-used alternative measure for the default risk is the credit default swap (CDS) premia. However, for our purposes it suffers from several shortcomings. First, a well-developed CDS market exists only for few countries in our sample, and even for those countries data are available only for the more recent years. Second, CDS premia are driven not only by credit risk considerations but also by counteparty risk. Third, during the crisis in some countries CDS markets were subject to policy interventions, such as short-selling bans, which are likely to have had an impact on CDS premia.

[8] The VIX, the Chicago Board Options Exchange Market Volatility Index, is a measure of the implied volatility of the S&P 500 stock index; it is considered a good indicator of the level of risk aversion in global capital markets.

[9] This measure of liquidity is common in the literature (see, among others, Codogno, Favero and Missale, 2003, and Favero, Pagano and von Thadden, 2010). Our variable is computed as the difference between the minimum bid yield and the maximum ask yield observed at daily frequencies for benchmark bonds; this computational method implies limited variability over time of this difference. Favero, Pagano and von Thadden (2010) use instead the best five bid and ask prices.

[10] As is well known this is not true of emerging economies (see e.g. the contributions in Eichengreen and Hausmann, 2004). Concerning debt duration, in our sample we observe moderate cross-country differences, but they are basically time-invariant and therefore mostly captured by the country fixed effects.

and problematic macroeconomic developments. Furthermore, external capital inflows (the mirror image of the current account deficit) may trigger a boom in the non-tradable sector (particularly the housing market), increasing the risk of a subsequent bust.[11] A similar line of reasoning can be applied to private sector debt: if households and firms turn out to be unable to repay their debt, this might jeopardize public finances, either because the government may bail them out directly or – as often happens – because it bails out the domestic banks that lent to households and firms in the first place. In any case, in the presence of substantial private liabilities, public debt might increase significantly and overnight. Notice that both variables are to be monitored at the European level under the new Macroeconomic Imbalances Procedure (European Commission, 2012).[12]

Fiscal and macroeconomic variables are taken from the Eurostat quarterly database. These data are generally released with a delay of one quarter. Our monthly series are obtained keeping the value of the variable constant in each month of the quarter. In our specification we thus assume that spreads react simultaneously to liquidity and volatility factors and with a 3-month lag to fiscal and macroeconomic variables. This also limits endogeneity problems and thus concerns about possible reverse causation between the current spread and the independent variables.

In Table 1 we report some descriptive statistics of the variables used in our benchmark specification, distinguishing between two sub-periods (before and during the crisis). In the upper part of each panel we summarize the evolution of our dependent variable, i.e. the average yield spread, and the financial factors that in our specification are assumed to influence it. In the bottom part we summarize the development of fiscal and macroeconomic fundamentals. Statistics refer to all countries except Germany and Greece.

The spread between the government bond yields of these nine euro-area countries and the German one increased on average from 19 basis points in the period before the crisis to 175 basis points from October 2009 onwards. The increase was significantly larger in the sub-group of peripheral countries (Portugal, Ireland, Italy and Spain), from 25 to 330 basis points. Liquidity, measured by the bid-ask spread, worsened on average in the second part of our sample period (on average the spread increased from 1 to 6 basis points). The evolution of the VIX shows that global risk aversion increased during the euro-area sovereign crisis; however, as acute financial markets tensions had already emerged following the Lehman Brothers bankruptcy, the difference across sub-periods is not appreciable.

Turning to fundamentals, both fiscal and macroeconomic conditions deteriorated significantly during the sovereign debt crisis. Among domestic imbalances, the average general government debt increased by 17 percentage points of GDP (almost 30 in the peripheral countries); the increase in

[11] This in turn would induce sizable output gaps and revenue shortfalls, increasing public debt and jeopardizing its sustainability. This is how Spaventa and Giavazzi (2011) interpret the EMU crisis.

[12] Concerning external imbalances, the European scoreboard also includes the net investment position (the stock counterpart of the current account balance), the change in export market shares, the change in unit labour costs, and the change in the real effective exchange rate. Concerning domestic imbalances, the scoreboard includes the private-sector credit flow (the flow counterpart of domestic debt), the change in the house price index, and the unemployment rate.

private debt was even larger (42 percentage points in the entire sample and 57 in the peripheral countries). GDP growth slowed on average from 1.8 to 1.1 per cent, reflecting a negligible acceleration in the "virtuous" countries and a marked slow down in the others (from 2 to almost 0). External positions also worsened: on average the current account deficit increased from 0.5 to 0.7 per cent of GDP; with respect to Germany the deterioration was greater (about 2.5 percentage points of GDP), reflecting strongly diverging competitiveness paths between Germany, on one side, and the other countries, on the other.

4 Empirical analysis

We use two alternative empirical models. The first (Section 4.1) is akin to equation (1), as it assumes that the spread is a stationary variable, even if it has an auto-regressive component. As stationarity is assumed by all the previous literature, we provide estimates of this model mainly for the sake of comparability. However, as we will argue below, there are good empirical reasons to question the stationarity hypothesis and also to conjecture the existence of a long-run cointegrating relationship between the spread and the other covariates (Section 4.2). Therefore, we will subsequently focus on the estimation of that long-run relationship (Section 4.3).

4.1 Stationary case

The empirical model. - We enrich the specification in (1) in order to take into account the three different kinds of contagion effects outlined in Section 2. We estimate the following model:

$$s_{it} = \alpha_{i0}+\alpha_1 s_{it-1}+\beta_0 Z_{it}+\beta_1 F_t+\gamma_0 D_t+\gamma_1 D_t s_{it-1}+\gamma_2 D_t Z_{it}+\gamma_3 D_t F_t+\varepsilon_{it},\ \ |\alpha_1|,|\alpha_1+\gamma_1|<1 \quad (2)$$

where the error term is assumed zero-mean, stationary and independent across countries (but we allow for heteroskedasticity and auto-correlation), and D_t is a dummy variable taking value one after the outbreak of the Greek crisis, which in our model coincides with the revision of the official public finance figures by the new government in October 2009.

Therefore, γ_0 captures "pure contagion", the vector of coefficients γ_2 captures the wake-up-call effect (a more pronounced post-crisis sensitivity to country-specific fundamentals), and γ_3 captures shift-contagion (an increased sensitivity to common factors).

Notice that in our specification we allow for country-specific fixed effects, to control for time-invariant unobserved characteristics. Indeed, the previous literature has pointed to some very slow-moving features that influence a sovereign's creditworthiness, such as the political system (Akitoby and Stratmann, 2008) or debt intolerance (Qian et al. 2011). We also allow for a change in the auto-correlation coefficient in the post-crisis period (γ_1).

Baseline results. - The Least Square Dummy Variables (LSDV) estimates of equation (2) shows that in the pre-crisis period the only statistically significant coefficients are those of GDP growth and of the VIX: both a slowdown in GDP and a decrease in global risk appetite widen the spread (Table 2, column 1).

Instead, during the crisis the relationship becomes significant for all the fundamental variables except private debt and the bid-ask spread. This suggests that a wake-up-call effect exists for EMU countries. In particular, current account imbalances and public debt are not relevant in the pre-crisis period, whereas in the crisis period they become positively related to the sovereign spreads. By contrast, neither "pure contagion" nor "shift-contagion" effects are present (both γ_0 and γ_3 are insignificant). Finally, the estimated auto-correlation parameter is relatively high (with no change in the coefficient after the Greek crisis), which points to possible non stationarity.

Considering only the peripheral countries. - The results could be different if one only considers peripheral euro area countries. First, it is more likely that investors' attention to these countries was already high before the crisis, given that their fiscal reputation was already undeniably worse. This reduces the probability of observing wake-up-call contagion. Second, the probability of observing pure contagion should increase as investors possibly consider these countries more similar to Greece.

However, even when we restrict the sample to Portugal, Spain, Ireland and Italy, we find no pure contagion. The results are quite similar to the baseline estimation (Table 2, column 2). While Portugal, Spain, Ireland and Italy are conventionally considered the "periphery" of the euro area, the results are qualitatively unchanged when we include Belgium or both Belgium and France together in the periphery

Bias-corrected estimates. - Since Nickell (1981), it is well known that the LSDV estimator is biased when used in dynamic panels. While the fact that this bias decreases with the length of the panel should be reassuring, given our very long sample period, we also experimented with the Kiviet (1995) estimation technique, which appears to be particularly appropriate for macroeconomic (i.e. big T/small N) panels (Judson and Owen, 1999). It turns out that the bias-corrected estimates are basically identical to our baseline.[13]

4.2 Testing for unit roots and cointegration

A legitimate issue with the econometric analysis presented in Section 4.1, given the observed high persistence of the spreads, is that they could actually be non-stationary. Indeed, performing common panel unit root tests such as those proposed by Levin Lin and Chu and by Pesaran, Im and Shin (see Banerjee, 1999, Baltagi, 2008, and Choi, 2006), we could not reject the null of integration for the sovereign spreads (Table 3, top panel). This result is robust even if we compute the relevant test statistics using different lag structures and different time spans. In particular, unit roots appear

[13] Results are not shown.

to be present not only if we look at the full sample, or at the post-crisis period, but also when we restrict the analysis to the pre-crisis period.[14]

We also tested for the existence of a cointegrating relationship between the spread and its determinants. In particular, we adopted the residual-based approach by Kao and Pedroni (see Banerjee, 1999, and Baltagi, 2008). While the results are consistent with the existence of a cointegrating vector, they are not very clear-cut (Table 3, bottom panel).

4.3 Non-stationary case

In this section we model the long-run relationship between spreads and fundamentals as:

$$s_{it} = \alpha_{i0} + \beta_0 Z_{it} + \beta_1 F_t + \gamma_0 D_t + \gamma_1 D_t Z_{it} + \gamma_2 D_t F_t + \varepsilon_{it}, \tag{3}$$

therefore allowing for a structural change in the relationship in the post-crisis period, and for the different kinds of contagion effects highlighted in the previous sections. As before, the error term is assumed independent across countries but possibly heteroskedastic and auto-correlated.

To estimate equation (3), we resort to different methods, in order to check the robustness of the results to different statistical assumptions.

First, we run a simple LSDV regression. Indeed, if spreads are I(1) and there is no cointegrating relationship between spreads and fundamentals, i.e., ε_{it} in equation (3) is I(1), the LSDV estimator delivers consistent estimates of the long-run average relationship between them, contrary to the pure time-series case (Phillips and Moon, 1999, Phillips and Moon, 2000, and Baltagi, 2008).

The results are qualitatively similar to those obtained with the stationary model, but much more pronounced and clear-cut (Table 2, column 3). Before the crisis, all the fundamentals are significant with economically meaningful signs, except GDP growth (which is not significant) and the current account surplus (which has the wrong sign). After the start of the crisis, the effect on the spread is magnified and with the expected sign for all the fundamentals. In particular, the effect of GDP growth and of the current account surplus becomes significant and negative, as it should be if markets correctly assess sovereign creditworthiness. Also, shift contagion (i.e. an increased post-crisis role of the VIX) emerges.

If spreads are I(1) but there exists a cointegrating relationship between spreads and fundamentals, i.e. ε_{it} in equation (3) is I(0), it can be shown that OLS estimates are inconsistent. We therefore estimate equation (3) using the panel dynamic least square (DOLS) estimator proposed by Kao and Chiang (2000), which extends to panel data the approach of Saikkonen (1991) and Stock and Watson (1993). That is, estimates of the coefficients of interest are found by running the

[14] This suggests some caution in interpreting the results of previous papers, which did not consider the issue.

following OLS regression:

$$s_{it} = \alpha_{i0} + \beta_0 Z_{it} + \beta F_t + \gamma_0 D_t + \gamma_2 D_t Z_{it} + \gamma_3 D_t F_t + \sum_{j=-2}^{2} \delta_{0j} \Delta Z_{it+j} + \sum_{j=-2}^{2} \delta_{1j} \Delta F_{t+j} + \varepsilon_{it}, \quad (4)$$

where the inclusion of ΔZ_{it+j} and ΔF_{t+j} among the regressors helps to get a consistent estimate of the βs and the γs. The results are remarkably similar to those of the previous exercise (Table 2, column 4).

As a final exercise, we consider a model with random, instead of fixed, individual effects. As shown by Baltagi et al. (2008, 2011), to this end the best available option is to estimate equation (3) with feasible generalized least squares (notice that this holds irrespective of whether ε_{it} is I(0) or I(1)). The results are qualitatively similar to those obtained with the fixed-effects specification (Table 2, column 5).

5 Robustness checks

5.1 Using different proxies

As a first robustness exercise, we consider two alternative measures of liquidity. One, used by Attinasi et al. (2009), among others, is the country's share of the euro-area long- and medium-term sovereign bond issuance. The other is the monthly average of the traded volumes of the country's government securities with maturity between nine and eleven years relative to Germany's, used for example by Codogno et al. (2003). In both cases, we found liquidity to be statistically insignificant, both alone and interacted with the crisis dummy.

As a second check, we experiment with a different proxy for global risk aversion and, following Codogno et al. (2003) and Bernoth et al. (2012), we substitute the VIX with the yield spread between low-rated (BBA) US corporate bonds and the US Treasuries of corresponding maturity, without any notable effect on the results.

5.2 Controlling for banking sector stress

As is commonly acknowledged, in several EMU countries worries about public debt sustainability were magnified by concerns about the state of the banking sector. While the role of banks in the EMU crisis is not the focus of this paper, it is important to control for this channel.

To do this, we first add to our baseline regressions a measure of domestic banks' credit risk, proxied by the CDS banking index, to account for the negative feedback effects from the banking to the government sector.[15] Both in the stationary and in the non-stationary models, the absence

[15] We define the CDS banking index as the simple average of all the CDS premia on banks resident in a given

of pure contagion and the presence of wake-up-call contagion are robust to the inclusion of the new variable. The latter is significant and has the expected sign, except for the stationary specification. That is, an increase in the country's CDS banking index increases the country's sovereign spread as well. However, the effect does not appear to have increased in the post-crisis period.

Alternatively, we introduced in our regressions, as a factor common to all countries (therefore included in the F_t vector together with the VIX), the spread between the three-month euro interbank offered rate (Euribor) and the corresponding OIS swap rate (which captures the market's expectations of the overnight funds rate). This difference is considered a gauge of fears of bank insolvency (see e.g. Thornton, 2009). Contrary to country-specific CDS premia, this regressor becomes much stronger after the crisis, suggesting that the crisis gave rise to widespread concern about the health of the European banking system as a whole. In any case, even in these richer specifications we still find wake-up-call contagion, while we do not find pure contagion.

5.3 The definition of the contagious event

A possible pitfall of our analysis is that it relies on a sharp hypothesis concerning the start of the EMU sovereign crisis, although we do find that changing the moment of the structural break from October 2009 to May 2010 (when the euro area countries launched the first Greek bail-out programme) or to November 2010 (when for the first time EU authorities officilally envisaged the possibility of private sector involvement in sovereign debt crises resolution) does not drastically change the estimation results. Moreover, a dichotomous crisis dummy cannot capture changes in the intensity of the crisis.

We address both problems by using, instead of our crisis dummy, a variable summarizing the Greek credit rating; we borrow this approach from Gande and Parsley (2005) and De Santis (2012). In particular, we transform the sovereign credit rating information (expressed in letters) of the three major credit rating agencies (Fitch, Moody's and Standard & Poor's) into a numerical variable using a linear scale. The variable takes 22 values from 1 (triple-A) to 22 (selective default). We also take credit-watch changes into consideration: a negative credit watch increases the value of the variable by 0.5 while a positive credit watch corresponds to a decrease of 0.5. We use the average of the numerical indicators computed for the three main rating agencies.

The results are analogous to our baseline regressions (Table 4, columns 1-4). In particular, the only fundamental variable which is statistically significant when taken in isolation is GDP growth. When interacted with the Greek rating variable, instead, government debt and the current account surplus also become significant, as in the baseline regression. In particular, the analysis shows that a worsening of the situation in Greece magnifies the positive effect of a current account surplus and the negative effect of public debt on the spreads of the other EMU countries. Finally, as in

country which are available in the Thomson Financial Reuters database. Due to lack of banks' CDS data, we drop Finland from the sample.

our baseline model, the Greek fiscal situation index, taken alone, has no effect on other countries' spreads.

5.4 EU policy-makers at work

In the months following the crisis, EU authorities announced and implemented several crisis-management interventions. While the efforts to improve the euro-area crisis management framework have continued after the end of our sample period (see e.g. the ECB's OMT, announced in August 2012), during our sample period three major policy episodes can be singled out.

- After several weeks of discussion, the turning-point in the EU authorities' approach to the Greek crisis came in the spring of 2010. On 2 May the euro-area countries agreed on a three-year financial support plan that provided bilateral loans to Greece. On 10 May, the EU Council established the EFSF, a vehicle empowered to issue securities guaranteed by euro-area countries and to provide loans to countries experiencing severe financial disturbance (loans are provided under conditions similar to those applied by the IMF). On the same day, the ECB launched the SMP, a programme of purchases of public and private debt securities issued in the euro area to support segments of the market especially hard hit by the crisis.
- On 28 November 2010, the euro-area finance ministers agreed to institute the ESM, a permanent crisis management tool, which is due to replace the EFSF, providing financial support to countries that request assistance subject to strict conditions. Assistance is also subject to a rigorous debt sustainability analysis. Member states considered insolvent would have to negotiate a restructuring plan with private creditors. On the same day, the finance ministers also decided to grant support to Ireland through the EFSF.
- On 21 July 2011, the Council agreed on a new Greek assistance programme, which included a sizable bail-in for private investors (with estimated losses amounting to €50 billion).

These policy actions may have influenced sovereign debt markets. To investigate this issue, we augment our empirical models with three event dummies, set equal to one in May 2010 (creation of the EFSF and launch of the SMP), December 2010 (creation of the ESM) and July 2011 (Greek private sector involvment), respectively. Introducing the event dummies does not change the economic and statistical significance of the other coefficients (Table 5, columns 1-4). However, non-conventional actions of EU policy-makers had an impact. In particular, as expected, the actions taken in May 2010 eased the tensions on the sovereign debt markets, and the involvement of the private sector in the Greek debt restructuring increased spreads. The results concerning the announcement of the ESM are somewhat less obvious, as that policy dummy is either insignificant (Table 5, columns 1 and 2) or significant with a positive sign (Table 5, columns 3 and 4). This indicates that the replacement of the temporary EFSF with the permanent ESM did not calm the markets, possibly owing to the news that Ireland as well as Greece had lost market access and had to be bailed-out, or to the official announcement that private sector involvement would be a permanent feature of the EU crisis resolution mechanism in the future.

5.5 A richer set of common factors

Ideally, one would like to control completely for unobserved time-varying common factors with a full set of time dummies. In practice, however, this would drastically reduce the degrees of freedom of our estimation. Moreover, the crisis dummy, which is the focus of our analysis, would be collinear with these dummies.[16] However, we can go some way in accounting for common time trends by enriching our vector of controls. In particular, we add to our F_t vector two further variables: (1) the monetary policy rate set by the ECB (i.e. the interest rate on main refinancing operations); (2) an index of economic policy uncertainty for Europe computed recently by Baker et al. (2013). This second addition is quite interesting for its own sake. According to this index, economic policy uncertainty increased on average by 48% in the crisis period. We show that this richer specification leaves our results unaffected (Table 6, columns 1-4). The two common factors appear significant in some but not all of the models that we estimate. They display the expected signs: both a tightening of monetary policy and an increase in policy uncertainty tend to increase sovereign spreads.

6 Computing the long-run level of sovereign spreads

Equation (3) can be rewritten applying the Oaxaca-Blinder decomposition to the crisis-induced change in spreads, as in Eichengreen and Mody (2000). That is, the difference between the pre-crisis and the crisis spread can be decomposed into two parts: one due to a change in the regressors, the other due to a change in the coefficients. The change in the constant term is what we identify as the "pure" contagion effect. Conditional on the occurrence of the crisis, one gets:

$$\begin{aligned} E(s_{it}^{LR}|D_{it} &= 0) = \alpha_{0i} + \beta_0 E(Z_{it}|D_{it}=0) + \beta_1 E(F_t|D_{it}=0)\ , \\ E(s_{it}^{LR}|D_{it} &= 1) = \alpha_{0i} + \gamma_0 + (\beta_0+\gamma_1)\, E(Z_{it}|D_{it}=1) + (\beta_1+\gamma_2)\, E(F_t|D_{it}=1), \end{aligned}$$

[16] Incidentally, this is why Bernoth at al. (2012) cannot allow for pure contagion.

where the LR superscripts serve as a reminder that we are considering here the long-run equilibrium values of the spread. Therefore, the post-crisis long-run value of the spread is equal to:

$$\begin{aligned} E(s_{it}^{LR}|D_{it} &= 1) = E(s_{it}^{LR}|D_{it} = 0) + \\ &\beta_0 \left[E(Z_{it}|D_{it} = 1) - E(Z_{it}|D_{it} = 0)\right] + \beta_1 \left[E(F_t|D_{it} = 1) - E(F_t|D_{it} = 0)\right] + \\ &+ \underbrace{\gamma_0}_{pure} + \underbrace{\gamma_2 E(Z_{it}|D_{it} = 1)}_{wake-up\text{-}call} + \underbrace{\gamma_3 E(F_t|D_{it} = 1)}_{shift} . \end{aligned} \tag{5}$$

Terms in the second row capture the post-crisis change in fundamentals, while terms in the third row capture the different kinds of contagion: γ_0 is what we call pure contagion and is unrelated to country characteristics; $\gamma_2 E(Z_{it}|D_{it} = 1)$ captures wake-up-call contagion, is country-specific and depends on fundamentals; $\gamma_3 E(F_t|D_{it} = 1)$ is the shift-contagion component.

We use the estimates presented in Section 4.3 to compute the various pieces of equation (5). We first consider, for each country, the estimated value of $E[s_{it}^{LR}|D_t = 0]$ (Table 7, column 1). We then add to this value the terms in the second line of equation (5) (Table 7, column 2). To compute those values it is necessary to assess the pre- and post-crisis values of the fundamentals and of the VIX. In the table, we put them equal to their respective sample counterparts. Finally, we add the contagion terms, and we get to $E[s_{it}^{LR}|D_t = 1]$ (Table 7, column 3).

According to our calculations, for most countries the spreads observed at the end of the sample period (December 2011) are very close to their estimated long-run levels. However, for two countries, namely Spain and Italy, they are considerably above their equilibrium values (Figure 2).

7 Conclusions and policy implications

The analyses presented in this paper suggest that investors largely ignored macroeconomic indicators when pricing sovereign bonds before October 2009. At that date they started to discriminate among sovereigns based on the quality of their fundamentals. In particular, countries with worse fiscal conditions and external positions recorded higher spread levels. In the terminology adopted in this paper, the sharp increase in spreads observed for some countries after the start of the Greek crisis was the result of a wake-up-call rather than of a pure form of contagion: the Greek crisis increased investors' sensitivity to the fundamentals of the other euro-area countries.

Concerning the policy implications of our results, the fact that for some countries the current spread levels are above their long-run values argues for policy measures to speed up the convergence of spreads towards their long-run levels. It must be stressed that the absence of pure contagion, per se, does not settle the normative issue concerning the investors' ability to price sovereign bonds correctly.[17] We cannot say, for example, whether the increased post-crisis sensitivity to fundamen-

[17] Symmetrically, the existence of contagion does not imply malfunctioning of the markets. This is particularly true

tals is “appropriate”: it could also be "too limited" or “excessive”. Answering this question would be important in implementing the OMT. More broadly, it would help settle the debate about the relative merits of market-based as against rules-based fiscal and macroeconomic discipline, which is as old as the very idea of EMU. Indeed, already in 1989 the Delors report worried that market forces "might be either too slow and weak or too sudden and disruptive". Further research on this issue, both theoretical and empirical, is warranted.

Another related question is the possible reoccurence of a regime in which investors do not pay attention to fundamentals. To avoid disruptive cycles of excessive complacency and sudden wake-up calls, it seems advisable to push for market-friendly policies that highlight the fundamental imbalances of EMU countries even in good times. This is the rationale behind the decision to periodically publish scoreboards prepared by the European Commission and the results of the Macroeconomic Imbalance Procedure. Needless to say, the variables included and the methodology adopted in such exercises should be based on sound economic principles.[18]

in the case of wake-up-call and shift contagion. For example, rational inattention stories would imply that markets are constrained-efficient, once the limits in information processing are taken into account. It appears more difficult, but not impossible, to reconcile "pure" contagion with market efficiency and/or with full rationality (Kyle and Xiong, 2001, Kodres and Pritsker, 2002).

[18] Another avenue for further research would be to investigate whether the risk of the break-up of the euro area influences sovereign debt spreads. Di Cesare et al. (2012) point out that this risk began to be perceived by investors in 2012, therefore after the end of our sample.

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8 Figures and tables

Figure 1 – Yield spreads between ten-year government bonds and the German Bund (basis points)

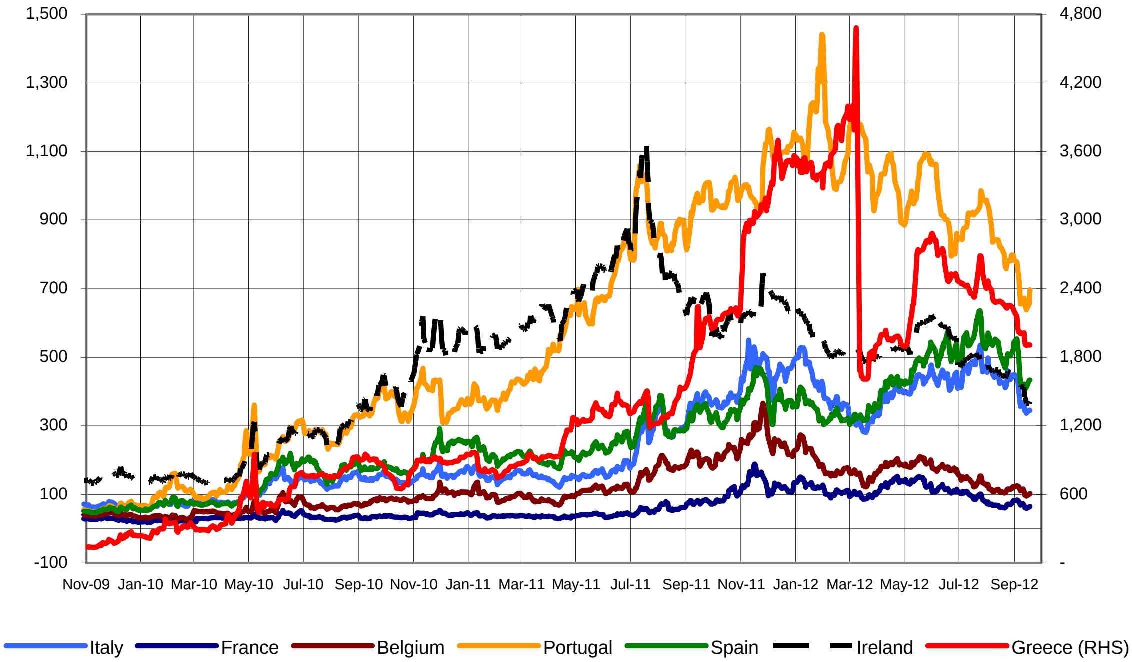

Table 1 – Descriptive statistics

	Mean	St.dev	Min	Max	Mean	St.dev	Min	Max
	January 2000 – October 2009				November 2009 – December 2011			
				Overall sample				
Sovereign spread (bp)	19.3	27.9	22.1	242.4	174.9	220.0	12.3	1109.3
Bid-Ask spread (bp)	1.0	0.8	0.2	6.0	5.5	14.8	0.3	85.4
Risk aversion (VIX)	25.9	10.9	12.6	63.3	28.0	7.1	20.1	45.6
Public debt/GDP×100	64.0	24.5	24.5	117.0	81.3	22.0	43.5	121.0
Private debt/GDP×100	162.0	42.8	75.2	303.1	204.4	49.3	125.3	303.4
GDP growth %	1.8	3.0	9.8	12.4	1.1	2.0	6.5	5.8
Current account surplus/GDP×100	0.5	5.5	13.3	11.9	0.7	5.0	13.3	11.7
				Ireland, Italy, Spain and Portugal				
Sovereign spread (bp)	25.7	35.6	22.1	242.4	328.1	253.5	52.7	1109.3
Bid-Ask spread (bp)	1.3	0.9	0.3	3.7	11.0	20.9	0.6	85.4
Risk aversion (VIX)	25.9	10.9	12.6	63.3	28.0	7.1	20.1	45.6
Public debt/GDP×100	63.2	28.8	24.5	117.0	92.1	22.2	53.9	121.0
Private debt/GDP×100	164.8	52.6	75.2	303.1	222.2	61.1	125.3	303.4
GDP growth %	2.0	3.3	8.3	12.4	0.1	1.5	5.5	2.2
Current account surplus/GDP×100	4.8	4.1	13.3	1.9	3.9	4.0	13.3	4.2
				Austria, Belgium, Finland, France and the Netherlands				
Sovereign spread (bp)	14.2	18.3	15.8	108.2	52.4	45.8	12.3	292.0
Bid-Ask spread (bp)	0.9	0.6	0.2	6.0	1.2	0.9	0.3	4.1
Risk aversion (VIX)	25.9	10.9	12.6	63.3	28	7.1	20.1	45.6
Public debt/GDP×100	64.7	20.4	29.9	115.6	72.7	17.7	43.5	100.0
Private debt/GDP×100	159.8	34.2	16.2	98.7	190.1	30.8	156.8	242.3
GDP growth %	1.7	2.6	9.8	6.4	1.9	1.9	6.5	5.8
Current account surplus/GDP×100	2.9	3.6	8.6	11.9	1.9	4.2	6.0	11.7

Table 2 – Regression results

	(1)		**(2)**		**(3)**		**(4)**		**(5)**	
spread(t-1)	0.927	***	0.930	***						
	(0.035)		(0.037)							
general government debt	-0.018		-0.088		1.211	***	1.120	***	0.337	***
	(0.116)		'(0.147)		(0.295)		(0.258)		(0.0671)	
private debt	0.050		0.043		0.926	***	0.939	***	0.167	***
	(0.040)		'(0.031)		(0.077)		(0.080)		(0.039)	
GDP growth	-0.542	**	-1.062	***	-0.077		-1.276		-2.341	***
	(0.27)		'(0.408)		(0.639)		(0.783)		(0.825)	
current account surplus	0.147		0.416		2.610	***	2.619	***	-0.351	
	(0.135)		'(0.308)		(0.369)		(0.392)		(0.246)	
liquidity (bid-ask)	0.422		1.480	*	7.751	***	7.659	***	10.998	***
	(0.561)		'(0.835)		(1.342)		(1.454)		(1.824)	
VIX	0.152	***	0.191	***	0.676	***	0.603	***	0.960	***
	(0.027)		'(0.046)		(0.077)		(0.107)		(0.131)	
Dummy crisis	-15.128		-43.819		-84.738	***	-85.365	***	-95.619	***
	(10.377)		'(35.894)		(25.716)		(23.346)		(15.467)	
spread(t-1)×crisis	0.083		0.061							
	(0.052)		'(0.073)							
public debt×crisis	0.151	*	0.543	*	1.381	***	1.300	***	1.388	***
	(0.091)		'(0.294)		(0.275)		(0.247)		(0.168)	
private debt×crisis	0.044		0.139		0.337	**	0.293	**	0.649	***
	(0.047)		'(0.115)		(0.138)		(0.121)		(0.080)	
GDP growth×crisis	-3.193		-7.274		-26.123	***	-21.603	***	-29.393	***
	(2.090)		'(5.019)		(3.614)		(3.231)		(1.965)	
current account surplus×crisis	-0.871	*	-1.909		-4.597	***	-4.249	***	-5.282	***
	(0.524)		'(1.333)		(1.219)		(1.124)		(0.673)	
liquidity×crisis	-0.594		-1.657		0.065		0.064		-2.470	
	(0.769)		'(1.018)		(1.507)		(1.463)		(1.840)	
VIX×crisis	0.198		0.192		2.174	**	2.204	***	2.007	***
	(0.345)		'(0.893)		(0.882)		(0.825)		(0.462)	
R^2	0.98		0.98		0.87		0.89		0.85	
Observations	1,269		564		1,269		1,242		1,269	

Notes: Columns 1,2,3: LSDV; Column 4: DOLS (1 lead and 1 lag added for each variable; country dummies incl.); Column 5: FGLS. All estimations except column 5: Huber-white robust standard errors in parentheses. All estimations except column 2: full sample (Column 2: sample limited to the periphery countries: PT, IT, IR, ES). *: significant at the 10% level; ** at the 5%; *** at 1%.

Table 3 – Unit root and Cointegration Tests

Panel unit root tests	
*Levin, Lin and Chou t**	15.940
H0: unit roots for all i's(H1: no unit root)	(1,000)
Im, Pesaran and Shin W-stat	11.970
H0: unit roots for all i's (H1: some unit roots)	(1,000)
Panel cointegration tests	
ADF statistic (Pedroni 1)	-1.642
H0: no cointegration (H1 assumes common autocorr. coefficient)	(0,0503)
ADF statistic (Pedroni 2)	-1.170
H0: no cointegration (H1 allows country-specific autocorr. coefficients)	(0,121)

Notes: P-values in parentheses; number of lags =1.

Figure 2 – Cointegrated model: predicted values (dashed lines: 95% conf. bands)

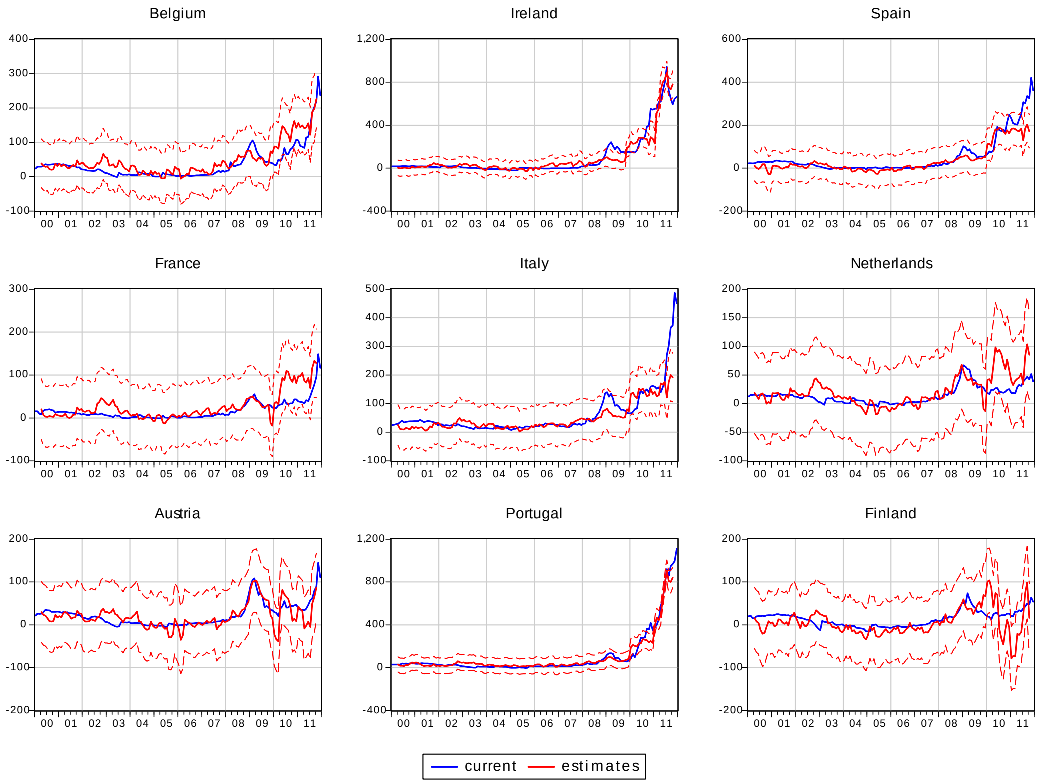

Table 4 – Regression results (continuous crisis variable)

	(1)		(2)		(3)		(4)	
spread(t-1)	0.947	***						
	(0.040)							
general government debt	0.179		2.088	***	0.625	***	2.000	***
	(0.112)		(0.278)		(0.057)		(0.256)	
private debt	0.087	**	1.117	***	0.388	***	1.102	***
	(0.043)		(0.073)		(0.032)		(0.070)	
GDP growth	-1.172	**	-3.301	***	-6.520	***	-3.483	***
	(0.516)		(0.825)		(0.705)		(0.870)	
current account surplus	0.068		1.599	***	-1.360	***	1.494	***
	(0.166)		(0.354)		(0.217)		(0.361)	
liquidity (bid-ask)	1.413		6.517	***	10.154	***	7.141	***
	(1,144)		(1.816)		(1.015)		(1.604)	
VIX	0.101		0.604	***	0.890	***	0.597	***
	(0.068)		(0.124)		(0.121)		(0.141)	
Greek rating	-0.238		-4.747	*	-4.120	**	-5.920	**
	(1,626)		(2.857)		(1.728)		(2.611)	
public debt $\times$ Greek rating	0.028	*	0.165	***	0.176	***	0.171	***
	(0.016)		(0.032)		(0.019)		(0.030)	
private debt $\times$ Greek rating	0.009		0.073	***	0.105	***	0.074	***
	(0.011)		(0.230)		(0.011)		(0.022)	
GDP growth $\times$ Greek rating	-0.743		-2.863	***	-2.958	***	-2.144	***
	(0.484)		(0.721)		(0.308)		(0.652)	
current account $\times$ Greek rating	-0.126	*	-0.898	***	-0.920	***	-0.881	***
	(0.076)		(0.128)		(0.084)		(0.132)	
liquidy $\times$ Greek rating	-0.132		-0.196		-0.403	***	-0.236	*
	(0.097)		(0.154)		(0.081)		(0.138)	
VIX*Greek rating	-0.010		0.055		-0.010		0.088	
	(0.044)		(0.079)		(0.043)		(0.074)	
R^2	0.98		0.91		0.87		0.93	
observations	1,269		1,269		1,269		1,242	

Notes: Columns 1,2: LSDV; Column 3: DOLS (1 lead and 1 lag added for each variable; country dummies incl.); Column 4: FGLS. All estimations except column 4: Huber-white robust standard errors in parentheses. *: significant at the 10% level; ** at the 5%; *** at 1%.

Table 5 – Regression results (policy dummies)

	(1)		(2)		(3)		(4)	
spread(t-1)	0.926	***						
	(0.034)							
general government debt	0.007		1.299	***	0.337	***	1.222	***
	(0.113)		(0.288)		(0.065)		(0.249)	
private debt	0.047		0.900	***	0.167	***	0.894	***
	(0.038)		(0.073)		(0.037)		(0.075)	
GDP growth	-0.550	**	-0.107		-2.340	***	-1.120	
	(0.269)		(0.637)		(0.795)		(0.743)	
current account surplus	0.102		2.338	***	-0.351		2.263	***
	(0.131)		(0.346)		(0.237)		(0.352)	
liquidity (bid-ask)	0.430		7.758	***	11.000	***	7.636	***
	(0.557)		(1.312)		(1.758)		(1.424)	
VIX	0.151	***	0.677	***	0.960	***	0.589	***
	(0.027)		(0.076)		(0.126)		(0.102)	
Dummy crisis	-20.962	**	-142.003	***	-154.590	***	-138.957	***
	(10.37)		(25.854)		(16.135)		(24.425)	
spread(t-1)×crisis	0.079							
	(0.052)							
Public debt×crisis	0.164	*	1.380	***	1.387	***	1.291	***
	(0.092)		(0.261)		(0.163)		(0.233)	
Private debt×crisis	0.053		0.376	***	0.686	***	0.330	***
	(0.046)		(0.131)		(0.078)		(0.114)	
GDP growth×crisis	-3.692	*	-27.587	***	-30.780	***	-22.798	***
	(2.043)		(3.607)		(1.904)		(3.207)	
Current account surplus×crisis	-0.986	*	-5.307	***	-6.055	***	-4.964	***
	(0.536)		(1.180)		(0.655)		(1.088)	
liquidity×crisis	-0.689		-0.504		-3.001	***	-0.448	
	(0.743)		(1.458)		(1.775)		(1.414)	
VIX×crisis	0.293		3.908	***	3.770	***	3.801	***
	(0.349)		(0.882)		(0.485)		(0.856)	
May 2010	8.488		-107.158	***	-107.390	***	-90.014	***
	(7.750)		(16.328)		(15.991)		(15.443)	
December 2010	-11.762		40.575		49.323	***	66.071	**
	(11.366)		(27.965)		(15.251)		(28.857)	
July 2011	54.110	***	102.693	***	106.139	***	107.377	***
	(16.942)		(30.333)		(15.216)		(27.415)	
R^2	0.98		0.88		0.86		0.91	
Observations	1,269		1,269		1,269		1,242	

Notes: Columns 1,2: LSDV; Column 3: DOLS (1 lead and 1 lag added for each variable; country dummies incl.); Column 4: FGLS. All estimations except column 4: Huber-White robust standard errors in parentheses. *: significant at the 10% level; ** at the 5%; *** at 1%.

Table 6 – Regression results (more common factors)

	(1)		(2)		(3)		(4)	
spread(t-1)	0.920	***						
	(0.035)							
general government debt	0.012		1.381	***	0.312	***	1.333	***
	(0.121)		(0.303)		(0.065)		(0.258)	
private debt	0.059		0.873	***	0.144	***	0.901	***
	(0.039)		(0.072)		(0.038)		(0.074)	
GDP growth	-0.418		-0.725		-2.515	***	-1.684	**
	(0.294)		(0.684)		(0.819)		(0.776)	
current account surplus	0.118		2.501	***	-0.299		2.411	***
	(0.132)		(0.357)		(0.241)		(0.347)	
liquidity (bid-ask)	0.167		6.702	***	9.508	***	6.657	***
	(0.537)		(1.304)		(1.820)		(1.399)	
VIX	0.076	***	0.250	***	0.252		-0.007	
	(0.029)		(0.093)		(0.225)		(0.139)	
policy unceirtanty	0.037	***	0.204	***	0.349	***	0.045	
	(0.013)		(0.038)		(0.090)		(0.062)	
monetary policy rate	0.910	***	-1.048		-0.001		0.832	
	(0.249)		(0.738)		(1.503)		(0.775)	
dummy crisis	-84.379	***	-244.561	***	-231.860	***	-262.364	***
	(21.928)		(55.081)		(28.368)		(50.019)	
spread(t-1)×crisis	0.082							
	(0.050)							
public debt×crisis	0.167	*	1.499	***	1.511	***	1.411	***
	(0.091)		(0.255)		(0.164)		(0.220)	
private debt×crisis	0.069		0.416	***	0.751	***	0.381	***
	(0.047)		(0.133)		(0.079)		(0.111)	
GDP growth×crisis	-3.053		-24.423	***	-28.462	***	-20.492	***
	(1.947)		(3.449)		(1.920)		(3.027)	
current account surplus×crisis	-1.067	**	-5.001	***	-5.701	***	-4.776	***
	(0.544)		(1.201)		(0.6576)		(1.108)	
liquidity×crisis	-0.455		0.314		-1.604		0.157	
	(0.729)		(1.484)		(1.837)		(1.424)	
VIX×crisis	-2.355	**	-1.025		-0.703		-0.855	
	(1.005)		(1.699)		(0.955)		(1.678)	
policy unceirtanty×crisis	0.669	**	0.168		0.049		0.170	
	(0.282)		(0.559)		(0.266)		(0.552)	
monetary policy rate×crisis	38.309	**	190.433		165.896	***	212.190	***
	(16.755)		(39.688)		(21.661)		(29.616)	
R^2	0.98		0.88		0.85		0.92	
Observations	1,269		1,269		1,269		1,242	

Notes: Columns 1,2: LSDV; Column 3: DOLS (1 lead and 1 lag added for each variable; country dummies incl.); Column 4: FGLS. All estimations except column 4: Huber-White robust standard errors in parentheses. *: significant at the 10% level; ** at the 5%; *** at 1%.

Table 7 – Long-run values of the spread (basis points)

	Fitted values		
Coefficients	*pre-crisis*	*pre-crisis*	*post-crisis*
Fundamentals	*pre-crisis*	*post-crisis*	*post-crisis*
Italy	24	47	247
Austria	23	43	131
Belgium	21	45	210
Finland	0	16	81
France	7	32	175
Ireland	35	335	558
Portugal	46	257	507
Spain	28	97	269
Netherlands	10	35	134

Notes: Spreads computed with coeff. from Table 2, col. 4 (DOLS).

The elasticity of demand for sovereign debt. Evidence from OECD Countries (1995-2011)

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May 2013

Abstract

Public debt levels in advanced economies have increased dramatically over recent years and could put considerable upward pressure on market yields. Using a novel identification approach based on financial accounts and focusing on 18 advanced economies over the period 1995–2011, this paper estimates the long-term elasticity of the demand for government securities. We find that public debt does matter: each percentage point increase in the ratio of public debt to GDP raises 10 year rates by about 3 basis points. The potential drag on growth caused by public debt through higher interest rates should thus not be overlooked.

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1. Introduction

The sharp run-up in public sector debt in advanced economies is likely to be one of the most enduring legacies of the 2007-09 global financial crisis.[1] A key policy question is at what interest rates foreign and domestic investors will be willing to hold such increasing amounts of government debt. So far, investors' preference for safe assets has sustained the demand for government securities, while in some countries unconventional monetary policies have contributed to relieve the pressure of bond supply on bond prices, thus diluting the effects of inflated public deficits over time. Eventually, however, all this newly created supply of government debt will be on the market and investors might start requiring higher yields in order to keep it in their portfolios.

To what extent could interest rates increase? To answer this question we need some measure of the elasticity of demand for sovereign debt to interest rates. The abundant empirical literature on the impact of fiscal variables on interest rates mostly relies upon reduced-form equations, which give biased estimates of the demand elasticity, especially in periods characterized by large shifts in the non-interest sensitive demand for bonds. The main contribution of this paper is to solve the identification problem by resorting to financial account statistics. We disentangle the long-term from the reduced-form demand curve by using as shifters of demand the financial accounts balances of three institutional sectors: households, non-financial firms and the foreign sector. We also control for foreign official reserves and the gross assets of financial institutions, following a recent strand of research highlighting the importance of gross (rather than net) capital flows in determining financial conditions.[2] Finally, we allow for shifters of demand (e.g., sovereign ratings) that capture the degree of substitutability of sovereign debt with other assets due to credit risk concerns.

For a panel of 18 advanced economies over the period 1995-2011, we find that the level of public debt does matter for interest rates: an increase by one percentage point in the public debt-to-GDP ratio raises 10-year rates by about 3 basis points. This is in line with results available for the United States, but based on reduced-form equations estimated on sample periods that do not extend beyond the mid-2000s.

The paper is structured as follows. Section 2 reviews the literature and Section 3 describes the data base. The identification framework is presented in Section 4, while Section 5 reports the econometric estimates. Some robustness checks are discussed in Section 6. Section 7 draws some conclusions.

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[1] Reinhart and Rogoff (2011). See also Cecchetti, Mohanty and Zampolli (2011).

[2] Borio and Disyatat (2011) and Shin (2012).

2. The relationship between fiscal variables and interest rates in previous studies

The subject of the impact of fiscal variables on interest rates has long been a major theme in macroeconomic theory and policy debate. Considering the last thirty years, studies flourished in the eighties and early nineties when in the United States public debt relative to GDP was raising rapidly. The debate was heavily influenced by debt sustainability considerations at that time, as in Blanchard (1984), Hamilton and Flavin (1986), Bohn (1995). Research was stimulated also by the rational expectation revolution in economic theory, which led macroeconomists to investigate public debt irrelevance propositions such as the Ricardian Equivalence hypothesis within dynamic rational equilibrium models (Barro (1989)). Interest in the issue was rekindled in the early 2000s, once again in a period characterized by a large expansion of public debt in the United States. A review of the debate can be found in Gale and Orszag (2004), Liungqvist and Sargent (2004), Engen and Hubbard (2004) and Haugh, Ollivaud and Turner (2009).

A worsening of public finances can affect medium- and long-term yields through three main channels.[3] First, if the supply of savings is not perfectly elastic, financing the budget deficit has to compete for resources with the demand for funding of the private sector, causing real interest rates to rise.[4] Second, an increase in the public debt may cause fears that even sovereign borrowers may default, leading to increased credit risk premiums on government bonds. Third, a larger deficit may fuel expectations of inflation or exchange-rate depreciation, with additional repercussions on interest rates.

While a strand of research focuses on sovereign credit risk premia,[5] most of the large empirical literature tries to assess the overall effect of fiscal imbalances on interest rates without distinguishing among the three channels. The econometric framework normally relies on reduced-form regressions. The fiscal variable of interest can be either public debt or public deficit; in several papers, both variables are interchangeably tried and compared. The majority of studies, however, focus on public deficit, because public debt is rarely significant at conventional confidence levels.

[3] See Ardagna, Caselli and Lane (2007), Balassone, Giordano and Franco (2004) and the Box "The effects of the public debt on long-term interest rates" in Banca d'Italia (2010).

[4] As pointed out by Ardagna, Caselli and Lane (2010), it is useful to distinguish between shorter- and longer-run effects. In an economy in which there is some degree of short-run nominal stickiness, a weakening in the primary fiscal balance adds to aggregate demand and leads to an increase in nominal and real short-term interest rates. Insofar as the adjustment of nominal prices is gradual and the primary fiscal balance's deterioration is perceived to be persistent, the increase in short-term interest rates feeds through medium- and long-term interest rates. In the longer run, to the extent that fiscal expansion crowds out private investment and results in a lower steady-state capital stock, it will be associated with a higher marginal product of capital and thus a higher real interest rate. For an analysis of the long-run implications of rising public debt for interest rates see Engen and Hubbard (2004).

[5] That approach is not pursued here. Reviews of recent studies related to the euro-area sovereign debt crisis can be found in, among others, Di Cesare, Grande, Manna and Taboga (2012) and Favero (2013). For earlier analyses, see Codogno, Favero and Missale (2003).

The econometric models differ considerably also in terms of the other explanatory variables considered, functional specification, estimation method, sample period and sample countries. Three of the most representatives studies are Engen and Hubbard (2004), Laubach (2009) and Ardagna, Caselli and Lane (2007).

Engen and Hubbard (2004) provide a useful discussion of the appropriate specification of the reduced-form equation. First, they argue that, in a closed production economy with a standard Cobb-Douglas technology, public debt affects interest rates because it replaces, or crowds out, productive physical capital and thus raises the marginal productivity of capital. For this reason, an appropriate specification is to regress the *level* of interest rates on the *stock* of public debt. An alternative specification is to regress the *change* in interest rates on the *change* in public debt (i.e. government borrowing or the public deficit). A third, widely used, specification in which the *level* of interest rates is regressed on the *change* in public debt is instead less consistent with what an economic model of crowding out would suggest and can be justified only by assuming sluggish nominal price adjustment and a persistent deterioration in the fiscal position.

Second, Engen and Hubbard (2004) make clear that, in open monetary economies, the substitution of public debt for capital may be less than one-to-one because part of the supply of government bonds may be met by the demand stemming from foreign investors and the domestic central bank. Moreover, since the supply and demand of loanable funds is also affected by private sector's endogenous behaviour, an increase in government debt (other things being equal) may be offset by increases in private saving, limiting its impact on the capital stock and the interest rate. They conclude that, because economic theory is not conclusive on the size of crowding-out effects, the issue must ultimately be addressed by empirical analysis. Engen and Hubbard (2004) then provide several estimates for long-term interest rates in the United States and find that the impact of public debt is statistically significant and economically relevant: about 3 basis points for one percentage point increase in the debt-to-GDP ratio. Similar results obtain if vector autoregression analysis is carried out in order to account for dynamic effects.

Laubach (2009) argues that spot interest rates are strongly influenced by the business cycle and the associated stance of monetary policy. If during recessions automatic fiscal stabilizers raise deficits, while at the same time long-term interest rates fall due to monetary easing, deficits and interest rates may be negatively correlated even if the partial effect of deficits on interest rates—controlling for all other influences—is positive. To control for business cycle and monetary policy effects on interest rates, he claims that one should focus on the relationship between long-horizon expectations of both interest rates and fiscal variables. Accordingly, his preferred specification for the United States is one in which the endogenous variable is the 5-year-ahead 10-year forward rate and the fiscal variable is the Congressional Budget Office's 5-year-ahead projection of deficit/GDP ratio or debt-to-GDP ratio. For the 30-year 1976-2006 for which these projections are available, Laubach finds that the estimated effects of government debt and deficits on interest rates are sizable: about 3 to 4 basis points for a one percentage point increase in the debt/GDP ratio and about 25 basis points per percentage point increase in the projected deficit/GDP ratio.[6]

Ardagna, Caselli and Lane (2007) focus on the international dimension by using a panel of 16 OECD countries that covers a maximum time span from 1960 to 2002. They find that, in a simple static specification, a one-percentage-point increase in the primary deficit relative to GDP increases contemporaneous long-term interest rates by about 10 basis points. They argue that their estimates tend to understate the effects of fiscal variables on interest rates, as they use current fiscal policy variables, rather than projected variables. As for debt, they find a non-linearity: only for countries with above-average levels of debt does an increase in debt affect the interest rate. They also find that world fiscal policy is important as well: an increase in total OECD government borrowing increases each country's interest rates. However, domestic fiscal policy continues to affect domestic interest rates even after controlling for worldwide debts and deficits. They argue that the latter finding can be explained either by a less-than-perfect degree of integration of advanced economies' government bond markets or by differences in perceived government default risks.

The issue of the impact of fiscal variables on long-term interest rates has been recently reexamined by Baldacci and Kumar (2010), who estimate a panel of 31 advanced and emerging market economies for the period 1980–2008. Like most previous studies, the econometric framework is based on reduced-form regressions and focuses on deficits (rather than debt). For a country experiencing an increase in the fiscal deficit of 1 percentage point of GDP, long-term interest rates could rise by 20 basis points in the baseline case. Taking into account also a combination of adverse factors (e.g., unfavorable initial fiscal conditions, weak institutions, and elevated global risk aversion), the authors argue that the effect could be as high as 50 basis points and that, according to their computations, such effect would be equivalent to a calculated debt elasticity of 5–6 basis points.

[6] He also argues that the fact that the estimated coefficients on the deficit/GDP ratio are six to seven times as large as those on the debt/GDP ratio is consistent with the view that investors perceive increases in projected deficit/GDP ratios as highly persistent, but not strictly permanent. This argument is however challenged by Engen and Hubbard (2004), who note that public debt is also serially correlated in U.S. data, so that investors should also expect increases in federal government debt to be persistent.

Over the last decade several studies have focused on the impact on long-term yields of the demand for government securities stemming from official reserve accumulation, changes in financial regulation or, more recently, large-scale asset purchases (LSAP) programs by the Federal Reserve and other central banks.[7] Beltran, Kretchmer, Marquez and Thomas (2012) find quite a sizable effect of foreign reserves.[8] They also argue that the estimated impact of the Fed's LSAP program tends to be lower, because the program was designed as a temporary stimulus program (and announced as such) and the LSAPs apparently increased the amount of uncertainty surrounding the level of future inflation, thus rising the inflation risk premium embedded in long-term interest rates.

Andritzky (2012) addresses the thorny issue of whether changes in the investor base (e.g., domestic versus non-resident investors, or leveraged versus unleveraged investors) matter. Using a new dataset on the composition of the investor base for government securities in selected G20 and euro-area countries, Andritzky estimates a reduced form regression of 10-year yields in which the explanatory variables also include the shares of government securities held by three typologies of investor: (1) non-residents; (2) private non-bank financial institutions (institutional investors); (3) public sector. He finds that an increase in the share of government securities held by institutional investors or non-residents (i.e. the ratio of the bonds held by that type of investors to the existing stock of bonds) by one percentage point is associated with a reduction in yields by about 2 or 4 basis points, respectively. In order to evaluate whether causality goes from yields to holding shares (pull effect) rather than the other way round (push effect), Andritzky carries out a panel VAR analysis and finds evidence of a pull effect, that is that lower yields attract non-resident investors. He observes, however, that the result could be driven by the fact that the sample period is characterized by falling yields and increasing non-resident holdings. Finally, Andritzky also pursues a structural approach and estimates a portfolio balance model for the US, the UK, Germany and Japan. He finds that a one percentage point increase in the share of statutory or regulatory (i.e. zero or low interest-rate sensitive) holdings of government securities causes expected annual bond returns to decline very little, by a minimum of 0.7 basis point in the UK to a maximum of 2.5 basis points in Japan.

A new perspective comes from a recent strand of the literature on the global financial crisis which emphasizes the role played by gross (rather than net) capital flows in determining financial

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[7] See, e.g., Chapter VI in BIS (2006), Greenwood and Vayanos (2010), Beltran, Kretchmer, Marquez and Thomas (2012), Andritzky (2012) and references therein.

[8] A $100 billion (about 0.7 per cent of US GDP in 2011) increase in foreign official flows into US Treasury notes and bonds would lower the 5-year yield by roughly 20 basis points.

conditions. Borio and Disyatat (2011) and Shin (2012) start from the observation that, in the global financial system, gross cross-border positions are huge and argue that a focus on current accounts and net capital flows is misleading. This is because net capital flows, by netting out the gross assets and liabilities, mask the underlying changes in gross flows and their contributions to existing stocks, including all the transactions involving only trade in financial assets, which make up the bulk of cross-border financial activity. Borio and Disyatat discuss the implications of this approach for the determination of market interest rates, mentioning as an example the downward pressure of gross capital inflows to the United States on US dollar long-term rates. Shin develops a theoretical model linking the total intermediation capacity of the banking sector and market risk premia.

Unconventional monetary policies and foreign or institutional demand for government securities certainly contribute to explain the low level of interest rates after the global financial crisis. An alternative explanation has been put forward by Krugman (2012), who argues that, because of the depressed levels of activity, business confidence in advanced economies is depressed as well and thus the private sector does not compete with the public sector for funds. Hence, budget deficits do not necessarily lead to soaring interest rates.

3. Data

The data used for the analysis are mainly obtained from the dataset published by the "OECD *Economic Outlook*". We concentrate on national macroeconomic and fiscal aggregates, for a panel of 20 countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, Norway, New Zealand, Portugal, Spain, Sweden, Switzerland, the United Kingdom and the United States. We use yearly data, from 1980. Much of the analysis concentrates on a shorter time-span – from 1995 – that provides complete information on national financial accounts for all countries in the panel. All macroeconomic aggregates are measured in terms of share of GDP.

In terms of data and methodology, the closest reference paper is Ardagna et al. (2007). Our panel differs in that it contains 20 countries, adding Finland, Norway, Portugal and Switzerland. It also differs in terms of estimation samples: we analyze the periods 1980-2011 and 1995-2011, whereas they concentrate on the periods 1960-2002 and 1975-2002. Our choice is motivated by two facts: firstly, few aggregate variables are available for all countries prior to 1980, while there is an almost perfectly balanced panel after that year. Secondly, 1980 is a year of structural break for public finance aggregate relationships, both in terms of monetary policy –Volcker's designation at the Federal Reserve in 1979 – and in terms of fiscal policy – elections of Thatcher (1979) in the UK and Reagan (1981) in the USA.

Variables are listed in Table 1.

Table 1: Variables

Variable Name	Description
YIELD_10Y	10-year government bond nominal yield
YIELD_3M	3-month treasury bill nominal yield
INFLATION	Current inflation rate, YoY
INFLATION_10Y	Modelled forecasted 10-year inflation, YoY
REAL_10Y	10-year government bond real yield
REAL_3M	3-month treasury bill real yield
DEBT	Gross government debt (% of GDP)
GOV_ASSET	Gross government assets (% of GDP)
WEALTH_HH	Net wealth of households (% of GDP)
WEALTH_NF	Net wealth of non-financial corporations (% of GDP)
ASSET_FF	Gross assets of the financial sector (% of GDP)
NF_DEBT	External debt (% of GDP)
RESERVES	Share of debt held as official reserves by foreign central banks (% of GDP)
AVG_LIFE	Average life to maturity of outstanding marketable debt (years)
RATING	Maximum rating grade

YIELD_10YR and YIELD_3M are the nominal yields of the 10-year benchmark government bond and the 3-month money market interest rate, respectively, both computed on yearly basis. The inflation rate enters the regressions in two different manners: either as "expected" 10-year rate (INFLATION_10Y), on yearly basis, or as yearly "spot" rate (INFLATION). In the former case, it is subtracted from the nominal yield to compute the real 10-year yield (REAL_10Y); in the latter, it is added to the r.h.s. of the regression, either directly or subtracted from the nominal yield to compute the 3-month real rate. Expected inflation rates on the 10-year horizon are available from "*Consensus Economic Forecasts*", for 15 of the 20 countries in the panel, from 1989. We have imputed those of the other countries based on a model that predicts future 10-year inflation rates based on short-term forecasts of the inflation rate and recent past rates. Details are reported in the Appendix.

Data on the financial accounts positions of the main sectors of the economies in the panel are drawn from the National Financial Accounts as reported by the "*OECD Economic Outlook*". In particular, DEBT and GOV_ASSET are the gross positions of the public sector at large; WEALTH_HH, WEALTH_NF, are the net financial positions of the Household and Non-Financial Corporations sectors; ASSET_FF are the gross assets of the Financial sector whereas their gross liabilities are excluded from the analysis. NF_DEBT is the net position of the Foreign sector, as reconstructed by Lane and Milesi-Ferretti (2007).

RESERVES is the amount of a country's currency held by foreign central banks as reserves. As this is normally all invested in government bonds, we include it in our regressions as a proxy of "high powered" net-foreign-debt. Data are drawn from IMF COFER and more details on the methodology are in the appendix. AVG_LIFE is the average life to maturity of the outstanding marketable debt, measured in years, as collected by the OECD. RATING is a categorical variable that summarizes the rating of the three major rating agencies, according to the methodology outlined in the appendix.

We test for unit root in panel data using the diagnostics of Im, Pesaran and Shin (2003) and a Fisher-type test as in Choi (2001), based on augmented Dickey-Fuller independent tests on each country, combined together. For most of the series with longer time-span (1980-2011), we are able to reject the null hypothesis of the presence of unit roots. Only DEBT appears to be I(1), according to both tests. However, in accordance with Engen and Hubbard (2005), we prefer to include this variable into the regressions in levels rather than in first differences. Considering the other series, included in our regressions as controls, we find some evidence for non-stationarity. In fact, both the IPS and the ADF fail to reject the null hypothesis in some cases. However, given the very short time span available for these series (only 16 data points), the power of both tests is extremely low and might invalidate our conclusions about the stationarity properties of the series. Thus we prefer to use all of them in levels.

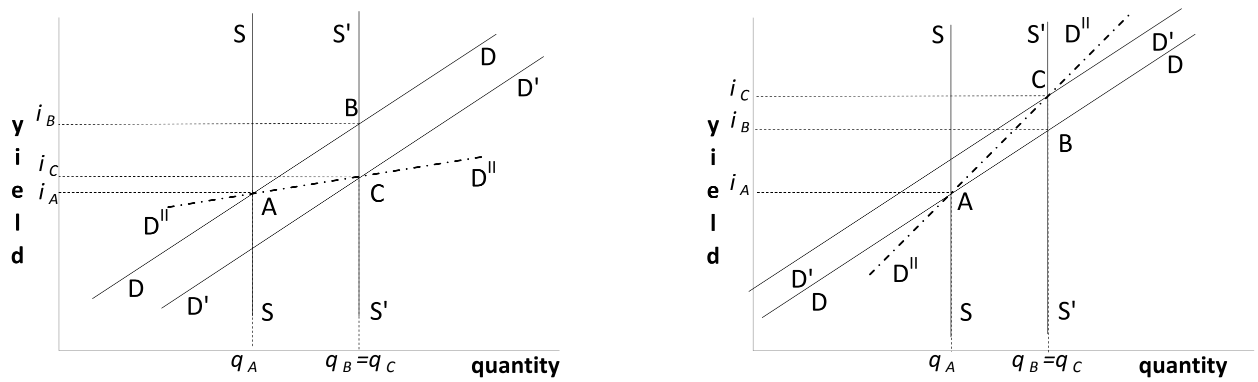
4. Identification Strategy

On bond markets yields clear demand and supply. If there is an increase in the supply of bonds, its impact on the market yield depends on the slope of the demand curve – i.e. on the so-called interest rate elasticity of demand. For a given increase in the supply of bonds, the higher the elasticity of demand the lower the increase in yield which is necessary to clear the market. The objective of the paper is to estimate the interest rate elasticity of the demand for sovereign debt for advanced countries. This is a key parameter, as it allows to quantify the potential impact, on long-term yields, of a change in the stock of sovereign debt.

Figure 1

Interest rate effect of a positive shock to the supply of bonds
in the presence of shifts in the demand schedule (1)

(*a*) Increase in the demand for bonds for given yield (*b*) Decrease in the demand for bonds for given yield



(1) Just for illustrative purposes, the assumption is made that the supply curve is perfectly inelastic.

The reduced-form equation for the market yield i_t at time t is

$$i_t = a_0 + a_1 * q_t + e_t, \tag{1}$$

where q_t is the outstanding amount of the bond at time t and e_t is a residual. The slope parameter a_1 provides an estimate of the interest rate elasticity of demand. The main problem in estimating equation (1) is illustrated in Figure 1.a (under the hypothesis of a perfectly inelastic supply of bonds). In case of a positive shock to the supply of bonds, represented by a shift of the supply schedule from the SS curve to the S'S' curve, the market yield should increase from i_A to i_B. However, if at the same time there is an increase in the "autonomous" (unrelated to yield) demand, due, for example, to larger capital inflows from abroad, the demand schedule shifts from the DD curve to the D'D' curve and the market yield rises less, from i_A to i_C. In that case, the slope estimate provided by the reduced-form equation does not relate to the D'D' curve – which is the true or "structural" demand curve –, but relates to the D''D'' curve. The reduced-form equation thus overestimates demand elasticity. In case of a decrease in autonomous demand, the reduced-form equation underestimates demand elasticity (Figure 1.b).

In order to control for changes in the demand for bonds that are unrelated to interest rates, we allow for exogenous shifts in the demand for sovereign debt. Our baseline equation is as follows:

$$yield\_10y_{it} = a_0 + a_1 * debt_{it} + a_2 * x_{it} + e_{it}, \tag{2}$$

where $yield\_10y_{it}$ is the yield on 10-year government bonds for country i in period t, $debt_{it}$ is the debt-to-GDP ratio, x_{it} is a set of controls and e_{it} is a disturbance term with standard assumptions.

Our parameter of interest is a_1. Depending on the specification, *yield_10y*$_{it}$ may be measured either in real or nominal terms, with proper adjustments to the explanatory variables to make the two sides of the equation consistent.

The key idea that we use to obtain identification is that of exploiting the national financial accounts (or flow of funds accounts) identity. As shifters of demand, we use the balances – i.e. the difference between the value of assets and liabilities – of the financial accounts of the main institutional sectors of the economy. More specifically, we use the financial accounts identity to saturate the regression with the balances of all but one of the sectors (so as to avoid collinearity). We thus control for the net financial balances of households, non-financial firms and the foreign sector, leaving aside the net balances of financial intermediaries (see also the appendix). In order to assess whether gross (rather than net) positions also have an impact, we control for the world reserves invested in the currency of the country and for the gross assets of the financial sector. The reason to include the latter variable also lies in the implicit burden that sovereigns might shoulder in case of financial crises. Gross assets of the financial sector should give a rough measure of this burden. Finally, we include general government's gross assets among the regressors. All financial balances are measured as a fraction of GDP.

In addition to allowing for changes in autonomous demand, we limit the range of possible slopes of the yield curve, controlling for the short-term rate. We also allow for other shifters of demand that capture the degree of substitutability of sovereign debt with other assets, such as the average life to maturity of the outstanding amount of government bonds and the ratings of sovereign issuers.

The above identification approach rests upon three key assumptions:

(1) *Bond supply is exogenous*. This is a strong assumption, and later on we show how to relax it. Bond supply can be regarded as inelastic to interest rates only in the short run. In the long-run, the supply of bonds is to some degree interest-rate sensitive (the higher the interest rate the lower the supply of bonds).

(2) *The financial positions (gross assets and liabilities and/or their balance) of the institutional sectors of the economy are exogenous*. This is also not necessarily true. The portfolio choices of households and foreign investors are likely to be affected, to some degree, by the level of sovereign yields. This observation is consistent with Andritzky (2012)'s finding that declines in yields would be followed by (rather than being a consequence of) inflows of foreign investments in government bonds. Similarly, the financing decisions of non-financial corporations are affected by the level of sovereign yields. However, for the purpose of this

analysis, the failure of this assumption is a second-order problem, as our parameter of interest is the demand elasticity (coefficient a_1 in equation (2)), not the coefficients associated with the demand shifters (vector a_2 in equation (2)).

(3) *Institutional sectors' asset allocation is assumed to be constant over time and across countries*, as reflected by the fixed portfolio coefficients. This is also a simplifying assumption, because one may argue that, for example, the share of households' financial wealth held in government bonds may change. However, we have too few observations to allow for time- or country-varying coefficients.

Our baseline estimation method is least squares with fixed effects and robust standard errors. The reduced-form equation (1) can be estimated for 20 countries over the period 1980 – 2011, totalling 562 observations in the sample (on average almost 28 observation per country). For equation (2), that makes use of financial accounts variables, we have data for 18 countries over the period 1995-2011, totalling 292 observations (about 16 observations per each country) [9].

Once our best model specifications are fixed, we are able to relax assumption (1) addressing the potential endogeneity of the bond supply. The supply of bonds is well likely to be influenced by the level of the interest rate: the lower this level, the higher the supply of bonds. In practice, the endogeneity problem might be not extremely relevant, because the share of bond supply that is actually interest-rate sensitive is limited. Every year, the bond supply – which can be proxied by the debt-to-GDP ratio, $debt_{it}$ – is in fact constrained by the realized debt-to-GDP ratio one year before ($debt_{it-1}$). The fiscal room to determine the supply of bonds at time t is further constrained by the amount of interest payments on the realized government debt. Finally, the automatic stabilizers that react to the cycle would further reduce the endogeneity of the supply of debt.

In order to correctly address the endogeneity problem, we then isolate the share of the debt supply that is actually discretional and use the strictly exogenous debt supply as instrumental variable for the actual debt-to-GDP ratio in a two-stage least-squares fixed-effect estimation. The exogenous component of the debt supply - $debt\_ex_{it}$ – is calculated as follows:

$$debt\_ex_{it} = debti_{t-1} + int\_pay_{it} + aut\_stab_{it}$$

where $debti_{t-1}$ is the realized debt-to-GDP ratio at time *t-1*, int_pay_{it} is the ratio of interest payments due at time *t* to GDP and aut_stab_{it} is the share of the primary balance attributable to non-discretional automatic stabilizers. The latter variable, being a typical cyclical component, is

[9] We drop New Zealand and Switzerland from the analysis, since financial accounts data for these two countries are missing.

calculated as the difference between the realized primary balance of each country and its cyclically-adjusted value, as calculated by the OECD.

5. Estimation Results

Our main results are summarized in Table 2, in which columns differ from one another either in terms of the sample period or the set of control variables.

Right-hand side variables always include a 3-month (real or nominal) interest rate, to control for parallel shifts of the yield curve. The underlying assumption that this rate is uncorrelated with the error term is based on the fact that this is a policy rate set by the central bank. The other explanatory variables only capture movements of the slope of the yield curve.

In column 1 of Table 2 we present a “Plain vanilla” fixed effects regression, with time dummies that capture any common time trend. In the literature, this is the workhorse model for most studies like this one. As an example, Ardagna, Caselli and Lane (2007) always have time dummies as controls. The fit is very good, because the common trend captures most of the variation. However, for this specification the demand elasticity is only 1 basis point per percentage point of GDP. Column 2 presents the same specification for a shorter sample: 1995-2010. This is the sample on which all the other regressions are estimated, hence we use this as a benchmark. Also in this case the demand elasticity is 1 basis point.

The main results of our paper are presented in columns 3 and 4 of Table 1, where the common trend is replaced with our economic restriction, namely that shifts in the demand schedule are driven by changes in the balances of the financial accounts of the main institutional sectors of the economy (as explained in the previous section). This amounts to giving each country its own “time trend”, driven by its fundamentals.

Table 2

Ten-year interest rates of advanced economies: Estimates of the elasticity of demand (1)

	[1]		[2]		[3]		[4]		[5]		[6]		[7]	
Public debt (% GDP)	0.013	[0.05]	0.018	[0.03]	0.0480	[0.00]	0.0309	[0.00]	0.0130	[0.02]	0.0481	[0.00]	0.0298	[0.01]
Inflation (%)	0.638	[0.00]	0.416	[0.00]	0.4502	[0.00]	0.4060	[0.00]	0.4723	[0.00]				
3-month real rate (%)	0.502	[0.00]	0.431	[0.00]	0.4977	[0.00]	0.4999	[0.00]	0.4577	[0.00]				
3-month nominal rate (%)											0.4840	[0.00]	0.4691	[0.00]
General gov't assets (% GDP)					-0.0286	[0.00]	-0.0242	[0.00]	-0.0032	[0.46]	-0.0282	[0.00]	-0.0228	[0.00]
Househ.ds' net fin. Wealth (% GDP)					-0.0287	[0.00]	-0.0240	[0.00]	-0.0065	[0.14]	-0.0279	[0.00]	-0.0223	[0.00]
Non-fin. corp.ns' net debt (% GDP)					-0.0129	[0.03]	-0.0090	[0.06]	-0.0020	[0.56]	-0.0123	[0.06]	-0.0080	[0.11]
Net foreign debt (% GDP)					-0.0067	[0.16]	-0.0064	[0.12]	0.0025	[0.53]	-0.0057	[0.28]	-0.0042	[0.36]
Foreign off. reserves (% GDP)					-0.1480	[0.00]	-0.1307	[0.00]	-0.0158	[0.51]	-0.1514	[0.00]	-0.1363	[0.00]
Fin. corp.ns' assets (% GDP)					-0.0002	[0.63]	-0.0006	[0.21]	0.0006	[0.05]	-0.0002	[0.69]	-0.0005	[0.21]
Average life to maturity (years)					-0.2165	[0.03]	-0.1643	[0.04]	-0.0472	[0.23]	-0.2267	[0.02]	-0.1837	[0.02]
AA + (dummy)							0.3361	[0.02]	0.1276	[0.51]			0.4210	[0.02]
AA (dummy)							0.5205	[0.02]	0.2957	[0.37]			0.6476	[0.01]
AA - (dummy)							1.6263	[0.00]	0.9858	[0.12]			1.7559	[0.00]
A + (dummy)							1.9507	[0.00]	2.1857	[0.00]			2.0284	[0.00]
BBB + (dummy)							5.0045	[0.00]	4.8252	[0.00]			4.8819	[0.00]
BB + (dummy)							6.4117	[0.00]	6.5398	[0.00]			6.3771	[0.00]
Constant	3.636	[0.00]	4.094	[0.00]	5.4568	[0.00]	5.9523	[0.00]	4.9441	[0.00]	5.4026	[0.00]	5.6213	[0.00]
Year dummies	Yes		Yes		No		No		Yes		No		No	
R-square	0.918		0.606		0.368		0.516		0.850		0.371		0.522	
Sample period	1980-2011		1995-2011		1995-2011		1995-2011		1995-2011		1995-2011		1995-2011	
Number of countries	20		20		18		18		18		18		18	
Number of observations	562		337		292		292		292		292		292	

<u>Legend of model specification</u>: [1] Common time trend; [2] Common time trend; [3] Heterogeneous time trend and economic restrictions; [4] Heterogeneous time trend, economic restrictions and ratings; [5] Common time trend and economic restrictions; [6] Heterogeneous time trend and economic restrictions; [7] Heterogeneous time trend, economic restrictions and ratings.

(1) Panel estimates with fixed effects, run on yearly data. For each specification, the table shows coefficient estimates and, in square bracket, the related *p*-values.

Starting from the specification shown in column (3), once the common trend is replaced with our economic restriction, the interest rate elasticity of the demand for government bonds becomes much smaller: an increase of one percentage point in the public debt-to-GDP ratio leads to an increase in the 10-year real rate on the order of 4 basis points. Moreover, the R^2 of the regression is fairly good and almost all of the other coefficients are significant and have the correct sign. An increase of one percentage point of GDP in general government's gross assets or households' net financial wealth lowers the 10-year real interest rate by about 3 basis points. A reduction of one percentage point of GDP in the net debt of non-financial firms or an increase of the same magnitude in the net foreign debt position[10] lower the 10-year real rate by about 1 basis point. A much stronger effect is found for foreign official reserves (a component of the net foreign debt position): one percentage point increase in the ratio of foreign reserves to GDP leads to a reduction of the 10-year real rate by more than 12 basis points. An increase in financial corporations' gross assets is also associated with a reduction of the 10-year real rate, but in this specification the effect is not statistically significant. Finally, a one-year increase in the average life to maturity of the outstanding amount of government bonds implies a decline of almost 19 basis points in the 10-year real rate.

The degree of substitutability between government bonds and alternative asset classes (e.g., corporate bonds and listed shares) is affected by changes in the creditworthiness of sovereign borrowers – i.e. sovereign credit risk. In the specification shown in column (3), the only variable that accounts for investors' sovereign debt sustainability concerns, in addition to the level of public debt as such, is the average life to maturity of the existing stock of bonds. The specification presented in column (4) of Table 1 tries to better capture investors' perception of the soundness of sovereign borrowers. It does so by including sovereign ratings dummies among the control variables, under the working hypothesis that the grades assigned by rating agencies to government bonds can be a rough indicator of financial markets participants' perceptions of sovereign credit risk.

In the specification shown in column (4), rating dummies turn out to have a strong effect on the 10-year real rate. Their coefficients are significant and proportional to the degree of riskiness associated with the rating grade. A comparison of column (4) with column (3) indicates that the inclusion of rating dummies tends to make demand elasticity higher than in the specification without rating dummies: one additional percentage point of public debt-to-GDP ratio increases the 10-year real rate by about 3 (instead of 4) basis points. The other coefficients are all remarkably stable, although most of them are a bit smaller in magnitude. The net foreign debt position is not

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[10] We are grateful to Philip Lane and Gian Maria Milesi-Ferretti for providing these data. The original reference is Lane and Milesi-Ferretti (2007).

anymore significant, while financial corporations' gross assets become significant at a 10 per cent confidence level.

In column (5), the year dummies are put back into the "financial account" regression of column (4). Not surprisingly, most of the control variables are not anymore significant and have a lower effect.

In column (6) and column (7) the short-term real rate and the inflation rate are replaced with the nominal short rate. Column (6) is without rating dummies. The demand elasticity is about 4 basis points, but most of the other coefficients are not significant. Column (7) also includes rating dummies. The demand elasticity is about 3 basis points, as in column (4), while net foreign debt and the rating dummies are not anymore significant.

All in all, a number of results stand out from our analysis. First of all, in advanced economies the long-run elasticity of demand for sovereign debt is on average quite high. Controlling for sovereign ratings, the estimated impact of bond supply is around 3 basis points per percentage point of public debt-to-GDP ratio. In the specifications without sovereign ratings, the demand elasticity is slightly lower, as the slope estimate is 4 basis points per percentage point of public debt-to-GDP ratio.

Despite being very strong, our restriction on demand shifters fits the data very well and we are able to quantify the effects of several factors in a consistent way. The largest coefficient is associated with official foreign exchange reserves. This agrees with the view that the sizable stock of reserves accumulated by emerging market countries since the late nineties (due to currency intervention, current account surpluses and other factors) has exerted a strong downward pressure on the yields of advanced economies. Other factors having an impact on the demand for sovereign debt are households' net financial wealth, general government's holdings of financial assets, non-financial corporations' net debt and the net foreign debt position (which takes into account all capital flows, including official foreign exchange reserves). The coefficient associated with banks' gross assets is scantly significant. This is not necessarily inconsistent with the hypothesis that gross (rather than net) capital flows matter, for two reasons. First, the acceleration in cross-border bank assets started around the middle of the 2000s[11] and thus weighs on a rather small fraction of our sample period. Second, gross (as opposed to net) flows effects are also captured by official reserves.

As for public debt sustainability indicators, which allow us to control for changes in financial markets' perception of sovereign credit risk, we find that the average maturity of public

[11] See, e.g., Figure 5 in Shin (2012).

debt does matter. The results for ratings are instead less clear-cut, as the dummies are significant only in equation (4).

A key aspect of demand elasticity estimates is the potential endogeneity of the supply of debt. As explained in Section 4, this is addressed through instrumental variables estimates. The results are shown in Table 3. Equations (3) and (4) of Table 2 are replicated with different instruments, namely (a) lagged debt, (b) exogenous debt (as specified in Section 4) and (c) exogenous debt together with its lagged value.

The coefficients for the debt variable range from 4 to 5 basis points for equation (3) and from 2 to 3 basis points for equation (4). These results are broadly in line with the baseline estimations – respectively, 5 and 3 b.p. for equations (3) and (4). The slight reduction in the magnitude of the coefficient in both specifications is consistent with the theory, which predicts a negative relation between the level of the interest rate and the supply of bonds. Moreover, the small change in the coefficients from the baseline strict-exogeneity estimations points to a limited sensitivity of the debt supply to interest rates.

6. Robustness

Our main robustness concern regards the reliability of financial accounts in accurately capturing demand shifts. As explained in Section 4, we saturated the equations with the net balances of the financial accounts identity of all institutional sector but one (that of financial intermediaries) and also included gross assets of financial intermediaries.

To test the robustness of our findings to this specification we run several replications of equations (3) and (4) including the net financial balance of financial intermediaries and in turn the gross assets of any other institutional sector. The results – shown in the Appendix (Table A.4) – are robust to changes in the chosen financial accounts specification. The debt coefficient remains significant and changes its value only marginally. The specification shown in Table 2 appears to be the most meaningful economically.

We conducted several other robustness checks (not shown). We thoroughly looked for non-linearities (trying many different parameterizations), but we didn’t find any type of non-linear effect. We also tried many different types of forward-looking variables (e.g., expected inflation and expected fiscal deficits), but we didn’t get any significant gain in terms of accuracy of the estimates.

Table 3

Estimations with instrumental variables (1)

	[3]		[4]		[3]		[4]		[3]		[4]		[3]		[4]	
Public debt (% GDP)	0.0480	[0.00]	0.0309	[0.00]	0.0443	[0.05]	0.0238	[0.00]	0.0442	[0.00]	0.0248	[0.00]	0.0478	[0.00]	0.0279	[0.00]
Inflation (%)	0.4502	[0.00]	0.4060	[0.00]	0.4512	[0.00]	0.4080	[0.00]	0.4512	[0.00]	0.4077	[0.00]	0.4453	[0.00]	0.4054	[0.00]
3-month real rate (%)	0.4977	[0.00]	0.4999	[0.00]	0.4987	[0.00]	0.4955	[0.00]	0.4987	[0.00]	0.4961	[0.00]	0.4868	[0.00]	0.4923	[0.00]
3-month nominal rate (%)																
General gov't assets (% GDP)	-0.0286	[0.00]	-0.0242	[0.00]	-0.0260	[0.00]	-0.0199	[0.00]	-0.0259	[0.00]	-0.0205	[0.00]	-0.0286	[0.00]	-0.0225	[0.00]
Househ.ds' net fin. Wealth (% GDP)	-0.0287	[0.00]	-0.0240	[0.00]	-0.0264	[0.00]	-0.0205	[0.00]	-0.0264	[0.00]	-0.0209	[0.00]	-0.0282	[0.00]	-0.0224	[0.00]
Non-fin. corp.ns' net debt (% GDP)	-0.0129	[0.03]	-0.0090	[0.06]	-0.0106	[0.02]	-0.0058	[0.14]	-0.0105	[0.02]	-0.0063	[0.11]	-0.0126	[0.01]	-0.0078	[0.05]
Net foreign debt (% GDP)	-0.0067	[0.16]	-0.0064	[0.12]	-0.0045	[0.34]	-0.0031	[0.46]	0.0044	[0.36]	-0.0035	[0.40]	-0.0058	[0.22]	-0.0047	[0.26]
Foreign off. reserves (% GDP)	-0.1480	[0.00]	-0.1307	[0.00]	-0.1415	[0.00]	-0.1201	[0.00]	-0.1414	[0.00]	-0.1216	[0.00]	-0.1488	[0.00]	-0.1261	[0.00]
Fin. corp.ns' assets (% GDP)	-0.0002	[0.63]	-0.0006	[0.21]	-0.0002	[0.67]	-0.0006	[0.16]	0.0002	[0.67]	-0.0006	[0.16]	-0.0002	[0.66]	-0.0006	[0.16]
Average life to maturity (years)	-0.2165	[0.03]	-0.1643	[0.04]	-0.2146	[0.03]	-0.1600	[0.01]	-0.2146	[0.00]	-0.1606	[0.01]	-0.2149	[0.00]	-0.1658	[0.01]
AA + (dummy)			0.3361	[0.02]			0.4316	[0.01]			0.4181	[0.02]			0.3826	[0.03]
AA (dummy)			0.5205	[0.02]			0.6950	[0.01]			0.6705	[0.01]			0.6357	[0.02]
AA - (dummy)			1.6263	[0.00]			1.7841	[0.00]			1.7619	[0.00]			1.4175	[0.01]
A + (dummy)			1.9507	[0.00]			2.1499	[0.00]			2.1219	[0.00]			2.0575	[0.00]
BBB + (dummy)			5.0045	[0.00]			5.2495	[0.00]			5.2150	[0.00]			5.0799	[0.00]
BB + (dummy)			6.4117	[0.00]			6.7654	[0.00]			6.7157	[0.00]			6.5956	[0.00]
Constant	5.4568	[0.00]	5.9523	[0.00]	5.4874	[0.00]	5.9970	[0.00]	5.4879	[0.00]	5.9908	[0.00]	5.4481	[0.00]	5.9705	[0.00]
Year dummies	No		No		No		No		Yes		No		No		No	
R-square	0.368		0.516		0.389		0.553		0.389		0.548		0.365		0.525	
Sample period	1995-2011		1995-2011		1995-2011		1995-2011		1995-2011		1995-2011		1995-2011		1995-2011	
Number of countries	18		18		18		18		18		18		18		18	
Number of observations	292		292		292		292		291		291		291		291	
Instruments					Debt (t-1)		Debt (t-1)		Debt_ex		Debt_ex		Debt_ex; Debt_ex (t-1)		Debt_ex ; Debt_ex (t-1)	

<u>Legend of model specification</u>: [3] Heterogeneous time trend and economic restrictions; [4] Heterogeneous time trend, economic restrictions and ratings;
(1) Panel estimates with instrumental variables, run on yearly data. For each specification, the table shows coefficient estimates and, in square bracket, the related *p*-values.

7. Conclusions

We obtain an estimate of the long-term elasticity of the demand for government securities in advanced economies. We use panel data on 18 countries covering the sixteen years from 1995 to 2011. The sample period includes not only the low interest rate phase in the mid-2000s but also the further downward trend in market yields observed in the three-year 2009-11 after the most acute phase of the global financial crisis. We find that, in the long run, each percentage point increase in the ratio of public debt to GDP raises 10-year rates by about 3 basis points. This is a sizeable effect, considering that, in the three-year 2009-11, the median value of the average annual change in the debt-to-GDP ratio was equal to almost 6 percentage points in the 18 countries considered and 7 percentage points in the G7 countries.

Previous estimates of the demand elasticity of government debt in the United States are in line with our results. Most of those estimates, however, are based on sample periods that do not extend beyond the mid-2000s. More importantly, they are based on reduced form regressions, while our empirical framework is able to identify the long-term elasticity of the demand for government bonds, because it controls for short-term shifts in demand and also addresses the potential endogeneity of debt supply to interest rates.

The use of financial accounts data also allows us to take into account in a consistent way some factors that have been said to affect market yields since the 2000s, namely the accumulation of foreign official reserves by emerging market and oil exporting countries and changes in financial regulation and accounting. Our results give insights also on the interest rate effects of unconventional monetary policies, although the latter matter only in the last two years of the sample and for a limited number of countries. We also find evidence that gross (rather than net) positions do have an impact on interest rates. These results are robust to the use of different combinations of financial accounts data as right-hand side variables. Finally, we control for changes in the perceived riskiness of government bonds by using debt sustainability indicators (average life to maturity and sovereign ratings).

In conclusion, public debt does affect long-term interest rates and its potential drag on growth through higher interest rates should not be overlooked. It must be underlined, however, that the interest rate burden and other costs of public debt have always to be carefully weighed against the overall short- and long-run benefits of government intervention, especially in depressed economies.

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Appendix

In this Appendix we provide further evidence of the methodologies we use to assemble the data for some variables. All the macroeconomic and fiscal variables did not need any special care, except presenting the main fiscal variables as a share of GDP, while the long-term expected inflation, the share of debt held in foreign official reserves and the rating variables needed some additional manipulation. Moreover, some further explanation on financial account variables might be useful.

Macroeconomic Variables

The macroeconomic and fiscal variables data were collected from the n. 88 OECD Economic Outlook. Here is presented the list of the main variables selected, with their OECD code and definition.

Table A.1: Macroeconomic and Fiscal Series

Code	Description
CBGDPR	Current account balance, as a percentage of GDP
CPIH_YTYPCT	Consumer price index, harmonised, year-on-year growth
EXCHER	Real effective exchange rate, constant trade weights (New)
GAP	Output gap of the total economy
GDPV	Gross domestic product, volume, market prices
GFAR	General government gross financial assets, as a percentage of GDP
GGFL	General government gross financial liabilities, value
IRL	Long-term interest rate on government bonds
IRS	Short-term interest rate
NLG	Government net lending, value

Expected long-term inflation

The most comprehensive and reliable data source of forecasts of macroeconomic indicators is provided by Consensus Economics, from 1989 onwards. Regarding short-term forecasts, CE provides monthly updates of forecasts for the year under review and one year ahead, for all the countries of the sample[12].

However data for 10-year ahead forecasted inflation are provided only for a subset of 12 countries out of the 20 we are interested in (Canada, France, Germany, Italy, Japan, the UK and the US from 1990; the Netherlands, Spain and Sweden from 1995 and Norway and Switzerland from 2000)[13]. We specify a regression model able to explain the 10-year forecasted inflation for these 12 countries. Then, through the regression coefficients, we build a modelled 10-year forecasted inflation variable for all the countries of the sample. We estimate a fixed-effect panel regression model, following the same econometric restrictions about the robustness of the errors as in the main regression in Section 5. The model specifications we estimate are shown in Table A.2. Trying to hold the model as simple as possible, we restricted the set of regressors to realized and short-term forecasted values of inflation, while the regressand is obviously the 10-year forecasted inflation.

[12] These forecasts structure is followed from February onwards, as the January release reports data for the year before and the one under review. Collecting short-term data we chose the forecasts released every February.

[13] Actually, the average 10-year inflation forecast is not readily available. Consensus Economics provides semi-annual forecasts of inflation 1-, 2-, 3-, 4-, 5-, and the average 6-to-10-year ahead. Thus, we have averaged the data through every maturity of the reference horizon and between the two issues of every year.

Our estimation results are presented in Table A.2. According with the AIC and BIC criteria, we can consider a simple model with only the 1-year forecasted inflation as regressor as our preferred specification, as in eq. (3).

Table A.2: 10-Year Ahead Inflation Forecasts Estimation

	Dep. Var. INFLATION_10Y Time Sample (1990-2010)			
	(1) OLS	(2) OLS	(3) OLS	(4) OLS
INFL_CUR_FCAST	0.480	-0.036		
	(9.96)**	(-1.07)		
INFL_1Y_FW		0.787	0.746	0.775
		(11.69)**	(20.40)**	(17.40)**
INFLATION				-0.027
				(-1.27)
CONSTANT	1.241	0.609	0.624	0.616
	(12.60)**	(6.91)**	(7.97)**	(7.58)**
Adjusted R^2	0.73	0.91	0.91	0.91
F-Test	99.13	286.70	416.35	202.95
AIC	183.10	-89.71	-89.72	-90.49
BIC	186.48	-82.95	-86.34	-83.74
N. of obs.	217	217	217	217

As a robustness check, we plugged realized and expected values of GDP growth in our regressions. However GDP growth does not appear to play a statistically significant role in shaping long-term inflation expectations. Thus we maintain eq. (3) as our preferred model specification. From its coefficients, we can build the modelled 10-year ahead inflation forecasts.

Financial Accounts

Financial balance sheets data are collected from the OECD Database "Financial Accounts", which belongs to the System of National Accounts (SNA 93). According to the OECD definition, the financial balance sheets "record the stocks of assets and liabilities held by the institutional sectors, and give a picture of their net worth, at the end of the accounting period"[14]. Even though data are available from 1970 onwards, we find adequately populated series for the 20 countries of our sample from 1995 onwards[15]. However New Zealand did not provide this set of data with continuity, thus restricting our sample to 19 countries when using financial accounts variables.

The institutional sectors, the economy is broke-down in, are: non-financial corporations (S11), financial corporations (S12), general government (S13), households and non-profit institutions serving households (S14-15). These sectors sum up to make the total economy sector (S1). Finally, another sector is added, accounting for the rest of the world sector (S2), which reflects asset and liabilities of non-residents. Similarly to the balance of payment identity, any net worth value of the total economy sector is balanced by a net worth value of opposite and equal size for the rest of the world, such that the following identity is always true:

$$NetWorth^{S11} + NetWorth^{S12} + NetWorth^{S13} + NetWorth^{S14\text{-}15} + NetWorth^{S2} = 0$$

[14] Further details can be found at: http://www.oecd.org/LongAbstract/0,2546,en_2649_34245_37366237_1_1_1_1,00.html. For a recent analysis based on these data, see Bruno, De Bonis and Silvestrini (2012).

[15] 1995 is the year of the introduction of the ESA95 standard, which makes the national accounts data comparable across countries. Only few countries provide data prior to this year.

We chose to use consolidated financial accounts. The reliability of consolidated accounts more accurately represents the financial position of the various sectors in the economy.

Debt held in foreign currency official reserves

Since data accounting for the share of government debt held by foreign central banks, through the allocation of their official foreign exchange reserves, are not available, a proxy variable mimicking this phenomenon is necessary. We rely on the IMF COFER database, which provides the currency composition of official foreign exchange reserves being held globally[16]. Moreover, these reserves data do not include holdings of a currency by its issuing country. Thus this dataset can serve as the best proxy for the demand of sovereign debt from foreign central banks.

Only the following currencies are identified in the database: US Dollar, Euro, Pound Sterling, Japanese Yen and Swiss Franc; and other currencies[17]. Thus we are able to attribute the share of foreign exchange reserves invested in sovereign debt only for four countries: the US, the UK, Japan and Switzerland. For the remainder of the sample – euro-area countries and others – we need some manipulation of the data. For the euro-area countries in our sample we choose to assign to each country a share of the reserves in euros equal to its share of the euro-area GDP[18]. It is hard to imagine an objective criterion able to account for the flight-to-haven phenomenon that affected the euro-denominated debt market in the aftermath of the global financial crisis of 2007-2010. Thus, our data might be slightly underestimated for core euro-area countries (e.g. Germany, France, etc.) and slightly overestimated for peripheral countries (e.g. Ireland, Portugal, etc.). For the remainder of the sample we follow a similar approach. The "other currencies" series is broke down according to the share of each country's GDP, relative to world GDP.

Rating Grades

Rating grades are collected from the three main rating agencies (Standard & Poor's, Moody's and Fitch Ratings), made comparable through the commutation criteria shown in Table A.3, and associated with the corresponding number, as reported in the Rank column, such that every country have three numbers, corresponding to three ratings, for every year. The lowest number, corresponding to the highest rating grade, is selected for every country and year. In the 1995-2010 sample the highest rating grade is always no less than 4 (that is AA-/Aa3/AA-).

[16] For further details, see: http://www.imf.org/external/np/sta/cofer/eng/index.htm

[17] Before the introduction of the euro in 1999, the COFER database also identified: Deutsche Mark, French Franc, Netherlands Guilder and the European Currency Unit (ECU).

[18] For the years preceding the introduction of the euro (1995-1998), when three national currencies and the ECU were identified, we choose a similar criterion. Deutsche Mark, French Franc and Netherlands Guilder are attributed to Germany, France and the Netherlands. In addition, a share of ECU holdings is assigned to each EU country according to its share of EU GDP.

Table A.3: Rating grades conversion table

S&P	Moody's	Fitch	Rank
AAA	Aaa	AAA	1
AA+	Aa1	AA+	2
AA	Aa2	AA	3
AA-	Aa3	AA-	4
A+	A1	A+	5
A	A2	A	6
A-	A3	A-	7
BBB+	Baa1	BBB+	8
BBB	Baa2	BBB	9
BBB-	Baa3	BBB-	10
BB+	Ba1	BB+	11
BB	Ba2	BB	12
BB-	Ba3	BB-	13
B+	B1	B+	14
B	B2	B	15
B-	B3	B-	16
CCC+	Caa1	CCC+	17
CCC	Caa2	CCC	18
CCC-	Caa3	CCC-	19
CC	Ca	CC	20
C	C	C	21

Table A.4

Robustness: Estimates of the elasticity of demand through different breakdowns of the financial accounts identity(1)

Gross assets included as regressors (% GDP)												
	Househ.ds' assets [3]		Non-fin. corp.ns' assets [3]		Foreign assets [3]		Househ.ds' assets [4]		Non-fin. corp.ns' assets [4]		Foreign assets [4]	
Public debt (% GDP)	0.0330	[0.01]	0.0376	[0.01]	0.0438	[0.00]	0.0197	[0.03]	0.0218	[0.03]	0.0266	[0.00]
Inflation (%)	0.4233	[0.00]	0.4617	[0.00]	0.4674	[0.00]	0.3698	[0.00]	0.4008	[0.00]	0.4162	[0.00]
3-month real rate (%)	0.4655	[0.00]	0.4885	[0.00]	0.4946	[0.00]	0.4798	[0.00]	0.4889	[0.00]	0.4980	[0.00]
3-month nominal rate (%)												
General gov't assets (% GDP)	-0.0116	[0.02]	-0.0168	[0.00]	-0.0221	[0.00]	-0.0121	[0.01]	-0.0149	[0.00]	-0.0172	[0.00]
Househ.ds' assets (% GDP)	-0.0191	[0.00]					-0.0183					
Househ.ds' net fin. Wealth (% GDP)			-0.0178	[0.00]	-0.0247	[0.00]		[0.00]	-0.0155	[0.00]	-0.0207	[0.00]
Non-fin. corp.ns' assets (% GDP)			-0.0014	[0.78]					-0.0051	[0.20]		
Non-fin. corp.ns' net Wealth (% GDP)	-0.0022	[0.62]			-0.0078	[0.11]	-0.0020	[0.58]			-0.0049	[0.08]
Foreign assets (% GDP)					-0.0005	[0.57]					-0.0014	[0.09]
Net foreign debt (% GDP)	0.0051	[0.28]	0.0043	[0.28]			0.0015	[0.70]	0.0013	[0.58]		
Foreign off. reserves (% GDP)	-0.0842	[0.06]	-0.1393	[0.01]	-0.1337	[0.00]	-0.0837	[0.02]	-0.1138	[0.00]	-0.1075	[0.00]
Fin. corp.ns' assets (% GDP)												
Fin. corp.ns' net Wealth (% GDP)	-0.0036	[0.72]	-0.0008	[0.92]	-0.0052	[0.44]	-0.0013	[0.86]	0.0005	[0.94]	-0.0043	[0.45]
Average life to maturity (years)	-0.1181	[0.21]	-0.1887	[0.04]	-0.1992	[0.05]	-0.1011	[0.23]	-0.1492	[0.05]	-0.1418	[0.09]
AA + (dummy)							0.2119	[0.16]	0.3225	[0.03]	0.3612	[0.01]
AA (dummy)							0.4605	[0.13]	0.5765	[0.01]	0.6380	[0.00]
AA - (dummy)							1.3928	[0.00]	1.7053	[0.00]	1.7443	[0.00]
A + (dummy)							2.0498	[0.00]	2.1132	[0.00]	2.1488	[0.00]
BBB + (dummy)							5.1739	[0.00]	5.2750	[0.00]	5.3083	[0.00]
BB + (dummy)							6.6748	[0.00]	6.6527	[0.00]	6.4626	[0.00]
Constant	6.0700	[0.00]	5.4568	[0.00]	5.2389	[0.00]	6.7892	[0.00]	6.0946	[0.00]	6.7892	[0.00]
Year dummies	No		No		No		No		No		No	
R-square	0.436		0.418		0.399		0.564		0.583		0.527	
Sample period	1995-2011		1995-2011		1995-2011		1995-2011		1995-2011		1995-2011	
Number of countries	18		18		18		18		18		18	
Number of observations	292		292		292		292		292		292	

Legend of model specification: [3] Heterogeneous time trend and economic restrictions; [4] Heterogeneous time trend, economic restrictions and ratings;
Panel estimates with fixed effects, run on yearly data. For each specification, the table shows coefficient estimates and, in square bracket, the related *p*-values

SECTION 2

THE TRANSMISSION OF GOVERNMENT BOND YIELDS TO BANK RATES AND LENDING CONDITIONS

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The impact of the sovereign debt crisis on the activity of Italian banks

Ugo Albertazzi*, Tiziano Ropele*, Gabriele Sene* and Federico M. Signoretti‡

May 2013

Abstract

We assess the effects of the sovereign debt crisis on Italian banks' activity using aggregate data on funding and loan rates, lending quantities and income statements for the period 1991-2011. We augment standard reduced-form equations for the variables of interest with the spread on 10-year sovereign bonds as an additional explanatory variable. We find that, even when controlling for the standard economic variables that influence bank activity, a rise in the spread is followed by an increase in the cost of wholesale and of certain forms of retail funding for banks and in the cost of credit to firms and households; the impact tends to be larger during periods of financial turmoil. An increase in the spread also has a direct negative effect on lending growth, beyond that implied by the rise in lending rates. Finally, we document a negative impact of the spread on banks' profitability, stronger for larger intermediaries.

JEL Classification: E44, E51, G21.

Keywords: sovereign spread, bank interest rates, bank lending.

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1. Introduction

After increasing gradually up to 200 b.p. between 2010 and the first half of 2011, the spread between the yield on the 10-year Italian government bond and the corresponding German one (henceforth called the BTP-Bund spread) rapidly increased throughout the summer of 2011, reaching a peak of 550 b.p. in November (Fig. 1). The widening of the sovereign spread reflected the sovereign debt crisis which first affected Greece in the first months of 2010, then involved Ireland and Portugal and finally reached Italy and Spain, assuming a systemic dimension.

As shown by the sharp increase recorded by CDS spreads (Fig. 2), the tensions in the sovereign debt market were swiftly transmitted to Italian banks, affecting both the cost and the availability of funding – especially on wholesale markets. The link between sovereign and bank risk reflected a number of different channels, such as the high exposure of banks to domestic sovereign debt,[1] the role of government securities as collateral in secured transactions and the connections between sovereign and banks' credit ratings.[2] The deteriorating economic outlook put additional strains on bank funding conditions.

The tensions on the funding side translated into a tightening of credit standards in the second half of 2011, as reported by the banks contributing to the *Euro Area Bank Lending Survey* (Fig. 3). Initially, the tightening was implemented mainly by increasing the margins applied to new loans, in particular to the riskier ones; in the most acute phase of the crisis, in the last quarter of 2011, the availability of credit to the private sector was also curtailed.

In this paper we seek to quantify the effect of the sovereign debt market tensions – proxied by the level of the 10-year BTP-Bund spread[3] – on the cost of funding for Italian intermediaries, the cost and availability of lending to firms and households, and the main items of banks' income and loss statements. In particular, for all the variables that we are interested in,

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[1] In June 2011, holdings of domestic government securities by Italian banks amounted to 6.3 per cent of total assets. Although this figure had declined during the years prior to the financial crisis, probably as a reflection of an increasing diversification allowed by the adoption of the single currency, it was still higher than what was observed in the other main euro area countries. It is worth mentioning however that Italian banks have a negligible exposure to sovereign borrowers of the other euro area countries under stress (Bank of Italy, 2011)

[2] In particular, banks' ratings tend to be downgraded shortly after sovereign ratings, because, among other reasons, the sovereign rating normally represents a ceiling for the ratings assigned to all other domestic borrowers (Bank of Italy, 2011). Once a bank is downgraded, "threshold effects" – such as the exclusion of a bank's liabilities from the basket of securities that certain categories of investor, such as pension funds and insurance companies, are allowed to purchase – can further worsen its funding conditions.

[3] The spread between 10-year BTP and Bund is the most common gauge of the risk premium demanded by investors for Italian government securities. It is worth noting that the dynamics of this indicator tends to overestimate the impact of sovereign strains on the cost of funds for the Italian government, as it also reflects flight-to-quality effects which tend to reduce the yield on German government securities. We present below a number of exercises, based on alternative measures of sovereign risk, to check the robustness of our findings.

we estimate reduced-form equations in which the sovereign spread is added, as an explanatory variable, to the standard determinants identified in the literature. Furthermore, in order to check for potential non-linear effects of the BTP-Bund spread on bank interest rates – in relation to the fact that the spread was basically zero throughout the 2000s – we also estimate equations in which the spread is interacted with dummy variables identifying two periods in which the level of the spread was high: the pre-EMU period (from 1991Q1 to 1997Q4) and the current sovereign debt crisis (since 2010Q2), on which our analysis will focus.

We draw on two data sources. One dataset contains quarterly aggregate information for the period 1991Q1-2011Q4, and includes previous episodes of tension on Italian sovereign debt, like those observed in 1992-93 and in 1995. This dataset permits an examination of the banks' cost of funding, average interest rates for loans to firms and households and the main items of banks' income and loss statement (net interest margin, other income and provisions), available on a quarterly basis. A second dataset, containing monthly information on bank lending and interest rates for a shorter period (January 2003 – December 2011), is used to assess the effects of the sovereign debt crisis on banks' activity with a finer sectorial breakdown. In particular, this dataset allows us to study separately loans of an amount up to €1 million, whose cost provides a measure of the interest rate paid by small and medium enterprises, and loans of a larger amount. Moreover, it permits a distinction to be made between fixed- and variable-rate loans to households for house purchase.

Italy is an especially good case for studying the effects of the sovereign risk on the banking sector. First, in Italy the causal relationship between the difficulties of the sovereign market and those of the banking sector during the current crisis is clear: unlike other European countries (Ireland and, to a large extent, Spain), problems originated in the public sector and then spilled over to the banking system.[4] This suggests that the sovereign spread can indeed be considered as an exogenous variable in our regressions. Second, the transmission of the tensions in the sovereign debt market to the banking sector is likely to be sizeable in Italy, due to the high level of public debt and to the heavy exposure of Italian banks to domestic sovereign bonds. Third, Italy experienced periods of tensions on its sovereign debt market also during the 1990s, which helps to identify the effects of the spread in the estimation.

Our analysis provides a number of results. *First*, variations of the BTP-Bund spread affect banks' funding cost rapidly and significantly: an increase in the spread is associated, at the

[4] Italian banks withstood the first phase of the financial crisis better than many foreign competitors mainly thanks to their greater reliance on a traditional business model and sound supervision.

latest with a one-quarter lag, with a sizeable rise of the remuneration on longer-term deposits, such as households' deposits with agreed maturity, as well as repurchase agreements and bonds. Such a relationship is strengthened during crises as opposed to "normal" times, suggesting the presence of a non-linearity. We instead find that the spread does not affect the return on overnight deposits, consistently with the sluggish adjustment of these yields to market conditions. *Second*, the BTP-Bund spread is found to exert a significant effect on the interest rate charged on loans to firms and on mortgages to households; we estimate that this effect largely reflects the increase of the marginal cost of funding – as proxied by the interest rate on term deposits. Also for loan rates, we find evidence of a non-linear transmission, as the effect of the spread is exacerbated during crises as opposed to normal times. The transmission is quantitatively larger than that observed on passive interest rates and occurs with a one-quarter lag. In a counterfactual exercise, where the spread is assumed to have remained constant at the level recorded in 2010Q1 (and all the other explanatory variables are assumed to have followed their actual development), we estimate that sovereign tensions contributed to increase interest rates on loans to firms and households by, respectively, 170 and 120 basis points. *Third*, changes in the BTP-Bund spread exert a significant direct effect on the dynamics of lending to both firms and households for house purchases, in addition to the indirect effect occurring through higher interest rates and the consequent lower demand for credit; in particular, we estimate that a 1 percentage point increase in the spread is directly associated with a 0.7 percentage point reduction of the annual growth rate on loans to firms. *Finally*, we find that tensions in the sovereign debt market have a significant negative impact on the profitability of the five largest Italian banking groups, affecting all the main items of the income and loss statement. For the Italian banking system as a whole, however, we find a negative effect only for loan-loss provisions, while we find a mildly positive relation for the net interest income and no effect on the other revenues; this finding is likely to reflect the lower importance of wholesale funding on the smaller intermediaries and the weaker responsiveness of their non-interest income to market conditions.

The rest of the paper is organized as follows. Section 2 briefly describes the different channels through which the sovereign debt crisis can affect banks' activity. Section 3 presents the analysis for banks' interest rates. Section 4 looks at the relationship between sovereign risk and lending volumes. Section 5 investigates the effects on the main items of banks' profit and loss statements. Section 6 draws conclusions.

2. The channels of transmission of sovereign risk to the banking sector

The tensions on sovereign debt, beyond influencing the general economic conditions of a country (through, for example, a fall in demand induced by fiscal adjustments or loss of confidence of households and firms) may also have more specific direct effects on the banking sector.

Following Panetta *et al.* (2011), González-Páramo (2011) and Holton *et al*. (2012), we can identify three main channels through which sovereign tensions may be transmitted to bank funding and credit supply conditions. First, a loss in the value of government bonds held in the portfolios of banks, through its effects on banks' income and possibly on their capital, can have an impact on a bank's funding ability and thus ignite a deleveraging process with a consequent reduction in credit supply (*balance sheet channel*). Second, given that government bonds in banks' portfolios are typically used as collateral in interbank transactions as well as in refinancing operations with central banks, the reduction in their value reduces banks' ability to borrow, and therefore to sustain credit supply (*liquidity channel*). A similar mechanism may operate when a bank's rating is downgraded following a reduction in sovereign rating, which is typically a ceiling for domestic private sector borrowers. As a consequence of the downgrading, the bank's liquidity position may be damaged for various reasons. For example, its liabilities may be excluded from the basket of securities that certain categories of investor, such as pension funds and insurance companies, are allowed to purchase, or it could receive calls for enhancing collateralization on ABS and covered bonds, or even lose the status of eligible counterparty for operations related to ABS. Third, the yield on sovereign debt may represent a benchmark for determining the cost of credit to the economy due to arbitrage-type mechanisms, given that government bonds are one of the most important investment opportunities available on the market (*price channel*). Moreover, the interest rate on bank deposits and bonds may also depend on the degree of solvency of the State, since this is perceived as the implicit guarantor of bank liabilities, and in particular of those not covered by an explicit private guarantee scheme[5].

All three channels imply that an increase in the yield of sovereign bonds can be expected to be associated with a rise in the cost of funding for national lenders and possibly a reduction in its availability, with repercussions on the cost and quantity of lending to the economy and on banks' profitability. The impact of the various channels may differ across banks' funding instruments, intermediaries with different characteristics or segments of the credit market. For

[5] All Italian banks participate in a deposit insurance scheme: the bigger intermediaries are compulsorily members of the *Fondo Interbancario di Tutela dei Depositi*; "mutual banks" (*Banche Di Credito Cooperativo*) are members of a separate fund.

instance, the *price channel* could be more important for bank bonds (which, unlike bank deposits, are not insured) and, among deposits, for longer-term ones, given that overnight deposits are primarily held for transaction purposes and exhibit rather sluggish remunerations. Mortgages may be expected to be less severely affected by the liquidity channel compared with corporate loans, as the former can be more easily pooled and used to guarantee ABS, which in turn can be sold to the market or used as collateral in refinancing operations with central banks.

According to the banks' answers to specific *ad hoc* questions introduced in the December round of the *Euro Area Bank Lending Survey*, concerning the fourth quarter of 2011, financing conditions of Italian intermediaries were markedly affected by the turmoil in the sovereign debt market; tensions were transmitted through all the three channels, and especially through the balance sheet channel (Fig. 4). Concerning the effects on credit supply, the banks declared that the business loans segment was the one most severely hit.

3. Bank interest rates

We start our analysis by examining to what extent banks change their interest rates in response to changes in the BTP-Bund spread. We first consider the interest rates paid on liabilities, i.e. the cost of different components of banks' funding, and then the interest rates charged on loans. Figures 5 and 6 show, respectively for passive and active rates, the time series of selected interest rates in Italy since 1990, together with the BTP-Bund spread and the monetary policy rate (for most of the time both passive and active interest rates shadow the latter). The figures show that, since the spring of 2010, marked interest rate increases were associated with the widening of the BTP-Bund spread.[6] On the liability side, large increases were observed for the interest rates on deposits with agreed maturity, repos and debt security yields, while interest rate on overnight deposits, which are typically less sensitive to market conditions, barely reacted to changes in the sovereign spread. As for lending rates, the cost of loans to firms and to households for house purchases increased in the second half of 2010 and, more markedly, in 2011; the cost of consumer credit increased very moderately and only in 2011.

In light of this preliminary descriptive inspection, we proceed with a formal econometric analysis to quantitatively assess the effect of the BTP-Bund spread.

[6] Over the same period, money market interest rates did not show movements of comparable size: short-term rates increased slightly in 2010, following the April and July official rate increases by the ECB, and then declined in the last months of 2011, when the ECB cut official rates.

3.1 The empirical methodology: an ARDL model

To assess the impact of the BTP-Bund spread on banking interest rates we use an autoregressive distributed lags (ARDL) model. Such an approach has been extensively used to study the transmission of changes in monetary policy rate to the banking rates (Cottarelli and Kourelis, 1994; Favero et al., 1997; Marotta, 2010).[7] For the scope of our analysis we specify the following ARDL model, which is estimated via OLS[8]:

$$\begin{aligned} i_t^B = c + \sum_{j=1}^{p} \alpha_i \; i_{t-j}^B + \sum_{j=1}^{p} \alpha_i \; i_{t-j}^M + \sum_{k=0}^{r} \beta_k \; X_{t-k} + \gamma \; spread \; + \\ + \theta_1 \, D_{SovCrisis} \times spread \; + \theta_2 \, D_{SovCrisis} + \phi_1 \, D_{preEMU} \times spread_{t-1} + \phi_2 \, D_{preEMU} + \varepsilon_t \end{aligned} \tag{1}$$

where the dependent variable i^B is the bank interest rate (passive or active) under examination; i^M is the relevant (policy or market) interest rate (possibly at various maturities) for the bank interest rate considered; X is a vector of macroeconomic variables (such as GDP growth, unemployment rate, disposable income) used to control for economic activity and to proxy borrowers' creditworthiness. In addition to these explanatory variables, we include the variable *spread*, calculated as the difference between the 10-year yield on Italian BTP and that on German Bund minus the difference between the 10-year swap rate in Italy and Germany (Fig. 7)[9]. In the regressions, the variable *spread* is also interacted with two dummy variables, in order to check for potential additional (or differentiated) effects of this variable when it reaches high levels, as opposed to "normal" times (when it fluctuates at low levels): the dummy D_{preEMU} identifies the pre-EMU period, and takes value 1 from the beginning of the sample (1991Q1) until 1997Q4 and zero elsewhere;[10] $D_{SovCrisis}$ identifies the sovereign debt crisis, and takes value 1 from 2010Q2 until the end of the sample (2011Q4) and zero elsewhere.

The choice of both the lag structure and the appropriate market/monetary policy rate to be included in the various regressions is based on a simple correlation analysis (De Bondt, 2005) and also takes into account the goodness of fit, the statistical significance of the coefficients and the presence of autocorrelation in the residuals.

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[7] Also depending on the nature of the dataset available, other empirical approaches are possible. For example, Sørensen and Werner (2006) investigate the heterogeneity in the pass-through process of money market rates to bank interest rates across euro area countries with panel-econometric methods.

[8] Standard errors are computed with the Newey-West correction for heteroskedasticity and autocorrelation.

[9] As in Favero, Giavazzi and Spaventa (1997), we adopt this adjustment of the BTP-Bund spread in order to make sure that in the pre-EMU period our measure of sovereign risk is not contaminated by other factors that may affect the rates on long-term bonds (such as expectations of the future conduct of monetary policy, inflation differentials, etc.). A thorough robustness check, considering alternative measures of sovereign risk, confirms our main results (see Table A2 in the Appendix).

[10] During that period (1991Q1-1997Q4) the (non-adjusted) spread (as quarterly averages) was always above the maximum level 40 b.p.) reached between 1999Q1 and 2008Q2 (before the Lehman collapse); in 1998 the spread was always lower than 40 b.p..

3.2 Cost of funding

One distinguishing feature of Italian intermediaries is their reliance on stable sources of funding, such as retail deposits and bonds placed with retail customers, whose cost is generally insensitive to market volatility. In particular, overnight deposits – whose rates are typically very sluggish – amounted to almost two-thirds of the sum between total deposits and bonds at the end of 2011 (Bank of Italy, 2012). The composition of funding for Italian banks is likely to have helped to moderate the increase in the average interest rate on deposits recorded since the beginning of 2010. In fact, the increase mainly reflected the marked rise of the rates for deposits with agreed maturity, which displayed a strong correlation with the sovereign spread over the last two years (see Fig. 5).

Table 1 reports the results for banks' cost of funding. In particular, we consider separately yields on: households' overnight deposits (columns *i* and *ii*), households' deposits with agreed maturity[11] (column *iii* and *iv*), repurchase agreements (column *v* and *vi*) and bank bonds (column *vii* and *viii*).

For all of the instruments considered, we find that the coefficients for the standard explanatory variables are significant and show the expected sign. In particular, the remuneration is closely related to the money market interest rates (the rate on three-month interbank transactions for all the instruments considered and also the three-year swap rate for bank bonds). Moreover, the yields on all these instruments tend to show a significant degree of persistence, as indicated by the large coefficient on the lags of the dependent variable.

As for the BTP-Bund spread, we find that its impact is different for the various instruments considered. It does not appear to affect the return on overnight deposits: its coefficient is very small (column *i*) and does not become significant when we add the time dummies. Three factors may help to explain this finding: demand deposits are covered by the deposit insurance, which reduces the influence that changes in the level of the risk perceived on banks' liabilities may have on their remuneration; unlike other types of deposits, the financial duration of the demand deposits is nil, which further attenuates the risk-premium component of their returns; overnight deposits are primarily held for transaction purposes and their remuneration is therefore less reactive to market returns.[12]

[11] Time series for the returns on overnight deposits and deposits with agreed maturity held by other sectors are not available with a long time-span.

[12] Very similar findings are obtained for the overnight deposits held by non-financial corporations, which, at the end of 2011, represented approximately one-fifth of the total amount of the overnight deposits of Italian banks.

The BTP-Bund spread plays however a relevant role when we consider the yields on the other funding instruments, for which we find a significant and sizeable effect, both when the spread alone is considered directly and when it is interacted with $D_{SovCrisis}$. The size of the estimated coefficients indicate that, in normal times and *ceteris paribus*, a temporary (i.e., lasting for one quarter) 100 b.p. increase in the spread is associated, within the same quarter, with a 34 and 21 b.p. increase, respectively, in the interest rate paid on households' deposits with agreed maturity and repurchase agreements; such an effect has been bigger during the sovereign debt crisis, reaching around 40 b.p. for both instruments.[13] The effect is even larger for the remuneration banks pay on newly issued bonds and the funding component is more sensitive to market conditions: 70 b.p. in normal times and over 100 b.p. in the crisis. This latter finding provides empirical support for the relevance of the price channel reviewed in section 2.[14]

Table A1 in the Appendix presents analogous estimations conducted on monthly data, running from January 2003 to December 2011: the results are qualitatively similar, though the coefficients are somewhat smaller (which could reflect, at least in part, the fact that the shorter monthly sample does not include the sovereign tensions experienced in the early 1990s).

3.3 Interest rates on loans

Table 2 reports the estimation results for the interest rates applied on short-term loans to firms (columns *i* and *ii*), new loans to households for house purchases (columns *iii* and *iv*) and consumer credit and other households' loans (columns *v* and *vi*). For all these rates we find a positive and significant effect for the monetary policy rate and for the autoregressive component; we also include GDP growth, the unemployment rate and households' disposable income in the regressions as controls for the macroeconomic outlook and changes of borrowers' creditworthiness.[15]

Turning to the effect of the sovereign spread, which enters these regressions with one-quarter lag, and considering the specifications without the time dummies (columns *i*, *iii* and *v*), we find that the coefficients are positive, significant and equal to around 20 b.p. Once we consider the regressions with the time dummies (columns *ii*, *iv* and *vi*), the estimates for the

[13] The sum between the coefficients on BTP-Bund spread and that on its interaction with *Dummy sov_crisis*, equals 38 and 35 b.p., respectively, for households' deposits with agreed maturity and for repos.

[14] These bonds include both the securities placed with retail customers as well as those sold on the international financial markets.

[15] In the case of interest rates on loans, the official monetary policy rate (the official discount rate of the Bank of Italy until 1998Q4 and the minimum interest rate on ECB main refinancing operations since 1999Q1) is the short-term rate that yields the best fit. Results are analogous if we use alternative measures, such as the three-month Euribor or three-month e-MID.

interacted terms are larger (and more significant) than the ones found with the previous regressions, in particular for firms, while the coefficients for the direct terms do not become statistically significant. This result suggests the presence of “non-linear” effects in the pass-through of the spread, as the effect becomes quantitatively more sizeable during periods when the spread is high.[16]

Based on the size of the estimated coefficients, we can calculate that *during the sovereign debt crisis* the response of loan rates to a temporary 100 b.p. increase in the BTP-Bund spread was around 50 b.p. for firms and 30 b.p. for households’ mortgages. In the case of a *permanent* increase of the sovereign spread, the pass-through after one year would be complete for loans to firms, while it would be 83 b.p. for mortgages, reflecting the higher persistence shown by the cost of these loans.[17]

As already pointed out, the spread between 10-year BTP and Bund could overestimate the impact of sovereign strains, as it may reflect flight-to-quality effects which tend to reduce the yield on German government securities. In order to check the robustness of our findings we run two exercises. First, we re-estimate the regressions for active interest rates considering different measures of the sovereign risk, namely: the adjusted BTP-Bund spread at shorter maturities and the corresponding unadjusted spread; the three-, five- and ten-year yield on BTP and on interest rate swaps (IRS); the spread between the BTP and the French government bonds (OAT) yield at different maturities.[18] Table A2 in the Appendix shows that our estimates are very robust.[19] In particular, the results are quantitatively similar, both when considering the specification without interaction terms and the interaction terms for the sovereign debt crisis period.

Second, we modify the benchmark model for the loan rates by including the level of the yield on the 10-year Bund and its interaction with the dummies for the two crisis periods (D_{preEMU} and $D_{SovCrisis}$) as additional explanatory variables. For loans to firms, the results (not reported) indicate that during the sovereign crisis the yield on 10-year Bunds is significant (although marginally) and with a positive coefficient. For loans to households it is not significant. This finding suggests that the transmission to loan rates of an increase in the

[16] It is interesting to note that while the size of the coefficient for the interacted term relative to the pre-EMU period is similar to the one observed in the sovereign debt crisis for mortgages, it is somewhat smaller for loans to firms; this finding probably reflects different structural conditions in the banking sector and more intense competition.

[17] The estimated autoregressive coefficient is 0.52 for loans to firms and 0.81 for household mortgages (see columns *ii* and *iv* in Table 2).

[18] The spread vis-à-vis French government bonds (OAT) provides a measure of sovereign risk which is likely to be less affected by the flight-to-quality phenomenon, whereby the demand for German Bunds increases due to the safe-haven status of these securities.

[19] The table reports only the estimated coefficients for the sovereign risk measure and its interactions with the pre-EMU and sovereign debt crisis dummy variables. The estimated coefficients on other explanatory variables are basically unchanged.

sovereign spread may be somewhat smaller than the one estimated in the baseline specification if such increase reflects a reduction in the Bund yield rather than an increase in the BTP yield.

As a further robustness check we carry out our estimation exercises using the monthly dataset, which also allows us to analyse the pass-through of the sovereign spread at a finer sectorial breakdown (Tables A3 and A4 in the Appendix). These results confirm the significant effect of the sovereign spread on lending rates, which has become stronger during the sovereign debt crisis. The pass-through is roughly similar for rates on small loans to firms (up to €1 million) and for larger ones (over €1 million). As for mortgages, the pass-through is significant and approximately of the same magnitude for variable-rate and for fixed-rate loans (though for the latter we find a significant coefficient for the direct effect but not for the interacted term).

The cost of lending during the sovereign debt crisis: a counterfactual exercise

In light of the results described above, we conducted a simple counterfactual exercise in order to see what would have happened to bank interest rates if the BTP-Bund spread had remained unchanged at the level observed in 2010Q1, i.e. at 70 b.p.. For this purpose, we rely on the estimated coefficients obtained using all the sample periods and we employ the observed time series for the main macroeconomic variables, namely unemployment rate, GDP growth and market rates.[20] The results indicate that, under the hypothetical scenario, the cost of lending to firms in 2011Q4 would have been about 170 b.p. lower than its actual value (Fig. 8), while the cost of new mortgages would have been about 120 b.p. lower. In both cases, half of the final effect is cumulated in the period 2010Q2 to 2011Q3, while the other half is attributable to just the fourth quarter of 2011, reflecting the large increase recorded by the BTP-Bund spread in the previous quarter (about 160 b.p., the largest quarterly increase in the sample period).

As already pointed out, part of the increase in the BTP-Bund spread is connected with flight-to-quality phenomena. Thus, we have performed an additional counterfactual exercise including in the specification the yield on the 10-year Bund and its interaction with the time dummy variables (see the robustness check above). In this case the counterfactual experiment is carried out holding unchanged the BTP-Bund spread as well as the yield on the 10-year German bond at their respective levels observed at the end of 2010Q1. The result shows that the cost of lending to firms in 2011Q4 would have been about 150 b.p. lower than its actual value. The

[20] By using the observed time-series for the main macroeconomic variables, we do not take into account the fact that also these variables would probably have had a different path in the counterfactual scenario, affecting the measure of the counterfactual loan rate. Nonetheless, the size of the estimated coefficients indicate that such indirect effects are likely to be of a second-order magnitude. For instance, in the equation for short-term loans to firms, a percentage point reduction of the unemployment rate is associated with an implied reduction of the interest rate of about 5 b.p., everything else being equal.

somewhat smaller estimated impact of the sovereign tensions – with respect to the counterfactual based on the benchmark specification – reflects the reduction by around 110 b.p. of the Bund yield (which enters with a positive, though marginally significant, coefficient in the equation) occurred between 2010Q2 and the end of 2011.

An exercise with banks' costs of funding

According to the results of sections 3.2 and 3.3, the estimated pass-through of the BTP-Bund spread on the loan rates is stronger than that on banks' funding costs, if we average out the estimated effect on the various funding components; the stronger impact on loan rates mainly reflects the fact that the greatest contribution to the average cost of funding comes from overnight deposits, whose yield is not significantly affected by the BTP-Bund spread. This finding seems to suggest that banks price their new loans by taking into account their *marginal* cost of funding, which is much more reactive to changes in funding conditions, rather than their *average* cost of funding.

A simple way to check whether this notion is correct is adding a proxy for the marginal cost of funding to the regressions for rates on loans to firms and for rates for household mortgages. If our explanation is correct, we should find that the cost of funding crowds out the effect of the BTP-Bund spread. The marginal cost of funding is proxied by the yield on the component of the funding which are subject to a more frequent repricing, such as term deposits, and for which we have indeed found a stronger impact reactivity to the sovereign spread, compared to other forms, such as overnight deposits.

Indeed, when we add the marginal cost of funding to the firm rate equation, the BTP-Bund spread turns out not to be significant in the regression without the interaction terms (Table A5 in the Appendix, column *ii*) and the overall impact diminishes in the regression with the interaction terms (column *iii*).[21] In those regressions, the coefficient for the marginal cost of funding is positive and highly significant. For loans to households for house purchase the impact of the cost of funding is not statistically significant, but its inclusion in the regression eliminates the significance of the coefficient of both the BTP-Bund spread and the interaction between this term and the sovereign debt crisis dummy (columns *vi* and *viii*).

[21] The sum between the coefficients on the BTP-Bund spread and that on its interaction with *Dummy sov_crisis*, equals 51 b.p. in the baseline regression (column *iii*), compared with 18 b.p. in the regression with the cost of funding (column *iv*).

The above results confirm that the impact of the sovereign spread on loan rates stems to a large extent from the increase of the marginal cost of funding. As already mentioned, the remaining impact of the spread on loan rates might result from the increase in firms' riskiness associated with the deterioration of the economic outlook and with the reduction in banks' willingness to lend, which is consistent with survey evidence relating to the most acute phase of the crisis, beyond what is captured by the macroeconomic indicators included in the regressions.[22]

4. Lending volumes

This section analyses the impact of sovereign risk on the amount of lending. An increase in the sovereign spread may reduce credit volumes via the *indirect* effect connected with the increase in the loan rates which, also depending on the coefficients of elasticity of demand and supply in the credit market, may in turn reduce the amount of credit in the economy in equilibrium. Moreover, spreads could also have a *direct* effect on loan quantities, to the extent that tensions in funding markets prompt banks to conduct an outright rationing of lending supply. For Italian banks, the *Bank Lending Survey* suggests that direct effects on lending volumes and indirect ones via cost of credit coexisted in the final part of 2011, reflecting the significant funding difficulties of intermediaries on wholesale financial markets. Understanding whether sovereign tensions directly reduced lending volumes at that time is important also for assessing the usefulness of the three-year refinancing operations launched by the Eurosystem in December 2011 and February 2012: by alleviating strains in banks' funding, these operations may have directly supported financing of the real economy beyond the indirect effect induced by the provision of cheap and very long-term funding.

In light of these considerations we explore the implications of the BTP-Bund spread on lending activity with the aim of distinguishing between the direct and the indirect effects. To this end, we specify two simple regressions, one for the 12-month growth rate of loans to firms and the other one for the 12-month growth rate of new loans to households for house purchases. The general specification is given by:

$$y_t^j = c + \alpha\, y_{t-1}^j + \beta\left(i_{t-1}^B - i_{t-1}^M\right) + \eta\, i_{t-1}^M + \sum_{k=0}^{q} \delta\, X_{t-k}^B + \gamma\, spread_{t-1} + \varepsilon_t \qquad (2)$$

[22] Moreover, to the extent that banks have suffered from rationing-type phenomena on their funding sources – such as those that arguably happened on selected wholesale funding markets in the most acute phase of the sovereign crisis – our measure of the marginal cost might underestimate their actual *shadow* cost of funding. In turn, also our estimate of the intensity of the transmission from funding costs to loan rate might be stronger than the one we estimate.

In (2) the dependent variable y_t^j is the growth rate of lending to sector *j* in the corresponding quarter. The explanatory variables are the autoregressive term (y_{t-1}^j), the difference between the cost of credit (i^B) and the short-term market interest rate (i^M)[23], the BTP-Bund spread, and a number of macroeconomic controls (X_{t-k}), which include GDP growth, unemployment rate, firms' financial needs, household consumption expenditures, house price growth (for the exact specification of the equations for firm and household loans, see Table 3).

Table 3 shows that all macroeconomic determinants exhibit the expected sign and are statistically significant. In particular, the growth of lending to firms is positively associated with firms' financing needs (calculated as the ratio between the corporate sector's investments and gross operating profit), and nominal GDP growth, while it is negatively associated with the three-month interbank interest rate and the spread between the cost of lending and the three-month interbank interest rate. The growth of new loans for house purchases is positively related to house price growth, and negatively to the spread applied on new mortgages and to the level of short-term interest rates; the dynamics of these loans is also significantly related to business cycle conditions.[24] Most importantly for the purpose of our paper, we find a significant and negative effect stemming from the BTP-Bund spread on the growth of loans to both firms and households.[25] The impact of such an effect can be quantified in a reduction of 0.7 percentage points of the annual growth rate on loans for every 100 b.p. increase in the sovereign spread.[26]

Loan developments during the sovereign debt crisis: a counterfactual exercise

Similarly to what we presented in Section 3, a simple counterfactual exercise, in which the BTP-Bund spread is assumed to remain at the level observed in 2010Q1 (70 b.p., and everything else being equal), indicates that the BTP-Bund spread affected the amount of lending

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[23] In particular, we use the three-month interbank interest rate on transactions conducted on the e-MID market; e-MID is a multilateral platform for interbank deposits in Europe. Intermediaries operating with e-MID are from about 30 countries, though a significant number are from Italy. The average interest rate on e-MID is therefore representative of the cost of interbank transactions specific to Italy.

[24] The specification adopted includes two controls for business cycle conditions. First, as expected, the growth of mortgages is positively related to that of GDP and consumption expenditure. Second, it is also positively influenced by the unemployment rate, possibly capturing some safe-haven effects related to house purchases.

[25] As for loan rates, we tried to include the interaction between the spread and the time dummies as explanatory variables also in these regressions. The results indicate that there are no significant non-linear *direct* effects of the spread on lending growth.

[26] It is important to remember that effect of the *spread* here only captures the *direct* effects on loan quantities, while the indirect effects connected with the increase in lending rates is captured, in the regression, through the coefficient on the loan interest rate.

significantly only in the second half of 2011. In particular for 2011Q4 we can quantify this impact in a reduction of about 2 percentage points of the (annual) growth of loans to both firms and households, considering the direct effect as well as the effect through the cost of lending (Fig. 9).[27] As shown by the decomposition in the figure, the largest part of this reduction can be attributed to the direct effect.

5. Income and loss statement

In the previous sections we have documented the impact of sovereign risk on banks' interest rates and lending volumes. In this section, we study the effect of the sovereign spread on the profitability of banks, analyzing separately the various components of banks' income statements: interest income, trading income and other revenues, loan-loss provisions.[28] The following comments apply as regards the expected effects.

The impact of the BTP-Bund spread on banks' net interest income is *a priori* ambiguous, also in light of the results that we found in sections 3 and 4. On the one hand, a rise in the BTP-Bund spread tends to reduce lending volume (through both direct and indirect effects) and to increase bank funding rates; both effects tend to compress the net interest margin. On the other hand, we have documented that banks increase loan rates in response to a rise in the spread and the increase is typically larger than the one in funding rates (in particular given the high incidence of demand deposits and, more generally, of retail funding for Italian banks); this mechanism may thus contribute to increasing the net interest income. The assessment of the overall effect of sovereign risk on the interest margin is therefore an empirical matter.

As regards non-interest income and other revenues, a depreciation of government bonds is likely to induce losses in proprietary trading; tensions could also possibly affect income from fees and commissions, for example, through a decline in trading volumes. In this respect, it is however important to bear in mind that only a small part of the government securities held by Italian banks are in the trading portfolio, for which changes in value directly affect the income

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[27] The counterfactual series are obtained by adding back to the actual values of lending growth the contribution of the difference between the actual and the counterfactual BTP-Bund spread. The *direct* effect is calculated using the coefficient for the spread in the loan equation; the *indirect* or *price* effect is the contribution occurring via the effect of the spread on the loan rate and is thus calculated using the product between the coefficient of the spread in the equation for the loan rate and the coefficient of the loan rate in the equation for lending. This methodology does not take into account the autoregressive structure of the loan rate and credit growth; if this were included, the difference between actual lending growth and growth in the counterfactual exercise would be larger.

[28] We do not consider banks' operating costs as they mainly reflect structural factors and thus are likely to be unresponsive, at least in the short and medium term, to sovereign debt tensions.

statement;[29] most of the sovereign debt is instead included in the available-for-sale portfolio, whose changes in value do not have direct repercussions on the income statement.

Finally, sovereign risk may also have a negative impact on banks' loan-loss provisions *beyond* the indirect effect connected with the deterioration of business cycle conditions and the ensuing worsening of credit quality.[30] For example, the increase in sovereign risk could worsen the financing position of firms or the scenario of future fiscal consolidation could depress the expected income of both households' and firms, weakening their debt repayment capacity.

For each of the above-mentioned components of banks' income statements we estimate the following OLS regression, on a quarterly dataset that runs from 1991Q3 to 2011Q4:

$$\log(y_t) = c + \sum_{i=1}^{P} \alpha_i \log(y_{t-1}) + \sum_{k=0}^{q} \beta_k X_{t-k} + \gamma\, spread_{t-1} + \varepsilon_t \qquad (3)$$

The baseline regression (3) generalizes the specification used in existing studies on the determinants of banks' profitability (see, for example, Demirgüç-Kunt and Huizinga, 1999; Casolaro and Gambacorta, 2005; Albertazzi and Gambacorta, 2009) in order to assess the role of sovereign risk for bank profitability. The set of explanatory variables (X_t) includes the stock market index and its volatility in addition to some of those used in the analysis of interest rates and credit, such as nominal GDP, unemployment rate, short- and long-term interest rates. As in the previous sections, the impact of the sovereign spread is evaluated by adding the 10-year BTP-Bund spread as an explanatory variable. All regressions include a set of dummy variables controlling for seasonal effects and outliers.

Table 4 shows the results. In column (*i*) we see that net interest income is positively affected by the nominal GDP, reflecting the increase in lending demand by the private sector in periods of higher economic activity. The interest margin also displays a positive relationship with the short-term interest rate, consistent with the faster reaction of loan rates, when compared with deposit rates, to changes in market rates. Also the long-term interest rates have a positive coefficient, probably reflecting the beneficial effect on income of an increase in the slope of the yield curve, connected with banks' maturity transformation activity.

[29] For the five largest banking groups, the share of domestic government securities held in the trading book at the end of 2011 was around one-quarter (Bank of Italy, 2012).

[30] See Bofondi and Ropele (2011) for a comprehensive study of the macroeconomic determinants of banks' loan quality in Italy in the past 20 years.

The effect of the BTP-Bund spread is positive and highly significant, suggesting that the stronger reaction of the loan rates to changes in the spread more than offsets its impact on funding cost and lending volumes; this is consistent with the increase in the interest margin observed in 2011, concentrated in the second half of the year, in parallel to the increase in the sovereign spread. A breakdown of the data by bank size, however, reveals that the whole increase was driven by the behaviour of smaller banks, for which the interest margin increased by 19 per cent, while banks belonging to the five biggest groups recorded a contraction of 6 per cent. We thus rerun our regression only for the first five groups, finding that the spread has a *negative* impact on interest margin (column *ii*), while all the other coefficients remain virtually unchanged. These findings can be explained by a higher impact of the sovereign debt tensions on the largest banks' funding costs, due to their greater reliance on wholesale funding sources, which were the instruments most affected by the crisis. The result for the five largest groups is also consistent with hard data on lending, which show that in 2011 loans decelerated more for these banks than for the rest of the system (Bank of Italy, 2012).

Columns (*iii*) and (*iv*) report the results for the non-interest income equation, for all banks and for the five largest groups only, respectively. In both samples, the coefficient of GDP is positive (though not statistically significant), reflecting the correlation between economic activity and the demand for banking services. Moreover, the effect of both the short- and long-term interest rates is negative, while the coefficient for the stock market index is positive (and highly significant). These results may reflect a negative correlation between trading income and asset prices; another possible interpretation is that, when interest rates are low, savers have more need of professional services provided by banks in order to manage their own portfolios, which increases income from fees and commissions.[31] As regards the effect of the BTP-Bund spread, we find a significant (negative) coefficient only when estimating the equation for the five largest groups, whose non-interest income is likely to be more responsive to financial market conditions.

Finally, the estimates for loan-loss provisions are shown in columns (*v*) and (*vi*); in this case, the results for the whole sample and for the five largest groups are very similar. As expected, loan-loss provisions are negatively related to GDP growth, while stock market volatility, which can be considered as a proxy of risk, has a positive effect (thus a negative effect on profitability). The BTP-Bund spread is also positively related to loan-loss provisions,

[31] For a discussion, see Albertazzi and Gambacorta (2009).

suggesting the presence of a direct transmission between sovereign risk and private non-financial borrowers' risk.

6. Conclusions

We have presented a comprehensive analysis of the effects of sovereign debt tensions on banking activity in Italy, focusing in particular on the crisis of 2010-11. The empirical analysis is based on aggregate data for the cost of funding of Italian intermediaries, the cost of credit they extend, the dynamics of lending as well as the main items of their profit and loss statements.

Our findings indicate that sovereign debt tensions, as measured by the evolution of the BTP-Bund spread, exert significant effects on most of the variables considered. Among funding rates, the strongest impact is on time deposits, repurchase agreements and newly issued bonds, while we find no effect of the spread on the yield of overnight deposits, which account for the bulk of banks' deposits in Italy. The effect on rates on lending to both firms and households are statistically and economically significant and reflect to a large extent the increase in bank marginal cost of funding. The sovereign spread also affects lending volumes directly, beyond the indirect effect exerted through its impact on the cost of credit. Finally, we find that the spread has a negative effect on the profitability of the largest banking groups, unfavorably affecting all the main items of their income statements; when we consider the whole banking system, we find a negative impact only for loan-loss provisions.

There is evidence of non-linearity in the effects of the BTP-Bund spread on active and passive interest rates: the estimated pass-through increases during periods characterized by a high level of the spread, such as the pre-EMU period or the current sovereign debt crisis, roughly doubling for interest rates on loans. In particular, during the sovereign debt crisis, a temporary 100 b.p. increase in the sovereign spread in a given quarter is associated with an increase (at the latest in the following quarter) of around 40 b.p. for the yield of retail time deposits and repurchase agreements, and of around 100 b.p. for the bond yields; no pass-through is observed on the return on retail overnight deposits, consistently with the sluggish adjustment of these yields to market conditions. The pass-through to new loan interest rates is, respectively, around 50 and 30 b.p. for loans to firms and to households for house purchases, with a one-quarter lag. If we consider a permanent 100 b.p. increase in the spread, we estimate that the rise in the loan rate after one year would be of the same magnitude for new loans to firms and of 80 b.p. for household mortgages.

A counterfactual analysis suggests that, at the end of 2011, loan rates would have been at least 170 and 120 b.p. lower, respectively for firm loans and household mortgages, and lending growth (for both credit market segments) about 2 percentage points higher than what was actually observed, had the spread remained unchanged at the level of 2010Q1.

Although our results point to a strong transmission of the sovereign spread to the cost of credit, one should be cautious in drawing implications for financial stability resulting from a rise in the Italian sovereign spread. In this regard, the key variable is the private sector debt-service burden, i.e. the cost of repaying the debt – principal and interest – for firms and households. At the current juncture, several factors tend to dampen the effect of an increase in the cost of lending on this variable. First, by construction, our estimates of the pass-through of the BTP-Bund spread are conditional on the risk-free interest rates used in our specifications (alternatively, the Eonia, the 3-month Euribor or the monetary policy rate); these rates decreased during the crisis, reflecting flight-to-quality phenomena or monetary policy decisions, attenuating the overall impact of the tensions on loan rates. In particular, the 3-month Euribor, to which most Italian household mortgages are indexed, stood, in 2012, at historically low levels; this significantly contributes to reduce payments on existing mortgages for Italian households. A related consideration is that the estimated BTP-Bund spread coefficient itself may overestimate the impact of sovereign strains, as it neglects flight-to-quality effects which tend to reduce the yield on German government securities. Second, the level of indebtedness of Italian households and firms is low in the international comparison (Bank of Italy, 2012), which helps attenuating the negative impact of any interest rate rise on these sectors on aggregate. Third, our analysis refers to *new* businesses or short-term loans, while the overall effect of the sovereign spread on the debt-service burden largely depends on the interest rates on *outstanding* loans; for fixed rate loans that were priced before the beginning of the sovereign tensions there is, by definition, no transmission. The estimated increase in the cost of new credit would fully translate to an increase in the debt-service burden of firms and households only in the case of a permanent increase in the sovereign spread.

We can think of at least two immediate follow-ups to the analysis developed in this paper. First, a methodological limitation of our approach is to only concentrate on the direct effects of sovereign market stress on banks' activities, while ignoring its potential general equilibrium effects; in particular, sovereign tensions are likely to bring about a weakening in macroeconomic conditions which, in turn, affects bank balance sheet conditions and income. It would thus be interesting trying to endogenize these feedback effects by resorting to alternative econometric techniques, such as VAR. Second, another direction for possible future research would aim at pointing out potential heterogeneity across banks in the transmission of sovereign risk; this would require an analysis based on bank-level data, which could also be useful in order to disentangle the relative importance of the different transmission channels.

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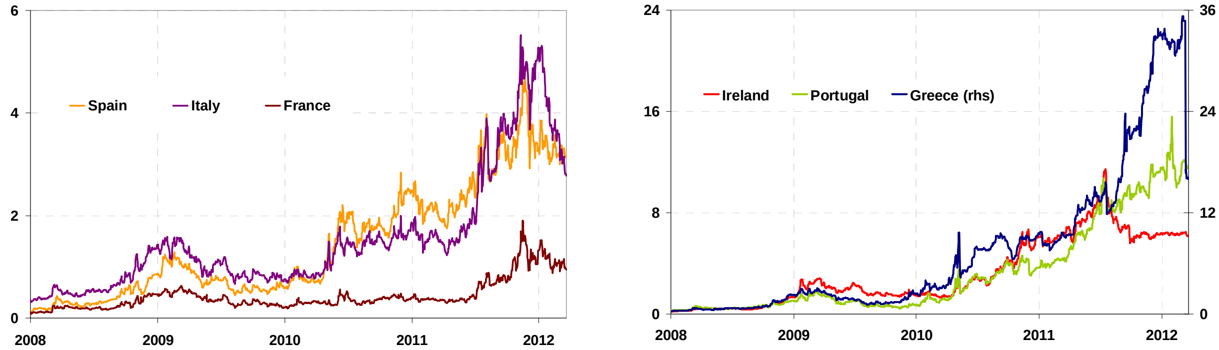
FIGURES AND TABLES

Figure 1

Sovereign spreads for selected euro area countries (1)

(*percentage points*)



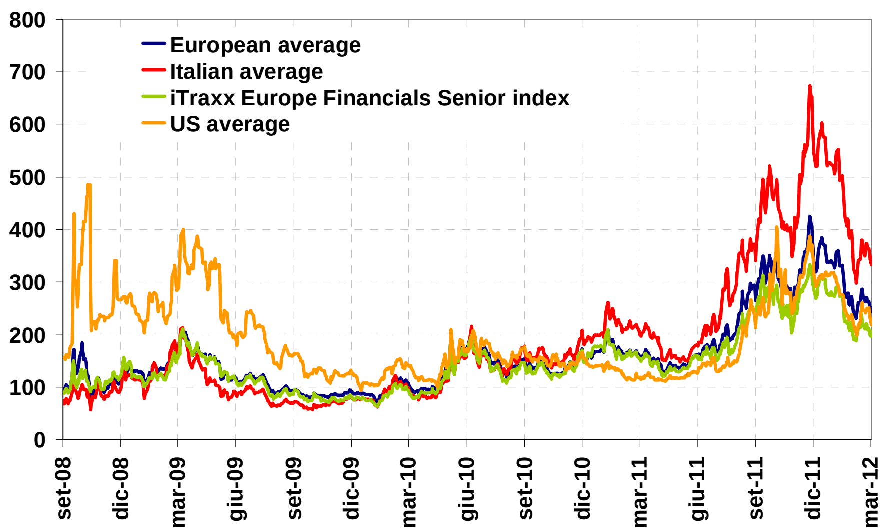
Source: Bloomberg.

(1) Spread on 10-year government bonds with respect to the corresponding German Bund.

Figure 2

Banks' CDS premia

(*basis points*)

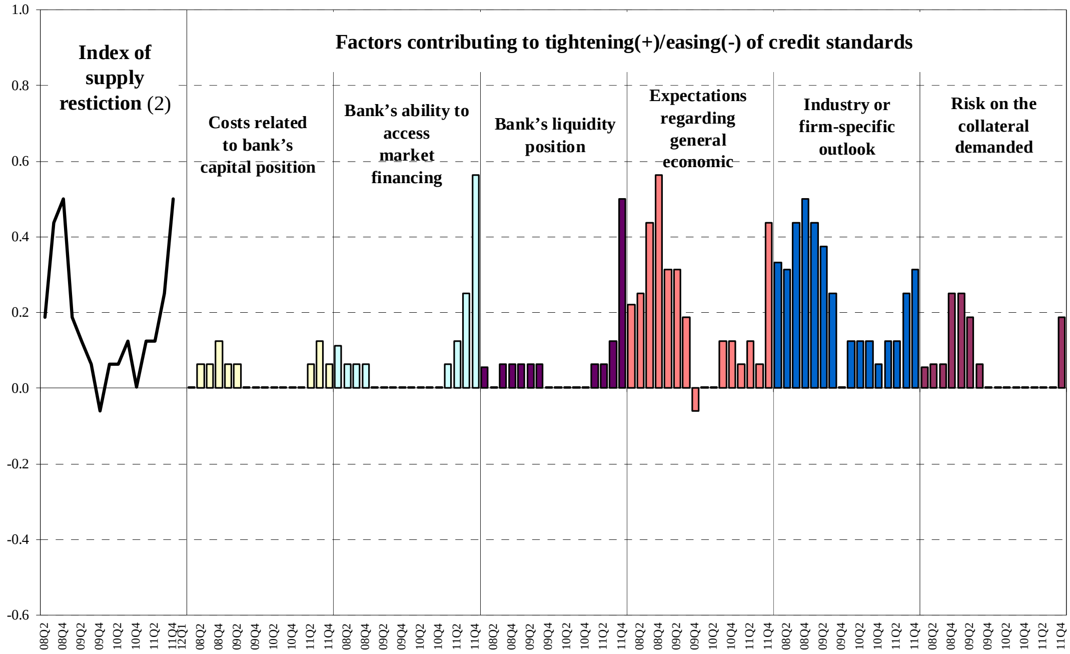


Source: Bloomberg.

Figure 3

Euro Area Bank Lending Survey: Italian panel (1)

(*diffusion indices*)



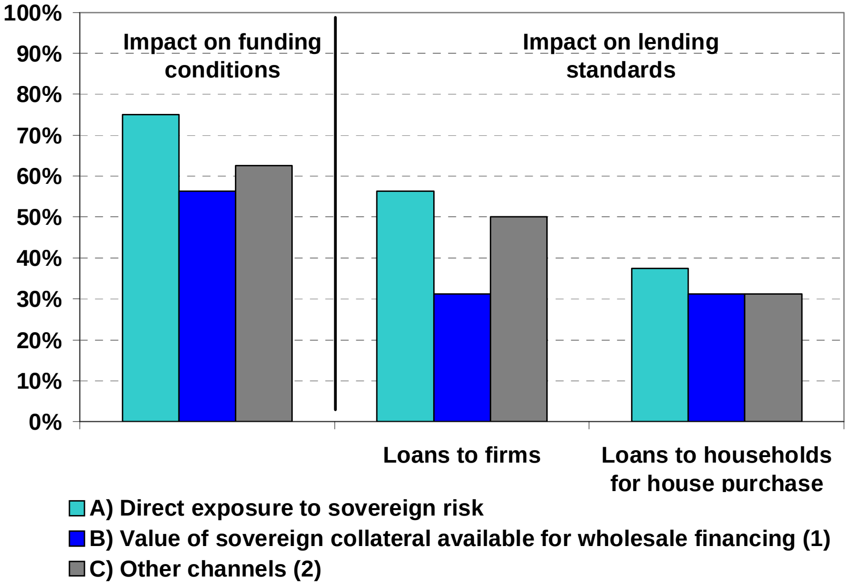
Source: Banca d'Italia.

(1) Positive values indicate supply restriction compared with the previous quarter. Diffusion indices are constructed on the basis of the following weighting scheme: 1 = tightened considerably, 0.5 = tightened somewhat, 0 = remained basically unchanged, -0.5 = eased somewhat, -1 = eased considerably. (2) Refers to the quarter ending at the time of the survey.

Figure 4

Transmission channels of the sovereign debt crisis on banks' funding conditions and lending standards, in 2011Q4

(*(+)/(–) = channel contributed to worsening/improving*)



Source: Banca d'Italia, Euro Area Bank Lending Survey (Italian panel).

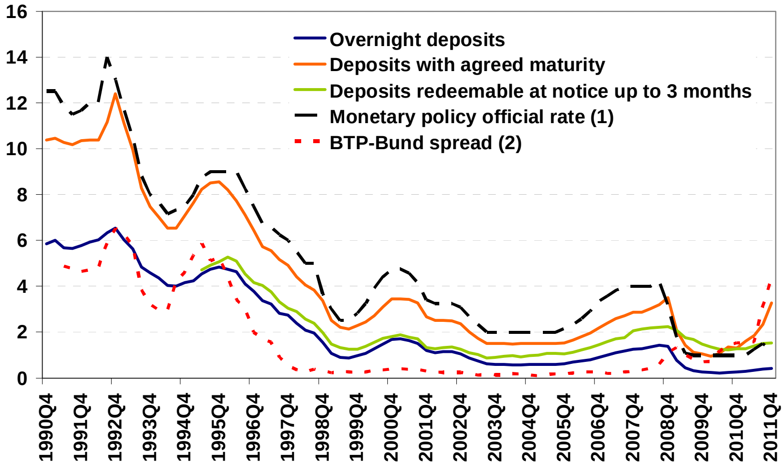
(1) For example, repos or secured transactions in derivatives. (2) For instance, any automatic rating downgrade affecting the bank following a sovereign downgrade or changes in the value of the domestic government's implicit guarantee, as well as spillover effects on other assets, including the loan book.

Figure 5

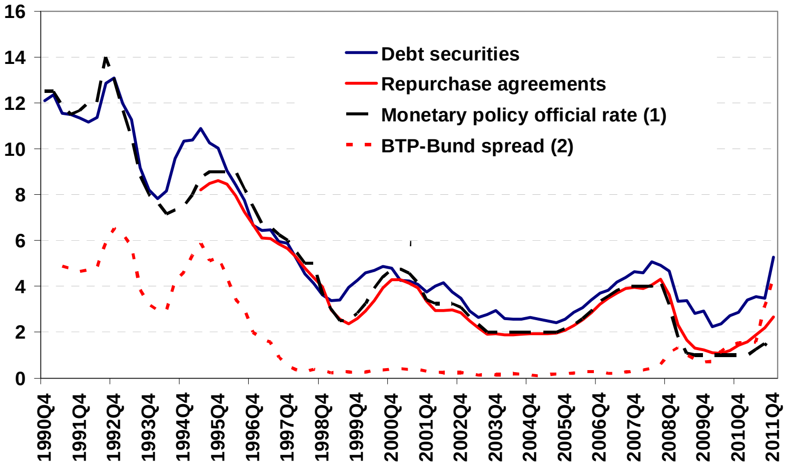
Banks' cost of funding: selected technical forms

(*percentage points; quarterly frequency*)

(a) Deposits held by households



(b) Other items



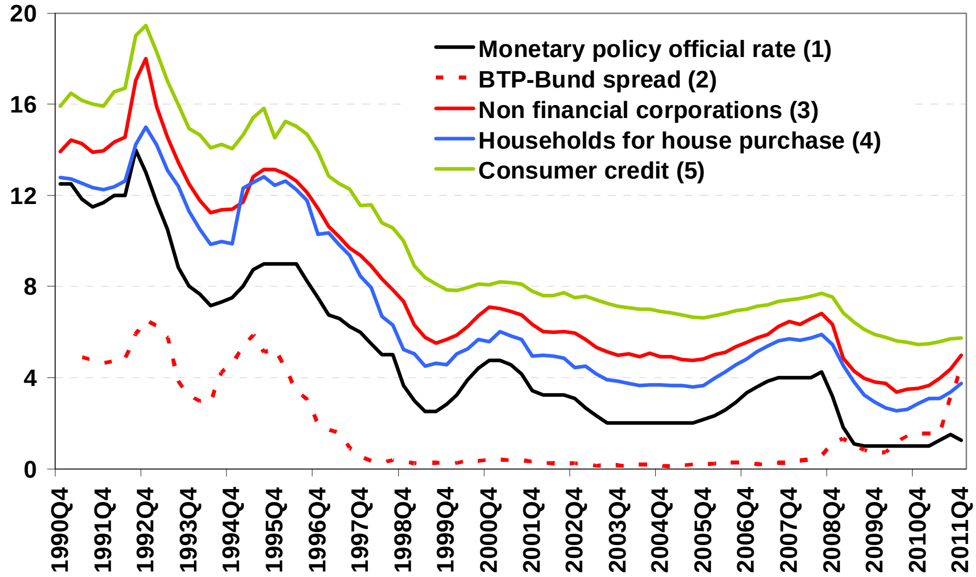
Source: Banca d'Italia.

(1) Until 1999Q4 official discount rate (TUS) set by Banca d'Italia; minimum/fixed bid rate on Eurosystem MROs thereafter. (2) Spread on 10-year government bonds with respect to the corresponding German Bund, corrected by the difference in the 10-year swap rate in Italy and Germany for the pre-EMU period.

Figure 6

Interest rates on loans

(*percentage points; quarterly frequency*)



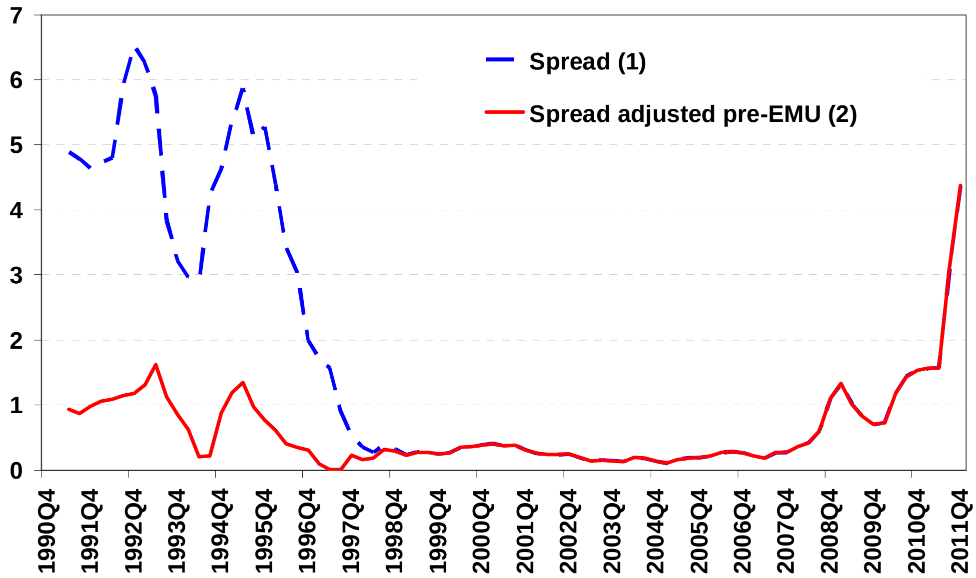
Source: Banca d'Italia

(1) Until 1999Q4 official discount rate (TUS) set by Banca d'Italia; minimum/fixed bid rate on Eurosystem MROs thereafter. (2) Spread on 10-year government bonds with respect to the corresponding German Bund, corrected by the difference in 10-year swap rate in Italy and Germany. (3) Average interest rate on outstanding loans in euro with maturity up to 1 year, including overdrafts. (4) Average interest rate on new loans to households for house purchases in euro, excluding overdrafts. (5) Average interest rate on outstanding loans in euro.

Figure 7

Italian 10-year sovereign spread

(*quarterly averages*; *percentage points*)

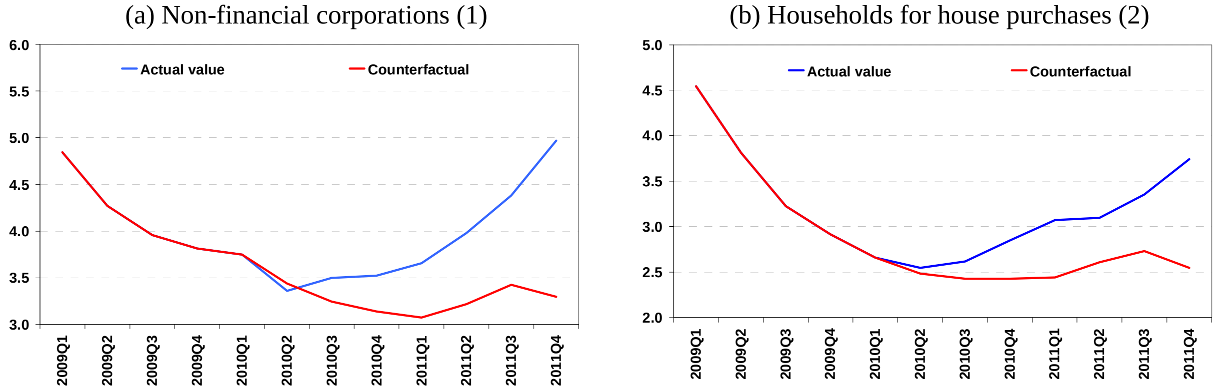


(1) Spread on 10-year government bonds with respect to the corresponding German Bund. (2) Spread corrected, for the pre-EMU period, with the difference between the Italian 10-year swap rate and the corresponding German rate for the pre-EMU period.

Figure 8

Counterfactual exercise: interest rates on loans

(*percentage points*)

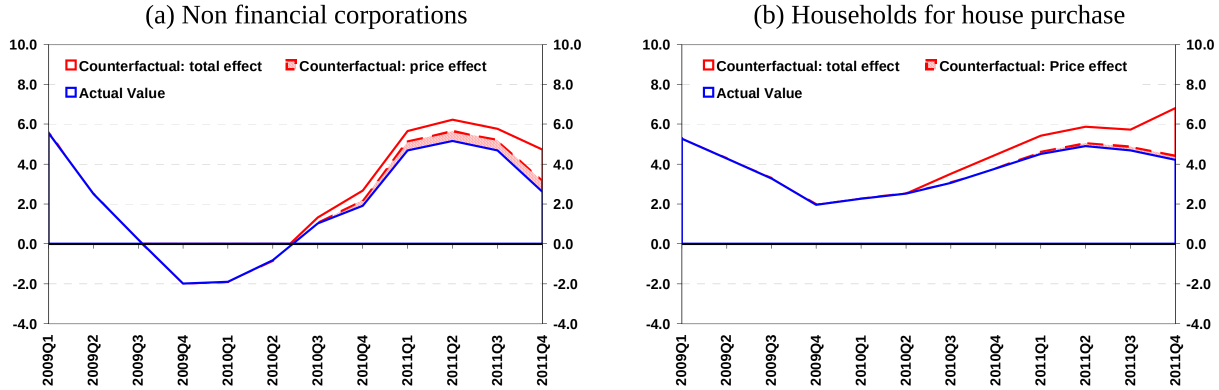


Source: Authors' calculations.
(1) Average interest rate on outstanding loans in euro with maturity up to 1 year, including overdrafts. (2) Average interest rate on new loans to households for house purchases in euro, excluding overdrafts.

Figure 9

Counterfactual exercise: lending growth (1)

(*12-month percentage changes*)



Source: Authors' calculations.
(1) The counterfactual series are obtained by adding back to the actual values of lending growth the contribution of the difference between the actual and the counterfactual BTP-Bund spread. The direct effect is calculated using the coefficient for the spread in the loan equation; the indirect or price effect is the contribution occurring via the effect of the spread on the loan rate and is thus calculated using the product between the coefficient of the spread in the equation for the loan rate and the coefficient of the loan rate in the equation for lending. This methodology does not take into account the autoregressive structure of the loan rate and credit growth; if these were included, the difference between actual lending growth and growth in the counterfactual exercise would be larger.

Table 1

Results for funding rates

Explanatory variables	Households' deposits: overnight (i)	overnight (ii)	with agreed maturity (iii)	with agreed maturity (iv)	Repurchase agreements (v)	Repurchase agreements (vi)	Bank bonds (vii)	Bank bonds (viii)
Lagged dependent (t-1)	0.76 ***	0.64 ***	0.56 ***	0.41 ***	0.48 ***	0.28 **	0.34 *	0.48 ***
Short-term interest rate (t)	0.29 ***	0.24 ***	0.52 ***	0.42 ***	0.62 ***	0.58 ***	0.29 **	0.26 **
Short-term interest rate (t-1)	-0.17 ***	-0.10 *	-0.19 **	-0.01	-0.21	0.02	-0.12	-0.16 **
Medium-term interest rate (t)							0.56 ***	0.51 ***
Medium-term interest rate (t-1)							-0.19	-0.13
BTP-Bund spread (t)	0.04 ***	0.02	0.34 ***	0.17 **				
BTP-Bund spread (t) * Dummy sov_crisis		-0.02		0.21 **				
BTP-Bund spread (t) * Dummy pre_EMU		0.15		0.20				
BTP-Bund spread (t-1)					0.21 ***	0.14 *	0.71 ***	0.58 ***
BTP-Bund spread (t-1) * Dummy sov_crisis						0.21 ***		0.44 *
BTP-Bund spread (t-1) * Dummy pre_EMU						-0.36 **		-0.70 ***
Dummy sov_crisis		0.08 **		-0.02		-0.06		-0.48
Dummy pre_EMU		0.17 **		0.26 *		-0.01		0.22
Adjusted R-squared	0.996	0.996	0.996	0.996	0.996	0.997	0.987	0.993
Sample (adjusted)	1993Q3 - 2011Q4		1993Q3 - 2011Q4		1995Q3 - 2011Q4		1993Q3 - 2011Q4	
	BTP-Bund spread pass-through							
After 1 year (during sovereign crisis)	0.14	0.00	0.73	0.63	0.39	0.49	1.05	1.85
Long-run	0.18		0.77		0.42		1.07	

Note. All regressions include a constant term and time dummies for outliers. *Short-term interest* rate is the 3-month interbank rate. *Medium-term interest rate* is the 3-year swap rate. The BTP-Bund spread is calculated at the 10-year maturity and is corrected by the difference between the 10-year swap rate in Italy and Germany. The 1-year BTP-Bund spread pass-through is calculated assuming *Dummy sov_crisis* = 1.

Table 2

Results for loan rates

Explanatory variables	Loans to firms (i)	Loans to firms (ii)	Loans to households for house purchase (iii)	Loans to households for house purchase (iv)	Consumer credit and other loans to households (v)	Consumer credit and other loans to households (vi)
Lagged dependent (t-1)	0.67 ***	0.52 ***	0.85 ***	0.81 ***	0.92 ***	0.93 ***
Short-term interest rate (t)	0.78 ***	0.74 ***	0.57 ***	0.56 ***	0.71 ***	0.69 ***
Short-term interest rate (t-1)	-0.43 **	-0.27	-0.49 ***	-0.46 ***	-0.62 ***	-0.64 ***
Long-term interest rate (t)			0.05	0.03		
Unemployment rate (t)	0.03	0.05 *	-0.03	-0.02		
GDP, 12-month growth (t-1)	0.01	-0.02			0.04 *	0.05 **
GDP, 12-month growth (t-2)			0.03 *	0.03		
Disposable income, 12-month growth (t-1)					0.01 ***	0.01 **
BTP-Bund spread (t-1)	0.21 ***	-0.15	0.17 ***	-0.01	0.16 ***	0.36 *
BTP-Bund spread (t-1) * Dummy sov_crisis		0.66 ***		0.29 **		-0.22
BTP-Bund spread (t-1) * Dummy pre_EMU		0.51 ***		0.30 *		-0.04
Dummy sov_crisis		-0.78 ***		-0.33 ***		-0.02
Dummy pre_EMU		0.18		0.17		0.01
Adjusted R-squared	0.994	0.995	0.995	0.995	0.997	0.997
Sample (adjusted)	1991Q3-2011Q4		1991Q3-2011Q4		1991Q3-2011Q4	
	BTP-Bund spread pass-through					
After 1 year (during sovereign crisis)	0.51	0.97	0.53	0.83	0.57	0.50
Long-run	0.64		1.12		2.00	

Note. All regressions include a constant term and time dummies for outliers. *Short-term interest rate* is the official monetary policy interest rate (the official discount rate of the Bank of Italy until 1998Q4 and the minimum interest rate on ECB main refinancing operations since 1999Q1). *Long-term interest rate* is the 10-year swap rate. The BTP-Bund spread is calculated at the 10-year maturity and is corrected by the difference between the 10-year swap rate in Italy and Germany. The 1-year BTP-Bund spread pass-through is calculated assuming *Dummy sov_crisis* = 1.

Table 3

Results for lending growth

Explanatory variables	Loans to firms	Loans to households for house purchase
	(i)	(ii)
Lagged dependent (t-1)	0.821 ***	0.842 ***
Loan rate - short-term interest rate spread (t-1)	-0.010 ***	
Short-term interest rate (t-1)	0.000	
Loan rate - short-term interest rate spread (t-2)		-0.008 *
Short-term interest rate (t-2)		-0.005 ***
Firms' financing needs, 3-month change (t-3)	0.077 **	
GDP, 12-month growth (t)	0.002 **	
GDP, 12-month growth (t-4)		0.003 **
House prices, 12-month growth (t-8)		0.002 **
Unemployment rate (t)		0.011 ***
BTP-Bund spread (t-1)	-0.007 ***	-0.010 **
Adjusted R-squared	0.893	0.942
Sample (adjusted)	1991Q3-2011Q4	1991Q3-2011Q4

Note. All regressions include a constant term and time dummies for outliers. *Short-term interest rate* is the e-MID interbank interest rate. The BTP-Bund spread is calculated at the 10-year maturity and is corrected by the difference between the 10-year swap rate in Italy and Germany. *Firms' financing needs* is the ratio between the corporate sector's investments and gross operating profit.

Table 4

Results for banks' income statements

Explanatory variables	Net interest income		Non-interest income and other revenues		Loan-loss provisions	
	All banks	5 largest groups	All banks	5 largest groups	All banks	5 largest groups
	(i)	(ii)	(iii)	(iv)	(v)	(vi)
Lagged dependent (t-1)	-0.08	0.02	-0.11	0.00	0.29 ***	0.26 **
Lagged dependent (t-2)	0.42 ***	0.37 ***				
Short-term interest rate (t-1)	0.03 ***	0.03 ***	-0.06 **	-0.05 *		
Long-term interest rate, 3-month change (t)	0.03 ***	0.01	-0.24 ***	-0.23 ***	0.00	0.00
GDP (t-1)	0.68 ***	0.36 ***	0.51	0.18		
GDP, 3-month growth (t-1)					-0.09 **	-0.08 **
Unemployment rate (t-1)					0.03	0.03
Stock price index (t-1)			0.55 ***	0.61 ***		
Stock market volatility/100					0.06 *	0.08 **
BTP-Bund spread (t-1)	0.02 *	-0.04 ***	0.02	-0.02 ***	0.26 ***	0.25 ***
Adjusted R-squared	0.838	0.682	0.870	0.829	0.616	0.584
Sample (adjusted)	1991Q3-2011Q4		1991Q3-2011Q4		1991Q3-2011Q4	

Note. All regressions include a constant term, seasonal dummies and time dummies for outliers. *Short-term interest rate* is the e-MID interbank interest rate. *Long-term interest rate* is the yield on 10-year Italian government bonds. The BTP-Bund spread is calculated at the 10-year maturity and is corrected by the difference between the 10-year swap rate in Italy and Germany. *Stock price* refers to the main Italian stock exchange index. *Stock market volatility* is calculated as the implicit standard deviation of the options on the *stock price*.

APPENDIX

Table A1

Results for funding rates (monthly data)

Explanatory variables	households' deposits: overnight (i)	overnight (ii) *memo: quarterly data*	with agreed maturity (iii)	with agreed maturity (iv) *memo: quarterly data*	Yield on banks' bond (v)	Yield on banks' bond (vi) *memo: quarterly data*
Lagged dependent (t-1)	0.78 ***	*0.64 ****	0.69 ***	*0.41 ****	0.48 ***	*0.48 ****
Short-term interest rate (t)	0.13 ***	*0.24 ****	0.21 ***	*0.42 ****	0.08	*0.26 ***
Short-term interest rate (t-1)	-0.05	*-0.10 **	0.00	*-0.01*	0.02	*-0.16 ***
Medium-term interest rate (t)					0.54 ***	*0.51 ****
Medium-term interest rate (t-1)					-0.19 **	*-0.13*
BTP-Bund spread (t)	0.00	*0.02*	0.09 *	*0.17 ***		
BTP-Bund spread (t) * Dummy sov_crisis	-0.01	*-0.02*	0.11 **	*0.21 ***		
BTP-Bund spread (t) * Dummy pre_EMU	0.07 *	*0.15*	0.11 *	*0.20*		
BTP-Bund spread (t-1)					0.55 ***	*0.58 ****
BTP-Bund spread (t-1) * Dummy sov_crisis					0.25	*0.44 **
BTP-Bund spread (t-1) * Dummy pre_EMU					-0.25 *	*-0.70 ****
Dummy sov_crisis	0.04 ***	*0.08 ***	0.00	*-0.02*	-0.57 *	*-0.48*
Dummy pre_EMU	0.15 ***	*0.17 ***	0.16 ***	*0.26 **	0.19 *	*0.22*
Adjusted R-squared	0.997	*0.996*	0.998	*0.996*	0.988	*0.993*
Sample (adjusted)	1993M4-2011M12		1993M4-2011M12		1993M4-2011M12	
	BTP-Bund spread pass-through					
After 1 year (during sovereign crisis)	0.00	*0.00*	0.63	*0.63*	1.53	*1.85*

Note. All regressions include a constant term and time dummies for outliers. *Short-term interest* rate is the 3-month interbank rate. *Medium-term interest rate* is the 3-year swap rate. The BTP-Bund spread is calculated at the 10-year maturity and is corrected by the difference between the 10-year swap rate in Italy and Germany. The 1-year BTP-Bund spread pass-through is calculated assuming *Dummy sov_crisis* = 1.

Table A2

Results for loan rates: alternative measures of sovereign tensions (quarterly data)

Alternative measures of sovereign risk (x)	Interest rate on loans to firms: Regression without interactions (column *i* of Table 2) x(t-1) (i)	Regression with interactions (column *ii* of Table 2) x(t-1) (ii)	Regression with interactions (column *ii* of Table 2) x(t-1) * dummy sov_crisis (iiii)	Interest rate on loans to households for house purchase: Regression without interactions (column *iii* of Table 2) x(t-1) (iv)	Regression with interactions (column *iv* of Table 2) x(t-1) (v)	Regression with interactions (column *iv* of Table 2) x(t-1) * dummy sov_crisis (vi)
10-year BTP-Bund spread (*Baseline*)	**0.21 *****	**-0.15**	**0.66 *****	**0.17 *****	**-0.01**	**0.29 ****
5-year BTP-Bund spread	0.21 ***	-0.14	0.57 ***	0.19 ***	0.16 *	0.09
3-year BTP-Bund spread	0.13	-0.01	0.41 **	0.09	0.03	0.22 *
10-year BTP yield	**0.08 ***	**-0.06**	**0.67 *****	**0.08**	**-0.09**	**0.34 ***
5-year BTP yield	0.15 ***	0.01	0.46 ***	0.11 *	-0.03	0.23 *
3-year BTP yield	0.14 ***	-0.04	0.45 ***	0.08	-0.08	0.28 ***
10-year BTP-IRS spread	**0.22 *****	**0.09**	**0.51 ****	**0.16 *****	**-0.28**	**0.64 *****
5-year BTP-IRS spread	0.28 ***	0.05	0.48 **	0.20 **	0.03	0.28
3-year BTP-IRS spread	0.15	-0.01	0.46 **	0.08	-0.04	0.31 **
10-year BTP-OAT spread	**0.22 *****	**-0.16**	**0.76 *****	**0.18 ****	**-0.13**	**0.48 ***
5-year BTP-OAT spread	0.26 ***	-0.08	0.61 ***	0.20 ***	0.15	0.16
3-year BTP-OAT spread	0.14	-0.06	0.49 **	0.08	0.02	0.25
Not adjusted 10-year BTP-Bund spread	**0.11 *****	**-0.17**	**0.68 *****	**0.12 *****	**-0.10**	**0.39 ****
Not adjusted 5-year BTP-Bund spread	0.13 ***	-0.15	0.58 ***	0.15 ***	0.15	0.11
Not adjusted 3-year BTP-Bund spread	0.10 ***	-0.10	0.50 ***	0.13 ***	-0.09	0.36 **

Note. The coefficients reported in the Table are obtained by re-estimating the equations in columns (*i*)-(*iv*) in Table 2 and substituting the 10-year BTP-Bund spread by the selected variables (all the remaining explanatory variables are kept unchanged). For the sake of comparison, the first row reports the coefficients for the baseline regression. *IRS* is the interest rate swap. *OAT* is the French sovereign bond.

Table A3

Results for loans rates to firms (monthly data)

Explanatory variables	Without interaction				With interaction			
	memo: quarterly data	All new loans	Loans up to €1 mil	Loans over €1 mil	*memo: quarterly data*	All new loans	Loans up to €1 mil	Loans o €1 mi
	(i)	(ii)	(iii)	(iv)	*(v)*	(vi)	(vii)	(viii)
Lagged dependent (t-1)	*0.67* ***	0.81 ***	0.74 ***	0.70 ***	*0.52* ***	0.70 ***	0.65 ***	0.64
Short-term rate (t)	*0.78* ***	0.54 **	0.63 ***	0.59 *	*0.74* ***	0.37 *	0.48 ***	0.40
Short-term rate (t-1)	*-0.43* **	-0.37 *	-0.36 **	-0.30	*-0.27*	-0.11	-0.15 *	-0.09
Unemployment rate (t)	*0.03*	-0.01	0.03	-0.02	*0.05* *	-0.05	0.00	-0.07
GDP or Industrial production, 12-month growth (t-1)	*0.01*	0.01 **	0.00	0.01 **	*-0.02*	0.00	0.00	0.00
BTP-Bund spread (t-1)	*0.21* ***	0.09 ***	0.11 ***	0.12 ***	*-0.15*	-0.23 ***	-0.13 **	-0.19
BTP-Bund spread (t-1) * Dummy sov_crisis					*0.66* ***	0.39 ***	0.30 ***	0.37
BTP-Bund spread (t-1) * Dummy pre_EMU					*0.51* ***			
Dummy sov_crisis					*-0.79* ***	-0.18 **	-0.16 **	-0.18
Dummy pre_EMU					*0.18*			
Adjusted R-squared	*0.994*	0.977	0.989	0.969	*0.995*	0.981	0.990	0.972
Sample (adjusted)	*1991Q3-2011Q4*	2003M2-2011M12			*1991Q3-2011Q4*	2003M2-2011M12		
	BTP-Bund spread pass-through							
After 1 year (during sovereign crisis)	0.51	0.44	0.42	0.39	0.97	0.40	0.41	0.41
Long-run	0.64	0.48	0.44	0.40				

Note. All regressions include a constant term and time dummies for outliers. *Short-term interest rate* is the official monetary policy interest rate (the official discount rate of the Bank of Italy until 1998Q4 and the minimum interest rate on ECB main refinancing operations since 1999Q1). As a measure of economic activity, *GDP* is used for regressions with quarterly data, *industrial production* for regressions with monthly data. The BTP-Bund spread is calculated at the 10-year maturity and is corrected by the difference between the 10-year swap rate in Italy and Germany. The 1-year BTP-Bund spread pass-through is calculated assuming *Dummy sov_crisis* = 1.

Table A4

Results for loan rates to households for house purchases (monthly data)

Explanatory variables	Without interaction					With interaction				
	memo: quarterly data	All new loans	Variable-rate loans	Fixed-rate loans	Variable-rate loans (extended sample)	*memo: quarterly data*	All new loans	Variable-rate loans	Fixed-rate loans	Variabl rate loa (extend sampl
	(i)	(ii)	(iii)	(iv)	(v)	*(vi)*	(vii)	(viii)	(ix)	(x)
Lagged dependent (t-1)	*0.85* ***	0.79 ***	0.80 ***	0.75 ***	0.89 ***	*0.81* ***	0.85 ***	0.73 ***	0.72 ***	0.89
Short-term interest rate (t)	*0.57* ***	0.37 ***	0.51 ***	0.16 ***	0.34 ***	*0.56* ***	0.22 ***	0.35 ***	0.21 ***	0.34
Short-term interest rate (t-1)	*-0.49* ***	-0.19 **	-0.31 ***	-0.16 ***	-0.23 ***	*-0.46* ***	-0.12 *	-0.11 **	-0.20 ***	-0.23
Long-term interest rate (t)	*0.05*			0.24 ***		*0.03*			0.25 ***	
Unemployment rate (t)	*-0.03*	-0.05 ***	-0.02	-0.07 ***	-0.02 **	*-0.02*	-0.08 ***	-0.06 ***	-0.06 **	-0.02
GDP 12-month growth (t-1) (1)	*0.03* *	0.002 **	0.005 ***	-0.002 **	0.005 ***	*0.03* *	0.001	0.001	-0.002	0.004
Sovereign spread (t-1)	*0.17* ***	0.07 ***	0.05 ***	0.10 ***	0.05 ***	*-0.01*	-0.11 ***	-0.15 ***	0.16 ***	0.02
Spread BTP-Bund (-1) * Dummy sov_crisis						*0.29* **	0.22 ***	0.26 ***	-0.07	0.07
Spread BTP-Bund (-1) * Dummy pre_EMU						*0.32* *				
Dummy sov_crisis						*-0.33* ***	-0.14 ***	-0.18 ***	0.06	-0.07
Dummy pre_EMU						*0.25*				
Adjusted R-squared	*0.995*	0.994	0.994	0.969	0.996	*0.995*	0.996	0.996	0.970	0.996
Sample (adjusted)	*1991Q3-2011Q4*	2003M2-2011M12			1995M2-2011M12	*1991Q3-2011Q4*	2003M2-2011M12			1995M 2011M
	BTP-Bund spread pass-through									
After 1 year (during sovereign crisis)	*0.53*	0.30	0.23	0.39	0.35	*0.83*	0.63	0.39	0.32	0.62
Long-run	*1.12*	0.32	0.25	0.40	0.47	*1.44*	0.74	0.40	0.32	0.82

Note. All regressions include a constant term and time dummies for outliers. *Short-term interest rate* is the official monetary policy interest rate (the official discount rate of the Bank of Italy until 1998Q4 and the minimum interest rate on ECB main refinancing operations since 1999Q1). *Long-term interest rate* is the 10-year swap rate. As a measure of economic activity, *GDP* is used for regressions with quarterly data, *industrial production* for regressions with monthly data. The BTP-Bund spread is calculated at the 10-year maturity and is corrected by the difference between the 10-year swap rate in Italy and Germany. The 1-year BTP-Bund spread pass-through is calculated assuming *Dummy sov_crisis* = 1.

Table A5

Results for loan rates with banks' marginal cost of funding (quarterly data)

Explanatory variables	Interest rate on loans to firms				Interest rate on loans to households for house purchase			
	Without interactions		With interactions		Without interactions		With interactions	
	Baseline	With cost of funding	*Baseline*	With cost of funding	*Baseline*	With cost of funding	*Baseline*	With cost of funding
	(i)	(ii)	*(iii)*	(iv)	*(v)*	(vi)	*(vii)*	(viii)
Lagged dependent (t-1)	*0.68* ***	0.56 ***	*0.52* ***	0.42 *	*0.85* ***	0.81 ***	*0.81* ***	0.72 ***
Short-term interest rate (t)	*0.78* ***	0.55 ***	*0.74* ***	0.50 ***	*0.57* ***	0.53 ***	*0.56* ***	0.54 ***
Short-term interest rate (t-1)	*-0.43* **	-0.40 **	*-0.27*	-0.27	*-0.49* ***	-0.48 ***	*-0.46* ***	-0.42 ***
Long-term interest rate (t)					*0.05*	0.14 **	*0.03*	0.09 *
Unemployment rate (t)	*0.03*	0.03	*0.05* *	0.06 *	*-0.03*	-0.01	*-0.02*	0.00
GDP, 12-month growth (t-1)	*0.01*	0.00	*-0.02*	-0.02				
GDP, 12-month growth (t-2)					*0.03* *	0.03 *	*0.03*	0.04 *
BTP-Bund spread (t-1)	*0.21* ***	-0.10	*-0.15*	-0.42 **	*0.17* ***	0.19	*-0.01*	0.10
BTP-Bund spread (t-1) * Dummy sov_crisis			*0.66* ***	0.60 ***			*0.29* **	0.25
BTP-Bund spread (t-1) * Dummy pre_EMU			*0.51* ***	0.52 ***			*0.30* *	0.41 *
Dummy sov_crisis			*-0.79* ***	-0.81 ***			*-0.33* ***	-0.43 **
Dummy pre_EMU			*0.18*	0.06			*0.17*	0.42
Marginal cost of funding (t)		0.38 **		0.40 **		0.01		-0.04
Adjusted R-squared	*0.994*	0.995	*0.995*	0.995	*0.995*	0.994	*0.995*	0.991
Sample (adjusted)	1991Q3-2011Q4		1991Q3-2011Q4		1991Q3-2011Q4		1991Q3-2011Q4	

Note. All regressions include a constant term and time dummies for outliers. *Short-term interest rate* is the official monetary policy interest rate (the official discount rate of the Bank of Italy until 1998Q4 and the minimum interest rate on ECB main refinancing operations since 1999Q1). *Long-term interest rate* is the 10-year swap rate. The BTP-Bund spread is calculated at the 10-year maturity and is corrected by the difference between the 10-year swap rate in Italy and Germany. *Marginal cost of funding* is the interest rate on households' term deposits. The BTP-Bund spread is calculated at the 10-year maturity and is corrected by the difference between the 10-year swap rate in Italy and Germany.

The impact of the sovereign debt crisis on bank lending rates in the euro area

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Abstract

Since the early part of 2010 tensions in the sovereign debt markets of some euro-area countries have progressively distorted monetary and credit conditions, hindering the ECB monetary policy transmission mechanism and raising the cost of loans to non-financial corporations and households. This paper makes an empirical assessment of the impact of the tensions on bank lending rates in the main euro-area countries, concluding that they have had a significant impact on the cost of credit in the peripheral countries. A counterfactual exercise indicates that if the spreads had remained constant at the average levels recorded in April 2010, the interest rates on new loans to non-financial corporations and on residential mortgage loans to households in the peripheral countries would have been, on average, lower by 130 and 60 basis points, respectively, at the end of 2011. These results are robust to alternative measures of the cost of credit and econometric techniques.

JEL Classification: C32, E43, G21.

Keywords: sovereign debt crisis, bank lending rates, seemingly unrelated regression.

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1. Introduction

Since early 2010 tensions in the sovereign debt markets of several euro-area countries have progressively distorted monetary and credit conditions, raising the cost of credit for non-financial corporations and of residential mortgage loans to households. The financial strains spread rapidly from Greece, whose budget situation had been concealed by the official statistics, to Ireland, which suffered the consequences of a profound banking crisis, and to Portugal, penalized by its balance-of-trade deficit. In the summer of 2011, tensions in the financial markets hit Italian and Spanish government bonds and assumed systemic proportions. Banks' heightened problems with wholesale funding worsened credit supply conditions (see the Eurosystem Bank Lending Survey for the fourth quarter of 2011).[1]

The sovereign debt crisis has increased the heterogeneity of financial conditions within the euro area. Given the predominant role of banks in financing the private sector, these cross-border differences have posed a serious challenge to the ECB's conduct of monetary policy, making the transmission of monetary impulses to euro-area countries less uniform.[2]

The debate on the origins of the euro-area debt crisis has brought out several factors: the macroeconomic weaknesses of some member states, which fuelled doubts about the sustainability of their public debt; the incompleteness of the European construction (see, among others, Visco, 2012); and some form of "wake-up contagion" (Giordano *et al*., 2013). This paper presents an econometric analysis of the transmission of the tensions in sovereign debt markets to the cost of new short-term loans to non-financial corporations and mortgage loans to households in a group of euro-area countries (Germany, France, Italy, Spain, Netherlands, Belgium, Austria, Greece, Portugal and Finland).[3] The analysis, based on monthly data for the 2003-2012 period, uses the seemingly unrelated regression approach (SUR), which provides more efficient estimation by taking account of the correlations between the disturbances in the equations. As a robustness check, the estimation is also performed by ordinary least squares and for the period 2008-2012. The paper complements the analysis by Albertazzi *et al*. (2012), which gauges the effects of the sovereign debt crisis on the activity of Italian banks. Neri and Ropele (2013) assess the macroeconomic impact of

[1] The report is available at http://www.ecb.int/stats/pdf/blssurvey_201201.pdf?a28f705bb565d645677ba8ce15ba6049.
[2] The sovereign crisis "[…] has made difficult the transmission of impulses coming from an accommodative monetary policy through adjustments in interest rates on loans to households and firms by banks. Interest rates do not have to be identical across the euro area, but it is unacceptable if significant differences arise because of the fragmentation of capital markets or the perception of a break-up of the euro area. […] The fragmentation of the single financial market has led to a fragmentation of the single monetary policy." See "The monetary policy of the European Central Bank and its transmission in the euro area", speech by Mario Draghi, President of the ECB, Università Bocconi, Milan, 15 November 2012.
[3] Ireland is excluded for lack of continuous series for confidence indexes.

the sovereign debt crisis for a set of euro-area countries using a factor augmented vector autoregressive (FAVAR) model.

The econometric analysis set out here shows that the sovereign debt tensions have had a substantial impact on bank lending rates in the peripheral countries (Italy, Spain, Greece and Portugal), but practically none in the core countries (Germany, France, Belgium, the Netherlands, Austria and Finland). A counterfactual exercise suggests that if sovereign spreads (defined as the difference between the yields on 10-year government bonds and the yield on swap contracts with the same maturity) had remained constant at their average levels of April 2010, at the end of 2011 the interest rates on new loans to non-financial corporations and residential mortgages in the peripheral countries would have averaged respectively 130 and 60 basis points lower than their actual values. For non-financial corporations in particular, rates would have been 180, 140, 130 and 50 basis points lower in Greece, Portugal, Italy and Spain respectively.

The remainder of the paper is organized as follows. Section 2 describes developments in sovereign spreads and bank rates in the course of the debt crisis, Section 3 studies the impact on bank rates, Section 4 presents some robustness checks and Section 5 concludes.

2. The sovereign crisis and bank lending rates in euro-area countries

In view of the importance of banks in financing the private sector, the pass-through of changes in ECB policy rates to bank lending rates is a key element of the euro-area monetary policy transmission mechanism (see ECB, 2008b and Draghi, 2012). Normally the marginal cost of funding for banks is related to the interest rates on unsecured borrowing in the interbank market, which in turn are related to policy rates. By means of its supply of reserves through refinancing operations, the ECB can influence short-term money market rates (such as EONIA, the average overnight rate) which in turn affect lending rates and ultimately the real economy and inflation.

2.1 The evolution of the sovereign debt crisis

The financial crisis has caused severe fragmentation of the European financial market along national lines, hampering the transmission of ECB monetary policy and preventing the uniform transmission of monetary impulses to the different euro-area countries. The Governing Council of the ECB responded to such malfunctions by adopting, after October 2008, a series of extraordinary measures (see Trichet, 2009, “The ECB’s enhanced credit support”) to sustain credit to the private sector and maintain the correct transmission of monetary policy (see Cecioni *et al*. 2011). These extraordinary measures helped to improve financial conditions in 2009.

Early in 2010, however, the government securities markets of some euro-area countries with weak fiscal and macroeconomic fundamentals were strained severely (Fig. 1). The sovereign spreads of several countries rose to unprecedented levels, and in some government bond markets liquidity nearly evaporated. These tensions spread to other market segments, including equity and money markets (see Trichet, 2010). The ECB responded by reinstating some measures that had been withdrawn in the preceding months and instituting the Securities Markets Programme (SMP) to ensure the proper transmission of monetary policy impulses to the economy through secondary-market purchases of government securities.

The strains spread from Greece, where they had been triggered by fears for the sustainability of the public finances, to the Irish and Portuguese government bond markets in the autumn and winter of 2010-2011. In July 2011, the spreads of Italian and Spanish government bonds vis-à-vis the German Bund increased significantly, following further downgrades of Greek bonds and the announcement by the European Council of the involvement of the private sector in the second bail-out package for Greece. In order to ease the tensions in financial markets and halt the contagion, the ECB revived the Securities Markets Programme. Banks' funding conditions worsened significantly, affecting their lending standards, particularly on loans to firms. The extraordinary measures taken in December by the ECB, including two three-year refinancing operations, and the lowering of policy rates, together with the fiscal adjustment measures of some governments, helped to allay the tensions and alleviate banks' funding difficulties.

Financial market tensions resurfaced in March 2012 as investor worries heightened over the political situation in Greece, the difficulties of the Spanish banking sector and the inability of governments to reform European Union governance arrangements or create effective crisis management tools for the euro area. The sovereign spreads of the peripheral countries increased steadily until August, when the Governing Council of the ECB announced extraordinary measures to address the bond market disruptions, which stemmed in part from concern that the euro might prove reversible ("redenomination risk"), and to preserve the singleness of the ECB monetary policy and the correct functioning of its transmission mechanism.[4] Di Cesare *et al.*, 2012 and Giordano *et al.*, 2013, have shown that in 2012 the sovereign spreads of several countries had reached levels that were higher than those consistent with fiscal and macroeconomic fundamentals, partly because of the perceived risk of a break-up of the euro area. After its meeting on 6 September 2012, the Governing Council described the new Outright Monetary Transactions programme for the

[4] From the introductory statement to the press conference of 2 August 2012: "The Governing Council, within its mandate to maintain price stability over the medium term and in observance of its independence in determining monetary policy, may undertake outright open market operations of a size adequate to reach its objective."

purchase of government securities in secondary markets in detail. The programme, which is fully within the remit of the ECB, helps to maintain price stability.[5]

2.1. Sovereign spreads and bank lending rates

Several channels have been suggested for the transmission of government bond market tensions to the banking system.[6] According to the so-called price channel thesis, the interest rate on government bonds is a benchmark for determining the cost of loans to households and non-financial corporations. The balance-sheet thesis is that a capital loss on banks' government bond portfolios, via its effects on profitability and capital, may induce a tightening of credit supply. And working via the liquidity channel, the loss of value of bonds, which are typically used as collateral for interbank transactions, could reduce banks' ability to procure funds in the money market.[7]

As a gauge of the impact of the sovereign tensions on lending rates, Figures 2 and 3 show the mean cost of new loans to non-financial corporations and of residential mortgages in the euro area and the interval between the 10th and the 90th percentile of the distribution of these rates across countries. The heterogeneity in these costs has increased significantly during the sovereign debt crisis. The standard deviation of interest rates on new loans to non-financial corporations increased from 62 basis points in January 2003-April 2010 to 139 points in May 2010-August 2011 (Fig. 2); for households the increase was more modest, from 45 to 56 basis points (Fig. 3).

The picture as regards real interest rates (net of inflation), the relevant variable for the investment and consumption decisions of firms and households, is similar. As for the most appropriate gauge of inflation, economic theory holds that lenders require compensation for the monetary erosion produced by the expected change in the price level. This suggests that the best indicator may be the ex-ante real interest rate. However, euro-area inflation expectations, such as those of Consensus Economics, are generally available at quarterly frequency only for the four largest countries and for the area as a whole. To compute real inflation rates for more countries, one can use rates based on the Harmonized Index of Consumer Prices (HICP), which have a high correlation with expected inflation rates, and so permit the calculation of an ex-post measure of the real interest rate. Figures 4 and 5 show the mean and median real rates on loans and the 10th and 90th percentile interval of the distribution across countries. The figures confirm that the

[5] Implementation of the OMT is subject to strict conditionality connected with the application to the European Financial Stability Facility (EFSF) or the European Stability Mechanism (ESM) for financial support.
[6] For a description of the relationship between sovereign risk and funding conditions for banks, see Committee on the Global Financial System (2011) and Trichet (2010).
[7] Sovereign downgrades are generally followed by downgrades of domestic banks, while a weakening of the public finances reduces the availability of implicit and explicit government guarantees.

heterogeneity in the real cost of new loans to non-financial corporations increased significantly after the beginning of the sovereign debt crisis, while for households the increase was more moderate.

The increase in the dispersion of lending rates reflects the asymmetrical nature of the sovereign debt crisis. Before the crisis, lending rates were not correlated with sovereign spreads in Greece, Portugal, Spain and Italy (Fig. 6). The correlation emerged only once the crisis had broken out. The correlation is analyzed more closely, through an econometric model, in Section 3.

Despite the increase in the dispersion of bank rates, the heterogeneity in trends in loans to non-financial corporations (Fig. 7) and households (Fig. 8) diminished, albeit to varying extent, in the course of the financial and the sovereign debt crises. This diminution mainly reflected the impact of the financial crisis on banking and on the real economy, which has tended to synchronize credit developments in euro-area countries. Before the crises, these variables were driven by country-specific factors such as buoyant housing markets (Spain and Ireland), or strong corporate cash-flow (Germany).

The standard deviation of the growth rates of credit to non-financial corporations fell from 6.4 percentage points in the period from 2003 to the summer of 2007 to 3.8 points between autumn 2007 and August 2012 (Table 1). By contrast, the sovereign debt crisis did not affect the volatility of credit dynamics significantly: the standard deviation for the period from May 2010 to August 2012 is identical to that between September 2007 and April 2010. As for residential mortgage loans, the decline in volatility between these two periods was comparable (from 7.5 to 4.8 percentage points), but such decline occurred mainly during the sovereign debt crisis (3.4 percentage points between May 2010 and August 2012). As for lending rates, the statistics in Table 1 confirm these results: for non-financial corporations the standard deviation increased from 0.6 percentage points between 2003 and the summer of 2007 to 0.7 during the financial crisis (from September 2007 to April 2010) and to 1.4 during the sovereign debt crisis. For households the increase in volatility was more muted.

To sum up, the financial crisis brought a decline in the standard deviation of credit growth and almost no change in the dispersion of bank lending rates, whereas the sovereign debt crisis increased the dispersion of lending rates, in particular to non-financial corporations, while further lowering the volatility of growth rates of credit to households.

3. Econometric analysis

The indicator used here for tensions in sovereign debt markets is the spread between the yields on government bonds and the 10-year swap rate of equal maturity. Unlike the German Bund yield, the swap rate as benchmark enables us to include an equation for lending rates in Germany and to assess the possible effect on these rates of the decline of the Bund yield due to the "flight to quality" during the worst of the financial and sovereign debt crises. The system describing lending rates to non-financial corporations (*k* = *f*) and households for house purchase (*k* = *h*) in country *j*:

$$R_{j,t}^{k} = \overline{R}_{j}^{k} + \alpha_1 D_t^{crisis} + \alpha_2 D_t^{2008} + \alpha_3 R_{t-1}^{ov} + \alpha_4 \left(R_{t-1}^{3m} - R_{t-1}^{ov} \right) + \alpha_5 \left(R_{j,t-1}^{10} - R_{t-1}^{10,swap} \right) + \alpha_6 Y_{t-1}^{k} + \alpha_7 R_{j,t-1}^{k} + \varepsilon_t \qquad (1)$$

is estimated by the SUR method.[8] Each equation in (1) is set up on the autoregressive distributed lags model (ARDL), which has been used extensively to assess the transmission of changes in monetary policy rates to lending rates (Cottarelli and Kourelis, 1994; Favero *et al*., 1997; Marotta, 2010).[9] The explanatory variables are: the constant $\overline{R}_{j}^{k}$, a dummy D^{crisis} that takes the value one from the Lehman Brothers default to the end of the sample period, a dummy D^{2008} taking value one between November 2008 and February 2009 (to avoid large residuals in some of the equations), the EONIA rate, the spread between three-month Euribor and EONIA $R^{3m} - R^{ov}$ (a proxy for credit risk in the money market), the spread between the yields on 10-year government bonds and that of swap contracts in euro of equal maturity, $R_j^{10} - R^{10,swap}$, a confidence indicator Y^k, and the lagged value of the bank rate.[10] The independent variables are introduced at time *t*-1 in order to avoid endogeneity problems. The sovereign spread can be thought of as a summary statistic for the impact of banks' funding difficulties on lending standards.[11] This thesis finds support in Del Giovane *et al*. (2013), according to which developments in banks' balance sheet and funding conditions mainly reflect those in the sovereign debt markets.

The interest rates are the rates on new loans (excluding overdrafts) to non-financial corporations and on residential mortgages, with initial maturity up to one year (see the Appendix for a detailed description). These data are available from January 2003 onwards on the ECB website.[12] The monetary financial institutions interest rate statistics cover all the interest rates that credit institutions apply to euro-denominated deposits and loans (outstanding amounts and new business)

[8] The rationale for the seemingly unrelated regression approach is that there could be common factors not included as independent variables affecting all the equations at once, inducing correlation between the error terms.

[9] Depending on the nature of the dataset available, other empirical approaches are also possible. For example, panel-econometric methods have been used (Sørensen and Werner, 2006) to study heterogeneity in the pass-through of changes in money market rates to lending rates in the euro area.

[10] Since mid-2011 Eurepo rates have been affected by the strong demand for high-quality collateral, which has pushed them down close to or even below zero. Data on Eonia swap rate contracts are available only after mid-2005. The use of the spread between Euribor and Eurepo and between Euribor and Eonia swaps results in only marginally different coefficients; and the coefficients on sovereign spreads and counterfactual simulations are identical (Section 3.1).

[11] See the ad hoc questions on the impact of the sovereign debt crisis on banks' funding conditions and credit standards in the Eurosystem Bank Lending Survey (http://www.ecb.int/stats/money/surveys/lend/html/index.en.html).

[12] See, for instance, http://www.ecb.int/stats/pdf/mir_general_description.pdf?a13000d3717c100bef08c0c137f28bcd.

to all households and resident non-financial corporations of all sizes.[13] These data have been used, among others, by Marotta (2009) to study the pass-through of changes in the ECB rate to lending rates. The rates used in the estimation are among the most responsive to changes in banks' funding costs and the policy rate.

Before estimating the model, the stationarity of the time series for lending rates was tested using the Augmented Dicky-Fuller test (Said and Dickey, 1984), which rejected the null of stationarity at the 5 per cent significance level for the interest rates on new loans to both non-financial corporations and households. These results were also confirmed by the Kwiatowski *et al*. (KPSS, 1992) and Phillips-Perron (PP, 1998) tests (see Table 2).[14] When the dependent variables follow unit root processes, the estimation of ARDL models is equivalent to estimating error correction models (Pesaran and Shin, 1999); the results are discussed in Section 4.4.

Tables 3 and 4 show the estimates for the period from January 2003 to August 2012.[15] For non-financial corporations (Table 3): (*i*) the adjustment of lending rates to changes in the explanatory variables is gradual (column h); (*ii*) both EONIA (column b) and the spread between three-month Euribor and EONIA (column e) are significant in all the equations; (*iii*) the transmission of changes in EONIA is incomplete in both the short and the long run; (*iv*) the sovereign spread (column f) is statistically significant in the peripheral but not the core countries. After one quarter, an increase of 100 basis points in these spreads raises lending rates by 33 basis points in Italy, 20 in Portugal, 16 in Spain and 7 in Greece.

For Italy, Albertazzi *et al*. (2012) estimate the effect on rates on new short-term loans (excluding overdrafts) at 35 basis points using monthly data over the period 2003-2012; the impact is greater, 50 basis points, when overdrafts are included, the period considered is 1991-2011, and the data used are quarterly rates on (outstanding) short-term loans.[16] Del Giovane *et al*. (2013) find that an increase of 100 basis points in the sovereign spread raises the interest rate on new short-term loans by around 30 basis points with a one- quarter lag;[17] Zoli (2013) estimates that between 30 and 40 per cent of the increase in the Italian sovereign spread is transmitted to the interest rates in loans

[13] The MFI interest rate statistics provide monthly data on 45 instrument categories of euro-denominated deposits and loans, disaggregated by original maturity, notice periods or periods of interest rate fixation. The statistics are produced for the euro area as a whole and individually for each member country.
[14] KPSS is often used to confirm the results obtained from other tests in which the null hypothesis is the non-stationarity of the time series. However, inferences based on the KPSS test can be sensitive to the number of lags in the estimation of the heteroskedasticity and autocorrelation consistent covariance matrix (Maddala and Kim, 1998). In fact, with more lags the assumption of stationarity cannot be rejected for some countries, including Italy and Spain.
[15] The correlation between the residuals (on average 0.2, in some cases 0.6) justifies the SUR approach.
[16] The difference between the estimates may reflect both the different sample periods and the types of loans used.
[17] Del Giovane *et al*. (2013) analyze the relative role of demand and supply factors in explaining credit developments, focusing on the differences between the global financial crisis and the sovereign debt crisis.

to non-financial corporations within three months. The positive coefficients on D^{crisis} indicate that the spread between the interest rate on loans to non-financial corporations and EONIA was larger after the collapse of Lehman Brothers than before.

For new residential mortgage loans, the estimation yields the following results (Table 4): (*i*) in all countries, the adjustment to changes in the explanatory variables is more gradual than for non-financial corporations (column h); (*ii*) EONIA is significant in all equations (column b); (*iii*) the transmission of changes in EONIA is incomplete both in the short and in the long run; (*iv*) the sovereign spread is statistically and economically significant in Spain, Italy and Portugal and, among the core countries, also in Belgium and Finland (column f).

Section 4 assesses the robustness of these estimations.

3.1. The impact of sovereign spreads on lending rates

Using the system estimated in (1), we can quantify the effect of increases in sovereign spreads on lending rates by means of two counterfactual exercises; namely, we assume that the spreads remain constant at the values recorded in April 2010, before the heightening of the tensions on Greek bonds, or, alternatively, at the level of June 2011, before the tensions hit the markets for Italian and Spanish bonds. All the other variables are assumed to follow their actual paths. By construction, the exercises do not consider that in the absence of the sovereign crisis the ECB would have followed different (conventional and unconventional) policies and the macroeconomic outlook of the euro-area countries would also have been better.[18] Therefore the counterfactual exercise offers a conservative estimate of the impact of the sovereign debt market tensions.

The differences between the actual values of bank rates and those obtained with the two counterfactual exercises are set out in Tables 5 and 6 and Figures 9-12. If sovereign spreads had remained constant at the levels recorded in April 2010, the average interest rates on loans to non-financial corporations in 2011 would have been lower than the actual rates by around 50 basis points in Italy and 40 in Spain; and at the end of the year the rates would have been lower than their actual values by 130 and 50 basis points respectively (Table 5). Albertazzi *et al*. (2012), using quarterly data, estimate that if the sovereign debt crisis had not occurred the interest rate on short-term (outstanding) loans in the last quarter of 2011would have been lower than actual rates by 170 basis points.[19] They would have been about 140 basis points lower in Greece and 180 points lower

[18] Using the actual data for business and household confidence and the Euribor-Eonia spread, we ignore the fact that if our counterfactual scenarios had been realized, these variables would probably have taken a different course.

[19] An update of these results, presented in Banca d'Italia, Financial Stability Report No. 4, November 2012, shows that if the spread had remained at the same level as in the first quarter of 2010, the interest rate on new loans to firms in the second quarter of 2012 would have been lower than the actual value by 160 basis points.

in Portugal. The average rate in the peripheral countries would have been 130 basis points lower, while the average dispersion of rates on new loans to non-financial corporations would have been equal to 1 percentage point, compared with an actual value of 1.39 (Table 1).

For households, the counterfactual exercise suggests that the effects of the tensions on lending rates were somewhat smaller than for non-financial corporations. Had the spreads remained constant at the levels recorded in April 2010, then in December 2011 the lending rates in Italy, Portugal and Spain would have been lower than the actual values by around 120, 110 and 30 basis points respectively. The average rate in the peripheral countries would have been 60 basis points lower, while the average dispersion of bank rates on new loans to households between May 2010 and August 2012 would have not changed much (0.60, compared with 0.56).

Qualitatively similar results both for non-financial corporations and for households are obtained with the counterfactual exercise taking July 2011 as reference level (Table 6). Comparing the two simulations, we see that in Italy a large part of the rise in banks' lending rates occurred after July 2011. In Spain, by contrast, a good part of the increase in lending rates came after the renewed tensions for Spanish bonds in the spring of 2012. In Greece and Portugal, the effects on lending rates are already largely incorporated in the exercise taking May 2010 as reference date, suggesting that the cost of credit in these countries increased mainly in the first part of the sovereign crisis.

4. Robustness checks

In this section we assess the robustness of the results to changes in sample period, to an alternative econometric approach and to different data on the cost of loans.

4.1. Sub-sample analysis

Here we give the results obtained by estimating our system of equations over the entire crisis period (2008-2012). Comparing the results on the cost of new loans to non-financial corporations (Table 7 and Table 3) suggests that the results are qualitatively robust to the choice of the sample period; the adjustment to changes in EONIA is, on average, slightly greater in the shorter sample, and the sensitivity of bank rates to changes in sovereign spreads is almost identical in the two periods. The speed of adjustment (measured by the coefficient of the lagged bank rate), is on average slower. The similarity in the coefficients of the system of equations translates into similar results for the counterfactual exercises described in Section 3.1 as well.

Similarly, a look at the results for the rates on new residential mortgages (Table 8 and Table 4) indicates very little difference between the two sample periods. The sensitivity of lending rates to sovereign spreads is slightly greater in the peripheral countries in the shorter sample, whereas the results of the counterfactual simulations are similar to those discussed in Section 3.1. Interestingly, if the system of equations in (1) is estimated over the period 2003-2007 the parameters measuring the pass-through of changes in sovereign spreads to bank lending rates in all the countries considered are not statistically different from zero. This is in accordance with the thesis that prior to the crisis government bond yields had little importance for banks' price-setting policies for short-term loans.

4.2. Single equation ordinary least squares

In this section we test for robustness of the results using ordinary least squares (OLS) instead of the seemingly unrelated regression approach to estimate the system of equations in (1). There are two reasons for choosing the SUR methodology: a gain in efficiency thanks to combining information from the different equations and the possibility of testing restrictions across equations (Moon and Perron, 2006). There is no theoretical result that applies to our case, as the equations have different explanatory variables and the disturbances are allowed to be correlated. However, it can be shown (Greene, 2011) that the gain in efficiency tends to be greater when the correlation among the disturbances is stronger and when the explanatory variables are not strongly correlated. One potential problem with the SUR approach is that it may propagate misspecification and inconsistency across equations. For if even one of the equations is misspecified then under SUR all the parameters are estimated inconsistently. This danger suggests the advisability of estimating each equation separately.

In our case, despite the potential disadvantages of the SUR methodology, comparison of Tables 3-4 with Tables 9-10 shows that our results are robust to the use of OLS. The coefficients on the sovereign spreads and EONIA are, on average, marginally higher with OLS, but the results of the counterfactual simulations are nearly identical both for non-financial corporations and for households.

4.3 Alternative measures of the cost of loans

The dependent variable in the foregoing analysis has been interest rates on new loans. But the overall impact of the sovereign spread on the debt-service burden of non-financial corporations and households depends on the interest paid on existing loans as well. While for firms the fraction of variable rates – which naturally react more rapidly and strongly to changes in banks' funding costs – is above 90 per cent in all the countries considered, for households the proportion averages only 60 per cent.

In this light it is important to test the robustness of our results to alternative definitions of lending rates. For non-financial corporations the system in (1) is estimated using the interest rate on outstanding loans (up to one year) including overdrafts; for households, the dependent variable is the average cost of outstanding loans at both variable and fixed rates. The results, which are not reported, are qualitatively similar albeit with some quantitative differences. The coefficient of the EONIA rate is lower, on average across countries, than in the estimation discussed in Section 3; the coefficient of the sovereign spreads is lower and that of the lagged value of the bank rate is higher. The counterfactual exercises confirm that the sovereign debt tensions had a significant effect on the cost of loans to non-financial corporations.

For households, when the dependent variable is the average rate on existing loans at variable and fixed rates, the coefficients of the sovereign spreads are positive and significant also in the core countries, while those of EONIA are much smaller. The former finding can be explained by the fact that fixed rates on loans to households are generally set on the basis of long-term yields, including government bond yields. In fact, if sovereign spreads are replaced with the yields, the associated coefficients are high and significant in some core countries as well.

4.4 Error correction models

Error correction models have been used extensively in the literature on the bank interest rate pass-through (Sørensen and Werner, 2006 and Marotta, 2010). Changes in a specific bank rate are regressed on simultaneous and lagged changes in a relevant market rate and on an error correction term that captures the divergence of the bank rate from its long-run equilibrium relationship with the market rate.

Section 3 shows that all lending rates are non-stationary and follow unit root processes. Here we test for the presence of cointegration relationships between lending rates and EONIA. The Engle and Granger (1987) tests reported in Table 11 find no cointegrating relationship between the interest rate on new loans to non-financial corporations and EONIA. A graphical analysis of the residuals of the first-step regression (Figures 11 and 12) suggests that the sovereign crisis may have caused a break in the cointegrating relationship in the peripheral countries. If instead of EONIA we use three-month Euribor, a cointegrating relationship is found for households only for Germany and Finland; for non-financial corporations, only for Austria, Germany, Finland and the Netherlands. In these cases, the estimation of an error correction model shows that the sovereign spreads are not statistically significant, confirming the findings of Section 3.

5. Conclusion

Since the beginning of 2010, the strains in some sovereign debt markets have progressively distorted monetary conditions in the euro area as a whole and in some member countries. The fragmentation of financial markets along national lines has prevented the homogeneous transmission of the ECB monetary policy stance to the interest rates on new bank loans to households and firms. The three-year refinancing operations, the expansion of eligible collateral in late 2011 and more recently the introduction Outright Monetary Transactions have been designed to preserve the singleness of monetary policy of the ECB and the correct functioning of its transmission mechanism.

This paper presents the results of an econometric assessment of the transmission of sovereign tensions to lending rates in the main euro-area countries. The model estimated finds that the sovereign debt tensions have had a significant impact on lending rates in the peripheral but not in the core countries. A counterfactual exercise suggests that, all else constant, if the sovereign crisis had not occurred, then the rates on new bank loans to non-financial corporations and residential mortgages loans to households in the peripheral countries would have been significantly lower.

The analysis is based on the estimation of simple interest-rate equations. Future research might explore the role of time variation in the parameters and the possibility of non-linear effects of the sovereign spreads. An extension of the analysis by means of Vector Auto-Regressive models, moreover, could quantify the impact of the sovereign debt crisis not only on the cost of credit but also on the main macroeconomic variables, including money, credit and economic activity.

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Appendix: Data

All the data are available in the ECB *Statistical Dataware House* (SDW) or in the OECD.StatExtracts webpage. Data in the SDW for each country can be obtained by replacing the percentage sign % with the following codes:

Austria	=	AT	Belgium	=	BE
Germany	=	DE	Spain	=	ES
Finland	=	FI	France	=	FR
Greece	=	GR	Italy	=	IT
The Netherlands	=	NL	Portugal	=	PT

- **Interest rate on loans to non-financial corporations**: Annualised agreed rate / Narrowly defined effective rate, credit and other institutions (Monetary Financial Institutions, MFIs, except money market funds and central banks) reporting sector - loans other than revolving loans and overdrafts, convenience and extended credit card debt, initial maturity up to one year maturity, total amount, new business coverage, euro, non-financial corporations sector.
  Statistical Dataware House: MIR.M.%.B.A2A.F.R.A.2240.EUR.N

- **Interest rate on residential mortgages (loans to households for house purchase)**: Annualised agreed rate / Narrowly defined effective rate, credit and other institutions (MFIs except money market funds and central banks) reporting sector - lending for house purchase excluding revolving loans and overdrafts, initial maturity up to one year maturity, new business coverage, euro, households and non-profit institutions serving households.
  Statistical Dataware House: MIR.M.%.B.A2C.F.R.A.2250.EUR.N

- **Loans to non-financial corporations**: Index of notional stocks, MFIs excluding the ESCB reporting sector - loans, total maturity, all currencies combined - euro area (changing composition) counterpart, non-financial corporations sector, annual growth rate, data not adjusted for seasonal or calendar effects.
  Statistical Dataware House: BSI.M.%.N.A.A20.A.I.U2.2240.Z01.A

- **Loans to households**: Index of notional stocks, MFIs excluding ESCB reporting sector - loans, total maturity, all currencies combined - euro area (changing composition) counterpart, households and non-profit institutions serving households sector, annual growth rate, data not adjusted for seasonal or calendar effects.
  Statistical Dataware House: BSI.M.%.N.A.A20.A.I.U2.2250.Z01.A

- **EONIA rate**: Eonia rate - Historical close, average of observations through period - provided by ECB
  Statistical Dataware House: FM.M.U2.EUR.4F.MM.EONIA.HSTA

- **Three-month EURIBOR rate**: Euribor three-month - Historical close, average of observations through period - Euro, provided by Reuters
  Statistical Dataware House: FM.M.U2.EUR.RT.MM.EURIBOR3MD_.HSTA

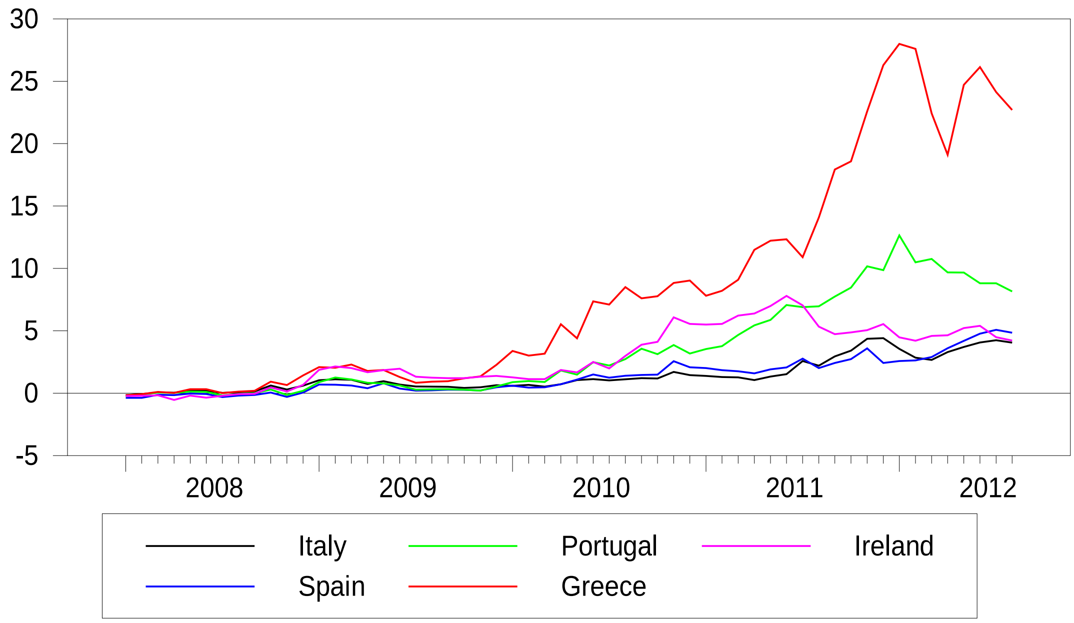
- **Non-financial corporations' confidence indicator**: Main Economic Indicators (OECD); monthly; business tendency surveys, seasonally adjusted. Available at http://stats.oecd.org/

- **Households' confidence indicator**: Main Economic Indicators (OECD); monthly; consumer opinion surveys, seasonally adjusted. Available at http://stats.oecd.org/

Figures and Tables

Fig. 1 Spreads between 10-year yields on sovereign bonds and the 10-year swap rate in euro

(a) Peripheral countries



(b) Core countries

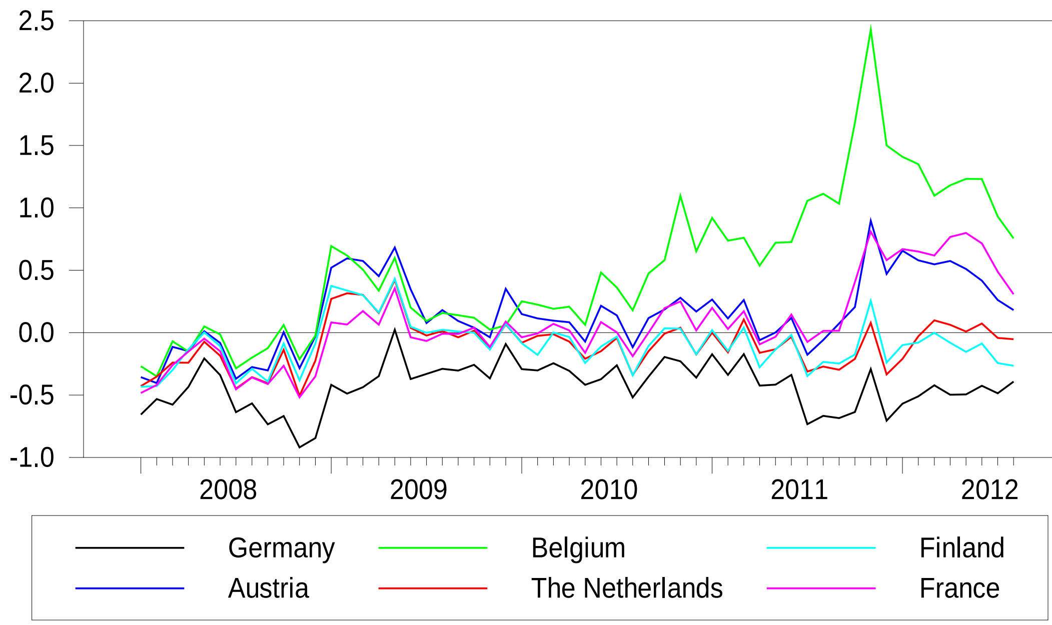
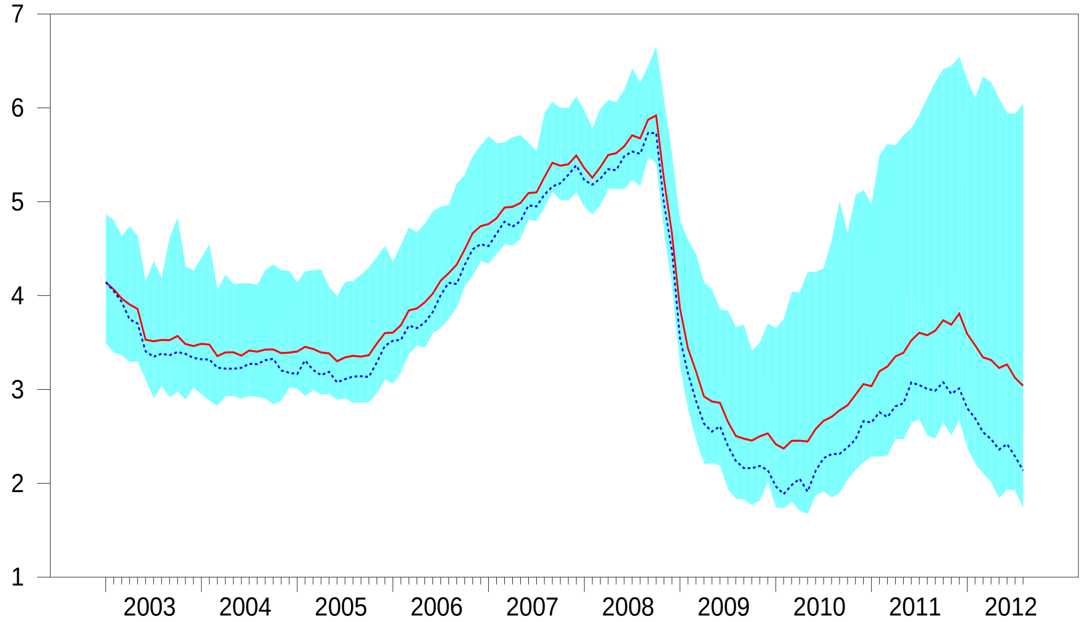
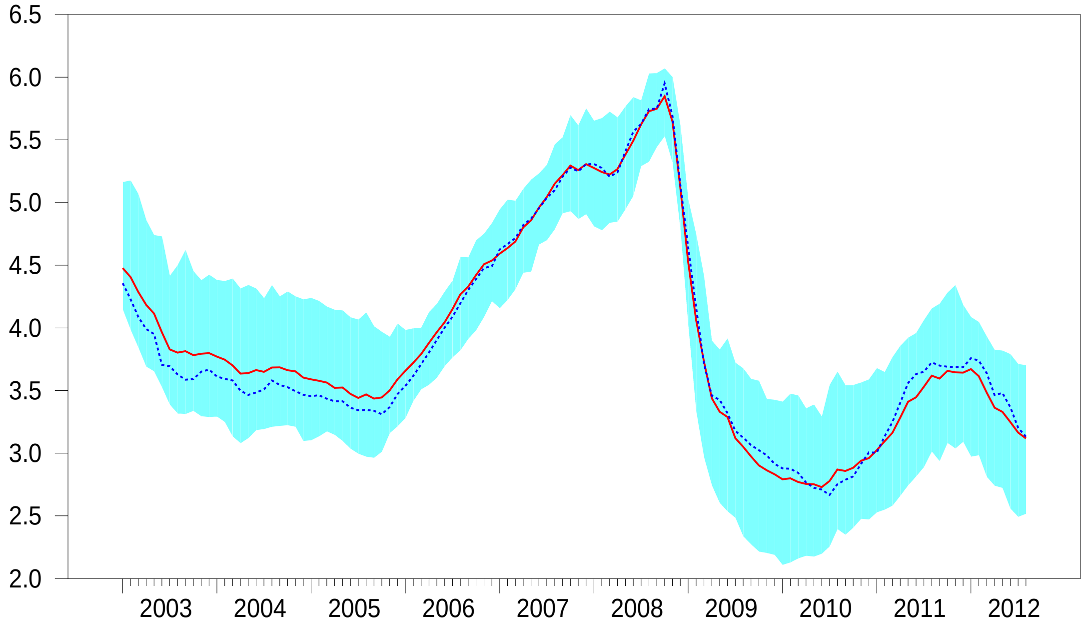


Fig. 2 Interest rates on new loans to non-financial corporations



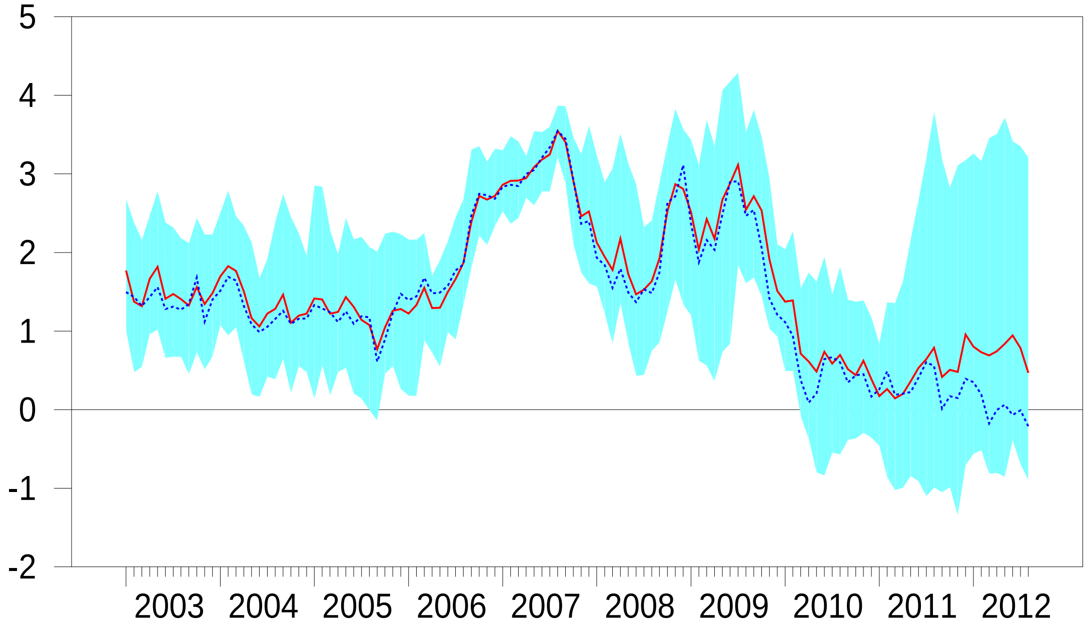
Note: blue shaded area: 10th-90th percentiles; blue dotted line: median; red line: mean. Percentage points.

Fig. 3 Interest rate on new loans to households for house purchases



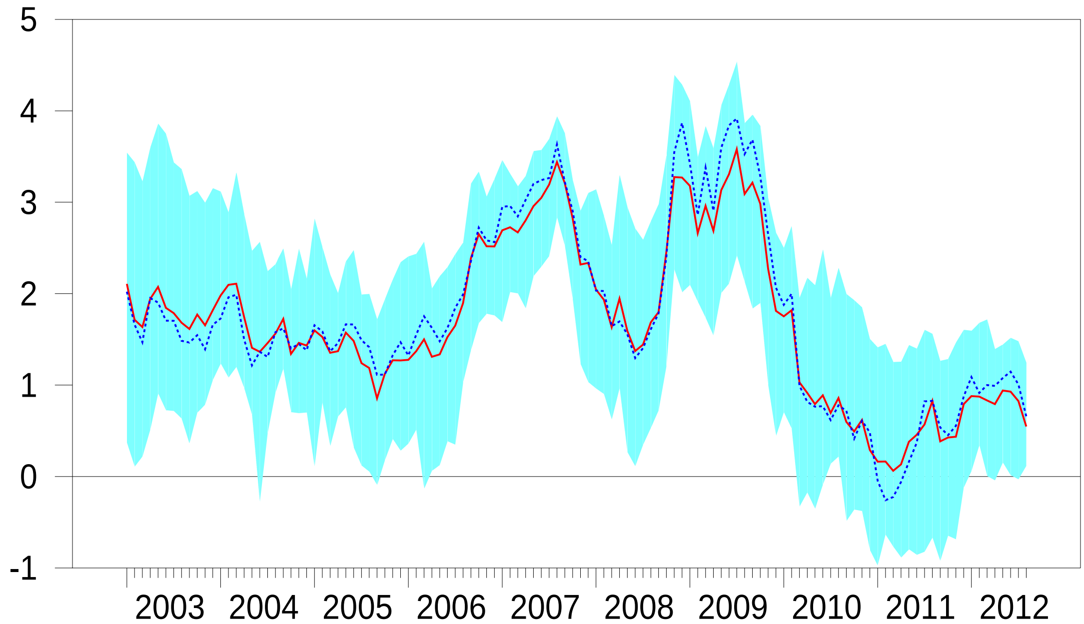
Note: blue shaded area: 10th-90th percentiles; blue dotted line: median; red line: mean. Percentage points.

Fig. 4 Real interest rates on new loans to firms, deflated with current HICP inflation rates



Note: blue shaded area: 10th-90th percentiles; blue dotted line: mean; red line: median. Percentage points.

Fig. 5 Real interest rates on new loans to households for house purchase, deflated with current HICP inflation rates

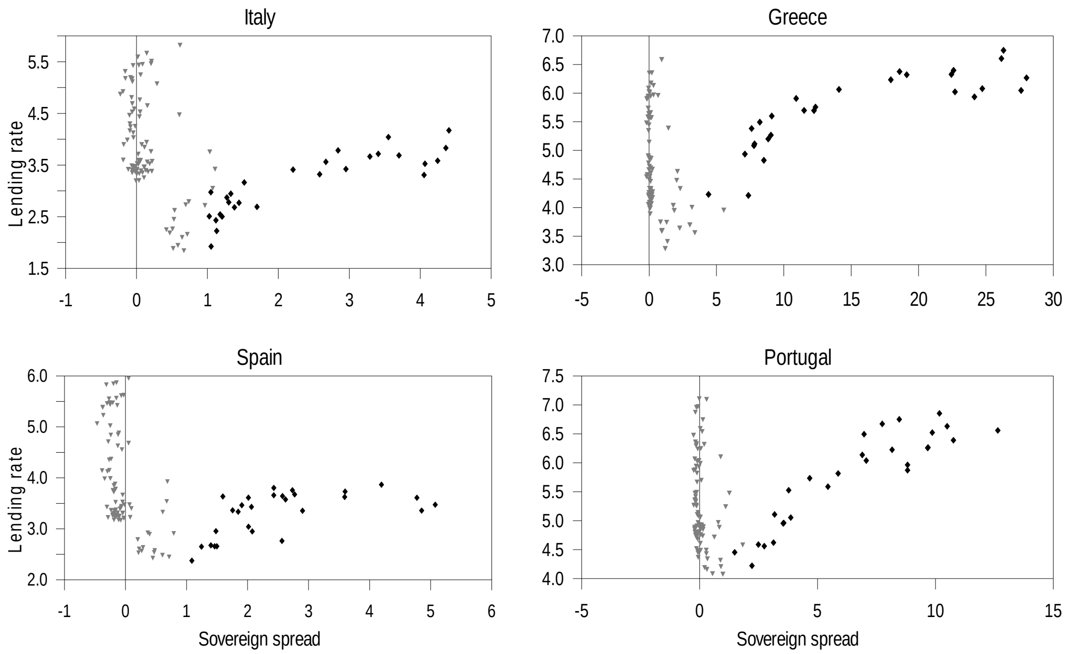


Note: blue shaded area: 10th-90th percentiles; blue dotted line: mean; red line: median. Percentage points.

Fig. 6 Sovereign spreads and interest rates on new loans

(*grey triangles: 2003:01 – 2010:04; black diamonds: 2010:5 – 2012:08; percentage points*)

a) Non-financial corporations



b) Households

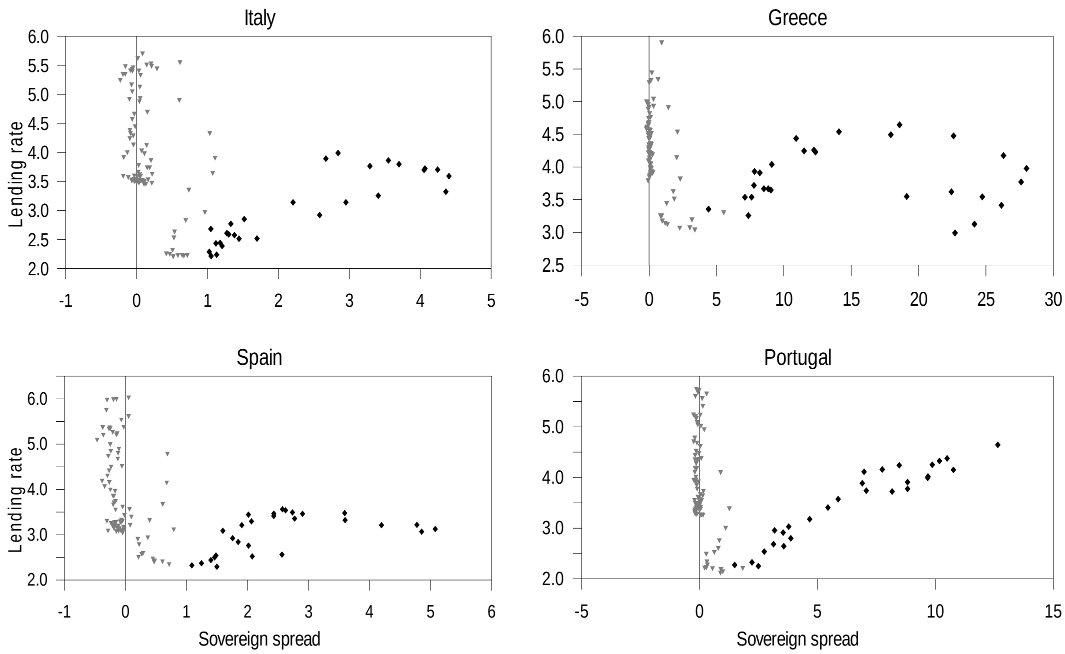
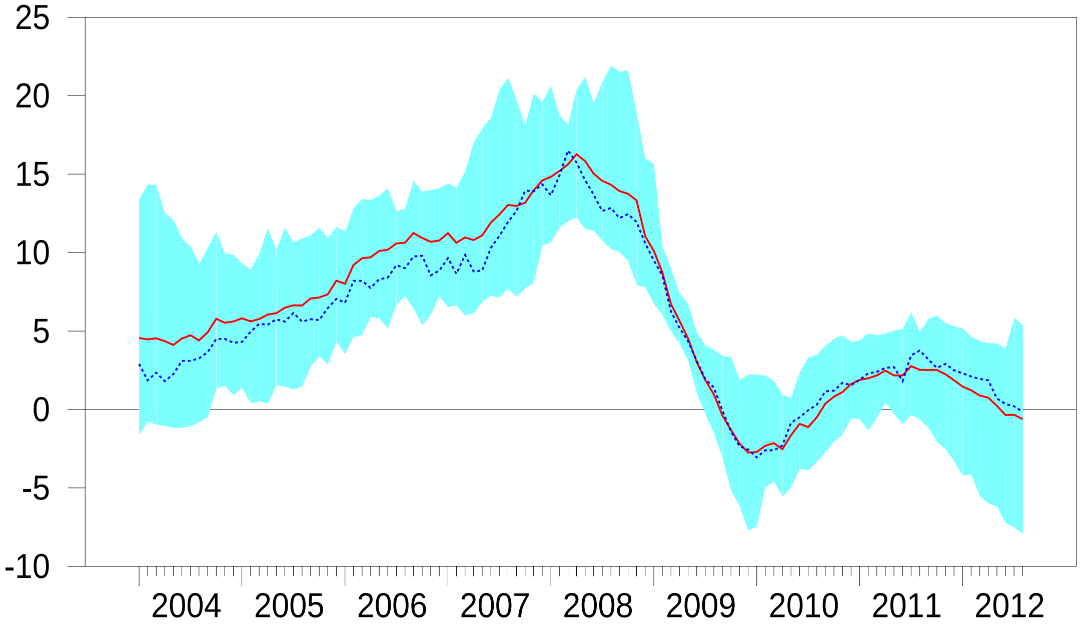
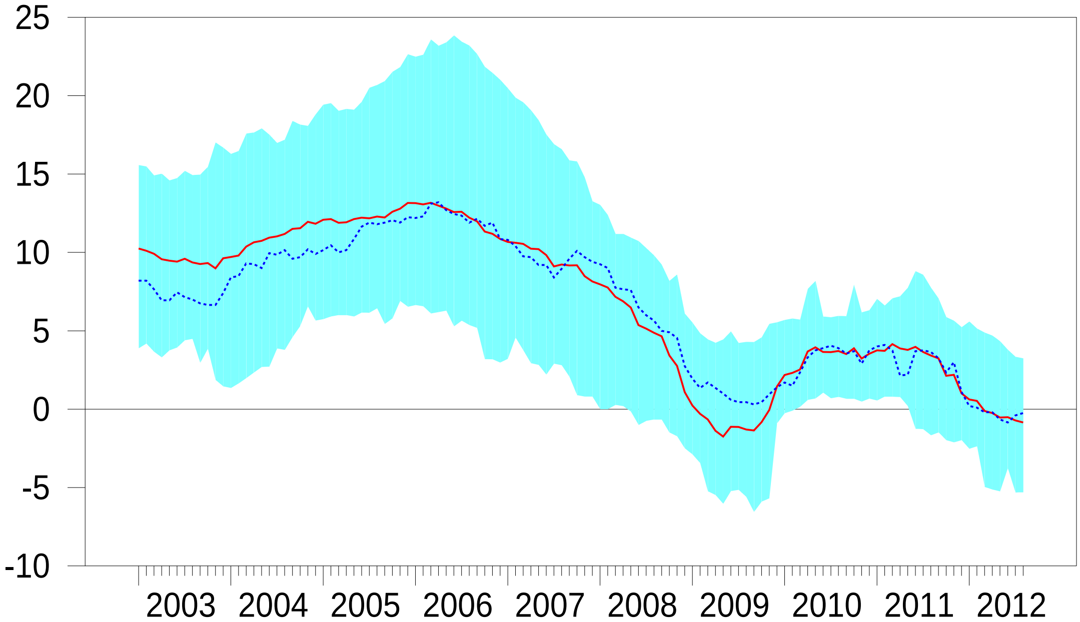


Fig. 7 Loans to non-financial corporations
(*twelve-month growth rates; per cent*)



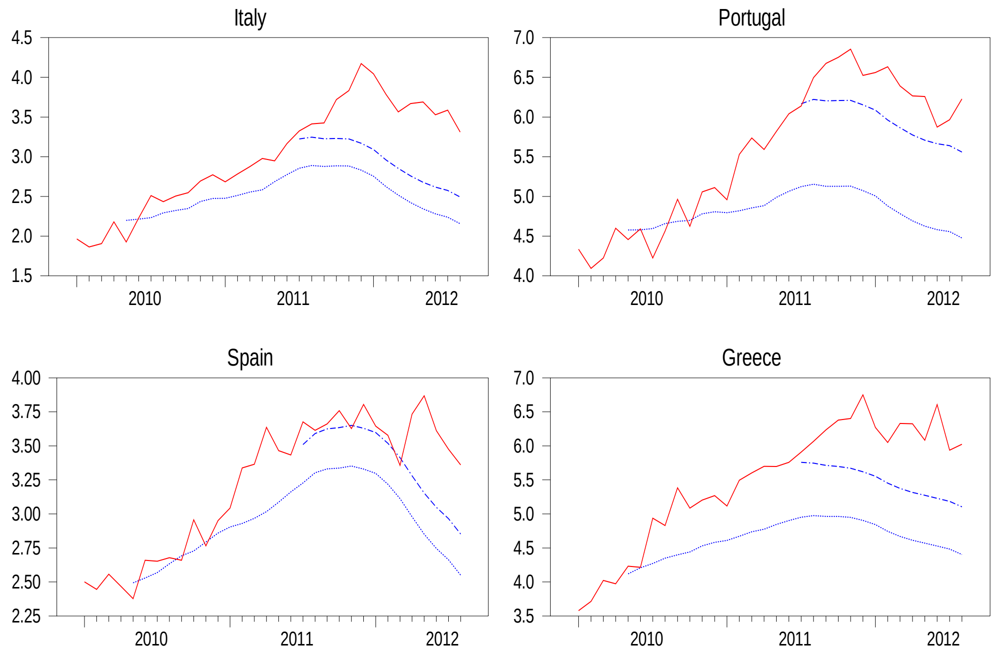
Note: blue area: 10th-90th percentiles; blue dotted line: median; red line: mean.

Fig. 8 Loans to households for house purchase
(*twelve-month growth rates; per cent*)



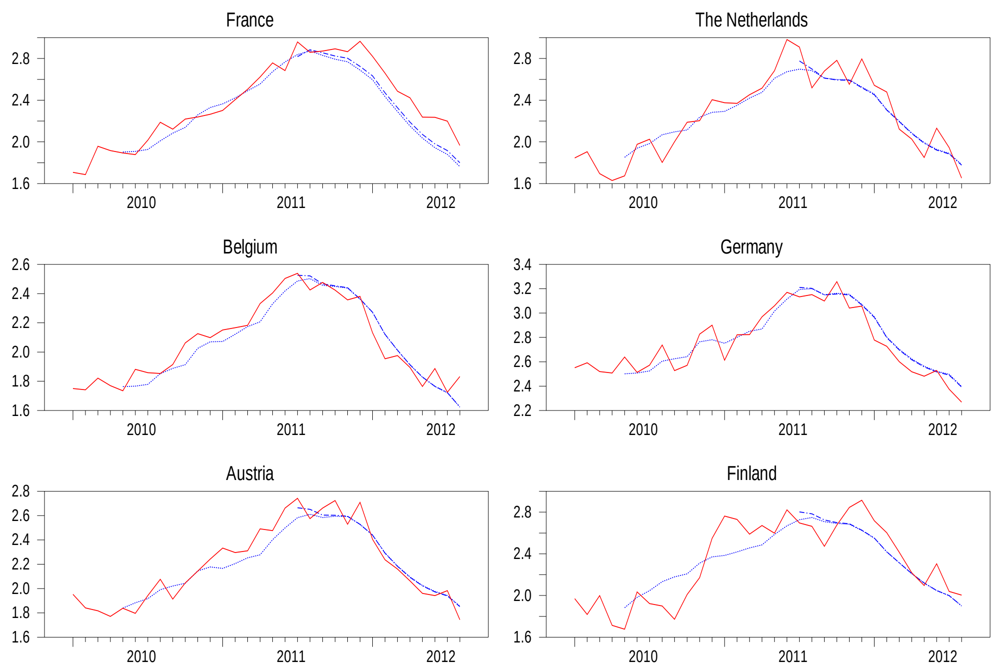
Note: blue area: 10th-90th percentiles; blue dotted line: median; red line: mean.

Fig. 9 Interest rates on new loans to non-financial corporations: counterfactual simulations - peripheral countries



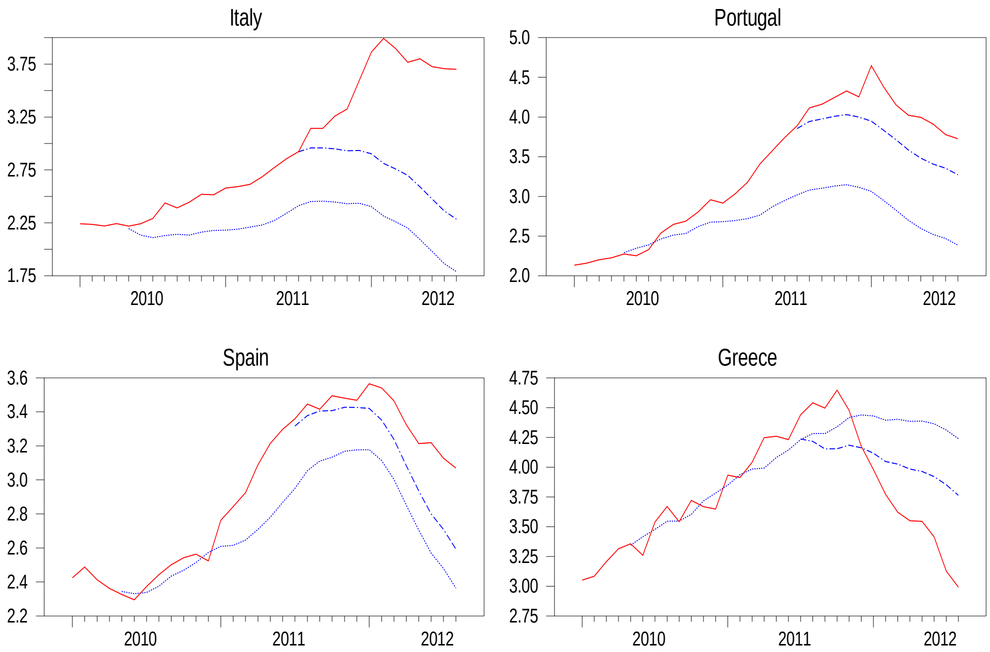
Note: red lines: actual data; blue dotted lines: simulated data starting from May 2010; blue dashed lines: simulated data starting from July 2011. Percentage points.

Fig. 10 Interest rates on new loans to non-financial corporations: counterfactual simulations - core countries



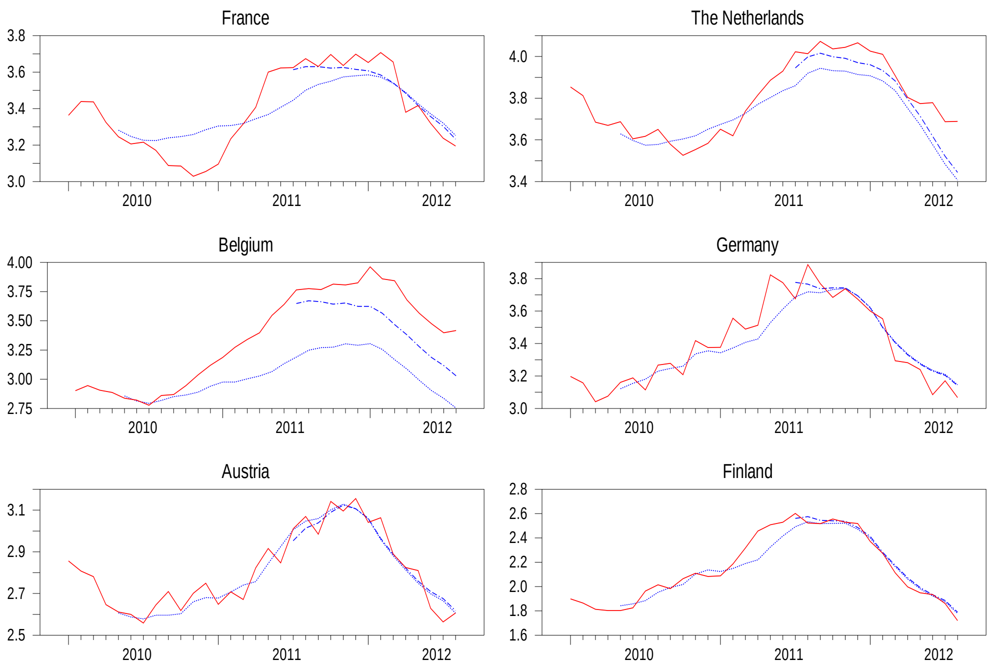
Note: red lines: actual data; blue dotted lines: simulated data starting from May 2010; blue dashed lines: simulated data starting from July 2011. Percentage points.

Fig. 11 Interest rates on new loans to households for house purchase: counterfactual simulations - peripheral countries



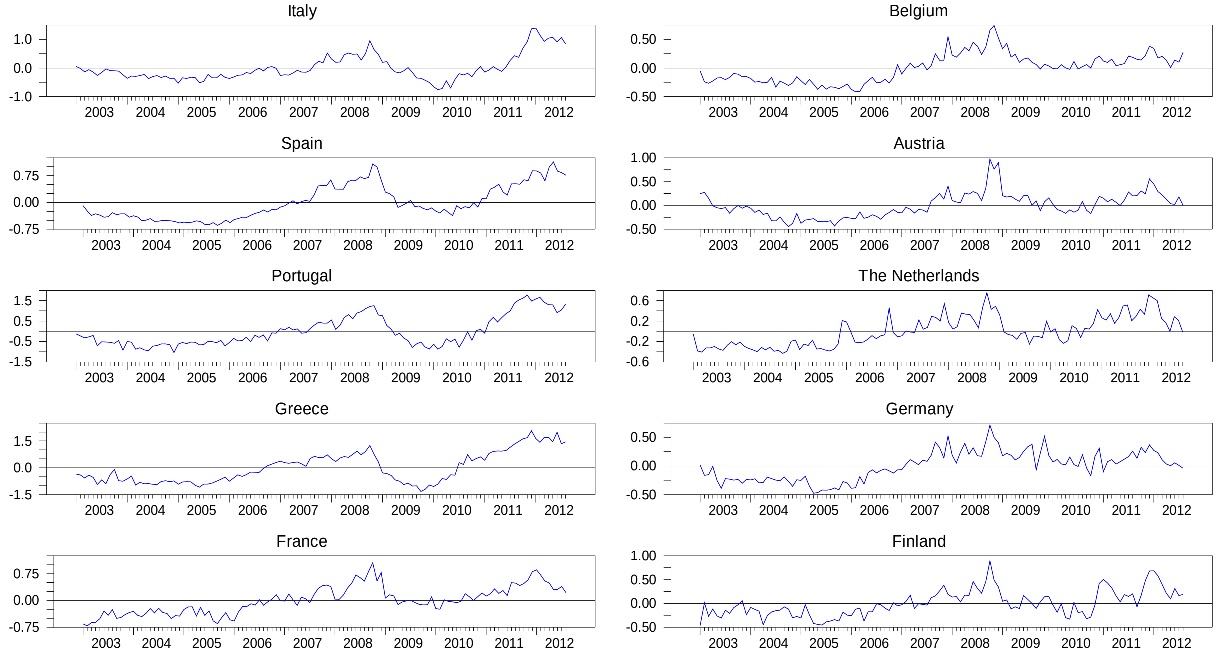
Note: red lines: actual data; blue dotted lines: simulated data starting from May 2010; blue dashed lines: simulated data starting from July 2011. Percentage points.

Fig. 12 Interest rates on new loans to households for house purchase: counterfactual simulations - core countries



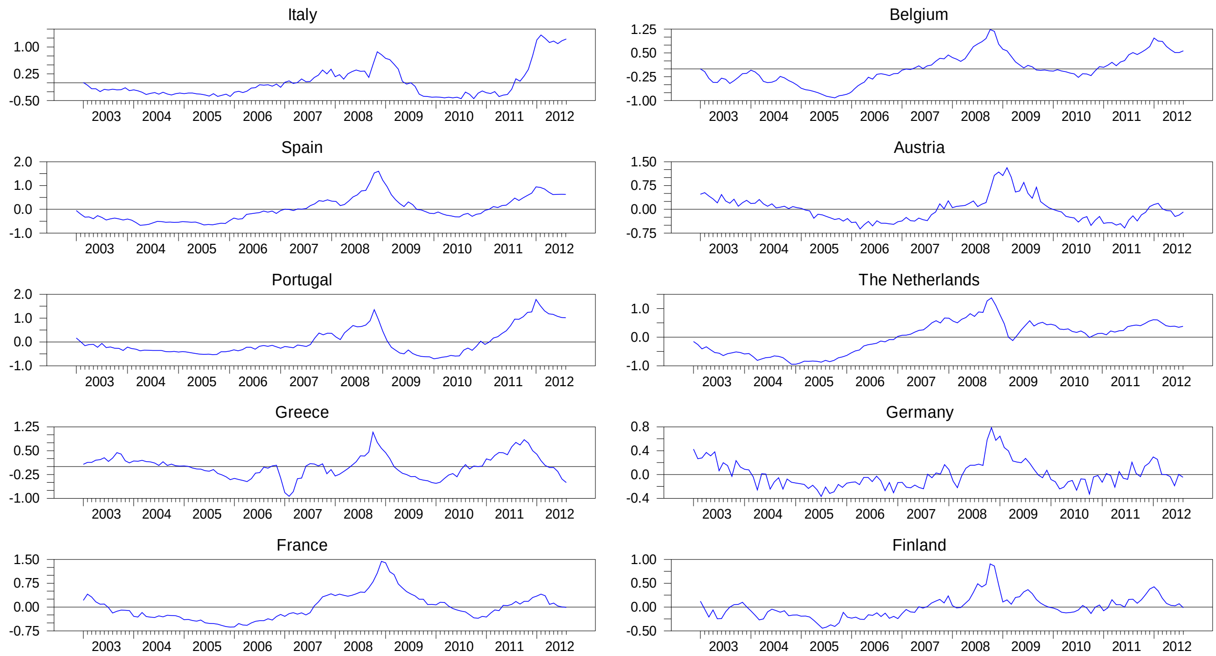
Note: red lines: actual data; blue dotted lines: simulated data starting from May 2010; blue dashed lines: simulated data starting from July 2011. Percentage points.

Fig. 13 Disequilibrium term: interest rates on new loans to non-financial corporations



Note: The graphs report the residual of the regression of each lending rate on EONIA. Percentage points.

Fig. 14 Disequilibrium term: interest rates on new loans to households for house purchase



Note: The graphs report the residual of the regression of each lending rate on EONIA. Percentage points.

Table 1. Standard deviation of lending rates and credit growth

	Interest rate on new loans to non-financial corporations	Interest rate on new loans to households	Loans to non-financial corporations	Loans to households for house purchase
2003:1-2012:8	0.81	0.47	5.03	6.08
2003:1-2007:8	0.58	0.43	6.36	7.48
2003:1-2010:4	0.62	0.45	5.41	6.95
2007:9-2012:8	1.02	0.51	3.78	4.78
2007:9-2010:4	0.69	0.46	3.74	6.02
2010:5-2012:8	1.39	0.56	3.82	3.35
2011:7-2012:8	1.60	0.58	4.41	3.67

Note: percentage points.

Table 2. Unit root tests

	ADF	KPSS	PP
Non-financial corporations			
Italy	-1.963	0.460	-1.474
Spain	-2.097	0.366	-1.448
Portugal	-1.995	0.540*	-1.521
Greece	-1.733	0.781**	-1.352
France	-2.019	0.556*	-1.215
Belgium	-2.042	0.781**	-0.995
Austria	-1.567	0.797**	-1.114
Netherlands	-1.486	0.691*	-1.102
Germany	-1.517	0.739**	-0.966
Finland	-1.478	0.757**	-1.144
Households			
Italy	-2.320	0.574*	-1.267
Spain	-2.174	0.434	-1.356
Portugal	-2.398	0.267	-1.672
Greece	-2.525	0.407	-1.264
France	-2.077	0.567*	-1.716
Belgium	-2.158	0.343	-1.519
Austria	-1.549	1.118**	-0.841
Netherlands	-1.844	0.567*	-1.488
Germany	-1.918	1.041**	-0.997
Finland	-1.410	0.845**	-1.092

Note: ADF = Augmented Dickey–Fuller; KPSS = Kwiatkowski–Phillips–Schmidt–Shin. PP = Phillips-Perron. * = significance at 5 per cent; ** = significance at 1 per cent. In all three tests 3 lags are used. The tests follow the procedures for WinRats 8.0 available on the webpage of Estima; dfunit.src for the ADF test, kpss.src for the KPSS and ppunit.src for the PP.

Table 3. Interest rates on loans to non-financial corporations: estimated coefficients

	(*a*)	(*b*)	(*c*)	(*d*)	(*e*)	(*f*)	(*g*)	(*h*)
	$\overline{R}$	R^{ov}	D^{crisis}	D^{2008}	$R^{3m}-R^{ov}$	$R^{10}-R^{10,swap}$	Y	R_{-1}
Italy	0.592**	0.360**	0.008	-0.518**	0.138*	0.165**	0.014	0.604**
Spain	0.480**	0.314**	0.176**	-0.586**	0.331**	0.079**	0.045**	0.648**
Portugal	2.040**	0.459**	0.352**	-0.200**	0.280**	0.130**	0.012	0.371**
Greece	0.816**	0.246**	0.443**	-0.651**	0.067	0.033**	0.047*	0.675**
France	0.481**	0.501**	0.259**	-0.655**	0.366**	0.151**	0.088**	0.475**
Belgium	0.556**	0.509**	0.140**	-0.502**	0.249**	0.001	0.034**	0.458**
Austria	0.539**	0.453**	0.194**	-0.612**	0.172**	-0.048	0.009	0.512**
Netherlands	0.875**	0.609**	0.229**	-0.558**	0.386**	-0.108	0.071**	0.268**
Germany	1.045**	0.564**	0.306**	-0.623**	0.200**	0.108*	0.010	0.368**
Finland	0.679**	0.466**	0.136**	-0.690**	0.295**	0.035	0.024*	0.476**

Note: Standard errors are computed using the heteroskedasticity and autocorrelation consistent estimate of the covariance matrix of the coefficients. * = significance at 10 per cent; ** = significance at 5 per cent. Sample period: 2003:1 – 2012:8.

Table 4. Interest rates on loans to households for house purchase: estimated coefficients

	(*a*)	(*b*)	(*c*)	(*d*)	(*e*)	(*f*)	(*g*)	(*h*)
	$\overline{R}$	R^{ov}	D^{crisis}	D^{2008}	$R^{3m}-R^{ov}$	$R^{10}-R^{10,swap}$	Y	R_{-1}
Italy	0.395**	0.230**	-0.135**	-0.272**	0.145**	0.155**	0.004**	0.750**
Spain	0.385**	0.287**	0.079*	-0.329**	0.373**	0.052**	0.001	0.689**
Portugal	0.809**	0.390**	-0.054	-0.310**	0.320**	0.081**	-0.001	0.521**
Greece	0.386**	0.091**	0.079	-0.453**	-0.149*	-0.010**	-0.004**	0.879**
France	0.421**	0.115**	0.091**	-0.110*	0.170**	-0.050	0.001	0.806**
Belgium	0.279**	0.170**	0.058	-0.515**	0.224**	0.097**	0.003**	0.810**
Austria	0.648**	0.249**	-0.020	-0.070	0.035	0.098	-0.001	0.713**
Netherlands	0.181**	0.099**	0.109**	-0.512**	0.204**	0.074	0.002**	0.883**
Germany	1.220**	0.343**	0.041	-0.299**	0.114*	0.075	0.000	0.563**
Finland	0.526**	0.412**	0.088**	-0.678**	0.217**	0.266**	0.001	0.558**

Note: Standard errors are computed using the heteroskedasticity and autocorrelation consistent estimate of the covariance matrix of the coefficients. * = significance at 10 per cent; ** = significance at 5 per cent. Sample period: 2003:1 – 2012:8.

Table 5. The impact of sovereign spreads on bank rates on loans to non-financial corporations and households: counterfactual exercise starting in May 2010

	(*a*)	(*b*)	(*c*)	(*d*)	(*e*)	(*d*)	(*e*)
	$R^{10}-R^{10,swap}$	NFCs	households	NFCs	households	NFCs	households
	April 2010	2011[(a)]	2011[(a)]	Dec. 2011	Dec. 2011	Aug. 2012	Aug. 2012
Italy	0.71	0.54	0.62	1.34	1.16	1.15	1.91
Spain	0.71	0.37	0.33	0.47	0.29	0.81	0.70
Portugal	1.82	1.08	0.80	1.45	1.14	1.75	1.34
Greece	5.53	1.07	0.12	1.85	-0.26	1.62	-1.25
peripheral[(b)]	*2.19*	*0.77*	*0.47*	*1.28*	*0.58*	*1.33*	*0.68*
France	0.02	0.05	0.08	0.28	0.12	0.20	-0.06
Belgium	0.21	0.03	0.45	0.02	0.53	0.21	0.67
Austria	0.10	0.10	0.00	0.18	0.05	-0.11	0.00
Netherlands	-0.10	0.09	0.07	0.27	0.15	-0.13	0.28
Germany	-0.30	-0.01	0.08	-0.01	-0.02	-0.13	-0.08
Finland	-0.05	0.10	0.07	0.29	0.05	0.10	-0.06
Euro area[(c)]	*0.87*	*0.35*	*0.26*	*0.61*	*0.32*	*0.55*	*0.35*

Note: Based on coefficients estimated over the period 2003:1 – 2012:8. The sovereign spreads are assumed to remain constant at their April 2010 levels over the period between May 2010 and August 2012. Columns from (*b*) to (*g*) report the differences between actual and simulated values. [(a)] Average in 2011. [(b)] Simple average over the peripheral countries. [(c)] Simple average.

Table 6. The impact of sovereign spreads on bank rate on loans to non-financial corporations and households: counterfactual exercise starting in July 2011

	(*a*)	(*b*)	(*c*)	(*d*)	(*e*)	(*d*)	(*e*)
	$R^{10} - R^{10,swap}$	NFCs	households	NFCs	households	NFCs	households
	June 2011	2011[(a)]	2011[(a)]	Dec. 2011	Dec. 2011	Aug. 2012	Aug. 2012
Italy	1.52	0.43	0.29	1.00	0.66	0.82	1.42
Spain	2.06	0.08	0.05	0.18	0.04	0.51	0.48
Portugal	7.06	0.38	0.20	0.37	0.25	0.67	0.45
Greece	12.34	0.59	0.28	1.13	0.01	0.92	-0.77
peripheral [(b)]	*5.74*	*0.37*	*0.20*	*0.67*	*0.24*	*0.73*	*0.39*
France	0.14	0.08	0.04	0.24	0.08	0.17	-0.04
Belgium	0.73	-0.03	0.14	0.02	0.20	0.21	0.39
Austria	0.12	0.05	0.02	0.18	0.05	-0.11	-0.01
Netherlands	-0.03	0.07	0.06	0.28	0.09	-0.13	0.24
Germany	-0.34	-0.03	-0.01	-0.01	-0.02	-0.12	-0.07
Finland	-0.02	-0.01	0.00	0.29	0.03	0.10	-0.07
Euro area[(c)]	*2.36*	*0.16*	*0.11*	*0.37*	*0.14*	*0.30*	*0.20*

Note: Based on coefficients estimated over the period 2003:1 – 2012:8. The sovereign spreads are assumed to remain constant at their June 2011 levels over the period between July 2011 and August 2012. Columns from (*b*) to (*g*) report the differences between actual and simulated values. [(a)] Average in 2011. [(b)] Simple average over the peripheral countries. [(c)] Simple average.

Table 7. Interest rates on loans to non-financial corporations: estimated coefficients

	(*a*)	(*b*)	(*c*)	(*d*)	(*e*)	(*f*)	(*g*)	(*h*)
	$\bar{R}$	R^{ov}	D^{crisis}	D^{2008}	$R^{3m} - R^{ov}$	$R^{10} - R^{10,swap}$	Y	R_{-1}
Italy	0.578**	0.409**	0.093	-0.496**	0.013	0.176**	0.027**	0.585**
Spain	0.713**	0.306**	0.077	-0.473**	0.202*	0.072**	0.057**	0.634**
Portugal	1.549**	0.360**	0.292**	-0.184**	0.141	0.103**	0.045*	0.527**
Greece	0.840**	0.108**	0.341**	-0.451**	-0.194	0.019**	0.098**	0.787**
France	0.598**	0.496**	0.206**	-0.601**	0.221	0.142**	0.078**	0.478**
Belgium	1.688**	1.031**	-0.052*	-0.256**	0.381**	0.009	0.038**	-0.167
Austria	1.039**	0.808**	0.292**	-0.479**	0.476**	0.025	0.036**	0.085
Netherlands	1.048**	0.680**	0.245**	-0.486**	0.387**	-0.123	0.081**	0.167*
Germany	2.050**	0.775**	0.140**	-0.458**	0.166*	0.169**	0.001	0.041
Finland	0.591**	0.428**	0.156**	-0.669**	0.277**	0.055	0.034**	0.525**

Note: Standard errors are computed using the heteroskedasticity and autocorrelation consistent estimate of the covariance matrix of the coefficients. * = significance at 10 per cent; ** = significance at 5 per cent. Sample period: 2008:1 – 2012:8.

Table 8. Interest rates on loans to households for house purchase: estimated coefficients

	(*a*)	(*b*)	(*c*)	(*d*)	(*e*)	(*f*)	(*g*)	(*h*)
	$\bar{R}$	R^{ov}	D^{crisis}	D^{2008}	$R^{3m} - R^{ov}$	$R^{10} - R^{10,swap}$	Y	R_{-1}
Italy	0.441**	0.228**	-0.149**	-0.229**	0.067	0.157**	0.004**	0.750**
Spain	0.403**	0.285**	0.065	-0.327**	0.444**	0.057**	0.001	0.675**
Portugal	1.005**	0.644**	0.085	-0.236**	0.484**	0.134**	0.000	0.271**
Greece	0.195	0.067	0.065	-0.469**	-0.286**	-0.012**	-0.004**	0.959**
France	0.687**	0.155**	0.095**	-0.054	0.215**	-0.025	0.000	0.713**
Belgium	0.510**	0.203**	0.020**	-0.438	0.067**	0.123**	0.001	0.764**
Austria	0.835**	0.303**	-0.017	-0.030	0.001	0.080	-0.002*	0.646**
Netherlands	0.592**	0.134**	0.075	-0.431**	0.189**	0.008	0.001*	0.779**
Germany	2.200**	0.616**	0.095**	-0.224**	0.466**	0.036	-0.001	0.168**
Finland	0.691**	0.537**	0.142**	-0.672**	0.287**	0.208**	0.000	0.410**

Note: Standard errors are computed using the heteroskedasticity and autocorrelation consistent estimate of the covariance matrix of the coefficients. * = significance at 10 per cent; ** = significance at 5 per cent. Sample period: 2008:1 – 2012:8.

Table 9. Interest rates on loans to non-financial corporations: estimated coefficients

	(*a*)	(*b*)	(*c*)	(*d*)	(*e*)	(*f*)	(*g*)	(*h*)
	$\overline{R}$	R^{ov}	D^{crisis}	D^{2008}	$R^{3m}-R^{ov}$	$R^{10}-R^{10,swap}$	Y	R_{-1}
Italy	0.599**	0.353**	0.016	-0.476**	0.153*	0.167**	0.035**	0.603**
Spain	0.500**	0.336**	0.192	-0.555**	0.357**	0.094**	0.059**	0.624**
Portugal	2.061**	0.463**	0.355**	-0.180**	0.277**	0.133**	0.016	0.364**
Greece	0.650**	0.192**	0.443**	-0.624**	0.050	0.027**	0.061**	0.735**
France	0.541**	0.560**	0.292**	-0.600**	0.412**	0.213**	0.115**	0.411**
Belgium	0.564**	0.508**	0.141**	-0.460**	0.245**	0.012	0.051**	0.455**
Austria	0.632**	0.518**	0.215**	-0.554**	0.191**	0.011	0.025**	0.437**
Netherlands	0.912**	0.621**	0.240**	-0.518**	0.394**	-0.037	0.085**	0.249**
Germany	1.028**	0.550**	0.318**	-0.564**	0.197*	0.184**	0.022**	0.385**
Finland	0.627**	0.433**	0.139**	-0.670**	0.275**	0.074	0.030**	0.515**

Note: Equation-by-equation estimation. Standard errors are computed using the heteroskedasticity and autocorrelation consistent estimate of the covariance matrix of the coefficients. * = significance at 10 per cent; ** = significance at 5 per cent. Sample period: 2003:1 – 2012:8. Each equation is estimated with ordinary least squares.

Table 10. Interest rates on loans to households: estimated coefficients

	(*a*)	(*b*)	(*c*)	(*d*)	(*e*)	(*f*)	(*g*)	(*h*)
	$\overline{R}$	R^{ov}	D^{crisis}	D^{2008}	$R^{3m}-R^{ov}$	$R^{10}-R^{10,swap}$	Y	R_{-1}
Italy	0.431**	0.260**	-0.143**	-0.249**	0.145**	0.180**	0.005**	0.721**
Spain	0.435**	0.350**	0.109**	-0.261**	0.416**	0.078**	0.003	0.631**
Portugal	0.807**	0.400**	-0.058	-0.287**	0.308**	0.088**	-0.001	0.515**
Greece	0.327**	0.091**	0.076	-0.449**	-0.165**	-0.009**	-0.004**	0.893**
France	0.483**	0.124**	0.100**	-0.102**	0.181**	-0.066**	0.001	0.783**
Belgium	0.234**	0.164**	0.036	-0.494**	0.234**	0.111**	0.004**	0.826**
Austria	0.632**	0.236**	-0.022	-0.068	0.051	0.065	0.000	0.723**
Netherlands	0.199**	0.094**	0.110**	-0.496**	0.222**	-0.011	0.002**	0.876**
Germany	1.238**	0.343**	0.038	-0.268**	0.116*	0.097	0.000	0.561**
Finland	0.422**	0.349**	0.069	-0.660**	0.211**	0.157**	0.002**	0.628**

Note: Equation-by-equation estimation. Standard errors are computed using the heteroskedasticity and autocorrelation consistent estimate of the covariance matrix of the coefficients. * = significance at 10 per cent; ** = significance at 5 per cent. Sample period: 2003:1 – 2012:8. Each equation is estimated with ordinary least squares (OLS).

Table 11. Testing for cointegration

	Interest rate on new loans to non-financial corporations [a)]	Interest rate on new loans to households [b)]
Italy	-1.395	-0.972
Spain	-0.979	-1.690
Portugal	-0.781	-1.527
Greece	-0.719	-3.377
France	-2.293	-2.145
Belgium	-1.831	-1.657
Austria	-2.718	-2.290
Netherlands	-2.971	-1.343
Germany	-2.484	-2.866
Finland	-3.040	-2.563

Note: test based on Engle and Granger (1987). Critical values from MacKinnon (1991). Tests carried out in WinRats with the procedure egtest.src. *a)* The number of lags is set to 1. *b)* The number of lags is set to 3. * = significance at 5 per cent; ** = significance at 1 per cent. Critical values are from MacKinnon, "Critical Values for Cointegration Tests", in R. F. Engle and C. W. J. Granger (eds.), Long-Run Economic Relationships, London, Oxford U. Press, 1991, pp 267–276.

Supply tightening or lack in demand:
Is the sovereign debt crisis different from Lehman?

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Abstract

This paper analyzes the relative role of demand and supply factors in explaining credit developments in Italy during the financial crisis, focusing on the differences between the "global crisis" and the "sovereign debt crisis". The identification of demand and supply is based on the individual banks' responses to the euro-area Bank Lending Survey. The results indicate that the contribution of weak demand conditions was similar in the two phases, while the supply tightening had a stronger effect during the sovereign debt crisis, as a result of a greater importance of factors related to strains in banks' balance-sheet and funding conditions. Larger effects of the supply tightening are obtained when the Italian sovereign spread is considered in addition to the BLS supply indicators. The impact of supply restrictions is stronger for lending to enterprises than for mortgages to households, reflecting the effect of credit rationing phenomena, rather than a higher elasticity of loan demand to the cost of credit.

JEL Classification: E30; E32; E51.

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1. Introduction

Over the last five years credit developments in the euro area have been heavily affected by the global financial crisis and the following sovereign debt crisis, and by the negative cyclical evolution which accompanied the financial strains. Credit market developments have been quite homogenous across countries during the first phase of the crisis, which was characterized by a generalized contraction of economic activity, a strong worsening of borrowers' creditworthiness and a sharp increase in risk aversion in financial and credit markets. On the contrary, significant heterogeneity characterized the sovereign debt crisis period. The fall in lending has been concentrated in the countries hit by the sovereign debt tensions, where banks' access to wholesale funding worsened abruptly, while credit continued to grow in the other countries.

This paper analyzes the relative role of demand and supply factors in explaining credit developments in Italy during the financial crisis, focusing on the differences between the first phase, which we will indicate as the "global crisis" and the second phase, the "sovereign debt crisis". It also investigates the relative importance of the different supply factors, distinguishing those relating to the cost of funds and balance sheet constraints ("pure supply" factors) on the one hand and those connected to borrowers' creditworthiness and banks' risk perception on the other. This distinction is important, and bears significant policy implications, because the factors driving the tightening of supply conditions may matter for both the effects on credit dynamics and the appropriate policy reactions.

The analysis is carried out on data for Italy, considering separately loans to firms and loans to households for house purchases. Italy is a particularly relevant case in the euro area, as bank lending slowed down sharply between 2008 and the first half of 2009 and, after a brief recovery in 2010-11, fell again in the final part of 2011 and still continues to contract.

The identification of demand and supply factors, which is crucial in the analysis of credit developments, is based on the information provided by the Italian banks participating to the *Bank Lending Survey* (BLS), the quarterly survey on credit conditions carried out in all countries of the euro area since the end of 2002.[1] In particular, we exploit the individual

[1] The survey includes questions on credit standards, loan demand, factors driving loan supply and demand, specific terms and conditions in the provision of loans (such as price and non-price supply conditions). The results are published regularly by the European Central Bank (ECB) for the euro area as a whole and by the Eurosystem national central banks for the respective countries. A detailed description of the survey can be found in Berg et al. (2005).

bank's survey responses (though with no disclosure of individual answers) on loan demand and supply conditions, for the latter distinguishing between the various categories of factors, as mentioned above. This information is combined with bank-level data on loan quantities and interest rates for the same banks, as well as with additional information on interest rates on selected liabilities. The sample period goes from the fourth quarter of 2002 (the first quarter for which the BLS is available) to the second quarter of 2012, which allows us to incorporate the sovereign debt crisis and to analyze whether the role of demand and supply factors differed compared to the global crisis.

The use of bank-level information differentiates this paper from most of the other studies based on the BLS or other lending surveys, which use aggregate data both for survey information and for credit developments.[2] An exception is the paper by Del Giovane, Eramo and Nobili (2011) – henceforth DEN (2011) – which analyzed the relative contribution of demand and supply factors in credit market dynamics in the first part of the financial crisis, up to the immediate aftermath of the Lehman collapse. With respect to DEN (2011), we use a longer sample period, which also includes the sovereign debt crisis, thus allowing us to compare the relative importance of demand and supply factors in the two phases of the crisis. In addition, our paper provides an original contribution in several other respects.

First, we significantly improve the methodology. The empirical strategy consists (for each credit market segment considered in this analysis) of a two simultaneous equation model, where the dependent variables are the growth rate of loans and the change in interest rate on such loans and the various BLS indicators are used as instruments to identify a credit demand and a credit supply curve. As compared to the reduced-form approach in DEN (2011), we provide a deeper structural interpretation of the estimated relationships between hard variables and BLS indicators. In particular, our empirical model is able to describe a standard imperfect competition framework in credit markets, in which the intermediaries set loan interest rates and fully accommodate the credit demand (see the survey of theoretical and empirical contributions in Freixas and Rochet, 2008 and Degryse, Kim and Ongena, 2009). In addition, the statistical model also nests a credit rationing framework, in the spirit of Jaffee and Modigliani (1969) and following the disequilibrium approach proposed by Fair and Jaffee (1972). In this regard, we consistently estimate the relative contribution to lending

[2] See Berg et al. (2005), de Bondt et al. (2010), Hempell (2004), Hempell and Kok Sorensen (2009) and Ciccarelli, Maddaloni and Peydró (2009) for analyses on the euro area based on the BLS data, and Lown, Morgan and Rohatgi (2000) and Lown and Morgan (2006) for analyses on the U.S. based on data from the Senior Loan Officer Opinion Survey carried out by the Federal Reserve System.

dynamics of the indirect effects of supply restrictions (i.e., via the elasticity of loan demand to the loan rates) *versus* the direct ("credit-rationing") effects stemming from non-price allocations of credit. This question cannot be properly addressed using a reduced-form approach due to the endogeneity of the cost of credit as regressor in a loan quantity equation and to a "simultaneity bias" between the BLS supply indicators and the cost of credit.

During the crisis, policymakers have been particularly interested in this issue, since different types of supply restriction may call for different policy responses, i.e., official interest rate changes versus the implementation of unconventional measures. To the extent that monetary policy operates through a "credit channel" (in which contractionary policy affects the economy through a decline in the supply of funds available for banks to lend), and to the extent that changes in the terms of lending include not only changes in loan pricing but also changes in the quantities of credit available to borrowers, credit rationing may play an important role in the transmission of monetary policy effects on the economy (Blinder and Stiglitz, 1983). Evidence that some rationing occurred during the crisis is provided by both the banks' answers to the BLS and firms' replies in business surveys concerning the different ways in which they experience difficulties in obtaining bank's credit (cost vs. quantity). Practical considerations also suggest that a tightening in credit standards induced by "pure-supply" factors, such as a worsening in banks' capital position or in their access to funding conditions, can be more likely interpreted as the evidence of a "credit crunch" affecting the economy.

Second, concerning the sovereign debt crisis, we also studied whether the inclusion of the spread between the yields on 10-year Italian and German government bonds – often regarded as a sort of "sufficient statistics" to measure the intensity of tensions – has marginal predictive content over the BLS indicators and helps explaining the evolution of the cost and dynamics of bank lending. In this regard, we provide alternative and complementary evidence on the effects of the sovereign debt crisis on credit in Italy with respect to other studies carried out on macroeconomic data (Albertazzi et al., 2012; Neri, 2012) or based on bank-firm relationship information (Albertazzi and Bottero, 2012; Bofondi et al., 2012).

Third, we test the robustness of our identification of loan demand and supply curves by using the BLS answers on the specific "terms and conditions" through which banks reported to have implemented changes in their credit standards (distinguishing, in particular, between price and non-price conditions). Estimating the same equations considered above by including these BLS supply indicators as regressors instead of factors provides a particularly valuable test, since the various terms and conditions appear to be much more clearly related to price and non-price allocations of credit than factors. Obtaining similar results with this alternative source of information – in particular as regards the estimated elasticity of loan demand and of credit rationing– would be a valid indication in support of our identification.

The rest of the paper is organized as follows. Section 2 describes the data and presents descriptive evidence. Section 3 provides new evidence on the information content of BLS indicators using standard reduced form equations. Section 4 illustrates the methodology used for the identification of loan demand and supply curves and discusses the main findings for a baseline specification – where the BLS supply and demand indicators are used as the main explanatory variables. Section 5 presents extended specifications that also include the sovereign debt spread as additional explanatory variable, while Section 6 discusses the results obtained by replacing the BLS supply indicators related to the factors behind credit standards with the BLS indicators related to the terms and conditions. Section 7 illustrates the counterfactual exercises carried out to assess the relative importance of demand and supply factors over the crisis and to compare the effects during the sovereign debt crisis with those observed in the global phase of the crisis. Section 8 offers some concluding remarks.

2. BLS indicators and lending to enterprises: data and descriptive evidence

This section provides information on the data used in the paper and some descriptive statistics. We carry out the study on data for the panel of Italian banking groups (henceforth "banks") participating in the BLS, which are among the largest in the country. The effects of mergers, which over time had tended to reduce their number, has been offset by subsequent additions. As a result, the dataset consists of an unbalanced panel of 11 Italian banks involved in the survey (with a maximum of 8 banks per quarter, including the more recent period) over a sample period of 39 quarters (from the fourth quarter of 2002 to the second quarter of 2012), providing a total of 287 observations.[3] For loans to enterprises and mortgage loans to households, the outstanding amounts granted by the banks participating in

[3] Merger and acquisitions that involved the banks participating in the survey over the sample period were carefully addressed. They were treated by using the standard reclassification methods in the computation of the lending growth rate for the acquirer, which is included in the panel over the entire sample period, while the target bank is excluded since the date of the operation (consistently with the treatment of individual bank data in the BLS). We checked the robustness of the results by also using a different approach, in which both of the banks involved are excluded from the panel since the date of the operation, and a new bank is included afterwards. The results do not change.

the survey corresponded at the end of the sample period to, respectively, around 60 and 63 percent of the total provided by the whole Italian banking system.[4]

Tables 1 and 2 show descriptive statistics for the indicators of supply and demand conditions provided by the Italian component of the BLS for, respectively, lending to enterprises and to households for house purchase. They are reported for the pre-crisis period (2002Q4–2007Q2), the crisis period (2007Q3–2012Q2), and, within the latter, the "global crisis" (2007Q3–2010Q1) and the most recent "sovereign debt crisis" (2010Q2-2012Q2). The tables report the frequency of individual banks' answers concerning supply conditions and their assessments of demand developments; all answers refer to the changes with respect to the previous three months.[5]

Concerning lending to enterprises, in the pre-crisis period, 80 percent of individual banks' responses on supply conditions fell in the "unchanged" category. Answers reporting that supply conditions had eased (either considerably or somewhat) were almost absent. Less than one fifth of responses indicated "tightened somewhat", while very few indicated "tightened considerably". In the crisis period the percentage of answers falling in the "tightened" category rose considerably, to 37 and 29 percent, respectively, in the two phases of the crisis. As to the demand assessments, extreme answers were virtually absent over the whole sample period. The frequency of responses indicating a "decrease" more than doubled in the crisis period, to 19 and 28 percent, respectively, in the two phases. Similar developments were observed for the answers concerning mortgage loans to households.

Figures 1 and 2 provide descriptive evidence on the relationship between the evolution of the BLS indicators of supply and demand conditions and that of, respectively, the dynamics and the cost of loans to both enterprises and households for house purchases.

Figure 1 shows that lending to enterprises recorded two phases of sharp slowdown: during the 2008-09 global crises and in the most recent sovereign debt crisis. In both cases the slowdown in lending went along with a fall in the BLS demand indicator and a tightening of the BLS indicator of supply conditions, the latter being particularly strong in the last two quarters of 2008 and in the last quarter of 2011, when all or almost all the

[4] The pattern of loan dynamics for the banks in the BLS panel is similar to that for the system as a whole, although the rate of growth is on average lower over of the sample period.

[5] Banks are asked the following question concerning supply conditions: "Over the past three months, how have your bank's credit standards as applied to the approval of loans or credit lines to enterprises changed?". As to demand conditions, the question is: "Over the past three months, how has the demand for loans or credit lines to enterprises changed at your bank, apart from normal seasonal fluctuations?". In both cases, they can choose their answer among five options, as reported in Tables 1 and 2.

participating banks reported a tightening. Figure 2 shows that the two phases of strongest tensions were also characterized by a strong rise of the average cost of credit to both enterprises and households, associated to the tightening of supply conditions.

The BLS supply indicator reported in Figures 1 and 2 refers to the change in the overall supply conditions reported by the banks. In the survey, however, banks are also asked to respond to more detailed questions concerning the importance of the various factors determining the changes in their supply policy, differentiating between: i) “cost of funds and balance sheet constraints” (in the case of loans to enterprises with a further distinction between “costs related to bank’s capital position”, “banks’ ability to access market financing” and “bank’s liquidity position”); ii) “pressure from competition”; iii) “perception of risk” (in turn relating to “expectations regarding general economic activity” or to more specific factors: in the case of loans to enterprises, “industry or firm-specific outlook” and “risk on collateral demanded”; for mortgage loans to households, “housing market prospects”).

Figure 3, based on the answers to these questions, shows that the relative importance of the factors affecting credit standards has been different during, respectively, the 2008-09 global crisis and the 2011 sovereign debt crisis. In the former period, the tightening mostly reflected an increase in perception of risk, while the role of banks’ cost of funds and balance sheet constraints was limited. In the latter period, risk perceptions have again played a role, but an even larger role has been played by the bank’s difficulties in obtaining market financing and by their liquidity position. This is a relevant difference, which we take into account in designing the empirical exercises of the following sections.

A general caveat, which applies to our study, as to any other analysis based on a survey, is that the quality of the results depends on both the reliability and the truthfulness of the respondents’ answers. In the case of lending surveys, on the one hand, banks may be inclined to report tighter credit standards than those actually applied. This hypothesis originates from the observation that indications of ‘‘tightening’’ have historically outnumbered those of ‘‘easing’’; in addition, banks may have an incentive to report tighter policies if they fear that the information could be exploited for supervisory purposes. On the other hand, during the crisis banks were exposed to public criticism and political pressure, being regarded as responsible for a credit crunch, and thus might have had an incentive to portray their policies as less restrictive.

3. The reduced-form representation of the credit market

3.1 Specification and methodological issues

In this section we sketch the methodology we use for the econometric analysis. The starting point is an extension of the approach proposed by DEN (2011), in which the information content of BLS indicators can be assessed by estimating reduced-form regressions in which the variables of interest, namely the growth rate of loans and the cost of credit, are regressed on BLS demand and BLS supply factors, as follows:

$$(1)\ \Delta bankspread_{it} = \alpha_{1i} + \gamma_1(L)\Delta bankspread_{it} + \beta_1(L)BLS\_S_{it} + \lambda_1(L)BLS\_D_{it} + \varepsilon_{it}$$

$$(2)\ \Delta loans_{it} = \alpha_{2i} + \gamma_2(L)\Delta loans_{it} + \beta_2(L)BLS\_S_{it} + \lambda_2(L)BLS\_D_{it} + \eta_{it}\ .$$

The variables $\Delta bankspread_{it}$ and $\Delta loans_{it}$ are, respectively, the first difference of the spread between bank *i* average rate on new loans and the Eonia rate[6] in quarter *t* and the quarter-on-quarter (henceforth q-o-q) rate of growth in bank lending for the same bank in the same quarter. BLS_S_{it} is a vector of credit supply indicators based on the bank *i*'s BLS answers concerning the influence that the various factors (cost of funds and balance sheet constraints, and risk perceptions) have had on its decisions on supply policy. The variable BLS_D_{it} is the overall indicator of credit demand conditions based on the banks' assessment.

In the equations we chose to separately include supply factors considered in the BLS, rather than the overall indicator of supply conditions – which has been used in most works based on the BLS – for a number of important reasons. First, it is important to distinguish whether banks modified their credit standards as a result of changes in their own conditions (balance sheet constraints, ability to access market financing) or instead in reaction to the risks connected with economic developments and borrowers' creditworthiness for a number of reasons. This distinction has indeed significant policy implications, since both the effects of the supply tightening on credit dynamics and the appropriate policy reactions can depend on the factors driving it. Factors belonging to the first group can unambiguously be interpreted as "pure" supply factors, while the case is less clear for the second group. A more prudent attitude on the part of banks may in fact reflect a reduction in banks' ability or willingness to assess borrowers' creditworthiness properly, or an increase in banks' risk aversion beyond what is warranted by economic developments; but it may also be the proper

[6] We consider the difference between the bank loan rate and the Eonia rate in order to rule out the effects of monetary policy.

reaction to a worsening in the borrowers' creditworthiness. Therefore, only in the former case it can be characterized as a "pure" supply factor.[7]

Second, DEN (2011) showed that the indicators for the specific factors can be more informative in certain phases. In the BLS, in fact, there is not always a clear correspondence between the banks' answers to the general question on the changes in their credit standards and the replies concerning the factors behind these changes. In particular, there are cases in which a bank signals no change in its own overall supply policy but reports that a specific factor has contributed to a change in its credit standards. This suggests that the banks' replies to the other specific questions concerning credit supply are not always "conditional" on their answers on the general question concerning their overall supply policy (although this is what the formulation of the questionnaire would imply).[8]

Since the BLS indicators are qualitative variables, BLS_S_{it} *and* BLS_D_{it} are defined as vectors of dummy variables, each of which corresponds to one of the possible alternative answers in the survey.[9] As shown by DEN (2011), this choice helps capturing non-linearity in the estimated relations between endogenous and exogenous variables, which may be particularly relevant in the case of the BLS supply indicators. Both the supply and the demand indicators may enter with the contemporaneous and/or lagged values; the lag order for each of them is chosen by trying a range between 0 and 4, on the basis of the fit of the regression and the indications derived from standard information criterion.

The equations are further enriched along several dimensions.[10] First, we consider dynamic models by including autoregressive components, if statistically significant. Second, we also add seasonal dummies and bank-specific fixed effects. The latter may be crucial to

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[7] We follow DEN (2011) in using the expression "pure" supply rather than the expression "credit crunch" since, as argued convincingly in that paper, there is no universally accepted definition of "credit crunch".

[8] Banks are asked the following question: "Over the past three months, how have the following factors affected your bank's credit standards as applied to the approval of loans or credit lines to enterprises (as described in question 1)?", where "question 1" is the general question concerning supply conditions (see footnote 6 in Section 2).

[9] It is worth noting that an alternative could be to include the cumulated levels of the BLS indicators, rather than the indicators themselves. As remarked by DEN (2011), this definition would be more consistent with a literal reading of the BLS questions and answers; however, the robustness analysis carried out in that paper shows that the inclusion of the cumulated indicators provides unclear results or worsens the fit of the estimates (depending on the approach), which argues against following this alternative specification.

[10] We also explored specifications aimed at investigating whether the effects found for the entire sample period were magnified during the crisis or during specific phases of tensions. To this purpose, we included in our equations a Crisis_dummy variable or, alternatively, a Lehman_dummy variable and a Sovereign_dummy variable – which take value 1, respectively, from 2008Q3 to 2012Q2, from 2008Q3 to 2010Q1 and from 2010Q2 onwards –, as well as the respective interaction terms with the BLS supply factors. However, none of the estimated coefficients of the interaction terms proved to be statistically significant.

control for unobserved bank-specific factors that may be correlated with the BLS variables, thus leading to inconsistent estimated coefficients. In this regard, banks could interpret the qualitative BLS questions in different ways or make systematic mistakes in their answers.

For the coefficients of equations (1) and (2) we report the SURE estimations in order to take into account the correlation between the residuals of the regressions and to gain in efficiency with respect to OLS. By construction, the unobserved panel-level effects may be correlated with lagged dependent variables, making standard estimators inconsistent. To test the robustness our estimates, we also report the coefficients using the generalized method of moments (GMM) estimator designed by Arellano and Bond (1991) for linear dynamic panel-data models. In this case, we perform single-equation regressions and the test statistics for the significance of the coefficients are based on robust standard errors.

3.2 The results for loans to enterprises

Panel (a) of Table 3 reports the results of the econometric analysis for the dynamic and the cost of loans to enterprises. In the loan spread equation, a worsening in banks' access to funding, as captured by the BLS answers concerning this specific factor, exerts a statistically significant effect on the cost of credit on impact and with a one-quarter lag: responses of a tightening related to this factor by all banks in the panel is associated with an increase in the bank mark-up by about 45 and 18 basis points, respectively, with respect to the spread that would have been observed had all banks signalled no such effect on credit standards. A supply tightening related to risk perception also exerted a significant effect: responses of a tightening related to this factor by all banks would be associated to a contemporaneous increase in the loan spread of 55 basis points when it was reported to have contributed with the "considerably" qualification to a tightening and of around 10 basis points when it was reported to have contributed with the "somehow" qualification to a tightening. Neither the BLS supply factor related to banks' capital position nor the BLS demand indicator resulted to be statistically significant in this equation. The coefficients of the lagged values of the loan spread are negative and highly significant. This likely results from the fact that the effects of the reductions of the spread observed over the sample period are not properly captured by the BLS supply indicators, which are characterized by a strong asymmetry (in several periods they point to a tightening of credit standards, never or very rarely, depending on the indicator, to an easing, a feature that the BLS shares with other similar surveys).

In the reduced-form equation for loan quantities, the BLS demand indicator enters significantly and with the expected sign. We do no find evidence of non-linearity in the estimated relationship. Even if distinguishing between responses of an “increase in demand” and “a decrease in demand”, the resulting estimated coefficients are opposite in sign but the same in magnitude: responses of an increase (decrease) in demand by all banks are related to an immediate increase (decrease) in credit growth by about 0.9 percentage points. The BLS supply indicators are also highly significant: a tightening in credit standards related to banks’ funding conditions, capital position or risk perception with the “considerably” qualification would be associated to an immediate decline in loan growth rate by about 1.1-1.3 percentage points. The negative effect is not statistically significant if the credit supply restriction is related to banks’ risk perception with the “somehow” qualification.[11]

3.3 The results for mortgage loans to households

Panel (b) of Table 3 reports the results of the econometric analysis for the dynamic and the cost of mortgage loans to households. In the loan-spread equation, the dependent variable is the difference between a weighted average of a mortgage rate on new fixed-rate contracts and that on new floating-rate contracts and the Eonia rate. Since the mortgage rate on fixed-rate contracts reflects mostly changes in long-term market risk-free rates rather than the policy rate, we also included in the regression the 10-year swap rate as additional regressor. The coefficient resulted to be negative and statistically significant, since long-term rates adjust less to policy rate changes.

As for the BLS indicators, a change in supply conditions related to the banks’ cost of funds and balance sheet conditions – they are considered as single factor in the BLS question concerning supply conditions for mortgage loans – exerts a positive effect on the cost of credit and a negative one on loan quantity with a one quarter lag. The estimated coefficients are about one half in magnitude than those recorded for BLS “pure-supply” factors in the regressions for loans to enterprises. In the case on loan quantity, they are not statistically significant. There is evidence, therefore, that the effects of the “pure-supply” factors transmit more rapidly and with a greater intensity to credit standards granted on loans to firms than on mortgage loans to households. A supply tightening due to a higher risk perception[12] is

[11] We also examined whether the inclusion of macro variables usually used in reduced-form equation for loans to Italian firms affects our identification scheme, just to be sure that there is no bias in the estimated coefficients due to omitted variables. Following Casolaro et al. (2006) and Albertazzi et al. (2012) we consider nominal GDP and firms’ financing needs for loans to enterprises. Neither the nominal GDP nor firms’ financing needs resulted to be statistically significant, meaning that the bank-specific BLS demand indicators dominates the aggregate variables. A similar result has been also found by DEN (2011).

[12] As for the indicator on “perception of risk”, in the case of mortgage loans to households, banks’ answers may refer to their expectations on general economic developments and housing market prospects. Similarly to the case of loans to enterprises, we collapsed all this information into just one variable. We also carried out a

associated to an immediate rise in the loan spread and to a decline in quantity, when it has been reported by banks with the "considerably" qualification. The estimated effect on the cost of credit is comparable in magnitude to that for loans to enterprises; the effect on loan quantity is instead much stronger. Risk perception considerations reported by banks with the "somehow" qualification seem to exert a significant effect only on the cost of credit.

The BLS demand indicator enters the loan quantity equation in a non-linear way: it is marginally significant both contemporaneously and with a one-quarter lag only when banks reported that they have experienced a "decrease" in demand. The estimated coefficients suggest a weaker effect with respect to loans to enterprises. Surprisingly, a decrease in the BLS demand is also significant in the loan spread equation but with the wrong sign, since a decrease in demand is associated to an increase in the loan spread. One possibility is that in some periods the BLS demand indicator may capture risk considerations that instead should be reported by banks in their answers to supply factors behind changes in credit standards. Figure 3 shows that, during the sovereign debt crisis, only very few banks reported that the risk perception related to the housing market prospects have contributed to a tightening in credit standards.[13] Finally, in the loan quantity equation, the lagged values of the dependent variable are highly significant reflecting the strong persistence in the dynamics of mortgage loans to households.

4. The structural representation of the credit market

4.1 Identification strategy in a partial equilibrium framework

The reduced-form approach, even if informative, cannot provide a deeper structural interpretation of the estimated relationships between the "hard" variables and the BLS indicators. During a financial crisis policymakers have been particularly interested in assessing whether the supply restriction affected lending dynamics through their increase in the cost of credit (i.e., via the elasticity of loan demand to the bank mark-up) or through a

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regression including the BLS supply factor related to competition among banks or non-banks; the coefficient, however, did not turn out to be statistically significant.

[13] We checked the robustness of our results by including aggregate variables related to the housing market dynamics. We considered the quarterly change in house prices, a variable that was often used in empirical studies with aggregate data in the case of Italy (see Casolaro and Gambacorta, 2005; Albertazzi et al., 2012; Nobili and Zollino, 2012). Alternatively, we used the quarterly change in housing transactions, as they usually react more promptly than prices to shocks in the economy and that was found to perform relatively better. These variables enter significantly the estimated regressions for loan quantity and tend to offset the BLS demand indicator, while leaving unchanged the coefficients of the BLS supply indicators. On the contrary, they have no significant effects on the cost of credit.

direct "credit-rationing". Different supply restrictions call for different policy responses, as, for example, changes in official interest rates vs. unconventional monetary policy measures.

A structural representation of the reduced-form equations can be described by the following system of simultaneous equations, in which the cost and the growth rate of lending become the endogenous variables linked by a two-way causality and expressed as functions of all the exogenous and predetermined variables:

(3) $\Delta bankspread_{it} = \alpha_{1i} + \gamma_1(L)\Delta bankspread_{it} + \beta_1(L)BLS\_S_{it} + \lambda_1(L)BLS\_D_{it} + \theta \cdot \Delta loans_{it} + \varepsilon_{it}$

(4) $\Delta loans_{it} = \alpha_{2i} + \gamma_2(L)\Delta loans_{it} + \beta_2(L)BLS\_S_{it} + \lambda_2(L)BLS\_D_{it} + \rho \cdot \Delta bankspread_{it} + \eta_{it}$.

This system of equations cannot be estimated using OLS, since it is not identified. In this paper identification is reached using exclusion restrictions on both exogenous regressors (the BLS indicators) and predetermined variables (the lagged values of the endogenous variables). The necessary and sufficient condition for identification of each equation is the *rank condition*, which states that the matrix of coefficients for the set of excluded variables from one equation must have full row rank in the other equation. A simpler way to think about the identification is in terms of the instrumental variable approach: an equation is identified if and only if there are enough instruments for the right-hand-side endogenous variables that are fully correlated with these variables. In general, exclusion restrictions are only necessary for identification (they satisfy the *order condition*), but in a two-equation system they also satisfy the rank condition (see, e.g., Wooldridge, 2002).

In our framework, the BLS indicators may represent an appropriate solution to the identification problem. Indeed, a reasonable strategy is to assume that the BLS demand indicators are excluded from equation (3), while the BLS supply factors are excluded from equation (4). More specifically, the system of equations would become:

(5) $\Delta bankspread_{it} = \alpha_{1i} + \gamma_1(L)\Delta bankspread_{it} + \beta_1(L)BLS\_S_{it} + \theta \cdot \Delta loans_{it} + \mu_{it}^S$

(6) $\Delta loans_{it} = \alpha_{2i} + \gamma_2(L)\Delta loans_{it} + \lambda_2(L)BLS\_D_{it} + \rho \cdot \Delta bankspread_{it} + \mu_{it}^D$.

The coefficients of the structural equations (5) and (6) can be estimated consistently and efficiently using the 3-Stage Least Squares (3SLS) estimator, which takes into account the endogeneity among the dependent variables, as well as the correlation between the estimated residuals of the two equations. If BLS_S_{it} and BLS_D_{it} are statistically significant (e.g. they are reliable instruments) and θ and ρ are, respectively, positive and negative, identification is reached. Accordingly, equation (5) can be interpreted as a credit supply curve where the bank mark-up on the monetary policy rate may vary reflecting the banks' balance sheet conditions and their perception of the borrowers' riskiness (these terms are captured by the BLS supply indicators), which act as credit supply "shifters". A tightening in credit standards implies an increase in banks' margins and a decline in loan growth rate via the elasticity of loan demand (e.g. the coefficient ρ). Equation (6) would instead represent the credit demand curve, where loan quantity depends on a demand shifter and negatively on the cost of credit. A downward (upward) shift in credit demand, as captured by the BLS indicator, leads to a reduction (increase) in both the loan growth rate and the bank mark-up via the elasticity of the loan supply (the coefficient θ).

A special case of this structural model occurs for θ=0, which is consistent with the widely used representation of the credit market in an imperfect competition framework. Accordingly, the credit supply is flat and the intermediaries set loan interest rates and fully accommodate the credit demand (Freixas and Rochet, 2008; Degryse, Kim and Ongena, 2009).[14] In this theoretical framework, a shift in credit demand would affect quantities while leaving unchanged the cost of credit. This assumption could be particularly debatable during a crisis when funding becomes sluggish and costly for banks, thus inducing the latter to accommodate an increase in demand by raising the mark-up. The distinction between a flat versus an upward sloping credit supply curve is, therefore, tested empirically in this paper.

Notice that actually we may deal with *over-identified* equations, in which the number of excluded instruments exceeds the number of endogenous variables. Precisely, in equation (5) we use not one but a set of BLS supply indicators (the various factors behind changes in credit standards) and, for some of them, different variables capturing the various categories of answers (for example, the qualification of "contributed considerably to a tightening" and "contributed somehow to a tightening"). This implies that the loan demand equation is over-identified. Similarly, in equation (6) we may potentially include the BLS demand indicator as distinguished between the "increase" and "decrease" qualification, thus implying that the loan supply equation is also over-identified. We discuss the reliability of the over-identified restrictions by reporting the Sargan-Hansen test for each structural equation. The joint null hypothesis tested is that the instruments are correctly excluded from the structural equation to be identified. Under the null hypothesis the test statistic is distributed as chi-squared in the number of over-identifying restrictions.

We also address the issue of "weak identification". When the excluded instruments are weakly correlated with the endogenous variables, the estimates may be not consistent, tests

[14] See Panetta and Signoretti (2010) for a simple illustration of this theoretical framework and its possible use to interpret credit developments during the global crisis.

of significance have large size distortions and the confidence intervals are wrong: the estimated variance of the estimator tend to be biased downward in finite samples and the bias become large when the instruments are weak, thus tending to reject too often the null hypothesis of a zero coefficient. Staiger and Stock (1997) formalized the definition of “weak instruments” and argued that if the F-statistic on the excluded instruments in the first stage regression is greater than 10, one need worry no further about weak instruments.

4.1 Identification strategy in a disequilibrium framework

The system based on equations (5) and (6) represents an empirical framework that is useful to describe a partial equilibrium model, in which the credit market clears continuously and the interest rate changes ensure that the supplied quantity equals the demanded quantity at each point in time. An important limitation is the inability of the system to capture “credit rationing” episodes. Broadly speaking, credit rationing occurs when lenders limit the supply of credit to borrowers, even if the latter are willing to pay higher margins. In the spirit of Jaffee and Modigliani (1969), “credit rationing” is a situation in which the demand for loans exceeds the supply of loans at the loan rate quoted by the banks. Key to this definition is that changes in the interest rate cannot be used to clear excess demand for loans in the market. In essence, this definition treats credit rationing as a supply side phenomenon, with the lender’s supply function becoming perfectly price inelastic at some point.

The seminal theory developed by Stiglitz and Weiss (1981), however, made a proper distinction between a situation in which a lender eventually restricts the size of loan to any potential individual borrower and one in which lenders fully fund some borrowers but deny loans to others, because of the presence of asymmetric information between lenders and borrowers. Banks may not raise lending rates above a certain level to avoid financing more risky borrowers (adverse selection) or to discourage firms to take more risk (moral hazard).[15] Albeit there are two main working definitions of “credit rationing”, it is more useful for the purpose of our analysis to consider a broader definition of “credit rationing”, in which other phenomena, such as regulatory constraints (for example, liquidity and capital requirements) or banks’ inability to access to market funding[16,] in addition to informational problems, lead to non-price allocations of credit.

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[15] See Freixas and Rochet (2008) for a survey of the theoretical contributions on this issue.

[16] Seminal theories may link “credit rationing” to a worsening in banks’ funding conditions. The experience of the Great Depression in the US suggests that banking crisis principally arose due to depositors’ panic, which caused a run on the banks. The source of a bank-run may emerge from liquidity shocks (Diamond and Dybvig,

In most recent econometric analyses, the authors followed the seminal disequilibrium approach for macroeconomics developed by Fair and Jaffee (1972) to assess loan dynamics and to identify credit crunch episodes in the aftermath of financial crises for a number of countries (Pazarbasioglu, 1996; Ghosh and Ghosh, 1999; Kim, 1999; Barajas and Steiner, 2002; Ikhide, 2003; Baek, 2005; Bauwens and Lubrano, 2007; Allain and Oulidi, 2009). These statistical models relied on the voluntary exchange principle, namely that the observed traded quantity in a specific good market is determined by a short-side rule, i.e. by the minimum of supplied and demanded quantity. The basic disequilibrium approach can be described by the following equations:

(7) $Q_t^S = D(X_t^S, p_t) + \mu_{it}^S$

(8) $Q_t^D = D(X_t^D, p_t) + \mu_{it}^D$

(9) $Q_t = min(Q_t^S, Q_t^D),$

where Q_t is the observed traded quantity, p_t is the price level, X_t^S, X_t^D vectors of exogenous variables. In this model the price level is assumed to be exogenous and no prior information about the excess demand state of the credit market is available, meaning that we not known the periods when loan quantity lies on the demand curve and the periods when it lies on the supply curve. Under the additional assumption that the error terms in equations (7) and (8) are uncorrelated and normally distributed, several maximum-likelihood estimation methods have been developed in order to provide the probabilities that each observation belongs to the supply or demand regime and the estimates of the structural parameters. [17]

However, the maximum likelihood estimation of models where the sample separation is unknown often leads to the likelihood function being unbounded in parameter space, which results in the computation procedure breaking down.[18] Moreover, without information about the interest rate adjustment the equilibrium set-up and the disequilibrium set-up will be two non-nested models (Quandt, 1978) and it would not be feasible to perform a statistical test to discriminate among the two cases. To overcome the unboundedness of the likelihood function problem, some authors employed limited maximum likelihood estimates where the unbounded regions of the parameter space are avoided by simply assuming that the change

1983) or shocks to the banks' asset value (Calomiris and Khan, 1991) in a theoretical framework of information asymmetry between banks and depositors.

[17] See Amemiya (1974), Fair and Kelejian (1974), Maddala and Nelson (1974), Goldfeld and Quandt (1975), Bauwens and Lubrano (2007).

[18] As argued by Maddala (1983), when it is not known which observations are on the demand function and which are on the supply function, too much may be asked of the data. Monte Carlo methods found that there is considerable loss of information if sample separation is not known.

in the price level is linearly related to the size of excess demand. The basic disequilibrium model is, therefore, extended by specifying a price adjustment equation (see Fair and Jaffee, 1972; Laffont and Garcia, 1977; Bowden, 1978) as follows[19]:

(10) $p_t - p_{t-1} = \gamma(Q_t^D - Q_t^S)$,

In analysis for the credit market, this implies the inclusion of a dynamic equation for the evolution of the interest rates that become an endogenous variable in the system (see Laffont and Garcia, 1977; Ito and Ueda, 1981). While not properly rooted in theoretical considerations, a price equation makes the statistical model more tractable and allows for a disequilibrium model to encompass an equilibrium one (see Quandt, 1978).

In this paper, we essentially rely on the quantitative approach proposed by Fair and Jaffee (1972) but use the BLS information for the identification of the periods characterized by excess demand or excess supply because. The model of equations (7) through (10) is, indeed, only one of possible way in which disequilibrium might be modelled. As already mentioned, we aim at capturing non-price allocation of credit. In particular, we assume that the credit markets may be characterized by excess demand when a bank reported a tightening in credit standards. Correspondingly, we consider as periods of excess supply those specific quarters in which a bank reported an easing in credit standards. Accordingly, the system is defined by the following equations:

(11) $\Delta bankspread_{it} = a_{1i} + \theta \cdot \Delta loans_{it}^S + \beta(L)BLS\_S_{it} + \mu_{it}^S$

(12) $\Delta loans_{it}^D = a_{2i} + \rho \cdot \Delta bankspread_{it} + \lambda(L)BLS\_D_{it} + \mu_{it}^D$

(13) $\Delta loans_{it}^D - \Delta loans_{it}^S = \sigma_1(L)BLS\_S\_tightening_{it}$

(14) $\Delta loans_{it}^S - \Delta loans_{it}^D = \sigma_2(L)BLS\_S\_easing_{it}$

(15) $\Delta loans_{it} = min(\Delta loans_{it}^S , \Delta loans_{it}^D)$.

In the system, $\Delta loans_{it}^S$ and $\Delta loans_{it}^D$ are the demanded and supplied quantities, which are not observable, while $\Delta loans_{it}$ are observed traded quantities. Notice that we are dealing with a linear framework with stationary variables. In general, the disequilibrium model of equations (7) through (10) reflects a valid theory only if variables are expressed in levels. In the empirical literature, the issue of non-stationarity is barely explicitly addressed and the

[19] Fair and Jaffee (1972) assumes that the change in the price level is directly proportional to the difference between demand and supply. Laffont and Garcia (1977) suggested a price-setting rule allowing for different downward and upward adjustment speeds. Bowden (1978) used a partial-adjustment scheme for the price level dynamics.

significant relationships may arise from spurious regressions[20]. With our dataset, it would be hard to defend a co-integration framework between hard quantitative variables for the credit market and qualitative survey data.[21] One may argue that imposing a minimum condition on variables in quarterly growth rates may lack of proper theoretical foundations, since the quarterly growth rate of credit demand may exceeds the growth rate of credit supply not necessarily when the credit supply level is the binding constraint.[22] In our specific case, since the disequilibrium indicators are the BLS supply variables, the quarterly growth rate of credit demand exceeds the growth rate of credit supply as a result of a binding constraint in the credit supply level.

Following the discussion in Fair and Jaffee (1972), the system of equations (11)-(15) can be reduced to a system with a single demand and a single supply equation, as follows:

(16) $\Delta bankspread_{it} = a_{1i} + \theta \cdot \Delta loans_{it} + \beta(L)BLS\_S_{it} - \sigma_2(L)BLS\_S\_easing_{it} + \mu_{it}^S$

(17) $\Delta loans_{it} = a_{2i} + \rho \cdot \Delta bankspread_{it} + \lambda(L)BLS\_D_{it} - \sigma_1(L)BLS\_S\_tightening_{it} + \mu_{it}^D$.

The system of equations (16) and (17) can be estimated consistently and efficiently over the entire sample period using a 3SLS approach. Notice that the interest rate-setting mechanism operates in each period but it does not necessarily clears the market. According to this framework, at a given point in time the credit market can exhibit temporary credit rationing owing to imperfect flexibility in the interest rates. The usual test of the statistical significance of the coefficients $\sigma_1(L)$ and $\sigma_2(L)$ can be, indeed, interpreted as a direct test for the presence of credit rationing in the credit market at each point in time: if these coefficients are statistically significant, changes in supply conditions exert a direct effect on lending dynamics beyond that occurring via the changes in interest rates. Since banks rarely reported an easing in credit standards, our model will be essentially a test of whether the credit market is in equilibrium or in a regime of excess demand. Our estimates will provide evidence about which of the different BLS supply factors capture credit rationing.

[20] Exceptions are in Ghosh and Ghosh (1999) and Allain and Oulidi (2009).

[21] One possibility, however, would be to specify the system in the levels of "hard" credit variables and using the cumulated version of the BLS indicators (see DEN, 2011).

[22] If the true data generating process of $Q(t)$ is such that $Q(t)=\min(S(t), D(t))$, the first-differences series $\Delta Q(t)$ follows a four-regime dynamics where there are two regimes characterized by a demand higher (lower) than supply at two points in time (t and $t-1$) and two regimes of switch from demand to supply and viceversa.

4.2 The market for loans to enterprises

Table 4 reports the results of the econometric analysis for the dynamic and the cost of loans to enterprises. We begin the discussion with the system specification reported in columns (a) and (a'), which represents a credit market equilibrium with no credit rationing.

Accordingly, all the BLS supply factors enter as instruments the loan supply equation (see column (a)), while the BLS demand indicators and the lagged values of loan quantities enter as instruments the loan demand equation (see column (b)). The estimated effects of the BLS supply variables on the cost of credit are remarkably very similar to those obtained with the reduced form equation. The loan growth rate enters the supply equation with a positive coefficient but it is not statistically significant, thus suggesting a flat credit supply curve. In the demand equation, the BLS demand indicators enter significantly and with the expected sign; the coefficient for the loan spread is highly significant and negative, suggesting that we are correctly identifying a downward sloping credit demand curve. The estimated loan demand elasticity to the bank mark-up is high: a 100 basis points increase in the cost of credit would lead to a reduction in the loan growth rate by more than 2 percentage points.

According to this structural representation, a tightening in credit standards related to banks' funding conditions would be associated to a decline in the loan growth rate by about 1.0 and 0.5 percentage points on impact and with a one quarter lag. The lagged estimated effect is much lower than the one obtained with the reduced-form equations. In the case of the BLS indicator related to banks' capital position there would be no significant effect. As for the risk perception, the negative effect on loan growth rate is higher when reported with the "considerably" qualification while lower with the "somehow" qualification (it was null in the reduced-form equations). The various diagnostics suggest that the identification scheme is only partly satisfactory. In particular, the Sargan-Hansen test suggests that not all the exclusion restrictions imposed in the credit demand equation are valid. The "difference-in-Sargan" statistics[23], which allow a test of the exclusion restrictions on each instrument separately, suggest that both the BLS supply factor related to capital position and the lagged value of the BLS indicator of funding conditions could be included in the demand equation.

We now explore this alternative system specification. This structural representation of the market would be consistent with the presence of credit rationing. Results are reported in

[23] The statistics is defined as the difference of the Sargan-Hansen statistic of the equation with the smaller set of instruments (valid under both the null and alternative hypotheses) and the equation with the full set of instruments, i.e., including the instruments whose validity is suspect. Under the null hypothesis the statistic is distributed as chi-squared in one degree of freedom.

columns (b) and (b’). The coefficient for BLS supply indicators, indeed, turns out to be negative and highly significant: responses of a tightening related to these factors by all banks in the panel is associated with a decline in the loan growth rate by about 1.4 percentage points. The various diagnostics fully support this specification, including the Sargan-Hansen test that now definitively accepts all the over-identifying restrictions at any significance level. We notice that the slope of the loan demand curve reduces, while the outcome of a flat credit supply curve is still confirmed. Overall, this specification appears to capture credit market conditions both in normal times and in a crisis period, also providing evidence of a direct credit rationing on loans to enterprises.

One concern regarding our estimated coefficients is that the BLS indicator of funding conditions, being a dummy variable, may capture only partially the banks’ difficulties in obtaining funds. In addition, the Eonia rate used for the computation of the bank mark-up may also depend on other bank-specific variables not included in the equation. Angelini et al. (2010) and Affinito (2011) showed that the interbank rates at longer maturities faced by Italian banks depend significantly on the specific characteristics of borrowers and lenders and that some of the estimated relationships dramatically magnified after the breakout of the 2007-08 crisis. This might also be the case for overnight interbank rates. If there are omitted variables correlated with both the bank mark-up and the supply conditions, as captured by the BLS indicators, we could obtain biased estimates.

To investigate whether this is indeed the case, and possibly improve the estimation, we include the individual bank’s marginal cost of funding as an additional instrument in the mark-up equation. This variable is computed as the difference between the weighted average of the interest rates paid by the bank on its sources of funding (customer deposits and bank debt securities) and the Eonia rate, with the weights reflecting the relative importance of each type of liability. The estimated coefficients for this alternative system specification are reported in columns (c) and (c’). They indicate that a one percentage point increase in the marginal cost of funding is associated to a rise of the cost of credit of about 10 basis points. The effect measured by the BLS supply indicators related to funding conditions and risk perception remain highly significant, meaning that these variables have marginal information content over that of the marginal cost of funding. The estimated loan demand elasticity to the bank mark-up remains broadly unchanged, thus suggesting no relevant bias in our previous specification.

It is interesting to note that demand conditions, as captured by the BLS indicator, have no effect on the cost of credit, as a result of the flat credit supply curve, an outcome that is fully consistent with what we found in the reduced form equations. During a crisis, demand conditions may affect directly the bank mark-up when they reflect changes in the borrowers' composition (i.e. banks may face a demand for loans characterized by a larger fraction or riskier borrowers, thus inducing them to increase margins). This issue has been better addressed for Italy by recent works using bank-firm data from the Credit Register (see Albertazzi and Marchetti, 2010; Dimitri, Gobbi and Sette, 2010). We cannot exclude that the BLS risk perception indicator also captures a change in the borrowers' composition, since this variable is related not only to the general economic activity but also to the outlook for some specific sectors or firms.[24]

4.3 The market for mortgage loans to households

In Table 5 we present the estimated regressions for the cost and the growth rate of mortgage loans to households. Similarly to the strategy followed for loans to enterprises, we first investigate an identification scheme in which the various BLS supply factors enter as instruments the loan supply equation, while the BLS demand indicators enter as instrument the loan demand equation. The estimated coefficients are reported in columns (a) and (a').

The main difference with respect to the reduced-form equations is that the estimated effects of the various BLS supply factors become highly significant, while very similar in magnitude. This outcome reflects the gains in efficiency stemming from the 3SLS estimation with respect to the SURE estimation. The BLS demand indicators enter significantly the loan demand equation only with a one quarter lag and in a non-linear fashion: it is, indeed, highly significant only when banks reported to have experienced a "decrease" in loan demand.

The estimated loan demand elasticity to the bank spread is somehow lower to that obtained for loans to enterprises, albeit characterized by a higher level of uncertainty. In the loan supply equation the coefficient for the loan growth rate is not statistically different from zero, thus, suggesting a flat supply curve also for mortgage loans to households. The various diagnostic tests suggest that the identification scheme is overall satisfactory, but the Sargan-Hansen test rejects the exclusion restrictions imposed on the loan supply curve, in particular, those related to the BLS demand indicators.

[24] We address more deeply this concern by exploring a system specification in which the BLS demand indicator also enters the credit supply equation. Only a decline in demand has a significant and positive effect on the cost of credit, which is estimated to be by 20 basis points. However, the fit of the supply equation dramatically worsens and the various test for a correct identification (especially the weak identification test) become largely unsatisfactory, thus, casting serious doubts on this alternative specification. However, all the estimated coefficients remain virtually unchanged.

We did not find a significant direct effect on mortgage loan dynamics neither for the BLS "pure-supply" factor nor for the "risk-perception" factor. This possibility has been tested by considering alternative specifications in which these factors enter, one at the time, the loan demand equation. As for the "pure-supply" factor, this outcome might reflect the fact that for mortgage loans to households the BLS collects under a single heading ("cost of funds and balance sheet constraints") factors which are instead investigated separately in the case of loans to enterprises (for the latter, as mentioned above, banks are asked to indicate separately the importance of "liquidity position", "access to funding conditions" and "capital position"). It is possible that this choice results in an imprecise identification of the factors that concurred to credit rationing phenomena with respect to those affecting the credit cost.

We report in columns (b) and (b'), the estimated coefficients based on an alternative specification in which the BLS indicator capturing a "decrease" in loan demand also enters the supply equation. The estimated coefficients are negative and highly significant, consistently with the findings with the reduced-form equations. The fit of the loan supply equation improves and the various diagnostics test accept this specification. As for the interpretation of these results, the same considerations expressed in Section 3.3 still hold. Overall, the estimated loan supply and demand elasticity remains slightly affected.

In columns (c) and (c') we report the estimated coefficients for an alternative system in which the individual bank's marginal cost of funding enters as instrument the bank mark-up equation. A one percentage point increase in the marginal cost of funding is associated to a contemporaneous and lagged rise in the bank mark-up by, respectively, about 15 and 8 basis points. The estimated elasticity of loan demand to the cost of credit declines considerably, thus suggesting a relevant bias in our previous specification. This outcome may also be related to the fact that the BLS indicators capture only partially the "pure-supply" factors.

5. Including the sovereign spread

During the sovereign debt crisis great attention has been paid by analysts and policymakers to the developments of the spreads between sovereign bond yields of the euro-area countries hit by the tensions and those of Germany. This spread has indeed been regarded as a sort of "sufficient statistics" to measure the intensity of tensions. Recently, Albertazzi et al. (2012) have analysed reduced-form relationships between the BTP-Bund spread and developments in various credit market segments in Italy using macro data for the entire banking system. Neri (2013) and Zoli (2013) also found a significant role played by the sovereign spread on bank loan rates for a number of countries, including Italy.

In the light of this – and of the fact that in answering ad hoc questions recently included in the BLS several banks reported a specific effect of the sovereign debt crisis on their credit policy – we deemed it useful to investigate the information content of the spread for both credit demand and supply curves. In addition, the economic theory suggests that, both in crisis and in non-crisis times, the yield on sovereign bonds and the cost of credit are imperfect substitutes one another, meaning that both credit demand and supply may be decreasing functions in the sovereign bond yields (see Bernanke and Blinder, 1988).

We first test the role of the sovereign spread as a credit supply shifter. To this end, we consider systems of equations for both loans to enterprises and mortgage loans to households in which the change in the difference between the yield on the 10-year Italian government bond and the yield on the German bond of the same maturity enter as instrument in the loan supply equation. Since the sovereign spread is a macro variable common across banks, the estimated standard errors for its coefficients are lower than the true ones (see Moulton, 1990). We, therefore, report the statistical significance referring to standard errors computed by clustering observations over time periods. Results are presented in Tables 6 and 7.

The coefficient of the sovereign spread is positive and highly significant in the loan-spread equation for both loans to enterprises and mortgage loans to households: a 100 basis point increase in the spread is associated with a pass-through after one quarter of, respectively, 60 and 45 basis points. The estimated effect on the cost of new loans to enterprises is essentially the same as those obtained by recent studies based on aggregate data (see Albertazzi et al., 2012; Zoli, 2013; Neri, 2013).[25]

In the estimates for loans to enterprises the coefficients of the BLS funding condition indicator and the bank-specific marginal cost are no longer significant, while the coefficient of the supply conditions related to risk perception remain significant with the "considerably" qualification. For mortgage loans to households, the coefficient of the BLS indicator of cost of funds and balance sheet conditions is no longer significant. The marginal cost of funding and the indicator of risk perceptions, instead, remain statistically significant, though the coefficient of the former becomes smaller. The model now would suggest an upward sloping supply curve, albeit characterized by a higher degree of uncertainty.

Interestingly, the sovereign spread appears to have predictive content over the entire sample period and not only during the more recent period of sovereign debt tensions (an

[25] Albertazzi et al. (2012) fond that the impact is larger (50 basis points) when data for the period 1991-2011 on bank rates on outstanding short-term loans to non-financial corporations (including credit lines) are used.

interaction term between the sovereign spread and the sovereign dummy variable resulted not to be statistically significant when included in the regression). By contrast, for both loans to enterprises and mortgages to households the sovereign spread does not appear to have played a significant direct effect in the loan quantity equation, beyond that occurring through its effects on the cost of credit. The Sargan-Hansen test accepts the restrictions used for the identification of the loan demand, including the exclusion of the sovereign spread from that equation. However, a direct credit rationing effect was also tested by including the sovereign spread in the loan demand equation: its coefficient resulted to not be statistically significant, even if interacted with a "sovereign dummy". For loans to enterprises, the coefficients of the BLS supply indicators remain significant and of approximately the same magnitude. The coefficient for the bank mark-up becomes not statistically significant and both the weak identification and under-identification test strongly reject this system specification.

All in all, these results suggest that the relationship between banks' funding difficulties and credit developments reflected to a large extent the strains in the sovereign debt market, in particular concerning the cost of loans to enterprises. For these loans, the common shock hitting the banking system, as captured by the changes in the sovereign debt spread, prevailed over the more idiosyncratic components as potentially captured by the individual bank's answers on their funding difficulties and their marginal cost of funding. The evidence is less clear cut for mortgage loans to households, also due to the absence (as recalled above) of a BLS indicator which only measures the cost of fund supply factor.[26]

A potential concern in the interpretation of these results is that the effects of changes in the spread could be different depending on whether they reflect variations in the yield of the Italian government bonds or in the yield of the German Bunds. Indeed, during the crisis the spread reflected both idiosyncratic factors related to economic and public finance evolution in Italy and more general "flight-to-quality" phenomena connected with the investors' worries about the possible reversibility of the euro (the so-called "redenomination risk"). Although both factors have likely affected both yields (given the link between doubts about the sustainability of member countries' public debts and redenomination fears), one could

[26] A potential concern is that the sovereign spread may be endogenous, as also reflecting the worsening in the outlook for economic activity, as well as the worsening in banks' balance sheet conditions. As a further robustness check, we run an alternative system in which the sovereign spread is also considered as endogenous variable and regressed on all the BLS supply and demand indicators. As a result, the contemporaneous effect of the sovereign spread on the cost of credit becomes not statistically significant, while the lagged coefficient remains highly significant and of the same magnitude. In general, however, the fit of the model worsens and the various diagnostic tests are not always fully satisfactory. This implies that the most of the changes in the sovereign spread can be considered exogenous in our systems of simultaneous equations.

expect idiosyncratic factors to have exerted a stronger effect on the BPT yield and the Bund yield to have mostly reflected the latter type of phenomenon. Taking this into account, we ran an alternative regression, in which the BTP and the Bund yields were included separately among the explanatory variables, in the place of the sovereign spread.

The results, reported in columns (b) and (b') of each table, show that the effect of an increase in the sovereign spread on the cost of loans to enterprises is somehow stronger if it reflects a rise in the BTP yield, compared to the case in which it reflects a reduction in the Bund yield: the implied one-quarter pass-through of a 100 basis points increase in the sovereign spread is about 35 basis points when the increase only reflects a rise in the BTP yield and 20 basis points if the increase is entirely determined by a decline in the Bund yield. For the cost of mortgage loans to households, instead, the coefficients of the BTP and the Bund yields implied a pass-through by 38 and 27 basis points, respectively, after one quarter.

6. Using BLS terms and conditions

In this section we carry out a robustness check of the results obtained above, concerning in particular the identification of the direct (credit-rationing) effects on the dynamics of loans vis-à-vis the indirect effects taking place through the cost of credit. To this purpose, we use the BLS responses regarding the specific terms and conditions through which banks report to have changed their credit standards.

In the BLS banks are asked to indicate whether a change took place through a variation in price (bank's margin on average or on riskier loans) or in other non-price conditions (for loans to enterprises, non-interest rate charges, size of the loan or credit line, collateral requirements, loan covenants, maturity; for mortgage loans to households, non-interest rate charges, collateral requirements, loan-to-value ratio, maturity). Estimating the same equations considered above by including these BLS supply indicators as regressors instead of factors provides a particularly valuable test, since the various terms and conditions appear to be much more clearly related to, respectively, the cost and the quantity of credit than factors. Obtaining similar results with this alternative source of information – in particular as regards the estimated elasticity of loan demand and supply – would be a valid indication in support of our identification.

We also include the bank marginal cost of funding in the system specification to control for bank-specific omitted variables correlated with the bank spread, as discussed in Section 4. Interestingly, in occasion of some rounds of the BLS we have informally asked the respondent banks to state how they interpret the notion of "margin" when replying to the questionnaire. It turned out that some groups included in the Italian sample define margins with respect to market rates. Others, on the contrary, define margins with respect to some notion of cost of funding.

Table 8 reports the results of the estimates of the structural system carried out with the BLS indicators of terms and conditions for loans to firms.[27] We start the analysis in a partial equilibrium framework and assume that all the various BLS terms and conditions are instruments for the identification of the loan demand equation. For loans to firms, the estimated coefficients for the loan supply equation suggest that a considerable tightening of margins on riskier loans – that dominates the indicator related to margins on average loans - reported by all banks in the panel would be associated with a rise of about 88 basis points of the spread on impact, while the effect of a moderate tightening would be smaller (about 13 basis points). For the change in marginal cost of funding the estimated effect is 10 basis points and similar to that obtained in Table 4. In the loan demand equation, the coefficients of the BLS demand and the elasticity to the bank mark-up are significant and similar to those reported with the system of equations based on the BLS factors. The Sargan-Hansen test again rejects this identification since, as expected, the BLS supply indicator related to size of the loan enters directly the loan demand equation.

Therefore, we moved to a disequilibrium framework as reported in columns (b)-(b'), in which the non-price condition enters directly the demand equation. This variable has a significant negative effect on the loan growth rate by about -4.0 and -1.6 percentage points, respectively, when banks reported a tightening with the "considerably" and the "somehow" qualification. The elasticity of demand to the loan rate becomes much lower and not statistically significant. However, the latter coefficient becomes significant (about -1.1) if one considers a specification which also includes the sovereign spread.

Table 9 reports the results of the estimates of the regressions carried out with the BLS indicators of terms and conditions for mortgage loans to households. The impact of the BLS indicator related to the price conditions (that has been reported by banks in the sample period only with the "somehow" qualification) is similar, when compared to loans to enterprises. A reported tightening of the margins on average loans by all banks would be associated with an increase of 15 basis points in the loan spread. We find that the estimated effect of the BLS demand indicator remains alike to those estimated in the different specifications of Table 5. Consistently with the results obtained with the regressions including the BLS factors, we

[27] The quarterly net percentage of each BLS term and condition during the crisis is reported in Figure 4.

find no evidence of a “direct” rationing effect: the “loan-to-value ratio” does not enter significantly in the loan demand equation. The elasticity of loan demand to the bank mark-up remains also very similar, especially when the sovereign spread is included in the system.

7. Assessing the role of supply and demand factors: is the sovereign debt crisis different from the global crisis?

In order to quantify the role played by supply and demand factors in credit developments in Italy during the two phases of the financial crisis, we performed counterfactual exercises in which we compared the fitted values obtained from our estimates with those we would have obtained had supply and demand indicators remained unchanged at their pre-crisis levels (i.e., at the levels observed in 2007q2).

We performed the analysis using three different specifications. The models differ in the set of supply indicators used for the exercise: *i*) based on the BLS supply factors that provide a simple structural interpretation of the forces driving the cost and dynamics of loans during the crisis and allows us to better distinguish between “pure” and “risk-related” supply factors; *ii*) based on the BLS supply factors and the sovereign spread, which are in general characterized by a better fit and allow us to assess the relative importance of the common shock captured by the sovereign spread vis-à-vis the idiosyncratic bank-specific factors; based on the BLS terms and conditions and the sovereign spread, which offers a comparison on the relative role of credit rationing and indirect supply effects on the growth rate of credit.

Figures 5 and 6 highlight the results for the cost and the amount of loans to firms and of mortgage loans to households. All charts show the impact of the various driving forces in each quarter of the period under consideration, by depicting the effects on the quarterly change in bank interest rate margin and on the quarterly growth rate of loan quantities. Complementary information is provided by Figures A1 and A2 in the Appendix, which show the corresponding cumulated effects on the same credit variables.

7.1 Loans to firms

The results of the counterfactual exercises indicate that supply factors – as measured by the BLS indicators – exerted a relevant effect on both the cost and the availability of credit throughout the crisis. The magnitude of these effects was, on average, somehow stronger during the sovereign debt crisis than in the aftermath of the Lehman collapse. The tightening of supply conditions is estimated to have determined a quarterly rise in the cost of credit of 60 basis points at the peak of the sovereign debt crisis (2011q4), compared to 25 basis points at the peak of the global crisis (2008q4). The cumulated effect since the beginning of the crisis (until 2012q2) is estimated at 165 basis points, of which one half occurred during the global crisis and one half during the sovereign debt crisis.

As to loan dynamics, both weak demand and tight supply exerted a relevant negative effect on loans to enterprises over both phases of the crisis, but the estimated supply effects were much stronger during the sovereign debt crisis. The impact of supply factors reached its peak in the last quarter of 2011, when it reduced the q-o-q growth rate of loans to enterprises by about 2.0 percentage points, compared to around 1.1 percentage points in 2008q4. At the end of the sample period considered (2012Q2), supply factors are estimated to have had a cumulated negative impact on the stock of loans of about 10 percent, of which one third occurred during the global crisis and two thirds during the sovereign debt crisis. The largest effect on loan volumes is ascribed by a credit rationing rather than to the adjustment of the loan demand to the increase in the cost of credit.

The two phases of the crisis have been characterized by a different relative importance of the "pure" supply factors with respect to "risk perception". During the global crisis supply effects were mostly related to risk perception, while the impact of "pure" supply factors was smaller (consistently with the findings in DEN, 2011). By contrast, during the sovereign debt crisis "pure" supply factors, related to difficulties in the access to funding and to the capital position, became much more relevant: on average, these factors have determined around two thirds of the increase in interest rates and three fourths of the reduction in granted loans that can be attributed, as a whole, to all supply factors. The effect of "pure" supply factors, after reaching a peak in the last quarter of 2011, decreased sharply in the following quarters, as a result of a large improvement in funding conditions (which eventually stopped exerting an unfavourable influence on credit standards); this is a clear indication of the effectiveness of the exceptional measures adopted by the ECB at the end of 2011.[28]

Demand provided a strong negative contribution to loan dynamics throughout the crisis. Its negative contribution was greater than that of supply in most quarters of the period considered, with the notable exceptions of the periods around the tension peaks, at the end of 2008 and in the second half of 2011. Demand conditions were particularly weak during 2009 and 2012. On average, the quarterly contribution of demand to the rate of change of loans has been similar in two phases of the crisis. At the end of the sample period (2012q2)

[28] In December 2011, the Governing Council of the ECB announced two longer-term refinancing operations with maturity at 3 years and full allotment (which took place on 21 December and 29 February 2012, respectively) and an expansion of the range of assets eligible as collateral in refinancing operations.

demand conditions are estimated to have determined a cumulated reduction of the stock of loans by about 11 percent.

The counterfactual exercise conducted with the specification including the sovereign spread provides a similar picture about the relative importance of the pure-supply factors and risk perception. However, the contribution of the supply restriction on the cost of loans during the sovereign debt crisis is estimated to be higher, reaching 80 basis points at the peak (2011Q4). The cumulated effect over the entire crisis period rises to 180 basis points. As a consequence, the negative contribution of supply factors to the growth rate of loans to enterprises is also larger, reaching to 2.3 percentage points in the last quarter of 2011 and implying a cumulated negative impact on the stock of loans by 11 percent. The estimated contribution of the BLS demand indicator remains almost unchanged when we introduce the spread, confirming our interpretation of the latter as a credit supply shifter.

The decomposition of the effect of the sovereign spread changes between the parts attributable, respectively, to the changes in the BTP yield and in the Bund yield indicates that the impact of the domestic yield changes was stronger, counting for roughly two thirds of the total. We estimate that in 2012q2, absent "flight-to-quality" effects, which were particularly strong in 2011 and 2012, in connection with the risk of a euro-area breakup and a currency redenomination, the level of the bank margin would have been 1 percentage point lower and the stock of loans would have been higher by roughly the same amount.

Finally, the counterfactual exercise based on the specification including the BLS terms and conditions and the sovereign spread suggests that during the global crisis the "credit rationing" effects played a more relevant role than suggested by the previous exercise based on the BLS factors, while they are very similar during the sovereign debt crisis. A visual comparison of Figures 3a and 4a shows that the former finding reflects the higher frequency with which banks have reported to have tightened supply conditions by acting on the "size of the loan or the credit line" during the global crisis, compared to the frequency with which they reported the factors "capital position" and "funding conditions" (these are the BLS supply factors capturing credit rationing) to have influenced their decisions. All in all, these results suggest that the credit rationing effects during the global crisis might be somewhat underrated by our estimates based on supply factor indicators.

7.2 Mortgage loans to households

The results of the counterfactual exercises for mortgage loans to households suggest that the supply factors played a significant role also on both the cost and the availability of credit to this sector. Since the interpretation of the BLS demand indicator in the loan supply equation is not clear, we did not consider this factor as a driving force in the counterfactual exercises for the cost of credit.

For the loan rate, the magnitude of these effects, as captured by only the BLS factors, was broadly similar in the two phases of the crisis and comparable to what observed in the case of loans to enterprises. At the two peaks of the supply restriction (2008q4 and 2011q4) the contribution to the increase in the bank margin was, respectively, 20 and 45 basis points. However, the effects were more persistent over time. As a result, the cumulated effect of the supply restriction since the beginning of the crisis (until 2012q2) is estimated at about 130 basis points, distributed in equal proportion among the two phases of the crisis. The "pure" supply factors result to have played the most important role over the entire crisis period.

When we conduct the counterfactual exercise with the specification that includes the sovereign spread the contribution of supply factors to the cost of credit during the sovereign debt crisis roughly doubles at the peak of the sovereign debt crisis (the quarterly contribution of supply restriction reaches 85 basis points in 2011q4). The cumulated effect in 2012q2 is estimated at 210 basis points. These results remain virtually unchanged when using BLS terms and conditions in the place of the BLS supply factors.

As to loan quantities, the most striking result is that loan demand conditions provided the strongest negative contribution to loan dynamics throughout the crisis period, with the usual exceptions at the peaks. At the end of the sample period (2012q2) demand conditions are estimated to have determined a cumulated reduction of the stock of loans by about 6 percent for all the considered specifications.

Notwithstanding the large effects on the cost of credit, the restriction of supply exerted a more muted effect on quantities, compared to what observed for loans to enterprises. At the end of 2012q2, the cumulated effect of supply restrictions on the stock of loans amounted to about -6 percent. Since the estimated loan demand elasticity to the cost of credit is similar in the two credit market segments, this smaller impact of the supply restrictions on mortgage loans to households is entirely explained by the lack of direct "credit rationing" effects.

8. Conclusions

In this paper we provided new evidence on the information content of the BLS for the cost and the growth rate of loans to enterprises and mortgage loans to households, with a special focus on the main differences between the "global crisis" and the "sovereign debt crisis". The analysis was performed using a system of simultaneous equations that allowed a structural identification of loan demand and supply curves and a deeper understanding of the functioning of the credit markets. In particular, the statistical models are able to describe a standard partial equilibrium framework in which the restrictions of credit supply lead to an adjustment of the traded quantities via the elasticity of loan demand to the cost of credit, as well as a disequilibrium framework in which the credit market is characterized by credit rationing phenomena typical of the financial crises.

We found that supply tensions, as captured by banks' difficulties in funding conditions and risk perception, mostly affected the dynamics of loans to enterprises via the elasticity of loan demand to the bank margin. A worsening in banks' capital position and access to funding exerted a direct negative effect on credit growth, thus suggesting credit rationing phenomena. In the case of mortgage loans to households the supply tensions affected loan quantities only through their impact on bank interest rates.

When the sovereign spread is included in the system specification, it tends to offset much of the significance of the BLS indicators in the cost of credit equations, thus, suggesting that the evolution in the sovereign debt markets dramatically affect banks' funding conditions and risk aversion. Interestingly, changes in the sovereign spread affect credit conditions when they reflect a rise in the yield of Italian government bonds (which reflects idiosyncratic factors relating to economic and public finance evolution in Italy together with common euro-area developments), as well as when it originates from a reduction in the yield on the German government bonds (e.g. as a result of a "flight to quality"). This is not surprising, given the strict interconnections between national and euro-area developments, in particular during the sovereign debt crisis.

A counterfactual exercise in which demand and supply conditions (as measured by the BLS indicators and by the sovereign spread) are assumed to have remained at their pre-crisis levels over the entire crisis period indicates that the negative contribution of supply factors to the growth rate of loans has been smaller than that of demand conditions on average, but much higher in the quarters around the peaks of the two crisis periods (2008q4 and 2011q4). The exercise suggests that at mid-2012, had supply conditions remained at the pre-crisis levels, interest rates would have been almost 200 basis points lower for both loans to firms and mortgages to households; the stock of credit would have been higher by 11 and 8 percent. The effect of "pure-supply" factors reached an unprecedented peak in the last quarter of 2011, while decreasing sharply in the next quarter, as the result of the large improvement in liquidity conditions following the exceptional measures adopted by the ECB at the end of 2011 (the introduction of three-year, full-allotment refinancing operations and the expansion of the range of assets eligible as collateral in the operations).

Overall, the supply factors exerted a stronger effect on credit developments during the sovereign debt crisis than during the global crisis, as the result of a larger influence of the "pure-supply" factors as opposed to "risk-perception" ones. The estimated effects on the cost of credit are very large and of comparable magnitude for both enterprises and households. On the contrary, the corresponding effects on loan quantities are significantly more muted in the case of mortgage loans to households. Our structural models suggest that this smaller impact is explained mostly by the lack of "credit rationing" effect in the households' sector. The estimated loan demand elasticity to the cost of credit is, indeed, similar in the two credit markets for all specifications we considered.

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Tables and Figures

Table 1

BLS supply and demand conditions for loans to enterprises: descriptive statistics

(frequency of responses and, in brackets, percentages with respect to total in each period)

Supply				Demand			
	Pre-crisis 02Q4-07Q2	During crisis 07Q3-12Q2			Pre-crisis 02Q4-07Q2	During crisis 07Q3-12Q2	
		global crisis 07Q3-10Q1	sovereign debt crisis 10Q2-12Q2			global crisis 07Q3-10Q1	sovereign debt crisis 10Q2-12Q2
1="eased considerably"	0 *(0.0)*	0 *(0.0)*	0 *(0.0)*	1="decreased considerably"	0 *(0.0)*	0 *(0.0)*	1 *(1.4)*
2="eased somewhat"	2 *(1.5)*	1 *(1.2)*	0 *(0.0)*	2="decreased somewhat"	12 *(9.2)*	16 *(19.0)*	19 *(26.4)*
3="basically unchanged"	105 *(80.2)*	52 *(61.9)*	50 *(69.4)*	3="basically unchanged "	88 *(67.2)*	54 *(64.3)*	37 *(51.4)*
4="tightened somewhat"	21 *(16.0)*	31 *(36.9)*	21 *(29.2)*	4="increased somewhat"	31 *(23.7)*	14 *(16.7)*	15 *(20.8)*
5="tightened considerably"	3 *(2.3)*	0 *(0.0)*	1 *(1.4)*	5="increased considerably"	0 *(0.0)*	0 *(0.0)*	0 *(0.0)*
Total observations	131 *(100.0)*	84 *(100.0)*	72 *(100.0)*	Total observations	131 *(100.0)*	84 *(100.0)*	72 *(100.0)*

Table 2

BLS supply and demand conditions for mortgage loans to households: descriptive statistics

(frequency of responses and, in brackets, percentages with respect to total in each period)

Supply				Demand			
	Pre-crisis 02Q4-07Q2	During crisis 07Q3-12Q2			Pre-crisis 02Q4-07Q2	During crisis 07Q3-12Q2	
		global crisis 07Q3-10Q1	sovereign debt crisis 10Q2-12Q2			global crisis 07Q3-10Q1	sovereign debt crisis 10Q2-12Q2
1="eased considerably"	0 *(0.0)*	0 *(0.0)*	0 *(0.0)*	1="decreased considerably"	0 *(0.0)*	4 *(4.8)*	5 *(6.9)*
2="eased somewhat"	18 *(13.7)*	3 *(3.6)*	2 *(2.8)*	2="decreased somewhat"	6 *(4.6)*	26 *(31.0)*	22 *(30.6)*
3="basically unchanged"	104 *(79.4)*	54 *(64.3)*	51 *(70.8)*	3="basically unchanged "	70 *(53.4)*	41 *(48.8)*	31 *(43.1)*
4="tightened somewhat"	9 *(6.9)*	27 *(32.1)*	19 *(26.4)*	4="increased somewhat"	50 *(38.2)*	13 *(15.5)*	14 *(19.4)*
5="tightened considerably"	0 *(0.0)*	0 *(0.0)*	0 *(0.0)*	5="increased considerably"	5 *(3.8)*	0 *(0.0)*	0 *(0.0)*
Total observations	131 *(100.0)*	84 *(100.0)*	72 *(100.0)*	Total observations	131 *(100.0)*	84 *(100.0)*	72 *(100.0)*

Table 3a

Reduced-form regressions for loans to enterprises

	Δmark-up (t)	Δloan (t)	Δmark-up (t)	Δloan (t)
	(a)	(b)	(a')	(b')
Δ(loan) (t-1)		0.008		0.032
Δ(loan) (t-2)		0.195 ***		0.181 ***
Δ(mark-up) (t-1)	-0.381 ***		-0.381 ***	
Δ(mark-up) (t-2)	-0.234 ***		-0.234 ***	
BLS demand, increase (t)	0.101	0.916 **	0.097	0.898 ***
BLS demand, decrease (t)	0.060	-0.912 **	0.061	-0.851 ***
BLS supply, capital position, tightening (t)	-0.081	-1.107 *	-0.085	-1.291 ***
BLS supply, funding conditions, tightening (t)	0.441 ***	-1.356 **	0.443 ***	-1.209 *
BLS supply, funding conditions, tightening (t-1)	0.176 *	-1.271 **	0.178 *	-1.068 **
BLS supply, risk perception, strong tightening (t)	0.545 ***	-1.709 *	0.542 ***	-1.778 ***
BLS supply, risk perception, moderate tightening (t)	0.095 *	-0.052	0.092	-0.042
Constant	0.147 **	1.557 ***	0.079	1.456 ***
Fixed-effects	yes	yes	yes	yes
Seasonal dummies	yes	yes	yes	yes
Estimation	SURE	SURE	GMM	GMM
Observations	247	247	247	247
R-squared	0.313	0.311	-	-

Notes: The dependent variables, *"Δloan"* and *"Δ(mark-up)"* are, respectively, the quarterly change of the loan quantities and the quarterly change in bank mark-up (computed as the difference between the average rate on new loans and the Eonia rate). "*BLS supply funding conditions, tightening", "BLS supply, risk perception, tightening", "BLS supply, capital position, tightening"* are dummy variables taking the value of 1 if the bank reported that the respective factor contributed to a tightening in credit supply conditions (also distinguishing, when applicable, whether the bank reported that the specific factor contributed considerably/somewhat to the tightening; *risk perception* is related to the general economic activity and/or industry of firm-specific outlook). "*BLS demand, decrease", "BLS demand, increase"* are dummy variables taking the value of 1 if the bank reported, respectively, decrease/increase in demand. *, ** and *** denote significance, respectively, at 10%; * 5% and 1% confidence level.

Table 3b

Reduced-form regressions for mortgage loans to households

	Δmark-up (t)	Δloan (t)	Δmark-up (t)	Δloan (t)
	(a)	(b)	(a')	(b')
Δ(loan) (t-1)	-0.018		-0.010	
Δ(loan) (t-2)	-0.229 ***		-0.230 ***	
Δ(mark-up) (t-1)		0.303 ***		0.258 ***
Δ(mark-up) (t-2)		0.283 ***		0.307 ***
Δ(10year swap rate) (t)	-0.282 ***		-0.281 ***	
BLS demand, decrease (t)	0.137 **	-0.722 *	0.138 **	-0.621 ***
BLS demand, decrease (t-1)	0.110 *	-0.773 **	0.111 *	-0.889 ***
BLS demand, increase (t)	-0.029	0.173	-0.032	0.162
BLS "pure-supply", tightening (t-1)	0.185 **	-0.737	0.185 **	-0.477 *
BLS supply, risk perception, strong tightening (t-1)	0.427 *	-3.274 ***	0.420 *	-3.170 ***
BLS supply, risk perception, moderate tightening (t)	0.126 **	0.023	0.122 **	0.099
Constant	0.075	1.676 ***	0.020	1.955 ***
Fixed-effects	yes	yes	yes	yes
Seasonal dummies	yes	yes	yes	yes
Estimation	SURE	SURE	GMM	GMM
Observations	247	247	247	236
R-squared	0.463	0.312	-	-

Notes: The dependent variables, "*Δloan*" and "*Δ(mark-up)*" are, respectively, the quarterly change in loan quantities and the quarterly change in bank mark-up (computed as the difference between the average rate on new loans and the Eonia rate). "*BLS pure-supply, tightening*", "*BLS risk perception, tightening*", are dummy variables taking the value of 1 if the bank reported that the respective factor contributed to a tightening in credit supply conditions (also distinguishing, when applicable, whether the bank reported that the specific factor contributed considerably / somewhat to the tightening; "*pure-supply*" is related to cost of funds and balance-sheet constraints; *risk perception* is related to the general economic activity and/or housing market prospects). "*BLS demand, decrease*", "*BLS demand, increase*" are dummy variables taking the value of 1 if the bank reported, respectively, decrease/increase in demand. *, ** and *** denote significance, respectively, at 10%; * 5% and 1% confidence level.

Table 4
Structural equations for loans to enterprises

	(a)	(a')	(b)	(b')	(c)	(c')
Endogenous variables:	**Δloan (t)**	**Δ(mark-up)(t)**	**Δloan (t)**	**Δ(mark-up)(t)**	**Δloan (t)**	**Δ(mark-up)(t)**
Δ(loan) (t)		**0.029**		**0.032**		**0.029**
Δ(mark-up) (t)	**-2.061 *****		**-1.567 ****		**-1.483 ****	
Predetermined variables:						
Δ(loan) (t-1)	0.038		0.013		0.015	
Δ(loan) (t-2)	0.216 ***		0.221 ***		0.218 ***	
Δ(mark-up) (t-1)		-0.368 ***		-0.377 ***		-0.331 ***
Δ(mark-up) (t-2)		-0.211 ***		-0.214 ***		-0.183 ***
Exogenous variables:						
BLS demand, increase (t)	1.002 **		1.002 **		0.960 **	
BLS demand, decrease (t)	-1.089 **		-0.759 **		-0.708 *	
BLS supply, capital position, tightening (t)		-0.021	-1.456 **		-1.526 **	
BLS supply, funding conditions, tightening (t)		0.490 ***		0.481 ***		0.381 ***
BLS supply, funding conditions, tightening (t-1)		0.242 **	-1.442 **	0.225 **	-1.377 **	0.269 **
BLS supply, risk perception, strong tightening (t)		0.578 ***		0.561 ***		0.486 ***
BLS supply, risk perception, moderate tightening (t)		0.107 *		0.106 *		0.093 *
ΔMarginal cost of funding (t)						0.106 ***
Constant term	1.533 ***	0.111	1.624 ***	0.111	1.638 ***	0.105
Fixed-effects	yes	yes	yes	yes	yes	yes
Seasonal dummies	yes	yes	yes	yes	yes	yes
Estimation technique	3SLS	3SLS	3SLS	3SLS	3SLS	3SLS
Number of observations (N)	247	247	247	247	245	245
Number of regressors (K)	8	11	10	10	10	11
Number of endogenous regressors (K1)	1	1	1	1	1	1
Number of instruments (L)	14	14	14	14	15	15
Number of excluded instruments (L1)	7	4	5	4	6	5
R-squared	0.210	0.270	0.281	0.265	0.282	0.294
Identification-diagnostics:						
Underidentification test	56.32	26.11	55.02	26.11	58.68	27.95
Chi-sq(L1-K1+1) P-value	0.00	0.00	0.00	0.00	0.00	0.00
Weak identification test (Cragg-Donald statistic)	9.94	6.90	14.50	6.14	12.22	5.94
Overidentification test (Sargan-Hansen statistic)	16.77	4.87	3.33	4.96	3.94	5.09
Chi-sq(L-K) P-value	0.01	0.18	0.50	0.29	0.56	0.28

Notes: The dependent variables, *"Δloan"* and *"Δ(mark-up)"* are, respectively, the quarterly change of the loan quantities and the quarterly change in bank mark-up (computed as the difference between the average rate on new loans and the Eonia rate). "*BLS supply funding conditions, tightening", "BLS supply, risk perception, tightening", "BLS supply, capital position, tightening"* are dummy variables taking the value of 1 if the bank reported that the respective factor contributed to a tightening in credit supply conditions (also distinguishing, when applicable, whether the bank reported that the specific factor contributed considerably/somewhat to the tightening; *risk perception* is related to the general economic activity and/or industry of firm-specific outlook). "*BLS demand, decrease", "BLS demand, increase"* are dummy variables taking the value of 1 if the bank reported, respectively, decrease/increase in demand. *, ** and *** denote significance, respectively, at 10%; * 5% and 1% confidence level.

The "*Underidentification test*" is a Lagrange-Multiplier test aiming at check if the equation is identified. The null hypothesis is that the equation is under-identified (i.e. the matrix of reduced form coefficients on the L1 excluded instruments has rank K1-1), while the alternative hypothesis is that the equation is identified (i.e. the matrix has rank exactly equal to K1). Under the null hypothesis, the test statistic is distributed as chi-squared in (L1-K1+1) degrees of freedom. A rejection of the null indicates that the equation is identified. The "*Weak identification test*" is the F-version of the Cragg-Donald Wald statistic. The "*Overidentification test*" is based on the Sargan-Hansen statistics; the null hypothesis tested is that the exclusion restrictions are valid. Under the null hypothesis, the test statistic is distributed as chi-squared in the number of (L-K) over-identifying restrictions. A rejection casts doubt on the validity of the instruments used for the equation identification.

Table 5

Structural equations for mortgage loans to households

	(a)	(a')	(b)	(b')	(c)	(c')
Endogenous variables:	**Δloan (t)**	**Δ(mark-up)(t)**	**Δloan (t)**	**Δ(mark-up)(t)**	**Δloan (t)**	**Δ(mark-up)(t)**
Δ(loan) (t)		**0.005**		**0.013**		**0.014**
Δ(mark-up) (t)	**-1.714 ***		**-1.968 ** **		**-1.257 ***	
Predetermined variables:						
Δ(loan) (t-1)	0.306 ***		0.312 ***		0.308 ***	
Δ(loan) (t-2)	0.286 ***		0.290 ***		0.295 ***	
Δ(mark-up) (t-1)		0.064		0.006		0.000
Δ(mark-up) (t-2)		-0.164 ***		-0.197 ***		-0.190 ***
Exogenous variables:						
ΔSwap10y(t)		-0.346 ***		-0.299 ***		-0.305 ***
BLS demand, decrease (t)				0.159 **		0.099 *
BLS demand, decrease (t-1)	-1.100 ***		-0.853 **	0.112 *	-0.906 **	0.172 **
BLS "pure-supply", tightening (t-1)		0.220 ***		0.200 ***		0.169 **
BLS supply, risk perception, strong tightening (t-1)		0.476 **		0.532 **		0.355 *
BLS supply, risk perception, moderate tightening (t)		0.175 ***		0.122 *		0.119 *
ΔMarginal cost of funding (t)						0.156 ***
ΔMarginal cost of funding (t-1)						0.083 ***
Constant term	1.813 ***	0.059	1.786 ***	0.014	1.714 ***	0.002
Fixed-effects	yes	yes	yes	yes	yes	yes
Seasonal dummies	yes	yes	yes	yes	yes	yes
Estimation technique	3SLS	3SLS	3SLS	3SLS	3SLS	3SLS
Number of observations (N)	247	247	247	247	247	247
Number of regressors (K)	7	10	7	12	7	14
Number of endogenous regressors (K1)	1	1	1	1	1	1
Number of instruments (L)	12	12	13	13	15	15
Number of excluded instruments (L1)	6	3	7	2	9	2
R-squared	0.418	0.270	0.408	0.306	0.436	0.398
Identification-diagnostics:						
Underidentification test	47.12	65.75	51.55	63.81	75.43	64.75
Chi-sq(L1-K1+1) P-value	0.00	0.00	0.00	0.00	0.00	0.00
Weak identification test (Cragg-Donald statistic)	9.31	28.84	8.90	41.32	11.57	41.90
Overidentification test (Sargan-Hansen statistic)	8.09	7.69	8.55	0.04	12.65	0.30
Chi-sq(L-K) P-value	0.15	0.02	0.20	0.84	0.12	0.59

Notes: The dependent variables, "*Δloan*" and "*Δ(mark-up)*" are, respectively, the quarterly change in loan quantities and the quarterly change in bank mark-up (computed as the difference between the average rate on new loans and the Eonia rate). "*BLS pure-supply, tightening", "BLS risk perception, tightening*", are dummy variables taking the value of 1 if the bank reported that the respective factor contributed to a tightening in credit supply conditions (also distinguishing, when applicable, whether the bank reported that the specific factor contributed considerably / somewhat to the tightening; *"pure-supply"* is related to cost of funds and balance-sheet constraints; *risk perception* is related to the general economic activity and/or housing market prospects). *"BLS demand, decrease", "BLS demand, increase"* are dummy variables taking the value of 1 if the bank reported, respectively, decrease/increase in demand. *, ** and *** denote significance, respectively, at 10%; * 5% and 1% confidence level.

The "*Under-identification test*" is a Lagrange-Multiplier test aiming at check if the equation is identified. The null hypothesis is that the equation is under-identified (i.e. the matrix of reduced form coefficients on the L1 excluded instruments has rank K1-1), while the alternative hypothesis is that the equation is identified (i.e. the matrix has rank exactly equal to K1). Under the null hypothesis, the test statistic is distributed as chi-squared in (L1-K1+1) degrees of freedom. A rejection of the null indicates that the equation is identified. The "*Weak identification test*" is the F-version of the Cragg-Donald Wald statistic. The "*Over-identification test*" is based on the Sargan-Hansen statistics; the null hypothesis tested is that the exclusion restrictions are valid. Under the null hypothesis, the test statistic is distributed as chi-squared in the number of (L-K) over-identifying restrictions. A rejection casts doubt on the validity of the instruments used for the equation identification.

Table 6

Structural equations for loans to enterprises including the sovereign spread

	(a)	(a')	(b)	(b')
Endogenous variables:	**Δloan (t)**	**Δ(mark-up)(t)**	**Δloan (t)**	**Δ(mark-up)(t)**
Δ(loan) (t)		**0.014**		**0.004**
Δ(mark-up) (t)	**-1.498 ****		**-1.279 ****	
Predetermined variables:				
Δ(loan) (t-1)	0.012		0.011	
Δ(loan) (t-2)	0.212 ***		0.205	
Δ(mark-up) (t-1)		-0.406 ***		-0.409 ***
Δ(mark-up) (t-2)		-0.148 ***		-0.140 ***
Exogenous variables:				
BLS demand, increase (t)	0.951 **		0.925 **	
BLS demand, decrease (t)	-0.796 *		-0.857 *	
BLS supply, capital position, tightening (t)	-1.497 **		-1.486 **	
BLS supply, funding conditions, tightening (t)		0.061		0.005
BLS supply, funding conditions, tightening (t-1)	-1.368 **	0.065	-1.400 **	0.040
BLS supply, risk perception, strong tightening (t)		0.343 **		0.347 **
BLS supply, risk perception, moderate tightening (t)		0.057		0.053
ΔMarginal cost of funding (t)		0.039		0.034
Δsovereign spread (t)		0.245 ***		
Δsovereign spread (t-1)		0.389 ***		
ΔItalian BTP yield (t)				0.310 ***
ΔItalian BTP yield (t-1)				0.464 ***
ΔGerman Bund yield (t)				-0.239 ***
ΔGerman Bund yield (t-1)				-0.306 ***
Constant term	1.670 ***	0.077	1.667 ***	0.111
Fixed-effects	yes	yes	yes	yes
Seasonal dummies	yes	yes	yes	yes
Estimation technique	3SLS	3SLS	3SLS	3SLS
Number of observations (N)	247	247	247	247
Number of regressors (K)	10	13	10	15
Number of endogenous regressors (K1)	1	1	1	1
Number of instruments (L)	17	17	19	19
Number of excluded instruments (L1)	8	5	10	5
R-squared	0.282	0.417	0.302	0.437
Identification-diagnostics:				
Underidentification test	83.27	29.02	87.01	25.04
Chi-sq(L1-K1+1) P-value	0.00	0.00	0.00	0.00
Weak identification test (Cragg-Donald statistic)	14.99	6.14	12.73	5.15
Overidentification test (Sargan-Hansen statistic)	4.03	5.20	7.30	6.05
Chi-sq(L-K) P-value	0.78	0.27	0.60	0.20

Notes: The dependent variables, *"Δloan"* and *"Δ(mark-up)"* are, respectively, the quarterly change in loan quantities and the quarterly change in bank mark-up (computed as the difference between the average rate on new loans and the Eonia rate). *"BLS supply funding conditions, tightening"*, *"BLS supply, risk perception, tightening"*, *"BLS supply, capital position, tightening"* are dummy variables taking the value of 1 if the bank reported that the respective factor contributed to a tightening in credit supply conditions (also distinguishing, when applicable, whether the bank reported that the specific factor contributed considerably/somewhat to the tightening; *risk perception* is related to the general economic activity and/or industry of firm-specific outlook). *"BLS demand, decrease"*, *"BLS demand, increase"* are dummy variables taking the value of 1 if the bank reported, respectively, decrease/increase in demand. *, ** and *** denote significance, respectively, at 10%; * 5% and 1% confidence level.

The *"Under-identification test"* is a Lagrange-Multiplier test aiming at check if the equation is identified. The null hypothesis is that the equation is under-identified (i.e. the matrix of reduced form coefficients on the L1 excluded instruments has rank K1-1), while the alternative hypothesis is that the equation is identified (i.e. the matrix has rank exactly equal to K1). Under the null hypothesis, the test statistic is distributed as chi-squared in (L1-K1+1) degrees of freedom. A rejection of the null indicates that the equation is identified. The *"Weak identification test"* is the F-version of the Cragg-Donald Wald statistic. The *"Over-identification test"* is based on the Sargan-Hansen statistics; the null hypothesis tested is that the exclusion restrictions are valid. Under the null hypothesis, the test statistic is distributed as chi-squared in the number of (L-K) over-identifying restrictions. A rejection casts doubt on the validity of the instruments used for the equation identification.

Table 7

Structural equations for mortgage loans to households including BLS supply factors and the sovereign spread

	(a)	(a')	(b)	(b')
Endogenous variables:	**Δloan (t)**	**Δ(mark-up)(t)**	**Δloan (t)**	**Δ(mark-up)(t)**
Δ(loan) (t)		**0.022 ***		**0.021 ***
Δ(mark-up) (t)	**-1.429 ****		**-1.408 ****	
Predetermined variables:				
Δ(loan) (t-1)	0.289 ***		0.289 ***	
Δ(loan) (t-2)	0.322 ***		0.322 ***	
Exogenous variables:				
ΔSwap10y(t)		-0.217 ***		-0.376 ***
BLS demand, decrease (t)		0.007		0.000
BLS demand, decrease (t-1)	-0.887 **	0.150 ***	-0.891 **	0.152 ***
BLS "pure-supply", tightening (t-1)		0.049		0.028
BLS supply, risk perception, strong tightening (t-1)		0.469 **		0.483 **
BLS supply, risk perception, moderate tightening (t)		0.082		0.092
ΔMarginal cost of funding (t)		0.089 ***		0.090 ***
ΔMarginal cost of funding (t-1)		0.036		0.027
Δsovereign spread (t)		0.135 **		
Δsovereign spread (t-1)		0.330 ***		
ΔItalian BTP yield (t)				0.116 *
ΔItalian BTP yield (t-1)				0.380 ***
ΔGerman Bund yield (t-1)				-0.267 ***
Constant term	1.653 ***	-0.066	1.650 ***	-0.057
Fixed-effects	yes	yes	yes	yes
Seasonal dummies	yes	yes	yes	yes
Estimation technique	3SLS	3SLS	3SLS	3SLS
Number of observations (N)	256	256	256	256
Number of regressors (K)	7	14	7	15
Number of endogenous regressors (K1)	1	1	1	1
Number of instruments (L)	15	15	16	16
Number of excluded instruments (L1)	9	2	10	2
R-squared	0.427	0.432	0.428	0.430
Identification-diagnostics:				
Underidentification test	92.34	66.61	91.71	66.45
Chi-sq(L1-K1+1) P-value	0.00	0.00	0.00	0.00
Weak identification test (Cragg-Donald statistic)	15.46	42.94	13.70	42.62
Overidentification test (Sargan-Hansen statistic)	8.82	0.51	9.01	0.37
Chi-sq(L-K) P-value	0.36	0.48	0.44	0.54

Notes: The dependent variables, "*Δloan*" and "*Δ(mark-up)*" are, respectively, the quarterly change in loan quantities and the quarterly change in bank mark-up (computed as the difference between the average rate on new loans and the Eonia rate). "*BLS pure-supply, tightening", "BLS risk perception, tightening*", are dummy variables taking the value of 1 if the bank reported that the respective factor contributed to a tightening in credit supply conditions (also distinguishing, when applicable, whether the bank reported that the specific factor contributed considerably/somewhat to the tightening; *"pure-supply"* is related to cost of funds and balance-sheet constraints; *risk perception* is related to the general economic activity and/or housing market prospects. *"BLS demand, decrease", "BLS demand, increase"* are dummy variables taking the value of 1 if the bank reported, respectively, decrease/increase in demand. *, ** and *** denote significance, respectively, at 10%; * 5% and 1% confidence level.

The "*Under-identification test*" is a Lagrange-Multiplier test aiming at check if the equation is identified. The null hypothesis is that the equation is under-identified (i.e. the matrix of reduced form coefficients on the L1 excluded instruments has rank K1-1), while the alternative hypothesis is that the equation is identified (i.e. the matrix has rank exactly equal to K1). Under the null hypothesis, the test statistic is distributed as chi-squared in (L1-K1+1) degrees of freedom. A rejection of the null indicates that the equation is identified. The "*Weak identification test*" is the F-version of the Cragg-Donald Wald statistic. The "*Over-identification test*" is based on the Sargan-Hansen statistics; the null hypothesis tested is that the exclusion restrictions are valid. Under the null hypothesis, the test statistic is distributed as chi-squared in the number of (L-K) over-identifying restrictions. A rejection casts doubt on the validity of the instruments used for the equation identification.

Table 8

Structural equations for loans to enterprises:
using BLS indicators on "terms and conditions"

	(a) Δ(loan)(t)	(a') Δ(mark-up)(t)	(b) Δ(loan)(t)	(b') Δ(mark-up)(t)	(c) Δ(loan)(t)	(c') Δ(mark-up)(t)
Endogenous variables:						
Δ(loan) (t)		**0.014**		**0.016**		**0.018**
Δ(mark-up) (t)	**-1.896 *****		**-0.729**		**-1.060 ***	
Predetermined variables:						
Δ(loan) (t-1)	0.037		0.033		0.035	
Δ(loan) (t-2)	0.207 ***		0.200 ***		0.204 ***	
Δ(mark-up) (t-1)		-0.286 ***		-0.304 ***		-0.434 ***
Δ(mark-up) (t-2)		-0.131 **		-0.140 **		-0.143 **
Exogenous variables:						
BLS demand, decrease (t)	-1.096 **		-0.905 **		-0.913 **	
BLS demand, increase (t)	0.926 **		0.960 **		0.975 **	
BLS supply, margin on riskier loans, strong tightening (t)		0.875 ***		0.797 ***		0.413 **
BLS supply, margin on riskier loans, moderate tightening (t)		0.130 **		0.138 ***		0.129 ***
BLS supply, size of the loan, strong tightening (t)		-0.144	-4.705 **		-4.519 **	
BLS supply, size of the loan, moderate tightening (t)		0.188 **	-1.627 ***	0.167 **	-1.550 ***	0.055
Δ(marginal cost of funding) (t)		0.104 ***		0.108 ***		0.037
Δ(sovereign spread) (t)						0.220 ***
Δ(sovereign spread) (t-1)						0.401 ***
Constant term	1.571 ***	0.051	1.710 ***	0.056	1.741 ***	0.016
Fixed-effects	yes	yes	yes	yes	yes	yes
Seasonal dummies	yes	yes	yes	yes	yes	yes
Estimation technique	3SLS	3SLS	3SLS	3SLS	3SLS	3SLS
Number of observations (N)	245	245	245	245	245	245
Number of regressors (K)	8	11	10	10	10	12
Number of endogenous regressors (K1)	1	1	1	1	1	1
Number of instruments (L)	14	14	14	14	16	16
Number of excluded instruments (L1)	7	4	5	5	7	5
R-squared	0.218	0.311	0.295	0.308	0.290	0.430
Identification-diagnostics:						
Underidentification test	58.29	27.12	51.34	27.87	84.20	27.78
Chi-sq(L1-K1+1) P-value	0.00	0.00	0.00	0.00	0.00	0.00
Weak identification test (Cragg-Donald statistic)	10.43	7.21	12.37	5.95	17.51	5.87
Overidentification test (Sargan-Hansen statistic)	17.57	4.97	5.39	5.19	5.77	6.72
Chi-sq(L-K) P-value	0.00	0.17	0.25	0.27	0.45	0.15

Notes: The dependent variables, "*Δloan*" and "*Δ(mark-up)*" are, respectively, the quarterly change in loan quantities and the quarterly change in bank mark-up (computed as the difference between the average rate on new loans and the Eonia rate). "*BLS supply margin on riskier loans, strong tightening*", "*BLS supply, margin on riskier loans, moderate tightening*", "*BLS supply, size of the loan, strong tightening*", "*BLS supply, size of the loan, moderate tightening*" are dummy variables taking the value of 1 if the bank reported that the respective term or condition was tightened/eased also distinguishing, when applicable, with considerably/ somewhat qualification. "*BLS demand, decrease*", "*BLS demand, increase*" are dummy variables taking the value of 1 if the bank reported, respectively, decrease/increase in demand. *, ** and *** denote significance, respectively, at 10%; * 5% and 1% confidence level.

The "*Under-identification test*" is a Lagrange-Multiplier test aiming at check if the equation is identified. The null hypothesis is that the equation is under-identified (i.e. the matrix of reduced form coefficients on the L1 excluded instruments has rank K1-1), while the alternative hypothesis is that the equation is identified (i.e. the matrix has rank exactly equal to K1). Under the null hypothesis, the test statistic is distributed as chi-squared in (L1-K1+1) degrees of freedom. A rejection of the null indicates that the equation is identified. The "*Weak identification test*" is the F-version of the Cragg-Donald Wald statistic. The "*Over-identification test*" is based on the Sargan-Hansen statistics; the null hypothesis tested is that the exclusion restrictions are valid. Under the null hypothesis, the test statistic is distributed as chi-squared in the number of (L-K) over-identifying restrictions. A rejection casts doubt on the validity of the instruments used for the equation identification.

Table 9

Structural equations for mortgage loans to households: using BLS indicators on "terms and conditions"

	(a) Δ(loan)(t)	(a') Δ(mark-up)(t)	(b) Δ(loan)(t)	(b') Δ(mark-up)(t)
Endogenous variables:				
Δ(loan) (t)		**0.015**		**0.019**
Δ(mark-up) (t)	**-1.716 ****		**-1.367 ****	
Predetermined variables:				
Δ(loan) (t-1)	0.290 ***		0.288 ***	
Δ(loan) (t-2)	0.320 ***		0.322 ***	
Exogenous variables:				
Δ(swap10y) (t)		-0.388 ***		-0.247 ***
BLS demand, decrease (t)		0.058		-0.014
BLS demand, decrease (t-1)	-0.831 **	0.119 **	-0.900 **	0.142 ***
BLS supply, margin on loans, moderate tightening (t)		0.152 ***		0.105 **
BLS supply, margin on loans, moderate easing (t)		-0.039		-0.022
BLS supply, loan-to-value ratio, moderate tightening (t)		0.083		0.093
BLS supply, loan-to-value ratio, moderate easing (t)		0.050		0.061
Δ(marginal cost of funding) (t)		0.177 ***		0.100 ***
Δ(marginal cost of funding) (t-1)		0.113 ***		0.050 *
Δ(sovereign spread) (t)				0.155 ***
Δ(sovereign spread) (t-1)				0.317 ***
Constant term	1.693 ***	0.000	1.643 ***	-0.065
Fixed-effects	yes	yes	yes	yes
Seasonal dummies	yes	yes	yes	yes
Estimation technique	3SLS	3SLS	3SLS	3SLS
Number of observations (N)	256	256	256	256
Number of regressors (K)	7	13	7	15
Number of endogenous regressors (K1)	1	1	1	1
Number of instruments (L)	14	14	16	16
Number of excluded instruments (L1)	8	2	10	2
R-squared	0.419	0.345	0.429	0.435
Identification-diagnostics:				
Underidentification test	66.56	68.01	92.50	67.69
Chi-sq(L1-K1+1) P-value	0.00	0.00	0.00	0.00
Weak identification test (Cragg-Donald statistic)	10.77	44.38	13.90	43.71
Overidentification test (Sargan-Hansen statistic)	9.25	0.05	10.06	0.27
Chi-sq(L-K) P-value	0.24	0.83	0.35	0.60

Notes: The dependent variables, "*Δloan*" and "*Δ(mark-up)*" are, respectively, the quarterly change in loan quantities and the quarterly change in bank mark-up (computed as the difference between the average rate on new loans and the Eonia rate). *"BLS supply margin on loans, moderate tightening"*, *"BLS supply, margin on loans, moderate easing"*, *"BLS supply, loan-to-value ratio, moderate tightening"*, *"BLS supply, loan-to-value ratio, moderate easing"* are dummy variables taking the value of 1 if the bank reported that the respective term or condition was tightened/eased also distinguishing, when applicable, with considerably/ somewhat qualification. *"BLS demand, decrease"*, *"BLS demand, increase"* are dummy variables taking the value of 1 if the bank reported, respectively, decrease/increase in demand. *, ** and *** denote significance, respectively, at 10%; * 5% and 1% confidence level.

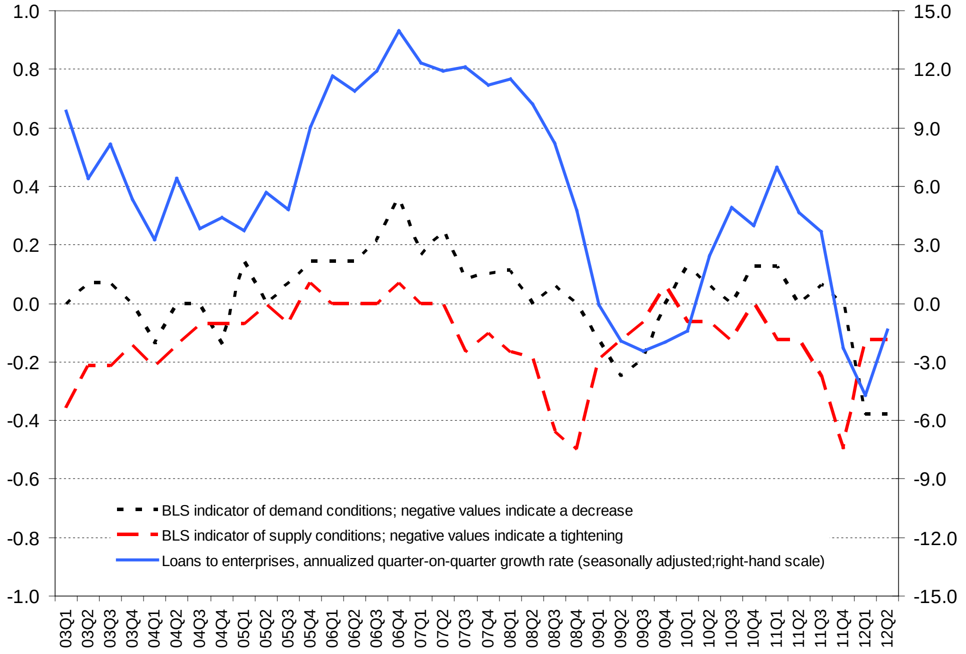
The *"Under-identification test"* is a Lagrange-Multiplier test aiming at check if the equation is identified. The null hypothesis is that the equation is under-identified (i.e. the matrix of reduced form coefficients on the L1 excluded instruments has rank K1-1), while the alternative hypothesis is that the equation is identified (i.e. the matrix has rank exactly equal to K1). Under the null hypothesis, the test statistic is distributed as chi-squared in (L1-K1+1) degrees of freedom. A rejection of the null indicates that the equation is identified. The *"Weak identification test"* is the F-version of the Cragg-Donald Wald statistic. The *"Over-identification test"* is based on the Sargan-Hansen statistics; the null hypothesis tested is that the exclusion restrictions are valid. Under the null hypothesis, the test statistic is distributed as chi-squared in the number of (L-K) over-identifying restrictions. A rejection casts doubt on the validity of the instruments used for the equation identification.

Figure 1

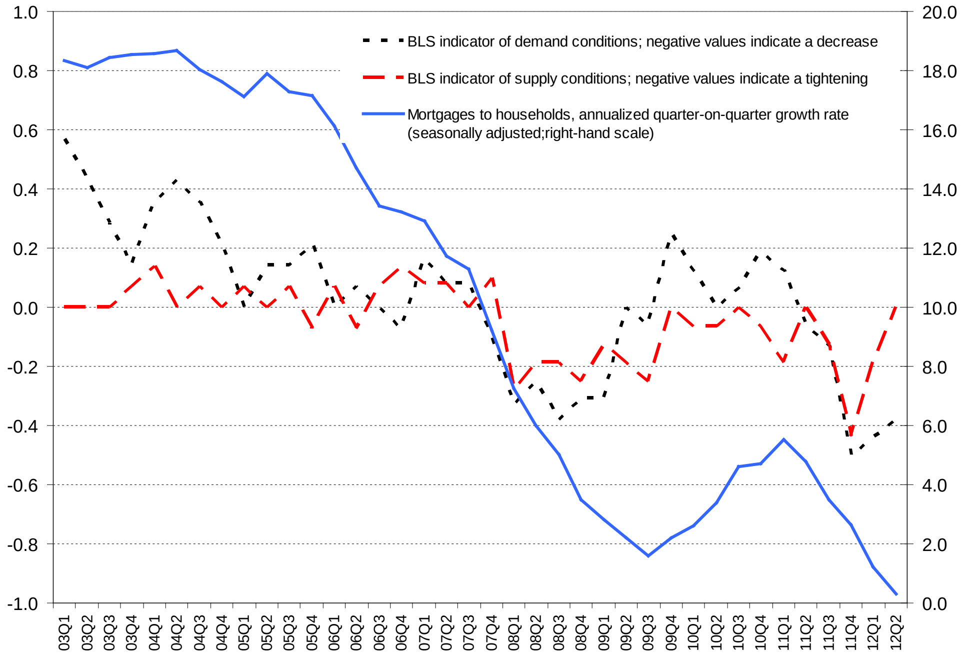
BLS supply and demand indicators and lending dynamics in Italy

(quarterly data; percentage points; diffusion indexes)

a) Loans to enterprises



b) Mortgage loans to households



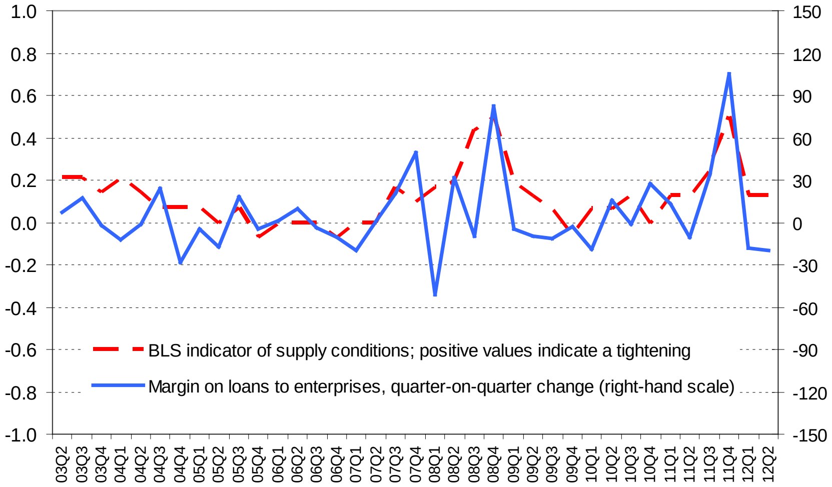
Source: Bank of Italy; the euro area bank lending survey.

Fugure 2

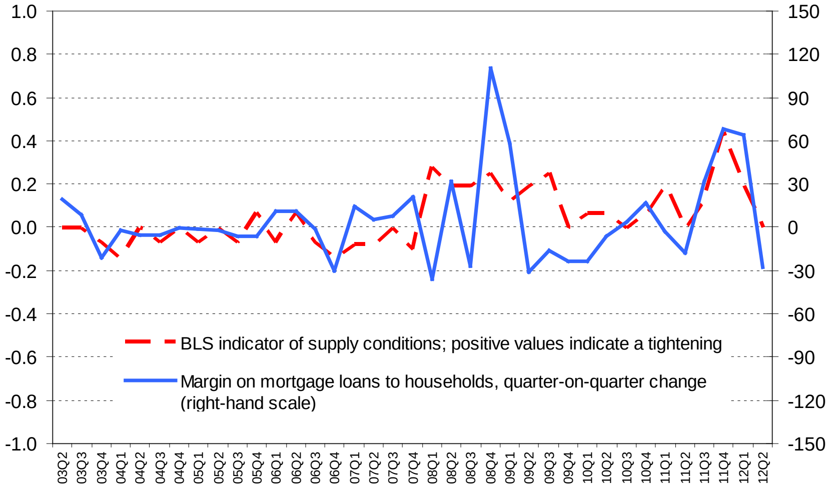
BLS supply indicator and margins on new loans

(quarterly data; basis points; diffusion indexes)

a) Loans to enterprises



b) Mortgage loans to households



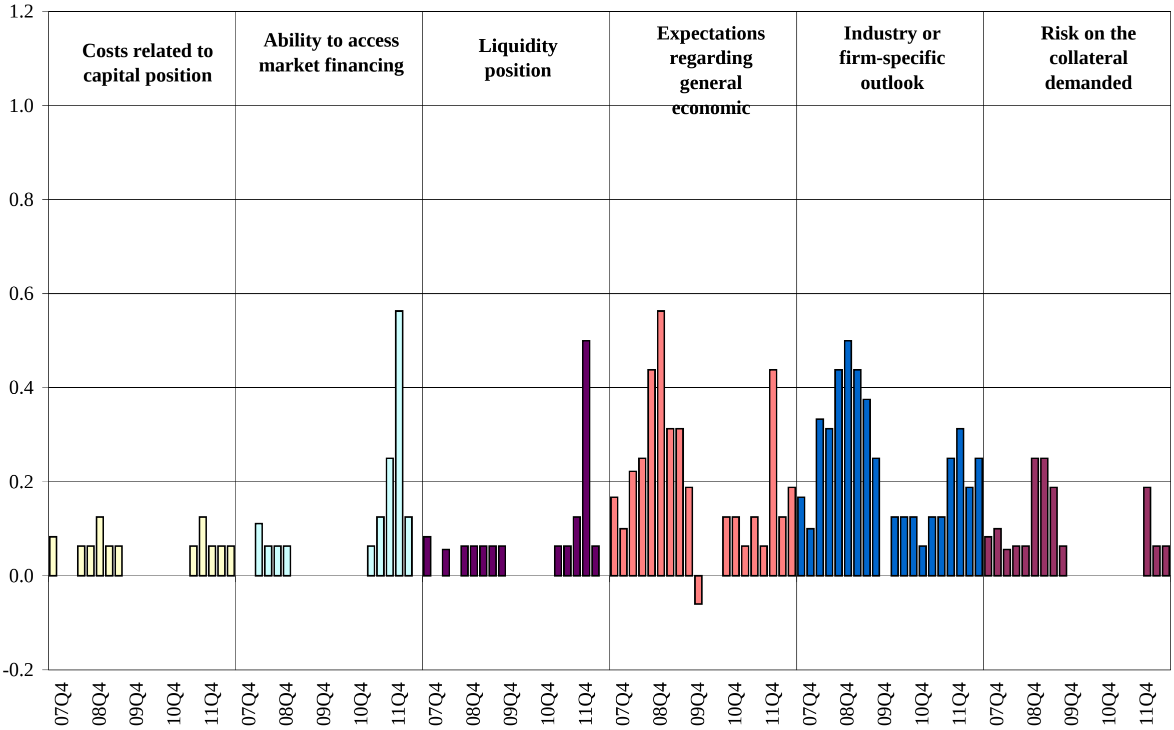
Source: Bank of Italy; the euro area bank lending survey.

Figure 3

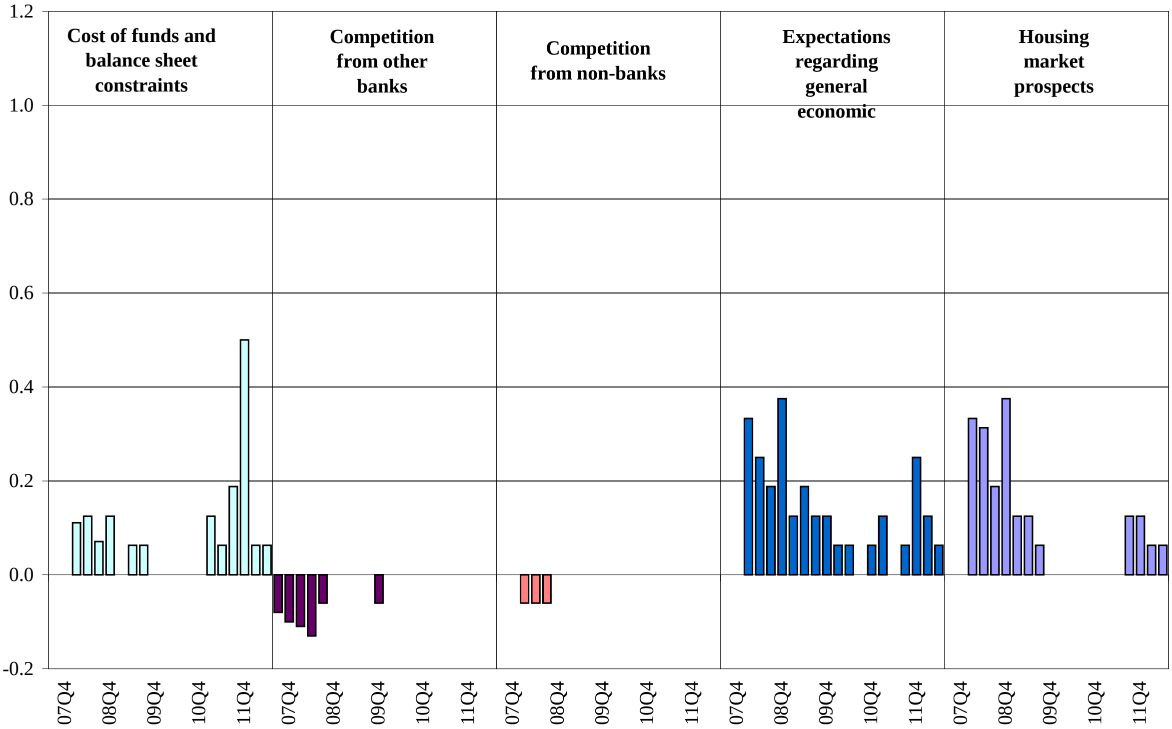
Factors behind changes in credit supply conditions in Italy

(diffusion indexes)

a) Loans to enterprises



b) Mortgage loans to households

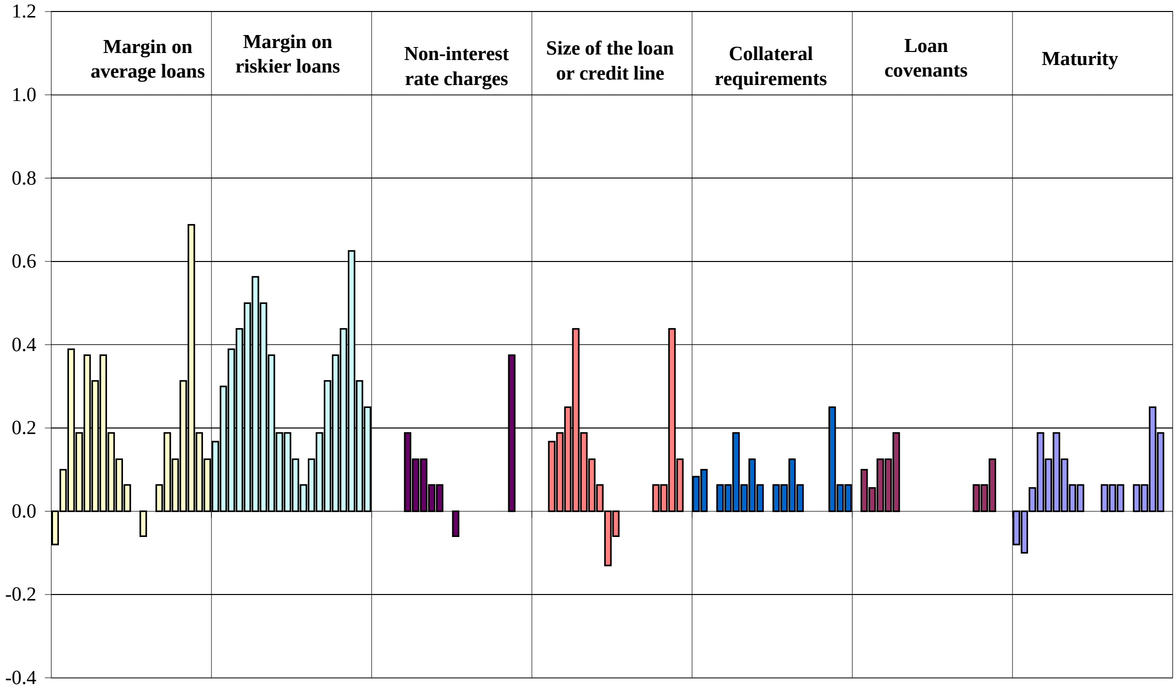


Source: The euro area bank lending survey.

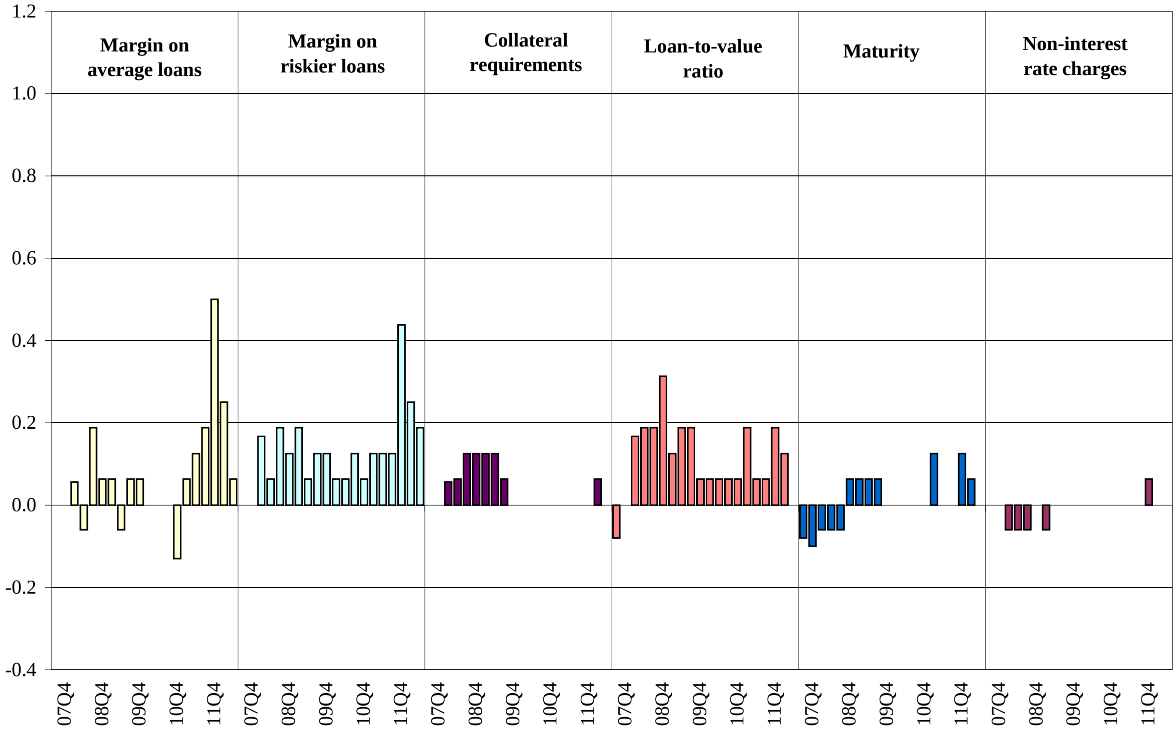
Notes: Positive values indicate supply restriction compared with the previous quarter. Diffusion indices are constructed on the basis of the following weighting scheme: 1 = contributed considerably to a tightening, 0.5 = contributed somewhat to a tightening, 0 = contributed to basically unchanged, -0.5 = contributed somehow to an easing, -1 = contributed considerably to an easing. The range of variation of the index is from -1 to 1.

Figure 4
Terms and conditions behind changes in credit supply conditions in Italy
(diffusion indexes)

a) Loans to enterprises



b) Mortgage loans to households



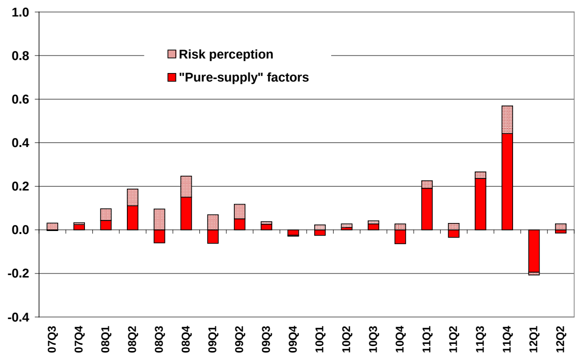
Source: The euro area bank lending survey.

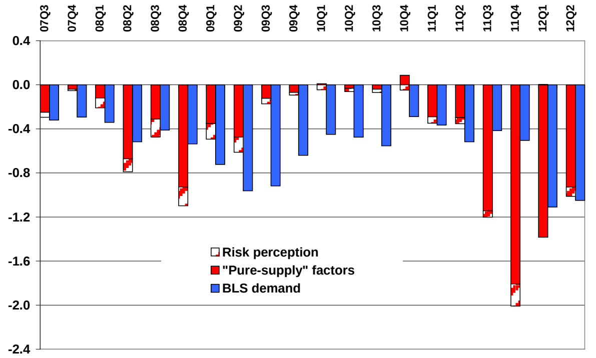
Notes: Positive values indicate supply restriction compared with the previous quarter. Diffusion indices are constructed on the basis of the following weighting scheme: 1 = tightened considerably, 0.5 = tightened somewhat, 0 = remained basically unchanged, -0.5 = eased somewhat, -1 = eased considerably. The range of variation of the index is from -1 to 1.

Figure 5
Estimated effects of supply and demand indicators on the cost and the amount of loans to enterprises
(quarterly data; percentage points)

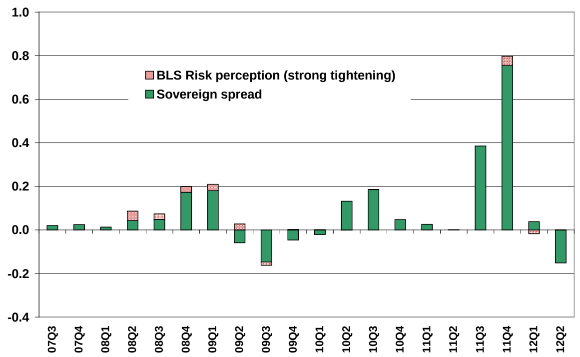
i) Cost of loans: q-o-q change in bank margin on new loans	ii) Amount of loans: q-o-q growth rate

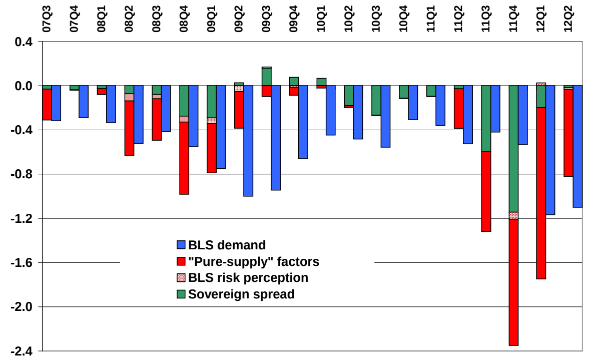
Using BLS supply factors
(specification (c)-(c') of Table 4)



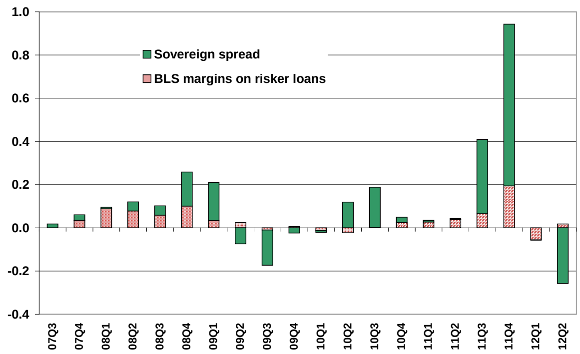


Using BLS supply factors and the sovereign spread
(specification (a)-(a') of Table 6)





Using BLS terms and conditions and the sovereign spread
(specification (c)-(c') of Table 8)



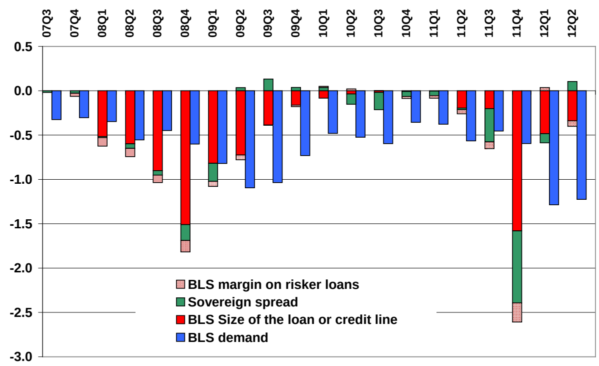


Figure 6

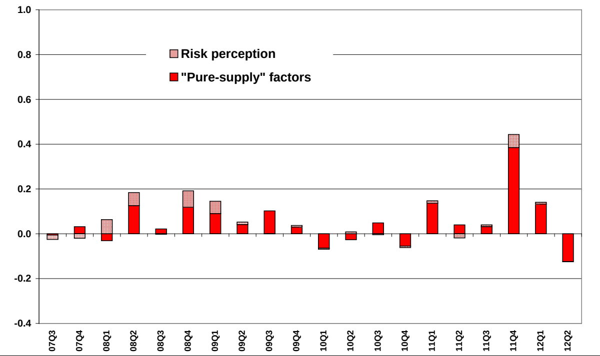
Estimated effects of supply and demand indicators on the cost and the amount of mortgage loans to households

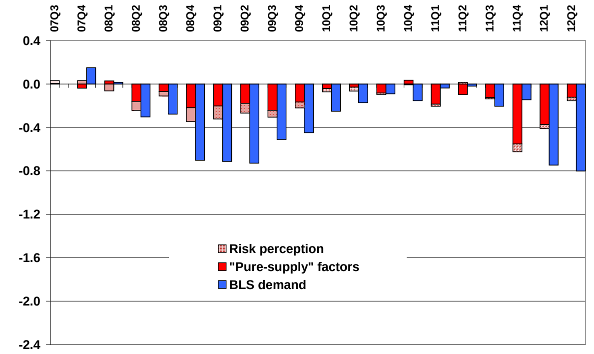
(quarterly data; percentage points)

i) Cost of loans: q-o-q change in bank interest rate margin	ii) Amount of loans: q-o-q growth rate

Using BLS supply factors

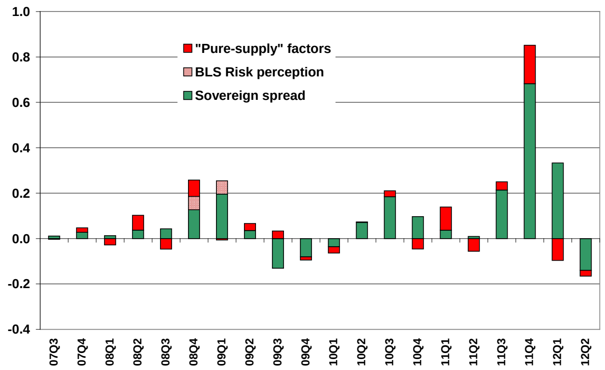
(specification (c)-(c') of Table 5)

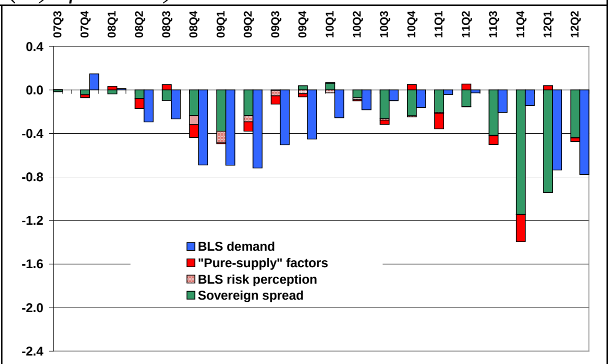




Using BLS supply factors and the sovereign spread

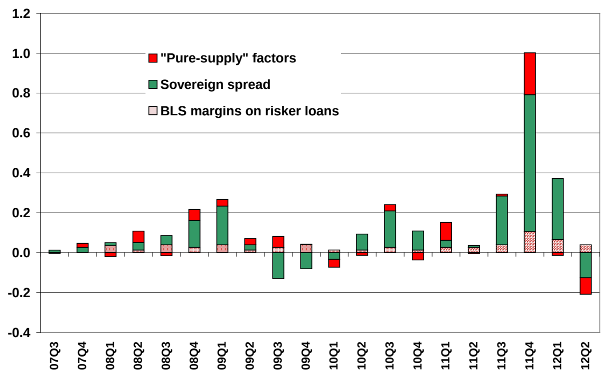
(specification (a)-(a') of Table 7)

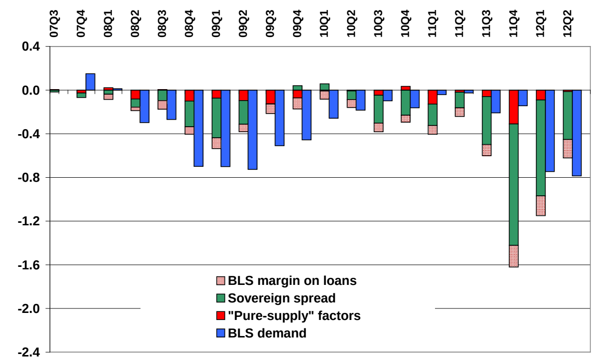




Using BLS terms and conditions and the sovereign spread

(specification (b)-(b') of Table 9)



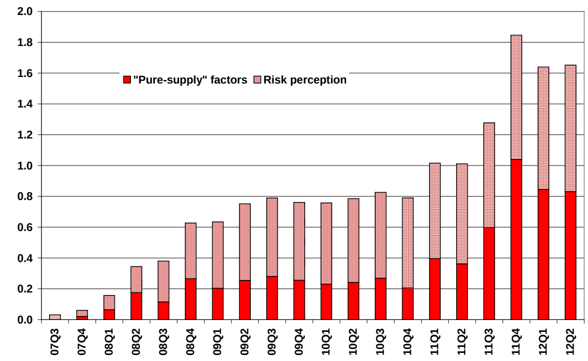


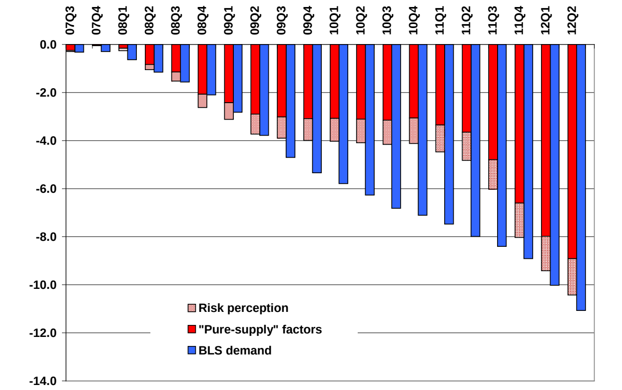
Appendix

Figure A1
Cumulated effects of BLS supply factors and demand indicators on the cost and the amount of loans to enterprises
(quarterly data; percentage points)

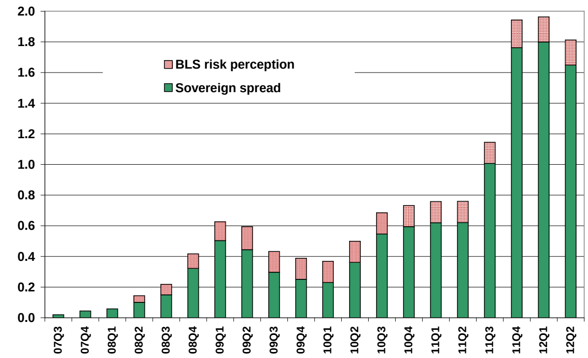
i) Cost of loans: q-o-q change in bank interest rate margin	ii) Amount of loans: q-o-q growth rate

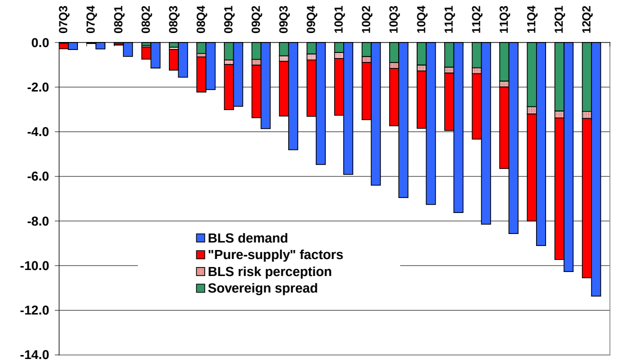
Using BLS supply factors
(specification (c)-(c') of Table 4)



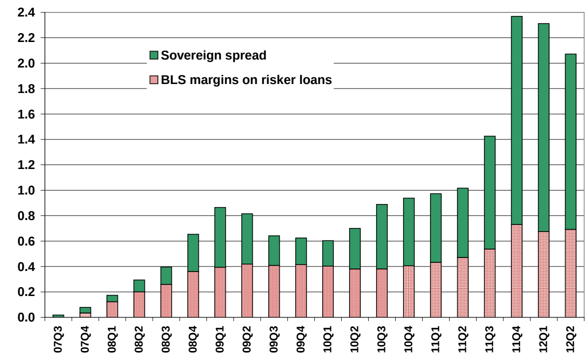


Using BLS supply factors and the sovereign spread
(specification (a)-(a') of Table 6)





Using BLS terms and conditions and the sovereign spread
(specification (c)-(c') of Table 8)



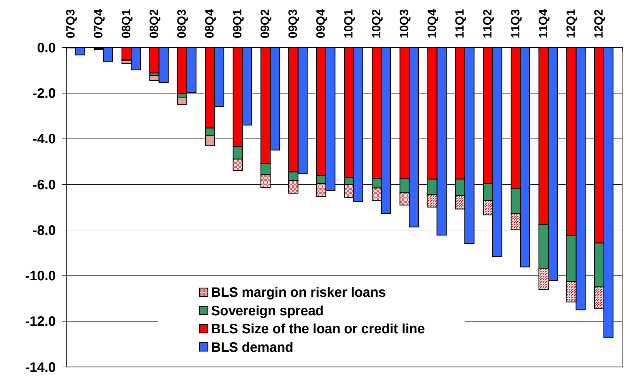
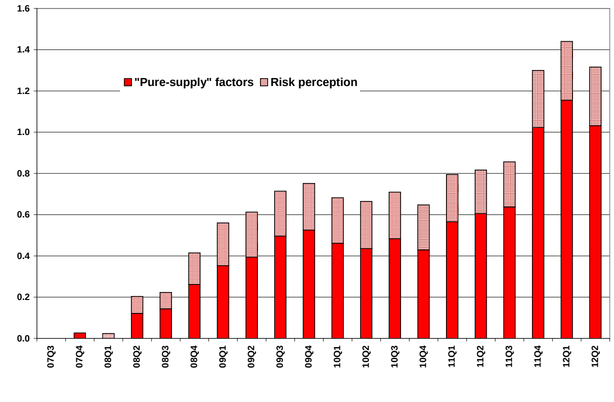
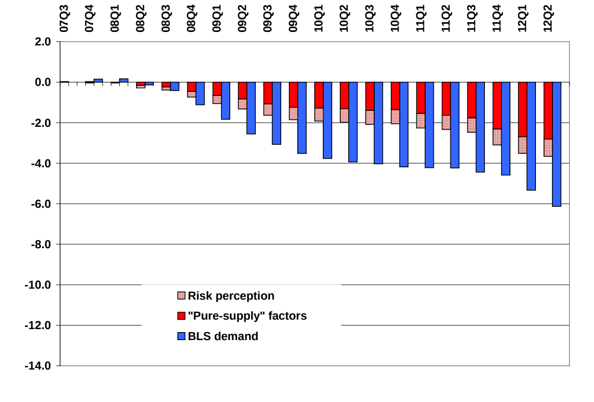


Figure A2
Cumulated effects of BLS supply factors and demand indicators on the cost and the amount of mortgage loans to households
(quarterly data; percentage points)

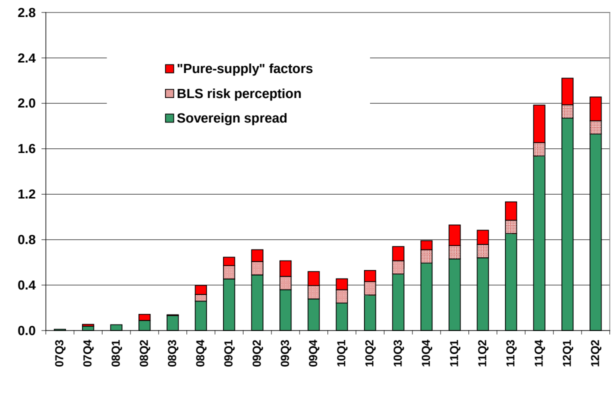
i) Cost of loans: q-o-q change in bank interest rate margin	ii) Amount of loans: q-o-q growth rate

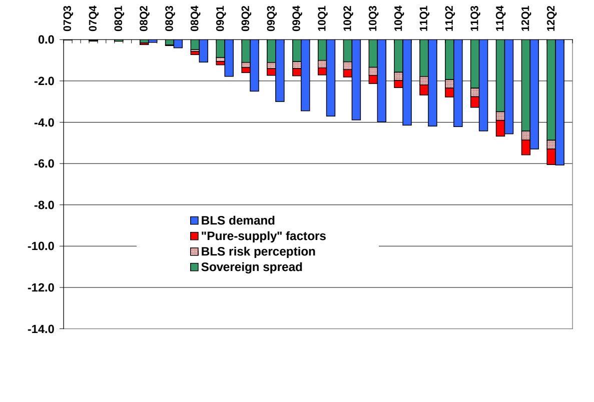
Using BLS supply factors
(specification (c)-(c') of Table 5)



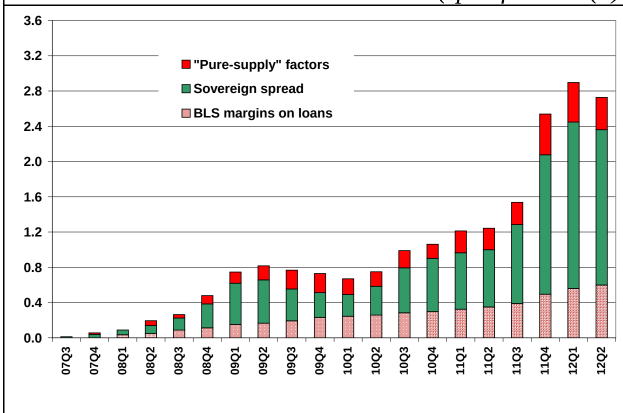


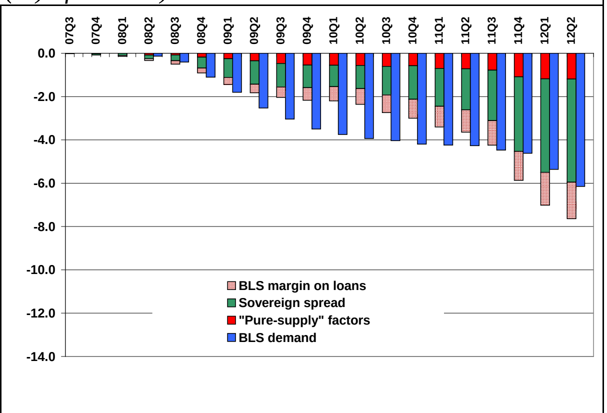
Using BLS supply factors and the sovereign spread
(specification (a)-(a') of Table 7)





Using BLS terms and conditions and the sovereign spread
(specification (b)-(b') of Table 9)





Credit supply during a sovereign debt crisis

Marcello Bofondi*, Luisa Carpinelli* and Enrico Sette*

April 2013

Abstract

We study the effect of the increase in Italian sovereign debt risk on credit supply on a sample of 670,000 bank-firm relationships between December 2010 and December 2011, drawn from the Italian Central Credit Register. To identify a causal link, we exploit the lower impact of sovereign risk on foreign banks operating in Italy than on domestic banks. We study firms borrowing from at least two banks and include firm x period fixed effects in all regressions to controlling for unobserved firm heterogeneity. We find that Italian banks tightened credit supply: the lending of Italian banks grew by about 3 percentage points less than that of foreign banks, and their interest rates were 15-20 basis points higher, after the outbreak of the sovereign debt crisis. We test robustness by splitting foreign banks into branches and subsidiaries, and then examine whether selected bank characteristics may have amplified or mitigated the impact. We also study the extensive margin of credit, analyzing banks' propensity to terminate existing relationships and to grant new loan applications. Finally, we test whether firms were able to compensate for the reduction of credit from Italian banks by borrowing more from foreign banks. We find that this was not the case, so that the sovereign crisis had an aggregate impact on credit supply.

JEL Classification: G21, F34, E44, E51.

Keywords: credit supply, sovereign debt crisis, multinational banks, bank lending channel.

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1 Introduction[1]

Since the outburst of the 2011 sovereign debt crisis, much debate has revolved around the impact that increased country risk could have on financial intermediaries' balance sheets, in particular on their funding costs and on their capacity to grant credit to firms and households for investment and consumption.

As sovereign bonds yields raise and sovereign ratings deteriorate, sources of funding become indeed more scarce and more costly: availability of wholesale funding markets, especially uncollateralized, becomes much thinner and banks' capacity to access collateralized lending decreases, as the value of eligible collateral, typically sovereign bonds, drops. Moreover, bank profitability may be reduced, in particular if sovereign bonds are held in banks' trading books which are marked-to-market. These factors all contribute to transmit tensions from the sovereign bond markets to banks' ability to supply credit and to the cost of credit for borrowers. Hence, a credit crunch may occur at a time in which governments may tighten fiscal policy to combat the sovereign tensions, triggering or amplifying a contraction in economic activity. Finally, higher sovereign yields may also impair the transmission mechanism of monetary policy, in particular within a monetary union: policy rate changes may not affect banks funding costs if the latter are increasingly driven by domestic sovereign yields.

Despite its relevance, there is limited empirical evidence on the direct and causal impact that sovereign shocks exert on credit supply. Identifying this effect is indeed particularly challenging, since banking and sovereign crisis tend to be intertwined, reinforcing each other through strong feedback effects (Reinhart and Rogoff 2009, Acharya et al. 2012).

First, it is difficult to isolate an exogenous sovereign shock: typical patterns suggest that sovereign debt crises are fuelled by banking crises, as governments disburse vast amounts of money to rescue troubled intermediaries. Second, sovereign and banking crises are often accompanied by recessions, when demand for credit typically drops, thus making difficult to disentangle supply from demand effects.

In this paper we overcome these identification challenges thanks to the nature of the shock and the richness of our data.

The outburst of the sovereign crisis in Italy was fairly exogenous with respect to the lending policies of Italian banks. Both low growth and high public debt are long-standing features of the Italian economy. The Italian banking system did not represent a source of instability for public finances (see, among others, IMF 2010 Article IV consultation on Italy) and Italy did not experience a housing bubble. Italian sovereign spreads increased sharply since the beginning of July 2011, without any specific domestic event triggering it: the stalemate in negotiations on Greek sovereign debt fuelled fears of a break-up of the Euro-area which were transmitted

[1] We thank Giorgio Albareto, Martin Brown, Elena Carletti, Nicola Cetorelli, Federico Cingano, Olivier De Jonghe, Domenico Depalo, Linda Goldberg, Giorgio Gobbi, Giuseppe Ilardi, Silvia Magri, Francesco Manaresi, Tommaso Oliviero, Steven Ongena, Alberto Pozzolo, Joao Santos, Koen Schoors, Neeltje Van Horen, participants at the 2012 CREDIT conference, at the workshop on "Macroeconomic policies, global liquidity, and sovereign risk", at the "6th CEPR Swiss Winter Conference in Financial Intermediation", at the "Third Mofir Workshop", at the 20th Finance Forum, at the Workshop "The Sovereign Debt Crisis and the Euro Area", seminar participants at the Bank of Italy and at the New York Fed, and an anonymous referee for helpful comments.

to Italian sovereign yields, while those of "core" European countries remained stable. Then, adopting a quasi-experimental methodology, we exploit the sudden and sharp increase in the yield on Italian sovereign debt of July 2011. The semester between December 2010 and June 2011 represents the pre-crisis period, and the one between June and December 2011 represents the crisis period.

Our data of about 670,000 bank-firm relationships from the Italian Credit Register allow us to properly distinguish supply from demand. We restrict our analysis to firms borrowing from at least two banks. In this way we fully control for firm observed and unobserved heterogeneity by plugging firm-fixed effects (with a methodology akin to that pioneered by Khwaja and Mian (2008)).

To identify the effect of the sovereign shock, we need to compare lending to the same firm by two or more banks that have been affected by the crisis to a different degree. To define "more" and "less" affected banks, we exploit the presence of foreign banks in the Italian market. Foreign banks, being headquartered in countries where the sovereign risk increased significantly less, were indeed way more shielded by the impact of sovereign tensions than Italian banks. Since the variation of the shock was primarily across countries, we believe that the heterogeneity between Italian and foreign banks is the dimension that most appropriately captures the differential impact of the shock. Although not fully insulated by the shock, foreign banks provide a good counterfactual to assess how the rise in sovereign spreads modifies credit supply decisions.

We find significant evidence of credit restrictions after the sovereign crisis. Italian banks decreased credit and increased interest rates charged to non-financial firms more than foreign banks. These results are confirmed if we use the change in the spread between yields on 10-year government bonds of headquarter's country and the German Bund of corresponding maturity to measure more directly the increase in funding cost by bank's nationality.

We also examine if the sovereign crisis had an impact on the extensive margin of credit. To this aim, we test whether Italian banks terminated relationships and rejected new loan applications more than foreign banks, as the risk on the Italian sovereign increased. We find that the sovereign debt crisis reduced the willingness of Italian banks to terminate existing relationships, whereas they drastically decreased the probability of accepting new applications. We also test if domestic banks charged higher interest rates on new term loans than foreign banks, and we find that this is the case.

Having found that there has been a significant credit tightening of Italian banks vis à vis foreign banks, we test whether this effect is in fact driven by bank characteristics that might have changed over time at a different extent across Italian and foreign banks. We then estimate the baseline model including a set of bank balance sheet characteristics: bank capitalization (the Tier 1 ratio), bank size, the ratio of sovereign securities from European troubled countries (GIIPS) to total assets, and the ratio between wholesale funding and total assets. The last two variables are especially important because they capture the extent to which banks might be affected by the sovereign crisis. We find that the interaction between the dummy domestic and the dummy crisis is still significant and its coefficient is of similar size as in the baseline regression; furthermore no bank variable is statistically significant. Therefore, there seems to be a country-specific effect common to Italian banks: even if they had the same capital position and funding structure as foreign banks, they would still be tightening credit to a larger extent.

Finally, we test whether firms were able to compensate the reduction of credit by Italian banks through increased credit from foreign banks. We estimate an aggregate effect of the sovereign shock on credit supply to Italian firms. We obtain an unbiased estimate of this aggregate effect by plugging firm effects estimated from our baseline regression at the bank-firm relationship-level into a firm-level equation in which the dependent variable is the growth of credit granted to firms by the full set of lending banks. Our results suggest that firms have not been able to fully substitute credit from domestic banks with credit from foreign banks. The sovereign crisis has therefore had a negative aggregate impact on credit supply.

The paper is structured as follows: the next section examines the related literature, section 3 discusses the empirical strategy, section 4 presents the dataset and the main descriptive statistics, section 5 contains the results of our baseline specification, section 6 illustrates results on alternative versions of baseline, section 8 examines the extensive margin, section 8 explores bank heterogeneity, section 9 presents the result on the aggregate effect, section 10 concludes.

2 Related Literature

Our contribution is related to three streams of literature. First, we contribute to the studies on the real effects of sovereign debt crises and sovereign defaults. Arteta and Hale (2008) examine how access to foreign credit to the private sector varies during sovereign debt crises. They group micro-level data on bond issuance and foreign syndicated bank loan contracts of firms into different export and non-export sectors. They find systematic evidence of a decline in foreign credit over the period between 1984 and 2004 for 30 emerging markets in the aftermath of a sovereign debt crisis. Borensztein and Panizza (2009) investigate whether default episodes give rise to a credit crunch: using industry-level data available for 149 countries over the period 1975-2000, they test whether defaults have a significantly larger effect on sectors that are more heavily dependent on external finance. Their results indicate that defaults have a limited impact on credit supply. Furceri and Zdzienicka (2011) evaluate the overall losses in terms of output that debt crises exert over the short and medium term, on a panel of 154 countries from 1970 to 2008. They find that the effects are sizeable, both the contemporaneous ones (6 percentage points) and those observed in the medium term over a 10 year horizon (up to 10 percentage points of GDP). De Paoli et al. (2009) look at the effects of debt crises on output; running a counterfactual analysis on 40 episodes of sovereign debt crises they also find that the output losses are prolonged and large. Yet, reductions in output seem to be significantly more pronounced when debt crises are associated with a banking and/or currency crises, which occur for over half of the crises in the sample. Albertazzi et al. (2012), in contemporaneous work on Italian data take a macro perspective. They run bank-level regressions of the volume of outstanding loans and of the level of interest rates on the level of the BTP-Bund spread. They find that a rise in the spread is followed by an increase in the cost of credit to firms and households, and by a reduction in lending growth.

We contribute to this literature by evaluating how the recent sovereign debt crisis, by increasing banks' funding cost, has been transmitted to bank lending, in terms of both quantities and prices. The originality of our contribution lies in three aspects. First, we study the effect of an episode of sovereign debt crisis that can be considered fairly exogenous with respect to the banking sector; in this way we are able to isolate the effect of sovereign tensions from the often concurring banking crises. Moreover, we provide evidence about a sovereign crisis affecting a developed country which is part of a currency union, where the risk of a currency crisis is basically non-existent and monetary policy is determined by all member countries. As a consequence, the analysis of the sovereign crisis in Italy represents an ideal laboratory for studying the impact of sovereign tensions on credit supply. Secondly, by relying on a unique dataset on bank-firm relationships, we are able to fully control for firm-level unobserved heterogeneity, thus isolating the impact of supply from the impact of demand factors and properly addressing the endogeneity issues that typically challenge the studies of the effect of financial crises based only on macro or bank-level data. Third, we concentrate on the initial phase of a sovereign crisis, and not of a country sovereign default. This allows us to zoom into the mechanisms that drive the transmission of sovereign tensions to the real sector, thus feeding back into larger public deficits.

Second, our paper is also broadly related to the literature on global banks and on the international transmission of shocks. This literature has mostly focussed on how foreign banks might have contributed to "export" tensions affecting the domestic market, thus highlighting a mechanism of international transmission of shocks. In their seminal papers, Peek and Rosengren (1997, 2000) examine the impact of the fall of Japanese stock prices of the 1990s on cross-border lending by Japanese banks. They show that Japanese bank branches operating in the U.S. tightened their credit supply. Popov and Udell (2010), based on survey data on SME financing on 14 CEE countries in the period 2005-2008, find evidence of international transmission of financial distress in the early stage of the crisis, with Western European banks restricting credit supply more than domestic banks. Cetorelli and Goldberg (2011) show that the transmission of shocks spurred by global banks to emerging economies in the 2007-2009 crisis was large. Using bilateral country-level data they show that the impact occurred not only through contraction of cross-border loan supply by foreign banks and foreign banks' affiliates, but also by domestic banks that suffered a funding shock due to the reduction of inter-bank cross–border lending. Schnabl (2012) examines the impact that a negative liquidity shock to international banks such as the 1998 Russian default had on credit to Peruvian firms. Using bank-level data, he finds that the impact was significant. The transmission of the shock occurred through foreign inter-bank funding and the effect was strongest for domestic firms that were borrowing internationally. Analyzing data on cross-border syndicated lending by 75 banks to 59 countries over the period 2000-2009, De Haas and Van Horen (2012) find that banks that were more severely affected by funding constraints have reduced their lending abroad significantly. Finally Kalemli-Ozcan et al. (2012) take a broader perspective and show that during the 2007-2009 crisis the impact of financial integration on output cycles has changed as opposed to the period 1970-2007: whereas before 2007 tighter financial linkages were associated with more divergent output cycles, in more recent years they were correlated with greater synchronization.

Our paper contributes to this field since, as a tool for identification of the effect of a sovereign shock on credit supply, it compares the patterns of credit granted by domestic and foreign banks. Hence we provide evidence on the lending policy of foreign banks in a country hit by a sovereign crisis showing that the presence of foreign banks may mitigate the impact of sovereign tensions on the supply of credit to domestic firms.

Third, from a methodological point of view, our paper relates to the empirical literature on the bank lending channel that uses credit registry data. Khwaja and Mian (2008) study the impact of an unexpected liquidity shock on credit supply on Pakistani data. They find that banks more exposed to the liquidity shock contracted their supply of credit more. Their paper also makes an important methodological contribution since they propose to control for firm-level unobserved characteristics including firm fixed effects. Jimenez, Ongena, Peydrò, Saurina (2011) and (2012) apply a similar technique to identify the banks' balance sheet channel of monetary policy and to study the effect of monetary policy on banks' risk taking.[2]

3 Empirical strategy

3.1 Issues for identification

Identifying a causal effect of sovereign tensions on credit supply poses important challenges.

First, the shock has to be exogenous with respect to the conditions of domestic banks. Yet sovereign spreads may rise as a consequence of a deterioration in domestic banks' balance sheets, or of the burst of an asset price bubble, which induces governments to bail out financial intermediaries (Acharya et al. 2012 show that government bail-outs of banks lead to higher sovereign spreads). We argue that this was not the case in Italy. During 2010 increasing concerns on the sustainability of public finances in Greece, Ireland and Portugal eventually led these countries to ask for international assistance from the European Union and the International Monetary Fund. Risk premiums on interbank and bond markets rose. Italian banks experienced an increase in the cost of wholesale funding, but their condition was not far from the one of their European peers. The situation changed dramatically from the June 2011, when rapidly deteriorating Greek economic conditions fuelled fears of a Euro-area break-up and triggered contagion to Italy. Between June and July 2011, indeed, S&P downgraded the Greek debt to CCC, the lowest rating for any country it reviews, Greek political instability rose, and announcements of an involvement of the private sector in Greek debt restructuring were made, characterizing it as a "selective default". Fearing that these events might have an impact on Italian sovereign risk, spreads on Italian government debt rose abruptly. Fig. 1 shows the magnitude of the increase in sovereign spreads on Italian 10 year government bonds with respect to the benchmark 10 year German Bund. All the action is concentrated in the second part of

[2] Other papers use a broadly similar identification strategy on Italian data: Bonaccorsi and Sette (2012) who study the bank lending channel during the 2007-2008 crisis and Albertazzi and Marchetti (2010) who study the presence of evergreening by banks after the Lehman default.

2011, when spreads increased sharply since June, reaching 370-390 basis points in September 2011 and a peak of 530 basis points in November. As opposed to what happened in other European countries the increase in sovereign yields can not be attributed to the instability of the financial sector. The weakness of Italian public finances is in fact driven by the high level of public debt and the low growth rate of the economy, which are both long standing features of the Italian economy (Bank of Italy 2011). Moreover, as opposed to what happened in Ireland or Spain, state aid to the banking sector was extremely limited and did not impact significantly on public deficit (see OECD 2009 among others). Fig. 2 shows primary net borrowing as a percentage of GDP of Greece, Ireland, Italy, Portugal and Spain. Public finances deteriorated markedly since 2008 in Ireland, Spain, Greece and Portugal, also as a consequence of bail-outs of troubled domestic banks. By contrast, primary deficit did not change much in Italy, also because the Italian financial sector needed little support to weather the crisis, and the high level of sovereign debt left little room to use fiscal policy to counteract the recession.[3] Finally, Italy did not experience an housing bubble.

On the contrary, increasing sovereign yields did have consequences on the banking system. The CDS spreads on the senior debt of the largest Italian banks rose abruptly leading to increasing difficulties in raising funds in the wholesale markets and rising interest rates on retail funding. The surge in the CDS spread was significantly higher than the one experienced by intermediaries in other developed countries (Fig. 3). Therefore the end of June 2011 can be reasonably identified as the moment in which the Italian banking system was hit by an unanticipated exogenous shock.[4]

A second crucial issue for identification is that sovereign tensions are accompanied by deteriorating economic conditions, inducing firms to scale down their investment plans and decrease demand for credit. Moreover, banks more exposed to sovereign tensions may lend to a different set of firms (e.g. firms with weaker balance sheets, riskier firms, etc.) than banks less exposed to sovereign tensions. Hence, it is critical to properly control for firm level demand for credit, for firms' riskiness, and, more generally, for firm unobserved heterogeneity. The richness of our dataset allows us to do so. Since Italian firms typically resort to multiple lenders (Detragiache et al. 2000, more recently Gobbi and Sette 2011), we identify the impact of sovereign risk on credit supply by comparing the pre-crisis and the crisis patterns of credit supplied to the same firm by two or more banks that have been affected by the sovereign crisis to different degrees. The inclusion of firm-period fixed effects in all regressions, similarly to what Khwaja and Mian (2008) or Jimenez et al. (2012) do, enables us to control for all firm-level unobserved heterogeneity that affects the dynamics of credit granted and of its cost in each period.

Another key condition for estimation of a supply effect is to identify banks, otherwise comparable, that have been differently affected by the shock. Since sovereign tensions were primarily country-specific, we consider Italian banks as the "more affected" group and foreign banks as the "less affected" one. The cross-country variability in the exposure to the shock is indeed quite

[3] Results are qualitatively similar if we use net government borrowing including interest expenses.

[4] Later developments during 2012 may discount deterioration in banks' access to funding, firms' profitability caused by the recession and government measures taken in the Autumn of 2011.

large. In particular, branches and subsidiaries of foreign banks, which hold 8% and 9% of total banking assets, were largely shielded from the Italian sovereign shock. Their lower exposure to the increased risk on the Italian government debt is attributable to a number of reasons. First of all, foreign banks operating in Italy are headquartered in countries where the sovereign risk was more contained, therefore making the chances of a downgrade on banks transmitted by lower ratings on domestic government debt limited. Second, given that the assets portfolio of foreign banks is less concentrated in government bonds of peripheral countries vis à vis Italian banks - holding mostly Italian debt-, the increase in riskiness of their asset side due to sovereign risk over the second half of 2011 was relatively milder. Third and most importantly, although lending to Italian firms, a significant fraction of their liabilities, 70% for branches and 40% for subsidiaries, are represented by interbank transfers from their headquarters that raise funds either in their home country or in the international wholesale markets. This contributed to a much lower increase in funding cost for foreign banks.

Since foreign banks cannot be considered as fully insulated by the sovereign shock, the effect we identify in the paper should be interpreted as a lower bound for the full causal impact of the crisis on lending, given that foreign banks do also tighten credit supply as a consequence of the shock, though modestly. As a robustness check, we also estimate a model using a continuous measure of the impact of the sovereign shock (the change in sovereign spread of the country where the banks is headquartered) which provides results that are both qualitatively and quantitatively consistent with those of the model comparing domestic and foreign banks.

In principle domestic and foreign banks may be different along several dimensions, and comparing them to assess the effect of the increase in sovereign spreads on credit supply may not be warranted. We argue that this is not the case for a number of reasons. First, the Italian banking system is rather sophisticated and Italian banks, especially larger banks, have similar business models, lending technologies, geographical scope as foreign banks, especially subsidiaries. Second, our identification strategy based on comparing lending by different banks to the same firm, allows us to fully control for possible differences in the composition of borrowers across domestic and foreign banks. Moreover, firms borrowing from very different types of banks, e.g. a domestic mutual bank, and a large international group, are rare. Last, but not least, we include bank fixed effects in our regressions, so that we can control for all unobserved heterogeneity among lenders, including notably differences in the ex-ante composition of loan portfolios, lending policies, extension of the network of outlets, etc.

3.2 The model

To identify the effect of the sovereign crisis on credit supply we estimate a model in which the observational unit is a credit relationship between a firm and a bank, and we compare two periods, the first half of 2011 (pre-crisis) and the second half of 2011 (crisis). Using a pre-crisis period allows to control for pre-crisis differences in the supply of credit by Italian and foreign banks. Moreover, it also allows us to include bank fixed effects to control for bank time-invariant unobservables.

The main models we estimate are as follows:

$$\Delta credit_{i,j,t} = \beta_1 domestic_j + \beta_2 domestic_j * crisis_t + \alpha_{i,t} + \varepsilon_{i,j,t} \quad (1)$$

$$\Delta APR_{i,j,t} = \gamma_1 domestic_j + \gamma_2 domestic_j * crisis_t + \alpha_{i,t} + \varepsilon_{i,j,t} \quad (2)$$

where $\Delta credit_{i,j,t}$ is the difference in the log credit granted by bank j to firm i in period t, and $\Delta APR_{i,j,t}$ is the change in the Annual Percentage Rate charged by bank j to revolving credit lines and to term loans granted to firm i in period t[5]. The dummy *domestic* equals 1 if bank j is Italian, zero if the bank is foreign, either as a branch or a subsidiary. The term $domestic * crisis$ is an interaction between the dummy *domestic* and the dummy variable *crisis* which equals 1 in the second half of 2011. We also include a full set of firm-period fixed effects, $\alpha_{i,t}$, which control for firm level unobserved heterogeneity in each period (including firm level demand for credit, firm balance sheet conditions, etc.). These fixed effects also absorb the dummy *crisis*, which therefore does not appear in the equations above. The effect is identified on firms that borrow from at least one Italian and one Foreign bank in at least one period.[6] We also run all regressions including bank fixed effects, which control for all bank time invariant unobserved heterogeneity, including systematic differences in banks' business models, geographical reach, etc.[7] Our focus is on the parameters β_2 and γ_2 which capture the differential behavior of Italian banks relative to foreign banks during the crisis.

All regressions also include variables intended to capture the specificity of the relationship between firm i and bank j. The first one is the share of total credit to firm i supplied by bank j (SHARE OF TOTAL CREDIT). Ex ante its expected sign is ambiguous: on the one hand, this variable measures the relative exposure of bank j towards firm i, and this is negatively correlated with loan growth and positively correlated with the change in the interest rate; on the other hand it could be interpreted as a proxy of the strength of the bank-firm relationship, therefore suggesting a positive relationship with credit quantities and possibly negative with interest rates. Moreover SHARE OF TOTAL CREDIT can also partially account for the initial size of the loan. The second variable is the share of drawn over credit granted by bank j to firm i (DRAWN OVER GRANTED). This control measures how intensively available credit lines are used. The third variable is the share of overdraft over total granted credit by bank j to firm i (OVERDRAFT). This regressor aims at controlling for the composition of total credit by different types of loan contracts (term loans, overdrafts, loans backed by account receivables).

[5] The reference rate for loans to non-financial corporations in Italy is the Euribor. In the case of revolving credit lines, this is the 1-month Euribor. Its movements are absorbed by firm*period fixed effects, so that our analysis, at least in the case of revolving credit lines, captures the effects of the sovereign crisis on spreads on loans to non-financial corporations. In the case of term loans, this is made more complicated by the lack of detailed data on the maturity of the loan (we only know whether its maturity is above or below 2 years).

[6] Suppose firm 1 borrows from Italian bank A, and Foreign bank B at June 2011. Our identification compares credit growth (and the interest rate changes) between June and December 2011 by bank A and B to the same firm 1. Then, we also add a pre-crisis period (December 2010-June 2011) to take care of possible different dynamics in credit supply by Italian and Foreign banks, but having repeated observations for the same firm-bank pairs is not strictly necessary for identification purposes.

[7] When we include bank fixed effects they absorb the dummy domestic, as no bank changes status (from domestic to foreign or viceversa) in our sample period.

3.3 Issues for empirical strategy

A key assumption underlying the validity of our identification strategy is that credit growth and the change in interest rate from Italian and foreign banks have a similar trend before the crisis, conditional on all controls.

A first graphical evidence on this assumption can be seen in Figures 4 and 5. Figure 4 shows the 6-month change in the log credit granted by Italian and Foreign banks. While prior to the crisis the two series moved similarly, since June 2011, credit from domestic banks decreased at a much faster rate than credit from foreign banks. Figure 5 shows the change in the Annualized percentage rates on revolving credit lines for domestic and foreign banks. Prior to June 2011, the two series moved together. After the crisis, both Italian and foreign banks raised the cost of credit, but Italian banks did so at a faster pace than foreign banks.

These graphs suggest that before the crisis Italian and foreign banks behaved similarly. However, no adjustment is made for the variability accounted for by the controls included in the regression, and in particular for the different composition of firms borrowing from the two types of banks. Hence, we also show the dynamics of credit granted and of its cost as deviations from firm-period averages. We expect credit from domestic and foreign banks, net of firm effects, to move similarly until June 2011, and to start diverging afterwards. This is precisely what happens, as shown in figure 6. Likewise, divergence in the patterns of cost of credit occurs after June 2011, as shown in figure 7. These are the graphical counterparts of equations 1 and 2 (see also Khwaja and Mian 2008 for a similar representation of the data).

It is important to keep in mind that all our regressions also include bank fixed effects, hence we are already controlling for bank-specific time-invariant trends. The requirement for a common trend then only applies to how much Italian and foreign banks' trends depart from their time-invariant component before and after the crisis.

4 Data and descriptive statistics

Dataset. We use a unique dataset containing information at the bank-firm relationship level on credit quantities and prices.

We obtain data on individual bank-firm relationships from the Italian Credit Register (CR). This source lists all outstanding loan amounts above 30,000 Euros (less than 40,000 USD) that each borrower (both firms and households) has with banks operating in Italy, including branches and subsidiaries of foreign banks. Intermediaries are required by law to report this information. Data are available at monthly frequency and are of very high quality since intermediaries use the CR as a screening and monitoring device for borrowers.[8] Loans are distinguished into three classes: revolving credit lines, term loans, and loans backed by account receivables. The dataset includes both granted and drawn amounts. We focus our study on credit granted, as this better

[8] The CR also contains information on the borrowers' sector of activity (industry, defined at the 4-digit Nace level), location (province), type of business entity (corporations, limited partnerships, general partnerships, sole proprietorships, etc.).

captures a decision of bank to supply credit. Drawn credit is influenced by the decision of the borrower to use available lines, and this is largely affected by demand.

We also use information on interest rates charged by a representative sample of banks (103 Italian banks and 10 branches and subsidiaries of foreign banks) to Italian borrowers. These data are included in a sub-section of the Credit Register ("Taxia database"), and are available at quarterly frequency.

Den consolidated and unconsolidated (in case of stand-alone banks) balance sheets for Italian banks from the Supervisory Reports submitted by the intermediaries to the Bank of Italy, which is in charge of banking supervision in the country. We obtain consolidated balance sheet data for foreign banks from Bankscope.

Finally, data on sovereign yields, which we use to compute spreads, are from Thomson Datastream.

We merge these different data using the unique bank identification number, and the data on sovereign yields using the bank headquarter home country code.

Data on credit quantity and interest rates are collected at December 31, 2010, June 30, 2011 and December 31, 2011. We do not extend our sample beyond December 2011, because on December 22nd the ECB enacted Long Term Refinancing Operations (LTRO), which eased tensions in funding markets, and thus confounded the effect of the sovereign shock. Yet, this may be a period worth studying as future research to assess the effect of the LTRO on credit supply. We do not extend the sample before 2010 to reduce the risk that our results are influenced by other events or developments occurring in previous periods. However, our results are robust to extending the sample to include 2010.

Bank balance sheet information refers to December 31 2010 and to June 30 2011.

Sample. We include all non-financial firms with outstanding credit in the CR, including very small firms, such as sole proprietorships. We exclude firms with bad loans outstanding at the beginning of each period, since these are officially classified as losses and banks will not grant further credit to these firms until the procedure to recover at least part of the outstanding amount is completed.

To control for firm unobservable heterogeneity we select only firms borrowing from at least two banks. Since our identification strategy relies on a comparison between the behavior of foreign and Italian banks lending to the same firm, we select firms that borrow from at least one Italian and one foreign bank. This yields 664,198 bank-firm relationships over the two periods (331,635 in the crisis period and 332,563 in the pre-crisis period), involving 164,470 firm-period couples (82,077 firms in the pre-crisis period, 82,393 in the crisis period, overall 92,620 distinct firms sampled at least in one period). Basic statistics of the firms included in the sample are shown in Column 1 of Table 1. The sample of firms borrowing from at least one domestic and at least one foreign bank is broadly representative of the population of firms with at least two lending relationships (Column 2 of Table 1). Firms included in our sample are larger (measured by the amount of credit granted), more located in the North of the country, the richest area of Italy where subsidiaries and branches of foreign banks are mostly based, active more in the industrial and agricultural sectors (this mainly reflects the geographical location

Table 1: Descriptive Stastistics of Firms in the sample

	Sample Firms	**Other firms in the CR** (with more than 1 bank)
Credit Granted - Median - December 2011 (euros)	870,470	417,485
Credit Granted - Median - June 2011 (euros)	814,225	403,644
Number of banks - December 2011	4.02	2.68
Number of banks - June 2011	4.05	2.68
Sector (percent of firms)		
Agriculture	8.31	5.20
Construction	11.59	14.25
Energy	0.56	0.43
Industry	29.28	27.82
Service	50.27	52.30
Area (percent of firms)		
North	62.97	59.22
Center	18.21	22.30
South	18.83	18.48

of firms in the North of the country) than the average firm in the CR that borrows from at least two banks.[9] Despite being larger than the average firm in the CR, firms in our sample are small. The median total credit granted is around 850,000 euros, the mean is around 6.5 million.

Dependent variables. We compute the log differences in outstanding credit in each bank-firm relationships between June 2011 and December 2010 and between December 2011 and June 2011 to obtain the growth rate of loans in the pre-crisis and in the crisis periods, respectively. We control for mergers and acquisition among banks, so that if a firm had a relationship with a bank, and the bank disappears because it is acquired or merged, we can track whether there is a new relationship with the newly formed bank, or with the acquirer, in which case we consider the relationship as still existing. We aggregate credit at the banking group level, so if a firm borrows from two banks belonging to the same banking group, we consider this as a single relationship. We do so since lending and funding policies are typically decided at the banking group level, and we believe this is the relevant unit of observation to analyze the dynamics of credit supply.

For the same periods we also compute the change in the Annual Percentage Rate (APR) on revolving and term loans. The APR is the actual interest rate paid by firms and is computed by dividing the amounts due (that may be gross or net of fees and commissions) by the products (outstanding amounts multiplied by the days the amount was outstanding). This gives an average annual percentage rate on the loan. Rates on term loans are a less precise measure

[9] Focusing on firms with at least two banks is not particularly restrictive, since multiple banking is mainly determined by firm size.

Table 2: Descriptive Stastistics of Main Dependent Variables

	Mean	**Median**	**p25**	**p75**	**StdDev**	**N Obs**
6-month log changes						
ΔLog Credit	-0.054	0	-0.086	0	0.422	664,198
ΔLog Credit - Pre crisis	-0.041	0	-0.081	0	0.412	332,563
ΔLog Credit - Crisis	-0.066	0	-0.092	0	0.431	331,635
6-month changes, percentage points						
ΔAPR - Revolving	0.61	0.54	0.08	1.12	1.40	203,042
ΔAPR - Revolving - Pre crisis	0.40	0.40	0.00	0.86	1.35	100,791
ΔAPR - Revolving - Crisis	0.82	0.77	0.22	1.37	1.40	102,251
6-month changes, percentage points						
ΔAPR - Term Loans	0.39	0.33	0.17	0.50	0.63	134,323
ΔAPR - Term Loans - Pre crisis	0.33	0.31	0.18	0.42	0.53	66,832
ΔAPR - Term Loans - Crisis	0.45	0.35	0.16	0.57	0.71	67,491

of cost of credit than rates on revolving credit lines, because they depend on the maturity of the loan, which we do not observe, and also on the collateral posted, since they are typically collateralized. Then, our main results are based on rates on revolving credit lines, and results on term loans provide additional supporting evidence. We choose to use APR net of fees and commissions, because these are typically applied on credit granted while the interest rates we observe are estimated on the basis of the actual usage of the credit line. Then, if a credit line is used for a relatively small amount and for a very short period of time both the flow of interest rates paid and the products are small. As a consequence fees and commissions are large relatively to both interest rates and products leading to extremely large APR. However, for robustness purposes, we also estimate our baseline regressions for interest rates gross of fees and commissions. Our preferred measure of cost of credit is the APR on revolving credit lines.

Descriptive statistics. Descriptive statistics for the three main measures of credit supply we use in the paper are shown in Table 2. Credit contracted, on average, in both periods, but the contraction was larger after the crisis. Interest rates increased more after the crisis than in the pre-crisis period. This is true for both revolving credit lines and for term loans. The former can be renegotiated at short notice by banks, and this explains why in the post-crisis period rates on revolving credit lines grow more than term loans, whose conditions are more stable over time.

The dynamic of both credit granted and interest rates charged by Italian banks has been different from that of foreign banks after the crisis. As shown in Table 3, the growth rate of credit granted by Italian banks dropped from -3.7 to -7.0%, while that by foreign banks stood at -5.5% after the crisis, just 0.3 percentage points less than prior to the crisis. This suggests that the sharp increase in the spread on Italian sovereign debt did not affect the lending supply of foreign banks very much, so that the effect we identify in equation 1 by comparing domestic and foreign banks represents mostly the reaction of the former to the shock.

By the same token, domestic banks increased interest rates sharply during the crisis. Foreign banks also raised rates on revolving credit lines, while those on term loans changed very little.

Table 3: Credit Supply by Italian and Foreign Banks (simple average)

	Italian	**Foreign**
6-month log changes		
ΔLog Credit - pre crisis	-0.0373	-0.0516
ΔLog Credit - post crisis	-0.0704	-0.0547
6-month changes, percentage points		
ΔAPR - Revolving - Pre crisis	0.43	0.34
ΔAPR - Revolving - Crisis	0.89	0.62
6-month changes, percentages		
ΔAPR - Term Loans - Pre crisis	0.34	0.30
ΔAPR - Term Loans - Crisis	0.52	0.30

Then, our estimates in equation 2 may represent a lower bound for the full effect the sovereign shock on rates on revolving credit lines.

Of course, this evidence is only suggestive, as firms borrowing from foreign banks may be different from firms borrowing from Italian banks, in terms of lower demand for credit and higher risk. Regression analysis takes care of these possibilities.

Table 4 shows the distribution of bank-firm relationships by home country of the lender. More than a quarter of the relationships are from foreign banks. The majority are French owned. Then, German, American, Austrian, Spanish, Dutch and British banks hold more than 2,000 relationships. Banks from Japan, Switzerland, and Slovenia are less represented. Table 4 also shows the change in the spread of the 10 year sovereign security over the 10 year German Bund, between the average of January and the average of March 2011 for the pre-crisis period, and between the average of July 2011 and the average of September 2011 for the crisis period. It can be seen that this spread increased sharply, by almost 200 basis points, for Italy (see also Figure 1), for Slovenia (110 basis points), Japan and Spain (98 and 83 basis points, respectively). Prior to the crisis, spreads changed little, and in some instances, they decreased.

Our sample includes 567 banks, 49 of which foreign. Descriptive statistics of banks' balance sheet variables are shown in Table 5.

There is large variability in banks' balance sheet structure and size. Larger banks rely more on interbank funding, are less capitalized, have a smaller exposure to troubled sovereign securities than smaller banks.

Table 6 shows descriptive statistics of the main bank variables distinguishing between Italian and Foreign banks. The statistics are computed over both the crisis and pre-crisis period (data shown in Table 5 indicate that there is little difference across periods).

Foreign banks are on average larger, less capitalized, rely more on interbank funding, are less exposed to troubled sovereign securities. The relatively low standard deviation and the small interquartile range of all variables suggest that foreign banks are a more homogeneous

Table 4: Home Country of Banks included in the sample and changes in spreads

Country	Number of relationships	%	ΔSpread - Pre crisis	ΔSpread - crisis
			basis points	basis points
Austria	8,395	1.26	-0.4	32.7
Switzerland	207	0.03	-9.4	45
Germany	22,846	3.44	0	0
Spain	4,353	0.66	3.2	83
France	134,954	20.32	-3.7	38
UK	2,312	0.35	-44	34
Japan	463	0.07	-13	98
Netherlands	2,908	0.44	5.1	15
Slovenia	42	0.01	-7.6	110
United States	9,339	1.41	-37	7.8
Total foreign	185,819	27.98		
IT	478,379	72.02	12	192

Table 5: Balance Sheet Variables of Banks

		Mean	Median	p25	p75	StdDev
	T1 Ratio %	17.1	13.9	11.1	18.5	14.0
Pre-Crisis	Interbank/Assets %	5.6	2.7	0.92	6.17	9.11
Period	Exposure to Giips/Assets %	13.8	11.9	6.8	18.4	10.2
(Dec 2010)	Log Assets	6.9	6.0	5.0	6.9	3.7
	T1 Ratio %	16.8	13.9	11.2	18.5	11.9
Crisis	Interbank/Assets %	5.3	2.7	0.82	6.7	8.2
Period	Exposure to Giips/Assets %	13.6	11.5	6.7	17.8	9.9
(June 2011)	Log Assets	6.9	6.0	5.0	6.9	3.7

Table 6: Balance Sheet Variables of Banks

		Mean	Median	p25	p75	StdDev
	T1 Ratio %	17.2	14.2	11.3	19.0	13.3
	Interbank / Assets %	4.6	2.4	0.75	5.55	7.94
Italian	Exposure to Giips / Assets %	14.4	12.3	7.5	18.7	9.8
	Log Assets	6.0	5.8	4.9	6.7	1.55
	T1 Ratio %	12.8	11.4	10.4	13.6	5.2
	Interbank / Assets %	18.3	17.7	11.2	23.9	9.3
Foreign	Exposure to Giips / Assets %	1.64	0.88	0.19	2.22	2.03
	Log Assets	19.7	20.3	18.1	20.9	1.6

Table 7: Descriptive Statistics of Relationship-Level Controls

		Mean	**Median**	**p25**	**p75**	**StdDev**
	Share %	24.4	17.6	8.5	34.7	21.1
whole sample	Drawn/Granted %	63.7	75.0	35.7	97.8	35.7
	Share overdraft %	23.7	9.1	1.5	30.7	31.9
	Share %	23.6	16.8	8.2	33.1	20.8
Italian	Drawn/Granted %	62.2	71.5	33.4	96.2	35.8
	Share overdraft %	24.4	10.0	2.3	32.2	32.0
	Share %	27.3	20.0	8.5	41.7	23.2
Foreign	Drawn/Granted %	69.6	87.6	43.2	100	13.3
	Share overdraft %	21.9	5.0	0	25.9	32.1

group than Italian banks. Larger Italian banks have a balance sheet structure similar to that of foreign banks. In our regressions, systematic differences across banks are controlled by bank fixed effects.

Finally, we describe basic statistics of the relationship-level control variables included in our regressions (Table 7). Banks hold on average one fourth of credit in each relationship. The median share stands at about 17%. Firms draw on average about 64% of available credit, but the median firm draws 74% of it. Finally, overdraft facilities are on average 24% of total credit, 9.1% at the median. Italian banks tend to have a lower share of credit, the ratio of drawn to granted credit is lower for Italian banks, the share of revolving credit lines is higher for Italian banks. The differences in the means of these variables between Italian and foreign banks, while not large in absolute value, are statistically significant. Then, we include these variables as controls in the regression analysis.

5 Baseline model

5.1 Credit quantity

Results from the estimation of equation 1 are displayed in table 8.[10]

Columns 1 and 2 show the effect of the dummy *domestic* on the growth of credit granted. Before the crisis there is no difference between Italian and foreign banks. During the crisis, the behavior of the two types of banks is in fact different: credit granted by Italian banks grew by about 3 percentage points less than credit granted by foreign banks. These results are robust to the inclusion of bank fixed effects (column 2), which absorb the dummy *domestic*. Bank fixed effects control for differences in bank balance sheet structure[11] (bank's balance sheet structure did not change much between December 2010 and June 2011), bank organizational structure, and other bank-level time invariant unobserved heterogeneity, including bank-specific trends in

[10] We double cluster standard errors at the bank and at the firm level.

[11] The inclusion of bank fixed effects allows us to totally control for time invariant differences in bank characteristics, such as the riskiness or sectoral concentration of bank loan portfolios.

loan growth. Yet we do not observe much difference in the coefficients in the two specifications, and this suggests that the "domestic bank" variable of column 1 is already accounting for most of the cross-sectional heterogeneity across banks.

5.2 Interest rates

We now move to study the impact of the sovereign crisis on the cost of credit, by comparing the behavior of foreign and Italian banks in the pricing of loans, estimating equation 2.

Table 9 shows results of regressions on the change in the Annual Percentage Rate (net of fees and commissions) on revolving credit lines in columns 1 and 2 and on term loans in columns 3 and 4, without and with bank fixed effects, respectively. Domestics banks increased rates on revolving credit lines by about 20 basis points more than foreign bank lending to the same firm. The size of the coefficient of the interaction *domestic*crisis* changes very little if bank fixed effects are included. We run the same regression on the change in interest rates on term loans. Domestic banks increased rates on term loans by about 15 basis points more than foreign banks lending to the same firm. Interestingly, the dummy *domestic* is not significant neither in regressions on the change in rates on revolving credit lines, nor on the change in rates on term loans, indicating that prior to the crisis, domestic and foreign banks did not price credit differently. Overall, these results indicate that after the crisis Italian banks increased the price of credit more than foreign banks.

Regarding relationship-level controls, the share of credit held by the bank is not statistically significant. By contrast, the share of credit granted by the bank as revolving credit lines is positive and significant. This captures the extent of bank's unsecured exposure to the firm, and this explains the positive sign of the control. Finally, the ratio of drawn to granted credit is also significant, although this has different sign in regressions on revolving credit lines (positive) with respect to those on term loans (negative). This has to do with the fact that regressions on the change in interest rates are conditional on credit being granted to the firm. Then, if a firm is already using extensively its available credit, it may obtain further term loans posting collateral, which yields lower rates; if instead it obtains revolving credit lines (unsecured) it faces higher rates.

5.3 Robustness

We perform a series of checks to test the robustness of our main results.

First, we use credit drawn as an alternative measure of credit growth. Credit drawn is much more affected by firm demand for credit than credit granted. Even including firm-period fixed effects, credit drawn still partly reflects a decision of the firm, rather than a supply-side (bank) decision. Results are shown in columns 3 and 4 of Table 8. Overall, credit is drawn less intensely from domestic banks, providing a picture consistent to the one coming from the analysis of credit granted.

Second, we perform a placebo experiment, using the periods before June 2011 to test whether the difference between domestic and foreign banks in fact occurred after the burst of the sovereign crisis. As regards credit quantity, we use data from 2010, setting the fictitious event at June 2010. Then, we add the first half of 2011, and we set the event at June 2010 or at December 2010. In all cases (Table 10) neither the dummy domestic, nor the interaction between the dummy domestic and the dummy post-event are significant. Coefficients are also small in size. As regards the cost of credit, our data start on March 2010. Then, we use the second half of 2010 and the first half of 2011, setting the event at December 2010.[12] Results for the change in the APR on revolving credit lines are broadly similar to those on quantities, and thus omitted.

These results also provide support to the common trend assumption, suggesting that prior to June 2011, credit supply from domestic and foreign banks was not different.

Third, we also estimate the baseline regressions on Annual Percentage Rates gross of fees and commissions. These are an important component of the cost of credit. Results are shown in Table 11.[13] It can be seen that estimates are essentially unchanged: the coefficient of the dummy Italian banks interacted with the dummy crisis in the regression on the gross APR on revolving credit lines is larger, since revolving credit lines are particularly prone to the effect of peaks of usage, which determine very large effective gross rates in our data. The coefficient of the dummy *domestic*, interacted with the dummy crisis, in the regression on the gross APR on term loans is instead similar to that of the regressions on the net APR.

We perform some additional robustness checks (not shown in the paper to contain its length, but available from the authors): we estimate the models excluding Spanish banks since these have also been affected by the crisis[14]; we trim or winsorize the change in log credit when it is above or below the 1st and 99th percentile; we estimate the models excluding the relationship level controls since these may be correlated with previous period growth of credit. In all cases results continue to hold.

6 Alternative specifications of baseline model

6.1 Subsidiaries and branches of foreign banks

We also investigate whether our results are driven by systematic differences between branches and subsidiaries of foreign banks, by running separate regressions, where either branches or subsidiaries represent the foreign banks group. Our results suggest that the overall mitigating effect of foreign banks was mostly due to subsidiaries, possibly because they are able to rely more upon soft information than branches.

In this section we test whether results are robust to a finer definition of foreign banks. These include both subsidiaries and branches. However, their operational and financial structures are

[12] We also run regressions including the second quarter of 2010, setting the event at June 2010 and nothing changes. However, in this case we compare a 3-month change in the APR between June 2010 and March 2010 with 6-month changes over the following periods.

[13] The change in the gross APRs on revolving credit lines is winsorized at the 5th-95th percentile: these correspond to -33.8 and 27.0 percent. The change in the gross APRs on term loans is winsorized at the 1th-99th percentile: these correspond to -1.92 and 3.25 percent.

[14] In the second half of 2011, the increase in the delta-spread of Spanish sovereign securities was much smaller than the corresponding rise on the Italian Btp , as Table 3 shows.

quite different. While subsidiaries are very similar to domestic banks in terms of extension of their network of outlets and business model, branches often are specialized in specific market segments (e.g. syndicated loans, leasing, etc.), and concentrate their activity in certain areas of the country. Subsidiaries and branches also differ in the way they obtain funding. Branches typically obtain most of their funding as transfers from the headquarter, while subsidiaries rely relatively more on retail funding.

Table 12 shows results. Columns 1 to 3 display estimates from regressions run on the subsample of firms borrowing from at least one domestic bank and at least one subsidiary of foreign banks (branches of foreign banks are excluded). Results are similar to those of the baseline regressions, both for credit growth and for the cost of credit. Domestic banks grant less credit, and raise the cost of term loans more than subsidiaries of foreign banks.

Columns 4 to 6 display estimates from regressions run on the subsample of firms borrowing from at least one domestic bank and at least one branch of foreign banks (subsidiaries are excluded). In this case, we find that domestic banks raise the cost of revolving credit lines more than branches of foreign banks, while there seems to be no difference in the credit quantity supplied and in the interest rate on term loans.

Overall these results suggest that the effect we find in the main regression on the growth of credit granted is mainly driven by a different behavior of Italian banks relative to subsidiaries of foreign banks. By contrast, we find no significant difference in credit supply between domestic banks and branches of foreign banks, despite the fact that the latter enjoy better access to funding than domestic banks. Results on the cost of credit indicate that foreign banks, both subsidiaries and branches, appear to increase the cost of credit less than Italian banks. We interpret these results as evidence that the type of presence in the Italian market is relevant for the decision about the quantity of credit granted. Subsidiaries of foreign banks have a more extensive network of outlets and have therefore the possibility to collect more soft information on borrowers than branches of foreign banks. Conditional on granting credit, the pricing policy depends mostly on the cost of funding which was lower for both subsidiaries and branches of foreign banks during the crisis.

6.2 A continuous measure of exposure to sovereign risk

We also test the model using a continuous measure of banks' exposure to sovereign tensions:

$$\Delta credit_{i,j} = \beta_1 \Delta spread_j + \alpha_i + \varepsilon_{i,j} \quad (3)$$

$$\Delta APR_{i,j} = \gamma_1 \Delta spread_j + \alpha_i + \varepsilon_{i,j} \quad (4)$$

where $\Delta spread$ is the change in the spread with the German Bund on the 10 year sovereign securities of the country in which bank j is headquartered.[15] For this purpose we limit our

[15] This is computed as the difference between the monthly average of September 2011 and the monthly average of June 2011. We do so in order to avoid possible endogeneity issues, as the burst of the sovereign debt crisis occurred during the third quarter of 2011, and later developments may have been affected by the worsening of the business cycle, at least in Italy.

attention to the June - December 2011 period. To identify the impact of a change in the sovereign risk premia it is indeed more useful to exploit the cross-sectional variation of the delta spread during the crisis. Our focus is on the parameters β_1 and γ_1, which capture the elasticity of lending and interest rates to increased home-country sovereign risk. This exercise is also useful to take care of the possibility that foreign banks react to the shock: this model estimates the effect of an increase in banks' home country spread on credit supply, and it amounts to compare the behavior of banks hit by shocks of different intensity.

Results are shown in Table 13 and are consistent with those found with the baseline model. A 100 basis points increase in the spread leads to a 1.3 percentage points lower credit growth. This is a sizable effect, as the mean log change in credit is -6.7 per cent. The same increase in spread leads to interest rates higher by 16 and 11 basis points for revolving credit lines and term loans, respectively.

Importantly, the model predicts that the increase in the Italian sovereign spreads between July and September (192 basis points) leads to a lower credit supply by -2.5 percentage points, and to a raise in rates on revolving credit lines and term loans by 31 and 20 basis points, respectively. These effects are very similar to those estimated in the baseline model. This suggests that the estimates of the baseline model are very close to the full effect of a rise in sovereign spreads on credit supply, and the effect of the shock on foreign banks, less affected (the "control" group) is very limited. Perhaps, only the effect on the change in rates on revolving credit lines is underestimated by the baseline model.

We also estimate the above model on our initial panel, including the pre-crisis period and bank fixed effects, and results are unchanged.

7 Extensive margin

The extent to which banks decide to terminate existing relationships and to start new relationships are important determinants of borrowers' access to credit. When an existing relationship is cut, borrowers may need to look for alternative funding sources or scale down investment. When a new relationship is started, borrowers get a significant boost in their access to credit; moreover, this may represent a positive signal of borrower's ability to stay in business for other financiers, suppliers and customers.

As an additional extension, we study whether the sovereign debt crisis also affected the propensity of banks to terminate relationships and to accept applications for new loans. We also study whether the sovereign debt crisis affected the interest rates charged on new term loans.

As a first step, we estimate equations for the probability that a relationship is terminated. To this aim, we define a dummy variable taking value 1 if a bank-borrower relationship had positive credit granted only at the beginning of the period and value 0 if credit granted was positive at both periods. We compare the probabilities that a foreign and an Italian bank terminate a relationship with the same firm, by estimating a linear probability model which allows to include firm-period fixed effects. Table 14 shows that domestic banks are less likely to cut credit than foreign banks (columns 1 and 2, the latter includes bank fixed effects). Italian banks are about 1.6 percentage points less likely to terminate a relationship than foreign banks after the sovereign crisis started (on average about 7.5 percent of the relationships in place at June 2011 have been terminated by December 2011).

As a second step we examine the "extensive margin" of credit, in particular whether Italian and foreign banks were more, less, or equally likely to grant loans to new clients. In line with Jimenez *et al.* (2012), we use data on loan applications recorded in the CR in order to analyze the probability of acceptance/refusal of new credit. Every time a bank requests information on a borrower, the query is recorded in the CR, together with the motivation of the request, typically a loan application by a new client. This allows us to recover the number of applications for a loan made by each borrower to each bank in every period. We collect data on all the requests recorded between October 2010 and March 2011 and between July 2011 and December 2011, pre-crisis and crisis period, respectively. For each application we check if the bank granted any credit to the loan applicant in the sample period and in the following three months. Hence, a loan application submitted to a bank, say, in December 2010, is classified as accepted if we observe that the bank grants credit to the borrower in any point in time between the time of the request and March 2011. Our dependent variable is a dummy equal to 1 if the application of firm j to bank i is accepted, 0 otherwise. A stand-out descriptive feature of the frequency of accepted applications is that overall it has sharply dropped during the crisis, to 9 per cent between June 2011 and March 2012 from the 37 per cent observed in the three previous quarters.

We estimate a linear probability model. We also include firm fixed effects in some specifications to fully control for firm heterogeneity. However, this may induce a selection bias since the effect is identified on firms that make loan applications to at least two banks over a relatively limited period. The reason for applying twice might precisely be that the first application has been denied. Results are shown in columns 3 and 4 of Table 14. All regressions include bank fixed effects. Column 3 shows results without firm-period fixed effects, thus including also firms that make only one loan application in each period. Column 4 include firm-period effects, and the analysis is done on firms that made loan applications to at least two different banks in each period.[16] Results indicate that after the crisis the willingness to accept a loan application by Italian banks decreased more than that of foreign banks. An inspection of descriptive statistics suggests that the effect comes from foreign banks remaining equally selective in accepting loan applications over time and Italian banks becoming way more selective after the sovereign crisis burst.[17]

The combination of the results for credit growth, for the probability that a relationship is terminated, and for the probability that a new loan is accepted provides an elaborate picture. Foreign banks, those that were less affected from increases in sovereign spreads, are more ag-

[16] In this case identification is achieved thanks to firms applying for a loan to at least one foreign and at least one italian bank in each period.

[17] This is corroborated by regressions excluding bank fixed effects, but including a dummy for domestic banks: the latter is positive and significant, indicating that domestic banks were more likely to accept a loan application than foreign banks in the pre-crisis period. After the crisis this gap was filled because Italian banks reduced significantly their willingness to accept a new loan application.

gressive in cutting credit relationships and, furthermore, before the crisis they were less likely to accept a loan application than Italian banks. However, conditional on relationships being in place, foreign banks provide more credit than Italian banks. This suggests that foreign banks became more selective with their borrowers, yet once they have established a relationship they support their borrowers more. Possibly foreign banks have a tougher budget constraint than Italian banks, and are more able to cut more fragile relationships. This finding can be interpreted in the perspective of relationship lending: since foreign banks have stepped into the Italian market only in the second half of the 2000s, they have had relatively less opportunities to develop long-term bank-firm relationships. This possibility is in line with the results of De Haas and Van Horen (2012), who show that after Lehman's default, foreign banks continued to lend more to countries where they have longer lending experience.

As a last step, we study whether domestic and foreign banks charged different interest rates on new term loans. We use the data included in the Taxia dataset on the Annual Percentage Rates gross of fees and commissions charged on new term loans. In this case, we study the level of interest rates, and not the change, since these data are relative to specific loans, and not to outstanding balances. To avoid the possible influence of seasonal effects, we compare the level of interest rates charged on loans granted in the fourth quarter of 2011 with those granted in the fourth quarter of 2010. Results are shown in column 5 and 6 of Table 14. Column 5 does not include firm fixed effects, and thus include all term loans granted. Column 6 includes firm-period fixed effect and thus is estimated on the subsample of firms that obtain two new term loans in a quarter. Results indicate that the interests charged by domestic banks on new term loans have been about 35 basis points higher than those charged by foreign banks. This is consistent with the results we found in the regressions on the change in the cost of existing loans. The other controls behave as expected: the dummy crisis is positive and highly significant, indicating that interest rates on new term loans increased during the crisis (the effect is large, about 130 basis points, although this is not a pure supply effect). The size of the loan is significant only in the regression that does not include firm-period effect. Therefore, it likely proxies for the size of the firm. The negative sign of its coefficient thus indicates that larger firms are charged lower rates.

8 Bank heterogeneity

We showed that the sovereign crisis, that hit domestic banks, had an effect on their supply of credit: their credit growth was lower than that of foreign banks after the crisis. We now proceed onto studying whether this effect was in fact driven by bank characteristics that might have changed over time with a different extent across Italian and foreign banks.

In particular, we focus on bank capitalization (the Tier 1 ratio), bank size, the ratio of sovereign securities from European troubled countries (GIIPS) to total assets, and the ratio between wholesale funding and total assets. The last two variables are especially important because they capture the extent to which banks might be affected by the sovereign crisis. The higher the exposure to European "peripheral" countries, the higher the losses banks recorded in their balance sheets, and the more the cost of funding increased, as fears mounted that banks could face large losses. However, portfolio holdings of government bonds constitutes a form of collateral available for refinancing from the central banks, and for collateralized interbank borrowing. Wholesale funding is the most volatile source of funding, and it dried-up sharply in the second half of 2011.

Hence, we test whether our result on credit tightening by Italian banks compared to foreign ones holds even including bank balance-sheet characteristics in our baseline equations. This should take into account the possibility that our results on the interaction domestic*crisis are due to a spurious correlation between being a foreign bank and having a balance-sheet structure changing over time. This is not the case, since, as shown in Table 15, the interaction remains significant and negative in the regression on credit quantity growth and significant and positive in the regression for the change in the interest rates on revolving credit lines.[18] This means that, even if they had the same capital position and funding structure at the onset of the crisis of foreign banks, Italian banks would still be restricting credit more after the crisis burst: there appears to be a country-specific effect common to all Italian banks.

9 The aggregate effect

The empirical analysis discussed so far shows that domestic banks contracted credit growth and increased the cost of credit more than foreign banks after the burst of the sovereign debt crisis. These results are based on coefficients estimated comparing the behaviour of a domestic and a foreign bank lending to the same borrower ("within"), and therefore reflect partial equilibrium outcomes. However, firms might compensate the reduction in credit from domestic banks with increased loans from foreign banks that were not directly hit by the sovereign debt crisis.

Estimates from a simple firm-level regression is likely to be biased, though, because changes in the log of total credit at the firm level also reflect firm-level demand for credit, changes in firm financial strength, etc. A method to estimate the unbiased firm-level ("aggregate") impact of the supply shock induced by the crisis on the growth of credit commitments has recently been proposed by Jimenez et al. (2010). However, their methodology does not allow to easily obtain standard errors of the firm-level effects, and thus to conduct inference. In this paper, we use an alternative estimation procedure.[19] We first estimate firm-fixed effects from our base model at the bank-firm level. Then we plug these estimates of firm effects in a firm-level equation in which the dependent variable is the growth of total credit granted to firms by banks (including new relationships) and bank balance sheet controls are computed as averages weighted by the initial credit granted. Standard errors are estimated by block-bootstrapping at the bank level, to take into account the fact that firm fixed effects are estimated regressors.[20]

Formally, from the base model (equation 1), we obtain an estimate of firm-period fixed effect

[18] For term loans, the interaction is not significant, although still positive.

[19] A first version of this methodology appears in the June 2012 version of Bonaccorsi di Patti and Sette (2012).

[20] This approach is similar in spirit to that proposed by Abowd, Kramarz and Margolis (1999) to estimate worker effects in their study of wage premia.

$\widehat{\alpha}_{i,t}$. As a second step we estimate

$$\Delta credit_{i,t} = \beta_1 \overline{domestic_i} + \beta_2 \overline{domestic_i} * crisis_t + \widehat{\alpha}_{i,t} + \varepsilon_{i,t}$$

where $\overline{domestic_i}$ is the average at the firm level of the dummy *domestic* weighted by the share of credit to the firm held by each bank. A more thorough description of this approach can be found in the Appendix.

Results are shown in Table 16. Column 1 shows results without the estimated firm effects. The interaction term between the dummy domestic and the dummy crisis is negative and significant. This indicates that firms are not able to fully substitute credit from domestic banks by increasing credit from foreign banks. However, as argued above, this result is likely biased. In column 2 we show estimates including the firm effect. Now, the dummy domestic is still negative and significant, although the size of the coefficient is smaller. This suggests that when taking into account firm unobservables, including firm-level demand for credit, the supply effect is smaller. It is nevertheless still large: if the share of credit a firm obtained before the crisis from domestic banks increases by one standard deviation (12 percentage points), credit growth after the crisis is about 0.4 percentage points lower. This is large as the median credit growth in the crisis period is -3.1 percent (the mean is -4.8 percent).

We also computed the aggregate effect on the basis of the methodology proposed by Jimenez et al. (2010). In this case, the coefficient of the dummy domestic bank is -0.042. The coefficient estimated through our two-step approach, -0.033, is not statistically different from this value.

Finally, the estimated firm fixed effect is highly significant and positive, indicating that this is likely capturing firm-level demand for credit.

These results suggest that firms have not been able to fully substitute credit from domestic banks with more credit from foreign banks, and the sovereign crisis has therefore had an aggregate impact on credit supply.

10 Concluding remarks

In this paper, we study the impact of the recent sovereign debt crisis on the lending activity of Italian banks. To this aim, we exploit the variability observed between different categories of banks operating in Italy in quantities lent, interest rates charged, willingness to accept new applications and to terminate existing relationships over the transition between the pre-crisis and the crisis periods. We exploit the heterogeneous impact of the crisis across Italian and foreign banks operating in Italy.

Our results show that Italian banks tightened their supply of credit after the sovereign crisis burst, both in terms of quantities and prices. Lending by Italian banks grew by 3 percentage points less and the interest rates charged were 15 to 20 basis points higher with respect to foreign banks operating in Italy. Our estimates fully control for firm unobserved heterogeneity, by including firm-time fixed effects, and also hold when capturing bank unobserved heterogeneity through bank fixed effects.

We also analyze whether firms have been able to fully substitute for the decrease in lending of Italian banks during the crisis by increasing lending by foreign banks, thus keeping firms' access to credit substantially shielded from sovereign tensions. We find that in fact this was not the case: substitution was not complete and therefore the sovereign crisis exerted a significant aggregate effect on credit supply.

We test our results across a wide set of robustness checks. In particular we find that the difference between Italian and foreign banks does not seem to be due to differences in banks balance sheet characteristics. We also find that Italian banks increased the growth of credit less than subsidiaries of foreign banks. By contrast, we find no significant difference in credit granted between domestic banks and branches of foreign banks, despite the fact that the latter enjoy better access to funding than domestic banks. By contrast, both subsidiaries and branches, appear to increase the cost of credit less than Italian banks.

Besides analyzing the terms of existing credit relationships, our investigation also explores the differential behavior of Italian and foreign banks in accepting new loan applications and terminating existing relationships as the sovereign crisis burst. These results are particularly insightful, as they show that foreign banks, while tightening credit less with respect to Italian banks, did not relax their selectivity criteria during the crisis; if any, they increased it, being more likely to cut credit and maintaining very high rejection rates. An interpretation of this finding could be that foreign banks "flew to quality" during the crisis, by concentrating on supporting less fragile borrowers. This story suggests an examination of firms' characteristics, which we intend to pursue as a further extension of our work, by studying whether foreign and Italian banks behave differently depending on firms' riskiness (z-score, leverage, profits), liquidity, and opacity (size, age, tangible to total assets).

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A Appendix

A.1 Derivation of the aggregate effect

The relationship level equation is the following

$$\Delta credit_{i,j,t} = \beta_1 domestic_j + \beta_2 domestic_j * crisis_t + \alpha_{i,t} + \varepsilon_{i,j,t}$$

where $\Delta credit_{i,j,t}$ is the growth rate of credit to firm i by bank j at time t. Then, we take the average of both sides of this equation weighted by the share of credit held by each bank as follows:

$$\sum_{j=1}^{n_i} \Delta credit_{i,j,t} * \frac{credit_{j,t}}{\sum_{j=1}^{n_i} \Delta credit_{i,j,t}} = \beta \sum_{j=1}^{n_i} domestic_j * \frac{credit_{j,t}}{\sum_{j=1}^{n_i} \Delta credit_{i,j,t}} +$$
$$\beta \sum_{j=1}^{n_i} domestic_j * crisis_t * \frac{credit_{j,t}}{\sum_{j=1}^{n_i} \Delta credit_{i,j,t}} + \sum_{j=1}^{n_i} \frac{credit_{j,t}}{\sum_{j=1}^{n_i} \Delta credit_{i,j,t}} \alpha_{i,t} + \sum_{j=1}^{n_i} \frac{credit_{j,t}}{\sum_{j=1}^{n_i} \Delta credit_{i,j,t}} \varepsilon_{i,j,t}$$

where $\sum_{j=1}^{n_i} \frac{credit_{j,t}}{\sum_{j=1}^{n_i} \Delta credit_{i,j,t}} = 1$. Simple algebra shows that the left hand side is the growth rate of total credit obtained by firm i at time t. Then this yields:

$$\Delta credit_{i,t} = \beta_1 \overline{domestic_i} + \beta_2 \overline{domestic_i} * crisis_t + \widehat{\alpha}_{i,t} + \nu_{i,t}$$

which is the equation for the growth of credit at the firm level we are interested to estimate. To obtain the $\widehat{\alpha}_{i,t}$ we estimate them from the relationship-level equation. These estimates are unbiased and consistent as the number of banks increases (provided that the number of firms does not go to infinity). As the $\widehat{\alpha}_{i,t}$ are estimated in the relationship level equation, standard errors need to be estimated by bootstrapping to obtain correct estimates of the variance-covariance matrix. This equation is exactly valid for the growth rate of credit. We approximate it by the log change in credit, in the estimation.

To estimate the full aggregate effect, we also take into account that part of the growth of credit is due to the starting of new credit relationships. Our approach is valid as long as the firm-specific effect is the same for old as for new relationships, possibly up to a noise term uncorrelated with both the other regressors and the firm effect. This is reasonably true for firm-specific characteristics such as firm riskiness. It must also be true for firm demand for credit, which must not be bank specific. This, however, is an identifying assumption that must hold throughout our analysis, also when we study credit supply at the bank-firm relationship level.

B Tables and figures

Figure 1: Spread between 10-year Italian Btp and German Bund (percentage points)

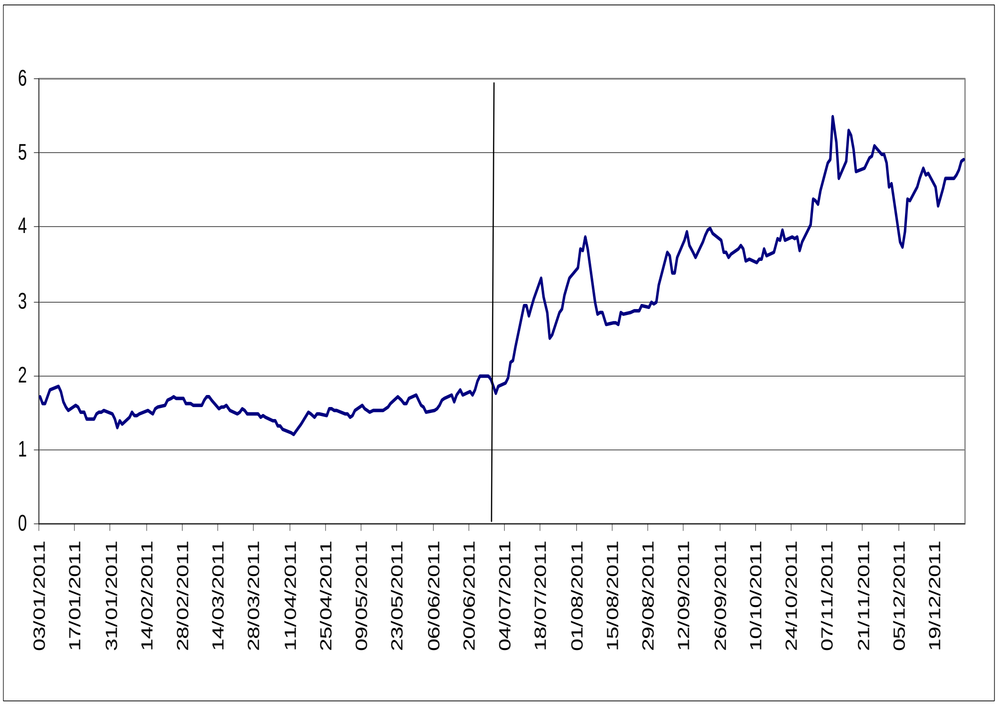

Figure 2: General government primary net borrowing / lending (percent of GDP)

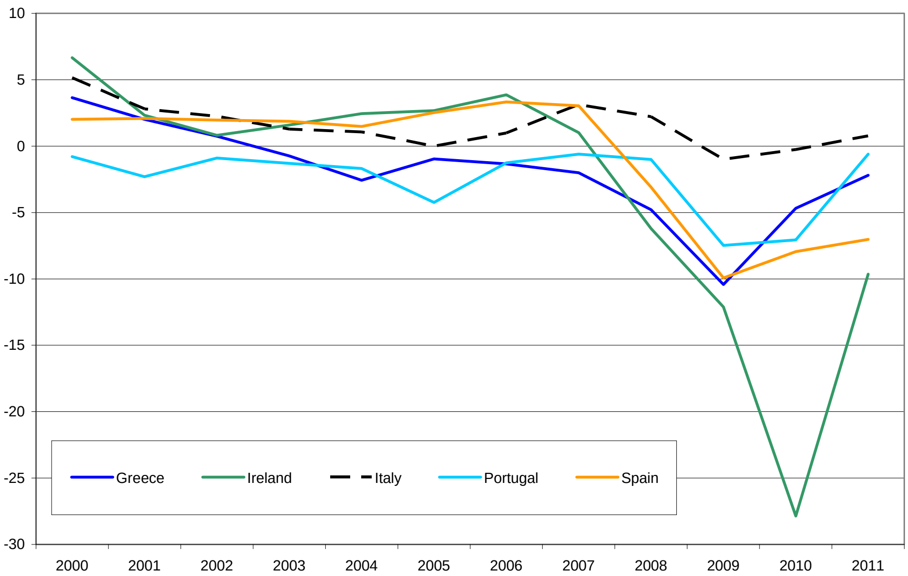

Figure 3: CDS spreads on 5-years senior debt of major banks (basis points)

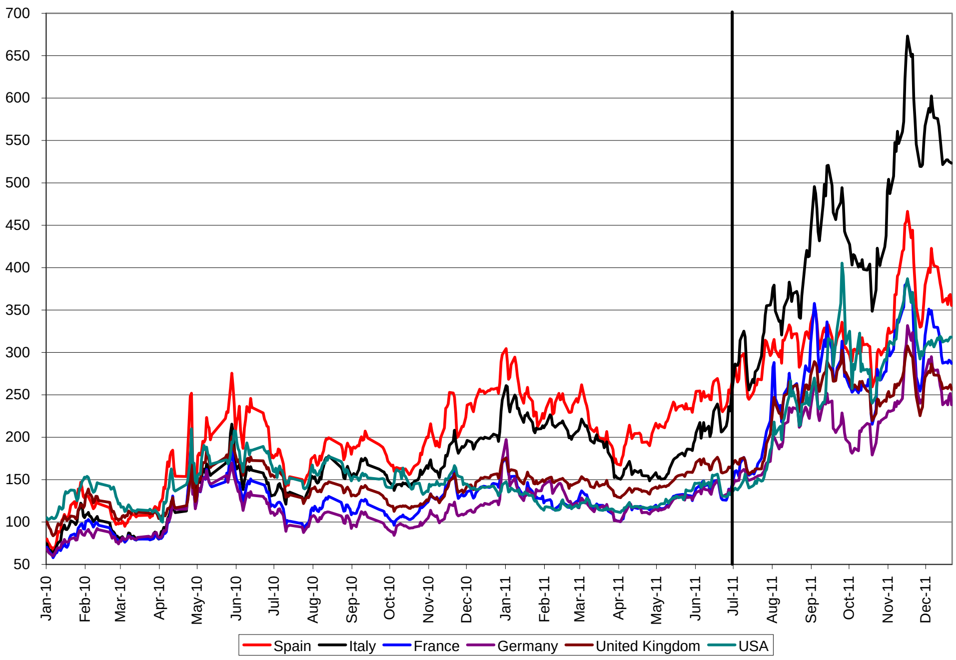

Figure 4: Change of credit granted by Italian and foreign banks (weighted average of log-changes of granted credit in each month relative to June 2011 - log points)

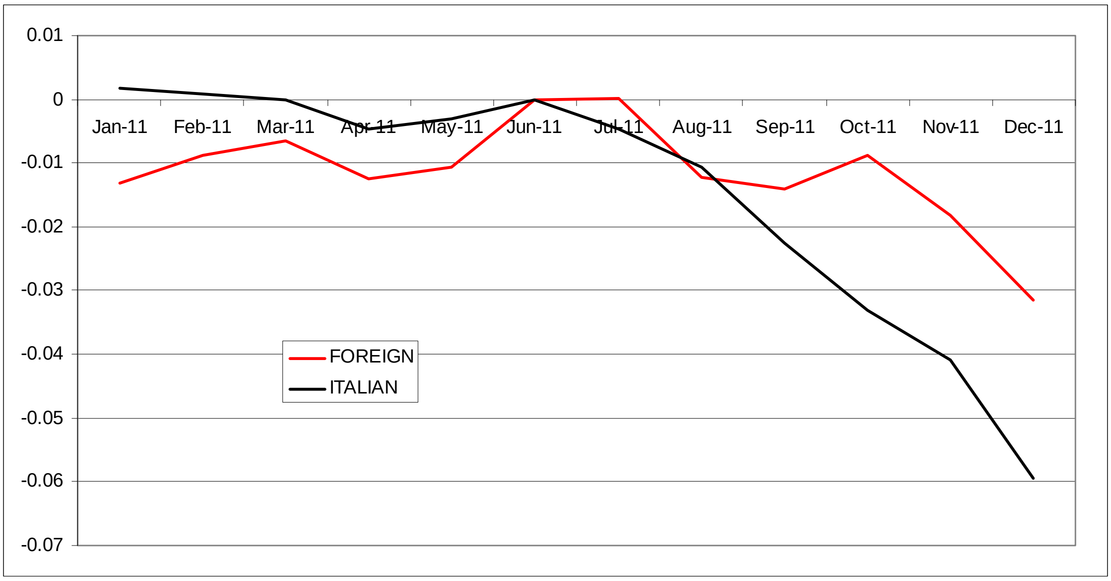

Figure 5: Change in the Annualized Percentage Rate on revolving credit lines (weighted average of changes of APR on revolving credit lines in each month relative to June 2011 - percentage points)

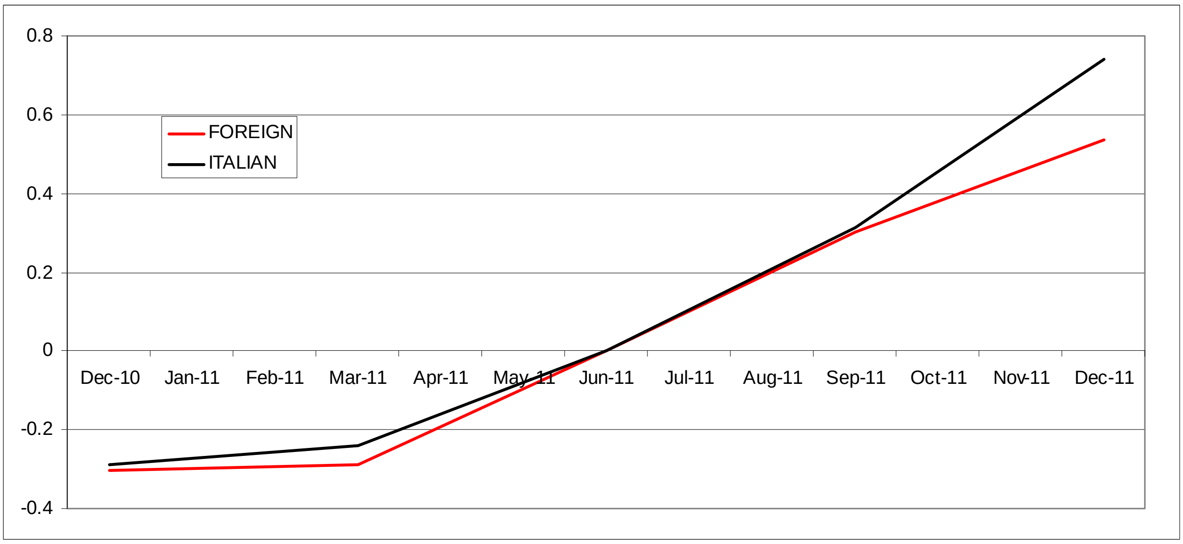

Figure 6: Change in credit granted, net of firms-period effects (growth rates of de-meaned credit granted in each month relative to June 2011 - percentage points)

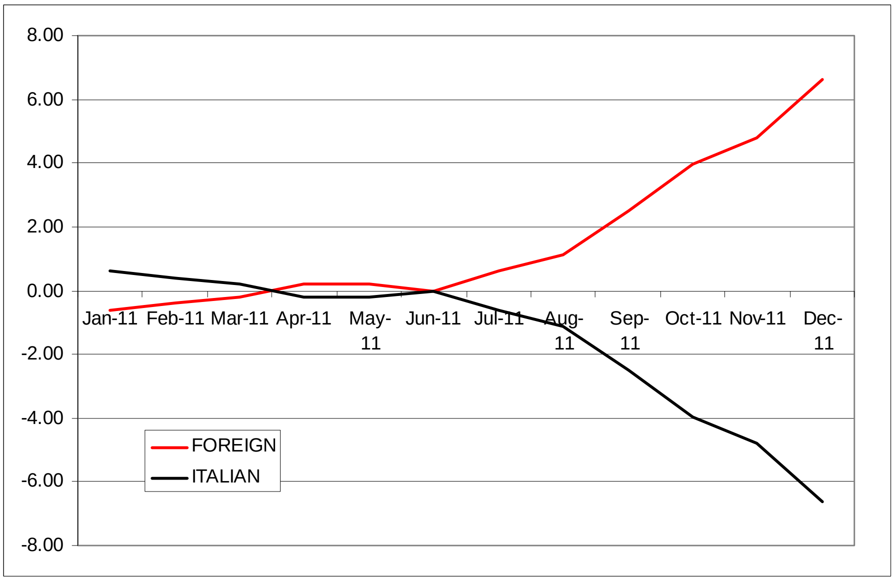

Figure 7: Change in Annualized Percentage Rate on revolving credit lines, net of firm-period effects (rate of change of de-meaned APR on revolving credit lines in each month relative to June 2011 - percentage points)

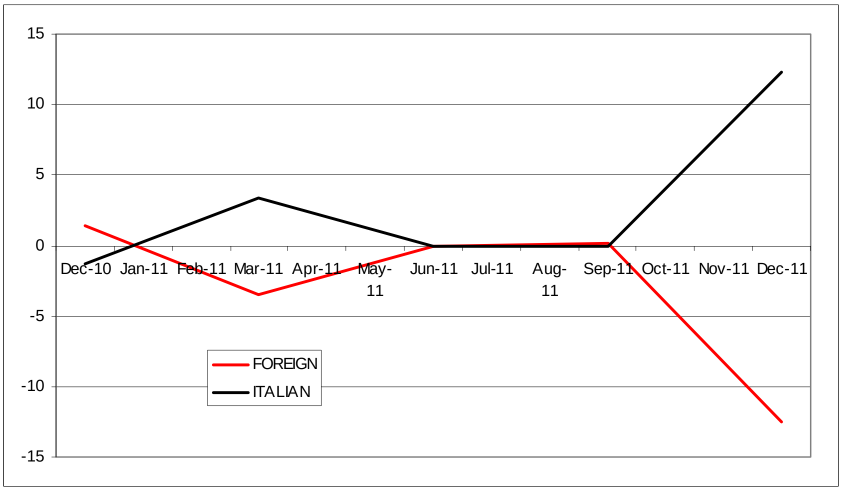

Table 8: Baseline - Credit quantity

DEP VARIABLE	$\Delta LOG(CREDIT) - granted$		$\Delta LOG(CREDIT) - drawn$	
	(1)	(2)	(3)	(4)
DOMESTIC BANK	0.00970		-0.0607***	
	(0.00836)		(0.0140)	
DOMESTIC BANK*CRISIS	-0.0298**	-0.0284**	-0.0525**	-0.0553**
	(0.0116)	(0.0116)	(0.0234)	(0.0233)
SHARE OF TOTAL CREDIT	-0.00119***	-0.00143***	-0.000521*	0.000183
	(0.000162)	(0.000149)	(0.000306)	(0.000248)
DRAWN OVER GRANTED	0.00833	0.0191***	-1.315***	-1.385***
	(0.00733)	(0.00726)	(0.0415)	(0.0379)
OVERDRAFT OVER TOTAL CREDIT	0.131***	0.119***	-0.0308	0.00546
	(0.00925)	(0.0107)	(0.0200)	(0.0227)
FIRM*TIME FIXED EFFECTS	yes	yes	yes	yes
BANK FIXED EFFECTS	no	yes	no	yes
Observations	664,198	664,198	569,608	569,608
R-squared	0.289	0.294	0.342	0.347
Number of Firm-Period Observations	164,470	164,470	146,886	146,886

standard errors clustered at bank and firm level in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 9: Baseline: Interest Rates

VARIABLES	$\Delta APR - REVOLVING$		$\Delta APR - TERM$	
	(1)	(2)	(3)	(4)
DOMESTIC BANK	0.0424		0.0142	
	(0.0614)		(0.0159)	
DOMESTIC BANK*CRISIS	0.208*	0.210*	0.154**	0.154**
	(0.114)	(0.113)	(0.0629)	(0.0626)
SHARE OF TOTAL CREDIT	0.000337	0.000292	9.11e-05	-0.000145
	(0.000409)	(0.000292)	(0.000161)	(0.000181)
DRAWN OVER GRANTED	0.0891***	0.106***	-0.0935***	-0.0620***
	(0.0237)	(0.0201)	(0.0234)	(0.0200)
OVERDRAFT OVER TOTAL CREDIT	0.139***	0.177***	0.0477**	0.0479**
	(0.0288)	(0.0218)	(0.0227)	(0.0205)
FIRM*TIME FIXED EFFECTS	yes	yes	yes	yes
BANK FIXED EFFECTS	no	yes	no	yes
Observations	203,042	203,042	134,323	134,323
R^2	0.319	0.320	0.373	0.370
Number of Firm-Period Observations	51,175	51,175	39,612	39,612

standard errors clustered at bank and firm level in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 10: Placebo Experiment

DEP VARIABLE	$\Delta LOG(CREDIT)$					
	Event June 2010				Event Dec. 2010	
	Only 2010		2010 and H1-2011		2010 and H1-2011	
	(1)	(2)	(3)	(4)	(5)	(6)
DOMESTIC BANK	0.00777		0.00777		0.0107	
	(0.0118)		(0.0119)		(0.0128)	
DOMESTIC BANK*CRISIS	0.00578	0.00651	0.0000939	0.00150	-0.00839	-0.00676
	(0.00452)	(0.00487)	(0.00416)	(0.00435)	(0.00803)	(0.00781)
FIRM*TIME FIXED EFFECTS	yes	yes	yes	yes	yes	yes
BANK FIXED EFFECTS	no	yes	no	yes	no	yes
Observations	688,562	688,562	1,045,143	1,045,143	1,045,143	1,045,143
R^2	0.258	0.263	0.270	0.273	0.270	0.273

standard errors clustered at bank and firm level in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 11: Gross Interest Rates

DEP VARIABLE	$\Delta APR - REVOLVING$	ΔAPR_TERM
	(1)	(2)
DOMESTIC BANK*CRISIS	1.228***	0.153**
	(0.382)	(0.0638)
SHARE OF TOTAL CREDIT	0.00642***	-0.000187
	(0.00146)	(0.000179)
DRAWN OVER GRANTED	0.882***	-0.0638***
	(0.191)	(0.0201)
OVERDRAFT OVER TOTAL CREDIT	1.142***	0.0528**
	(0.164)	(0.0205)
FIRM*TIME FIXED EFFECTS	yes	yes
BANK FIXED EFFECTS	yes	yes
Observations	203,042	134,323
R-squared	0.337	0.372

standard errors clustered at bank and firm level in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 12: Branches and Subsidiaries of Foreign Banks

	SUBSIDIARIES			BRANCHES		
VARIABLES	ΔLOG(CREDIT)	ΔAPR-REVOLV	ΔAPR-TERM	ΔLOG(CREDIT)	ΔAPR-REVOLV	ΔAPR-TERM
	(1)	(2)	(3)	(4)	(5)	(6)
DOMESTIC BANK*CRISIS	-0.0319**	0.208*	0.156**	0.00456	0.374***	0.0841
	(0.0126)	(0.114)	(0.0632)	(0.0222)	(0.136)	(0.0654)
SHARE OF TOTAL CREDIT	-0.00149***	0.000298	-0.000156	-0.00158***	-0.000116	-0.00136***
	(0.000154)	(0.000289)	(0.000183)	(0.000176)	(0.000455)	(0.000247)
DRAWN/GRANTED	0.0207***	0.107***	-0.0617***	0.0140**	0.125***	-0.0846***
	(0.00732)	(0.0199)	(0.0203)	(0.00714)	(0.0171)	(0.0288)
OVERDRAFT/TOTAL CREDIT	0.122***	0.177***	0.0482**	0.131***	0.140***	0.0486
	(0.0109)	(0.0219)	(0.0207)	(0.0128)	(0.0220)	(0.0324)
firm period fe	yes	yes	yes	yes	yes	yes
bank fe	yes	yes	yes	yes	yes	yes
Observations	639,569	202,494	133,513	445,464	132,601	58,819
R^2	0.303	0.320	0.372	0.367	0.336	0.402

standard errors clustered at bank and firm level in parentheses

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 13: Delta Spread

VARIABLES	$\Delta LOG(CREDIT)$	$\Delta APR-REVOLVING$	ΔAPR_TERM
	(1)	(2)	(3)
DELTA SPREAD	-0.0126**	0.161**	0.107***
	(0.00609)	(0.0795)	(0.0330)
SHARE OF TOTAL CREDIT	-0.00107***	0.00128*	0.000429
	(0.000173)	(0.000741)	(0.000290)
DRAWN OVER GRANTED	0.0136	0.0975***	-0.160***
	(0.0111)	(0.0280)	(0.0351)
OVERDRAFT OVER TOTAL CREDIT	0.149***	0.126***	0.0695
	(0.0146)	(0.0347)	(0.0438)
FIRM FIXED EFFECTS	yes	yes	yes
Observations	331,635	102,251	67,491
R^2	0.280	0.291	0.371

standard errors clustered at bank and firm level in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 14: Extensive Margin

DEP VARIABLE	Prob(cut=1)	Prob(cut=1)	Prob(accept=1)	Prob(accept=1)	APR_TERM	APR_TERM
	(1)	(2)	(3)	(4)	(5)	(6)
DOMESTIC BANK	-0.0190					
	(0.0120)					
DOMESTIC BANK*CRISIS	-0.0156**	-0.0167**	-0.109***	-0.0694***	0.361*	0.307**
	(0.00736)	(0.00754)	(0.0364)	(0.0216)	(0.203)	(0.119)
CRISIS			-0.154***		1.331***	
			(0.0221)		(0.183)	
SHARE OF TOTAL CREDIT	-0.00253***	-0.00207***				
	(0.000149)	(0.0000962)				
DRAWN OVER GRANTED	-0.0125*	-0.0375***				
	(0.00719)	(0.00645)				
OVERDRAFT OVER TOTAL CREDIT	-0.0813***	-0.0531***				
	(0.00617)	(0.00377)				
SIZE OF THE LOAN					-0.594***	-0.0620
					(0.0984)	(0.0714)
FIRM*TIME FIXED EFFECTS	yes	yes	no	yes	no	yes
BANK FIXED EFFECTS	no	yes	yes	yes	yes	yes
Observations	762,478	762,478	926,736	142,940	191608	60147
R-squared	0.407	0.429	0.088	0.110	0.342	0.651
Number of Firm-Period Observations	188,077	188,077		52,521		

standard errors clustered at bank level in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 15: Bank balance sheet structure

VARIABLES	$\Delta LOG(CREDIT)$	$\Delta APR-REVOLVING$	$\Delta APR-TERM$
	(1)	(2)	(3)
TIER 1 RATIO	-0.00661**	-0.0191	0.0443**
	(0.00296)	(0.0883)	(0.0189)
INTERBANK/ASSETS	0.000440	0.0632*	-0.00667
	(0.00223)	(0.0328)	(0.0189)
EXPOSURE TO GIIPS	-0.00477	0.0611	-0.0316*
	(0.00348)	(0.0458)	(0.0182)
TIER 1 RATIO * CRISIS	0.000177	0.0373*	-0.00818
	(0.00129)	(0.0196)	(0.00666)
EXPOSURE TO GIIPS * CRISIS	0.00100	-0.0222	0.00891
	(0.00143)	(0.0156)	(0.00730)
INTERBANK/ASSETS*CRISIS	-0.000221	0.0210	0.00834
	(0.000743)	(0.0134)	(0.00525)
DOMESTIC BANK * CRISIS	-0.0335**	0.432*	0.0787
	(0.0159)	(0.246)	(0.0890)
LOG ASSETS	-0.0551	-4.503	0.369
	(0.151)	(2.819)	(0.723)
SHARE OF TOTAL CREDIT	-0.00145***	0.000409	-0.000720***
	(0.000152)	(0.000270)	(0.000172)
DRAWN OVER GRANTED	0.0196***	0.0948***	-0.0592***
	(0.00733)	(0.0183)	(0.0199)
OVERDRAFT OVER TOTAL CREDIT	0.119***	0.182***	0.0400
	(0.0108)	(0.0219)	(0.0251)
FIRM*TIME FIXED EFFECTS	yes	yes	yes
BANK FIXED EFFECTS	yes	yes	yes
Observations	654,578	210,440	101,150
R^2	0.384	0.414	0.479
Number of Firm-Period Observations	162,128	53,914	32,426

standard errors clustered at bank and firm level in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 16: Aggregate Effect

VARIABLES	$\Delta LOG(CREDIT)$	$\Delta LOG(CREDIT)$
	(1)	(2)
DOMESTIC	-0.0109	0.00328
	(0.0178)	(0.0107)
DOMESTIC*CRISIS	-0.0436***	-0.0331***
	(0.0131)	(0.0103)
CRISIS	0.0104	0.00489
	(0.00931)	(0.00783)
SHARE OF TOTAL CREDIT	-0.00109***	-0.00110***
	(0.000118)	(6.75e-05)
DRAWN OVER GRANTED	-0.0176***	0.0149***
	(0.00419)	(0.00307)
OVERDRAFT OVER TOTAL CREDIT	0.0710***	0.105***
	(0.00590)	(0.00248)
FIRM EFFECT		0.689***
		(0.0265)
Observations	164,470	164,470
R^2	0.017	0.609

Block-bootstrapped (cluster at bank level) standard errors in parentheses

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

SECTION 3

THE MACROECONOMIC IMPACT OF THE SOVEREIGN DEBT CRISIS

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The macroeconomic effects of the sovereign debt crisis in the euro area

Stefano Neri and Tiziano Ropele*

May 2013

Abstract

Since the spring of 2010 tensions in the government bond markets in the euro-area have led to diverging dynamics in the cost of loans and credit developments among euro-area countries. These heterogeneous credit conditions, together with fiscal consolidations in some countries, have led to diverging trends in economic activity and employment. This paper studies the macroeconomic effects of the sovereign debt crisis focusing on a subset of euro-area countries using a Factor Augmented Vector Autoregressive (FAVAR) model. The analysis suggests that the sovereign tensions have led to an increase in the cost of new loans and a contraction in credit which has been particularly strong in the countries most affected by the crisis. The higher cost of credit and the contraction in lending have exerted a negative and significant effect on industrial production in both the peripheral and core countries. In the latter countries, the contraction in economic activity has reflected the strong trade link with the peripheral countries. The findings are robust to an alternative transformation of the data and measure of sovereign tensions.

JEL codes: E52, F41.

Keyword: sovereign debt crisis; FAVAR; Bayesian methods.

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* Bank of Italy, Economic Research and International Relations Area, Economic Outlook and Monetary Policy Department. We thank Fabio Canova, Francesco Nucci, Marco Lippi, Giulio Nicoletti, Marco Lombardi, Frank Smets, Giancarlo Corsetti and Gabriel Perez-Quiros for their comments and suggestions. The views expressed in the paper do not necessarily reflect those of the Banca d'Italia. All errors are the responsibility of the authors.

1. Introduction

The sovereign debt crisis that erupted in early 2010 has led to dramatic economic and social consequences. The widening of sovereign spreads in several euro-area countries has been accompanied by divergent financial and macroeconomics developments. While in the “core” countries (Germany, Netherlands, France, Austria, Belgium and Finland) financing conditions remained broadly in line with the European Central Bank (ECB) official rates, industrial production continued expanding and unemployment barely increased, in the “peripheral” countries (Greece, Ireland, Portugal, Spain and Italy) the picture was literally reversed. Credit became more costly and scantier, economic activity fell and unemployment increased. At the apex of the financial tensions in the last quarter of 2011, the survival of the euro area was at risk. The ECB and the national governments intervened with bold measures to restore confidence in financial markets, to support the flow of credit to the economy and to guarantee the sustainability of public finances.

In this paper we propose a quantification of the impact of the sovereign debt crisis on a number of macroeconomic variables, pertaining not only to the euro area as a whole but also to individual countries. Our interest in documenting the effects of the sovereign tensions on country-specific variables not only arises naturally because the debt crisis has exerted heterogeneous effects on member countries but also because shedding light on its transmission mechanisms may help shaping policy-makers’ interventions.

The outburst of the sovereign debt crisis in the euro area has attracted a lot of attention from central banks, international institutions, and academic researchers. However, to the best of our knowledge no paper has yet attempted a quantification of the macroeconomic impact of the sovereign crisis. In order to fill this gap, three important issues need to be resolved beforehand. First, how can one identify a “sovereign tensions” shock? Second, how can one deal with a potentially very large set of macroeconomic variables? Third, given that the tensions started at the end of 2009, how can one study the effects of the crisis using a limited sample size?

With regard to the first issue, given the uniqueness of the euro-area sovereign debt crisis, there is not a commonly and widely accepted indicator of sovereign tension. For this reason we propose three possible indicators, all of which are based on sovereign bonds yields. The first indicator builds on the so-called "wake-up call" hypothesis and consists of the spread between the 10-year Greek bonds yield and the interest rate swap of corresponding maturity. The second one exploits the results of a principal component analysis (PCA) on the whole set of sovereign spreads and is constructed as a weighted average of the spreads between the 10-year Italian, Greek, Irish, Portuguese, Spanish, Belgian, French and Austrian bonds yield and the interest rate swap of corresponding maturity. The last indicator consists of the first principal component of the whole set of sovereign spreads.

With regards to the second issue on how to handle a high-dimension data set, we conduct our empirical investigation estimating a Factor Augmented Vector Autoregressive (FAVAR) model with Bayesian methods. The FAVAR methodology, popularized by Bernanke, Boivin and Eliasz (2005), has been shown to be a suitable tool to examine the effects of shocks on high-dimension dataset. Our analysis employs a total of 139 time series regarding industrial production, bank interest rates, credit and monetary aggregates, inflation and unemployment rates at the national level and a set of aggregate variables for the euro area.

Finally, on the third issue we base our analysis solely on monthly time series observed from January 2008 to September 2012 and we use Bayesian methods.[1] In particular, Bayesian estimation allowed us to estimate the VAR over a short sample by employing prior distributions for the parameters of the model.

Our results can be summarised as follows. An adverse "sovereign tensions" shock, besides increasing sovereign spreads in the peripheral countries, leads to significantly heterogeneous credit conditions and credit dynamics across euro-area member countries. However, the impact on industrial activity is rather similar across countries, possibly reflecting the strong trade links

[1] As explained more in detail in Section 4.1, the estimation of the factors has been done using the sample period from January 2003 to September 2012.

between peripheral and "core" countries. The diverging trends in unemployment may be related to structural differences, including job protection and labour market flexibility. These findings are robust to a different transformation of the data and to alternative indicators of sovereign debt tensions.

2. The macroeconomic effects of sovereign debt crises

Historically, episodes of sovereign debt crises have been frequent in emerging as well as in developed countries.[2] Sovereign crises that spiralled out of control have often resulted in broader financial and banking crises and in some cases in a major macroeconomic meltdown.

Sovereign debt crises affect the economy through multiple channels. First, sovereign debt crises – and especially sovereign debt defaults – may lead to the exclusion of a country from international capital markets, with adverse effects on trade and investment activities. Richmond and Dias (2008) examine the duration of exclusion from international capital markets between 1980 and 2005 by a sample of sovereigns which defaulted and find that countries regained access to international capital markets after about five years. Second, sovereign crises usually entail a collapse in international trade. Rose (2005) uses a large panel data set for over 200 trading partners from 1948 to 1997 and finds a significant reduction in bilateral trade of approximately 8 per cent per year following the occurrence of a sovereign default. Third – and this is particularly relevant for the euro area – sovereign debt crises have direct effects on the banking sector and thus on the economy at large. Banks are major creditors of governments and thus their balance sheets and financial stability may be put at risk if governments may be expected to default on their debt. In these cases, banks' access to funding, especially on international wholesale markets, deteriorates, hampering their ability to provide credit to the economy and impairing the transmission of monetary policy.[3]

[2] See Reinhart and Rogoff (2009) for a history of financial crises; such episodes share remarkable similarities across time and countries. Recent episodes in developed countries are Russia in 1998 and Argentina in 2001.
[3] See Committee on the Global Financial System (2011).

Other channels may also be at work and feedback-loop effects may take place. Sovereign debt crises may be accompanied by a currency crisis and cause a deterioration in businesses' and households' confidence. Furthermore, measures of fiscal consolidation that are typically taken to restore confidence on the long-run sustainability of public debt may have short-run negative effects on the economy, and thus unintentionally exacerbate the crisis. Furthermore, banking crises are usually resolved through the injection of fresh capital by the national governments and thus the problems in the banking sector may end up as a further liability for the government (recall the Irish experience during the 2008-09 financial crisis).

While a sovereign debt crisis may unfold through multiple channels simultaneously – thus making difficult the chronological reconstruction of the events – the ultimate outcome is a contraction in output, a loss in the number of employees, a weaker financial system and, more generally, a decline in living standards. De Paoli, Hoggarth and Saporta (2009) document that sovereigns that faced debt crises have gone through deep recessions. The median output loss in their sample is nearly 5 per cent a year of pre-crisis annual output. Moreover, they show that a debt crisis commonly coincides with banking or currency crises and that when it coincides with both, it tends to be considerably more costly. More recently, Furceri and Zdzienicka (2012) analyze the short and medium-run effects of debt crises on output for a panel of 154 countries from 1970 to 2008. In the short-run, the results suggest that debt crises are very damaging, reducing output growth by 6 percentage points. Debt crises have also negative effects on output growth in the medium term. In particular, debt crises are associated with protracted output losses: 8 years after the occurrence of a debt crisis, output has fallen by around 10 per cent.

3. Macroeconomic developments in euro-area countries during the sovereign debt crisis

Before presenting the econometric analysis, in this Section we briefly outline the developments of some macroeconomic variables during the sovereign crisis.[4]

We emphasise two results: 1) the sovereign debt crisis has had dramatic financial and macroeconomic consequences and 2) the crisis has clustered the countries considered in the analysis into core and peripheral ones.

The financial crisis that erupted in September 2008 with the bankruptcy of Lehman Brothers, marked a halt in the trend towards more homogenous financial conditions in the euro area. Secured and unsecured money markets became increasingly impaired, especially along national borders. Soon after, sovereign bond yields started to diverge (Figure 1); this pattern became more pronounced following the onset of the sovereign debt crisis in May 2010, which caused significant heterogeneity in financial conditions across euro-area countries and more generally in macroeconomic developments. The underlying causes of the increase in heterogeneity originate in the accumulation of fiscal, macroeconomic and financial imbalances in several euro area countries prior to the crisis. When the crisis erupted, the unsustainable nature of these imbalances became evident.

Changes in banks' funding conditions have been extremely important to assess the ability of financial intermediaries to supply credit to the real economy. Figure 2 (top panels) reports the cost of deposit with agreed maturity up to one year held by households.[5] Since the beginning of 2009 the remuneration on these deposits in core countries remained broadly unchanged although with a larger dispersion compared to the pre-crisis period. For most of the peripheral countries from 2010 until the end of 2011 the cost of these deposits increased significantly, reflecting banks' difficulties

[4] We decided to keep this Section as brief as possible because there are now uncountable sources (central banks' websites, academic articles and working papers, newspapers, blogs, etc…) where one can read in great detail the chronology of the events that have characterized the sovereign debt crisis in the euro area.

[5] We concentrate on the interest rate paid on deposits with agreed maturity up to one year held by households, rather than an average rate on all deposits, for two main reasons. First, the remuneration on overnight deposits, which represent an important fraction of total deposits, has remained virtually unchanged in most countries. Second, the cost of deposits with agreed maturity, which have represented an important form of stable funding for banks, may well represent the marginal cost of stable funding.

in obtaining funding via market sources; throughout 2012 it slightly decreased owing to an improvement in market confidence, which was partly triggered the cuts in the key ECB interest rates in November and December 2011, as well as the non-standard monetary policy measures announced by the ECB in December 2011 (in particular, the two three-year longer-term refinancing operations) which aimed at further alleviating euro area banks' funding conditions.

Banks' funding difficulties in the peripheral countries also adversely affected the financing conditions to non-financing corporations (NFCs) and to households (HHs). As clearly illustrated in Figure 2 (middle and bottom panels) the cost of short-term loans to NFCs in the peripheral countries increased abruptly until the end of 2011; in some cases, namely Greece and Portugal, the interest rates reached the same levels observed in the third quarter of 2008 when the ECB policy rates were at their pre-crisis levels. Between May 2010 and December 2011 interest rates on short-term loans to NFCs increased by, respectively, 225, 140, 250 and 210 basis points in Italy, Spain, Greece and Portugal while in Germany the increase was remarkably smaller (40 basis points) and substantially in line with the increase (30 basis points) of the overnight (Eonia, Euro OverNight Index Average) rate. The interest rates on new loans to households for house purchase followed similar patterns. In either cases, the time dispersion across peer countries significantly increased during the sovereign debt crisis. The cost of new loans to HHs for house purchases at floating rate also increased rapidly in tandem with intensification of sovereign tensions. Only in 2012 the financing conditions in the peripheral countries became more favorable, still remaining at significantly higher levels than those observed in core countries.

Heterogeneous developments are also detectable when looking at credit developments (Figure 3). Although in 2010 the annual growth of loans to NFCs in core countries followed divergent patterns, from 2011 onwards growth rates varied in the range 0-5%. In the peripheral countries, however, despite a generalized recovery in 2010, the annual growth rate of loans to NFCs declined in 2011 and entered in negative territory in 2012. As for households, loan growth in the peripheral countries, while most of the correction of past excesses (before the financial crisis loans to households for house purchases grew at double digits in Greece, Spain and Italy) took place in the period 2006-2009, during the sovereign debt crisis developments in loans to households continued to remain subdued and in several cases became negative. Developments in lending to NFCs and to HHs reflected demand as well as supply factors. As for the former one can cite the weakness in economic activity, in investment spending by firms and housing markets. As for supply factors, especially towards the end of 2011, access conditions to credit markets tightened substantially in the peripheral countries. The Bank Lending Survey in the euro area conducted by the Eurosystem reported a remarkable tightening of credit standards to NFCs as well as the households.

The marked differences in financing conditions across countries and the scarcity of credit in the peripheral countries have been reflected in a poor economic performance. As reported in Figure 4, while in some core countries industrial production expanded and the unemployment rate either remained broadly stable or even declined (as in Germany), in the peripheral economies more affected the sovereign tensions, industrial activity sharply contracted and labour market conditions deteriorated markedly.[6] In most of these countries, industrial production stood at levels reached in 2009. From May 2010 to September 2012, industrial production contracted by 16 per cent in Greece, 7 in Spain and 3.5 in Italy. In contrast, in Germany as well as in other core countries industrial production grew by 10 per cent. In the summer of 2011, industrial production in Germany basically returned to the levels recorded before the 2009 Great Recession; in Italy and Spain, industrial production was lower by, respectively, 16 and 20 per cent.

Together with the decline in economic activity, the unemployment rate sharply increased in the peripheral countries, in particular in Greece, Portugal and Spain. In Italy, unemployment started rising in the summer of 2011.

[6] From mid-2009 to the beginning of 2012 unemployment in Germany steadily declined from 8 per cent to 5.5, as a result of the recourse during the crisis to more generous public subsidies and the achievement of agreements at the firm level that ensured employment levels in exchange for wage concessions or arrangements (see the box "Recent developments in the macroeconomic framework in Germany", in Economic Bulletin of the Bank of Italy, No. 62).

4. A Factor Augmented Vector Autoregression (FAVAR) model

In order to assess the effect of the sovereign debt crisis on key macroeconomic variables we rely on a Factor Augmented Vector Autoregressive (FAVAR) model that allows studying the effect of shocks on a large set of variables. As thoroughly documented in several recent contributions (e.g. Bernanke and Boivin, 2003; Bernanke, Boivin and Eliasz, 2005), the FAVAR approach builds on the idea the information contained in a *large number* of variables can be accurately captured by a *small number* of unobservable factors, which can then be included in a standard VAR model and to allow for a better characterization of the dynamics of the macroeconomic variables.

In what follows, we briefly go through the main steps of the FAVAR approach, highlighting some methodological issues we are confronted with.

4.1. Extraction of the factors: A Principal Component Analysis

Let X_t denote an $(n_x \times 1)$ vector (with n_x large) of economic variables for which we want to assess the dynamic response after a sovereign debt shock (to be defined shortly). To resolve the dimensionality problem that would make the analysis infeasible, the first step of the FAVAR approach is to run a Principal Component Analysis (PCA) on X_t, obtaining the following decomposition of X_t:[7]

$$\underset{(n_x\times 1)}{X_t} = \underset{(n_x\times n_f)}{\Lambda^F}\ \underset{(n_f\times 1)}{F_t} + \underset{(n_x\times 1)}{\xi_t} \tag{1}$$

where F_t is a vector of unobservable factors (with n_f small), ξ_t is a vector of idiosyncratic components and Λ^F is a matrix of factor loadings.[8] However, since the main goal of our analysis is to assess the effects of an unanticipated increase in the measure of sovereign debt tensions (see

[7] The PCA is a statistical technique that is used to determine whether the observed correlation between a given set of variables can be explained by a smaller number of unobserved and unrelated common factors.
[8] The number of factors to be retained is chosen on the basis of scree plot of the eigenvalues against the factor numbers. In particular, the number of factors coincides with the number of factors before the plotted line turns flat.

section 4.3) and the role of monetary policy we impose some observable factors Y_t. Formally, the decomposition of X_t now reads

$$\underset{(n_x \times 1)}{X_t} = \underset{(n_x \times n_f)}{\widetilde{\Lambda}^F} \underset{(n_f \times 1)}{\widetilde{F}_t} + \underset{(n_x \times n_y)}{\Lambda^Y} \underset{(n_y \times 1)}{Y_t} + \underset{(n_x \times 1)}{\widetilde{\xi}_t} \tag{2}$$

where Λ^Y is a matrix of observable factor loading. Note that in (2) $\widetilde{\Lambda}^F$, $\widetilde{F}_t$ and $\widetilde{\xi}_t$ have a superscript "~". In particular, the unobserved factors need to be orthogonalised with respect to Y_t and thus for this reason F_t is replaced by the residual of the following simple linear regression:

$$F_t = \beta Y_t + \widetilde{F}_t$$

and a new matrix of factor loading $\widetilde{\Lambda}^F$ is computed. Consequently, the vector $\widetilde{\xi}_t$ of idiosyncratic components is also re-computed. Note that these transformations are necessary in order to ensure a contemporaneous orthogonality between $\widetilde{F}_t$ and Y_t and thus making it possible to examine the effects of a shock to a variable in Y_t onto F_t and thereafter onto X_t (Bernanke, Boivin and Eliasz, 2005).[9]

Throughout our analysis, we assume that vector Y_t is bi-dimensional and includes the sovereign debt tensions indicator s_t (see section 4.3) as well as the monetary policy interest rate r_t (the rate on the main refinancing operations, MRO). The inclusion of the sovereign debt tensions indicator is trivially dictated by the fact that the scope of our analysis is to examine the effects of a sovereign debt tensions shock and thus we need have it in Y_t. There are two reasons why we treat the monetary policy rate as an observable factor. First, it turns out that one of the factors obtained in (1) visibly resembles the monetary policy rate or a very short-term money market interest rate. This is clearly illustrated in Figure 6. We thus have decided to impose the monetary policy rate as an observable factor. The second reason is that we are interested in studying the response of

[9] For this procedure to work one has also to ensure that the idiosyncratic components are orthogonal to Y_t.

conventional monetary policy to the shock. Likewise for this purpose we need have the monetary policy rate among the observable factors.[10]

Estimation of the unobserved factors

The unobserved factors are extracted from a series of 139 macroeconomic variables from January 2003 to September 2012. As for the factors, we decided to use a longer sample period in order to better capture the co-movements among the variables. On this regard, however, it is worth noting that the factors estimated over the 2003-2012 period are virtually identical to those estimated using the shorter 2008-2012 sample.

The set of variables includes industrial production for the total industry excluding construction, unemployment rates for the whole economy, loans to non-financial corporations, loans to households for house purchase, interest rates on new loans to non-financial corporations and households for house purchase at floating rate, interest rates on households deposits with agreed maturity up to one year, national contributions to the euro area monetary aggregate M3, inflation rates based on the Harmonized Indices of Consumer Prices (HICP), a group of variables for the euro area as a whole and, finally, sovereign spreads. The transformation applied to the time series are reported, together with the list of variables, in Table 1.

The transformed data are expressed in deviation from the sample mean and divided by their sample standard deviation. The number of factors extracted from the cross section of variables X_t turns out to be four. They account for around 80 per cent of the cross sectional variance. To assess the goodness of fit of the four latent factors as well as the two observable factors in explaining the variability of X_t we run linear regressions for each of the variables in X_t on the factors and compute the R^2. As reported in Table 1, the six factors provide a reasonable characterization of the set of variables with only one exception, industrial production in Ireland. Overall the average R^2 is equal to 0.89 and reaches 0.98 for bank rates, 0.92 for credit, 0.80 for industrial production and 0.91 for

[10] If we left the monetary policy rate in the vector X_t and did not include it in the VAR there would not be any feedback effect from the policy rate to the other variables and this in turn would the counterfactual exercise meaningless.

the unemployment rate; the R^2 for the other categories are between 0.76 and 0.86. Figure 6 reports both the PCA factors and the orthogonalized ones. Importantly, based on standard unit root tests, the idiosyncratic components are stationary and this allows us to avoid a detailed investigation on whether the factors are I(0) or some of them are I(1) (see Bai and Ng, 2004).

Similar results concerning the estimation of the unobserved factors are obtained if one first removes from all the variables in *X* the contemporaneous relation with the measure of sovereign debt tensions as well as with the monetary policy rate and then extracts the latent factors using the residuals of such regressions (see Buch, Eickmeier and Prieto, 2010).[11]

4.2. A VAR model

Once we have the unobservable and the observable factors we construct and estimate a reduced-form *p*-order VAR model:

$$z_t = \sum_{k=1}^{p} A_p Z_{t-k} + \varepsilon_t \text{ , with } E\left(z_t z_t^{'}\right) = \Sigma_{zz} \tag{3}$$

where $z_t = \left[\widetilde{F}_t \quad s_t \quad r_t\right]$ and s_t and r_t are the two observable factors, the indicator of sovereign debt tensions and the monetary policy rate, respectively.

In order to ensure identification of structural shocks, we assume – in line also with previous FAVAR analyses – a Cholesky factorization of the covariance matrix of the reduced-form residuals

$$\Sigma_{zz} = A_0^{'}\left(A_0^{-1}\right)^{'}$$

where the matrix A_0, which describes the contemporaneous relations among the elements of vector z_t, is lower triangular.[12] Such a recursive structure implies that the unobservable factors do not contemporaneously respond to either the sovereign debt tensions or to the monetary policy rate; the

[11] We thank Fabio Canova for suggesting this alternative method for estimating the latent factors. The results are available upon request.

[12] Sign restrictions are not useful as it is not easy to derive meaningful and theoretically justified restrictions to be imposed on the impulse responses there are only two observable variables, the MRO and the indicator of sovereign debt tensions, the responses of which can be restricted to have given signs.

sovereign debt tensions instead does respond contemporaneously to the latent factors but not to the policy rate; finally, the policy rate is left free to respond contemporaneously to measure of sovereign debt tensions as well as to the latent factors. Thus, it is clear the importance of ensuring the contemporaneous orthogonality between $\widetilde{F}_t$ and Y_t, which then allows use to use the identification strategy sketched above.[13]

Estimation of the VAR model

The VAR model is estimated using a Bayesian approach and in particular assuming a normal-diffuse prior for the coefficients (Kadiyala and Karlsson, 1997). This approach has several advantages. First, it provides an easier and more accurate assessment of the uncertainty. Second, it easily allows incorporating *a priori* information. Third, it better copes with the presence of unit roots (Sims and Uhlig, 1991). Finally, in-sample over-fitting is less problematic with Bayesian VAR models which have also good forecasting properties (Doan, Litterman and Sims, 1984).

The VAR model is estimated from January 2008 to September 2012 in order to focus on the financial turmoil that followed Lehman Brothers' collapse and on the euro-area sovereign debt crisis. The order of the VAR is set to 2 based on the Schwarz's Bayesian information criterion (BIC) and the outcome of serial correlation tests of the estimated residuals. Inspection of the residuals does not reveal any sign of heteroskedasticity, which is confirmed by formal tests.

As for the choice of the prior, we rely on the Litterman prior and set the mean of the coefficients on the first lag of each variable (except the indicator of the sovereign debt tensions) to 0.95, which corresponds to the average of the AR(1) coefficients over the VAR estimation period.[14] Given a normal distribution of the error terms in (2), the posterior distribution is normal-Wishart. Inference is conducted using the Gibbs algorithm.[15] Inference on the impulse responses is carried

[13] Strictly speaking, since we are interested in the effects of a sovereign debt tensions shock it would be enough to ensure orthogonality between $\widetilde{F}_t$ and s_t.

[14] Assuming a unit root prior yields similar results.

[15] Draws from the posterior distribution for which the companion of the VAR has eigenvalues larger than one in absolute value are discarded.

out with Monte Carlo methods drawing from the posterior distribution (which is normal conditional on the covariance matrix of the residuals) of the reduced-from parameters and that the covariance matrix of the residuals (which has a Wishart distribution).

4.3. Measuring the sovereign debt tensions

There exists a vast literature on the identification of structural shocks in VAR models. Just to name a few, Christiano, Eichenbaum and Evans (1999) have examined the effects of a monetary policy shock in the U.S. while Galí (1999) have concentrated on technology shocks. While the literature has converged on a particular set of assumptions for identifying the former shocks, the identification of an exogenous "sovereign debt tensions" shock is an unchartered territory.

The task of identifying a sovereign debt tensions shock is hard for several reasons. First, the sovereign debt tensions have materialized in the euro area only in late November 2009, when the new Greek government disclosed a significant revision to budget deficit and nearly doubled the 2009 forecast to roughly 13 per cent of GDP. Second, the financial turmoil initially confined to Greece rapidly spread to Ireland, whose economy suffered the consequences of a severe banking crisis, and to Portugal, whose economy was penalized by a sizable external trade imbalance and characterized by weak growth prospects. Tensions in financial markets assumed systemic proportions in the summer of 2011, when the European Council announced the private-sector involvement (PSI) to resolve the Greek crisis. Financial tensions then quickly passed onto Italian and Spanish government bonds. The Spanish economy was weakened by the burst of a housing bubble and by a fragile banking sector while the Italian economy was held back by weak prospects of economic growth and by an extremely high level of public debt to GDP ratio.

Figure 1 plots the 10-year sovereign spreads for the major euro-area countries. The spreads are computed with respect to the 10-year interest swap rate in order to account for flight-to-quality effects towards the German Bund during the financial and sovereign debt crises (De Sanctis, 2012). From 2003 to mid-2009, the average 10-year sovereign spread ranged from 22 basis points for

Greece to -28 for Germany. Thereafter, sovereign spreads, especially in Greece, Portugal, Ireland, Spain and Italy started to rise reaching unprecedented levels.

A convenient tool to examine the comovement among these spreads is the principal component analysis (PCA). Figure 7 reports pair-wise scatter plots between the R^2 of the regression of each of the spreads on the first three principal components at a time. It is interesting to note that regardless of which principal component one considers countries are clearly clustered into two groups: Germany, The Netherlands and Finland on one side and Greece, Portugal, Spain, Ireland, Italy, Belgium, France and Austria on the other side. This finding suggests that the assessment of market participants through prices clearly selected the most virtuous countries in the euro area as safe investment opportunities. Qualitatively similar results are obtained if the analysis is carried out over the 2008-2012 period.

Within the debate on the origins of the euro-area sovereign crisis some have emphasized the importance of “fundamentals” while others have stressed the role of “contagion”. For instance, Arghyroua and Kontonikasb (2011) find that during the crisis period market participants priced – much more heavily than they used to earlier – macro-fundamentals and international risk conditions on a country-by-country basis (mainly fiscal and other macroeconomic imbalances). Pericoli, Giordano and Tommasino (2013) and De Sanctis (2012) are recent studies that provide empirical support to the contagion hypothesis. More specifically, Pericoli Giordano and Tommasino (2013) find evidence of “wake-up call” contagion, which is defined as a situation where the outburst of a crisis in one country provides new information that triggers investors and market participants to update the default risk of other countries. De Sanctis (2012) shows that rating events concerning Greek sovereign bonds led to strong increases of sovereign yields in Ireland and Portugal and less noticeable, but still statistically significant, effects on Italian, Spanish, Belgian and French sovereign yields, suggesting that the spillover effect from Greece is predominant. Likewise, Arghyroua and Kontonikasb (2011) also find that several euro-area countries had experienced contagion from Greece supporting the view that Greek bond yield has become a proxy for EMU-specific systemic risk, increasing borrowing costs in other EMU countries beyond the level justified by the common international risk factor and their idiosyncratic fundamentals.[16]

In light of the above considerations, the following measures of sovereign debt tensions are used in the estimation of the VAR: (*i*) the spread between the 10-year Greek bonds yield and the 10-year interest rate swap, based on the evidence of "wake-up call" contagion; (*ii*) the weighted average of the spreads between the 10-year Italian, Greek, Irish, Portuguese, Spanish, Belgian, French and Austrian bonds yields and the 10-year interest rate swap, based on the results of the PCA reported in Figure 7; (*iii*) the first principal component on the whole set of sovereign spreads. Figure 8 shows our three indicators. Interestingly, they show remarkably similar patterns as also confirmed by the large pair-wise correlations (between 0.90 and 0.98).

5. The macroeconomic effects of the sovereign debt crisis

In this Section we describe the impulse responses of the macroeconomic variables to a sovereign tensions shock, identified using the Greek spread as measure of the tensions, and with the FAVAR outlined in Section 4.1 and employing the identification strategy discussed in Section 4.2. In particular, we examine the effects of an unexpected increase of 350 basis points in the Greek spread, which roughly corresponds to the increase observed from August to September 2011 after the announcement of the PSI. The results obtained using the other two indicators of the sovereign tensions are discussed in the Section 5.3.

[16] Several recent studies have put forward theoretical models with the aim to shed light on the possible causes and propagations mechanisms of the sovereign debt crisis in the euro area. For example, Arghyrou and Tsoukalas (2010) develop a theoretical model of the euro-area sovereign debt crisis combining elements from the second- and third generation currency crisis models, by Obstfeld (1996) and Krugman (1979) respectively. Guerrieri, Iacoviello and Minetti (2012) examine the international propagation of a sovereign debt default in a two-country microfounded economic model calibrated to data from the euro area, with the two countries representing the periphery (Greece, Italy, Portugal and Spain) and the core.

Figures 9 to 16 report the median impulse responses of the country-specific variables based on 10,000 draws. Figures 17 to 20 report the corresponding responses for the euro-area wide variables, together with 0.68 probability intervals (as suggested by Sims and Zha, 1999).[17]

5.1. The effects on individual countries

To begin with, we examine the impulse responses of the sovereign spreads to an unexpected increase of 350 basis points in the Greek spread. As reported in Figure 9, the sovereign tensions shock is immediately transmitted to the peripheral countries, while it barely affects the core economies. The Portuguese spread shows the largest increase (around 150 basis points on impact), followed by the Irish (around 80 b.p.) and by the Italian and Spanish (around 60). Among the core countries, Belgium shows the largest increase (20 basis points).

The transmission of tensions in government bonds markets works through several channels: the “price channel” (i.e. the lower the price of government securities (the higher their yield), the larger is the cost of credit since government bonds are one of the most important investment opportunities available on the market) (Albertazzi, Ropele, Sene and Signoretti, 2012 and Neri, 2013), the “balance sheet channel” (i.e. losses on government bonds lower banks’ profits and capital); the “liquidity channel” (i.e. the lower the prices, the lower the ability to borrow in the money market and from the central bank).

Following a sovereign debt tensions shock, the cost of short-term loans to non-financial corporations increases on impact in peripheral countries (Figure 10, top panels). The effect is particularly pronounced in Greece and Portugal where the interest rate rises by 50 and 40 b.p. whereas in Italy and Spain the effect is somewhat more muted (20 b.p. in either country). Turning to the core countries, the interest rate increases by 10 b.p. in France while it remains virtually unchanged in the other core economies. The probability of observing an increase on impact in

[17] According to Sims and Zha (1999) 0.68 probability error bands for impulse responses are more accurate and reliable than 0.95 ones.

Germany, Austria and Belgium is close to 0.60, while it is basically 1 in the peripheral countries, with the only exception of Ireland for which it is around 0.90.

The sovereign debt tensions shock also exerts a significant effect on the interest rate on new loans to households for house purchase in peripheral countries (Figure 10, bottom panels). The impact response is large in Portugal (around 30 b.p.) and relatively smaller in Italy and Spain (16 and 15 b.p., respectively). Among the core countries, it is interesting to note that while the shock raises the cost of new mortgages in Belgium and Netherlands (16 and 15 b.p., respectively) it decreases it in Germany and in Austria. As for non-financial corporations, the probability of observing an increase on impact in peripheral countries is one.

Figure 11 reports the impulse responses of annual growth rate of loans to non-financial corporations and households. The decline in lending to non-financial corporations and households is large in the peripheral countries, and particularly in Spain and Ireland for which the annual growth rates fall by respectively 2 and 3 per cent on impact. The decline is rather fast in the former two countries, while it is more delayed in Italy, where it is also more muted (-0.2 percentage points on impact), and Greece and Portugal (-1.1 percentage points in both countries on impact). While the model does not allow distinguishing between supply and demand factors underlying credit developments, qualitative evidence from the Bank Lending Survey of the Eurosystem shows that demand as well supply factors influenced the dynamics of credit to non-financial corporations and to households in the period 2011-2012. Demand factors were mostly related to the weakness in the economic outlook; supply factors reflected the banks' funding difficulties on wholesale markets and the deterioration in borrowers' creditworthiness. In the core countries, credit to non-financial corporations expands for nearly a year and then starts contracting.

Credit to households also fall significantly in the peripheral countries, with Greece and Ireland showing the largest decline on impact (-3.4 and -3 percentage points, respectively) and Italy and

Portugal the smallest (-1.3 and -0.9 percentage points, respectively). Among the core countries, the decline is significant only in Belgium.[18]

The annual growth rates of the national contributions to the euro-area M3 closely mirror those of credit, exhibiting large declines in peripheral countries: about -3 per cent in Greece and Ireland and about -1 per cento in Italy, Spain and Portugal (Figure 12). The impulse responses in the core countries point in general to small positive growth rates (except in Belgium). The dynamics in the peripheral countries are likely to reflect the weakening in economic activity and employment and, to some extent, deposit outflows.

The sovereign tensions shock reduces industrial activity in all countries (Figure 13). The decline in annual growth rate of industrial production is particularly large in Italy and Spain (-2.6 and -2.3 per cent after one year, respectively) but also in Germany (-2.2), Belgium (-2.4) and Finland (-2.0). The decline in the core countries reflects the fall in export to the rest of the euro area, including the peripheral countries (Figure 14).

While the fall in industrial production is pretty much of the same magnitude across euro-area countries, the responses of the unemployment rate exhibit a marked heterogeneity (Figure 15). One year after the shock, the unemployment rate increases by 2 percentage points in Greece and Spain, and 1.3 and 0.9 in Ireland and Portugal, respectively. In Italy the increase is smaller (0.4 percentage points). In the core countries, the unemployment rate remains almost unchanged in all the countries except Germany, where it declines by 0.7 percentage points after a year.

Finally, the sovereign tensions shock raises the annual HICP inflation in all countries except in Greece (Figure 16). In the peripheral countries, the rise in inflation may reflect variations in indirect taxes and administrative prices; in the core ones, it may reflect loose credit conditions in a context of low unemployment rates.

[18] The sharp decline in Belgium might reflect the large impact of securitisation (see also Figure 3). Unfortunately, data adjusted for securitisation are available only after 2010.

To sum up, an unexpected increase in sovereign debt tensions leads a rise in sovereign spreads of all the peripheral countries and to significantly heterogeneous credit conditions and credit dynamics across countries. The impact on industrial activity is similar across countries, possibly reflecting the strong trade links. The diverging responses in the unemployment may be related to structural differences in labour markets, including job protection and labour market flexibility.

The next Section discusses the impact on euro-area aggregate variables.

5.2. The effects on the euro area

In this Section we quantify the macroeconomic impact of an unexpected increase in the sovereign debt tensions on the euro area as a whole. The responses of the variables which are not included in the data set for the area as a whole are aggregate using the weights of countries real GDP in 2010 (normalized to 1 the sum of the GDP of the 11 countries considered).

After a sovereign debt tensions shock, the aggregate cost of new loans to non-financial corporations increases by slightly more than 10 basis points, while that of new loans to households by slightly less (Figure 17). The increase in lending rates is associated with a fall in credit to non-financial corporations and households (Figure 18); the former declines slowly and persistently reaching a maximum fall by 1.5 percentage points after two years while the latter falls by one percentage points after a year. As for monetary aggregates, M1 falls on impact by more than 1.5 percentage points, M2 by 0.4 and M3 by 0.2, suggesting that stable funding decreases after the shock. These dynamics are consistent with a reallocation within broad money from the more liquid component (M1) to the less liquid and more reactive money market rates deposits (M2-M1 which includes deposits with agreed maturity up to two years and deposits redeemable at notice up to three months). Indeed, money market rates increase after the shock, reflecting the rise in credit risk in interbank markets (Figure 19), which is measured by the spread between the three-month Euribor rate and the EONIA. The increase in the cost of money market funding and the decline in banks' stable funding exert a negative impact on banks' profitability as shown by the large decline in the average price of banks' stocks.

Following the shock, the ECB gradually reduces the rate on the main refinancing operations by around 30 basis points after two years, in an attempt to stabilise the euro-area economy and sustain the flow of credit to the private sector (Figure 19). The analysis does not consider the response of the ECB carried out through its unconventional monetary policy measures. The fall in the policy rate, the decline in banks' stock prices and the rise in sovereign spreads determine and immediate and persistent depreciation of the exchange rate of the euro against the major trading partners (Figure 19).

Finally, the transmission of the shock reaches the real economy causing a decline in industrial activity by 2.0 percentage points after a year; the response is hump-shaped and output turns positive after three years (Figure 20). The fall in activity is mirrored by the fall in confidence. Unemployment slowly increases, reaching a maximum of around 0.3 percentage points after two years, while inflation increases by 0.3 percentage points in annual terms.

5.3. Robustness: alternative indicators of the tensions and specifications of the VAR

In this Section we briefly discuss some robustness checks that we have carried out to assess the robustness of the results described in Section 4.

The first and most natural test is to repeat the analysis using the other two indicators of sovereign debt tensions described in Section 4.3. The strong correlation among the three indicators suggests that the results may be very similar, not only qualitatively but also quantitatively; and this is, indeed, the case.

The second test is to replace the MRO rate with the Eonia rate. The adoption of fixed-rate full allotment procedures in all the refinancing operations and the introduction of longer-term operations up to three year have determined a surge in excess liquidity which has pushed the Eonia rate to the lower bound of the corridor of ECB official rates, given by the rate on the deposit facility.

Therefore, the Eonia partly reflects changes in the MRO, and more precisely in the rate on the deposit facility, and partly the effects of the unconventional measures. Also in this case, the results are qualitatively similar.

The third test is to perform the analysis using the three-month differences for the variables which have been transformed using twelve-month changes and first differences for unemployment, lending rates and the sovereign spreads.[19] The results, which are broadly in line with those discussed in the previous sections, are reported in Figure 21 to 30. The results are broadly in line with those obtained with the baseline model and discussed in Sections 5.1 and 5.2. The larger persistence that characterizes almost all the responses is due to the transformation of the variables. Bank rates increase and lending and M3 fall substantially in the peripheral countries (Figures 22, 23 and 24) The fall in industrial production is more muted in Italy and Spain as well as in the core countries (Figure 25). Unemployment increases sharply in Greece while does not respond in the core countries (Figure 26). The increase in the other peripheral countries is smaller than in the baseline model. As for the aggregate variables, the impulse responses are in line with the baseline results (Figures 27 to 30). The only qualitative difference arises for the response of the nominal exchange rate, which appreciates following the sovereign shock.

Fourth, we have introduced the debt-to GDP ratios in the set of variables X_t. This has the advantage of allowing assessing the impact of the sovereign tensions on public debt and better identifying the sovereign tension shock by taking into account a key determinant of sovereign spreads (see, among others, Pericoli, Giordano and Tommasino, 2012). Overall, the impulse responses do not change. The responses of the debt-to-GDP ratios (Figure 31) are in line with our expectations; the peripheral countries experience a rapid increase while core countries are basically unaffected. In particular, public debt increases sharply in Ireland (by 12 percentage points), Portugal and Greece (8 percentage points), and slightly less in Spain and Italy (4 and 2, respectively). Finally, we have estimated the VAR over the period 2003:1-2007:12 and found, in line with our

[19] The exchange rate and the Euribor-Eonia swap spread enter in level while oil prices are first-differenced.

expectations, that an unexpected increase in sovereign spreads does not have exert any visible effect on any of the macroeconomic variables.

6. Conclusions

Almost four years have passed since the outburst of the global financial crisis, the euro area is struggling with an unprecedented crisis that has its roots, on the one hand, in the weak fiscal positions and macroeconomic imbalances of the peripheral countries and, on the other hand, in the lack of adequate instruments for managing and resolving the crisis, coupled with the incompleteness of the euro area architecture. The strong interconnection between sovereigns and banks has been a powerful transmission mechanism of the tensions in government bond markets. Two years and a half after the beginning of the tensions in government bond markets, economic activity is weakening in all euro-area member countries.

Surprisingly enough, there is very little evidence on the macroeconomic impact of the sovereign debt crisis on the euro area as a whole and on the individual countries. The aim of our research, which has to be though of as a starting point for a more structural analysis, is to fill this gap by resorting on state-of-the-art econometric tools, known as Factor Augmented Vector Autoregressive (FAVAR) models.

The empirical analysis confirms that the crisis, that started at the end of 2009 in Greece, rapidly spread to other countries with weak fiscal and macroeconomic conditions, namely Portugal and Ireland, and with some delay to Spain, which had suffered the consequences of a fall in property prices, and Italy, with its high public debt and weak growth prospects in the medium term. Credit conditions have become significantly heterogeneous, with the cost of credit raising sharply in the peripheral countries, and posing important challenges to the ECB monetary policy. The tensions in sovereign debt markets have caused a decline in economic activity in all countries and also at the aggregate euro-area level.

While our analysis helps understanding the real effects of the sovereign crisis, a lot more needs to be done, in particular along two dimensions. On the empirical side, more elaborated models, possibly allowing for time variation in parameters (Primiceri, 2005 and Koop and Korobilis, 2012), might be more useful to complete our task to fully capture the transmission of the sovereign debt crisis to the euro-area economy. The Large Bayesian VAR approach suggested by by Bańbura, Giannone and Reichlin (forthcoming) is an interesting and appealing alternative to deal with the high dimension of the data. On the theoretical side, structural models with open economy features allowing for the possibility of sovereigns' and banks' defaults may be extremely useful to analyse the channels through which the fear of unsustainable fiscal dynamics end up hitting the real economy and spilling over to the rest of the global economy. Needless to say, such models need to incorporate not only a "conventional" role for monetary policy but, most importantly, its "unconventional" dimension.

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Table 1

List of variables, transformation and properties of the factors

Variable		Transformation	Unit root test idiosyncratic term	Persistence idiosyncratic term	R-squared Factor 1	Factor 2	Factor 3	Factor 4	Sovereign spread	MRO rate	All factors
		(a)	(b)	(c)	(d)						
Rates on loans to NFCs											
	IT	1	-5.51	0.54	0.03	0.01	0.01	0.02	0.04	0.81	0.98
	ES	1	-6.48	0.35	0.01	0.04	0.00	0.00	0.04	0.82	0.99
	DE	1	-7.58	0.24	0.00	0.01	0.00	0.00	0.26	0.97	0.99
	FR	1	-7.16	0.35	0.01	0.01	0.00	0.00	0.15	0.93	0.98
	PT	1	-4.15	0.42	0.03	0.09	0.01	0.00	0.14	0.23	0.95
	GR	1	-4.96	0.56	0.01	0.01	0.01	0.01	0.27	0.14	0.95
	NL	1	-6.63	0.25	0.01	0.00	0.00	0.00	0.20	0.95	0.98
	FI	1	-5.74	0.42	0.01	0.00	0.00	0.00	0.22	0.96	0.98
	AT	1	-5.78	0.49	0.01	0.01	0.00	0.00	0.26	0.96	0.98
	BE	1	-2.65	0.46	0.01	0.01	0.00	0.00	0.26	0.98	0.99
	IE	1	-5.21	0.26	0.01	0.00	0.00	0.00	0.22	0.96	0.98
Rates on loans to HHs											
	IT	1	-3.27	0.80	0.02	0.01	0.04	0.03	0.07	0.84	0.98
	ES	1	-5.59	0.64	0.01	0.09	0.04	0.00	0.08	0.82	0.98
	DE	1	-2.81	0.66	0.02	0.01	0.01	0.01	0.33	0.93	0.97
	FR	1	-4.59	0.71	0.01	0.29	0.13	0.00	0.13	0.63	0.93
	PT	1	-4.89	0.58	0.04	0.02	0.00	0.02	0.00	0.65	0.97
	GR	1	-3.76	0.76	0.09	0.01	0.01	0.02	0.12	0.61	0.78
	NL	1	-4.86	0.51	0.01	0.19	0.03	0.15	0.02	0.56	0.97
	FI	1	-5.71	0.68	0.01	0.01	0.00	0.00	0.28	0.96	0.98
	AT	1	-3.05	0.79	0.02	0.05	0.05	0.02	0.39	0.84	0.95
	BE	1	-3.72	0.79	0.00	0.09	0.02	0.01	0.04	0.76	0.96
	IE	1	-4.32	0.66	0.00	0.01	0.00	0.00	0.16	0.94	0.99
Rates on HHs' deposits up to 1 year											
	IT	1	-6.48	0.41	0.02	0.07	0.04	0.00	0.14	0.26	0.96
	ES	1	-2.06	0.81	0.00	0.16	0.01	0.10	0.01	0.56	0.95
	DE	1	-5.48	0.49	0.00	0.02	0.00	0.00	0.20	0.95	0.99
	FR	1	-6.16	0.35	0.01	0.02	0.00	0.01	0.10	0.89	0.99
	PT	1	-4.54	0.77	0.01	0.10	0.00	0.01	0.08	0.32	0.94
	GR	1	-5.59	0.72	0.01	0.14	0.00	0.05	0.30	0.06	0.93
	NL	1	-5.18	0.51	0.00	0.12	0.03	0.04	0.02	0.65	0.93
	FI	1	-5.91	0.48	0.00	0.01	0.00	0.02	0.13	0.91	0.99
	AT	1	-6.04	0.50	0.01	0.03	0.01	0.00	0.17	0.93	0.99
	BE	1	-6.02	0.37	0.00	0.01	0.00	0.00	0.31	0.98	0.99
	IE	1	-5.76	0.52	0.00	0.04	0.00	0.02	0.04	0.81	0.98
Contribution to euro area M3											
	IT	2	-4.44	0.83	0.01	0.00	0.10	0.01	0.60	0.34	0.78
	ES	2	-4.08	0.65	0.00	0.01	0.03	0.07	0.51	0.77	0.97
	DE	2	-4.91	0.68	0.14	0.16	0.05	0.04	0.00	0.43	0.90
	FR	2	-4.69	0.73	0.09	0.09	0.04	0.08	0.12	0.63	0.82
	PT	2	-5.44	0.44	0.11	0.14	0.10	0.09	0.35	0.56	0.91
	GR	2	-4.34	0.67	0.01	0.02	0.08	0.02	0.79	0.57	0.98
	NL	2	-5.14	0.50	0.08	0.01	0.07	0.00	0.09	0.35	0.63
	FI	2	-3.56	0.70	0.05	0.05	0.00	0.15	0.09	0.59	0.79
	AT	2	-5.66	0.72	0.07	0.00	0.01	0.08	0.16	0.79	0.91
	BE	2	-3.55	0.78	0.01	0.06	0.00	0.11	0.48	0.40	0.74
	IE	2	-5.05	0.50	0.01	0.27	0.01	0.12	0.31	0.21	0.91
Loans to NFCs											
	IT	2	-2.43	0.82	0.03	0.00	0.00	0.02	0.32	0.87	0.90
	ES	2	-2.67	0.86	0.01	0.17	0.00	0.12	0.46	0.55	0.94
	DE	2	-4.05	0.73	0.04	0.36	0.11	0.00	0.02	0.48	0.95
	FR	2	-3.67	0.75	0.06	0.01	0.01	0.02	0.14	0.78	0.87
	PT	2	-4.94	0.50	0.01	0.07	0.01	0.02	0.60	0.71	0.96
	GR	2	-4.64	0.57	0.16	0.00	0.02	0.02	0.40	0.68	0.90
	NL	2	-3.58	0.75	0.02	0.08	0.00	0.02	0.12	0.69	0.84
	FI	2	-4.46	0.62	0.24	0.09	0.04	0.16	0.02	0.45	0.93
	AT	2	-3.68	0.79	0.03	0.20	0.12	0.00	0.08	0.60	0.87
	BE	2	-4.41	0.69	0.02	0.04	0.06	0.01	0.10	0.76	0.88
	IE	2	-5.63	0.77	0.01	0.12	0.01	0.20	0.42	0.50	0.96
Loans to HHs											
	IT	2	-2.97	0.72	0.02	0.56	0.16	0.08	0.32	0.04	0.94
	ES	2	-3.66	0.89	0.03	0.28	0.01	0.19	0.47	0.38	0.97
	DE	2	-3.91	0.77	0.02	0.33	0.17	0.09	0.22	0.48	0.89
	FR	2	-2.31	0.82	0.06	0.29	0.10	0.10	0.28	0.56	0.95
	PT	2	-4.65	0.68	0.00	0.03	0.03	0.01	0.64	0.66	0.89
	GR	2	-5.16	0.53	0.04	0.32	0.04	0.19	0.48	0.34	0.99
	NL	2	-4.44	0.71	0.16	0.40	0.39	0.09	0.06	0.00	0.73
	FI	2	-3.76	0.69	0.05	0.28	0.02	0.26	0.42	0.38	0.99
	AT	2	-3.23	0.79	0.09	0.33	0.03	0.38	0.30	0.14	0.92
	BE	2	-6.39	0.49	0.01	0.79	0.52	0.03	0.03	0.08	0.94
	IE	2	-4.32	0.63	0.08	0.41	0.04	0.37	0.28	0.19	0.98

Note: column *(a)*: 1 denotes level, 2 12-month percentage changes; column *(b)* reports the results of the Augmented Dickey-Fuller test for the idiosyncratic component of each variable; column *(c)* reports the sum of the first four autoregressive coefficients estimated for the idiosyncratic component of each variable; column *(d)* reports the R^2 (*R-squared*) of the regression of each variable on each factor separately and on all of them (last column).

Table 1 (cont'd)

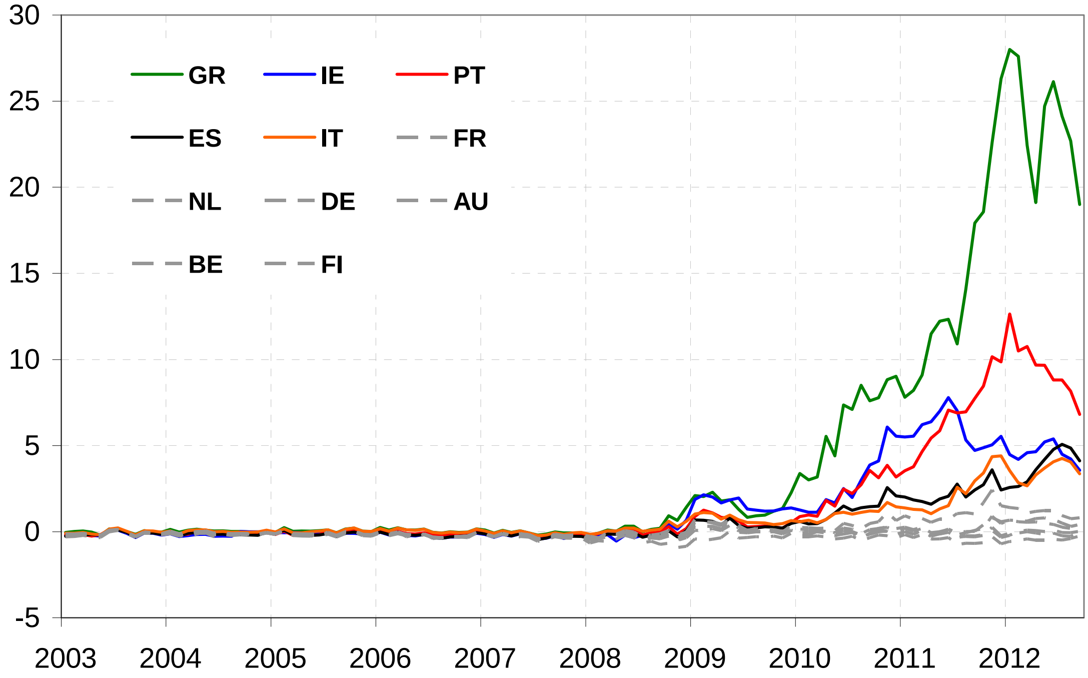
List of variables, transformation and properties of the factors

	Variable	Transformation	Unit root test idiosyncratic term	Persistence idiosyncratic term	R-squared Factor 1	Factor 2	Factor 3	Factor 4	Sovereign spread	MRO rate	All factors
		(a)	*(b)*	*(c)*	*(d)*						
Industrial production											
	IT	2	-7.15	0.17	0.01	0.62	0.74	0.09	0.01	0.04	0.95
	ES	2	-8.37	0.37	0.01	0.83	0.58	0.01	0.02	0.03	0.95
	DE	2	-8.04	0.38	0.00	0.56	0.83	0.07	0.01	0.02	0.95
	FR	2	-5.43	0.41	0.01	0.66	0.79	0.06	0.00	0.01	0.95
	PT	2	-4.75	0.49	0.08	0.36	0.40	0.20	0.00	0.03	0.73
	GR	2	-9.34	0.11	0.00	0.35	0.11	0.04	0.17	0.23	0.63
	NL	2	-8.60	0.35	0.03	0.64	0.68	0.08	0.00	0.02	0.91
	FI	2	-4.66	0.40	0.03	0.57	0.60	0.00	0.00	0.12	0.89
	AT	2	-4.28	0.60	0.00	0.64	0.74	0.00	0.00	0.05	0.91
	BE	2	-4.27	0.30	0.00	0.46	0.71	0.10	0.01	0.04	0.84
	IE	2	-5.96	0.00	0.00	0.11	0.09	0.00	0.00	0.00	0.13
Unemployment rate											
	IT	1	-3.28	0.83	0.00	0.00	0.00	0.02	0.56	0.64	0.84
	ES	1	-4.17	0.78	0.02	0.09	0.00	0.08	0.69	0.59	0.99
	DE	1	-5.59	0.72	0.03	0.21	0.00	0.25	0.56	0.10	0.98
	FR	1	-2.90	0.84	0.01	0.06	0.00	0.01	0.35	0.81	0.92
	PT	1	-2.88	0.88	0.03	0.02	0.00	0.07	0.81	0.38	0.91
	GR	1	-3.24	0.67	0.00	0.01	0.01	0.01	0.92	0.40	0.97
	NL	1	-4.04	0.74	0.02	0.44	0.07	0.23	0.10	0.35	0.92
	FI	1	-3.38	0.81	0.01	0.11	0.02	0.04	0.01	0.40	0.84
	AT	1	-5.38	0.65	0.00	0.13	0.03	0.18	0.08	0.09	0.82
	BE	1	-6.45	0.78	0.01	0.23	0.01	0.06	0.18	0.08	0.84
	IE	1	-3.03	0.81	0.03	0.11	0.00	0.19	0.59	0.51	0.97
Inflation rate											
	IT	2	-3.99	0.73	0.13	0.01	0.07	0.08	0.22	0.07	0.87
	ES	2	-4.24	0.67	0.18	0.21	0.41	0.05	0.02	0.29	0.91
	DE	2	-4.59	0.64	0.11	0.07	0.20	0.04	0.03	0.24	0.87
	FR	2	-4.64	0.84	0.14	0.08	0.30	0.03	0.04	0.13	0.78
	PT	2	-5.57	0.54	0.13	0.28	0.38	0.07	0.09	0.09	0.94
	GR	2	-4.67	0.62	0.04	0.09	0.54	0.11	0.10	0.10	0.74
	NL	2	-6.63	0.40	0.11	0.01	0.01	0.10	0.29	0.02	0.84
	FI	2	-4.14	0.71	0.02	0.27	0.00	0.07	0.22	0.01	0.90
	AT	2	-3.94	0.71	0.11	0.02	0.26	0.00	0.11	0.09	0.85
	BE	2	-4.04	0.70	0.21	0.01	0.28	0.00	0.05	0.11	0.86
	IE	2	-4.47	0.56	0.16	0.10	0.03	0.28	0.01	0.49	0.97
Euro area aggregate variables											
	Inflation	2	-6.62	0.61	0.00	0.00	0.00	0.00	0.97	0.29	0.98
	Inflation net of energy and unprocessed food	2	-2.64	0.25	0.00	0.01	0.01	0.00	0.93	0.33	0.95
	M1	2	-4.08	0.62	0.00	0.00	0.00	0.00	0.88	0.38	0.92
	M2	2	-6.01	0.10	0.01	0.00	0.03	0.03	0.69	0.45	0.83
	M3	2	-5.68	0.53	0.00	0.03	0.00	0.00	0.80	0.39	0.87
	Loans to NFCs	2	-8.03	0.17	0.01	0.20	0.02	0.10	0.15	0.02	0.64
	Loans to HHs	2	-6.28	0.16	0.01	0.09	0.14	0.05	0.04	0.32	0.57
	Dow Jones Euro Stoxx index	2	-6.55	0.17	0.00	0.13	0.13	0.00	0.44	0.45	0.74
	Dow Jones Euro Stoxx bank index	2	-8.92	0.08	0.00	0.16	0.15	0.02	0.01	0.22	0.49
	Oil price in euro	1	-4.52	0.82	0.00	0.02	0.03	0.04	0.68	0.41	0.87
	Unemployment rate	1	-8.67	0.22	0.03	0.71	0.66	0.06	0.00	0.00	0.92
	Retail trade	2	-8.97	0.25	0.07	0.32	0.55	0.25	0.01	0.00	0.76
	Three month Euribor - Eonia spread	1	-3.60	0.44	0.01	0.68	0.66	0.03	0.02	0.02	0.88
	Confidence index	1	-11.54	-0.09	0.01	0.52	0.65	0.11	0.00	0.00	0.82
	Nominal effective exchange rate	1	-6.20	0.54	0.01	0.54	0.53	0.02	0.03	0.01	0.71
	Bank rates on loans to NFCs	1	-9.47	0.22	0.05	0.18	0.22	0.01	0.01	0.03	0.31
	Bank rates on loans to HHs	1	-4.77	0.75	0.01	0.55	0.52	0.00	0.03	0.02	0.71
	Bank rates on HHs' deposits up to 1 year	1	-3.69	0.54	0.01	0.42	0.44	0.00	0.00	0.06	0.62
	Industrial production	2	-4.23	0.57	0.04	0.52	0.71	0.13	0.00	0.00	0.87
Sovereign spread											
	PT	1	-3.58	0.46	0.01	0.63	0.59	0.02	0.05	0.03	0.82
	IT	1	-8.00	-0.23	0.00	0.00	0.02	0.05	0.00	0.01	0.09
	ES	1	-3.67	0.74	0.17	0.07	0.27	0.04	0.04	0.19	0.93
	FR	1	-3.80	0.83	0.09	0.02	0.02	0.12	0.00	0.44	0.77
	BE	1	-3.14	0.69	0.05	0.10	0.02	0.02	0.27	0.04	0.92
	DE	1	-3.22	0.75	0.05	0.00	0.03	0.10	0.43	0.79	0.96
	NL	1	-4.51	0.74	0.03	0.01	0.02	0.09	0.32	0.84	0.96
	AT	1	-3.44	0.84	0.03	0.00	0.01	0.03	0.32	0.90	0.95
	FI	1	-2.28	0.88	0.05	0.46	0.13	0.14	0.34	0.32	0.97
	IE	1	-4.43	0.53	0.06	0.64	0.27	0.00	0.10	0.00	0.82
Export with rest of euro area											
	IT	2	-4.92	0.60	0.07	0.50	0.16	0.00	0.21	0.00	0.82
	ES	2	-3.20	0.86	0.00	0.00	0.03	0.13	0.38	0.01	0.63
	DE	2	-2.85	0.73	0.00	0.00	0.00	0.00	0.58	0.88	0.98
	FR	2	-10.94	0.02	0.02	0.54	0.41	0.01	0.07	0.25	0.86
	PT	2	-6.45	0.44	0.03	0.41	0.06	0.13	0.08	0.02	0.79
	GR	2	-4.73	0.72	0.00	0.56	0.55	0.02	0.00	0.16	0.89
	NL	2	-3.85	0.77	0.01	0.28	0.10	0.12	0.22	0.11	0.59
	FI	2	-6.59	0.32	0.01	0.01	0.00	0.00	0.14	0.94	0.99
	AT	2	-4.41	0.71	0.01	0.06	0.02	0.00	0.14	0.89	0.99
	BE	2	-3.49	0.78	0.00	0.09	0.00	0.05	0.00	0.58	0.97
	IE	2	-7.29	0.34	0.00	0.65	0.81	0.06	0.00	0.03	0.98

Note: column *(a)*: 1 denotes level, 2 12-month percentage changes; column *(b)* reports the results of the Augmented Dickey-Fuller test for the idiosyncratic component of each variable; column *(c)* reports the sum of the first four autoregressive coefficients estimated for the idiosyncratic component of each variable; column *(d)* reports the R^2 (*R-squared*) of the regression of each variable on each factor separately and on all of them (last column).

Figure 1.

Sovereign spreads (1)
(percentage points)



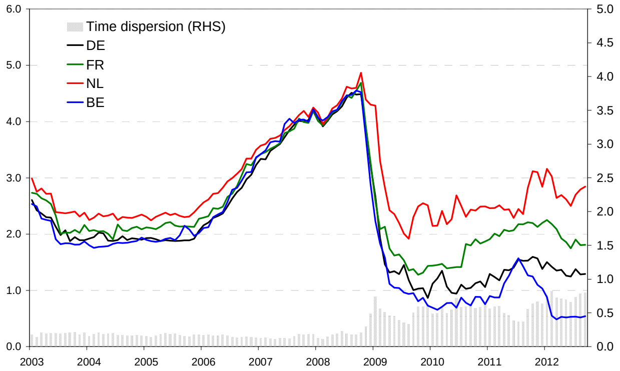
(1) Spread on 10-year government bonds with respect to the 10-year interest rate swap.

Figure 2.

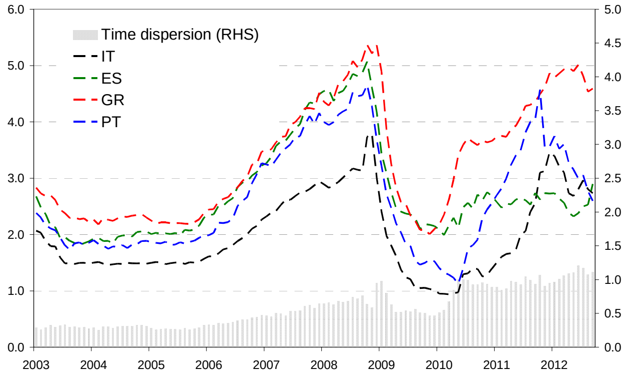
Interest rate on households' deposits with agreed maturity up to one year

(*percentage points*)

Core countries



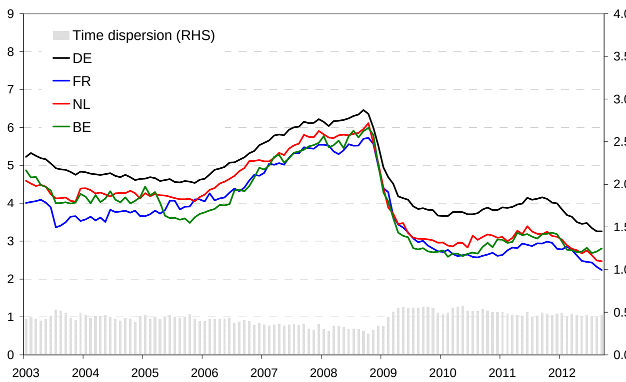
Peripheral countries



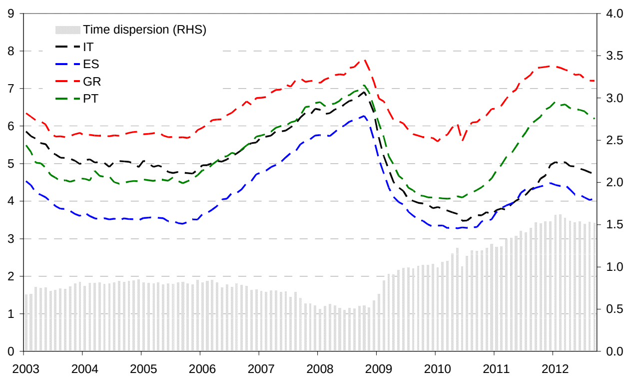
Interest rate on short-term loans to non-financial corporations

(*outstanding loans with maturity up to one year, including overdrafts; percentage points*)

Core countries



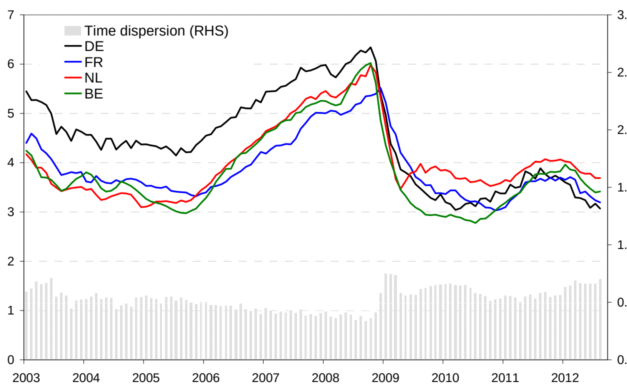
Peripheral countries



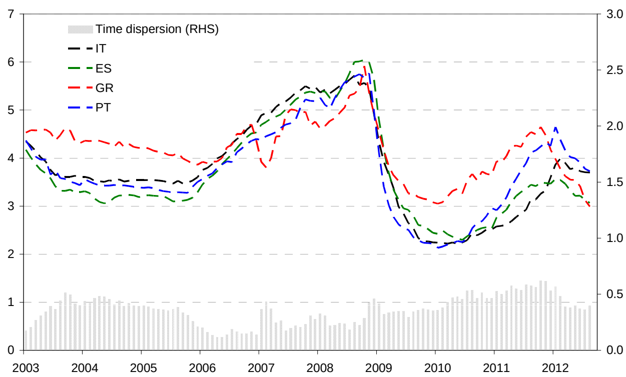
Floating interest rate on new loans to households for house purchases

(*percentage points*)

Core countries



Peripheral countries



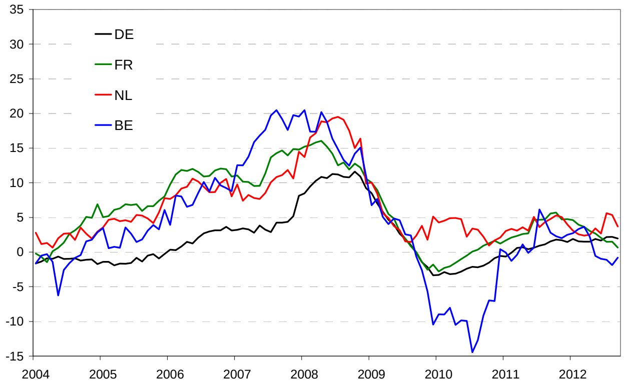
Note. DE = Germany, FR = France, NL = Netherlands, BE = Belgium, IT = Italy, ES = Spain, GR = Greece, PT = Portugal. The measure of time dispersion is calculated as the standard deviation of the interest rates at each point in time among countries.

Figure 3.

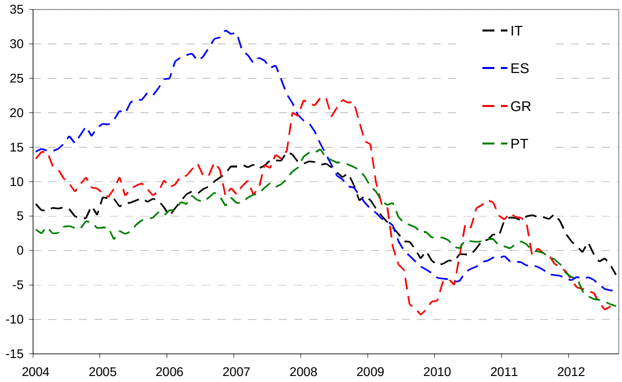
Twelve-month growth rate of loans to non-financial corporations

(*percentage points*)

Core countries



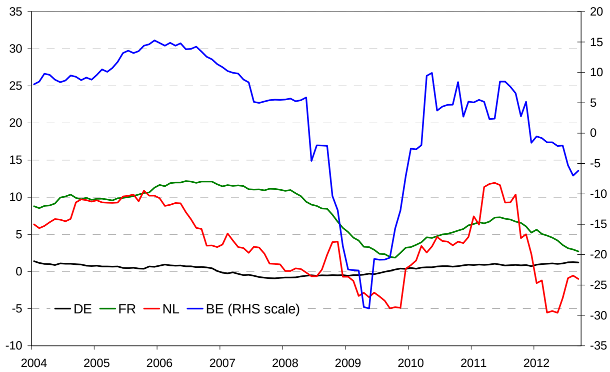
Peripheral countries



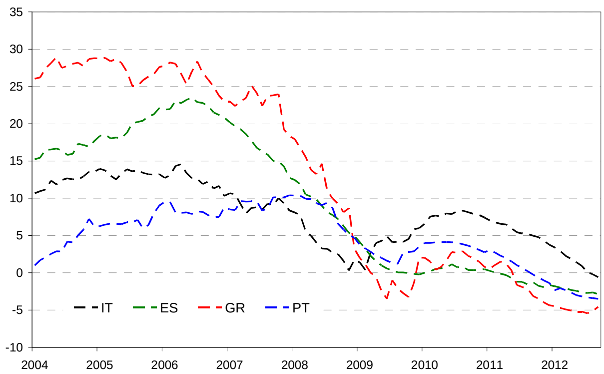
Twelve-month growth rate of loans to households for house purchases

(*percentage points*)

Core countries



Peripheral countries



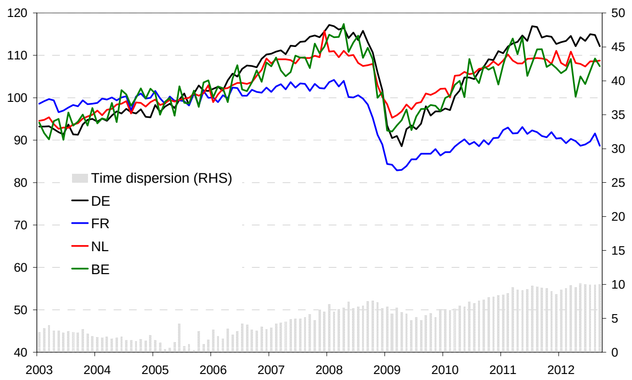
Note. Loans are not adjusted for sales and securitizations.

Figure 4.

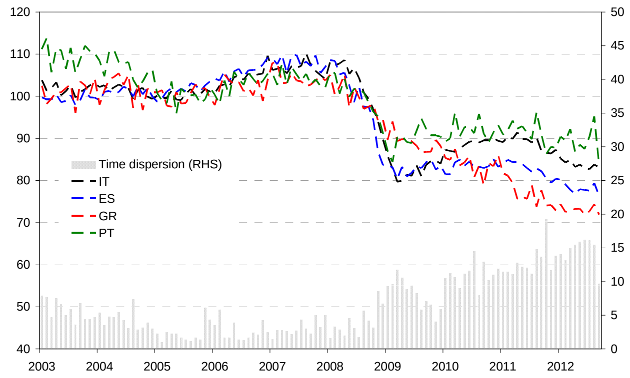
Industrial production

(base year = 2005)

Core countries



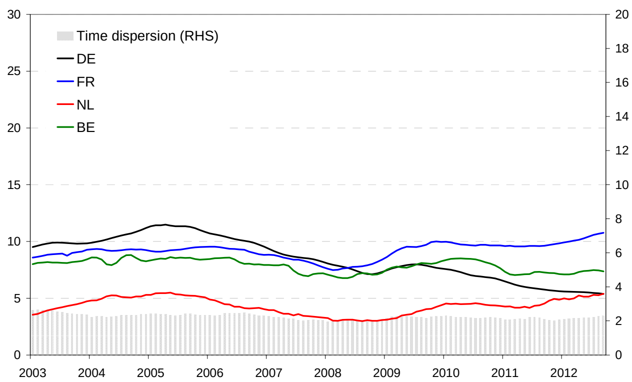
Peripheral countries



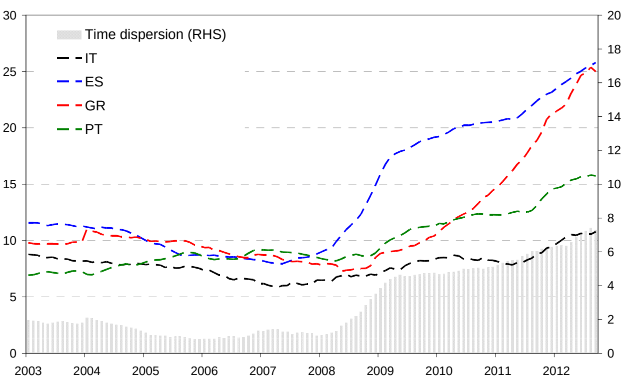
Unemployment rate

(percentage points)

Core countries



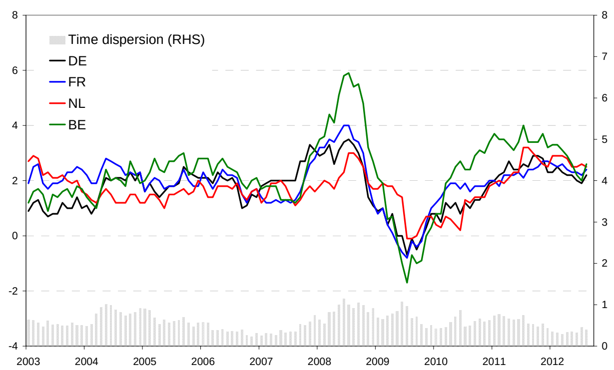
Peripheral countries



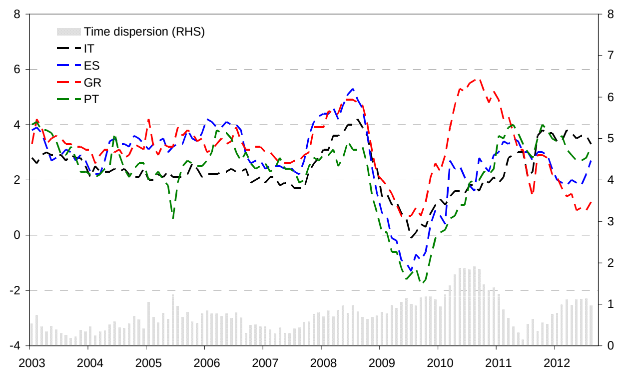
Inflation rate

(percentage points)

Core countries



Peripheral countries



Note. Inflation rate is based on the Harmonized Index of Consumer Prices.

Figure 5.

Eigenvalues

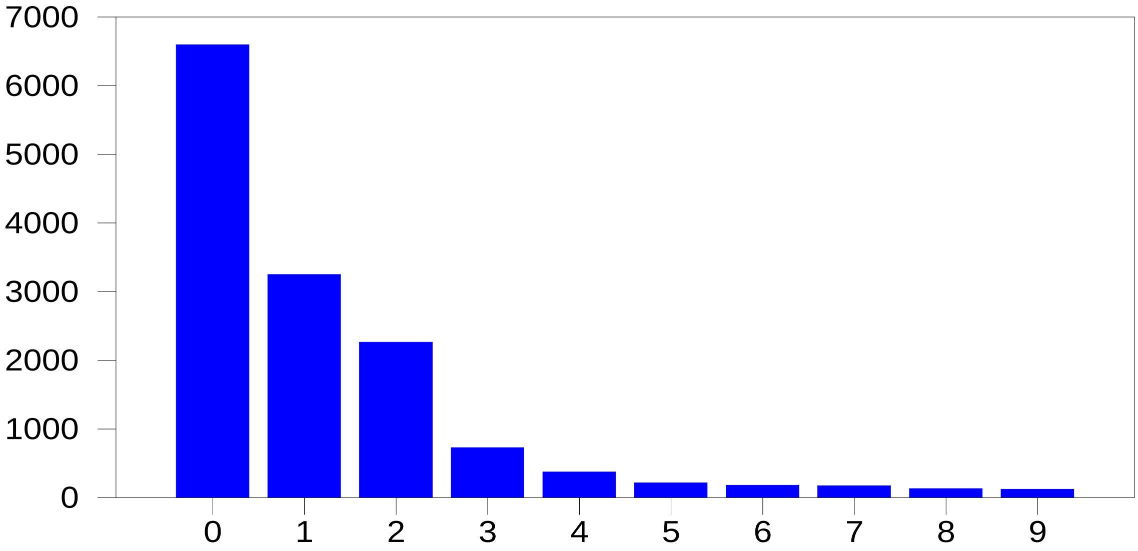
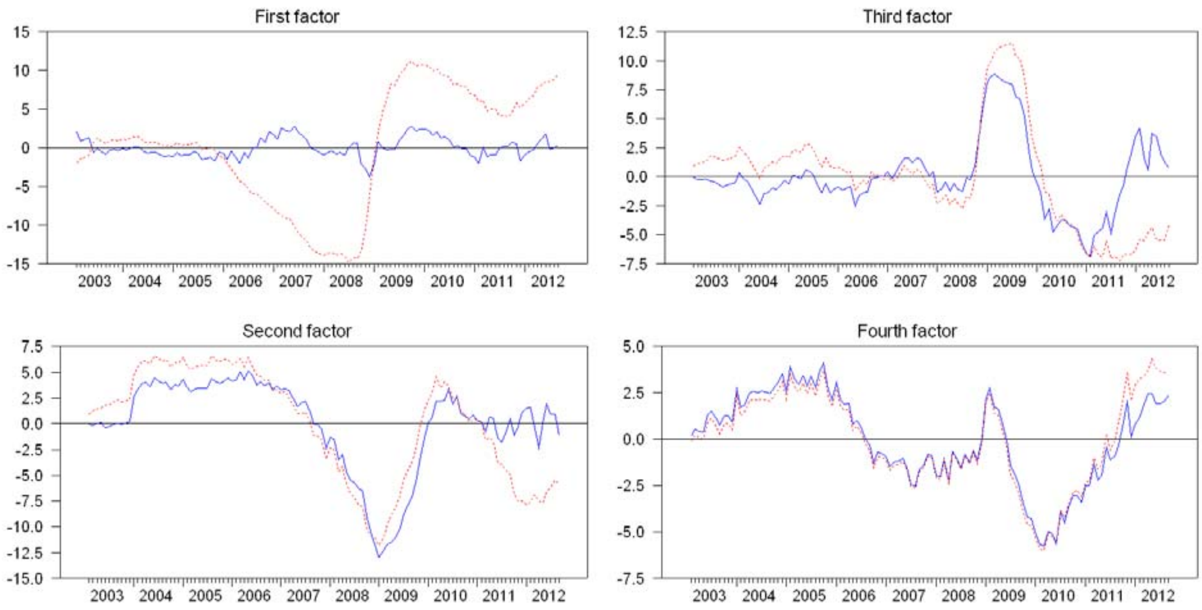


Figure 6.

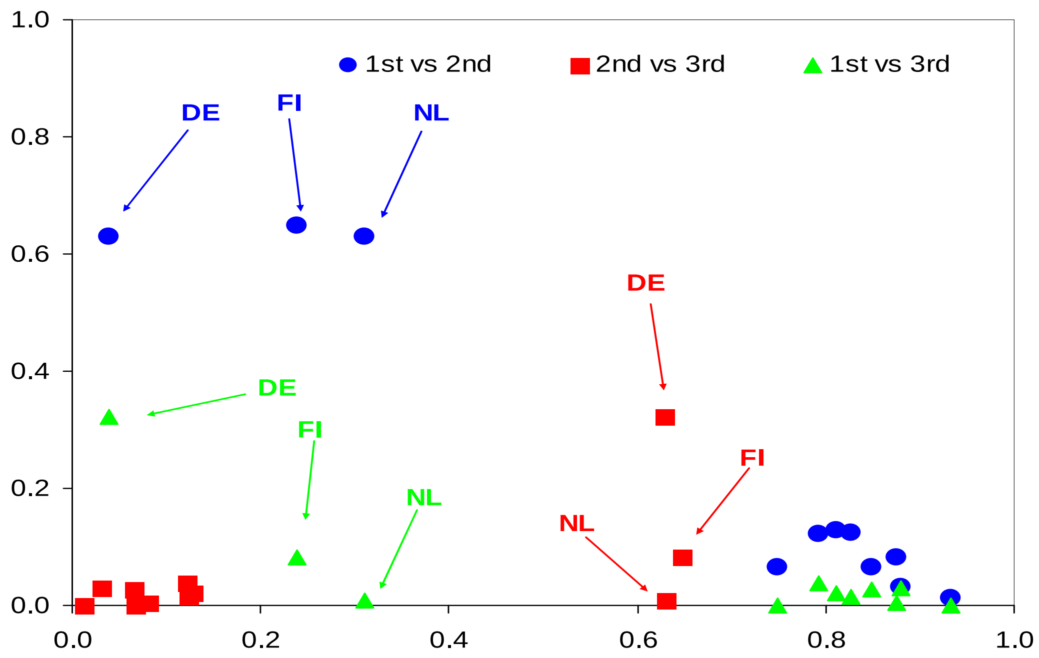
Latent factors



Note. Red dashed lines: unobserved factors extracted with principal components analysis; blue solid lines factors orthogonalized with respect to the MRO rate and the indicator of sovereign debt tensions.

Figure 7.

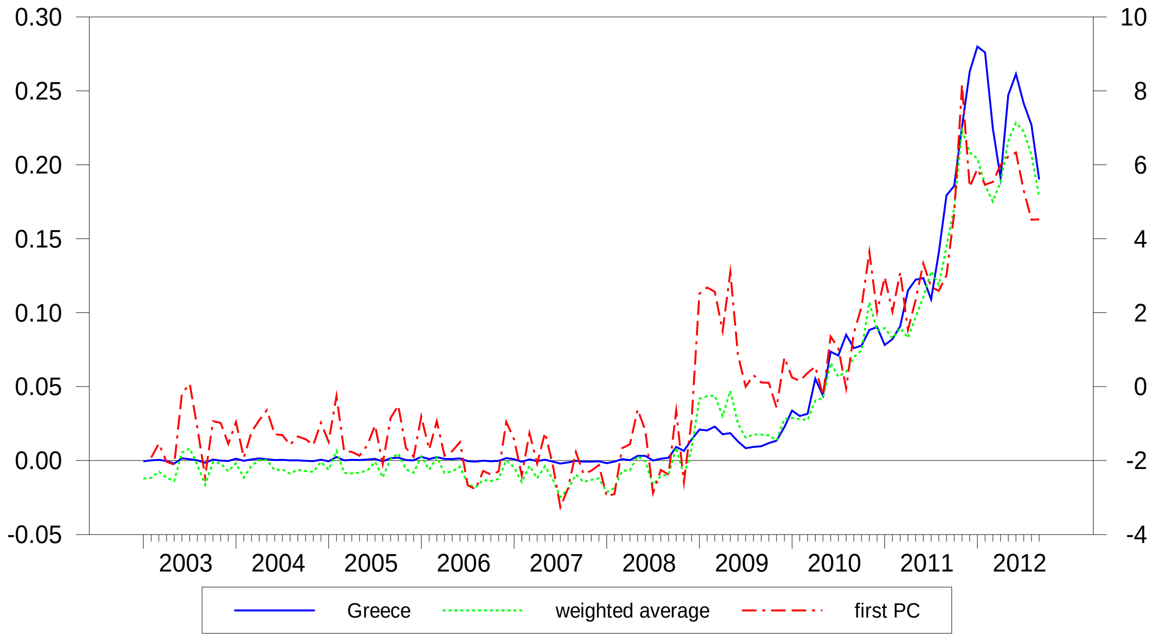
Principal component analysis of sovereign spreads



Note. Each pair measures the R^2 of a regression of sovereign spreads on each of the first three components separately.

Figure 8.

Indicators of sovereign debt tensions



Note. First PC = first principal component of the cross section of sovereign spreads; weighted average of sovereign spreads, multiplied by 10 (for ease of comparison), with the following weights: Italy = 0.172, Spain = 0.115, Portugal = 0.019, Greece = 0.021, Ireland = 0.019, France = 0.217, Belgium = 0.044, Austria = 0.033.

Figure 9.

The impact on sovereign spreads

(*percentage points; deviation from the baseline*)

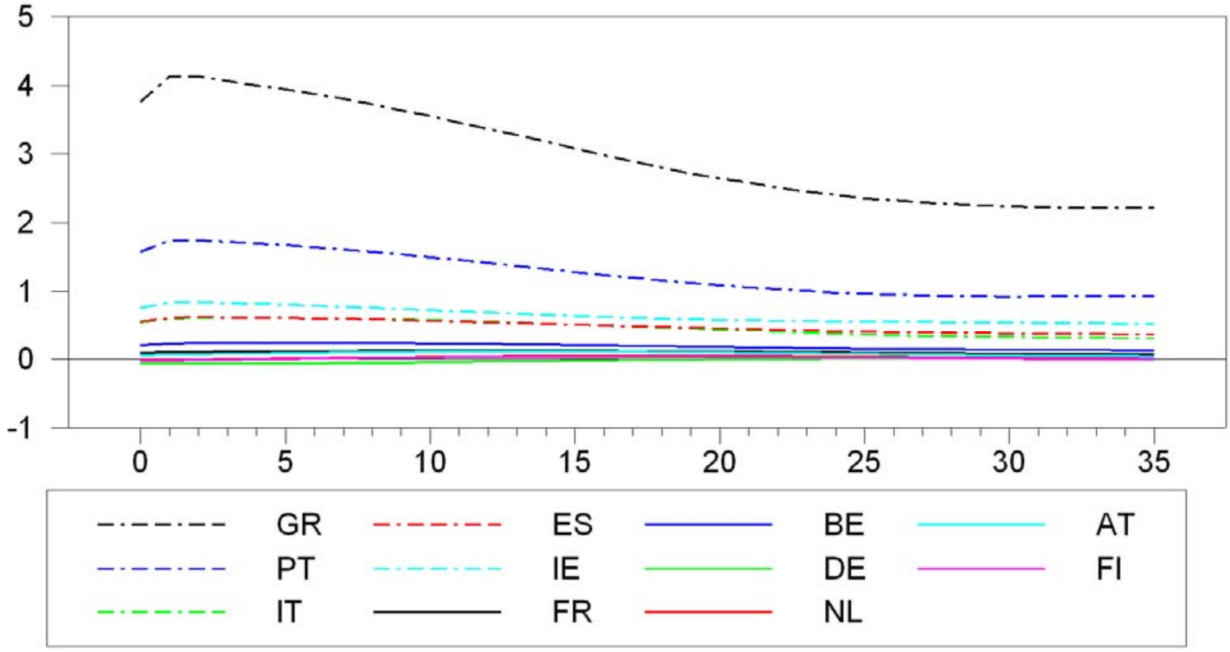


Figure 10.

The impact on the cost of new loans to non-financial corporations and households

(*percentage points; deviation from the baseline*)

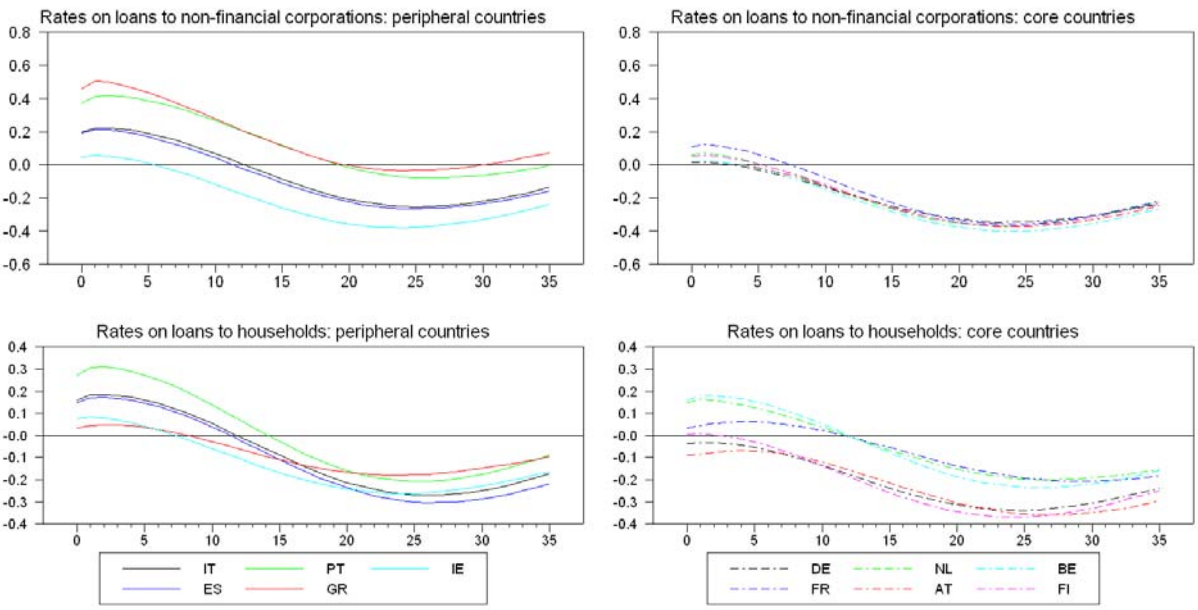


Figure 11.

The impact on loans to non-financial corporations and households

(*twelve-month percentage changes; deviation from the baseline*)

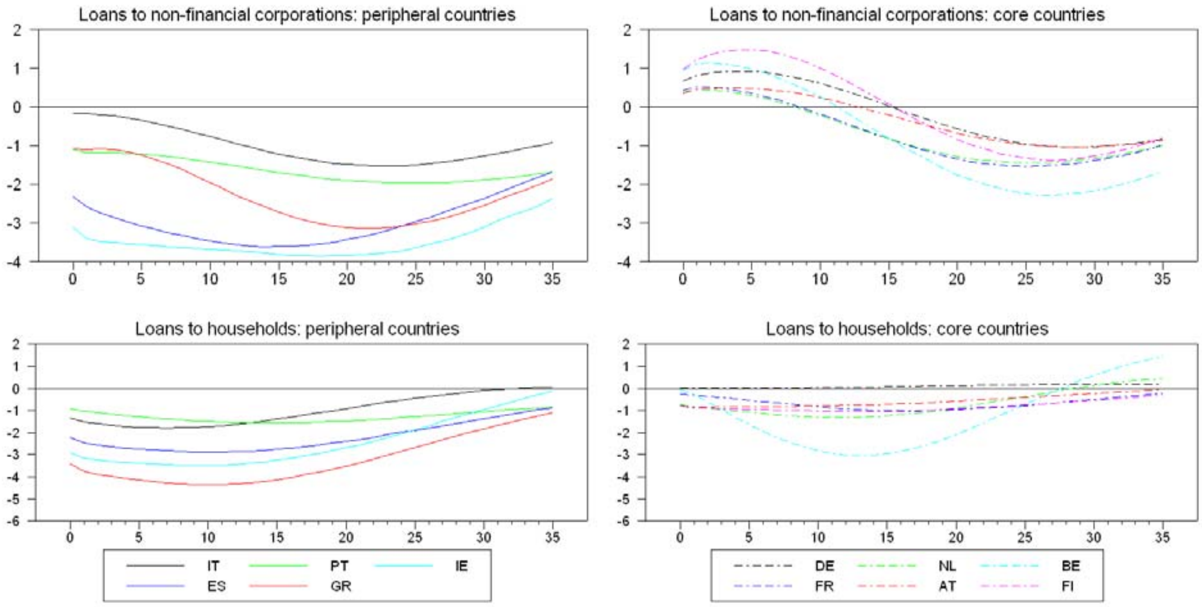


Figure 12.

The impact on M3

(*twelve-month percentage changes; deviation from the baseline*)

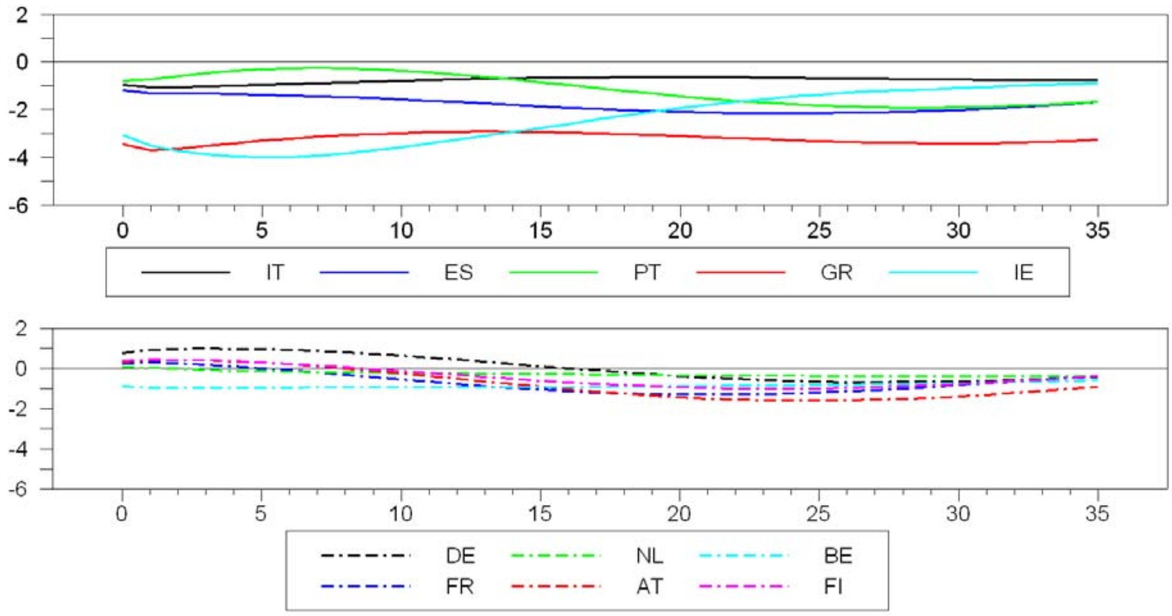


Figure 13.

The impact on industrial production

(*twelve-month percentage changes; deviation from the baseline*)

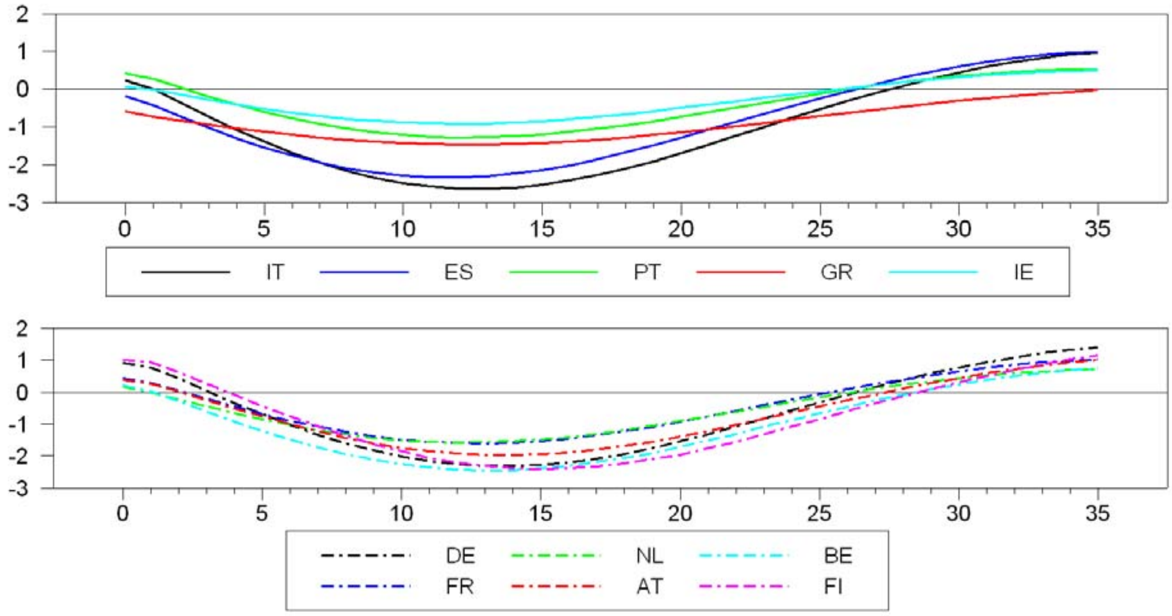
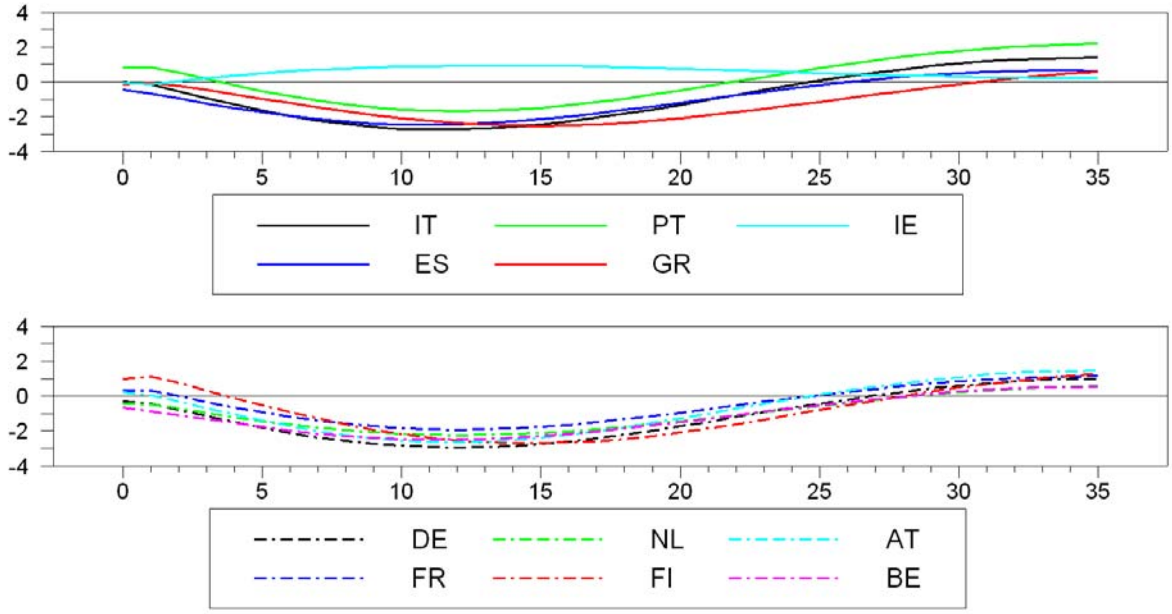


Figure 14.

The impact on exports within the euro area

(*twelve-month percentage changes; deviation from the baseline*)



Note. Response of each country exports to other euro-area member countries.

Figure 15.

The impact on unemployment

(*percentage points; deviation from the baseline*)

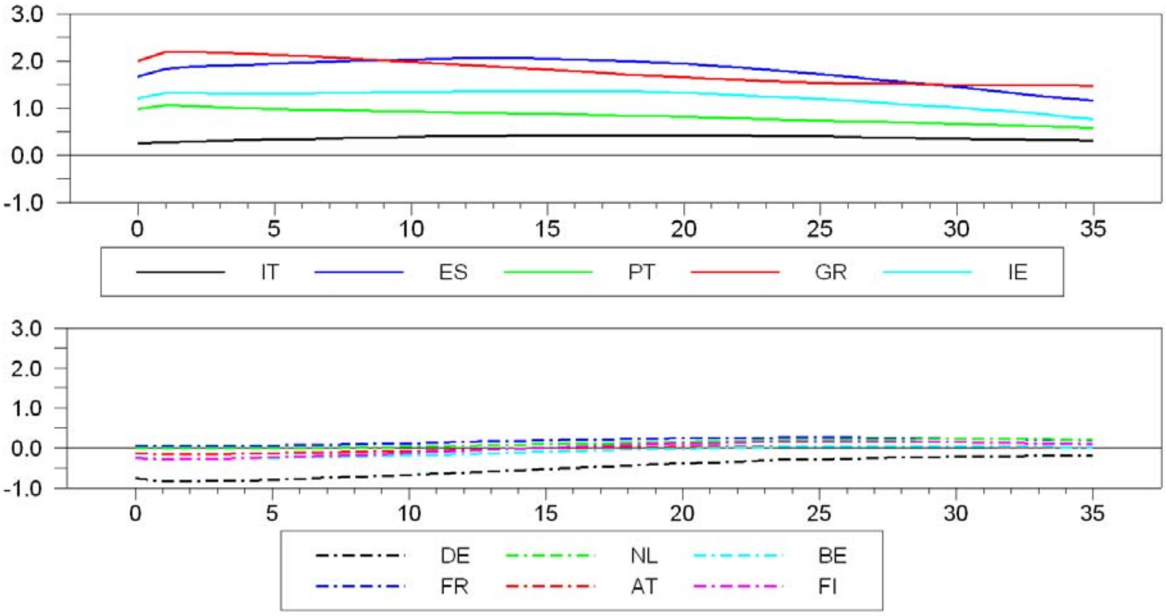


Figure 16.

The impact on inflation

(*percentage points; deviation from the baseline*)

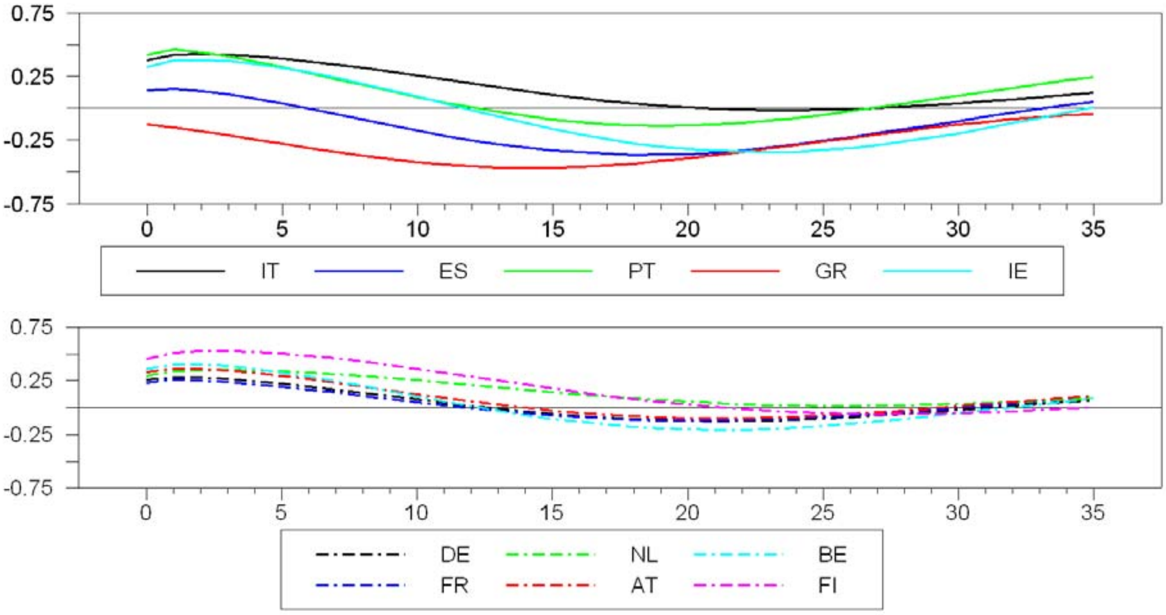


Figure 17.

The impact on the cost of credit to non-financial corporations and households in the euro area

(*percentage points; deviation from the baseline*)

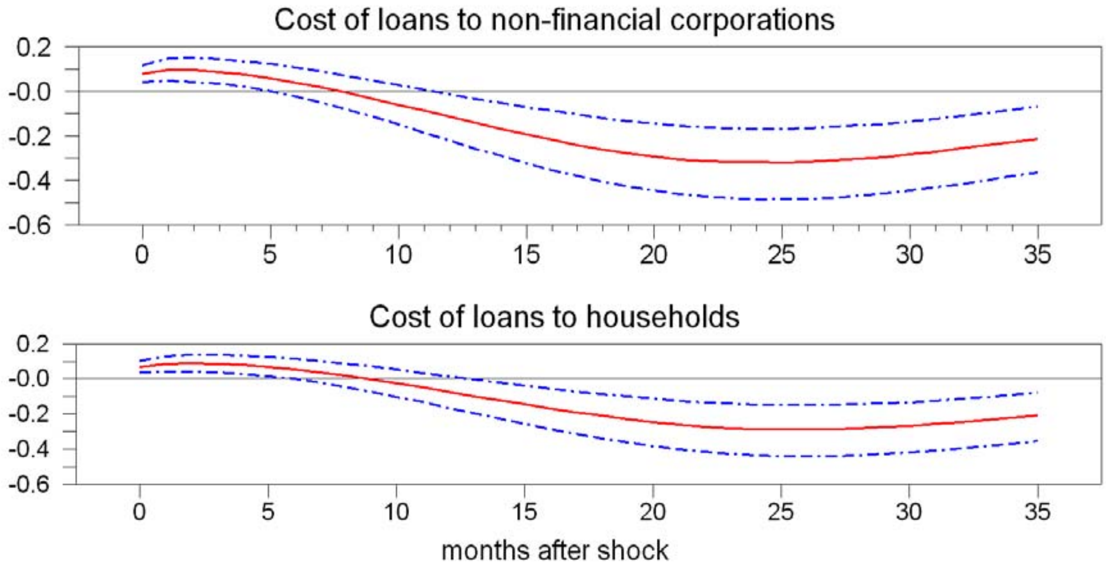


Figure 18.

The impact on monetary and credit aggregates in the euro area

(*twelve-month percentage changes; deviation from the baseline*)

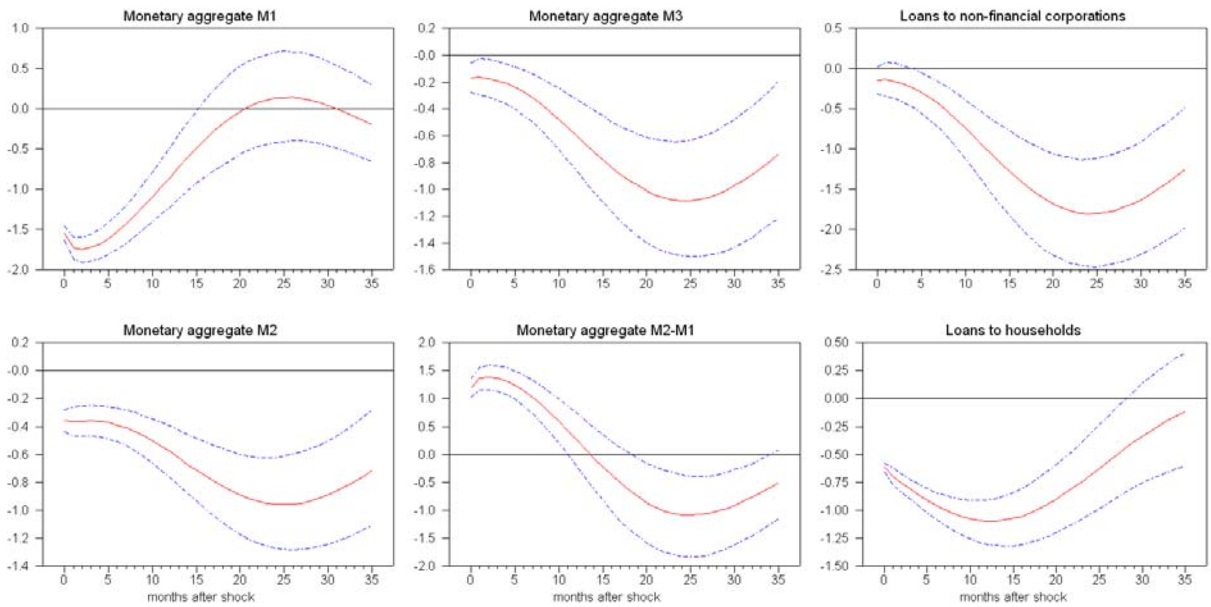
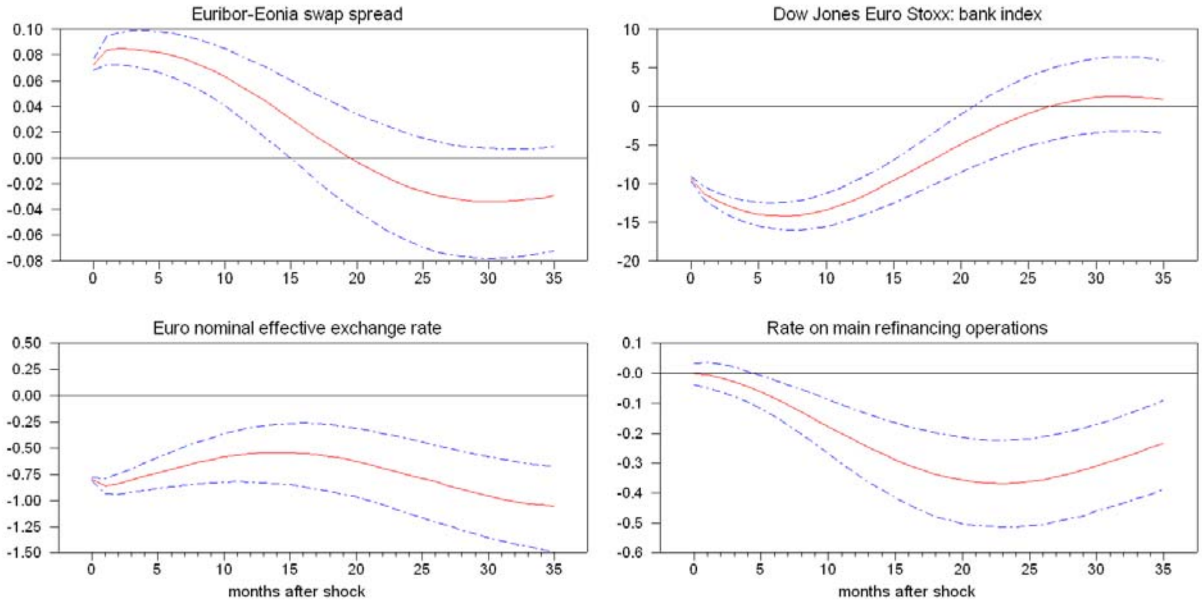


Figure 19.

The impact on the money market spread, banks' stock prices, the exchange rate of the euro and the ECB policy rate

(*twelve-month percentage changes and percentage points; deviation from the baseline*)



Note. A decrease in the nominal effective exchange rate denotes a depreciation.

Figure 20.

The impact on industrial production, unemployment, inflation and confidence in the euro area

(*twelve-month percentage changes, percentage points, per cent and net percentages; deviation from the baseline*)

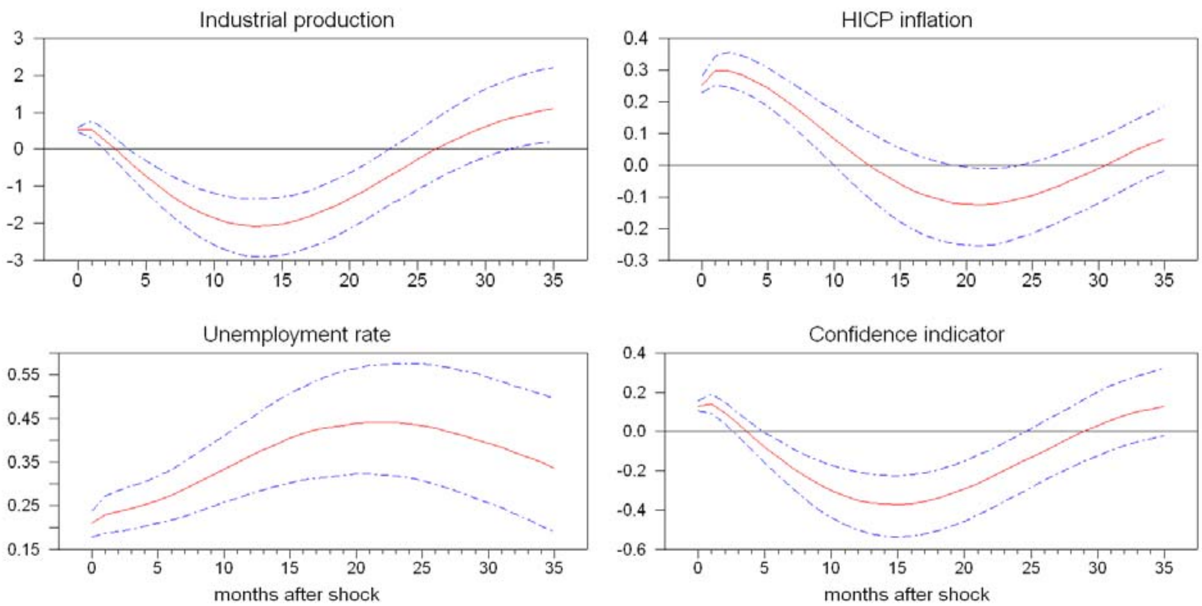


Figure 21.

The impact on sovereign spreads

(*percentage points; deviation from the baseline*)

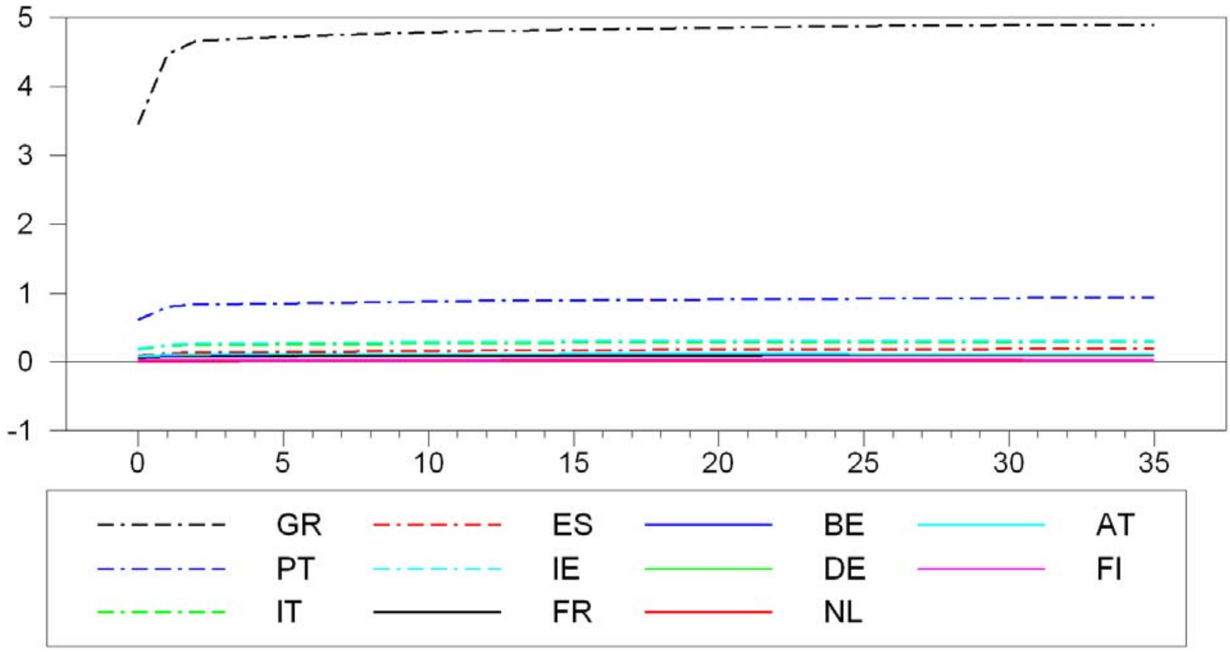


Figure 22.

The impact on the cost of new loans to non-financial corporations and households

(*percentage points; deviation from the baseline*)

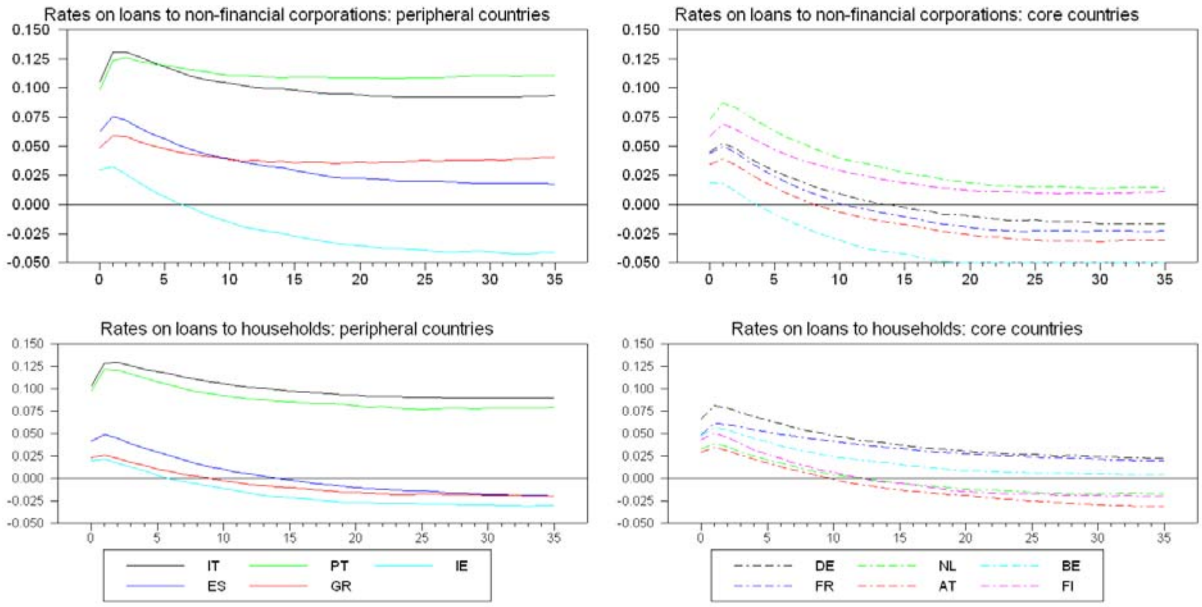


Figure 23.

The impact on loans to non-financial corporations and households
(*twelve-month percentage changes; deviation from the baseline*)

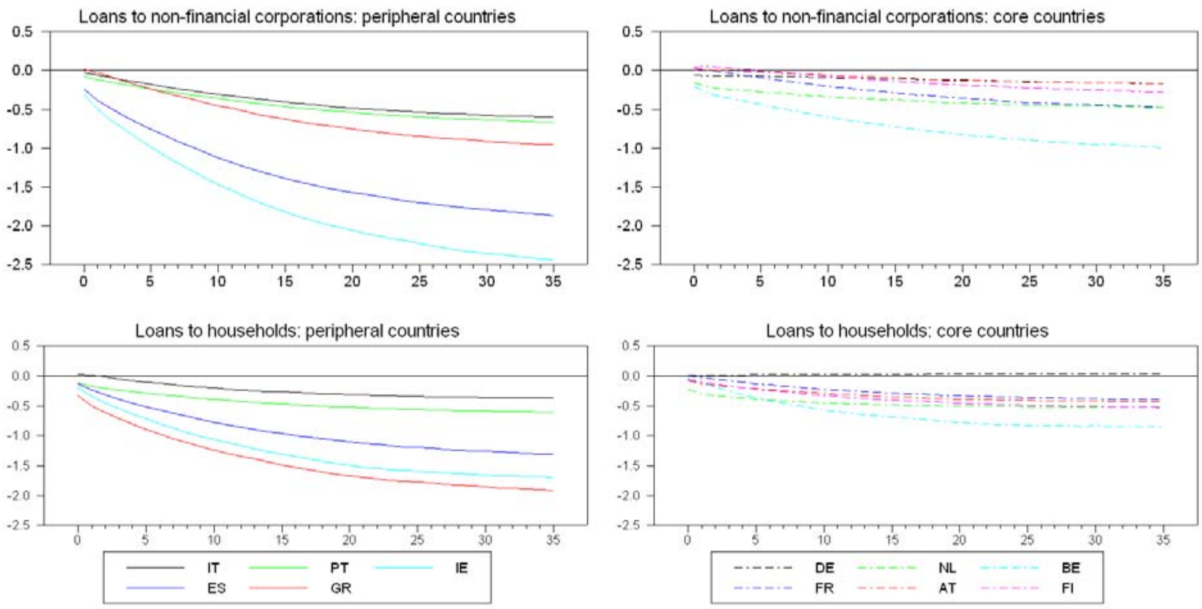


Figure 24.

The impact on M3
(*twelve-month percentage changes; deviation from the baseline*)

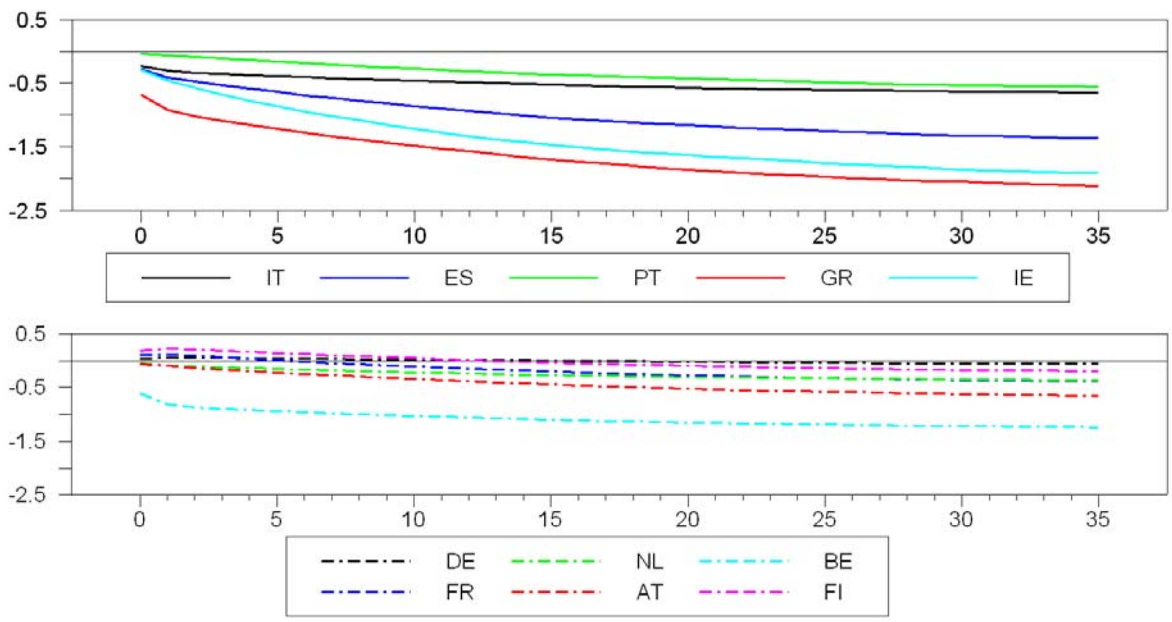


Figure 25.

The impact on industrial production
(*twelve-month percentage changes; deviation from the baseline*)

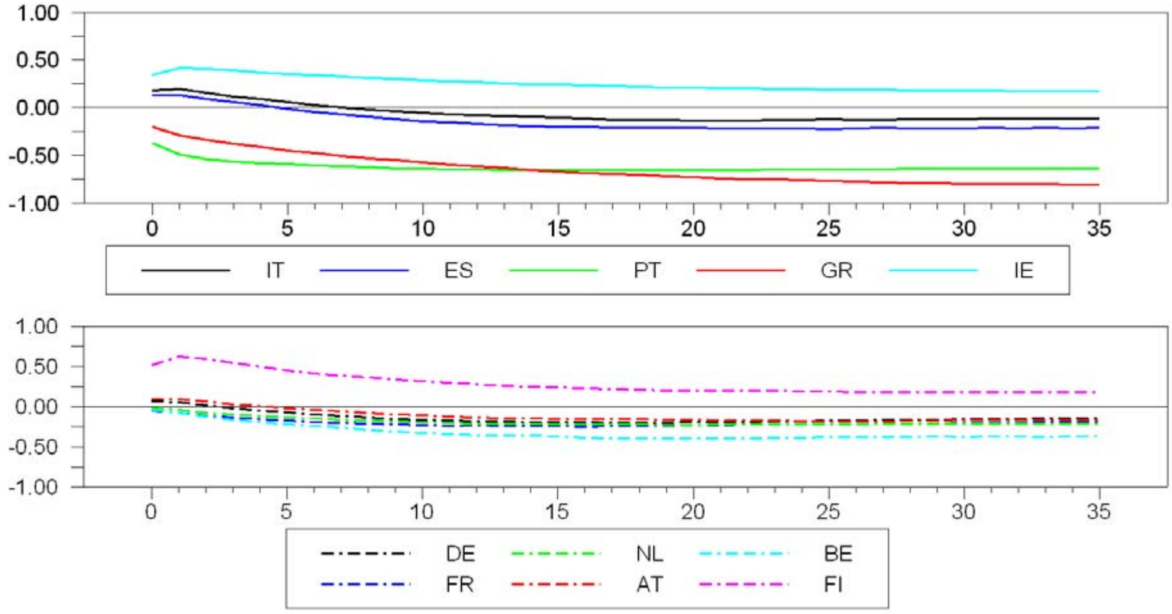


Figure 26.

The impact on unemployment
(*percentage points; deviation from the baseline*)

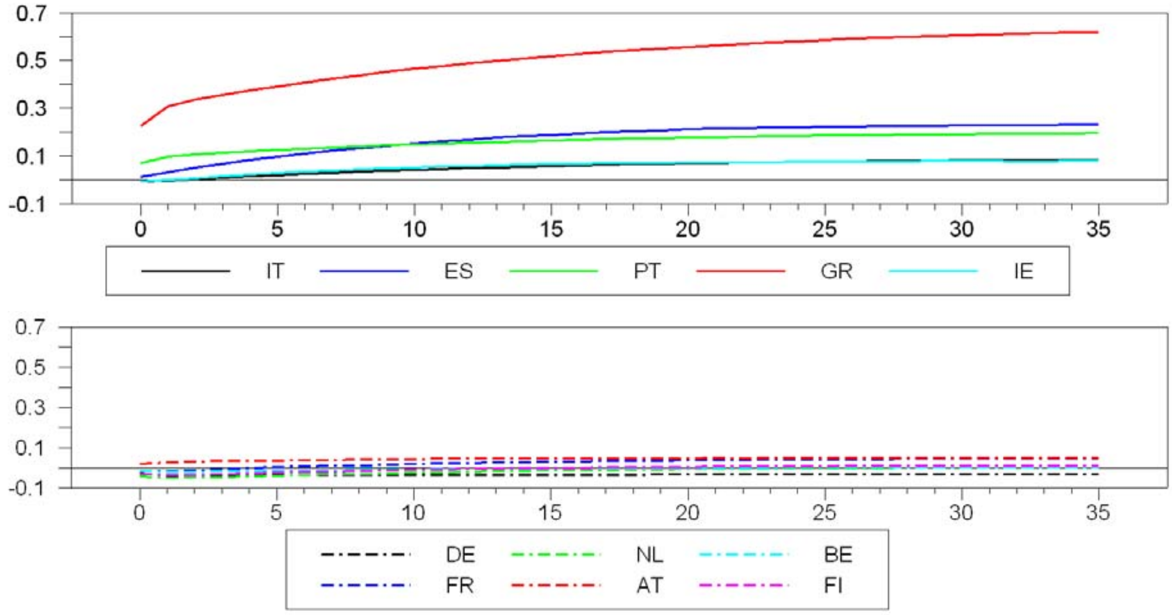


Figure 27.

The impact on monetary and credit aggregates in the euro area

(*twelve-month percentage changes; deviation from the baseline*)

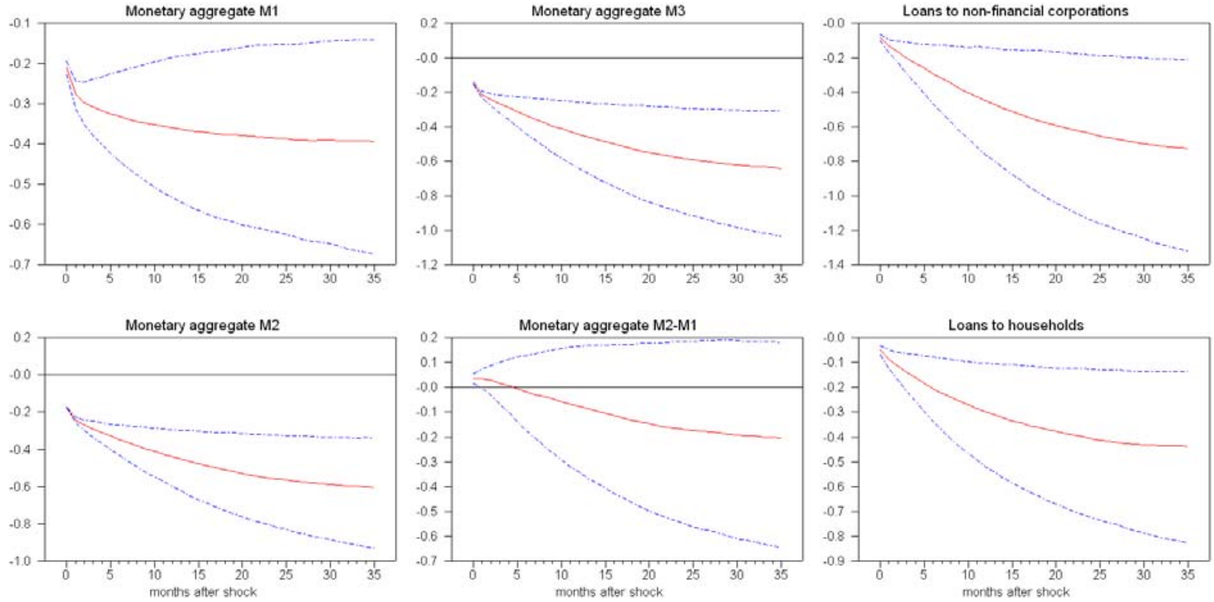
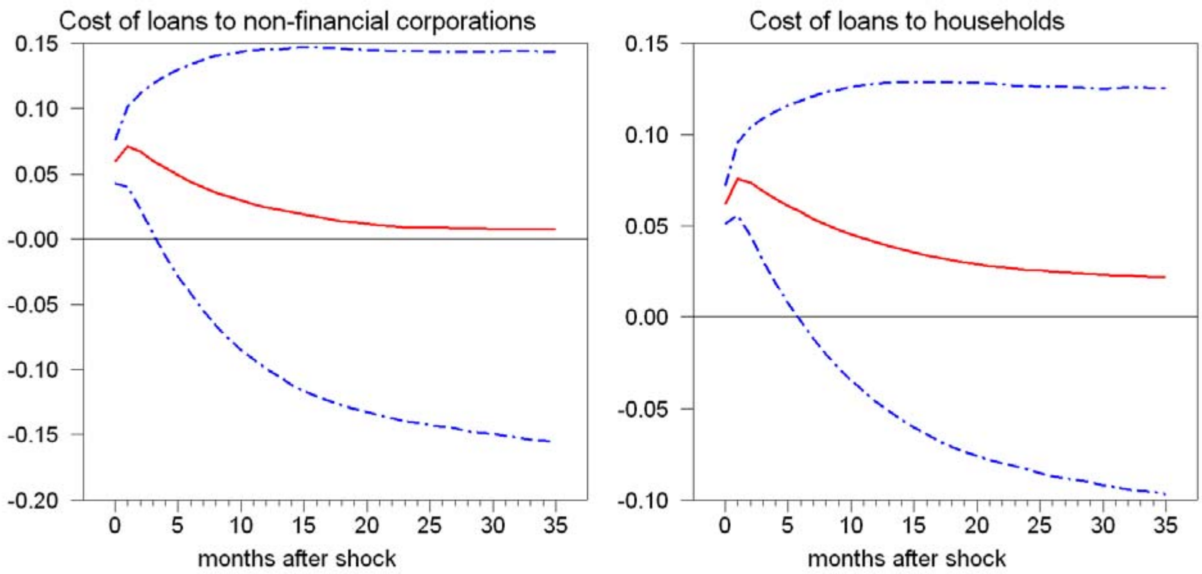


Figure 28.

The impact on the cost of credit to non-financial corporations and households in the euro area

(*percentage points; deviation from the baseline*)



Note. Weighted average of the responses at the country level.

Figure 29.

The impact on industrial production, unemployment, inflation and confidence in the euro area

(*twelve-month percentage changes, percentage points, per cent and net percentages; deviation from the baseline*)

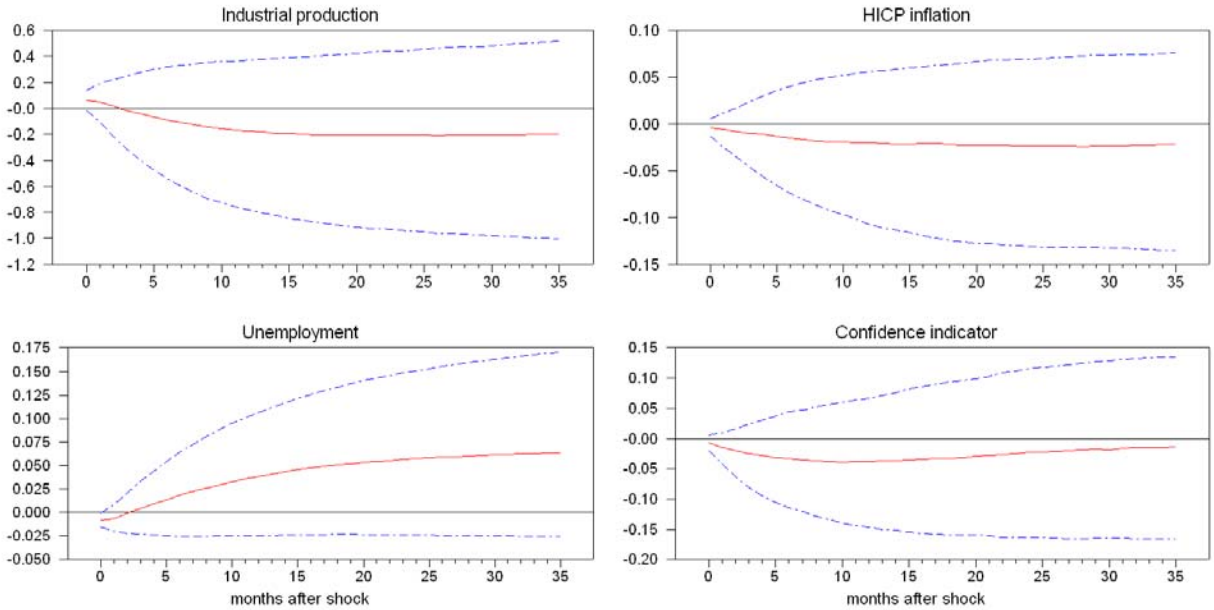
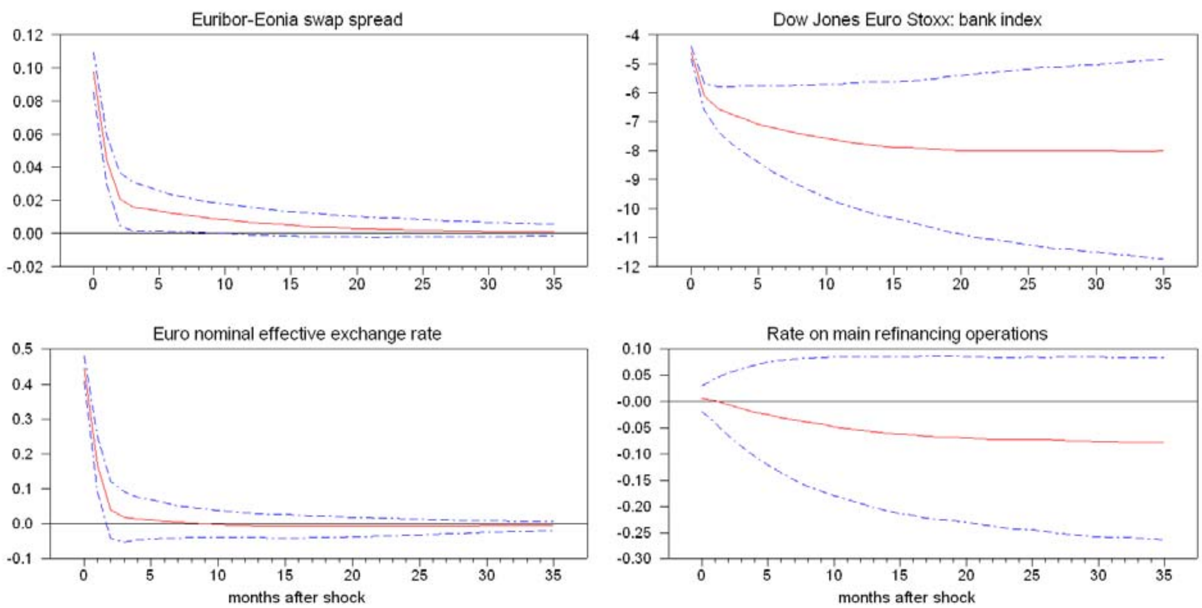


Figure 30.

The impact on the money market spread, banks' stock prices, the exchange rate of the euro and the ECB policy rate

(*twelve-month percentage changes and percentage points; deviation from the baseline*)

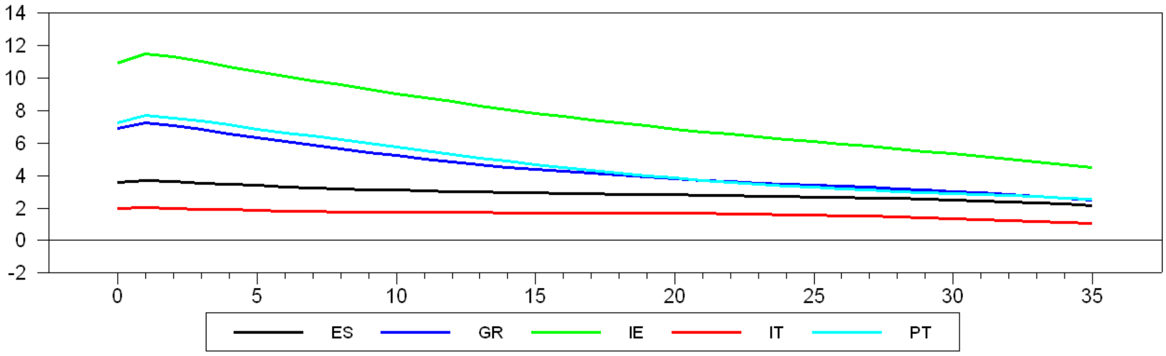


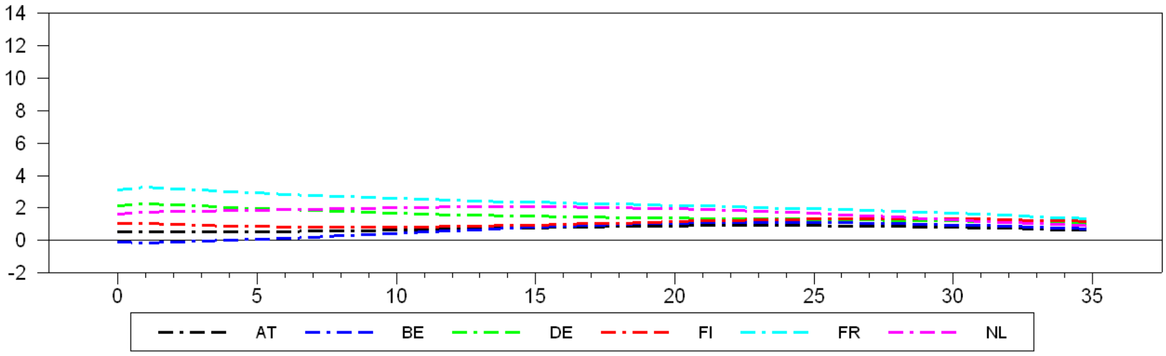
Note. A decrease in the exchange rate denotes a depreciation.

Figure 31.

The impact on public debt

(*percentage points; deviation from the baseline*)





Fiscal Multipliers, Monetary Policy and Sovereign Risk: A Structural Model-Based Assessment

Alberto Locarno*, Alessandro Notarpietro*, Massimiliano Pisani*

June 2013

Abstract

This paper provides a brief summary of the literature on fiscal multipliers and presents results for the Italian economy. The question of interest is whether an increase in government purchases leads to a greater than one-for-one increase in output, taking into account whether the zero lower bound is binding and whether initial conditions on public finances matter. Using a dynamic general equilibrium model for the Italian economy, we find that in general the government spending multiplier is lower than one, except when the monetary policy is stuck at the zero lower bound for an extended period of time. However, when the cost of borrowing for the sovereign is tightly linked to public finance conditions, the size of the government spending multiplier becomes much smaller. Tax multipliers are in all cases smaller than government consumption. Finally, we provide a tentative assessment of the fiscal consolidation measures adopted in Italy in 2011-2012 and find that the available evidence points to a much weaker impact on GDP that envisaged by the International Monetary Fund.

JEL Classification: E32, E52, E62.

Keywords: Fiscal multiplier, monetary policy, zero lower bound, sovereign risk.

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1 Introduction

Until the outbreak of the 2007 financial crisis there was wide consensus that discretionary fiscal policy was an ineffective tool for stabilizing aggregate demand and fighting recessions. This view was justified by the realization that the lags in implementing fiscal policy were typically too long to be useful for combating cyclical downturns and was reinforced by the econometric evidence on the size of the fiscal multiplier, generally estimated to be small, especially if the fiscal stimulus was eventually tax-financed. The crisis shattered these beliefs and when monetary policy interest rates hit the dreaded "zero lower bound (ZLB)" in several countries, it became abundantly clear that the report about the death of discretionary fiscal policy had been greatly exaggerated.

This paper reconsiders the question of the effectiveness of fiscal policy as a demand-management tool, by evaluating the size of the fiscal multiplier under different macroeconomic conditions. The question of interest is whether an increase in government purchases leads to a greater than one-for-one increase in output.[1] Particular attention is devoted to explaining whether there are reasons to believe that the size of the fiscal multiplier is larger under Depression-like circumstances and whether the initial stock of public-sector debt matters.[2] Our contribution is twofold. First, we survey the main theories about the size of the fiscal multiplier and discuss the existing empirical evidence supporting or disconfirming the competing views on the effectiveness of government spending. Second, using the Banca d'Italia dynamic stochastic general equilibrium (DSGE) model we estimate the size of the fiscal multiplier in Italy under alternative assumptions about the stance of monetary policy, the financing of the fiscal expansion and the role of sovereign risk.[3] In so doing, we also provide a tentative assessment of the macroeconomic effects of the fiscal consolidation measures adopted in 2011-2012 in Italy.

We are mainly concerned with the short-term impact of fiscal expansions, but we provide also an assessment of their long-term effects. Our main conclusions are the following. First, short-run fiscal multipliers are typically below one and, in particular, multipliers associated with taxation are lower than those associated with public expenditure. Second, public spending multipliers are substantially larger than one when the monetary policy rate is kept constant at the ZLB, but only if the policy rate remains at the ZLB for a sufficiently long period of time (at least five years in our simulations). Third, under conditions similar to those currently prevailing in the euro area (EA), in countries with a high public debt, the stimulus induces a deterioration of public finances and hence a rapid increase in the sovereign risk premium, which in turn substantially reduces the size of the multiplier and the effectiveness of fiscal policy. Fourth, the short-run contractionary effects of fiscal consolidation efforts can be partially mitigated by a reduction in

[1]Government spending is treated as pure waste in the analysis, in order to focus on the pure macroeconomic effects of fiscal policy as a determinant of aggregate demand in the short run.

[2]The exchange rate regime may also affect the strength of a fiscal stimulus. Corsetti et al. (2012b) mention other factors, in particular trade openness and the health of the financial system. None of them is extensively considered in this paper.

[3]We consider Italy as an illustrative case. Our results should hold to a large extent also for other euro area countries characterized by high public debt and precarious financial conditions.

the risk premium.

The rest of the paper is organized as follows. Section 2 provides a summary of the literature on fiscal multipliers and presents related empirical evidence. Section 3 presents the model used in the simulations and elaborates on model calibration. Section 4 illustrates the simulation exercises and shows fiscal multipliers under different macroeconomic scenarios. Section 5 concludes.

2 Theory and evidence on the effects of government spending

Government spending can boost economic activity only if it increases hours worked: as the capital stock cannot instantaneously adjust and technical progress is unresponsive to fiscal stimuli, in the short run output can increase only if more labour is used in production. Thus, the value of the fiscal multiplier is tightly linked to the effect of government spending on hours worked, though the channels through which the former affects the latter vary according to whether a Keynesian or Neoclassical (viz. real-business cycle) perspective is adopted. The value of the fiscal multiplier depends on (i) the length of the policy stimulus; (ii) how the budget slippage is financed; (iii) whether the monetary policy responds or not (e.g. because the binding ZLB keeps the policy rate well above the desired level); (iv) which are the country's initial conditions (namely the extent of resources left idle by the lack of aggregate demand and the size of outstanding public debt). Each of them must be properly accounted for in order to provide a reliable assessment of the macroeconomic impact of a change in the discretional component of government spending.

2.1 Neoclassical approaches

According to the neoclassical paradigm, a debt-financed increase in government purchases - unexpected when it occurs but known to be permanent as soon as it is implemented - has a negative wealth effect on households, related to the expected payment of higher taxes in the future. Individuals respond by reducing consumption and leisure, as long as both are normal goods.[4] Because the increase in the labour input shifts up the marginal product schedule for capital, investment rises and remains higher than in the no-stimulus scenario; it stops increasing only when the pre-shock level of the capital-labour ratio is restored. In response to the upward jump in labour supply, the real wage declines and the rental rate of capital increases symmetrically; these factor-price movements are however temporary, as the accumulation of capital ultimately restores the original situation. According to Baxter and King (1993), the long-run fiscal multiplier is 1.16, corresponding to a 0.2 percentage point fall in consumption and a 0.3 rise

[4]The most-cited reference on this regard is Baxter and King (1993). The numbers for the fiscal multipliers quoted in this section refer to their paper. Under fairly general conditions, there is no difference between a debt-financed and a tax-financed fiscal stimulus, provided the latter is based on lump-sum taxes. Baxter and King (1993) consider a fiscal expansion financed by lump-sum taxes.

in investment;[5] welfare is unambiguously lower, as the representative agent consumes less and works more. Starkly different results are obtained when the increase in government spending is temporary.[6] As before, agents, who are hit by a negative wealth shock, save and work more; unlike before, investment falls, due to the increased government absorption of resources. On impact, output increases, though less that in the previous case. After T years, when public spending is back to the pre-stimulus level, investment jumps above its long-run level and gradually declines thereafter; consumption and leisure remain below the steady-state equilibrium level and so does output. Eventually, all variables revert to their steady-state level and the original equilibrium is restored. It is worth stressing that when the fiscal stimulus is withdrawn, output falls below the pre-shock level, reducing the growth rate of the economy, and stays there indefinitely.

Government purchases that are financed by means of distortionary taxes have radically different effects, as they lessen rather than boost output. The "stimulus" works as follows. First, the increase in tax rates creates a gap between marginal productivity and (net) factor compensation and hence reduces individuals' incentives to work and invest. Second, the fall in labour supply and capital accumulation compresses the tax base and calls for higher tax rates to balance the public-sector budget. Third, heavier fiscal duties depress output even more and force the government to find additional revenues. The downward spiral that ensues brings output well below the pre-shock level. According to Baxter and King (1993), the fiscal multiplier may become as low as -2.5, implying that private-sector spending is completely crowded out and tax distortions discourage work effort and capital accumulation.

In case of a tax-financed fiscal stimulus that lasts for T years, temporarily low after-tax factor rewards induce households and firms to increase leisure and postpone investment; output declines and remains below baseline for as long as the stimulus lasts. When the stimulus ends (and tax rates return to their normal level), hours worked and capital accumulation immediately increase, pushing output slightly above baseline. Eventually, the initial equilibrium is restored. To summarize, the Neoclassical theory provides three main insights: (1) permanent changes in government purchases exhibit a multiplier that is greater than 1;[7] (2) temporary fiscal stimuli are less effective, even in terms of the impact multiplier, which tends to be smaller than 1;[8] (3) financing decisions are of the utmost importance, as they can not only reduce the size of the multiplier, but also change its sign.

[5]As for the change in government spending, variations in consumption and investment are measured in terms of units of output.

[6]Once again it is assumed that the increase in government spending is unanticipated but, as soon as it occurs, it is know to last for T years.

[7]A short-run multiplier greater than 1 is also possible if the labour supply is highly elastic.

[8]The finding that temporary stimuli are less effective than permanent ones is not trivial. Barro (1981) and Hall (1980) reach opposite conclusions.

2.2 Keynesian approaches

Keynesian analysis focuses on situations in which aggregate demand is the binding constraint on production and employment. The essential policy implication of Keynesian analysis is that any increase in aggregate spending, from whatever source, will induce firms to expand production and draw workers into employment without necessitating any change in wages or prices. Under the assumptions that (i) the economy is closed; (ii) there is no capital; (iii) monetary policy does not respond to the fiscal stimulus; (iv) government spending is debt financed, then the multiplier corresponding to a permanent increase in government purchases is equal to the reciprocal of the marginal propensity to save. Allowing for foreign trade or for a response of the monetary policy maker, reduces the size of the output expansion; the opposite happens if capital accumulation is taken into account. Even if the fiscal stimulus is tax-financed, the multiplier remains positive and big, as the Haavelmo theorem shows.[9] If instead it is temporary, the size remains the same as in the permanent case, but it becomes zero as soon as the government stops spending. Accordingly, a temporary fiscal stimulus simply shifts aggregate demand from one period to another: first it provides a boost to growth, then it subtracts from it.

New Keynesian models generate predictions that are in-between those consistent with the Neoclassical and the Keynesian theories. Since New Keynesian models add sticky prices and other frictions to the real business cycle theory, neoclassical features tend to mute the Keynesian multiplier. Galí et al. (2007) show however that the traditional Keynesian predictions can be restored if two ingredients are added, namely: (1) a sufficiently high proportion of rule-of-thumb consumers, that helps increasing the marginal propensity to consume;[10] (2) an elastic labour supply, that makes workers willing to offer as many hours as firms demand.[11] Both assumptions however boil down to make the models heavily dependent on non-optimising behavior and are therefore not entirely appealing.

2.3 ZLB, hysteresis and (other) initial conditions

Monetary policy reacts to demand shocks that drive output and inflation up; accordingly, in normal times the value of the fiscal multiplier is low, as the fiscal stimulus is to a large extent offset by the response of the central bank. In severely depressed economies, in which the policy interest rate is well above the desired level because of the ZLB, this is no longer the case. A

[9] See Haavelmo (1945).

[10] Rule-of-thumb consumers are non-Ricardian. They consume just what they earn, regardless of the impact of government spending on the inter-temporal budget constraint. The higher the share of these non-optimizing agents, the lower the (negative) impact of wealth effects on consumption and the higher the multiplier.

[11] Households' labour-supply decision is driven by the intra-temporal equilibrium condition, which states that the marginal utility of leisure (u_l) must be equal to the (real) wage rate (w) times the marginal utility of consumption (u_c), i.e. $u_l = wu_c$. Because of the negative wealth effect of additional government spending, consumption falls and its marginal utility increases; to restore the equilibrium, either leisure has to diminish (i.e. hours worked have to increase) and/or the real wage has to fall. In the standard Neo-classical (i.e. real business cycle model) both things happen. By preventing the real wage to change, all the adjustment is born by the labour supply, that accordingly has to increase more, boosting the output response to a fiscal stimulus.

stream of the literature has recently resumed the Keynesian argument that government spending is likely to boost aggregate demand much more in times of recession than during booms, in particular when the monetary policy interest rate is stuck at the ZLB. Examples are Christiano et al. (2011), Eggertsson (2001) and Woodford (2011). The story runs as follows. An increase in government spending when the ZLB is strictly binding leads to a rise in output, marginal cost and expected inflation; with the nominal interest rate stuck at zero, the rise in expected inflation drives down the real interest rate, which drives up private spending; this rise in spending leads to a further rise in output, marginal cost, and expected inflation and a further decline in the real interest rate. The net result is a large rise in output and a large fall in the rate of deflation: the increase in government consumption counteracts the deflationary spiral associated with the ZLB state. The value of the government-spending multiplier depends on how long the ZLB is expected to be binding. Christiano et al. (2011) also respond to the practical objection that using fiscal policy to counteract a contraction associated with the ZLB state is unfeasible, as there are long lags in implementing increases in government spending: they state that the case for fiscal stimulus while the constraint binds applies only to the case in which the increased government purchases will be terminated as soon as the constraint ceases to bind.[12] Christiano et al. (2011) provide also some estimates, obtained with a DSGE model, of the size of the fiscal multiplier: under the assumption that government spending lasts for 12 quarters and the nominal interest rate remains constant, the impact multiplier is roughly 1.6 and reaches a peak value of about 2.3. However, large estimates of the spending multiplier implicitly rely upon the assumption that non-standard monetary policy measures are incapable of stimulating aggregate demand and preventing a deflationary spiral.

Another factor that may affect the size of the fiscal multiplier is the presence of hysteresis, especially in the labour market. The concept of hysteresis was borrowed by economists from its original application to physical systems. The key idea is that transitory causes may have permanent effects. The concept of hysteresis, first used by Blanchard and Summers (1986), has been recently revived by DeLong and Summers (2012). Their main claim is that in a depressed economy hysteresis is important and once it is taken into account, the impact on output of additional government purchases can become so large as to be self-financing. They define a depressed economy as one in which many workers are without employment for an extended period of time and as a consequence, many see their skills and their morale decay. A depressed economy is also one in which investment is low, the capital stock is growing slowly, if at all, and entrepreneurial exploration is low. These factors may affect future potential output, implying that a temporary shortage of aggregate demand may generate a permanent reduction in aggregate supply. Any policy that may avert such an outcome is therefore worth of being pursued; in particular, a temporary increase in government spending can not only have a large impact on output and help to end the recession, but can also ensure permanent output gains at no financial

[12] Woodford (2011) adds an additional condition, namely that the tax increase required to finance the budget deficit also occurs while the constraint binds.

cost.[13] As stressed by Blanchard and Leigh (2013), hysteresis effects bear upon the transmission of fiscal impulses in general, but are particularly strong during severe downturns.

Besides business cycle conditions, other initial conditions matter, in particular the state of public finances and the level of government debt. Blanchard (1990) proposes a model based on the idea that the size of the fiscal multiplier may be inversely related to the debt-to-gross domestic product (GDP) ratio. When a government consolidates its budget position, it affects expectations and thus consumption in two ways. First, the inter-temporal redistribution of taxes from the future to the present is likely to increase the tax burden of current taxpayers and reduce their consumption. This effect is the conventional one, and its strength depends on how much the economy departs from the benchmark of Ricardian equivalence. Second, by taking measures today, the government eliminates the need for larger, much more disruptive adjustments in the future; in so doing, it removes the danger of low output, increasing consumption as a consequence. Third, consolidation may be associated with a substantial drop in uncertainty, leading to a decrease (i) in precautionary savings and (ii) in the option value of waiting by consumers (to buy durables) and firms (to take investment decisions). The last two mechanisms are unconventional and may justify non-Keynesian effects of tighter fiscal policies. Symmetrically, if an increase in government purchases is perceived as putting under threat sustainability of public finances, it may have very small or even negative effects on output. Sutherland (1997) presents a model that shows how the power of fiscal policy to affect consumption can vary depending on the level of public debt. At moderate levels of debt, fiscal policy has the traditional Keynesian effects: current generations of consumers discount future taxes because they may not be alive at the time of the next debt stabilization policy. But when debt reaches extreme values current generations of consumers know that there is a high probability that they will be alive when the next stabilization programme is implemented. In these situations a fiscal deficit can have a contractionary effect on consumer spending. Nickel and Tudyka (2013) provide empirical evidence on the negative correlation between the size of the fiscal multiplier and the level of public debt. According to the authors, the private sector increasingly displays Ricardian features as the degree of indebtedness rises: for low levels of the debt-to-GDP ratio, consumers and firms neglect the inter-temporal budget constraint of the Government, while for higher levels of borrowing they appear to internalize the tax burden that is invariably associated with an expansion of government spending.

[13]DeLong and Summers (2012) provide an example showing that an incremental 1 dollar of government spending raises permanently future output by 0.015 if (1) the fiscal multiplier is 1.5; (ii) the average income tax rate is 33 percent; (iii) the real interest rate on long-term government debt is fixed at 1 percent.

2.4 Empirical evidence on the size of the fiscal multiplier

2.4.1 Pre-crisis estimates of the multiplier

Until recently, it seemed to be a well established fact that the government-spending multiplier was not substantially larger than one. Authors such as Hall (2009) argue that in the US the multiplier is in the range of 0.7 to 1.0, while authors such as Ramey (2011a) estimate the multiplier to be closer to 1.2.[14] In both studies the estimates are obtained by using structural VAR (SVAR) models, which suffer from difficult-to-solve identification problems.[15] Moreover, studies using aggregate data measure what happens on average when government spending changes: to assess the effect of a deficit-financed stimulus, one needs either to focus on periods in which taxes did not change significantly or to control for tax effects, which is far from easy, as the estimates of tax multipliers range from -0.5 to -5.0.[16]

Similar evidence is obtained with DSGE models: in standard new-Keynesian models the government-spending multiplier can be somewhat above or below one depending on the exact specification of agent's preferences, while in frictionless real-business-cycle models this multiplier is typically less than one.[17] Accordingly, due to limited fire-power, implementation lags and financing costs, fiscal policy was viewed until just a few years ago as a poor tool for aggregate-demand management. Things have changed since the 2007-2008 financial crisis, due also to the perceived powerlessness of monetary policy, stuck at the ZLB.

2.4.2 Recessions, depressions and the ZLB

The evidence on the size of the multiplier when monetary policy is at the ZLB is based on both calibrated DSGE models and more standard (and data-based) econometric techniques. Christiano et al. (2011) use a DSGE model whose parameters match the response of ten US

[14] Leigh et al. (2010) present estimates for 15 developed countries, including the US. They however consider not the standard government-purchases multiplier, but average multipliers, referring to fiscal packages consisting of a mixture of changes in transfers, taxes and purchases of goods and services. They find that a 1 percentage point of GDP fiscal consolidation on average reduces output after 2 years by half a percentage point (and increases the unemployment rate by 0.3).

[15] The critical issue is to distinguish changes in government spending that genuinely represent changes in the fiscal policy stance from those that are caused by economic events. One solution is to focus on military buildups, under the assumption that this type of spending is the least likely to respond to economic events. Nevertheless, as Ramey (2011b) points out, there is always the possibility that the events that lead to these buildups – e.g. the start of World War II and the start of the Cold War – could have other influences on the economy, apart from the effects on government spending, that could bias the estimates of the multiplier. For example, during World War II increased patriotism could have raised labour supply more than would be predicted by economic incentives and hence could have raised the multiplier. In contrast, rationing and capacity constraints during the world wars could have dampened the multiplier.

An additional factor complicating identification is that government spending shocks are most often anticipated, implying that the econometrician does not have all the information that individual agents may have. Thus, expectations of individual agents may not be based just on past information from the variables in the empirical model. Hence, the expectation or forecast errors cannot be the residuals of the model set up by the econometrician and, thus, the shocks of interest may not be forecast errors and may be nonfundamental. See Ramey (2011b) and Perotti (2011).

[16] Ramey (2011b) lists a number of studies tackling this issue.

[17] See e.g. the evidence presented in Cogan et al. (2010) and Coenen et al. (2012).

macro variables to (i) a neutral technology shock; (ii) a monetary impulse; and (iii) a capital-embodied technology shock. Their key findings are the following: first, when the central bank follows a Taylor rule, the value of the government spending multiplier is less than one, in line with most of the literature; second, when the nominal interest rate does not respond to the rise in government spending, the multiplier is much larger;[18] third, the value of the multiplier depends critically on how much government spending occurs in the period during which the nominal interest rate is constant. The evidence provided by Christiano et al. (2011) has been criticized on the grounds that it unduly relies on linearization around the steady-state for a case-study - i.e. the effects of fiscal policy when interest rates are at the ZLB - that is necessarily some distance from the steady-state. According to Braun et al. (2012), this mistake accounts for roughly one half of the estimated size of the fiscal multiplier. Auerbach and Gorodnichencko (2011) use regime-switching models and find large differences in the size of spending multipliers in recessions and expansions: the response in expansions never rises above 1 and soon falls below zero, while the response in recessions rises steadily, reaching a value of over 2.5 after 20 quarters.[19] Some aspects of their analysis are however unconvincing and cast a shadow on the reliability of their results: first, the peak of the GDP response is reached 20 periods after the shock, at the end of the forecast window, when output is apparently gaining further momentum; second, the government shock is still 1 percentage point of GDP higher than in the baseline after 20 periods, suggesting that the shock is permanent rather than transitory; third, the output and tax responses in expansions seem utterly unbelievable: at period 4, with taxes 1.5 percentage point of GDP below and government spending 2 percentage points of GDP above baseline, output is by and large unchanged.

The evidence presented in Ramey (2012) does not support the claim that the multiplier is higher when there is slack in the economy or when interest rates are at the ZLB. She studies the period 1933 to 1951, which is characterized by very low interest rates as well as very high unemployment rates. She estimates on monthly data the following regression:

$$\frac{\Delta Y_t}{Y_{t-1}} = \beta_0 + \beta_1 \frac{\Delta G_t}{Y_{t-1}} + \beta_2 \frac{\Delta Y_{t-1}}{Y_{t-2}} + I_t \left[\beta_3 + \beta_4 \frac{\Delta G_t}{Y_{t-1}} + \beta_5 \frac{\Delta Y_{t-1}}{Y_{t-2}} \right] + \varepsilon_t$$

where Y_t is output, G_t government spending and I_t a dummy variable equal to 1 in periods with high unemployment rates (i.e. larger than 7 percent) and zero otherwise. Unlike Auerbach and Gorodnichencko (2011), she finds that $\beta_4 \simeq 0$. Evidence reported in Ramey (2012) is supported by Owyang et al. (2013), who estimate essentially the same model but use (i) a longer sample period and (ii) a "news" variable (*viz.* the change in the expected present value of government spending in response to military events) rather than G_t: the multiplier is always below unity and, if anything, is slightly lower during the high unemployment state. Owyang et al. (2013)

[18] For example, for a 12-quarter hike in government spending the impact multiplier is roughly 1.6, with a peak value of about 2.3.

[19] It is worth stressing that none of the recessions in the sample (but maybe the last one) qualifies as a depression, in which the policy interest rate is at (or close to) the zero lower-bound.

estimate the same model also on Canadian data, finding this time results that are closer to those of obtained by Auerbach and Gorodnichencko (2011).

More recently, an article in the October 2012 World Economic Outlook of the International Monetary Fund (IMF), written by Blanchard and Leigh, presents evidence that the fiscal multiplier in the advanced economies may be considerably larger than had been assumed when fiscal austerity plans were set in train in most economies in 2010.[20] Using a sample including 28 advanced economies, Blanchard and Leigh regress the forecast error for real GDP growth during 2010-11 on forecasts of fiscal consolidation for 2010-11 that were made in early 2010. Under rational expectations, and assuming that the correct forecast model has been used, the coefficient on planned fiscal consolidation should be zero. Blanchard and Leigh find the coefficient on planned fiscal consolidation to be large, negative, and significant: the baseline estimate suggests that a planned fiscal consolidation of 1 percent of GDP is associated with a growth forecast error of about 1 percentage point (the estimates are in the range of 0.4 to 1.2 percentage points). As the multipliers underlying the growth forecasts made in early 2010 were about 0.5, their results indicate that multipliers have actually been in the 0.9 to 1.7 range. The Blanchard and Leigh study drew a lot of attention and a lot of criticisms. First, the estimates seem to be highly dependent on the inclusion in the sample of Greece and Germany. Second, the results were presented as general, but are limited to the specific time period chosen: the 2010 forecasts of deficits are not good predictors of errors in growth forecasts for 2010 or 2011 when the years are analyzed individually; its 2011 forecasts are not good predictors of anything.[21] Third, the size of the fiscal consolidation efforts assumed by the IMF in early 2010 underestimates the extent of the measures actually implemented. Fourth, the correlation between growth forecast errors and changes in the fiscal stance breaks down when increases in sovereign bond yields are included in the regression.[22] Fifth, the IMF analysis does not distinguish between budget expansions (in place in 2010) and fiscal tightenings (mostly enacted in 2011): usually the former are temporary, while the latter are permanent. The European Commission (2012a) estimates the same regression as the IMF for consolidating countries only and finds no correlation between growth forecast errors and changes in the fiscal stance. Sixth, multipliers differ greatly across countries and take different values depending on the credibility of the consolidation effort and on the response of sovereign risk premia.[23]

Blanchard and Leigh (2013) answer some, but not all, of the critiques raised against their analysis. They claim that their results are extremely robust and in particular do not depend on the inclusion of Germany and Greece in the sample; moreover, they assert that it is no surprise that estimating their model in different periods yields inconsistent results, as economic theory itself predicts that the fiscal multiplier depends on business cycle conditions and on the monetary

[20] The fiscal multiplier in this case does not refer to government purchases. It measures the output response to all the fiscal consolidation measures – on both the revenue and the expenditure side of the public-sector budget – adopted in the countries included in the IMF sample.

[21] On the two points, see Financial Times (2012).

[22] On the third and fourth point, see European Commission (2012a).

[23] See European Central Bank (2012).

policy stance; finally, they posit that sovereign risk premia respond to growth prospects, not to the fiscal stance, and accordingly consolidation measures, by weakening aggregate demand and economic activity, raise the cost of borrowing for governments and increase the multiplier.

2.4.3 Hysteresis

Regarding hysteresis, the evidence is scant at best. Concerning the DeLong and Summers (2012) example, one thing is worth stressing: the size of the hysteresis effects they assume - just 0.015 for each dollar of additional temporary government purchases - seems small, but it is not. In their example, the gains that can be reaped from a fiscal stimulus are permanent and their present value - based on a discount rate that is the same as the real interest rate they use for US long-term bonds - is 1.5, which is larger than the size of the shock itself. Is so large a gain achievable? According to an economist not insensitive to the virtues of expansionary fiscal policies, "massive, unsustainable deficit spending in the hopes that this will somehow generate a self-sustained recovery can be justified only by exotic stories about multiple equilibria, the sort of thing you would imagine only a professor could believe".[24]

2.4.4 Fiscal multipliers in high-debt countries and the sovereign risk channel

The evidence on the relevance of the debt/deficit position of a country on the size of the fiscal multiplier is mostly casual. The sovereign debt crisis has clearly shown that the leeway for governments in setting the stance of fiscal policy is limited: any action that is perceived as jeopardizing debt sustainability triggers immediately a punitive response of financial markets. In particular, for countries with dangerously weak fiscal finances it is to be expected that any attempt to increase public expenditures may spark an upward jump in the risk premium charged on their debt, reducing the output response to the fiscal stimulus, while the contrary is likely to happen for fiscal consolidation attempts.

Two studies by Perotti (1999) and Corsetti et al. (2012b) are however worth mentioning. Perotti (1999) lays out a simple model where government expenditure shocks have a positive, "Keynesian" correlation with private consumption in "normal" times, and a negative, "non-Keynesian" correlation in bad times. Symmetrically, tax shocks have a negative, Keynesian correlation in "normal" times and a positive, non-Keynesian correlation in bad times. What is needed to rationalize state-dependent fiscal multipliers of the type described above is a model in which the correlation between private consumption and shocks to government expenditure and revenues changes, depending on the initial conditions. The empirical model uses a 30-year long panel of 19 country members of the Organization for Economic Cooperation and Development (OECD) and distinguishes "good periods" and "bad periods", defined according to size of the cyclically-adjusted public debt and the probability of re-election of the incumbent government. The empirical evidence supports the claim that expenditure shocks have Keynesian effects at low

[24] Krugman (1999).

levels of debt, and non-Keynesian effects in the opposite circumstances. The evidence of a similar switch in the effects of tax shocks is less strong. Corsetti et al. (2012b) carry out an empirical exploration on a sample of 17 OECD countries (for the period 1975-2008) into the determinants of government spending multipliers, by studying how the fiscal transmission mechanism depends on the economic environment. In terms of conditioning factors, they focus on the exchange rate regime, the level of public debt and the deficit, and the occurrence of a financial crisis. They find that: (1) multipliers are virtually zero under "normal" conditions; (2) the exchange rate regime matters; (3) the fiscal multiplier increases markedly during times of financial crises, being 2.3 on impact and 2.9 at peak; (4) fiscal strains may take the multiplier into negative territory: the cumulative effects over the first 2 years are strongly negative, although the effects become weaker over longer horizons. The usual caveat on cross-country studies with small samples applies to the Corsetti et al. (2012b); moreover, the finding on the impact of financial crises on multipliers may be due to reverse causality, i.e. it may simply reflect the fact that in times of financial crisis, countries experience a large drop in output and government spending; finally, the response to a crisis should be quite different across countries, as larger ones have more fiscal rooms to implement counter-cyclical policies.

While it is clear that there are times and circumstances in which an increase in spending (or a reduction in taxation) may not only boost aggregate demand, but also raise borrowing costs, thus reducing the size of the fiscal multiplier, the evidence on this link is limited. Most empirical studies focus on countries with negligible default risk and postulate linear relationships, as if a country's initial conditions on the stock of outstanding debt were irrelevant. For the United States Laubach (2009) finds that a 1 percentage point increase in the projected deficit-to-GDP (debt-to-GDP) ratio raises long-term yields on Treasury bonds by 20 to 30 (3 to 4) basis point. Gruber and Kamin (2012) obtain similar results for OECD countries, but find no support for the hypothesis that changes in fiscal balances affect yields through their effect on perceived default risk. Attinasi et al. (2010) for the pre-2010 period estimate even lower responses of EA sovereign spreads to anticipated changes in government deficit and debt. A higher elasticity of sovereign risk premia to public finance conditions is found by Belhocine and Dell'Erba (2013), who estimate for 26 emerging countries the response of the yield to maturity of sovereign bonds to changes in the primary balance (to GDP ratio), allowing the response to depend on the level of the outstanding debt. They find that, for countries that have debt levels higher than 45 percent of GDP, a 1 percentage point worsening of the primary balance from its debt-stabilising level translates into a 53.69 basis-point increase in the cost of borrowing.

3 The model setup

The previous section reports the results for fiscal multipliers obtained in the literature. In particular, it is stressed that the size of the multipliers depends upon the stance of monetary policy and on the response of credit spread to changes in public debt and deficit. To further assess the role of these channels, in the results (see section 4) we will show the fiscal multipliers obtained by simulating a DSGE model of the Italian economy. Its main features are illustrated in this section.

The model represents a world economy composed by three regions: Italy, rest of the EA (REA) and rest of the world (RW). In each region there is a continuum of symmetric households and symmetric firms. Italian households are indexed by $j \in [0; s]$, households in the REA by $j^* \in (s; S]$, households in the RW by $j^{**} \in (S; 1]$.[25]

Italy and the REA share the currency and the monetary authority, that sets the nominal interest rate according to EA-wide variables. The presence of the RW outside the EA allows to assess the role of the nominal exchange rate and extra-EA trade in transmitting the shocks. In each region there are households and firms. Households consume a final good, which is a composite of intermediate nontradable and tradable goods. The latter are domestically produced or imported. Households trade a one-period nominal bond, denominated in euro. They also own domestic firms and use another final good (different from the final consumption good) to invest in physical capital. The latter is rented to domestic firms in a perfectly competitive market. All households supply differentiated labor services to domestic firms and act as wage setters in monopolistically competitive labor markets by charging a markup over their marginal rate of substitution between consumption and leisure.

On the production side, there are perfectly competitive firms that produce the two final goods (consumption and investment goods) and monopolistic firms that produce the intermediate goods. The two final goods are sold domestically and are produced combining all available intermediate goods using a constant-elasticity-of-substitution (CES) production function. The two resulting bundles can have different composition. Intermediate tradable and nontradable goods are produced combining domestic capital and labor, that are assumed to be mobile across sectors. Intermediate tradable goods can be sold domestically and abroad. Because intermediate goods are differentiated, firms have market power and restrict output to create excess profits. We also assume that markets for tradable goods are segmented, so that firms can set three different prices, one for each market. Similarly to other DSGE models of the EA (see, among the others, Christoffel et al. 2008 and Gomes et al. 2012), we include adjustment costs on real and nominal variables, ensuring that, in response to a shock, consumption, production and prices react in a gradual way. On the real side, habit preferences and quadratic costs prolong the adjustment of households consumption and investment, respectively. On the nominal side, quadratic costs make wages and prices sticky.[26]

In the following section we describe in detail the fiscal policy setup (the public sector budget constraint and the sovereign spread), the monetary policy setup, and the household's problem for the case of Italy. Similar equations, not reported to save on space, hold for other regions.

[25] The parameter s is the size of the Italian population, which is also equal to the number of firms in each Italian sector (final nontradable, intermediate tradable and intermediate nontradable). Similar assumptions holds for the REA and the RW.

[26] See Rotemberg (1982).

The only exception is the equation of the spread, that holds for Italy only.[27]

3.1 The fiscal authority

We report initially the budget constraint, the fiscal rule of the public sector and, subsequently, the sovereign spread.

3.1.1 Budget constraint and fiscal rule

Fiscal policy is set at the regional level. The government budget constraint is:

$$\left[\frac{B_{t+1}^{g}}{R_t^H} - B_t^g\right] = (1+\tau_t^c)P_{N,t}C_t^g + Tr_t - T_t \tag{1}$$

where $B_t^g \geq 0$ is nominal public debt. It is a one-period nominal bond issued in the EA wide market that pays the gross nominal interest rate R_t^H. The variable C_t^g represents government purchases of goods and services, $Tr_t > 0$ (< 0) are lump-sum transfers (lump-sum taxes) to households. Consistent with the empirical evidence, C_t^g is fully biased towards the intermediate nontradable good. Hence it is multiplied by the corresponding price index $P_{N,t}$.[28]

We assume that the same tax rates apply to every household. Total government revenues T_t from distortionary taxation are given by the following identity:

$$T_t \equiv \int_0^s \left(\tau_t^\ell W_t\left(j\right)L_t\left(j\right) + \tau_t^k\left(R_t^k K_{t-1}\left(j\right) + \frac{\Pi_t^P}{s}\right) + \tau_t^c P_t C_t\left(j\right)\right)dj - \tau_t^c P_{N,t}C_t^g \tag{2}$$

where τ_t^ℓ is the tax rate on individual labor income $W_t\left(j\right)L_t\left(j\right)$, τ_t^k on capital income $R_t^k K_{t-1}\left(j\right) + \Pi_t^P/s$ and τ_t^c on consumption $C_t\left(j\right)$. The variable $W_t\left(j\right)$ represents the individual nominal wage, $L_t\left(j\right)$ is individual amount of hours worked, R_t^k is the rental rate of existing physical capital stock $K_{t-1}\left(j\right)$, Π_t^P stands for dividends from ownership of domestic monopolistic firms (they are equally shared across households) and P_t is the price of the consumption bundle.

The government follows a fiscal rule defined on a single fiscal instrument to bring the public debt as a percent of domestic GDP, $b^g > 0$, in line with its target $\bar{b}^g$ and to limit the increase in public deficit as ratio to GDP (b_t^g/b_{t-1}^g):[29]

$$\frac{i_t}{i_{t-1}} = \left(\frac{b_t^g}{\bar{b}^g}\right)^{\phi_1}\left(\frac{b_t^g}{b_{t-1}^g}\right)^{\phi_2} \tag{4}$$

[27] In the Appendix we lay down the rest of the model.

[28] See Corsetti and Mueller (2006, 2008).

[29] The definition of nominal GDP is:

$$GDP_t = P_t C_t + P_t^I I_t + P_{N,t}C_t^g + P_t^{EXP}EXP_t - P_t^{IMP}IMP_t \tag{3}$$

where P_t, P_t^I, P_t^{EXP}, P_t^{IMP} are prices of consumption, investment, exports and imports, respectively.

where i_t is one of the five fiscal instruments among three tax rates $(\tau_t^\ell, \tau_t^k, \tau_t^c)$ and the two expenditure items (C_t^g, Tr_t). Parameters ϕ_1, ϕ_2 are lower than zero when the rule is defined on an expenditure item calling for a reduction in expenditures whenever the debt level is above target and for a larger reduction whenever the dynamics of the debt is not converging. To the contrary, they are greater than zero when the rule is on tax rates.

3.1.2 Sovereign spread

The interest rate paid by the Italian government and Italian households when borrowing is determined as a spread over the EA risk-free nominal interest rate (which is set by the central bank of the EA). Following Corsetti et al. (2012a), the (gross) spread reflects the risk of sovereign default and is linked to (expected) variations in the fiscal stance as follows:

$$spread_t^H \equiv E_t \left[\left(\frac{b_{t+1}^g}{b_t^g} \right)^{\psi_b} \right] \tag{5}$$

The term on the right-hand side includes (expected) changes in the Italian public debt-to-GDP ratio, where $0 < \psi_b < 1$ is a parameter and $b_{t+1}^g > 0$ is the Italian public debt-to-GDP ratio at the beginning of period $t+1$. As such, the (gross) interest rate R^H paid by the Italian government is:

$$R_t^H \equiv R_t * spread_t^H \tag{6}$$

where R_t is the (gross) risk-free nominal interest rate. The spread also affects the intertemporal choices of the Italian households through the standard Euler equation, as reported later.

3.2 Monetary authority

The monetary authority controls the short-term policy rate R_t according to a Taylor rule of the form:

$$\left(\frac{R_t}{\bar{R}} \right) = \left(\frac{R_{t-1}}{\bar{R}} \right)^{\rho_R} (\Pi_{EA,t})^{(1-\rho_R)\rho_\pi} \left(\frac{GDP_{EA,t}}{GDP_{EA,t-1}} \right)^{(1-\rho_R)\rho_{GDP}} \tag{7}$$

The parameter ρ_R $(0 < \rho_R < 1)$ captures inertia in interest rate setting, while the term $\bar{R}$ represents the steady state gross nominal policy rate. The parameters ρ_π and ρ_{GDP} are respectively the weights of EA CPI inflation rate $(\Pi_{EA,t})$ and GDP $(GDP_{EA,t})$. The CPI inflation rate is a geometric average of CPI inflation rates in Italy and the REA (respectively Π_t and Π_t^*) with weights equal to the correspondent country size (as a share of the EA):

$$\Pi_{EA,t} \equiv (\Pi_t)^{\frac{s}{s+S}} (\Pi_t^*)^{\frac{S}{s+S}} \tag{8}$$

The EA GDP, $GDP_{EA,t}$, is the sum of the Italian and REA GDPs (respectively GDP_t and GDP_t^*):

$$GDP_{EA,t} \equiv GDP_t + rer_t * GDP_t^* \tag{9}$$

where rer_t is the Italian-to-REA bilateral real exchange rate, defined as the ratio of REA to Italian consumer prices. In some simulations, the interest rate will be held constant at its steady state value for several periods, instead of following the Taylor rule (7), that eventually kicks in the future. In this way it is possible to assess the role of the monetary policy stance for the size of fiscal multipliers.

3.3 Households

Households' preferences are additively separable in consumption and labor effort. The generic Italian household j receives utility from consumption C and disutility from labor L. The expected value of the lifetime utility is:

$$E_0 \left\{ \sum_{t=0}^{\infty} \beta^t \left[\frac{(C_t(j) - hC_{t-1})^{1-\sigma}}{(1-\sigma)} - \frac{L_t(j)^{1+\tau}}{1+\tau} \right] \right\} \tag{10}$$

where E_0 denotes the expectation conditional on information set at date 0, β is the discount factor ($0 < \beta < 1$), $1/\sigma$ is the elasticity of intertemporal substitution ($\sigma > 0$) and $1/\tau$ is the labor Frisch elasticity ($\tau > 0$). The parameter h ($0 < h < 1$) represents external habit formation in consumption.

The budget constraint of the household j is:

$$\begin{aligned} \frac{B_t(j)}{\left(1+R_t^H\right)} - B_{t-1}(j) \quad \leq \quad & (1-\tau_t^k)\left(\Pi_t^P(j) + R_t^K K_{t-1}(j)\right) + \\ & +(1-\tau_t^\ell)W_t(j)L_t(j) - (1+\tau_t^c)P_tC_t(j) - P_t^I I_t(j) \\ & +Tr_t(j) - AC_t^W(j) \end{aligned}$$

Italian households hold a one-period bond, B_t, denominated in euro ($B_t > 0$ is a lending position). The short-term nominal rate R_t^H is paid at the beginning of period t and is known at time t.[30] We assume that government and private bonds are traded in the same international market. Households own all domestic firms and there is no international trade in claims on firms' profits. The variable Π_t^P includes profits accruing to the Italian households. The variable I_t is the investment bundle in physical capital and P_t^I the related price index, which is different from the price index of consumption because the two bundles have different composition.[31] Italian households accumulate physical capital K_t and rent it to domestic firms at the nominal rate R_t^k.

[30] A financial friction μ_t is introduced to guarantee that net asset positions follow a stationary process and the economy converge to a steady state. Revenues from financial intermediation are rebated in a lump-sum way to households in the REA. See Benigno (2009).

[31] See the Appendix for more details.

The law of motion of capital accumulation is:

$$K_t(j) = (1-\delta) K_{t-1}(j) + \left(1 - AC_t^I(j)\right) I_t(j) \tag{11}$$

where δ is the depreciation rate. Adjustment cost on investment AC_t^I is:

$$AC_t^I(j) \equiv \frac{\phi_I}{2} \left(\frac{I_t(j)}{I_{t-1}(j)} - 1 \right)^2, \ \phi_I > 0 \tag{12}$$

Finally, Italian households act as wage setters in a monopolistic competitive labor market. Each household j sets her nominal wage taking into account labor demand and adjustment costs AC_t^W on the nominal wage $W_t(j)$:

$$AC_t^W(j) \equiv \frac{\kappa_W}{2} \left(\frac{W_t(j)}{W_{t-1}(j)} - 1 \right)^2 W_t L_t, \ \kappa_W > 0 \tag{13}$$

The costs are proportional to the per-capita wage bill of the overall economy, $W_t L_t$.

The sovereign risk channel (see equation 6) does affect the choices of the Italian households through the interest rate R_H in the Euler equation (obtained by maximizing utility subject to the budget constraint with respect to the bond holdings B_t):

$$(C_t(j) - hC_{t-1})^{-\sigma} = \beta E_t \left(R_t^H (C_{t+1}(j) - hC_t)^{-\sigma} \right) \tag{14}$$

The higher the spread, the higher the interest rate R_t^H and the larger the incentive for Italian households to postpone consumption.

Similar relations hold in the REA and in the RW. The only exceptions are two, as we make two simplifying assumptions. First, the spread paid by Italian households and government are rebated in a lump-sum way to households in the REA. Second, neither the public sectors nor the private sectors in the REA and RW pay the spread when borrowing. So it is the riskless interest rate to appear in the corresponding Euler equations.

Finally, it is assumed that the bond traded by households and governments is in worldwide zero net supply. The implied market clearing condition is:

$$-B_t^g + \int_0^s B_t(j)\, dj - B_t^{g*} + \int_s^S B_t(j^*)\, dj^* - B_t^{g**} + \int_S^1 B_t(j^{**})\, dj^{**} = 0 \tag{15}$$

where B_t^{g*},$B_t^{g**} > 0$ are respectively the amounts of borrowing of the REA and RW public sectors, while $B_t(j^*)$ and $B_t^{**}(j^{**})$ are respectively the per-capita bond positions of households in the REA and in the RW.

3.4 Calibration

The model is calibrated at quarterly frequency. We set some parameter values so that steady-state ratios are consistent with 2010 national account data, which are the most recent and complete available data. For remaining parameters we resort to previous studies and estimates available in the literature.[32]

Table 1 contains parameters that regulate preferences and technology. Parameters with "$*$" and "$**$" are related to the REA and the RW, respectively. Throughout we assume perfect symmetry between the REA and the RW, unless differently specified. We assume that discount rates and elasticities of substitution have the same value across the three regions. The discount factor β is set to 0.9927, so that the steady state real interest rate is equal to 3.0 per cent on an annual basis. The value for the intertemporal elasticity of substitution, $1/\sigma$, is 1. The Frisch labor elasticity is set to 0.5. The depreciation rate of capital δ is set to 0.025. Habit is set to 0.6.

In the production functions of tradables and nontradables, the elasticity of substitution between labor and capital is set to 0.93. The bias towards capital in the production function of tradables is set to 0.56 in Italy and, in the REA and in the RW, to 0.46. The corresponding value in the production function of nontradables is set to 0.53 in Italy and 0.43 in the REA and RW. In the final consumption and investment goods the elasticity of substitution between domestic and imported tradable is set to 1.5, while the elasticity of substitution between tradables and nontradables to 0.5. In the consumption bundle the bias towards the domestic tradeable is 0.68 in Italy, 0.59 in the REA and 0.90 in the RW. The bias towards the composite tradeable is set to 0.68 in Italy, to 0.5 in the REA and the RW. For the investment basket, the bias towards the domestic tradable is 0.50 in Italy, 0.49 in the REA and 0.90 in the RW. The bias towards the composite tradable is 0.78 in Italy, 0.70 in the REA and in the RW. The biases towards the domestically produced good and composite tradable good are chosen to match the Italy and REA import-to-GDP ratios.

Table 2 reports gross markup values. In the Italian tradable and nontradable sectors and in the Italian labor market the markup is set to 1.08, 1.30 and 1.60, respectively (the corresponding elasticities of substitution across varieties are set to 13.32, 4.44 and 2.65). In the REA tradable and nontradable sectors and in the REA labor market the gross markups are respectively set to 1.11, 1.24 and 1.33 (the corresponding elasticities are set to 10.15, 5.19 and 4.00). Similar values are chosen for the corresponding parameters in the RW.

Table 3 contains parameters that regulate the dynamics. Adjustment costs on investment change are set to 6. Nominal wage quadratic adjustment costs are set to 200. In the tradable sector, we set the nominal adjustment cost parameter to 300 for Italian tradable goods sold domestically and in the REA; for Italian goods sold in the RW, the corresponding parameter is set to 50. The same parameterization is adopted for the REA, while for the rest of the world we set the adjustment cost on goods exported to Italy and the REA to 50. Nominal price adjustment

[32] Among others, see Forni et al. (2009, 2010a, 2010b).

costs are set to 500 in the nontradable sector. The parameters are calibrated to generate dynamic adjustments for the EA similar to those obtained with the New Area Wide Model (NAWM, see Christoffel et al. 2008) and Euro Area and Global Economy Model (EAGLE, see Gomes et al. 2010). The two parameters regulating the adjustment cost paid by the private agents on their net financial position are set to 0.00055 so that they do not greatly affect the model dynamics.

Table 4 reports parametrization of the systematic feedback rules followed by the fiscal and monetary authorities. In the fiscal policy rule (4) we set $\phi_1 = \pm 0.05$, $\phi_2 = \pm 1.01$ for Italy and $\phi_1 = \phi_2 = \pm 1.01$ for the REA and the RW. Their sign is positive when the fiscal instrument in the rule is a tax rate, it is negative when the instrument is a public expenditure. The central bank of the EA targets the contemporaneous EA wide consumer price inflation (the corresponding parameter is set to 1.7) and the output growth (the parameter is set to 0.1). Interest rate is set in an inertial way and hence its previous-period value enters the rule with a weight equal to 0.87. Same values hold for the corresponding parameters of the Taylor rule in the RW.

Table 5 reports the actual great ratios and tax rates, which are matched in the model steady state under our baseline calibration. We assume a zero steady state net foreign asset position of each region. This implies that for each region - in steady state - the net financial position of the private sector is equal to the public debt. The size of Italian and REA GDPs, as a share of world GDP, are set to 3 percent and to 17 percent, respectively.

As for fiscal policy variables, the public consumption-to-GDP ratio is set to 0.20. The tax rate on wage income τ^{ℓ} is set to 42.6 per cent in Italy and to 34.6 in the REA. The tax rate on physical capital income τ^{k} is set to 34.9 in Italy and 25.9 in the REA, while the tax rate on consumption τ^{c} is equal to 16.8 in Italy and to 20.3 in the REA. The public debt-to-yearly GDP ratio is calibrated to 119 percent for Italy and to 0.79 for the REA. Variables of the RW are set to values equal to those of corresponding REA variables.

Finally, we have to calibrate for Italy the relationship between the fiscal policy stance and the spread on sovereign debt, as defined in the definition (5). Absent operational estimates of the link between fiscal conditions and risk premia, we resort to our personal reading of the literature on the subject. We refer in particular to Belochine and Dell'Erba (2013) and posit that a 1 percentage point of GDP increase in government spending maps into a 75bp step-up in the sovereign risk premium. The higher sensitivity of borrowing costs with respect to their estimates is justified on the grounds that the Italian debt-to-GDP ratio is much higher than the threshold they find for emerging economies. Moreover, such value is in a way consistent with market developments since mid 2011. In June 2011 the spread between Italian and German 10-year bond yields was about 180 basis points, close to the level reached in the aftermath of the Lehman crisis. During the summer it suddenly started increasing. The intensification of the EA sovereign debt crisis fuelled fears concerning the sustainability of public finances in peripheral countries; the dithering political handling of the crisis failed to prevent market tensions from heightening and by mid November the spread on Italian bonds had reached 553bp, some

370bp more than five months before.[33] It took three fiscal consolidation packages, amounting on aggregate to 4.8 percentage points of GDP to stop the escalation of borrowing costs. By assuming that a budget adjustment of that size is what financial markets expected to keep at just 370bp the re-pricing of Italian sovereign risk, we can gauge in some 75bp the cost (gain) of increasing (reducing) the public-sector deficit by 1 percentage point of GDP.[34] The estimate is admittedly rough, highly tentative and does not distinguish sovereign risk from redenomination risk, but seems nonetheless reasonable and more plausible than the available alternatives.

A number of assumptions are required in order to map observed variations in long-term government bond yields into our model-based quarterly interest rate R_t^H. As a preliminary step, we follow common practice and focus on the return on 10-year government bonds as the most representative long-term market rate. Next, we lay out a procedure to map a given change in the yield on 10-year bonds into variations in R_t^H. We assume for simplicity that changes in the return on a given maturity are equally transmitted to all maturities, so that the shape of the term structure is unchanged. Hence, a higher (lower) return on 10-year government bonds would simply correspond to an upward (downward) shift in the whole yield curve, with no effect on its steepness. The assumption mirrors the implicit definition of the model-based long-term interest rate as a weighted average of expected future short-term rates, via the expectation hypothesis and the Euler equation. An expected change in the short-term rate would equally affect the returns paid at different maturities in our model, so that the shape of the term structure of interest rates would remain unchanged.

4 Results

In what follows we simulate the model to assess the fiscal multipliers for Italy under, alternatively, standard monetary policy, constant monetary policy rate, responses of the credit spread. All simulations are run under perfect foresight. All shocks are fully anticipated by households and firms, with the exception of the shock perturbing the economy in the first (initial) period.

4.1 Benchmark fiscal multipliers

Table 6 shows the short-term (first and second year) results of increasing Italian public consumption by one percent of (pre-stimulus) baseline GDP. For the case of the permanent fiscal shock, the table also reports the long-run multipliers.[35] Monetary policy is conducted according to the

[33] Spikes were observed immediately after (i) the downgrade of Portugal in July; (ii) the release of the plan for the Private Sector Involvement during the EU summit on 21-22 July; (iii) the announcement of the Greek referendum on 1 November. Domestic events, i.e. the tensions generated by the uncertainty on the fiscal consolidation measures and the political vacuum created by the falling apart of the ruling coalition, played a role as well, though a more limited one. For a detailed account of the impact of news on the Italian BTP-German Bund spread between June 2011 and March 2012, see Pericoli (2012).

[34] The decrease in the Italian BTP-Bund spread observed in the initial months of 2012 and since August is not considered in the computation, as it is most likely due to monetary policy.

[35] Long-run multipliers are zero in the case of temporary fiscal shocks.

Taylor rule (7), while public debt is stabilized by rising lump-sum taxes according to the fiscal rule (4).[36] After the end of the stimulus, public spending is immediately brought back to its initial steady state value.

The first two columns of Table 6 report multipliers of the Italian public consumption when the latter is increased for one year. In the first year the Italian GDP increases by 0.87 percent of its baseline value. Italian households' consumption and investment slightly decrease. The nominal policy rate does not increase, because it is set at the EA level and reacts to EA-wide inflation and output. The latter are not greatly affected by the increase in Italian GDP and - to a lesser extent - CPI. Following the small increase in the Italian prices, the real exchange rates of Italy against the REA and the RW slightly appreciate. Similarly, the terms of trade of Italy against the REA and the RW slightly improve. Consistently, tradable goods produced in the REA and in the RW become cheaper than those produced in Italy. The Italian net exports decrease (Italian gross exports and imports decrease and increase, respectively).[37] Spillovers towards the REA and RW are small, because of the relatively small size of Italy in the world economy and the relatively large home bias in the REA and RW consumption and investment baskets.[38]

The remaining columns of Table 6 report multipliers for the first two years in correspondence of two-, five-year and permanent fiscal stimuli (for the last one the long-run multiplier is also reported). In the first year the Italian GDP increases by 0.81, 0.79 and 0.69 percent, respectively; in the second year by 0.68, 0.56 and 0.52. In the case of a permanent fiscal stimulus, the long-run multiplier is 0.59. Responses of the output components quantitatively change across the different scenarios. The longer the duration of the stimulus, the larger the decrease in private consumption and the smaller the decrease in private investment; the latter increases when the stimulus lasts for five years or longer. Differences across the responses of households' demand are associated with the strength of the negative wealth effect of current and expected future public spending. The larger the amount of resources appropriated for public consumption, the larger the negative wealth effect, the more Italian households reduce consumption and increase labor supply. The increase in the latter makes capital more productive and induces higher investment and capital accumulation. Accordingly the aggregate supply can match the persistently higher public demand for consumption.

For comparison, Table 7 reports the values of the public consumption multipliers when both public spending and labor income taxes are simultaneously increased. The increase in the labor tax rate is such that the corresponding revenues are equal to one percent of pre-stimulus GDP, so that the fiscal stimulus is ex ante revenue-neutral. The multiplier is now lower than in the case of higher lump-sum taxes. There is a lower incentive to increase labor effort than in the

[36] The implications of distortionary taxation for the spending multiplier will be considered later in this section.

[37] This is true for (bilateral) exports and imports to and from REA and RW (to save on space we do not report them). Exports decrease more towards the RW as their prices increase more than those of the exports towards the REA (the former are more flexible than the latter).

[38] The REA and RW consumption and investment (not reported) slightly decrease to finance the increase in Italian borrowing associated with the fiscal stimulus and the consumption smoothing of Italian households.

previous case, as the increase in distortionary labor taxes reduces the after-tax real wage. The differences are large for the second year, in particular for long-lasting stimuli.

The previous simulations have shown the multipliers associated with public consumption spending. Table 8 reports the multipliers associated with stimuli based on reducing tax rates on labor income, capital income and consumption. The reduction in tax revenues is equal to one percent of (pre-stimulus) GDP and lasts for, alternatively, one, two, five years or is implemented on a permanent basis. After the stimulus, the public debt is stabilized by increasing lump-sum taxes according to the fiscal rule (4); public consumption is held constant at its pre-stimulus level. Results show that in the short run tax multipliers are lower than one and lower than public consumption multipliers. Moreover, they are larger in the second year than in the first year (the only exception is consumption tax), because households' consumption and investment react in a smooth way, given the assumptions of habit persistence and adjustment costs on investment. Finally, in the case of labor and capital income taxes, the longer the duration of the stimulus, the larger the multipliers, because households have a larger incentive to increase labor effort when the lower taxation on their (labor and capital) income is closer to being permanent. In particular, in the long run the GDP multiplier associated with a permanent reduction in the capital tax rate is larger than one.

Figure 1 shows the dynamic response of the main macroeconomic variables in the benchmark case of a public consumption increase financed with lump-sum taxes. Figure 2 reports responses to the labor tax reduction. In both cases the stimulus lasts for one year. It is interesting to note that the increase in public consumption immediately raises GDP, while the reduction in labor tax boosts GDP only gradually, as consumption and investment increase smoothly because of external habit formation in consumption and adjustment costs for investment.

Overall, results suggest that fiscal multipliers are below 1 and that multipliers associated with taxation are lower than those associated with public spending for short-lived shocks, but larger (in the long-run) when fiscal expansion is permanent.

4.2 Constant monetary policy rate

Previous simulations show fiscal multipliers under the assumption that monetary policy follows the Taylor rule (7). Table 9 reports results for increases in public consumption (by one percent of GDP) for one, two, five years and on a permanent basis when the (nominal) policy rate stays constant ("accommodative" monetary policy stance) during the fiscal stimulus; for the permanent stimulus, the accommodative stance lasts for five years. After the stimulus, monetary policy is conducted in a standard way (the Taylor rule 7 kicks in). As in previous simulations the public debt is stabilized by increasing lump-sum taxes according to rule (4).

In the case of one- and two-year stimuli the GDP multiplier is similar to the one obtained under the standard stance of monetary policy (Table 6). The multiplier increases well above one when the stimulus lasts for 5 years. It is equal to 1.42 in the first year (1.16 in the second year).

When the fiscal stimulus is permanent and the monetary policy is accommodative for five years, the lack of full overlap between the monetary and fiscal policy implies that the multiplier is only slightly larger than in the case of standard monetary policy (0.80 and 0.63 in the first two years vs. 0.69 and 0.52). The results are qualitatively in line with those reported in Woodford (2011), who finds that both fiscal stimulus and accommodative monetary policy have to be held in place for a sufficiently long amount of time to generate large multipliers, as inflation expectations need to be large enough to reduce the ex ante real interest rate. Note also that the 5-year mix of expansionary Italian fiscal policy and constant EA policy rate positively affects the REA activity and inflation, through trade leakages. When the interest rate is held constant for sufficiently long and there is full overlap with the fiscal stimulus, inflation expectations of REA households become high enough to widely reduce the ex ante real interest rate, stimulating households' demand for consumption and investment. The latter favors Italian exports, by partially counterbalancing the loss of competitiveness associated with the appreciation of the Italian real exchange rate.

Table 10 reports results under the assumption that the policy rate remains constant for a number of periods half as large as the number of periods of the fiscal stimulus (2.5 years in the case of the 5-year and permanent fiscal stimuli). Multipliers are lower than those reported in Table 9, as now the monetary policy accommodates to a lower extent the public consumption shock. If the latter lasts for 5 years or is permanent, the corresponding multipliers are equal to 0.89 (instead of 1.42) and 0.72 (instead of 0.80) in the first year, respectively; to 0.66 (instead of 1.16) and 0.55 (instead of 0.63) in the second year.

Table 11 reports results for tax-rate multipliers. For 1- and 2-year stimuli the tax multipliers under the assumption of constant interest rate are similar to those in the case of standard monetary policy (Table 8). For 5-year and permanent stimuli the tax multipliers are larger under no monetary-policy response than in the case of standard monetary policy. In particular, the multiplier associated with the capital income tax becomes larger than one. To the opposite, the multiplier associated with the labor income tax decreases in the case of 2- and 5-year stimuli: the reason is the large initial positive response of the supply side of the economy, which reduces inflation expectations and, given the absence of monetary policy response, increases the real interest rate. Consumption and investment fall accordingly, as monetary policy is not accommodative anymore.

Overall, the multiplier associated with public consumption is well above one only when the stance of the monetary policy is accommodative for sufficiently long; otherwise the value of the multipliers do not change much relatively to the case of standard monetary policy response and remains generally below 1.

4.3 Sovereign risk premium

The macroeconomic effects of a fiscal stimulus can be affected not only by the monetary policy stance, but also by the response of financial markets. As stressed in the survey (Section 2), if investors are concerned about the solvency of the government, they will ask for a higher premium in response to a fiscal expansion. Moreover, the sovereign risk premium will be transmitted rather quickly to the cost of borrowing of domestic households and firms, thereby crowding out their spending decisions (so called "sovereign risk channel of fiscal policy", see Corsetti et al. 2012a). As a consequence, the operation of the sovereign risk channel can reduce the size of the fiscal multiplier in times of financial turbulence. This conjecture is supported by the empirical evidence reported in recent studies. Laubach (2012) studies the dependence of the sovereign spread on the current level of fiscal indicators (such as the surplus-to-GDP or the debt-to-GDP ratios) for a panel of EA countries and finds that the estimated elasticity is small or nil in non-crisis periods, but increases rapidly and dramatically in times of financial stress.

This section reports results of simulating a 1 percent (of pre-stimulus GDP) increase in public consumption spending for, alternatively, 1, 2, 5 years and on a permanent basis. We assume that the fiscal expansion entails an immediate 75 basis points rise in the sovereign risk premium and hence in interest rate paid on Italian government bonds (see section 3.4). The effects of the stimulus crucially depend on the dynamics of the sovereign risk premium after the initial increase. We adopt an agnostic approach and assume that, after the initial rise, the Italian sovereign spread linearly declines and returns to its baseline level by the time the stimulus is withdrawn.[39] Consistent with the empirical evidence for Italy (see Albertazzi et al. (2012) and Zoli (2013)), the increase in sovereign risk is fully passed-through (in a quarter) to the borrowing rate paid by the Italian private sector. Monetary policy follows the standard Taylor rule and public debt is stabilized by lump-sum taxes after the end of the fiscal stimulus. The output multipliers are reported in Table 12. In the first year they are equal to 0.79, 0.62, 0.27 and 0.18 when, respectively, the stimulus lasts for 1, 2, 5 years or is permanent; in the second year, they fall to -0.11, 0.58, 0.07 and 0.02 . The values are lower compared to the scenarios that do not include the sovereign channel (Table 6), because of the larger crowding-out effect on private-sector spending, associated with the increase in the borrowing rates. Moreover, the longer the stimulus, the slower the spread decrease, the larger the reduction in the multiplier. In the case of the 5-year stimulus, private consumption decreases by 1.19 percent in the first year (1.28 in the second), private investment by 1.51 percent (2.36 percent). Absent the sovereign risk channel (see Table 6), private consumption would decrease by 0.15 percent in the first year (0.33 in the second), while private investment would increase.

Figure 3 sums up the results reported in this and the previous section. It shows that the size of government-consumption multiplier strongly depends on the monetary policy response and on the change of the sovereign risk premium; in particular, it can be larger than one only if the monetary policy rate is held constant for a rather long amount of time. Moreover, monetary policy should remain accommodative for as long as the duration of the fiscal stimulus (compare the bars "accommodative monetary policy"and "partial overlap"). Otherwise, multipliers are

[39] In the case of a permanent stimulus, we assume it takes 5 years for the sovereign risk premium to return to the baseline value.

lower - and possibly much lower - than one, if the sovereign risk premium increases.

4.4 Fiscal consolidation and sovereign risk

The results reported in the previous sections may be interpreted as suggesting that, in times of financial stress, fiscal consolidation may reduce borrowing costs for households and firms. If the consolidation is credible, financial markets might ask for a lower sovereign risk premium, as investors anticipate that public finances have become fully sustainable. With a quick and complete pass-through of the sovereign premium to the private sector borrowing rate, the lower borrowing cost of households and firms might partially counterbalance the contractionary effects of the consolidation.

This section analyzes the output effects of fiscal consolidation in the presence of a sovereign risk channel. As in the case of Italy in late 2011, the policy tightening may be induced by an abrupt increase in the sovereign spread, as a result of financial market turbulence. Our results about fiscal consolidation should be ideally compared to a benchmark scenario where the surge in sovereign risk premia is not accompanied by any fiscal plan.

Table 13 reports results for the case of permanently reducing the public debt-to-GDP ratio by 1 percentage point. In line with the composition of the fiscal package adopted in Italy in the second half of 2011, public spending is permanently reduced by 0.25 percentage points of pre-stimulus GDP, while taxation (on labor income, capital income and consumption) is increased by 0.75 percentage points.[40] The response of the sovereign spread is designed as follows: the spread decreases on impact by 75 basis points - in line with the reduction observed in Italy after the announcement of the fiscal consolidation plan in the fall of 2011 - then starts increasing and returns to its baseline value after, alternatively, 1, 2, 3 or 5 years. We simulate both the case of a standard monetary policy response and the case of the ZLB binding for five years.[41] Results suggest that the largest output decrease is equal to -0.70 (-0.80) in the first (second) year. It is obtained when the monetary policy is constrained by the ZLB and there is no sovereign risk channel. The smallest output decrease is equal to -0.03 (-0.15) in the first (second) year. It is obtained when the decrease in the risk premium lasts for a relatively large period of time (5 years). In this case, households benefit from a relatively low real interest rate, partially counterbalancing the increase in distortionary taxation.[42] Note that, as a limiting case, the effect on output can be positive in the first year if the spread returns to its baseline value in three and five years.

Simulation results suggest that under conditions of financial stress, when the sovereign risk channel is active, the negative impact of fiscal consolidation can be quite modest and is certainly

[40] See Ministero dell'Economia e delle Finanze (2012).

[41] Note that we do not posit an exogenous recessionary shock that takes the monetary policy rate down to the ZLB, as opposed to common practice in the literature, (see for example Corsetti et al. 2012a). The reason is that the ZLB holds at EA level and, hence, it can be taken as exogenous with respect to changes in the Italian economic conditions.

[42] Note that public spending decreases. As such, it contributes to crowding in households' spending for consumption and investment.

lower than under "normal" conditions.

4.5 Sensitivity analysis

This section reports the sensitivity analysis on the multipliers associated with higher public consumption when the sovereign risk premium augments by 0.75 percentage points on impact (first quarter) and then gradually returns to zero in an amount of time equal to the length of the fiscal stimulus. Differently from the benchmark results illustrated in section 4.3 we assume that the share of liquidity constrained households is 30 percent of the Italian population and, alternatively, that the increase in public consumption is implemented simultaneously in Italy and the REA, under standard or accommodative monetary policy.

4.5.1 Share of liquidity constrained households

Table 14 shows results when in Italy the share of liquidity constrained households is increased from zero to 30 percent of the overall Italian population.

Following Campbell and Mankiw (1989) and Galí et al. (2004, 2007), we assume that in each period liquidity constrained households consume their after-tax disposable income. That is, the budget constraint of the generic liquidity-constrained household j is:

$$(1+\tau_t^c)P_tC_t\left(j\right) = (1-\tau_t^\ell)W_t\left(j\right)L_t\left(j\right)$$

We assume liquidity-constrained households wage and hours are the same as those of unconstrained households. We also assume that tax rates on labor income and consumption are the same for both constrained and unconstrained households.

Multipliers are larger than in the case of no liquidity constrained households. The higher values are due to the income effect associated with the liquidity constrained households, who immediately increase their consumption, as they do not save but spend all their available wage income. The latter increases because firms expand employment, to help production meet the higher aggregate demand. Differences with respect to the benchmark scenarios are not extremely large. More importantly, the multipliers continue to be below one.

4.5.2 Simultaneous fiscal stimulus in the EA

We also assess to which extent Italian fiscal multipliers do change when the stimulus is implemented simultaneously in Italy and the REA. We consider the case of an increase in public consumption by 1 percent of pre-shock GDP for two years. The monetary policy is conducted according to the Taylor rule or is accommodative (policy rate held at its baseline level during the fiscal stimulus).

Table 14 reports the results. Under standard monetary policy, they are slightly smaller in the case of EA-wide stimulus than in the case of Italian unilateral stimulus. As in the case of unilateral Italian stimulus, the multiplier is lower than one. Italian net exports (not reported) now decrease less, because the Italian bilateral exchange rate against the REA appreciates to a lower extent. The monetary policy rate now increases to a larger extent, given the increase in the EA-wide aggregate demand. As such, the real interest rate decreases to a lower extent when the fiscal stimulus is coordinated, contributing to crowding-out relative more Italian demand of households and firms.

Italian multipliers are larger than one when the monetary policy is accommodative for two years. The constant interest rate stimulates REA aggregate demand, because of the reduction in the REA real interest rate. Italian (gross) exports decrease less, driven by the higher aggregate demand in the REA. The Italian GDP multiplier is equal in the first year to 1.34 percent. It is 0.88 percent in the case of unilateral Italian stimulus and accommodative monetary policy (Table 9). This suggests that the accommodative monetary policy is more effective in driving the multiplier above one when the fiscal stimulus is simultaneously implemented at the EA level.

5 Conclusions

This paper provides estimates for Italy of the size of fiscal multipliers under alternative assumptions about the reaction of the central bank and the response of the sovereign risk premium. Our main conclusions are the following. First, short-run fiscal multipliers are typically below one and, in particular, multipliers associated with taxation are lower than those associated with public expenditure. Second, public spending multipliers are substantially larger than one when the monetary policy rate is kept constant at the ZLB, but only if the policy rate remains at the ZLB for a sufficiently long period of time (at least five years in our simulations). Third, under conditions similar to those currently prevailing in the euro area (EA), in countries with a high public debt, the stimulus induces a deterioration of public finances and hence a rapid increase in the sovereign risk premium, which in turn substantially reduces the size of the multiplier and the effectiveness of fiscal policy. Fourth, the short-run contractionary effects of fiscal consolidation efforts can be partially mitigated by a reduction in the risk premium. Overall, results suggest that the size of fiscal multipliers in times of financial distress can be different from that in "normal" times, as the initial public finance conditions and the monetary policy stance can greatly matter for the financing conditions of the private sector.

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Table 1. Parametrization of Italy, the rest of the Euro Area and the rest of the world

Parameter	IT	REA	RW
Discount rate β	0.993	0.993	0.993
Intertemporal elasticity of substitution $1/\sigma$	1.0	1.0	1.0
Inverse of Frisch elasticity of labor supply τ	2.0	2.0	2.0
Habit h	0.6	0.6	0.6
Depreciation rate of (private and public) capital δ	0.025	0.025	0.025
Tradable Intermediate Goods			
Substitution between factors of production $\xi_T, \xi_T^*, \xi_T^{**}$	0.93	0.93	0.93
Bias towards capital $\alpha_T, \alpha_T^*, \alpha_T^{**}$	0.56	0.46	0.46
Nontradable Intermediate Goods			
Substitution between factors of production $\xi_N, \xi_N^*, \xi_N^{**}$	0.93	0.93	0.93
Bias towards capital $\alpha_N, \alpha_N^*, \alpha_N^{**}$	0.53	0.43	0.43
Final consumption goods			
Substitution between domestic and imported goods $\phi_A, \phi_A^*, \phi_A^{**}$	1.50	1.50	1.50
Bias towards domestic tradable goods a_H, a_F^*, a_G^*	0.68	0.59	0.90
Substitution between domestic tradables and nontradables $\rho_A, \rho_A^*, \rho_A^{**}$	0.50	0.50	0.50
Bias towards tradable goods a_T, a_T^*, a_T^{**}	0.68	0.50	0.50
Final investment goods			
Substitution between domestic and imported goods $\phi_E, \phi_E^*, \phi_E^{**}$	1.50	1.50	1.50
Bias towards domestic tradable goods υ_H, υ_F^*	0.50	0.49	0.90
Substitution between domestic tradables and nontradables ρ_E, ρ_E^*	0.50	0.50	0.50
Bias towards tradable goods υ_T, υ_T^*	0.78	0.70	0.70

Note: IT=Italy; REA=rest of the euro area; RW=rest of the world.

Table 2. Gross Markups

	Markups and Elasticities of Substitution		
	Tradables	Nontradables	Wages
IT	1.08 ($\theta_T = 13.32$)	1.30 ($\theta_N = 4.44$)	1.60 ($\psi = 2.65$)
REA	1.11 ($\theta_T^* = 10.15$)	1.24 ($\theta_N^* = 5.19$)	1.33 ($\psi^* = 4$)
RW	1.11 ($\theta_T^{**} = 10.15$)	1.24 ($\theta_N^{**} = 5.19$)	1.33 ($\psi^{**} = 4$)

Note: IT=Italy; REA=rest of the euro area; RW=rest of the world; source: OECD (2012).

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Table 3. Real and Nominal Adjustment Costs

Parameter	IT	REA	RW
Real Adjustment Costs			
Investment ϕ_I, ϕ_I^{*}, ϕ_I^{**}	6.00	6.00	6.00
Households' financial net position ϕ_{b1}, ϕ_{b2}	0.00055, 0.00055	-	0.00055, 0.00055
Nominal Adjustment Costs			
Wages κ_W, κ_W^{*}, κ_W^{**}	200	200	200
Italian produced tradables κ_H, k_H^{*} k_H^{**}	300	300	50
REA produced tradables κ_H, k_H^{*} k_H^{**}	300	300	50
RW produced tradables κ_H, k_H^{*} k_H^{**}	50	50	300
Nontradables κ_N, κ_N^{*}, κ_N^{**}	500	500	500

Note: IT=Italy; REA=rest of the euro area; RW=rest of the world.

Table 4. Fiscal and Monetary Policy Rules

Parameter	IT	REA	EA	RW
Fiscal policy rule				
$\phi_1, \phi_1^{*}, \phi_1^{**}$	±0.05	±1.01	-	±1.01
$\phi_2, \phi_2^{*}, \phi_2^{**}$	±1.01	±1.01	-	±1.01
Common monetary policy rule	-	-		
Lagged interest rate at t-1 ρ_R, ρ_R^{**}	-	-	0.87	0.87
Inflation ρ_Π, ρ_Π^{**}	-	-	1.70	1.70
GDP growth $\rho_{GDP}, \rho_{GDP}^{**}$	-	-	0.10	0.10

Note: IT=Italy; REA=rest of the euro area; EA=euro area; RW=rest of the world.

Table 5. Main macroeconomic variables (ratio to GDP) and tax rates

	IT	REA	RW
Macroeconomic variables			
Private consumption	61.0	57.1	64.0
Private Investment	18.0	16.0	20.0
Imports	29.0	24.3	4.3
Net Foreign Asset Position	0.0	0.0	0.0
GDP (share of world GDP)	0.03	0.17	0.80
Public expenditures			
Public purchases	20.0	20.0	20.0
Interests	4.0	2.0	2.0
Public investment	2.0	3.0	3.0
Debt (ratio to annual GDP)	119	79	79
Tax Rates			
on wage	42.6	34.6	34.6
on rental rate of capital	34.9	25.9	25.9
on price of consumption	16.8	20.3	20.3

Note: IT=Italy; REA=rest of the euro area; RW=rest of the world. Sources: European Commission (2012b); tax rates (in percent) are from Eurostat (2012).

Table 6. Public consumption multipliers

	1 year-stimulus		2 year-stimulus		5 year-stimulus		permanent stimulus		
	1st year	2nd year	1st year	2nd year	1st year	2nd year	1st year	2nd year	LR
Italian variables									
GDP	0.87	-0.09	0.81	0.68	0.79	0.56	0.69	0.52	0.59
Consumption	-0.04	-0.07	-0.08	-0.18	-0.15	-0.33	-0.53	-0.83	-0.82
Investment	-0.05	-0.13	-0.03	-0.25	0.34	0.31	0.56	1.01	0.55
Export (volumes)	-0.42	-0.18	-0.56	-0.74	-0.63	-0.99	-0.48	-0.68	-0.28
Import (volumes)	0.09	-0.01	0.13	0.09	0.26	0.33	0.03	0.11	-0.15
Terms of Tr. REA (+=deterior.)	-0.13	-0.11	-0.22	-0.36	-0.27	-0.54	-0.20	-0.37	-0.19
Terms of Tr. RW (+=deterior.)	-0.35	-0.12	-0.45	-0.56	-0.48	-0.70	-0.36	-0.49	-0.19
Real Exc. Rate REA (+=depr.)	-0.06	-0.05	-0.09	-0.16	-0.13	-0.28	-0.09	-0.19	-0.14
Real Exc. Rate RW (+=depr.)	-0.06	-0.05	-0.10	-0.17	-0.15	-0.30	-0.12	-0.21	-0.14
Inflation(annualized)	0.08	-0.04	0.15	0.02	0.20	0.11	0.14	0.06	0.00
Real.Int.Rate (annualized)	-0.03	0.04	-0.12	0.04	-0.18	-0.07	-0.12	-0.04	0.00
Nominal Int. Rate (annualized)	0.01	0.00	0.01	0.02	0.01	0.02	0.01	0.00	0.00
Labor	1.36	-0.16	1.24	1.01	1.19	0.75	1.03	0.66	0.47
Pub.Def.(%gdp)	0.74	-0.10	0.76	0.86	0.78	0.92	0.93	0.93	-0.39
Prim.Pub.Def.(%gdp)	0.75	-0.13	0.78	0.84	0.80	0.91	0.95	0.91	-0.41
REA GDP	0.00	0.00	0.00	0.00	-0.01	-0.02	-0.02	-0.01	0.00
RW GDP	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

Note: LR=long run, REA=rest of the euro area; RW=rest of the world. All variables as % dev. from steady state, inflation and interest rate as % point dev. from steady state.

Table 7. Public consumption multipliers. Labor tax-based financing

	1 year-stimulus		2 year-stimulus		5 year-stimulus		permanent stimulus		
	1st year	2nd year	1st year	2nd year	1st year	2nd year	1st year	2nd year	LR
Italian variables									
GDP	0.84	-0.14	0.75	0.55	0.67	0.28	0.50	0.15	-0.33
Consumption	-0.05	-0.09	-0.11	-0.24	-0.22	-0.50	-0.85	-1.33	-1.77
Investment	-0.08	-0.18	-0.09	-0.42	0.23	-0.07	0.45	0.67	-0.14
Export (volumes)	-0.48	-0.26	-0.69	-0.98	-0.83	-1.47	-0.58	-0.98	-1.32
Import (volumes)	0.09	-0.01	0.13	0.09	0.26	0.30	-0.16	-0.18	-0.43
Terms of Tr. REA (+=deterior.)	-0.16	-0.15	-0.27	-0.48	-0.38	-0.82	-0.26	-0.57	-0.89
Terms of Tr. RW (+=deterior.)	-0.40	-0.18	-0.54	-0.73	-0.62	-1.04	-0.43	-0.69	-0.88
Real Exc. Rate REA (+=depr.)	-0.07	-0.07	-0.12	-0.23	-0.18	-0.43	-0.13	-0.32	-0.65
Real Exc. Rate RW (+=depr.)	-0.07	-0.07	-0.13	-0.23	-0.22	-0.46	-0.16	-0.33	-0.65
Inflation(annualized)	0.10	-0.03	0.20	0.04	0.30	0.20	0.20	0.13	0.00
Real.Int.Rate (annualized)	-0.05	0.04	-0.17	0.03	-0.29	-0.15	-0.20	-0.12	0.00
Nominal Int. Rate (annualized)	0.01	0.00	0.01	0.02	0.01	0.01	0.00	0.00	0.00
Labor	1.31	-0.24	1.13	0.78	0.98	0.24	0.69	0.02	-0.68
Pub.Def.(%gdp)	-0.29	-0.12	-0.27	-0.20	-0.25	-0.12	0.03	0.16	-0.10
Prim.Pub.Def.(%gdp)	-0.27	-0.11	-0.24	-0.18	-0.22	-0.09	0.06	0.18	-0.11
REA GDP	0.00	0.00	0.00	0.00	-0.02	-0.03	-0.02	-0.02	-0.01
RW GDP	0.00	0.00	0.00	0.00	0.00	-0.01	0.00	0.00	0.00

Note: LR=long run, REA=rest of the euro area; RW=rest of the world. All variables as % dev. from steady state, inflation and interest rate as % point dev. from steady state.

Table 8. Tax multipliers. Italian GDP and inflation

	labor tax			capital tax			consumption tax		
	1st year	2nd year	LR	1st year	2nd year	LR	1st year	2nd year	LR
GDP									
1 year-stimulus	0.03	0.04	0.00	0.02	0.02	0.00	0.34	0.07	0.00
2 year-stimulus	0.06	0.13	0.00	0.08	0.11	0.00	0.30	0.36	0.00
5 year-stimulus	0.11	0.29	0.00	0.23	0.47	0.00	0.28	0.30	0.00
permanent stimulus	0.19	0.37	0.90	0.15	0.52	2.49	0.08	0.15	0.36
Inflation									
1 year-stimulus	-0.02	0.00	0.00	0.00	0.00	0.00	0.06	-0.01	0.00
2 year-stimulus	-0.04	-0.02	0.00	0.00	-0.01	0.00	0.09	0.03	0.00
5 year-stimulus	-0.09	-0.08	0.00	0.04	-0.03	0.00	0.11	0.07	0.00
permanent stimulus	-0.06	-0.07	0.00	-0.01	-0.06	0.00	-0.03	-0.03	0.00

Note: LR=long run. GDP as % dev. from steady state, inflation as annualized % point dev. from steady state.

Table 9. Public consumption multipliers. Constant monetary policy rate

	1 year-stimulus		2 year-stimulus		5 year-stimulus		permanent stimulus		
	1st year	2nd year	1st year	2nd year	1st year	2nd year	1st year	2nd year	LR
Italian variables									
GDP	0.89	-0.08	0.88	0.74	1.42	1.16	0.80	0.63	0.59
Consumption	-0.02	-0.05	-0.01	-0.11	0.52	0.24	-0.42	-0.73	-0.82
Investment	-0.02	-0.09	0.09	-0.09	1.47	1.84	0.77	1.28	0.54
Export (volumes)	-0.41	-0.17	-0.51	-0.71	-0.15	-0.71	-0.39	-0.63	-0.28
Import (volumes)	0.11	0.01	0.20	0.16	0.87	1.00	0.13	0.23	-0.15
Terms of Tr. REA (+=deterior.)	-0.14	-0.11	-0.22	-0.36	-0.29	-0.56	-0.20	-0.38	-0.19
Terms of Tr. RW (+=deterior.)	-0.36	-0.12	-0.48	-0.56	-0.78	-0.70	-0.41	-0.49	-0.19
Real Exc. Rate REA (+=depr.)	-0.06	-0.05	-0.09	-0.16	-0.13	-0.28	-0.09	-0.19	-0.14
Real Exc. Rate RW (+=depr.)	-0.03	-0.04	0.01	-0.11	0.79	0.18	0.05	-0.12	-0.14
Inflation(annualized)	0.10	-0.03	0.22	0.05	0.84	0.45	0.25	0.12	0.00
Real.Int.Rate (annualized)	-0.06	0.03	-0.19	0.00	-0.80	-0.36	-0.23	-0.09	0.00
Nominal Int. Rate (annualized)	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Labor	1.39	-0.14	1.36	1.11	2.29	1.68	1.22	0.83	0.48
Pub.Def.(%gdp)	0.71	-0.12	0.72	0.80	0.44	0.55	0.91	0.92	-0.40
Prim.Pub.Def.(%gdp)	0.74	-0.14	0.75	0.80	0.50	0.62	0.93	0.92	-0.43
REA GDP	0.02	0.01	0.07	0.06	0.61	0.55	0.10	0.09	0.00
RW GDP	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

Note: LR=long run, REA=rest of the euro area; RW=rest of the world. All variables as % dev. from steady state, inflation and interest rate as % point dev. from steady state.

Table 10. Public consumption multipliers. Partial monetary policy accommodation

	1 year-stimulus		2 year-stimulus		5 year-stimulus		permanent stimulus		
	1st year	2nd year	1st year	2nd year	1st year	2nd year	1st year	2nd year	LR
Italian variables									
GDP	0.88	-0.09	0.82	0.69	0.89	0.66	0.72	0.55	0.59
Consumption	-0.03	-0.06	-0.06	-0.16	-0.04	-0.23	-0.50	-0.81	-0.82
Investment	-0.04	-0.11	0.00	-0.22	0.53	0.56	0.62	1.08	0.55
Export (volumes)	-0.41	-0.17	-0.55	-0.73	-0.55	-0.94	-0.46	-0.67	-0.28
Import (volumes)	0.10	0.00	0.15	0.11	0.36	0.44	0.05	0.14	-0.15
Terms of Tr. REA (+=deterior.)	-0.14	-0.11	-0.22	-0.36	-0.27	-0.55	-0.20	-0.37	-0.19
Terms of Tr. RW (+=deterior.)	-0.36	-0.12	-0.46	-0.56	-0.53	-0.70	-0.37	-0.49	-0.19
Real Exc. Rate REA (+=depr.)	-0.06	-0.05	-0.09	-0.16	-0.13	-0.28	-0.09	-0.19	-0.14
Real Exc. Rate RW (+=depr.)	-0.04	-0.05	-0.07	-0.16	0.01	-0.22	-0.07	-0.19	-0.14
Inflation(annualized)	0.09	-0.03	0.17	0.03	0.30	0.16	0.16	0.08	0.00
Real.Int.Rate (annualized)	-0.05	0.03	-0.15	0.03	-0.29	-0.13	-0.15	-0.06	0.00
Nominal Int. Rate (annualized)	0.00	0.00	0.00	0.01	0.00	0.00	0.00	0.00	0.00
Labor	1.38	-0.15	1.27	1.03	1.37	0.91	1.08	0.70	0.47
Pub.Def.(%gdp)	0.72	-0.11	0.74	0.84	0.71	0.85	0.92	0.92	-0.39
Prim.Pub.Def.(%gdp)	0.74	-0.13	0.77	0.83	0.75	0.86	0.94	0.92	-0.42
REA GDP	0.01	0.01	0.02	0.01	0.09	0.08	0.01	0.01	0.00
RW GDP	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

Note: LR=long run, REA=rest of the euro area; RW=rest of the world. All variables as % dev. from steady state, inflation and interest rate as % point dev. from steady state.

Table 11. Tax multipliers. Constant monetary policy rate. Italian GDP and inflation

	labor tax			capital tax			consumption tax		
	1st year	2nd year	LR	1st year	2nd year	LR	1st year	2nd year	LR
GDP									
1 year-stimulus	0.03	0.04	0.00	0.02	0.02	0.00	0.35	0.08	0.00
2 year-stimulus	0.05	0.12	0.00	0.09	0.12	0.00	0.36	0.42	-0.03
5 year-stimulus	0.00	0.17	0.00	0.47	0.70	0.00	0.82	0.82	0.00
permanent stimulus	0.39	0.56	0.90	1.47	1.77	2.50	0.16	0.23	0.36
Inflation									
1 year-stimulus	-0.02	0.00	0.00	0.00	0.00	0.00	0.07	-0.01	0.00
2 year-stimulus	-0.05	-0.03	0.00	0.02	-0.01	0.00	0.15	0.06	0.00
5 year-stimulus	-0.22	-0.15	0.00	0.27	0.09	0.00	0.66	0.37	0.00
permanent stimulus	0.14	0.04	0.00	1.34	0.66	0.00	0.06	0.02	0.00

Note: LR=long run. GDP as % dev. from steady state, inflation as annualized % point dev. from steady state.

Table 12. Public consumption multipliers. Spread increase

	1 year-stimulus		2 year-stimulus		5 year-stimulus		permanent stimulus		
	1st year	2nd year	1st year	2nd year	1st year	2nd year	1st year	2nd year	LR
Italian variables									
GDP	0.79	-0.11	0.62	0.58	0.27	0.07	0.18	0.02	0.59
Consumption	-0.22	-0.10	-0.48	-0.37	-1.19	-1.28	-1.57	-1.78	-0.75
Investment	-0.20	-0.21	-0.51	-0.67	-1.51	-2.36	-1.29	-1.65	0.59
Export (volumes)	-0.37	-0.15	-0.43	-0.63	-0.15	-0.33	0.00	-0.02	-0.35
Import (volumes)	-0.08	-0.06	-0.32	-0.20	-1.15	-1.37	-1.38	-1.58	-0.05
Terms of Tr. REA (+=deterior.)	-0.12	-0.10	-0.17	-0.31	-0.09	-0.25	-0.02	-0.08	-0.24
Terms of Tr. RW (+=deterior.)	-0.31	-0.10	-0.34	-0.47	-0.10	-0.21	0.01	0.00	-0.24
Real Exc. Rate REA (+=depr.)	-0.05	-0.05	-0.08	-0.14	-0.05	-0.15	-0.02	-0.07	-0.17
Real Exc. Rate RW (+=depr.)	-0.05	-0.05	-0.08	-0.14	-0.06	-0.16	-0.03	-0.07	-0.18
Inflation(annualized)	0.07	-0.03	0.12	0.03	0.07	0.09	0.00	0.05	0.00
Real.Int.Rate (annualized)	-0.03	0.03	-0.10	0.02	-0.07	-0.12	-0.01	-0.08	0.00
Nominal Int. Rate (annualized)	0.01	0.00	0.01	0.01	0.00	-0.01	-0.01	-0.02	0.00
Labor	1.23	-0.18	0.93	0.89	0.36	0.13	0.20	0.04	0.45
Pub.Def.(%gdp)	1.29	-0.04	1.47	1.32	1.74	1.91	1.70	1.54	-0.34
Prim.Pub.Def.(%gdp)	0.80	-0.14	0.89	0.88	1.09	1.15	1.05	0.80	-0.33
REA GDP	0.00	0.00	-0.01	-0.01	-0.03	-0.03	-0.03	-0.03	0.00
RW GDP	0.00	0.00	0.00	0.00	-0.01	-0.01	-0.01	0.00	0.00

Note: LR=long run, REA=rest of the euro area; RW=rest of the world. All variables as % dev. from steady state, inflation and interest rate as % point dev. from steady state.

Table 13. Fiscal consolidation and spread reduction. Italian GDP

	standard monetary policy		5 year-ZLB	
	1st year	2nd year	1st year	2nd year
No spread	-0.28	-0.39	-0.70	-0.80
Spread: -75 bp on impact, 0 bp after 1 year	-0.20	-0.38	-0.62	-0.78
Spread: -75 bp on impact, 0 bp after 2 years	-0.09	-0.30	-0.51	-0.70
Spread: -75 bp on impact, 0 bp after 3 years	0.02	-0.17	-0.38	-0.56
Spread: -75 bp on impact, 0 bp after 5 years	0.23	0.10	-0.03	-0.15

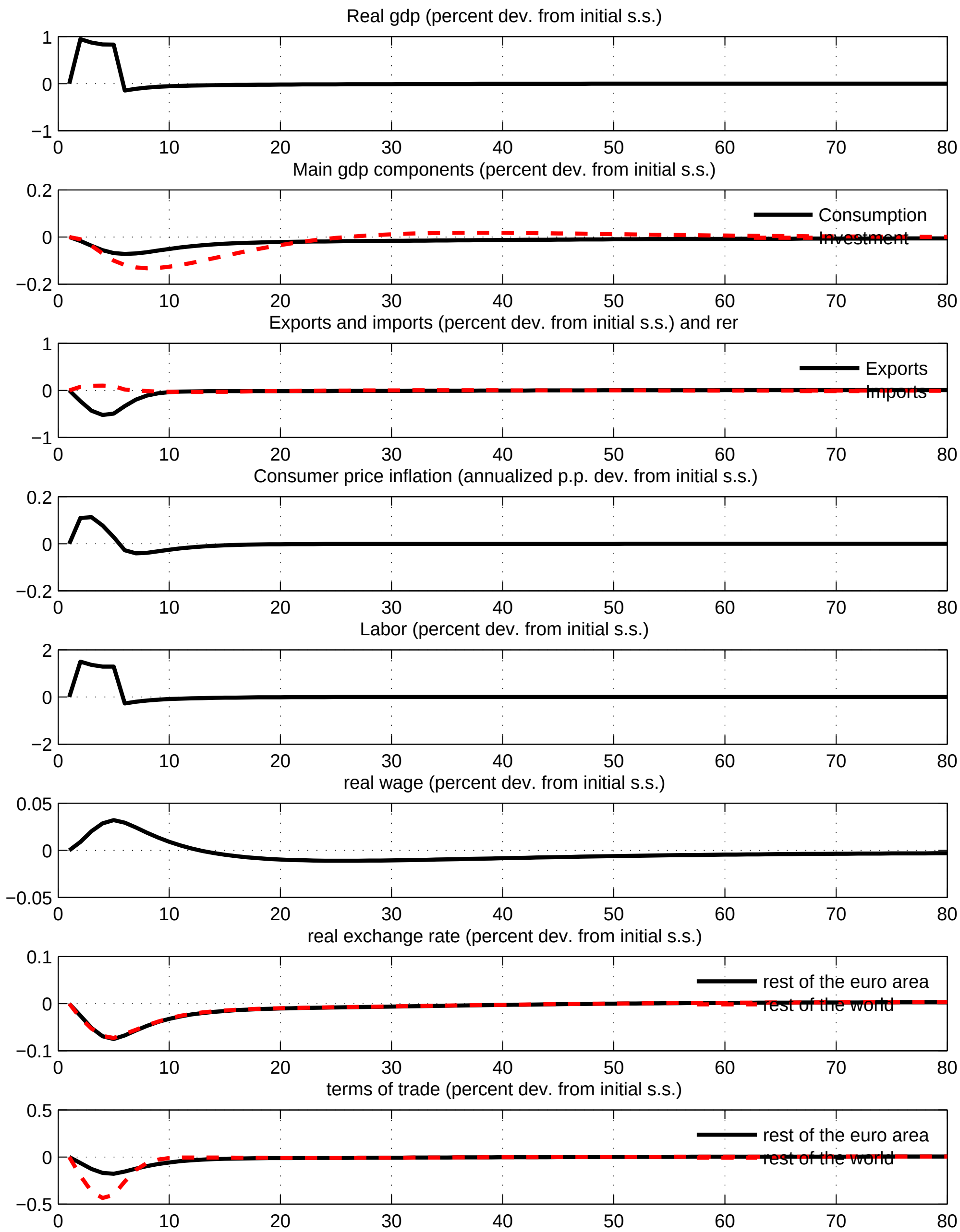
Note: GDP as % dev. from steady state.

Table 14. Sensitivity on public consumption multipliers. Italian GDP and inflation

	benchmark			ROT households			fiscal coord.			fiscal coord.+ZLB		
	1st year	2nd year	LR	1st year	2nd year	LR	1st year	2nd year	LR	1st year	2nd year	LR
GDP												
1 year-stimulus	0.87	-0.09	0.00	0.92	-0.14	0.00	0.87	-0.09	0.00	1.01	0.01	0.00
2 year-stimulus	0.81	0.68	0.00	0.73	0.66	0.00	0.76	0.64	0.00	1.34	1.16	0.00
5 year-stimulus	0.79	0.56	0.00	0.39	0.13	0.00	0.61	0.36	0.00	7.35	6.50	0.00
permanent stimulus	0.69	0.52	0.59	0.26	0.07	0.68	0.53	0.37	0.56	1.61	1.38	0.58
Inflation												
1 year-stimulus	0.08	-0.04	0.00	0.08	-0.04	0.00	0.12	-0.02	0.00	0.12	-0.02	0.00
2 year-stimulus	0.15	0.02	0.00	0.14	0.03	0.00	0.20	0.07	0.00	0.20	0.07	0.00
5 year-stimulus	0.20	0.11	0.00	0.09	0.11	0.00	0.17	0.17	0.00	0.17	0.17	0.00
permanent stimulus	0.14	0.06	0.00	0.00	0.04	0.00	0.04	0.03	0.00	0.04	0.03	0.00

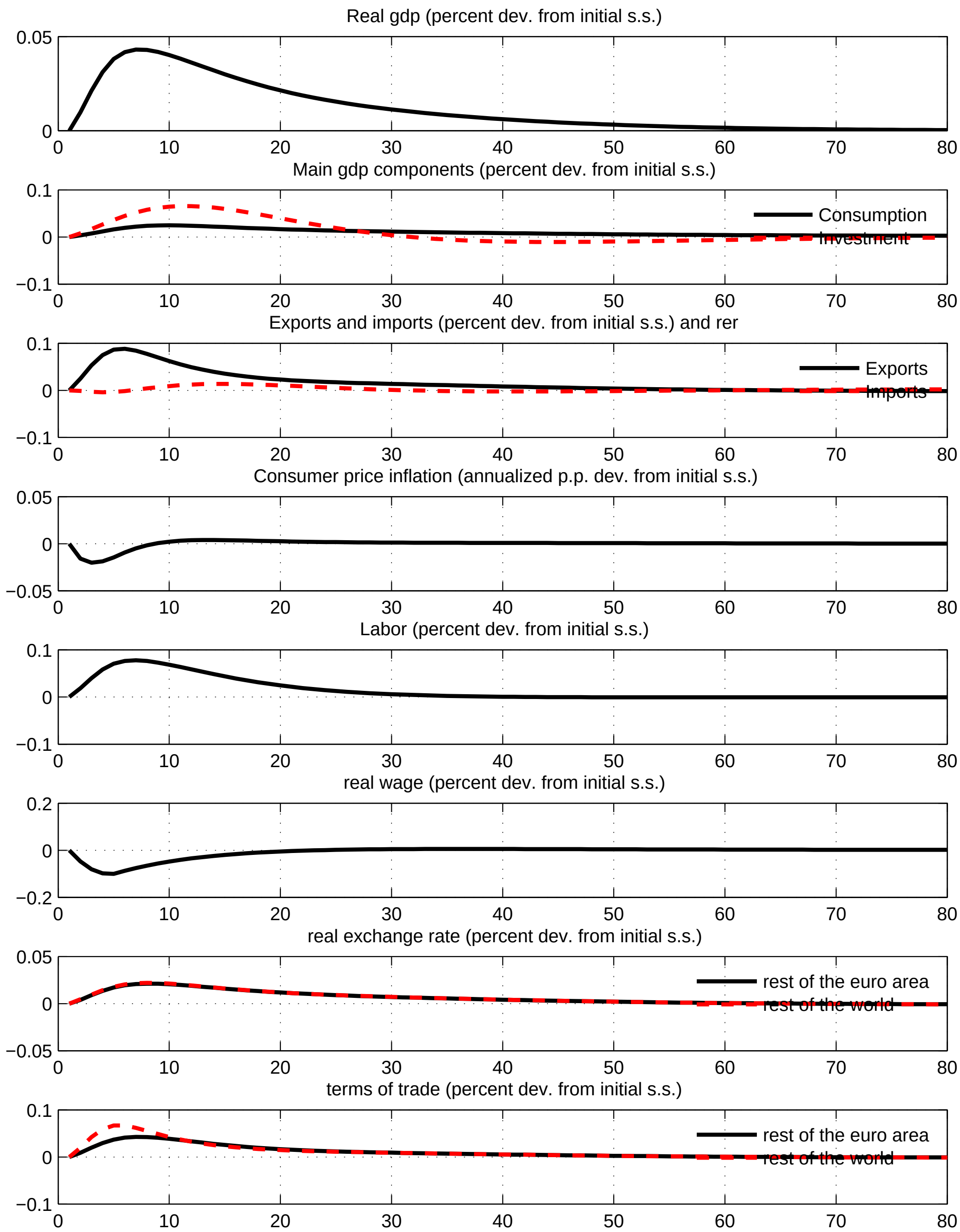
Note: LR=long run; GDP as % dev. from steady state, inflation as annualized % point dev. from steady state.

Figure 1. Italian public consumption shock



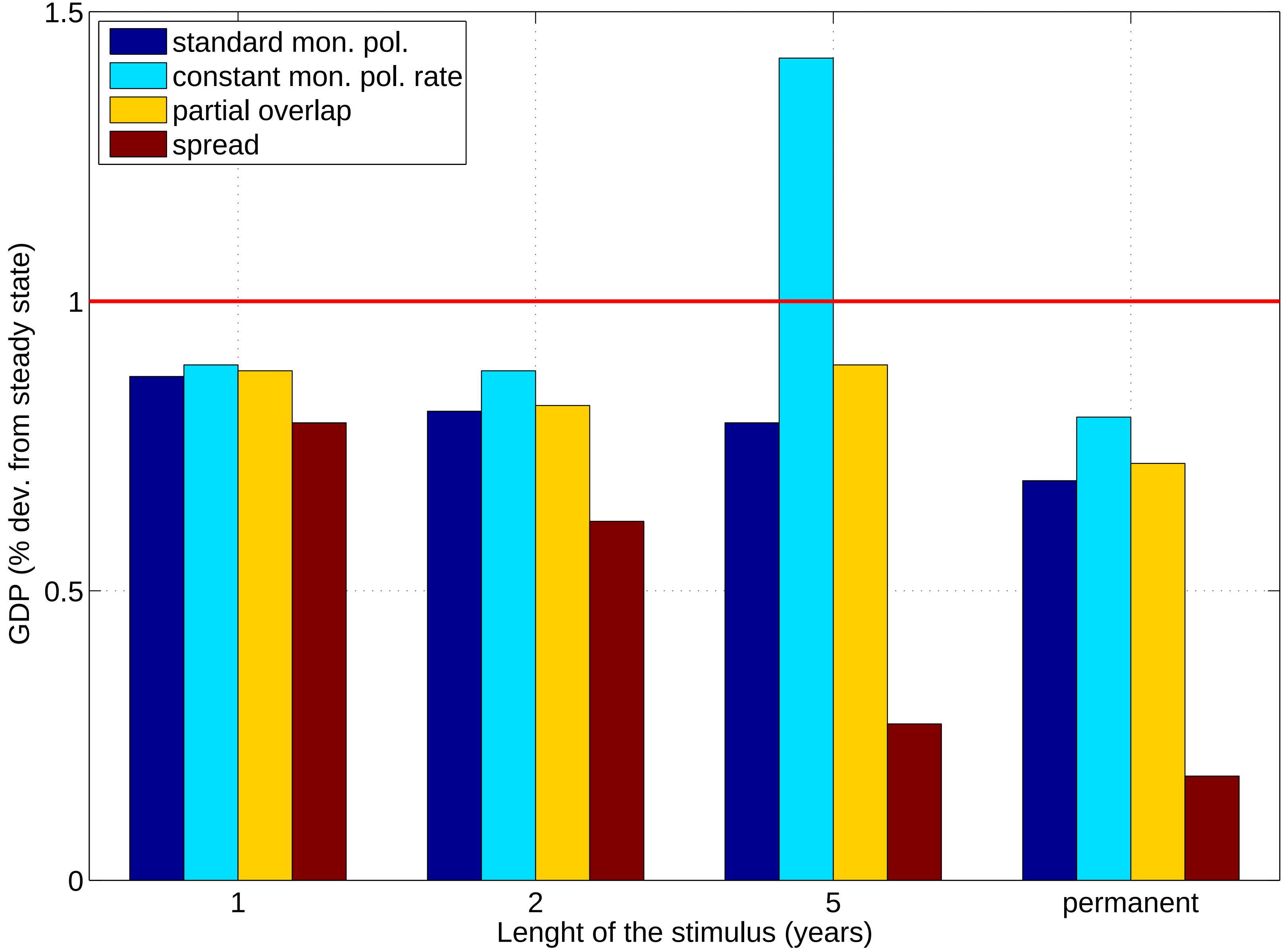
Note: 1-year increase in public consumption of 1% of (pre-shock) Italian GDP. Horizontal axis: quarters.

Figure 2. Italian labor tax shock



Note: 1-year reduction in labor taxation of 1% of (pre-shock) Italian GDP. Horizontal axis: quarters.

Figure 3. First year−GDP multiplier associated with public consumption



Appendix

In this Appendix we report a detailed description of the model, excluding the fiscal and monetary policy part and the description of the households optimization problem that are reported in the main text.[43]

There are three countries, Italy, the rest of the euro area (REA) and the rest of the world (RW). They have different sizes. Italy and the REA share the currency and the monetary authority. In each region there are households and firms. Each household consumes a final composite good made of nontradable, domestic tradable and imported intermediate goods. Households have access to financial markets and smooth consumption by trading a risk-free one-period nominal bond, denominated in euro. They also own domestic firms and capital stock, which is rent to domestic firms in a perfectly competitive market. Households supply differentiated labor services to domestic firms and act as wage setters in monopolistically competitive markets by charging a markup over their marginal rate of substitution.

On the production side, there are perfectly competitive firms that produce the final goods and monopolistic firms that produce the intermediate goods. Two final goods (private consumption and private investment) are produced combining all available intermediate goods according to constant-elasticity-of-substitution bundle. The public consumption good is a bundle of intermediate nontradable goods.

Tradable and nontradable intermediate goods are produced combining capital and labor in the same way. Tradable intermediate goods can be sold domestically or abroad. Because intermediate goods are differentiated, firms have market power and restrict output to create excess profits. We assume that goods markets are internationally segmented and the law of one price for tradables does not hold. Hence, each firm producing a tradable good sets three prices, one for the domestic market and the other two for the export market (one for each region). Since the firm faces the same marginal costs regardless of the scale of production in each market, the different price-setting problems are independent of each other.

To capture the empirical persistence of the aggregate data and generate realistic dynamics, we include adjustment costs on real and nominal variables, ensuring that, in response to a shock, consumption and production react in a gradual way. On the real side, quadratic costs and habit prolong the adjustment of the investment and consumption. On the nominal side, quadratic costs make wage and prices sticky.

In what follows we illustrate the Italian economy. The structure of each of the other two regions (REA and the RW) is similar and to save on space we do not report it.

[43]For a detailed description of the main features of the model see also Bayoumi (2004) and Pesenti (2008).

5.1 Final consumption and investment goods

There is a continuum of symmetric Italian firms producing final nontradable consumption under perfect competition. Each firm producing the consumption good is indexed by $x \in (0, s]$, where the parameter $0 < s < 1$ measures the size of Italy. Firms in the REA and in the RW are indexed by $x^* \in (s, S]$ and $x^{**} \in (S, 1]$, respectively (the size of the world economy is normalized to 1). The CES production technology used by the generic firm x is:

$$A_t(x) \equiv \left(\begin{array}{c} a_T^{\frac{1}{\phi_A}} \left(a_H^{\frac{1}{\rho_A}} Q_{HA,t}(x)^{\frac{\rho_A - 1}{\rho_A}} + a_G^{\frac{1}{\rho_A}} Q_{GA,t}(x)^{\frac{\rho_A - 1}{\rho_A}} (1 - a_H - a_G)^{\frac{1}{\rho_A}} Q_{FA,t}(x)^{\frac{\rho_A - 1}{\rho_A}} \right)^{\frac{\rho_A}{\rho_A - 1} \frac{\phi_A - 1}{\phi_A}} \\ + (1 - a_T)^{\frac{1}{\phi_A}} Q_{NA,t}(x)^{\frac{\phi_A - 1}{\phi_A}} \end{array} \right)^{\frac{\phi_A}{\phi_A - 1}}$$

where Q_{HA}, Q_{GA}, Q_{FA} and Q_{NA} are bundles of respectively intermediate tradables produced in Italy, intermediate tradables produced in the REA, intermediate tradables produced in the RW and intermediate nontradables produced in Italy. The parameter $\rho_A > 0$ is the elasticity of substitution between tradables and $\phi_A > 0$ is the elasticity of substitution between tradable and nontradable goods. The parameter a_H $(0 < a_H < 1)$ is the weight of the Italian tradable, the parameter a_G $(0 < a_G < 1)$ the weight of tradables imported from the REA, a_T $(0 < a_T < 1)$ the weight of tradable goods.

The production of investment good is similar. There are symmetric Italian firms under perfect competition indexed by $y \in (0, s]$. Firms in the REA and in the RW are indexed by $y^* \in (s, S]$ and $y^{**} \in (S, 1]$. Output of the generic Italian firm y is:

$$E_t(y) \equiv \left(\begin{array}{c} v_T^{\frac{1}{\phi_E}} \left(v_H^{\frac{1}{\rho_E}} Q_{HE,t}(y)^{\frac{\rho_E - 1}{\rho_E}} + v_G^{\frac{1}{\rho_E}} Q_{GE,t}(y)^{\frac{\rho_E - 1}{\rho_E}} + (1 - v_H - v_G)^{\frac{1}{\rho_E}} Q_{FE,t}(y)^{\frac{\rho_E - 1}{\rho_E}} \right)^{\frac{\rho_E}{\rho_E - 1} \frac{\phi_E - 1}{\phi_E}} \\ + (1 - v_T)^{\frac{1}{\phi_E}} Q_{NE,t}(y)^{\frac{\phi_E - 1}{\phi_E}} \end{array} \right)^{\frac{\phi_E}{\phi_E - 1}}$$

Finally, we assume that public expenditure C^g is composed by intermediate nontradable goods only.

5.2 Intermediate goods

5.2.1 Demand

Bundles used to produce the final consumption goods are CES indexes of differentiated intermediate goods, each produced by a single firm under conditions of monopolistic competition:

$$Q_{HA}(x) \equiv \left[\left(\frac{1}{s}\right)^{\theta_T} \int_0^s Q(h,x)^{\frac{\theta_T-1}{\theta_T}} dh\right]^{\frac{\theta_T}{\theta_T-1}} \tag{16}$$

$$Q_{GA}(x) \equiv \left[\left(\frac{1}{S-s}\right)^{\theta_T} \int_s^S Q(g,x)^{\frac{\theta_T-1}{\theta_T}} dg\right]^{\frac{\theta_T}{\theta_T-1}} \tag{17}$$

$$Q_{FA}(x) \equiv \left[\left(\frac{1}{1-S}\right)^{\theta_T} \int_S^1 Q(f,x)^{\frac{\theta_T-1}{\theta_T}} df\right]^{\frac{\theta_T}{\theta_T-1}} \tag{18}$$

$$Q_{NA}(x) \equiv \left[\left(\frac{1}{s}\right)^{\theta_N} \int_0^s Q(n,x)^{\frac{\theta_N-1}{\theta_N}} dn\right]^{\frac{\theta_N}{\theta_T-1}} \tag{19}$$

where firms in the Italian intermediate tradable and nontradable sectors are respectively indexed by $h \in (0,s)$ and $n \in (0,s)$, firms in the REA by $g \in (s,S]$ and firms in the RW by $f \in (S,1]$. Parameters θ_T, $\theta_N > 1$ are respectively the elasticity of substitution across brands in the tradable and nontradable sector. The prices of the intermediate nontradable goods are denoted $p(n)$. Each firm x takes these prices as given when minimizing production costs of the final good. The resulting demand for intermediate nontradable input n is:

$$Q_{A,t}(n,x) = \left(\frac{1}{s}\right)\left(\frac{P_t(n)}{P_{N,t}}\right)^{-\theta_N} Q_{NA,t}(x) \tag{20}$$

where $P_{N,t}$ is the cost-minimizing price of one basket of local intermediates:

$$P_{N,t} = \left[\int_0^s P_t(n)^{1-\theta_N} dn\right]^{\frac{1}{1-\theta_N}} \tag{21}$$

We can derive $Q_A(h,x)$, $Q_A(f,x)$, $C_A^g(h,x)$, $C_A^g(f,x)$, P_H and P_F in a similar way. Firms y producing the final investment goods have similar demand curves. Aggregating over x and y, it can be shown that total demand for intermediate nontradable good n is:

$$\begin{aligned} & \int_0^s Q_{A,t}(n,x)\,dx + \int_0^s Q_{E,t}(n,y)\,dy + \int_0^s C_t^g(n,x)\,dx \\ = & \left(\frac{P_t(n)}{P_{N,t}}\right)^{-\theta_N} \left(Q_{NA,t} + Q_{NE,t} + C_{N,t}^g\right) \end{aligned}$$

where C_N^g is public sector consumption. Italy demands for (intermediate) domestic and imported tradable goods can be derived in a similar way.

5.2.2 Supply

The supply of each Italian intermediate nontradable good n is denoted by $N^S(n)$:

$$N_t^S(n) = \left((1-\alpha_N)^{\frac{1}{\xi_N}} L_{N,t}(n)^{\frac{\xi_N-1}{\xi_N}} + \alpha^{\frac{1}{\xi_N}} K_{N,t}(n)^{\frac{\xi_N-1}{\xi_N}} \right)^{\frac{\xi_N}{\xi_N-1}} \tag{22}$$

Firm n uses labor $L_{N,t}^p(n)$ and capital $K_{N,t}(n)$ with constant elasticity of input substitution $\xi_N > 0$ and capital weight $0 < \alpha_N < 1$. Firms producing intermediate goods take the prices of labor inputs and capital as given. Denoting W_t the nominal wage index and R_t^K the nominal rental price of capital, cost minimization implies:

$$L_{N,t}(n) = (1-\alpha_N) \left(\frac{W_t}{MC_{N,t}(n)} \right)^{-\xi_N} N_t^S(n) \tag{23}$$

$$K_{N,t}(n) = \alpha \left(\frac{R_t^K}{MC_{N,t}(n)} \right)^{-\xi_N} N_t^S(n)$$

where $MC_{N,t}(n)$ is the nominal marginal cost:

$$MC_{N,t}(n) = \left((1-\alpha) W_t^{1-\xi_N} + \alpha \left(R_t^K \right)^{1-\xi_N} \right)^{\frac{1}{1-\xi_N}} \tag{24}$$

The productions of each Italian tradable good, $T^S(h)$, is similarly characterized.

5.2.3 Price setting in the intermediate sector

Consider now profit maximization in the Italian intermediate nontradable sector. Each firm n sets the price $p_t(n)$ by maximizing the present discounted value of profits subject to the demand constraint and the quadratic adjustment costs:

$$AC_{N,t}^p(n) \equiv \frac{\kappa_N^p}{2} \left(\frac{P_t(n)}{P_{t-1}(n)} - 1 \right)^2 Q_{N,t} \quad \kappa_N^p \geq 0$$

paid in unit of sectorial product $Q_{N,t}$ and where κ_N^p measures the degree of price stickiness. The resulting first-order condition, expressed in terms of domestic consumption, is:

$$p_t(n) = \frac{\theta_N}{\theta_N - 1} mc_t(n) - \frac{A_t(n)}{\theta_N - 1} \tag{25}$$

where $mc_t(n)$ is the real marginal cost and $A(n)$ contains terms related to the presence of price adjustment costs:

$$\begin{aligned} A_t(n) &\approx \kappa_N^p \frac{P_t(n)}{P_{t-1}(n)} \left(\frac{P_t(n)}{P_{t-1}(n)} - 1 \right) \\ &\quad -\beta \kappa_N^p \frac{P_{t+1}(n)}{P_t(n)} \left(\frac{P_{t+1}(n)}{P_t(n)} - 1 \right) \frac{Q_{N,t+1}}{Q_{N,t}} \end{aligned}$$

The above equations clarify the link between imperfect competition and nominal rigidities. As emphasized by Bayoumi et al. (2004), when the elasticity of substitution θ_N is very large and hence the competition in the sector is high, prices closely follow marginal costs, even though adjustment costs are large. To the contrary, it may be optimal to maintain stable prices and accommodate changes in demand through supply adjustments when the average markup over marginal costs is relatively high. If prices were flexible, optimal pricing would collapse to the standard pricing rule of constant markup over marginal costs (expressed in units of domestic consumption):

$$p_t(n) = \frac{\theta_N}{\theta_N - 1} mc_{N,t}(n) \tag{26}$$

Firms operating in the intermediate tradable sector solve a similar problem. We assume that there is market segmentation. Hence the firm producing the brand h chooses $p_t(h)$ in the Italian market,a price $p_t^*(h)$ in the REA and a price $p_t^{**}(h)$ in the RW to maximize the expected flow of profits (in terms of domestic consumption units):

$$E_t \sum_{\tau=t}^{\infty} \Lambda_{t,\tau} \left[\begin{array}{c} p_\tau(h) y_\tau(h) + p_\tau^*(h) y_\tau^*(h) + p_\tau^{**}(h) y_\tau^{**}(h) \\ -mc_{H,\tau}(h) \left(y_\tau(h) + y_\tau^*(h) + y_\tau^{**}(h) \right) \end{array} \right]$$

subject to quadratic price adjustment costs similar to those considered for nontradables and standard demand constraints. The term E_t denotes the expectation operator conditional on the information set at time t, $\Lambda_{t,\tau}$ is the appropriate discount rate and $mc_{H,t}(h)$ is the real marginal cost. The first order conditions with respect to $p_t(h)$, $p_t^*(h)$ and $p_t^{**}(h)$ are:

$$\begin{aligned} p_t(h) &= \frac{\theta_T}{\theta_T - 1} mc_t(h) - \frac{A_t(h)}{\theta_T - 1} && (27) \\ p_t^*(h) &= \frac{\theta_T}{\theta_T - 1} mc_t(h) - \frac{A_t^*(h)}{\theta_T - 1} && (28) \\ p_t^{**}(h) &= \frac{\theta_T}{\theta_T - 1} mc_t(h) - \frac{A_t^{**}(h)}{\theta_T - 1} && (29) \end{aligned}$$

where θ_T is the elasticity of substitution of intermediate tradable goods, while $A(h)$ and $A^*(h)$ involve terms related to the presence of price adjustment costs:

$$
\begin{aligned}
A_t(h) &\approx \kappa_H^p \frac{P_t(h)}{P_{t-1}(h)}\left(\frac{P_t(h)}{P_{t-1}(h)}-1\right) \\
&\quad -\beta\kappa_H^p \frac{P_{t+1}(h)}{P_t(h)}\left(\frac{P_{t+1}(h)}{P_t(h)}-1\right)\frac{Q_{H,t+1}}{Q_{H,t}} \\
A_t^*(h) &\approx \theta_T - 1 + \kappa_H^p \frac{P_t^*(h)}{P_{t-1}^*(h)}\left(\frac{P_t^*(h)}{P_{t-1}^*(h)}-1\right) \\
&\quad -\beta\kappa_H^p \frac{P_{t+1}^*(h)}{P_t^*(h)}\left(\frac{P_{t+1}^*(h)}{P_t^*(h)}-1\right)\frac{Q_{H,t+1}^*}{Q_{H,t}^*} \\
A_t^{**}(h) &\approx \theta_T - 1 + \kappa_H^p \frac{P_t^{**}(h)}{P_{t-1}^{**}(h)}\left(\frac{P_t^{**}(h)}{P_{t-1}^{**}(h)}-1\right) \\
&\quad -\beta\kappa_H^p \frac{P_{t+1}^{**}(h)}{P_t^{**}(h)}\left(\frac{P_{t+1}^{**}(h)}{P_t^{**}(h)}-1\right)\frac{Q_{H,t+1}^{**}}{Q_{H,t}^{**}}
\end{aligned}
$$

where κ_H^p,${\kappa_H^p}^*$,${\kappa_H^p}^{**} > 0$ respectively measure the degree of nominal rigidity in Italy, in the REA and in the RW. If nominal rigidities in the (domestic) export market are highly relevant (that is, if is relatively large), the degree of inertia of Italian goods prices in the foreign markets will be high. If prices were flexible ($\kappa_H^p = \kappa_H^{p*} = \kappa_H^{p**} = 0$) then optimal price setting would be consistent with the cross-border law of one price (prices of the same tradable goods would be equal when denominated in the same currency).

5.3 Labor Market

In the case of firms in the intermediate nontradable sector, the labor input $L_N(n)$ is a CES combination of differentiated labor inputs supplied by domestic agents and defined over a continuum of mass equal to the country size ($j \in [0, s]$):

$$
L_{N,t}(n) \equiv \left(\frac{1}{s}\right)^{\frac{1}{\psi}} \left[\int_0^s L_t(n,j)^{\frac{\psi-1}{\psi}} dj\right]^{\frac{\psi}{\psi-1}} \tag{30}
$$

where $L(n,j)$ is the demand of the labor input of type j by the producer of good n and $\psi > 1$ is the elasticity of substitution among labor inputs. Cost minimization implies:

$$
L_t(n,j) = \left(\frac{1}{s}\right)\left(\frac{W_t(j)}{W_t}\right)^{-\psi} L_{N,t}(j), \tag{31}
$$

where $W(j)$ is the nominal wage of labor input j and the wage index W is:

$$
W_t = \left[\left(\frac{1}{s}\right)\int_0^s W_t(h)^{1-\psi} dj\right]^{\frac{1}{1-\psi}}. \tag{32}
$$

Similar equations hold for firms producing intermediate tradable goods. Each household is the monopolistic supplier of a labor input j and sets the nominal wage facing a downward-sloping demand, obtained by aggregating demand across Italian firms. The wage adjustment is sluggish because of quadratic costs paid in terms of the total wage bill:

$$AC_t^W = \frac{\kappa_W}{2}\left(\frac{W_t}{W_{t-1}} - 1\right)^2 W_t L_t \tag{33}$$

where the parameter $\kappa_W > 0$ measures the degree of nominal wage rigidity and L is the total amount of labor in the Italian economy.

5.4 The equilibrium

We find a symmetric equilibrium of the model. In each country there is a representative agent and four representative sectorial firms (in the intermediate tradable sector, intermediate nontradable sector, consumption production sector and investment production sector). The equilibrium is a sequence of allocations and prices such that, given initial conditions and the sequence of exogenous shocks, each private agent and firm satisfy the correspondent first order conditions, the private and public sector budget constraints and market clearing conditions for goods, labor, capital and bond holdings.

The macroeconomic effects of expenditure shocks during good and bad times

Francesco Caprioli and Sandro Momigliano*

Abstract

We study how the effects of expenditure shock on economic activity are influenced by the state of the economy on the basis of various autoregressive models and indicators of cyclical conditions. For Italy over the period 1982-2011 we find some, but not conclusive, evidence that expenditure multipliers tend to be higher in recessions than in expansions.

JEL Classification: C13, H30, H63.

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1 Introduction

The large stimulus packages implemented by governments in most advanced countries to contrast the global recession that begun in mid-2008 stimulated a large debate (see Corsetti *et al.*, 2010; Romer and Romer, 2010) and brought renewed attention to the old question of the usefulness of fiscal policy to smooth cyclical fluctuations. More recently, a similar debate stemmed from fiscal consolidation policies and focused on the size of fiscal multipliers (IMF, 2102).

The theoretical literature provides limited guidance on these issues, as the qualitative effects of fiscal policy are model-dependent (see Cogan *et al.*, 2009); the empirical evidence is still not conclusive either, although it suggests that fiscal expansions generally boost private consumption and output.[1]

It has been often pointed out that the effects of fiscal policy may depend on the state of the economy (e.g., Parker, 2011; IMF, 2012), but there is still little empirical research trying to assess how the size of fiscal multipliers varies over the cycle. Indeed, most of the existing empirical literature uses linear models which, by construction, are unable to capture any dependence of fiscal multipliers on the level of aggregate demand.

In this paper we contribute to the debate by estimating for the Italian economy various threshold VARs, which allow to analyze the influence of the state of the economy on the effects of expenditure shocks. In particular, we study whether the traditional Keynesian argument of higher effectiveness of fiscal stimulus in periods characterized by slack of resources is supported by the Italian data. Similar analyses have been carried out for other countries (e.g. Canova and Pappa, 2011; De Cos *et al.*, 2013).

Our starting point is the Structural VAR (SVAR) employed in Caprioli and Momigliano (2011). Fiscal shocks are identified using the methodology developed by Blanchard and Perotti (2002), which delivers relatively efficient estimates in small samples as recently stressed by Chahrour *et al.* (2010).[2] The model includes two additional variables – government debt and foreign demand – with respect to the standard model found in the literature (5 variables: private GDP, inflation, interest rates, net revenue and government consumption). The inclusion of debt is important because it allows to take into account its influence on the fiscal authorities' decisions, particularly important in the case of Italy.[3] The inclusion of foreign demand is warranted by its strong influence on economic activity, Italy being a small open economy.

To take into account the state of the economy we first estimate the SVAR described above splitting the sample into "recessions" and "expansions" on the basis of the official chronology of the Italian economy published by Isco/Isae/Istat (Altissimo *et al*, 1999; Istat, 2011), which

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1 See Coenen *et al.* (2010). The two main empirical approaches that attempt to assess the effects of fiscal policy have specific limits. Reliable and non-interpolated quarterly fiscal data over a sufficiently long period of time, a prerequisite for the VAR approach, exist only for a few countries. The "narrative" approach (*i.e.*, Ramey and Shapiro, 1997, and Edelberg *et al.*, 1999) is resource-intensive and intrinsically subjective, making it almost impossible to apply across countries.

2 Other approaches most commonly used to identify structural shocks are the sign restrictions on impulse responses (see Mountford and Uhlig, 2002), the dummy variable one (see, e.g., Romer and Romer, 2010) and the Choleski ordering one, see, e.g., Fatás and Mihov, 2001. The literature about the effects of fiscal policy using Vector Autoregression is large and to offer a comprehensive survey goes beyond the scope of this paper. See Blanchard and Perotti (2002), Perotti (2004), Fatás and Mihov (2001), Mountford and Uhlig (2002), Giordano *et al.* (2008), Ramey and Shapiro (1997), Edelberg *et al.* (1999), and Burriel *et al.* (2010) among many others.

3 Other researchers have included public debt in a SVAR exercise examining fiscal multipliers. We broadly follow the methodology of Favero and Giavazzi (2007), who add a deterministic equation linking debt dynamics to the government budget balance. Chung and Leeper (2007) employ a conceptually similar approach. Creel *et al.* (2005) include public debt as an additional variable. This second approach allows the analysis of the effects of direct shocks on government debt. This, however, comes at the cost of estimating a higher number of parameters than actually needed, as the government budget constraint is disregarded.

identifies recessions following the methodology used for the US by the National Bureau of Economic Research.

Secondly, to assess the robustness of our results to different ways of identifying recessions and also to overcome the problem arising from the limited number of observations for such regime in the official chronology we estimate an Endogenous Threshold VAR (Fazzari *et al.*, 2012). This approach allows to endogenously identify recessions and expansions, based on an indicator of cyclical conditions. As in the previous analysis, the sample is split and two sets of parameters are estimated. We apply this method to 2 alternative indicators of cyclical conditions: the past private GDP growth and the output gap measured by the difference between the private GDP and its H.P. filter. These two indicators, commonly used in central banks, are chosen as they proxy the slack of resources in the economy.

Finally, we estimate the Smooth Transition VAR model proposed by Auerbach and Gorodnichenko (2012), which allows a gradual transition between recessions and expansion: in each quarter, the parameters of the model are a linear combination of two sets of values (corresponding to each regime), weighted by the degree of being in each regime.

The main results of this paper can be summarized as follows.

In the SVAR analysis which doesn't take into account the state of the economy, the response of private GDP to an expenditure shock is positive, hump-shaped and highly significant for approximately two years. The median value of the expenditure multiplier is equal to 1.04 on impact and reaches its peak (1.8) after three years.

When we split the sample in "expansion" and "recession" regimes, either on the basis of the official chronology or applying the ETVAR approach to the cyclical indicators, for both regimes we obtain impulse response functions (IRFs) broadly similar to those estimated for the full sample. Under the recession regime, the response of private GDP is larger and more prompt and the cumulative multiplier is stronger on average. However, the confidence bands of the estimate are relatively large, and the difference in the median value of the multiplier across regimes is not statistically significant. This difference is instead statistically significant if we apply the ETVAR analysis to the 3-variable model (namely, private GDP, government consumption and net revenue) originally proposed in Blanchard and Perotti (2002), which is more parsimonious in terms of parameters.

Estimates are generally less precise and results are less clear-cut when we apply the STVAR model: in recessions the median value of the expenditure multiplier can be higher or lower depending on the cyclical indicator used.

In conclusion, our empirical investigation shows some weak evidence that expenditure multipliers are higher in recessions than in expansions. While this result is influenced by the limited size of the sub-samples, it seems to suggest that the differences in the multipliers have not been extremely large, at least in the period under examination.

The paper proceeds as follows. Section 2 describes the data. In Section 3 we outline the specification of the VAR, ETVAR and STVAR models and our identification strategy. In Section 4 we analyze the effects of government consumption shocks without distinguishing between states of the economy. In Section 5 we discuss the results when we distinguish between the two regimes (expansions and recessions). We conclude with Section 6.

2 Data and variables

We extend up to 2011:2 the database of quarterly cash fiscal data used in Caprioli and Momigliano (2011) on the basis of the Italian Ministry for the Economy and Finance Quarterly Report and the general government borrowing requirement published by the Bank of Italy. The specification includes seven variables: private GDP, *i.e.*, total GDP net of government consumption (y_t); the inflation rate (π_t) based on the private GDP deflator; the nominal interest rate on government debt (i_t); government consumption (g_t); net taxes (t_t); the debt-to-GDP ratio (d_t); and foreign demand (f_t).

As in Caprioli and Momigliano (2011), we include GDP net of government consumption instead of total GDP. This choice stems from the fact that cash government consumption has a different quarterly profile from the corresponding national accounts aggregate, which complicates somewhat the interpretation of the effects on total GDP of a shock to (cash) government consumption, as it cannot be assumed (contrary to the case of national accounts fiscal data) to have a one-to-one impact on aggregate demand. Moreover, excluding the government component of aggregate demand from total GDP allows us to answer directly the most relevant policy question, that is how the private sector reacts to a fiscal shock.

We construct the interest rate on government debt as a weighted average of the yield on short-term and on long-term government debt, where the weight is given by the share of debt obligations with maturity shorter than one year. Government consumption is the sum of government spending on goods and services and government wages. Net taxes are computed by subtracting government consumption, interest payments and investment from the borrowing requirement; therefore this variable includes monetary transfers as well as revenue.[4]

All variables, apart from inflation, interest rate and the debt-to-GDP ratio, are log-transformed, converted in real terms using the private GDP deflator and seasonally adjusted using the TRAMO-SEATS procedure.

To identify expansions and recessions, we use the following indicators: i) the official chronology of the Italian economy, based on the National Bureau of Economic Research, produced by Istat-Isae-Isco (cfr. Altissimo *et al.*, 1999); ii) past 4-quarters private GDP growth; iii) output gap, computed on the basis of the Hodrick Prescott filter.

3 The models and the identification strategy

3.1 The SVAR

The reduced-form VAR is specified in level (as shown by Sims *et al.*, 1990), in large samples it is possible to ignore the cointegrating vector) and can be written as follows:

$$X_t = \sum_{i=1}^{k_1} C_i \, X_{t-i} + \sum_{i=1}^{k_2} \gamma_i \, d_{t-i} + \sum_{i=0}^{k_3} \delta_i \, \log(f_{t-i}) + U_t \qquad (1)$$

where:

[4] We exclude public investment from our benchmark specification (as in Giordano *et al.*, 2008), because we are not confident enough about the quality of the data. Results do not qualitatively change as a result of adding investment to either government consumption or net revenue, as shown in Subsection 4.2.

$$X_t = \begin{bmatrix} \log(y_t) \\ \pi_t \\ i_t \\ \log(t_t) \\ \log(g_t) \end{bmatrix} \tag{2}$$

k_1, k_2 and k_3 are the number of lags for the variables included in the VAR, for the debt-to-GDP ratio and for the foreign demand variable respectively.

U_t is the vector of reduced-form residuals. k_1, k_2 and k_3 are set to the minimum number of lags that delivers serially uncorrelated reduced-form residuals. In particular, they are set equal to 2, 1 and 1 respectively. A constant and a deterministic linear trend are included. According to equation (1), past values of the debt-to-GDP ratio influence the current values of macroeconomic variables, which conversely influence the current value of the debt-to-GDP ratio according to the following law of motion:

$$d_t = \frac{1+R_t}{(1+\pi_t)\left(\frac{y_t}{y_{t-1}}\right)} d_{t-1} + \frac{g_t - t_t}{y_t} \tag{3}$$

where:

$$R_t = \sum_{j=0}^{N} \frac{i_{t-j}}{N} \tag{4}$$

Equation (3) represents the period-by-period government budget constraint, expressed as a ratio to total GDP. Changes in the interest rate on government debt i_t only gradually affect its average cost R_t in equation (4); we set $N = 20$, as 5 years is approximately the financial duration of the debt at the end of our sample.

Compared with Favero and Giavazzi (2007), we add equation (4) and include in equation (1) the actual yield at issuance instead of the average cost of servicing public debt. We do so to identify more precisely the reaction of financial markets to the state of the public finances. In fact, the yield at issuance responds immediately to investors' sentiments, while the average cost adjusts with a relatively long delay, depending on the maturity structure of government obligations. Moreover, the yield at issuance is more directly relevant for investment decisions in the private sector.

We assume that, while current and past values of foreign demand affect the current values of macroeconomic and fiscal variables, the reverse is not true. This assumption seems appropriate as Italy is a relatively small open economy. As a measure of foreign demand, we follow Busetti *et al.* (2011), who compute the demand of Italian goods from abroad as:

$$f_t = \sum_{j=1}^{N} M_{j,t} \, \overline{q_j} \tag{5}$$

where $M_{j,t}$ corresponds to the total imports of goods by country j in volume at time t weighted by $\overline{q_j}$, the average ratio over the period 1999-2001 between Italian exports towards country j and total Italian exports. Busetti *et al.* (2011) construct this index for commercial partners both belonging to the Euro area and outside the EU. As a measure of global foreign demand, we consider the sum of the two indices.[5]

3.3 *The ETVAR and STVAR models*

In the ET-VAR model, equation (1) is substituted by equations (6) and (7), while equations (2)-(5) remain unchanged.

$$
\begin{aligned}
X_t =&(1-F(z_{t-1}))[\sum_{i=1}^{k_1^E} C_i^E X_{t-i} + \sum_{i=1}^{k_2} \gamma_i^E d_{t-i} + \sum_{i=0}^{k_3} \delta_i^E log(f_{t-i}) + U_t^E]+ \\
&+F(z_{t-1})[\sum_{i=1}^{k_1^R} C_i^R X_{t-i} + \sum_{i=1}^{k_2} \gamma_i^R d_{t-i} + \sum_{i=0}^{k_3} \delta_i^R log(f_{t-i}) + U_t^R]
\end{aligned} \tag{6}
$$

Equation (7) states that the economic system is described by two piecewise linear models with different sets of coefficients. The recessionary and expansionary regimes are identified by the transition indicator function *F*(z_{t-1}), defined by:

$$
F(z_{t-1}) = I(z_{t-1} \leq r) = \begin{cases} 1, & if z_{t-1} \leq r \\ 0 & if z_{t-1} > r \end{cases} \tag{7}
$$

where z_{t-1} is the threshold variable and *r* is the threshold value. The threshold variable is lagged one quarter - to avoid contemporaneous feedback effects from the model to the regime-switching probability – and normalized to have zero mean and unit variance. The model, which allows different lags for the autoregressive part across regimes and regime-specific covariance matrices, is estimated in two steps. First, for a given value of the threshold *r*, the regime-specific coefficients and covariance matrices are estimated by OLS using observations from each regime; second, the threshold value is estimated by minimizing the conditional likelihood over a grid of values, namely:

$$
\hat{r} = arg \min_{r \epsilon R} (\frac{T_R}{2} log|\hat{\Omega^R}_t| + \frac{T_E}{2} log|\hat{\Omega^E}_t|) \tag{8}
$$

In the STVAR model the transition function takes a more general form, given by:

$$
F(z_{t-1}) = \frac{exp(-\gamma z_{t-1})}{1+exp(-\gamma z_{t-1})}, \gamma > 0 \tag{9}
$$

which allows a smoother transition across the two regimes than what prescribed by equation (7). The weighting function *F*(z_{t-1}), the probability to be in a recession, depends on the business cycle indicator z_{t-1}; as it is imposed $\gamma>0$ in equation (9), the lower the z_{t-1}, the higher *F*(z_{t-1}), as shown in Figure 1, for different values of the γ parameter. As for the case of the threshold value in ETVAR, in the STVAR the γ parameter is estimated by minimizing the likelihood function.

[5] As a robustness check, we use also the world trade series obtained from IMF International Financial Statistics. The use of this series to measure foreign demand does not change results.

3.4 Identification strategy

The identification strategy is identical for the three models described above. The only difference is that in the ETVAR the procedure is applied to the residuals of each regime.

Reduced-form residuals associated with the fiscal variables, u_t^g and u_t^t can be written as linear combinations of the structural fiscal shocks and of the reduced-form residuals of the other variables in the VAR:

$$u_t^g = \alpha_y^g \, u_t^y + \alpha_\pi^g \, u_t^\pi + \alpha_i^g \, u_t^i + \beta_t^g \, \varepsilon_t^t + \varepsilon_t^g \tag{7}$$

$$u_t^t = \alpha_y^t \, u_t^y + \alpha_\pi^t \, u_t^\pi + \alpha_i^t \, u_t^i + \beta_g^t \, \varepsilon_t^g + \varepsilon_t^t \tag{8}$$

The α coefficients contain both the automatic elasticity and the discretionary change to the macro variables innovations, while the *β* coefficients measure the response of the fiscal variables to a structural shock. To estimate the *α* and *β* coefficients in equations (7)-(8) we follow the approach in Blanchard and Perotti (2002). First, we assume that, within a quarter, the discretionary change of fiscal variables to innovations in the macro variables is zero. Using quarterly data, this assumption can be justified on the ground of decision lags in fiscal policy-making which last longer than three months. Secondly, we estimate the *α* in equations (7)-(8) using external information on the elasticities of government consumption and taxes to output, inflation and interest rate. Following Giordano *et al.* (2008) (Appendix B therein) in this paper, we set $\alpha_\pi^g = -0.9$, $\alpha_y^t = 0.3$, $\alpha_\pi^t = -0.4$ and all the other α equal to zero. In addition, we assume that government consumption does not contemporaneously adjust to revenues, *i.e.*, we set β_t^g equal to zero. Consequently, we estimate β_g^t from equation (8) using OLS. We verify that even sizeable changes in these parameters do not significantly affect our results.

Finally, we estimate the coefficients relating the reduced-form macro variables residuals to the fiscal ones by instrumental variables, using as instruments for u_t^g and u_t^t their corresponding structural shocks, uncorrelated by definition.

It is important to notice that the identification strategy for structural shocks does not depend on the presence of the debt-to-GDP ratio, as the latter follows a deterministic law of motion. In other words, equation (3) holds as an identity and therefore it does not add any shock to the ones already included in the VAR model specified in equation (1).

A problem with the fiscal shocks identified using the VAR approach is that they may be anticipated by economic agents, owing to the delay between the announcement of fiscal measures and their actual implementation. In order to check for this possibility, we run Granger causality tests between the fiscal shocks estimated with the benchmark model and survey expectations about future policy actions and macro variables. The results do not support the hypothesis that fiscal shocks were anticipated.[6]

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[6] As for survey expectations, we use the Consensus mean forecasts of *i*) the annual growth rate of real GDP, private consumption, gross fixed investment, industrial production, consumer and producer prices, *ii*) unemployment rate (as a percentage of the labor *(continues)*

4 The effects of government consumption shocks in a SVAR model

Figure 2 show the response of the fiscal and macroeconomic variables to an exogenous shock (equal to 1% of private GDP) to government consumption in our benchmark model. In each panel the solid line represents the median response, while the dashed lines represent two sets of lower and upper bands, corresponding to the 5^{th}, 16^{th}, 84^{th} and 95^{th} percentiles of the distribution of the responses at each horizon, as commonly done in the literature.[7]

Concerning the reaction of fiscal variables, two points are worth mentioning. The first is that the government consumption shock is largely short-lived, being equal to 0.1% of private GDP already after four quarters. The second is that the higher public consumption is rapidly financed by higher revenues, which increase already in the first quarter, remain broadly constant at 0.2% of GDP for two years and then slowly decrease. The rise in net revenue, ensuring that the initial surge in the debt is fully absorbed within three years, reflects their direct stabilizing discretionary reaction to the debt and, to a lesser extent, to the increase in private GDP (see below).

After a shock to public consumption, the response of private GDP is positive and highly significant for approximately two years. The peak, reached at the fourth quarter, is equal to 0.25% of GDP. Positive and significant effects of government consumption shocks on economic activity represent a relatively common result of the VAR literature (e.g., Giordano *et al.*, 2008; Perotti, 2004; Mountford and Uhlig, 2002; and Neri, 2001). The output response to government consumption reflects the low persistence of the shock. To make our results more directly comparable with analyses which focus on total GDP (instead of its private component) and analyze shocks with a different persistence, we compute the cumulative multiplier (*i.e.*, the ratio of the cumulative change in total GDP to the cumulative change in total government consumption)[8] charted in Figure 3. The median value is equal to 1.04 on impact, reaches its peak (1.8) after three years and remains roughly constant thereafter. The confidence bands are relatively narrow compared with similar studies, with the 95^{th} and the 5^{th} percentiles of the distribution remaining above 1.3 and below 2.4 after the fifth quarter.

The median value for the long-run fiscal multiplier lies in the upper part of the wide range of estimates provided by the empirical literature. As shown in Spilimbergo *et al.* (2009), the relatively high value of the multiplier may be due to the debt-stabilizing reaction of fiscal variables. The transitory nature of the government consumption shock, rapidly compensated by higher revenues, and the small – and delayed – increase in interest rates do not pose a threat to the sustainability of the Italian public debt, notwithstanding its high level, making any precautionary savings by households unnecessary. The response of private GDP is robust across alternative specifications of the model.[9]

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force), current account and state sector budget balance, and *iii*) three-month euro-area interest rate and 10-year Italian government bond yield. Following Ramney (2008) and Kirchner *et al.* (2010a), the fiscal shocks at time t are regressed on a constant, its own lag and the previous forecasts made in period t–1 for period t.

[7] We compute confidence bands for IRF by bootstrapping. After estimating equation (1), we obtain fitted residuals $\hat{u}_1, \dots, \hat{u}_T$ normally distributed with zero mean and covariance matrix Ω. We draw errors from this distribution to simulate the system of equations (1)-(5) L times. For each draw we compute the IRF as described in the previous footnote. Finally, we collect the α^{th} and $1-\alpha^{th}$ percentile across the 1000 draws.

[8] Following Giordano *et al.* (2008), we compute total GDP in this context by adding the cash-based government consumption included in the model to private GDP.

[9] As robustness checks, we considered the following model specifications in which: i) we include the interest rate only on debt obligations with a maturity shorter than one year; ii) we use the gross yield on debt obligations with a maturity longer than three

(continues)

The reaction of inflation to a government consumption shock is not statistically significant. This is in line with the analyses of Marcellino (2006), King and Plosser (1985) and Henry *et al.* (2004). The response of interest rates is relatively small, hump-shaped and never statistically significant. The existence of a positive relationship between interest rates and the level of government debt can be found in many empirical studies (see Bernheim, 1987, 1989; Gale and Orzag, 2002; Miller and Russek, 1996; and Engen and Hubbard, 2004).[10]

4.1 The role of government debt and foreign demand

The model includes two additional variables – government debt and foreign demand – with respect to the standard model found in the literature (5 variables: private GDP, inflation, interest rates, net revenue and government consumption). The inclusion of debt is important because it allows to better understand the fiscal framework associated with the shock. In particular, the reaction of fiscal variables – namely, government spending and net revenue – to changes in public debt can be analyzed.[11] Empirical evidence (see Bohn, 2007; Trehan and Walsh, 1991; Hamilton and Flavin, 1986; and Golinelli and Momigliano, 2008) suggests that this feedback effect is generally important. In the case of a high-debt country like Italy, the influence of debt on the fiscal authorities' decisions is likely to be particularly large.[12] The inclusion of foreign demand is warranted by its strong influence on economic activity, Italy being a small open economy. As it can be safely assumed that foreign demand, measured by world demand, is not significantly influenced by Italian macro or fiscal variables, its inclusion in the VAR comes at a relatively small cost in terms of additional parameters to be estimated.

The left and right panels of Figure 4 show the impact of including public debt and/or foreign demand in the model respectively on the median response of private GDP and on the accuracy of this estimate, measured by the distance between the 95th and the 5th percentiles of the distribution.

Compared with a five-variable model that excludes both public debt and foreign demand, adding public debt determines a stronger (twice larger on average in the first two years) and longer lasting response of private GDP to a consumption shock (left panel). These results give support to the argument of Favero and Giavazzi (2007) that omitting debt in the model can result in biased estimates of the effects on GDP of fiscal shocks. The authors stressed the need to take into account the reactions of fiscal variables to changes in debt. In our case, these reactions would dampen the effects on output. On the contrary, we find a larger effect on private GDP, which comes from

years; iii) the specification of the VAR includes a quadratic trend instead of a linear one; iv) we include government investment in our definition of government consumption; v) net revenues come first when identifying the shocks (in the benchmark model, government consumption is ordered first); vi) the reduced-form residuals of fiscal variables depend explicitly on the level of government debt; and vii) the average financial duration is set equal to two years instead of its end-of-sample value (five years). We do not report these robustness checks, as estimates stay almost unchanged with respect to the benchmark specification. The results obtained with these alternative specifications confirm the hump-shaped pattern of private GDP and, apart from the "quadratic trend" specification for the quarters 6-10, they are well within the upper (95th percentile) and lower (5th percentile) bands of the GDP response in the benchmark specification. In the case of the "quadratic trend" specification, the lower impact on private GDP largely reflects the shorter persistence of the expenditure shock. The cumulative multiplier is very close to that for the benchmark specification.

[10] The results for inflation and interest rates are also robust across the alternative specifications described in the previous footnote.

[11] Recent research suggests that, depending on whether or not an expenditure shock is reabsorbed in the medium-long term, fiscal multipliers may have different values (see Corsetti *et al.*, 2009; and Ilzetzki *et al.*, 2009).

[12] Other researchers have included public debt in a SVAR exercise examining fiscal multipliers. We broadly follow the methodology of Favero and Giavazzi (2007), who add a deterministic equation linking debt dynamics to the government budget balance. Chung and Leeper (2007) employ a conceptually similar approach. Creel *et al.* (2005) include public debt as an additional variable. This second approach allows the analysis of the effects of direct shocks on government debt. This, however, comes at the cost of estimating a higher number of parameters than actually needed, as the government budget constraint is disregarded.

allowing a direct influence of debt on output.[13] Adding also the foreign demand (so as to reach our benchmark specification) does not instead have a sizeable effect on the response of private GDP.

Compared with a five-variable model that excludes both public debt and foreign demand, adding public debt determines a very large improvement in the precision of estimates: the confidence band of the response shrinks almost to a third, on average (right panel of Figure 4). This is not a surprise, given its major influence on Italian macroeconomic developments. Adding the debt also improves the accuracy of the estimates further, but to a lesser extent.

5 Distinguishing across states of the economy: the effects of government consumption shocks

Compared to the analysis presented in Section 4, here we estimate the effects of expenditure shocks distinguishing between states of the economy. There is no consensus in the literature on the most appropriate indicator for the business cycle. The official chronology of the Italian economy (Altissimo *et al.*, 1999; Istat, 2011) which follows the methodology followed in the US by the National Bureau of Economic Research, identifies 28 quarters (out of 118) as "recessions" in the period 1982-2011. We label the other periods as "expansions". As a dichotomy variable cannot be used in the ETVAR and STVAR models, we employ as business cycle indicators the 4-quarters private GDP growth and the output gap, measured by the difference between the private GDP and its H.P. filter.

Figure 5 shows the recession periods based on the official chronology and those identified by the ETVAR model using the alternative indicators just mentioned. The ETVAR-based recessions are generally shifted forward with respect to the official chronology. Also, with ETVAR, the sample is more evenly split between the two regimes.

Nevertheless, Figure 5 allows to readily identify the four major recessions in our sample: the one at the beginning of the eighties, triggered by the second oil shock; the one at the beginning of the nineties, determined by the financial crisis; the strong slowdown in the initial years of the last decade and, finally, the last episode, influenced by the Lehman Bros' collapse.

5.1 Identifying recessions on the basis of the official chronology of the Italian economy

Figures 6 and 7 show the response of the fiscal and macroeconomic variables to a government consumption shock in the recession periods identified by the official chronology and in the other periods ("expansions"), respectively. As in the IRFs previously discussed, the shock is equal to 1% of private GDP and the solid line represents the median response, while the dashed lines represent the 5^{th}, 16^{th}, 84^{th} and 95^{th} percentiles of the distribution of the responses.

When constructing impulse responses for a given regime, we assume that the state of the economy when the shock occurs does not change; in particular, we ignore any feedback effect from the fiscal shock to the type of regime. As this assumption becomes stronger the more we extend the time horizon of our analysis, we narrow it to 8 quarters in these IRFs and in the following ones.

The results are relatively close to those described in Section 4: in both regimes, the response of private GDP is positive and hump-shaped; revenues show a positive reaction making the initial surge in the debt to be gradually absorbed. However, under the recession regime, the response of

[13] Another possible explanation for the greater response of private GDP could be that the inclusion of debt led to a better identification of the exogenous fiscal shocks (as the endogenous reactions of fiscal variables to changes in debt were excluded). However, we compared estimated fiscal shocks obtained with and without debt and differences were negligible.

private GDP is more prompt (the peak effect of 0.4% is reached in the second quarter) and stronger on average in the first year. The response fades faster than in the expansion regime, but this is due to the fact that the expenditure shock goes to zero already in the second quarter, while in the expansion it diminishes more gradually.

Figure 8 compares the cumulative multiplier in the two regimes. In recessions, the median multiplier is constantly higher than in expansions; however, due to the very large confidence bands in the recession regime, the two bands largely overlap (note that the graph includes only one standard deviation bands), indicating that the difference between the two estimates is not statistically significant.

The very large size of the confidence bands estimated in the recession regime reflects the limited number of observations for this regime in the official chronology. To try to overcome this problem, we also run the SVAR adding to each recession the semester following it, as it is likely that substantial slack remains at the start of a recovery (in this way the number of observations for recessions increases to 40). This change improves the precision of estimates (in particular, the effects on private GDP in recession becomes statistically significant for 4 quarters) and reduces, but it does not eliminate, the overlap between the confidence bands of the estimates of the multipliers (Figure 9).

5.2 Identifying recessions on the basis of an ETVAR

While in the previous analysis recessions and expansions were identified outside the model, in this section the two regimes are endogenously identified in the ETVAR model on the basis of our indicators of cyclical conditions. The expenditure multipliers are reported in Figure 10. Results broadly confirm the analysis based on the official chronology: under the recession regime, the response of private GDP is more prompt and stronger. Compared to the previous analysis, the estimates for the recession regime (which includes 32 and 45 observations for the private GDP growth rate and the output gap respectively) are more accurate, but there is still a large overlap between the confidence bands of the fiscal multipliers. In order to overcome the problem of the limited number of observations in each sub-sample, we estimate the 3-variable model (namely, private GDP, government consumption and net revenue) originally proposed in Blanchard and Perotti (2002), which is more parsimonious than our benchmark model in terms of number of parameters to be estimated; in this case the difference across fiscal multipliers in the two regimes is positive and statistically significant (1-standard-deviation confidence bands do not overlap), but only using the 4-quarter private GDP growth rate (Figure 11).

5.3 Identifying recessions on the basis of an STVAR

In the previous two sections, the sample was split into two regimes (recessions and expansions) and two sets of parameters were estimated. The STVAR model allows instead a smooth transition between the two regimes; in each quarter the parameters of the model are a linear combination of two sets of values (corresponding to each regime), weighted by the degree of being in each regime (Auerbach and Gorodnichenko, 2012).

In order to produce IRFs with this approach (and also to identify fiscal multipliers), we need to select two benchmarks, representative for recessions and expansions. We select the benchmarks so to leave outside them 20 per cent of the observations. For the 4-quarter moving average of the private GDP annualized growth rate, these benchmarks corresponds to, respectively, –2,5 and 1.5. Figure 12 shows that the estimated degree of being in recession increases in those periods identified as recessions on the basis of the official chronology, especially when using the 4-quarters private GDP growth. The estimates of the expenditure multipliers are reported in Figure 13. Results are far less clear-cut than in the previous analysis. Depending on the cyclical indicator used, in recessions the expenditure multiplier is constantly higher (4-quarters private GDP growth) or constantly lower (output gap). Estimates are generally less precise than those obtained with the ETVAR model, as shown by the larger confidence bands compared to those in Figure 10 (imprecise estimates, not reported, are also obtained in the case of the 3-variable model).

8 Conclusions and future research

In this paper we study how the effects of expenditure shock on economic activity are influenced by the state of the economy on the basis of various autoregressive models and two commonly used indicators of cyclical conditions. We rely on quarterly cash-basis fiscal data for the Italian economy covering the period 1982:1-2011:2.

The main results can be summarized as follows.

Independently of the method that we use and whether we distinguish or not between states of the economy, the expenditure shocks that we estimate tend to be largely transitory and revenues show a positive reaction, making the initial surge in the debt to be gradually absorbed.

In the analysis which doesn't take into account the state of the economy, the response of private GDP to an expenditure shock is positive and highly significant for approximately two years. The government consumption multiplier (1.04 on impact and 1.8 at the peak) lies in the upper part of the wide range of estimates provided by the empirical literature.[14]

When we split the sample in "expansion" and "recession" regimes, either on the basis of the official chronology or applying the ETVAR approach, the regime-specific IRFs are broadly similar to those estimated for the full sample but generally less precise. Under the recession regime, the response of private GDP is larger and more prompt; the cumulative multiplier is also stronger on average. However, the confidence bands of the estimates are relatively large, and the difference in the median across regimes is not statistically significant. Only using a more parsimonious model and the private GDP growth rate as indicator of cyclical conditions the differences across regimes are statistically significant.

Median results are less clear-cut when we estimate the STVAR model: in recession the median value of the expenditure multiplier is higher or lower depending on the cyclical indicator used,. This result is associated to estimates that are even less precise than those obtained with the ETVAR approach. This is somewhat unexpected, as this method is deemed to use more efficiently the information of the sample; a possible explanation is that weighting function $F(z_{t-1})$, in the presence of highly imperfect indicators of cyclical conditions, represents a sort of unwarranted straightjacket for the data.

Results do not seem to give a clear support to the idea that higher multipliers may be due, at least partly, to a different behaviour of interest rates, as their response differs across methods and indicators; also the negative wealth effect (Barro, 1982) and/or the credit-crunch channel associated to public debt during recessions (De Bonis and Stacchini, 2009) can lower the direct stimulus provided to the economic activity by higher public consumption.

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[14] This may be due to the debt-stabilizing reaction of revenues, in line with the idea that the effects of fiscal stimulus on economic activity depend positively on the soundness of fiscal policy (see, e.g., Corsetti *et al.*, 2009). The transitory nature of the shocks that we observe and their small size may also have a bearing on the value of the multiplier.

In conclusion, our empirical investigation shows some weak evidence that expenditure multipliers tend to be higher in recessions than in expansions. While this result is influenced by the limited size of the sub-samples, it seems to suggest that the differences in the multipliers have not been extremely large, at least in the period under examination.

Finally, our empirical analysis could be strengthened along at least two lines. First, the assumption that the initial state of the economy does not change when constructing impulse responses for a given regime should be relaxed. Second, a more thorough selection and discussion of business cycle indicators should be conducted.

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FIGURES

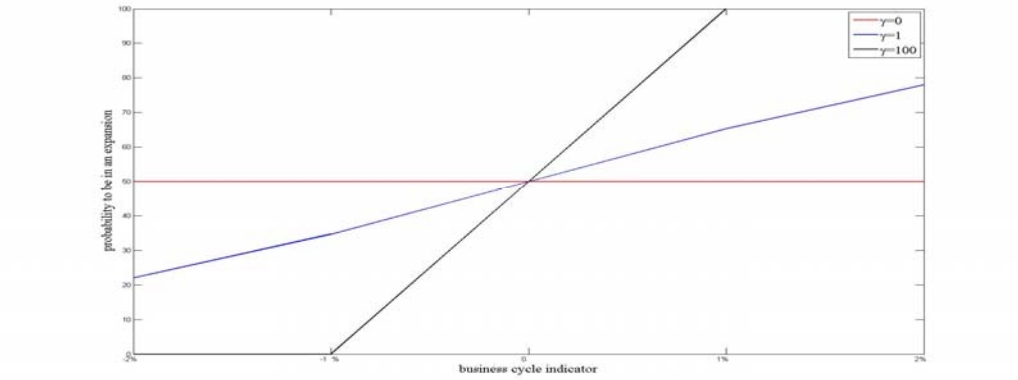

Figure 1: Probability to be in an expansion

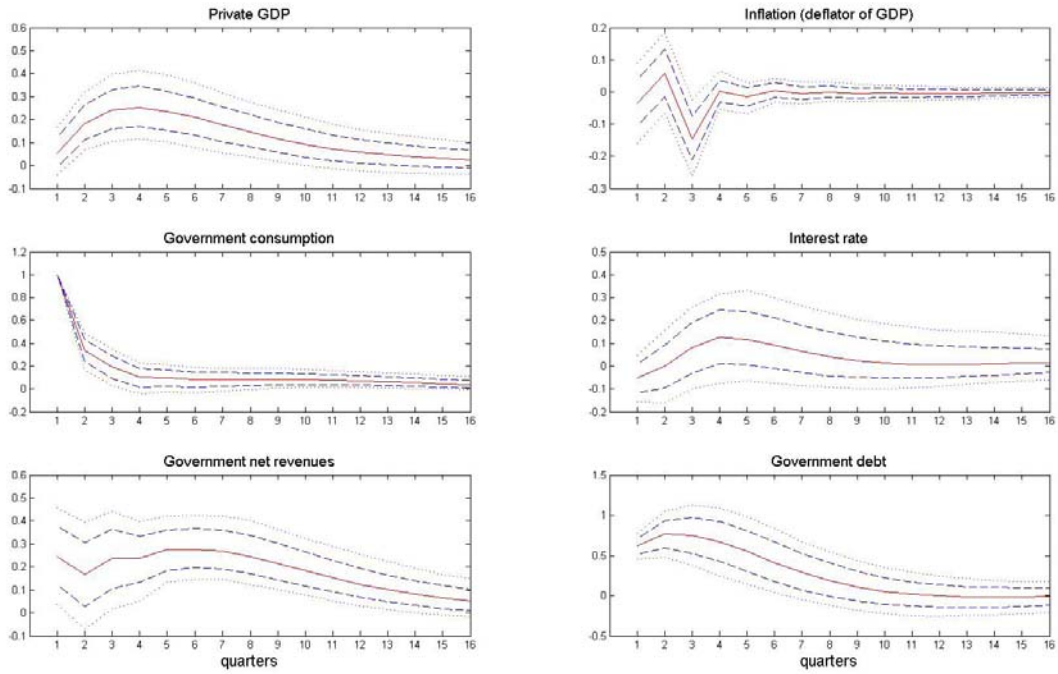

Figure 2: Impulse responses to a positive government consumption shock equal to 1% of private GDP: SVAR model. The curves represent the median and two sets of lower and upper bands, corresponding to the 5th, 16th, 84th and 95th percentiles of the distribution. Responses, except for inflation and interest rate, are deviations from the baseline and expressed in percentage points of private GDP. Inflation and interest rate responses are deviations from the baseline and expressed in percentage points.

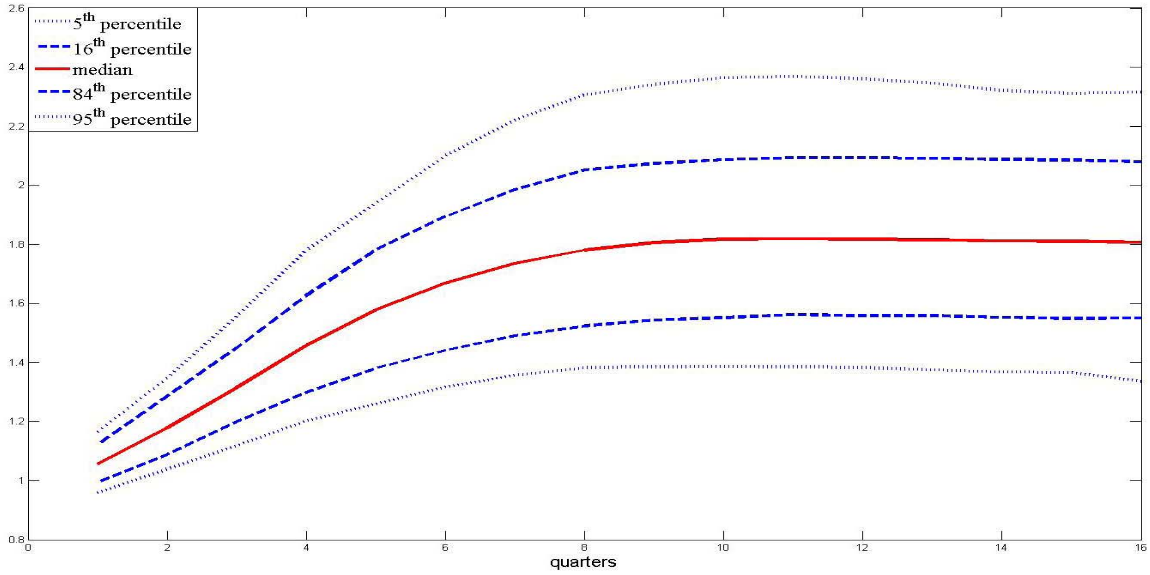

Figure 3: Cumulative multiplier of government consumption on GDP: SVAR model. The curves represent the median and two sets of lower and upper bands, corresponding to the 5th, 16th, 84th and 95th percentiles of the distribution.

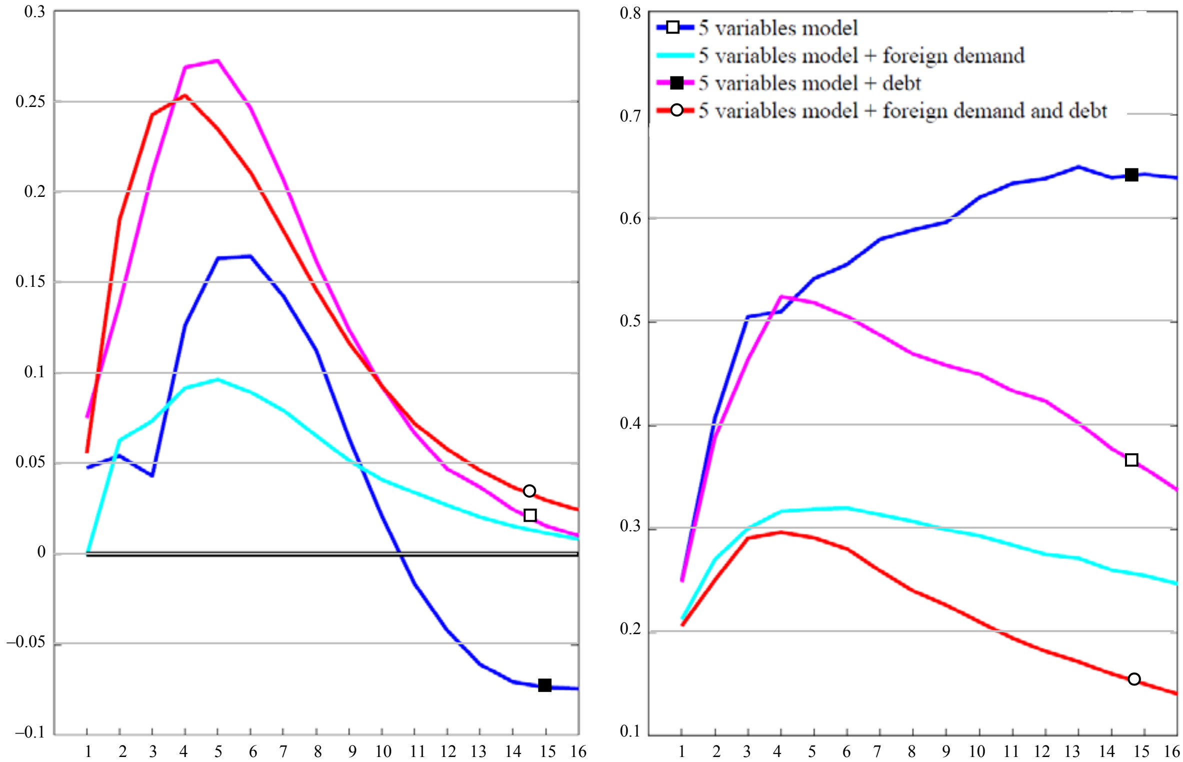

Figure 4: Left panel: Effects on private GDP of a shock to government consumption: SVAR model and alternative models which exclude debt and/or foreign demand (median values; % of private GDP); Right panel: Size of confidence bands of the estimates of the effects on private GDP of a shock to government consumption (difference between the 95th and 5th percentiles of the distribution of the private GDP responses; % of private GDP): benchmark specification and alternative models which exclude debt and/or foreign demand.

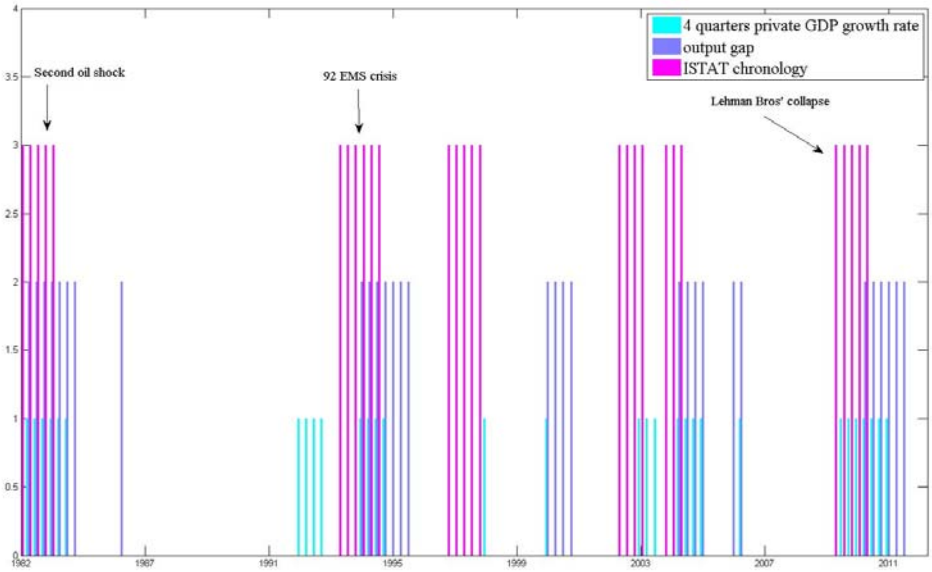

Figure 5: Recession periods based on: 1) the official chronology of the Italian economy published by Isco/Isae/Istat; 2&3) ETVAR analyses using as cyclical indicators the 4 quarters private GDP growth rate and the output gap, respectively.

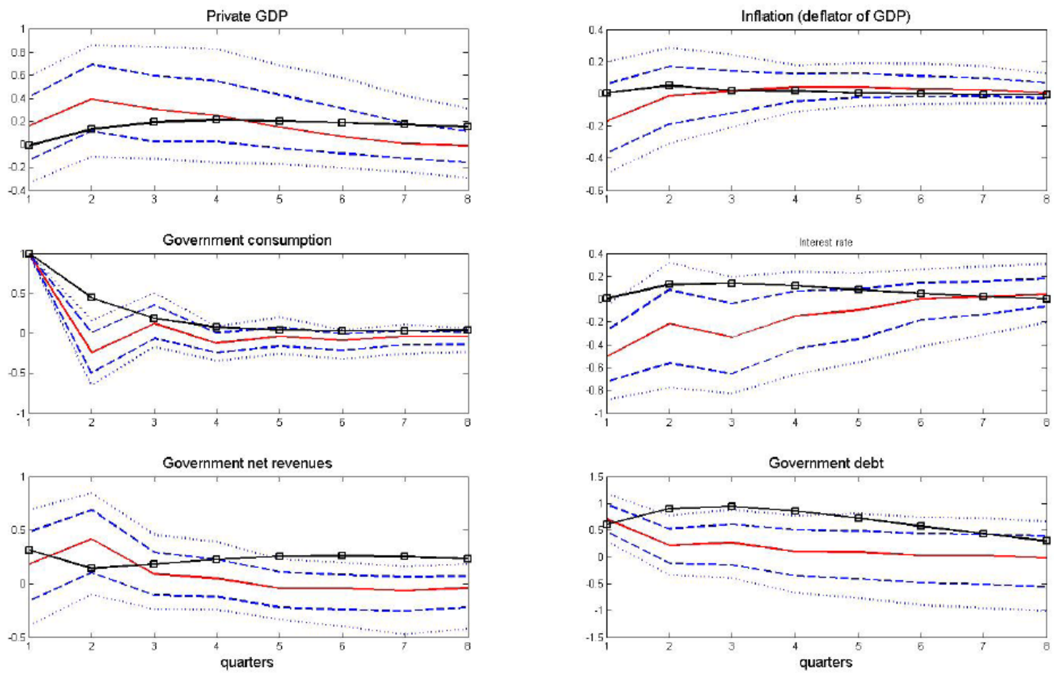

Figure 6: Impulse responses in recession to a positive government consumption shock equal to 1% of private GDP: SVAR model using the ISTAT chronology. The curves represent the median and two sets of lower and upper bands, corresponding to the 5th, 16th, 84th and 95th percentiles of the distribution. ***The solid curve with bullets represents the median response in expansion****. Responses, except for inflation and interest rate, are deviations from the baseline and expressed in percentage points of private GDP. Inflation and interest rate responses are deviations from the baseline and expressed in percentage points.*

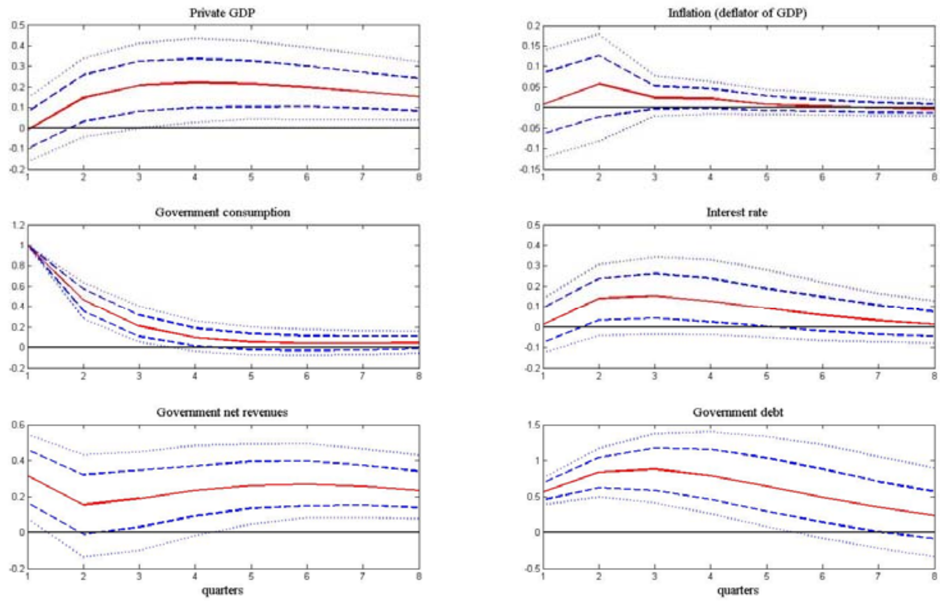

Figure 7: Impulse responses in expansion to a positive government consumption shock equal to 1% of private GDP: SVAR model using the ISTAT chronology. The curves represent the median and two sets of lower and upper bands, corresponding to the 5th, 16th, 84th and 95th percentiles of the distribution. Responses, except for inflation and interest rate, are deviations from the baseline and expressed in percentage points of private GDP. Inflation and interest rate responses are deviations from the baseline and expressed in percentage points.

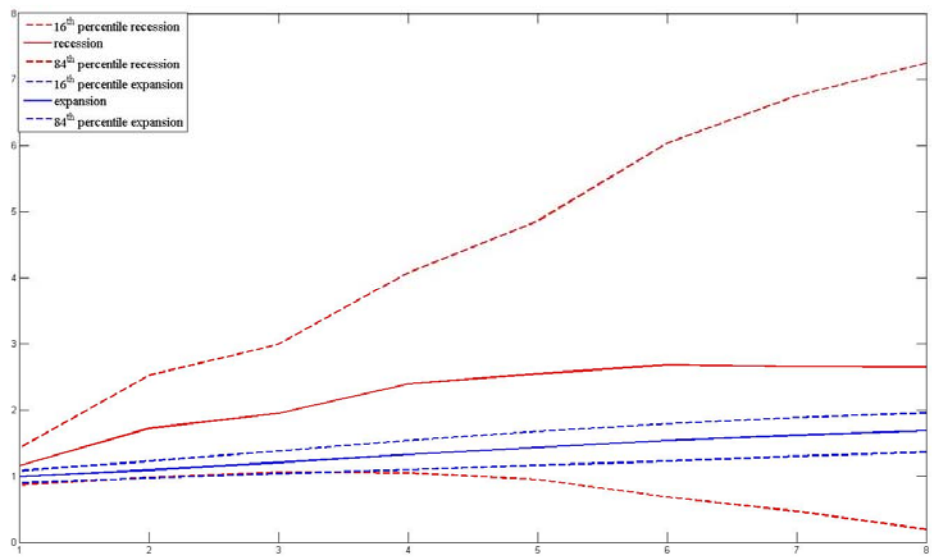

Figure 8: Fiscal multipliers in recession (red lines) and expansion (blue lines) using the ISTAT chronology.

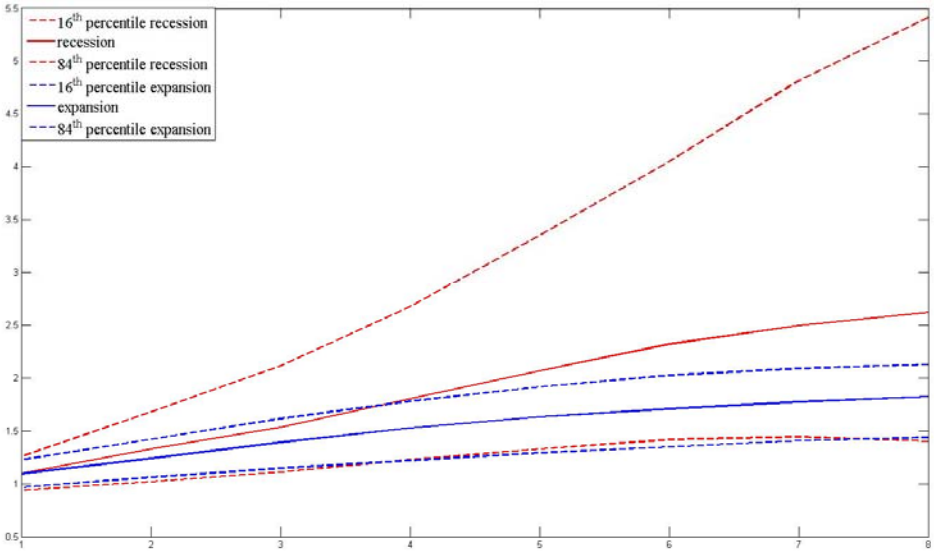

Figure 9: Fiscal multipliers in recession (red lines) and expansion (blue lines) using the ISTAT chronology adding two quarters at the end of each recession date.

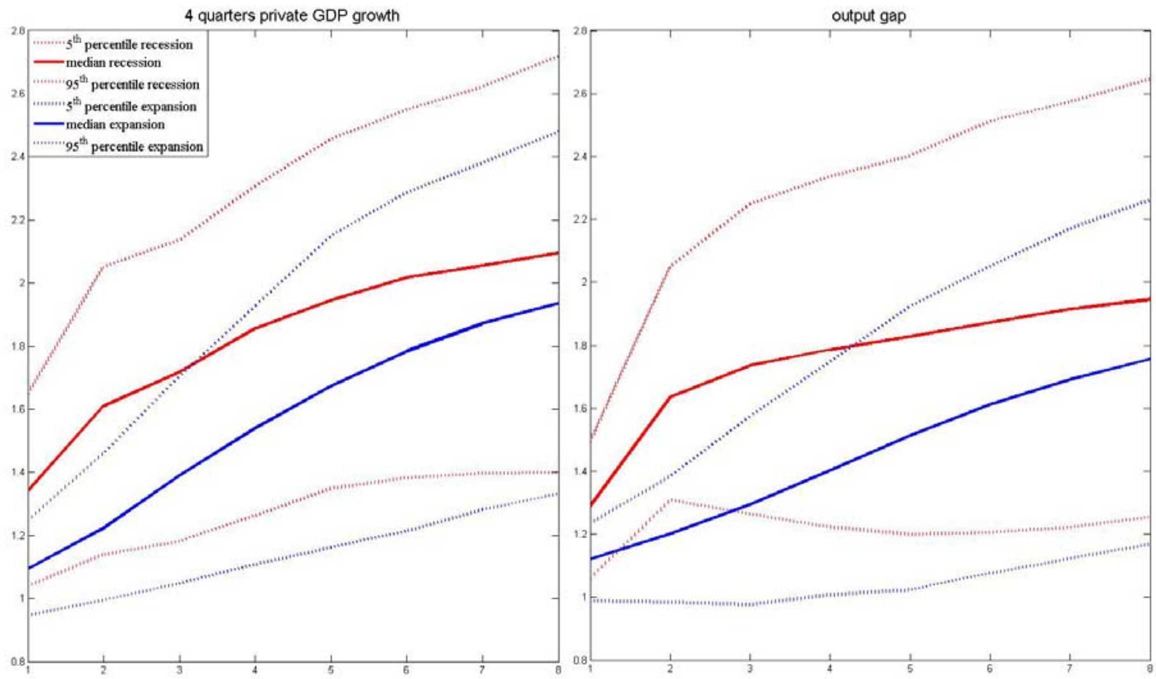

Figure 10: ETVAR analysis: Fiscal multipliers in recession (red lines) and expansion (blue lines) in our benchmark model; indicators of cyclical conditions: 4 quarters private GDP growth rate (left panel) and output gap (right panel).

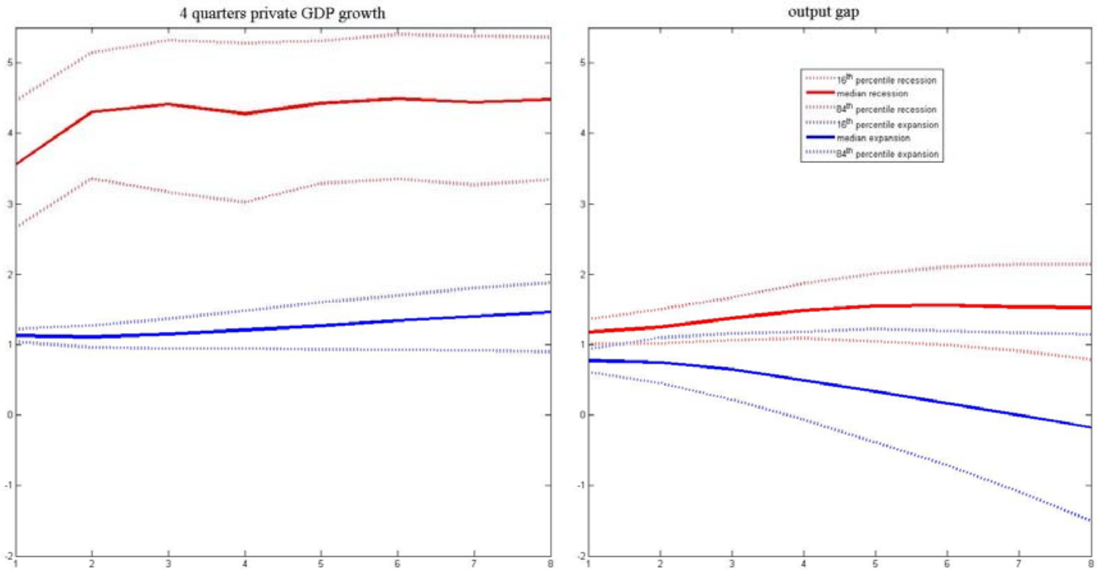

Figure 11: ETVAR analysis Fiscal multipliers in recession (red lines) and expansion (blue lines) in the 3-variable model; indicators of cyclical conditions: 4 quarters private GDP growth rate (left panel) and output gap (right panel).

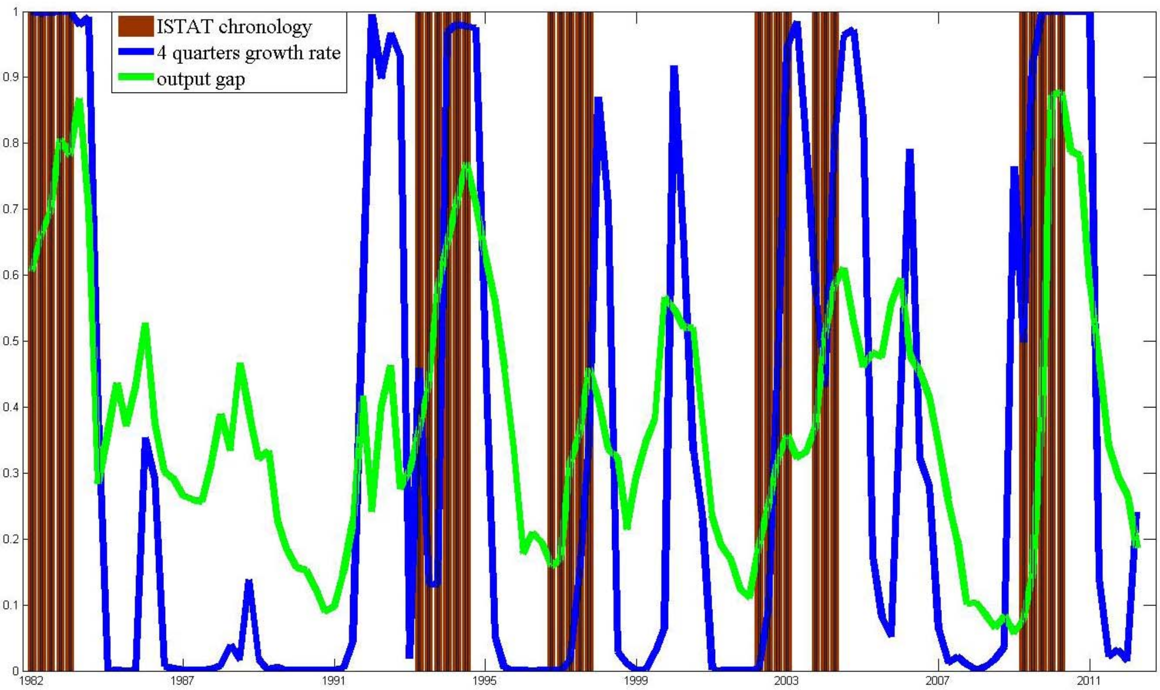

Figure 12: Probability to be in a recession based on: 1) the official chronology of the Italian economy published by Isco/Isae/Istat; 2&3) STVAR analyses using as cyclical indicators the 4 quarters private GDP growth rate and the output gap, respectively.

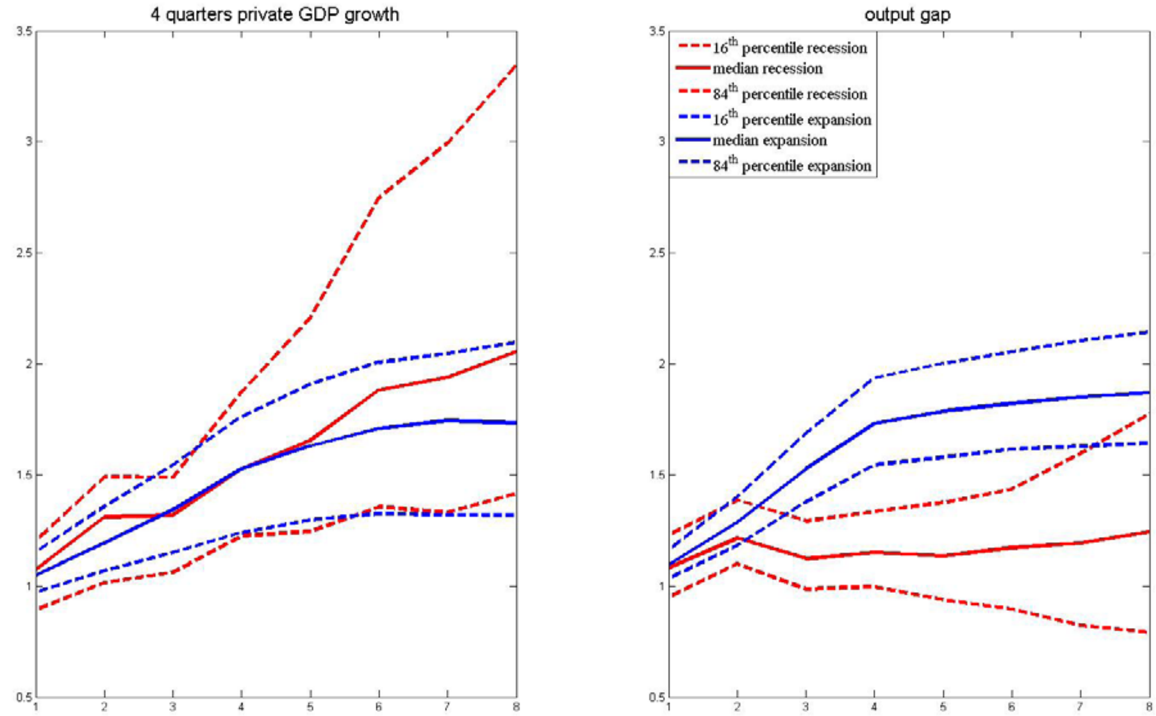

Figure 13: STVAR analysis: Fiscal multipliers in recession (red lines) and expansion (blue lines) in our benchmark model; indicators of cyclical conditions: 4 quarters private GDP growth rate (left panel) and output gap (right panel).