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FIRMS' AND HOUSEHOLDS' INVESTMENT IN ITALY: THE ROLE OF CREDIT CONSTRAINTS AND OTHER MACRO FACTORS

by Claire Giordano*, Marco Marinucci* and Andrea Silvestrini*

Abstract

Using significantly under-exploited data from institutional sector accounts, we assess the main drivers of both firms' and households' investment in Italy over the past two decades. We estimate a vector error correction model separately for firms and for households. Our findings support the existence in both institutional sectors of a long-run equilibrium relationship between investment, income and the user cost of capital, as predicted by the flexible neoclassical model, as well as adjustment dynamics towards the equilibrium level. Moreover, we find evidence that an increase in uncertainty and a decline in economic sentiment have a dampening effect on investment. Furthermore, high indebtedness, measured by financial accounts data, and tight credit constraints, based on survey data for firms, are found to have significantly hindered both firms' and households' capital accumulation, again in the short run. This leads us to conclude that studies that disregard the role of debt or financing constraints are unable to fully explain investment dynamics in Italy, especially in the most recent years of sharp contraction.

JEL Classification: E22, G01, G31.
Keywords: gross fixed capital formation, institutional sectors, credit constraints.

Contents

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1 Introduction[1]

Following the outbreak of the global financial crisis, euro-area countries experienced a large fall in gross fixed capital formation, which was even sharper than that recorded in GDP. Even though it is on the road to recovery, investment activity remains weak relative to pre-crisis years, in spite of low financing costs stemming from the strongly expansionary stance of monetary policy in the euro area.

Motivated by these facts, in the past few years there has been a resurgence of research in the determinants of capital accumulation and the influence of financial factors on investment.[2] Recent macroeconomic analyses (Bacchini, Bontempi, Golinelli and Jona-Lasinio [2017]; Banerjee, Kearns and Lombardi [2015]; Barkbu et al. [2015]; Busetti, Giordano and Zevi [2016]; De Bonis, Infante and釘Paternò [2014]) have explored the determinants of investment dynamics in advanced economies, often concentrating on the period since the eruption of the global financial crisis.

This paper fits into this recent strand of the empirical literature by investigating the main factors influencing investment in Italy in the period 1995–2016, including financial determinants. Amongst the main euro-area countries, Italy is an interesting case-study since, after a pronounced downturn during the recent double recession, in 2016 the investment rate of the total economy was still over three percentage points below its pre-crisis average. Gauging the determinants of investment performance in this country is therefore an insightful exercise.

[1]We thank an anonymous referee, Giorgio Albareto, Marta Bańbura, Marco Bernardini, Lorenzo Braccini, Riccardo De Bonis, Silvia Fabiani, Giuseppe Grande, Luigi Infante, Alberto Locarno, Taneli Mäkinen, Marcello Pericoli, Luigi Federico Signorini, Ignazio Visco, Giordano Zevi, Roberta Zizza and Francesco Zollino for useful comments on previous drafts of this paper. We are grateful to Davide Fantino for sharing some data with us. This version of the paper was the one presented at the Conference "How financial systems work: evidence from financial accounts" held at Banca d'Italia in 2017; we thank our discussant Marcello Messori and all conference participants for their valuable comments. Any error remains however the responsibility of the authors. The views expressed herein are those of the authors and do not necessarily reflect those of the institution they represent.

[2]This topic has recently been singled out as one of the five knowledge gaps in economic research by the Federal Reserve: "More empirical work would be useful to disentangle the spending effects that result from changes in credit conditions from those that result from movements in interest rates [...]. Most empirical models [...] do not explicitly control for the influence of non interest credit factors on consumption and investment; as a result, estimated interest rate effects will partially reflect the influence of these factors to the extent these factors are correlated with interest rates" (Yellen [2016], p. 8).

Compared to the afore-mentioned articles, this paper differs at least in three respects, mainly concerning the data and the empirical methodology employed. First, while the mentioned studies – by employing national account data – have focused on aggregate investment or on either a breakdown by asset type (construction, machinery and equipment, etc.) or by economic sector (e.g., manufacturing, services, etc.), we instead investigate investment dynamics of both non-financial corporations (which we call interchangeably and loosely "firms" in the course of the paper) and households by employing the highly under-utilised institutional sector account data. This allows exploring the (potentially different) factors affecting investment dynamics of both Italian firms and households, where the latter, although accounting for a non-negligible share of overall investment, are often disregarded in the existing literature on capital accumulation.

Second, in order to assess the drivers of firms' and households' gross fixed capital formation, we estimate a vector error correction model that mimics the flexible neoclassical model of investment put forward by Hall and Jorgenson [1967]. Differently from single-equation models (e.g., Barkbu et al. [2015]), this multivariate framework allows taking into account the dynamic linkages among several endogenous variables (investment, output, user cost of capital), as well as any feedback among them, both for firms and households. Furthermore, we augment the vector error correction model with several determinants that are weakly exogenous relative to the cointegrating relationship, such as economic uncertainty and the business climate/consumer confidence, with the aim of better describing the short-run dynamics of investment. Amongst these short-term weakly exogenous variables, we also include financing constraints.

The latter aspect of the model is our third contribution to the existing literature, in which we attempt to consider financial determinants that could affect investment in Italy, other than the user cost of capital, which is most commonly included in standard investment equations. In this respect, we first exploit financial account, which provide information on liabilities of both firms and households to construct institutional sector-specific indicators of indebtedness. To our knowledge, few papers have employed the flow-of-funds data to examine the indebtedness-investment link for both firms and households.[3] Next, owing to the fact that liabilities in financial accounts are the result of the matching of demand and supply of external finance and therefore a potentially unsatisfactory, albeit common, measure of financing constraints, for firms (for which the necessary data are available) we are able to build a measure of actual credit restrictions using the Bank of Italy's Survey of Industrial and Service Firms, in an attempt to bridge the macro-micro gap in measuring financing frictions.

Our main findings are the following. For both firms and households we find a long-run equilibrium relationship between real investment, real value added/disposable income and the real user cost of capital, a result that is in line with the predictions of the flexible neoclassical model of investment. Amongst the weakly exogenous, short-run variables affecting capital accumulation, a rise in economic uncertainty is confirmed to significantly dampen firms' investment, as does a deteriorated business climate. Financing constraints too matter significantly in the short-run. Indeed, high levels of indebtedness are associated with significantly lower investment for both non-financial corporations and households. Moreover, for firms tighter credit conditions are also found to be negatively correlated with capital accumulation. Finally we show that, by accounting for financial variables, the unexplained component of Italy's investment dynamics during the recent slump can be reduced, especially for firms.

The structure of the paper is as follows. Section 2 examines the main developments in firms' and households' investment in Italy since 1995, based on national account data broken down by institutional sector. Section 3 describes the measurement and the evolution of the potential factors influencing investment of both firms and households. Based on the aggregate investment theory, we first assess developments in output and in the user cost

[3] See, for example, Jaeger [2003], which uses financial accounts for Germany and the U.S. for this purpose, yet only for firms, and employing a very simple empirical framework. More recently, see Ruscher and Wolff [2012] for a descriptive analysis of the consequences of (again only) corporate balance sheet adjustment.

of capital. Additional variables affecting investment in the short-run include uncertainty and the economic climate. We close the section by analysing the role of financial frictions, measured by a macro indebtedness indicator and by a micro proxy of credit constraints. Section 4 first describes the vector error correction specification, conducts *ex ante* weak exogeneity tests, displays unit root and cointegration test results, then discusses the main econometric results for both firms and households and finally draws some conclusions on the goodness-of-fit of our alternative investment model specifications. Section 5 concludes.

2 Developments in Italy's firm and household investment

Since 1995 and until the outbreak of the global financial crisis, the investment rate of the Italian economy, measured as the ratio of total gross fixed capital formation to GDP at current prices, was on average 20 per cent, thereby just under that recorded in the other two largest euro-area economies, France and Germany (21 per cent) and the average euro-area rate (22 per cent; Figure 1). Spain, on the other hand, recorded an impressive hike in its investment rate from already non-negligible levels in the same period, scoring an average rate of 26.5 per cent. This rise was however driven by the impressive boom in the Spanish housing sector, which the other main euro-area countries did not experience.[4]

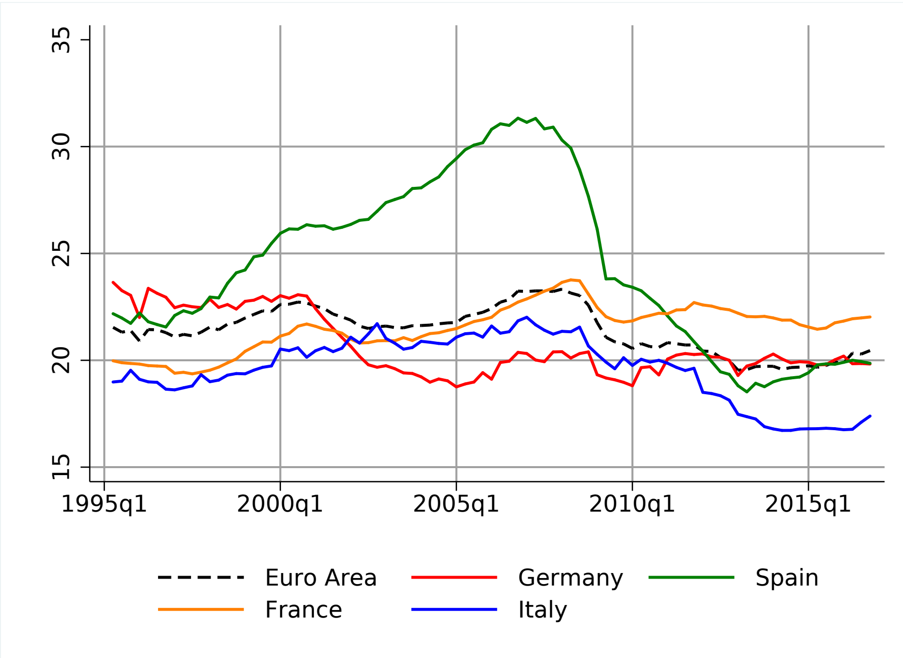
Excluding Spain, the drop in the investment rate recorded in Italy after mid–2008 was the most severe and the most persistent. Concerning the duration of the downturn, the beginning of the recovery was staggered across countries, with Italy recording a trough in its investment rate in mid–2014, against, for example, Germany's investment rate bottoming out already at the beginning of 2010. Regarding the severity of the slump, which was sharp also in a historical comparison (Busetti, Giordano and Zevi [2016]), in 2016, the last year for which annual data are currently available, Italy's investment rate stood at 17 per cent, whereas in the three other countries it was around or above 20 per cent. This implies an "investment gap" of the Italian economy relative to its pre-crisis average

[4] See Busetti, Giordano and Zevi [2016] for similar computations net of investment in construction, in which the pre-crisis investment rate in Spain, when excluding the construction item, is found to be lower than that in the other three main euro-area countries.

of 3.3 percentage points. Taking a long-run perspective, based on annual data published in Baffigi [2015], Italy's investment rate in recent years has been the lowest since 1950.[5]

Figure 1: Investment rates in the euro area and in the four largest euro-area economies

(ratio of nominal and seasonally adjusted total investment to GDP at market prices; percentage shares)



Source: Authors' calculations on Eurostat data.

Disaggregating investment dynamics helps in assessing these aggregate trends. The existing literature has mainly focused on investment by asset or by economic sector of activity. Thanks to the availability of national accounts by institutional sector, we instead break down investment rates into those of non-financial corporations and those of households in order to gain further insights. The latter include mainly "consumer" households, which invest in residential construction, but also "producer" households, i.e., small firms with up to five employees, and non-profit institutions serving households. Firms

[5]This broad picture is confirmed if one considers real, as opposed to, nominal investment rates. Differences in real versus nominal dynamics may, however, emerge when breaking data down by asset type, owing to differing investment price trends across assets. This issue, specifically for Italy, has been dealt with in Giordano, Marinucci and Silvestrini [2016] and [2017], to which we refer, using annual data.

account for nearly half of total investment expenditure in Italy, whereas total households contribute by more than a third, thereby together explaining the bulk of aggregate gross fixed capital formation in Italy (Table 1).

Table 1: Total gross fixed capital formation: breakdown by institutional sector

(percentage shares computed on current price series; annual data)

	Non-financial corporations (1)	Households (2)	General government (3)	Financial corporations (4)
1995–1999	49.5	34.6	14.4	1.6
2000–2007	50.2	34.7	13.7	1.6
2008–2016	49.7	34.9	14.1	1.5

Source: Authors' calculations on Istat data. Notes:

(1) Non-financial corporations include all private and public corporate enterprises that produce goods or provide non-financial services to the market.

(2) Households include "consumer" households, as well as "producer" households (i.e., household firms with up to five employees) and non-profit institutions serving households.

(3) General government includes central, regional and local government and social security funds.

(4) Financial corporations include both financial and insurance firms.

In this paper we do not consider the investment dynamics of financial corporations and of the general government, which are instead discussed in Giordano, Marinucci and Silvestrini [2016] and [2017]. This choice is dictated by the fact that public investment is often counter-cyclical and driven by different factors to private investment, and thus deserves a separate analysis. Moreover, financial firms' gross fixed capital formation is extremely volatile, warrants a specific analysis, but anyhow accounts for a negligible share of overall investment in Italy. This, however, implies that we lose out on any inter-sectoral linkages, especially between public and private investment.

By comparing quarterly developments of both firms' and households' investment rates since 1995,[6] the investment rate of households (about 7 per cent in the pre-crisis period) is

[6] Official quarterly Istat series are only available since 1999. However, for 1995–1998 we disaggregate

structurally lower than that of firms (10 per cent; Table 2). The pronounced expansion in gross capital formation until mid–2008 was of comparable magnitude across the two largest institutional sectors (Figure 2). The overall decline thereafter was also broadly similar, although the contraction in 2009 was sharper and the subsequent temporary recovery before the further fall associated with the sovereign debt crisis was steeper for firms relative to households.[7] Relative to the pre-crisis average, in 2016 firms' investment rate was 1.1 percentage points lower, against a comparable shortfall for households. Households' investment rates, in particular, are currently still at significantly lower levels than those registered in 1995 (when our time series begin), although, similarly to firms' rates, are now mildly on the rise. The more sluggish recovery of households' investment rates is plausibly connected to the less vibrant pickup of housing investment relative to business investment, which may be observed in national account data disaggregated by asset type (here not shown).

the annual sectoral data released by Istat into quarterly series, by applying standard statistical techniques (i.e., the original Chow-Lin procedure as proposed by Chow and Lin [1971] to disaggregate quarterly observations to monthly interpolations, adapted by Barbone, Bodo and Visco [1981] and subsequently by Abeysinghe and Lee [1998] to convert annual aggregates into quarterly values). In this way, our dataset covers the overall period 1995:Q1–2016:Q4.

[7] In particular, the temporary pick-up in investment after the global financial crisis was only experienced by firms. At the same time, the decline in household investment was particularly pronounced as of 2011.

Table 2: Investment rates: breakdown by institutional sector

(ratio of nominal total investment to GDP at market prices; percentage shares, unless otherwise specified)

	Investment rate 1995:Q1–2008:Q3 (1)	2008:Q4–2016:Q4	2016	Investment gap (2)
Non-financial corporations	10.0	9.1	8.9	-1.1
Households	6.9	6.3	5.7	-1.2

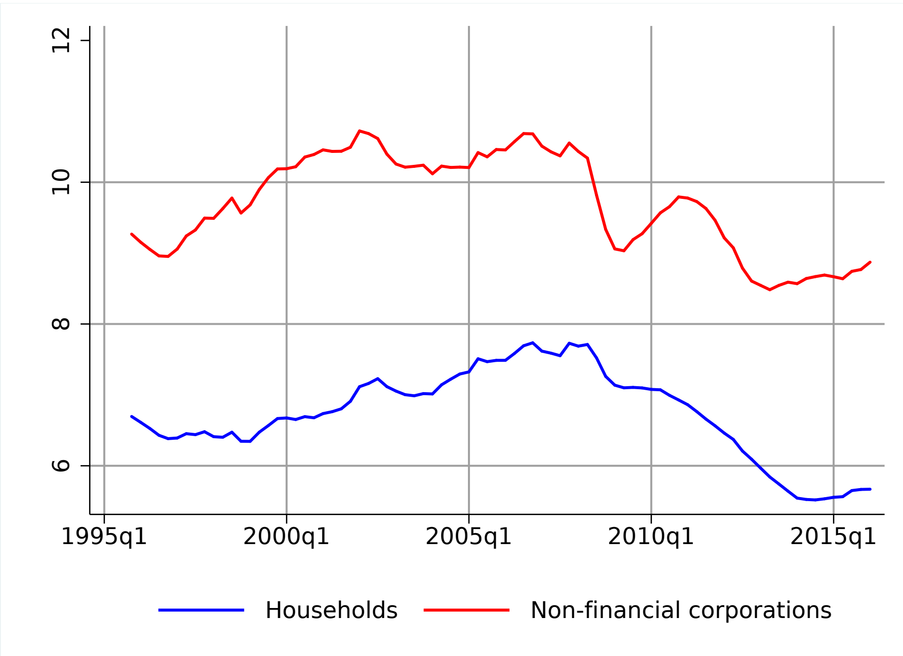
Source: Authors' calculations on Istat data. Notes:

(1) In 2008:Q3 the total economy investment rate peaked in Italy.

(2) The investment gap is computed as the difference in percentage points between the average investment rate in 2016 and that observed in the pre-crisis (1995:Q1–2008:Q3) period.

Figure 2: Non-financial corporations' and households' investment rates

(ratio of nominal investment to GDP at market prices; percentage shares)



Source: Authors' calculations on Istat data. Notes: The nominal investment series at the numerator is here smoothed by taking a 4-term moving average.

An assessment of the drivers of capital accumulation in Italy of both firms and households is therefore warranted in order to better understand the recent negative developments. Hence we next turn to discuss the possible variables affecting both institutional sectors' investment, suggested by economic theory. We then describe how we measure these variables and investigate their developments in Italy over the past two decades.

One issue with using macroeconomic data is that it is not possible to disentangle the intensive and extensive margin of investment or, in other terms, to assess whether the investment downturn in Italy was mainly driven by a reduction in the investment intensity of individual economic agents or to their exit from the market. Indeed, micro analyses focused on Italy have shown that the number of non-financial corporations (in particular, limited liability companies for which data are available) exiting the market soared after 2011 (De Socio and Michelangeli [2015]), possibly accounting for part of the investment drop. However, on the other hand, it has also been found that this crisis-driven firm turnover has been productivity-enhancing (Linarello and Petrella [2016]), possibly implying that the surviving firms were not only the most productive, but also those investing most, thereby limiting the negative effect of the fall in the extensive margin. The issue cannot be appropriately tackled with institutional sectoral account data, but we conduct some robustness analysis in our empirical analysis, which we discuss further on.

3 The analytical framework and the potential factors affecting investment in Italy

3.1 The flexible neoclassical model: the role of output and of the user cost of capital

One of the workhorse models of investment is the flexible neoclassical model as described in Hall and Jorgenson [1967]. The model's two main components are an equation for the desired level of capital stock, which is determined by real output and the user cost of capital, as well as an equation describing the time structure of the investment process.

The first equation is derived from the equilibrium condition according to which the value of the marginal product of capital is equal to its real rental price. Under the assumptions of a Cobb-Douglas production function, profit maximization and perfect competition, the desired level of capital K^* is equal to:

$$K^* = \alpha \frac{Y}{r}, \tag{1}$$

in which Y is real output, r is the real rental price (or real user cost) of capital and α is the elasticity of output with respect to capital. The higher the real rental value of capital and the lower the level of output, the lower is the desired capital stock.

The second component of the model rests on the assumption that investment depends on changes in the desired level of capital stock (K^*) in previous periods:

$$\begin{aligned} I_t &= \sum_{s=0}^{\infty} \mu_s (K^*_{t-s} - K^*_{t-s-1}) + \delta K_t \Rightarrow \\ I_t - \delta K_t &= N_t = \sum_{s=0}^{\infty} \mu_s \Delta K^*_{t-s}, \end{aligned} \tag{2}$$

where $\Delta = 1 - L$ is the first-difference operator, L is the lag operator, and where it is also assumed that replacement investment is proportional to existing capital stock (with a constant rate of replacement equal to δ). In equation (2), net investment is modelled as an infinite weighted average of past changes in the desired capital stock, where the weight μ_s measures the proportion of the change in desired capital at time $t - s$ that results in investment at time t.

The earliest flexible accelerator model that appears in the work of Koyck [1954] imposes the following distributed lag structure to the right-hand side of equation (2):

$$N_t = (1 - \lambda) \sum_{s=0}^{\infty} \lambda^s \Delta K^*_{t-s}, \tag{3}$$

with $\lambda < 1$, according to which investment at time t is a geometric average of all past changes in the desired stock capital stock. In order to estimate the parameters of the distributed lag function in equation (3), Hall and Jorgenson [1967] place restrictions on the infinite sequence of weights $\{\mu_s\}_{s=0}^{\infty}$. In particular, the first two weights are estimated as separate parameters, while successive weights decline geometrically, in the spirit of Koyck

[1954]. The resulting investment function takes the following form (Hall and Jorgenson [1967], p. 397):

$$\begin{aligned} N_t &= \gamma_0 \Delta K_t^* + \gamma_1 \Delta K_{t-1}^* - \omega N_{t-1} + \varepsilon_t \Rightarrow \\ (1-\omega L) N_t &= \alpha\gamma_0 \Delta \left(\frac{Y_t}{r_t}\right) + \alpha\gamma_1 \Delta \left(\frac{Y_{t-1}}{r_{t-1}}\right) + \varepsilon_t, \end{aligned} \tag{4}$$

in which α, γ_0, γ_1 and ω are unknown parameters to be estimated. Given these parameters, changes in real output and the user cost of capital affect investment spending.

In order to take this model to the data, investment is proxied by gross fixed capital formation available for firms and households at current prices from the institutional sector accounts. In the absence of official, disaggregated deflators in institutional sector accounts,[8] constant-price series are obtained by deflating the nominal expenditure of each institutional sector with the total non-housing investment deflator for firms and the residential investment deflator for households, respectively.

Y_t in equation (4) is proxied by value added for firms and by disposable income for households, both at current prices, taken from Istat's national accounts.[9] Firms' value added is deflated using the non-financial private economy value added deflator, while households' disposable is deflated by the consumer price index.[10]

Both series, and real disposable income in particular, increased significantly until the beginning of 2008 (Figure 3; left-hand side panel). Thereafter, the post-global financial crisis decline, again particularly pronounced for households, was interrupted by an albeit temporary recovery in 2010–2011. Both series started gradually picking up after 2013, to a greater extent for households.[11]

Finally, the real user cost of capital (r_t) can be proxied by the sum of the real interest rate and the depreciation rate of capital. The real interest rate is here computed as a

[8]See Giordano, Marinucci and Silvestrini [2016] and [2017] for a thorough investigation of the topic.

[9]Disposable income is included as a factor affecting the demand of households for investment goods, similarly to Nobili and Zollino [2012] and Loberto and Zollino [2016], which model investment in residential construction equations).

[10]Since they are not seasonally adjusted, both output series for each institutional sector are then smoothed with a one-sided 4-term moving average filter.

[11]One reason why households' real disposable income increased so pronouncedly in recent quarters is due to the very muted dynamics of the deflator used, the consumer price index.

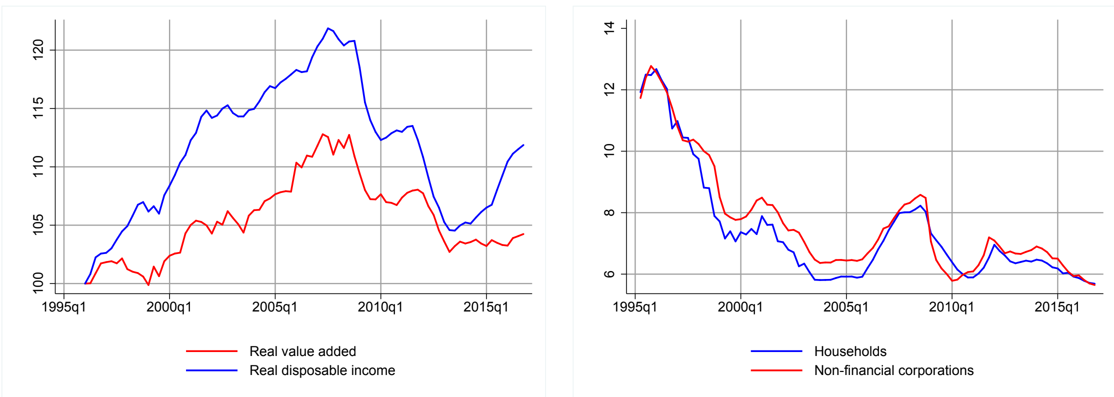
weighted average of the short and long-term nominal bank lending rates to either non-financial corporations or households (for house purchase), sourced from Banca d'Italia, net of one-year ahead inflation expectations provided by Consensus Economics. The depreciation rate too is calculated separately for firms and households, by dividing the amount of consumption of fixed capital of the non-financial private economy by the total gross capital stock.[12]

A downward trend in the real user cost of capital is apparent in the first half of the sample, linked to the accession to the euro area (Figure 3; right-hand side panel). Then, the real user cost of capital began to creep up as of the mid–2000s, reaching its peak in connection with the outbreak of the global financial crisis. The overall expansionary monetary policy in the euro area contributed to dampen real interest rates thereafter, with rates currently more or less at their lowest levels since 1995.

Figure 3: The neoclassical model determinants

(1996:Q1=100 in the left-hand side panel; percentage points in the right-hand side panel)

(a) Real output **(b) Real user cost of capital**



Source: Authors' calculations on Banca d'Italia, Consensus Economics and Istat data.

Beyond output dynamics and the user cost of capital, other drivers of capital accumulation have been singled out in the theoretical and empirical literature. These determinants

[12]The total amount of fixed capital is used for firms, only residential capital is employed for households.

are discussed in the following two sections.

3.2 The role of additional factors: economic uncertainty and confidence

The "real options" theory of investment (Dixit and Pindyck [1994]) suggests that economic uncertainty exerts a depressive effect on investment expenditure. Firms decide to invest in an irreversible investment project whenever the marginal product of capital exceeds a threshold value depending on the volatility of expected demand and on future returns to capital. Uncertainty about future economic outcomes increases this threshold value, resulting in firms postponing investment decisions until uncertainty has declined or new information has become available. Uncertainty can similarly influence decisions of households on housing purchases: high uncertainty can induce households to increase precautionary savings as an alternative to investment.

The empirical literature provides support for these theoretical predictions and advocates uncertainty as a relevant determinant to explain negative short-run capital fluctuations. Notably, Bloom, Bond and van Reenen [2007] show that uncertainty shocks delay investment in the UK manufacturing sector. Similarly, using survey data on the Italian manufacturing sector, Guiso and Parigi [1999] and Bontempi, Golinelli and Parigi [2010] find a negative relationship between investment and uncertainty, a result which was more recently confirmed by Busetti, Giordano and Zevi [2016] using national account data.

In the case of non-financial corporations, following the recent empirical literature, we proxy demand uncertainty with the cross-sectional dispersion in the subjective expectations of manufacturing firms interviewed in the monthly Istat Business Survey.[13] This measure is computed as:

$$unc_t = \sqrt{frac_t^+ + frac_t^- - (frac_t^+ - frac_t^-)^2}, \tag{5}$$

[13] This measure has been constructed and employed in Fuss and Vermeulen [2008], Bachmann, Elstner and Sims [2014], Busetti, Giordano and Zevi [2016] and Gamberoni, Giordano and Lopez-Garcia [2016]. Similarly to the first three studies we only consider expectations of manufacturing firms, since the resulting uncertainty measure is available for the longest time-span, it is satisfactorily representative of the private business economy in Italy and is the most reliable. The fourth mentioned study instead undertakes an economic sector analysis on input misallocation and therefore constructed measures of uncertainty also for distribution, construction and other services, for five different countries, albeit for a shorter timespan. Alternative proxies of uncertainty proposed in the literature are discussed in Appendix A.

where $frac_t^+$ and $frac_t^-$ are the fractions of firms with increase and decrease responses at time t, respectively. The survey questions we consider are those referring to future production and order expectations of firms relative to their current situation, taking quarterly averages of the mean of the two computed monthly dispersion measures.

With respect to households, we begin by computing a similar indicator using the question referring to expectations on consumers' personal situation, taken from the monthly Istat Consumer Survey. Since households also include micro-businesses, as mentioned in Section 2, we construct a weighted average of the firm and consumer indicators, where the weights are the share of households' investment accruing to producer and consumer households, respectively.[14] To our knowledge this is the first time a similar uncertainty measure for households has been computed.

Both firms' and households' survey-based uncertainty measures, which have been normalised for comparability reasons, are plotted in Figure 4 (left-hand side panel). They report striking peaks connected with the burst of the dot.com bubble for firms and then again during the global financial and sovereign debt crises for both firms and households (with the latter recording an even more pronounced hike in the second recessionary episode and the highest ever since 1995). For households uncertainty was particularly high in 2008 and then in 2011–2015, while for firms peaks in uncertainty were more unevenly distributed across time, including the quarters in the run-up to the adoption of the euro, the 2001 recession and then 2008–2009 and 2013.[15]

Finally, for both institutional sectors we also rely on a measure of economic sentiment that may affect their investment behaviour. Indeed, a weak current and prospective eco-

[14] According to annual data (the only disaggregated data available), "consumer" households' investment accounted for about 70 per cent of total household capital expenditure on average over the period considered.

[15] The recent reduction in firms' uncertainty has been driven by an increase in the share of enterprises expecting stable production and orders' growth, in line with the general improvement of the manufacturing business climate in Italy observed in recent quarters, shown in the right-hand side panel of Figure 4. The more general downward trend in firms' uncertainty, observed in the chart, could be due to the fact that there is evidence that, since the sovereign debt crisis, Italian firms have modelled their expectations according to a "new normal" of a generalised lower level of production (Conti and Rondinelli [2015]) and are less divergent across units in their replies.

nomic or personal climate are likely to deter firms and households from investing in new capital (Parigi and Siviero [2001]). Moreover, as this variable captures future expectations on the state of the economy and on economic agents' personal situation, its inclusion in an econometric model for investment is a means to (partially) tackle the claim that expected, as opposed to past, output growth better explains investment dynamics of firms (see, for example, Bussière, Ferrara and Milovich [2015]) and that permanent disposable income, as opposed to current income, better accounts for households' investment developments (see, for example, Duca, Muellbauer and Murphy [2012]).

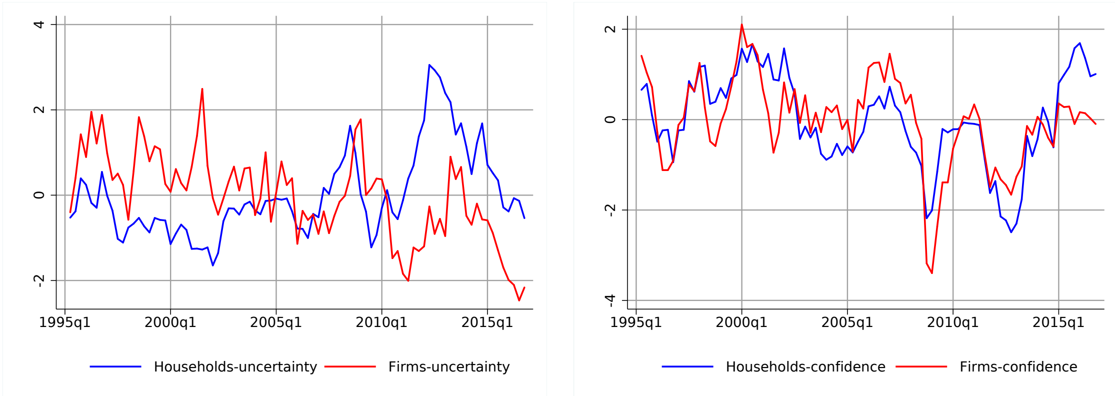
For non-financial corporations we consider an indicator of firms' business climate, sourced from the same manufacturing business survey used to construct the uncertainty indicator. As in Busetti, Giordano and Zevi [2016], this variable can be seen as capturing the "first moment" of firms' business outlook, whereas the uncertainty indicator captures the "second moment".[16] Similarly, for households we consider the consumer confidence indicator from the Istat consumer survey, also employed to construct our measure of uncertainty. Again, in order to take into account the presence of micro firms in the household institutional sector we construct a weighted average of both consumer and business confidence, as done for uncertainty. Whereas both firm and households' confidence levels were broadly stationary until the outbreak of the global financial crisis, these indicators then dropped dramatically after 2008, and then again during the sovereign debt crisis (Figure 4; right-hand side panel). Since then, they have been set on a broadly upward trend.

[16]Not necessarily periods of high uncertainty are connected to low confidence episodes; indeed, uncertainty can be high also in high-growth and confidence years, when returns to investment may be uncertain as entrepreneurs take more risks. It is noteworthy that, whereas in the case of households, uncertainty and confidence are highly negatively correlated (-0.75), in the case of firms the correlation is negligible (-0.20).

Figure 4: Firms' and consumers' uncertainty and confidence

(standardised dispersion measures in the left-hand side panel; standardised indices of confidence in the right-hand side panel; seasonally unadjusted data)

(a) Uncertainty **(b) Confidence**



Source: Authors' calculations on Istat data.

3.3 The role of additional factors: financing constraints

Neoclassical investment models, such as that described in Section 3.1, assume perfect information and competition in capital markets, in line with the influential work of Modigliani and Miller [1958]. The Modigliani-Miller theorem states that debt versus equity financing has no impact on the total market value of the firm, so that corporate financial policy is irrelevant for investment decisions. The latter should thus depend solely on the profitability of investment opportunities and/or on changes of the real user cost of capital. This explains why in many mainstream investment models real interest rates are the only transmission channel from the financial sector to real economic activity.

However, since the contribution by Stiglitz and Weiss [1981], it has been argued that the availability of external finance and not just the interest rate charged matters for business-cycle fluctuations. When financial frictions arise due to asymmetric information between firms and lenders (Tirole [2006]), credit rationing may occur, defined as in Stiglitz and Weiss [1981] as the case in which economic agents "*would not receive a loan even if they offered to pay a higher interest rate*" (p. 395). A complementary way of measuring financing constraints is, as discussed in Farre-Mensa and Ljungqvist [2016], by the size of the wedge between the cost of new debt and equity (i.e., funds raised externally) and the opportunity cost of internal finance generated through cash flows and retained earnings (Fazzari, Hubbard and Peterson [1988]). The cost of external finance may be higher than that of internal finance due to the existence of an "external finance premium" (Bernanke and Gertler [1989]), which arises when lenders incur a cost in order to monitor the entrepreneur's performance. Taking these considerations into account, a "financial accelerator" model points to credit developments propagating and amplifying exogenous shocks to the real economy (Bernanke, Gertler and Gilchrist [1999]).[17]

A related strand of the literature has examined the effect of a debt overhang, resulting from excessive leverage growth, on output dynamics and on investment spending. Increasing debt holdings raises default probabilities, in turn leading to financial distress, which is reflected in higher external financing premia or credit rationing. While low-leveraged firms face low financing constraints and can therefore access credit if profitable investment opportunities arise, highly indebted firms are more concerned about default risks and their financial status, thereby potentially giving up valuable investment opportunities when internal sources of funds are not sufficient (Myers [1977]), especially in times of financial turbulence (Bernanke, Gertler and Gilchrist [1999]). Indeed, during expansions households and firms may find it optimal to borrow less than their credit limit, whereas during recessions financial constraints become binding. In this framework, Occhino and Pescatori [2015] point to the existence of a "debt-overhang distortion", which reduces the benefit that firms achieve by investing and leads them to invest less than would be optimal if they had fewer liabilities. Indeed, when the burden of outstanding debt grows beyond

[17] As explained by Bernanke, Gertler and Gilchrist [1999], in the presence of credit frictions and with credit demand held constant, standard models of lending with asymmetric information imply that the external finance premium depends inversely on borrowers' net worth (liquid assets plus the collateral value of illiquid assets): when borrowers have little wealth for investment, the potential divergence of interests between the borrower and the suppliers of external funds is greater, implying heightened agency costs. To the extent that borrowers' net worth is procyclical (because of the procyclicality of profits and asset prices, for example), the external finance premium will be countercyclical, enhancing the swings in borrowing and thus in investment and economic activity in general.

a certain limit, a firm's risk of default increases. In the event of default, the benefit from investment entirely accrues to creditors. This on average lowers the marginal return of new investment and reduces the incentive to invest for borrowers.

On the empirical side, the literature has examined the role of leverage and of financial constraints in explaining investment, mainly relying on firm-level data,[18] but also from a macroeconomic perspective. For example, Barkbu et al. [2015] analyse gross fixed capital formation in selected euro-area countries, finding a large effect of demand expectations on investment, which is, however, compounded by uncertainty, financial constraints and corporate leverage. Likewise, focusing on Italy, Busetti, Giordano and Zevi [2016] show that credit supply restrictions accounted for about one third of the fall in non-construction investment of the private business sector during the periods 2008–2009 and 2012–2013. The interaction between the debt level and business-cycle dynamics is not limited to the corporate sector. A large empirical literature has shown that high leverage at the household level may also hamper macroeconomic performance. Notably, Mian, Sufi and Verner [2015] document that an increase in household leverage predicts lower output growth in a panel of 30 countries. Moreover, other studies (for example, Poterba [1984] and Goodwin [1986], based on US data), find a negative impact of credit rationing specifically on housing investment.

The existence of a link between level of indebtedness, financing constraints and the real economy motivates us to assess whether these financial variables have also played a significant role in explaining Italy's investment dynamics since 1995. The main issue we face when addressing this topic, as does the existing literature, is how to properly measure financing constraints. The credit supply curve is not readily observable to the econometrician nor is the opportunity cost of internal funds (underlying the wedge definition) easy to estimate, as discussed in Farre-Mensa and Ljungqvist [2016].

For this purpose, we exploit financial account data, sourced from Banca d'Italia,[19]

[18] See Hubbard [1998] for a general survey and Gaiotti [2013], Bond, Rodano and Serrano-Velarde [2015] and Cingano, Manaresi and Sette [2016] for recent evidence specifically on Italy.

[19] Financial accounts or flow-of-funds data are national statistics that provide a unified view of stocks

which allow constructing an indirect measure of financing constraints for both institutional sectors, namely their degree of indebtedness. This proxy is indirect for two reasons (Johnson and Li [2010]; Ferrando and Mulier [2013]; Barrero, Bloom and Wright [2017]). First, it is a forward-looking proxy of financing constraints, in the sense that high indebtedness indicates the ability to borrow in the past, but also that an economic agent is close to its credit limit and may not be able to borrow much in the future. Second, it captures both the demand and supply side of the availability of external finance, in that it is less desirable, but also more difficult and costly, for a highly indebted agent to be granted new debt.[20] Specifically, we consider firms' debt-to-GDP ratio and the household debt-to-disposable income ratio (Figure 5).[21] For firms, debt corresponds to the stock of short and long-term loans received[22] and of securities issued by non-financial corporations. For households, debt refers solely to short and long-term loans,[23] insofar as households cannot issue debt securities. Figure 5 clearly points to a substantial increase in indebtedness, both for Italian firms and households, in the years leading up to the global financial crisis and until 2010. The spike in firms' debt–to–GDP ratio in 2009 was due to the exceptional crash in GDP in that years (see, for example, De Socio and Finaldi Russo [2016] for a decomposition of the two components underlying the ratio). Some deleveraging is instead visible from 2012 onwards, albeit to a different extent across the two institutional

and flows of assets and liabilities, classified by institutional sector and by financial instrument, where the latter is also broken down by maturity at issue. As such, these accounts provide information on loans extended to both firms and households by all other institutional sectors (mainly banks, but also other financial institutions, the general government and the rest of the world) and on debt securities and equities issued by firms.

[20]However, using cross-country firm survey and balance-sheet data, Ferrando and Mulier [2013] find that firms with higher leverage are more likely to perceive access to finance as their most pressing problem, as well as to face actual credit constraints. The latter are measured by a categorical variable based on the replies to the questions in the Survey on the Access to Finance of small and medium-sized Enterprises on whether firms had applied or not for a bank loan, whether they were successful in getting any type of financing, and what was the reason not to have applied for external finance.

[21]These measures have been used in the macroeconomic literature as a standard metric for assessing the sustainability of debt (see, for example, Buttiglione et al. [2014]).

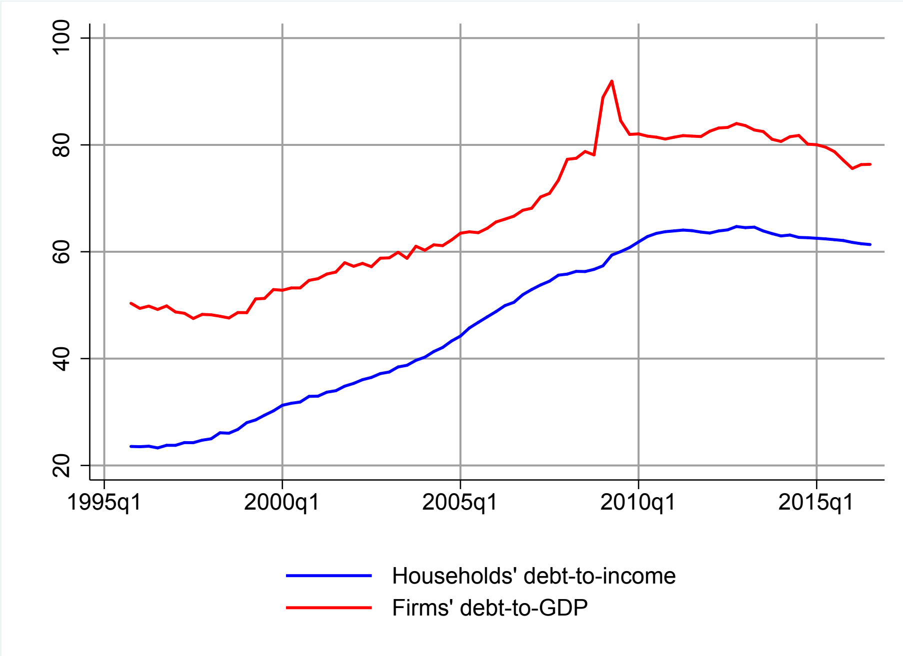
[22]It is worth mentioning that the loans obtained by firms are mainly granted by the banking system: in 2016 bank loans accounted for 70 cent of total loans granted to non-financial corporations. This explains why a measure of indebtedness based solely on bank (as opposed to total) loans, available upon request, records very similar dynamics to that reported herein.

[23]Loans obtained by households are mainly granted by the banking system: in 2016 bank loans accounted for 90 per cent of total loans.

sectors.[24]

Figure 5: Non-financial corporations' and households' indebtedness

(percentage points)



Source: Authors' calculations on Istat and Banca d'Italia data.

An issue with these indebtedness measures is that they do not specifically identify frictions in the credit market that may hamper investment decisions. From a macro perspective indeed, credit aggregates do not convey enough information to disentangle supply and demand (Bernanke and Gertler [1989]). Given these limitations, the literature has tried to improve identification using micro data. Following Guiso and Parigi [1999] and Gaiotti [2013], we exploit direct information on the limits of credit availability taken from a firm survey.[25] Credit access conditions refer to the entire pool of possible borrowers

[24]Similar developments are observed for firms if a financial leverage measure such as the ratio of debt to its sum with equity (suggested by Rajan and Zingales [1995], for example) is employed.

[25]An alternative would have been to construct a similar measure based on bank surveys. Indeed, Busetti, Giordano and Zevi [2016] construct a synthetic indicator of credit rationing based on the annual Bank Lending Survey for Italy. In particular, they use the index of supply tightening for firms, that is a weighted average of the values assigned to the qualitative answers obtained from the banks involved in the survey as follows: -1 = tightened considerably, -0.5 = tightened somewhat, 0 = basically unchanged, 0.5

and therefore these data are not subject to sample bias as is the case for macroeconomic data (and, in particular, our financial account data) on granted loans.

In particular, we construct an indicator of borrowing constraints of both industrial (net of construction) and service firms based on data extracted from Banca d'Italia's Survey of Industrial and Service Firms (SISF).[26] This variable is defined as the share of firms that were unable to obtain external finance from banks or other financial institutions out of all firms participating in the survey.[27] Firms are "financially constrained" when either their loan request is (partially or totally) refused by the bank or the loan conditions are deemed to be excessive by the firm and therefore the loan is not extended.

Figure 6 points to a progressive increase in the share of credit-rationed firms until the sovereign debt crisis (from around 2 to about 14 per cent, in line with figures reported in Gaiotti [2013] and in Bond, Rodano and Serrano-Velarde [2015]) and a loosening of credit constraints thereafter.[28] The chart also confirms that firms' indebtedness records dynamics that are, on the whole, similar to those of the financing constraint variable obtained from survey data, suggesting a high correlation between the two proxies of financing constraints.[29]

= eased somewhat, 1 = eased considerably. This indicator is, however, available only since 2003 and could not therefore be used in our analysis. For the years for which it is available, Gaiotti [2013] anyhow finds a strong correlation between the BLS indicator of credit constraints and the firm survey-based indicator we discuss later on.

[26] The SISF collects information on an annual basis from a panel of Italian firms with more than 50 employees operating in both industry net of construction and services since 1972 and with more than 20 employees since 2002. In the period of analysis in our paper, the number of participating firms significantly increased, passing from around one thousand firms in 1995 to almost 4,200 in 2016, thereby increasing the representativeness of this survey. As discussed in Gaiotti [2013], the SISF sample of firms tends to over-represent large firms in that it excludes firms with less than 20 employees. However, the under-representation of small firms, if anything, introduces a bias against finding a significant relationship between credit constraints and investment, since the former constraints are generally more binding for smaller enterprises.

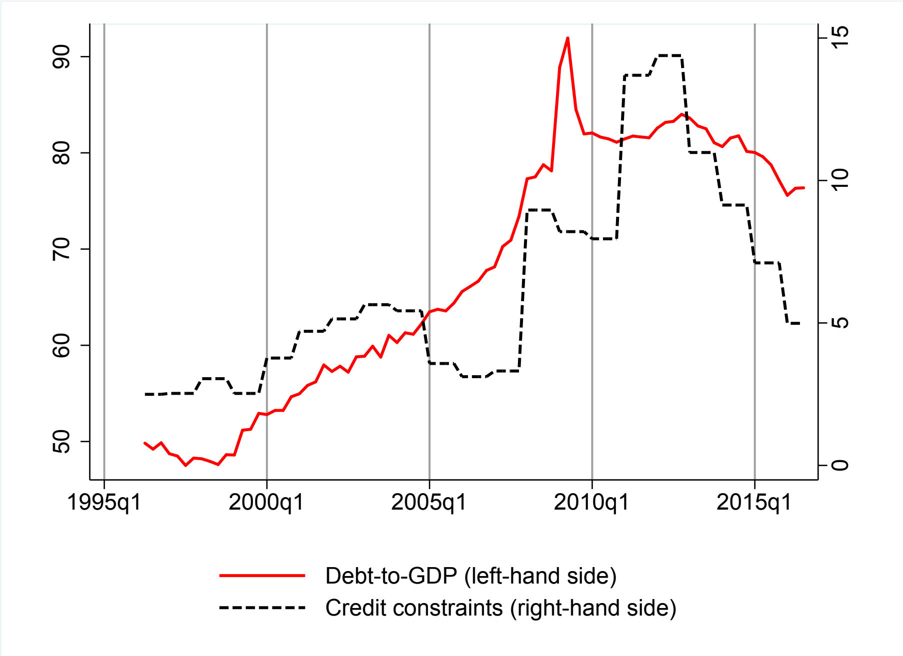
[27] From the denominator we exclude the firms that did not give any answer to questions in the financing section of the survey. However, even when we include them, the indicator constructed presents very similar dynamics.

[28] As mentioned earlier, this recent development may also be due to the market selection process enacted during the crisis years, where surviving firms were the most productive and less indebted than exiting firms.

[29] As for any survey data, self-reporting and the lack of incentive for respondents to reply truthfully can undermine the reliability of the answers given. However, this documented correlation helps ensuring the reliability of the information available and the overall credibility of the SISF.

Figure 6: Non-financial corporations' indebtedness and financing constraints

(percentage points)



Source: Authors' calculations on Istat and Banca d'Italia data.

Unfortunately, differently to firms, it is not possible to construct an actual financing constraint variable for households. The Banca d'Italia Survey on Household Income and Wealth (SHIW) is indeed biannual, presenting too little time variation for our quarterly analysis. At the same time, information from Istat's Consumer survey (opinions on a household's financial situation) is found to be a very imperfect proxy of financing constraints. Finally, it has been found that the estimated correlation between the Bank Lending Survey indicators on loan supply to households, anyhow available only as of 2003, and actual credit market developments is much weaker than that for firms (Del Giovane, Nobili and Signoretti [2013]), suggesting that this measure may not be a satisfactory proxy of credit constraints for households.

4 An empirical model of firm and household investment for Italy

4.1 The econometric specification

In this section we sketch an empirical investment model consistent with the theory outlined in Section 3.1. We adopt a multivariate framework that is particularly suitable to examine the dynamic linkages among the variables in equation (4), as well as the feedback effects among them, which single-equation models cannot capture. To this aim, we define an unrestricted vector autoregression model (VAR) of order p as:

$$\boldsymbol{A}(L)\boldsymbol{y}_t \quad = \quad \boldsymbol{C}\boldsymbol{D}_t + \boldsymbol{\varepsilon}_t, \quad t = 1, \ldots, T, \tag{6}$$

where $\boldsymbol{y}_t = (y_{1t}, \ldots, y_{nt})'$ is a vector of n endogenous I(1) variables. Furthermore, $\boldsymbol{C}$ is a matrix of coefficients for the deterministic variables in $\boldsymbol{D}_t$ (such as, for instance, a constant, time dummies, etc.), while $\boldsymbol{A}(L) = (\boldsymbol{I} - \boldsymbol{A}_1 L - \ldots - \boldsymbol{A}_p L^p)$ is a matrix polynomial in the lag operator L. Finally, $\boldsymbol{\varepsilon}_t$ is a vector of Gaussian white noise stochastic errors.

The VAR in equation (6) can be conveniently represented in the form of a "vector error correction" (VECM) model:

$$\Delta\boldsymbol{y}_t \quad = \quad \boldsymbol{C}\boldsymbol{D}_t + \boldsymbol{\Pi}\boldsymbol{y}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta\boldsymbol{y}_{t-i} + \boldsymbol{\varepsilon}_t, \quad t = 1, \ldots, T-1, \tag{7}$$

where $\boldsymbol{\Pi} = -\boldsymbol{A}(1) = -\left(\boldsymbol{I}_n - \sum_{i=1}^{p} \boldsymbol{A}_i\right)$ and $\boldsymbol{\Gamma}_i = -\sum_{j=i+1}^{p} \boldsymbol{A}_j$, for $(i = 1, \ldots, p-1)$.

In equation (7), whenever $\boldsymbol{\Pi}$ has reduced rank ρ with $0 < \rho < n$, there are ρ linearly independent combinations of the n variables comprised in $\boldsymbol{y}_t$ that are stationary (cointegrating relations), with $n - \rho$ common stochastic trends (Johansen, [1995]).[30] Then, it is possible to decompose $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are both $n \times \rho$ matrices (with full column rank ρ):

$$\Delta\boldsymbol{y}_t \quad = \quad \boldsymbol{C}\boldsymbol{D}_t + \boldsymbol{\alpha}\boldsymbol{\beta}'\boldsymbol{y}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta\boldsymbol{y}_{t-i} + \boldsymbol{\varepsilon}_t. \tag{8}$$

[30] The advantage of this representation is that the hypothesis of cointegration can be formulated entirely as a restriction on the matrix Π, leaving the other parameters unrestricted.

Empirically, the rank ρ can be determined by applying tests on the maximum eigenvalue or trace statistics.

In equation (8), $\boldsymbol{\beta}'\boldsymbol{y}_{t-1}$ is the vector of the long-run level cointegrating relationships among the variables, $\boldsymbol{\beta}$ is the matrix of cointegrating parameters and $\boldsymbol{\Gamma}_i$ are the matrices of parameters which account for short-run dynamics, based on the lagged changes in the endogenous variables. The elements of the $\boldsymbol{\alpha}$ matrix are loading factors measuring the speed of adjustment towards the long-run equilibrium relationships among the variables in levels. The larger their absolute value, the greater the response of the endogenous variables to the previous period's deviation from the long-run equilibrium, and the quicker the adjustment of the system towards the target. Conversely, if a loading is zero, then there is no adjustment back to equilibrium and the corresponding variable is defined as "weakly exogenous" (Engle, Hendry and Richard, [1983]), implying that it does not react to disequilibrium errors but may still react to lagged changes of the endogenous variables. In the presence of weak exogeneity the cointegrating vector does not enter the equation determining the i-th variable: in formal terms, if $\alpha_{ij} = 0$ $(i = 1, \ldots, n;\ j = 1, \ldots, \rho)$, then variable i is weakly exogenous with respect to the long-run parameters of interest in vector j. This definition of weak exogeneity matters essentially with the efficiency of estimation. A variable is indeed weakly exogenous for the purpose of estimating the parameters of interest if it entails no loss of information to confine ones attention to the conditional distribution of the endogenous variables given the weakly exogenous variable and to ignore the marginal distribution of the latter.

The parameterisation of $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$ implies identification issues on both $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$. These matrices are not identified since $\boldsymbol{\Pi}$ can be factorised as $(\boldsymbol{\alpha}\boldsymbol{H}^{-1})(\boldsymbol{H}\boldsymbol{\beta}') = \hat{\boldsymbol{\alpha}}\hat{\boldsymbol{\beta}}'$, choosing any full rank $\boldsymbol{H}$ matrix (of dimension $\rho \times \rho$). In general, ρ^2 restrictions for identification have to be imposed on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$. A solution is to set exact-identifying restrictions on $\boldsymbol{\beta}$, leaving $\boldsymbol{\alpha}$ unrestricted. For instance, following Johansen's [1995] identification scheme, we set:

$$\boldsymbol{\beta}^{'} = (\boldsymbol{I}_\rho \quad \tilde{\boldsymbol{\beta}}^{'}), \tag{9}$$

where $\tilde{\boldsymbol{\beta}}$ is a $(n-\rho)\times\rho$ matrix of identified cointegrated vectors and $\boldsymbol{I}_\rho$ is the $\rho\times\rho$ identity matrix.

4.2 Model formulation and testing for weak exogeneity

The model presented in Section 3.1 includes three variables, namely: 1) real investment, 2) real output (real value added for firms and real disposable income for households) and 3) the real user cost of capital. Furthermore, as we have seen in Sections 3.2 and 3.3, the empirical literature also suggests 4) uncertainty, 5) business climate (firms) or consumer confidence (households) and 6) financing constraints, however measured, as relevant determinants of capital accumulation. The variables 1) and in 2) are expressed in logs, while all the others are not transformed, given that they represent shares or percentages. Our dataset is quarterly and covers the 1995–2016 period.

We first set up a multivariate model containing all six variables. Admittedly, this model is over-parameterised, leading to declining precision in the estimates and, most likely, to a large number of insignificant parameters. As a consequence, our aim is to formulate and test particular restrictions on the data generating process allowing us to determine which variables can be considered as "weakly exogenous", and hence can be treated as conditioning variables. Ultimately, this modelling procedure is designed to deliver a more parsimonious model for performing statistical inference and estimation.

Starting with firms, the lag order of the VAR is selected using standard information criteria. We estimate a VAR(2) model for the six-variable system, using indebtedness as a measure of financial constraints. This model can be mapped into a VECM(1) representation with unrestricted intercepts and restricted trend coefficients in the cointegrating relations, as some of the variables appear to contain trends. Both trace and max-eigenvalue test statistics point to one cointegrating relation at conventional statistical levels.[31] We

[31] All estimation results in this section are available from the authors upon request.

next estimate the model parameters by maximum likelihood. All the parameters in the cointegrating equation are statistically significant, whereas some of the speed of adjustment coefficients (specifically, on uncertainty, business climate and indebtedness) are not. Therefore, we test for the null hypothesis of weak exogeneity with respect to the $\boldsymbol{\beta}$ vector along the lines described in Johansen [1992], Boswijk [1995] and recently applied by Bacchini, Bontempi, Golinelli and Jona-Lasinio [2017]. The test is first conducted for each of the six variables separately. In this case the test is a likelihood ratio statistics which, with a single restriction on $\boldsymbol{\alpha}$, is asymptotically distributed as a Chi-square with one degree of freedom. The null hypothesis of weak exogeneity is rejected at standard levels of confidence for investment, output and the user cost of capital, whereas it cannot be rejected for the other three variables (Table 3).

Table 3: Firms – Likelihood ratio statistics for testing weak exogeneity of each of the six variables (1)

	$\alpha_{1,1} = 0$	$\alpha_{2,1} = 0$	$\alpha_{3,1} = 0$	$\alpha_{4,1} = 0$	$\alpha_{5,1} = 0$	$\alpha_{6,1} = 0$
Chi-square(1)	4.3977	16.3545	15.9951	0.1315	0.3561	2.3685
P-value	0.0360	0.0001	0.0001	0.7169	0.5507	0.1238

Notes: The system includes real investment, real output, the real user cost of capital, uncertainty, the business climate and indebtedness. The asymptotic distribution is Chi-square(1), for which the 95% percentile is 3.8415.

We further test the hypothesis that uncertainty, the business climate and indebtedness are jointly weakly exogenous with respect to $\boldsymbol{\beta}$. This test can be performed by imposing multiple restrictions on $\boldsymbol{\alpha}$, such that $\alpha_{4,1} = \alpha_{5,1} = \alpha_{6,1} = 0$. In this case the likelihood ratio statistics is asymptotically distributed as a Chi-square with three degrees of freedom. As shown in Table 4, the test statistics is not significant, confirming that uncertainty, the business climate and indebtedness are (jointly) weakly exogenous with respect to the long-run cointegrating vector.

Table 4: Firms – Likelihood ratio statistics for testing joint weak exogeneity of uncertainty, business climate and indebtedness (1)

	$\alpha_{4,1} = \alpha_{5,1} = \alpha_{6,1} = 0$
Chi-square(3)	4.2399
P-value	0.2367

Notes: The system includes real investment, real output, the real user cost of capital, uncertainty, the business climate and indebtedness. The asymptotic distribution is chi-square(3), for which the 95% percentile is 7.8147.

We then estimate a VAR model for the same six-variable system, but with the survey-based credit constraint variable instead of indebtedness. In this set-up standard information criteria suggest a VAR(1) representation, which is mapped into a VECM(0) model with a VECM(0) model with unrestricted intercepts and restricted trend coefficients in the cointegrating relations. Both trace and max-eigenvalue test statistics indicate one cointegrating relation at the 5% confidence level. Conditioning on a single cointegrating relationship, we estimate the parameters by maximum likelihood. All the long-run parameters are significant. Concerning the loading factors, they are all statistically different from zero except those associated with uncertainty and credit constraints. We then test for the null hypothesis of weak exogeneity for each of the six variables separately (Table 5). The null hypothesis of weak exogeneity is rejected at standard levels of confidence for investment, output and the user cost of capital, whereas it cannot be rejected for uncertainty and for our proxy of credit constraints. In this case, results are less clear-cut for business confidence in the sense that the corresponding α coefficient is significantly different from zero at the 5% confidence level (but not at the 1% level).

Table 5: Firms – Likelihood ratio statistics for testing weak exogeneity of each of the six variables (2)

	$\alpha_{1,1}=0$	$\alpha_{2,1}=0$	$\alpha_{3,1}=0$	$\alpha_{4,1}=0$	$\alpha_{5,1}=0$	$\alpha_{6,1}=0$
Chi-square(1)	58.2430	65.7809	26.4882	0.0096	6.3292	1.6734
P-value	0.0000	0.0000	0.0000	0.9220	0.0119	0.1958

Notes: The system includes real investment, output, user cost of capital, uncertainty, business climate, financing constraints. The asymptotic distribution is Chi-square(1), for which the 95% percentile is 3.8415.

We again also test for joint exogeneity restrictions (Table 6). The null hypothesis of weak exogeneity is jointly rejected at the 5% confidence level, due to the role of the business climate. However, a joint restriction according to which uncertainty and financing constraints are weakly exogenous, i.e., $\alpha_{4,1} = \alpha_{6,1} = 0$, is supported by the data (test statistics equal to 1.6773, with a p-value equal to 0.4323).

Table 6: Firms – Likelihood ratio statistics for testing joint weak exogeneity of uncertainty, the business climate and leverage (2)

	$\alpha_{4,1}=\alpha_{5,1}=\alpha_{6,1}=0$
Chi-square(3)	9.0772
P-value	0.0283

Notes: The system includes real investment, real output, the real user cost of capital, uncertainty, the business climate and credit constraints. The asymptotic distribution is chi-square(3), for which the 95% percentile is 7.8147.

Thus, based on the above, inconcluding, results, we consider a smaller-scale four-variable system featuring real investment, real output, the real user cost of capital and the business climate. We select a VAR(2) specification based on information criteria. When performing the cointegration test, we assume a VECM(1) model with unrestricted intercepts and restricted trend coefficients in the cointegrating relations. Both trace and max-eigenvalue test statistics indicate a single cointegrating relationship. The long-run parameters are all significant, as well as the speed of adjustment coefficients (barring that associated with business climate).

We then test for weak exogeneity separately for each of the four variables (Table 7). The null hypothesis of weak exogeneity is strongly rejected for investment, output, while it is rejected at the 10% level for the user cost of capital; conversely, the weak exogeneity hypothesis is clearly not rejected for the business climate. Therefore, according to this evidence, in our final empirical framework we treat business climate as a weakly exogenous variable, just like uncertainty and financing constraints. Investment, output and the user cost of capital are instead considered endogenous variables.

Table 7: Firms – Likelihood ratio statistics for testing weak exogeneity of each of four selected variables (3)

	$\alpha_{1,1} = 0$	$\alpha_{2,1} = 0$	$\alpha_{3,1} = 0$	$\alpha_{4,1} = 0$
Chi-square(1)	17.3330	26.5949	3.0816	0.3635
P-value	0.0000	0.0000	0.0792	0.5465

Notes: The system includes real investment, real output, the real user cost of capital and the business climate. The asymptotic distribution is Chi-square(1), for which the 95% percentile is 3.8415.

Finally, we consider households and repeat the same procedure as above. We select a VAR(2) specification for the six-variable system, which we map into a VECM(1) model with unrestricted intercepts and restricted trend coefficients in the cointegrating relations. Both trace and max-eigenvalue test statistics suggest at least one cointegrating relation at conventional statistical levels. After having identified a single cointegrating vector, we estimate the model by maximum likelihood. All the parameters in the cointegrating vector are statistically significant, except for the coefficient on consumer confidence. Moreover, the coefficient on the user cost of capital is marginally not significant. The speed of adjustment coefficients are all statistically different from zero at the 10% level, except for uncertainty.

We then test for weak exogeneity with respect to the cointegrating vector, first separately for each of the six sector-specific variables (Table 8). For households, too, we find that uncertainty, consumer confidence and indebtedness are weakly exogenous with respect to $\boldsymbol{\beta}$, whereas the null hypothesis of weak exogeneity is rejected for investment, disposable income and the user cost of capital.

Table 8: Households – Likelihood ratio statistics for testing weak exogeneity of each of the six variables

	$\alpha_{1,1}=0$	$\alpha_{2,1}=0$	$\alpha_{3,1}=0$	$\alpha_{4,1}=0$	$\alpha_{5,1}=0$	$\alpha_{6,1}=0$
Chi-square(1)	4.3367	3.3405	10.7442	2.1687	1.0671	2.6762
P-value	0.0373	0.0676	0.0010	0.1408	0.3016	0.1019

Notes: The system includes real investment, real output, the real user cost of capital, uncertainty, consumer confidence and indebtedness. The asymptotic distribution is Chi-square(1), for which the 95% percentile is 3.8415.

Lastly, we test the joint hypothesis that uncertainty, consumer confidence and indebtedness are jointly weakly exogenous (Table 9). The test statistics is not jointly rejected, hence also according to this joint test the weak exogeneity zero restriction is supported by the data.[32]

[32] If, on top of these weakly exogeneity restrictions, we place an additional zero restriction on the long-run parameter associated to consumer confidence, results are confirmed. Specifically, the joint restriction is not rejected (the corresponding test statistics is equal to 4.8727, with a p-value of 0.3006).

Table 9: Households – Likelihood ratio statistics for testing joint weak exogeneity of uncertainty, consumer confidence and indebtedness

	$\alpha_{4,1} = \alpha_{5,1} = \alpha_{6,1} = 0$
Chi-square(3)	4.3754
P-value	0.2237

Notes: The system includes real investment, real output, the real user cost of capital, uncertainty, consumer confidence and indebtedness. The asymptotic distribution is chi-square(3), for which the 95% percentile is 7.8147.

Overall, these findings enable us to reduce the parameter space of the VAR model in equation (6), which therefore includes only three endogenous variables (investment, output and the user cost of capital), as well as three additional exogenous variables (contained in the z_t vector), which do not affect long-run dynamics:

$$z_t = (Climate_t, Uncertainty_t, Fin_t)', \tag{10}$$

where Fin_t is either a measure of indebtedness or of credit constraints, according to the case. Clearly, this choice allows decreasing the number of parameters to estimate and increasing the precision surrounding the estimates. This is of paramount importance, especially when working with a relatively short dataset, as in the case under analysis.

In the following section we maintain the hypothesis of a three-variable VAR/VECM model with exogenous variables and we proceed to examine more closely the presence of cointegration among investment, output and the user cost of capital.

4.3 Unit root tests and cointegration analysis

In a first step, we implement a unit root pretesting of the variables to determine whether investment, value added or disposable income and the user cost of capital are I(1), as usually done prior to testing for cointegration in a VAR framework. Then, in a second step, we focus on VAR lag order selection and on the cointegration analysis.

We examine the order of integration of the series by applying standard unit root tests. The null hypothesis of the Augmented Dickey-Fuller (ADF) test is the presence of a unit root (Dickey and Fuller [1979]). To select the appropriate number of lagged first-difference terms to add (k), we apply the recursive procedure proposed by Ng and Perron [1995].[33] Regressions are run with an intercept, with both an intercept and a linear time trend, and with neither of them. Most of the times, the time trend is found to be statistically significant. Therefore, we specify the test regression with a constant and a linear time trend. The ADF test is often criticised for its low power in rejecting the null hypothesis, especially when the sample size is small. To overcome this problem, unit root tests are often complemented with stationarity tests, such as the KPSS test (Kwiatkowski, Phillips, Schmidt and Shin [1992]).

The outcomes of the ADF and KPSS tests for real investment, real value added and the real user cost of capital of firms and households are presented in Appendix B.[34] ADF test results imply acceptance of the presence of a unit root for the series under study at the 5% significance level. A similar outcome is confirmed by the KPSS test. In particular, in all cases the LM statistic is greater than the 5% critical value, and therefore the null hypothesis of stationarity is not accepted.[35] We therefore conclude that all variables under consideration are I(1) and thus stationary after first differencing. This finding enables us to proceed to the second step of testing the existence of cointegrating relationships, using the methodology developed by Johansen [1995].

We initially estimate unrestricted VAR(p) models as in equation (6) for firms and households. Table C1 in Appendix C displays results of the VAR lag order selection test for firms. The maximum lag to test for is four. Several criteria are employed, such as the

[33] We start from a hypothetical maximum lag length, equal to 8 (two years of data). Then, this procedure requires to test downward, progressively reducing k by one unit until the last lag becomes statistically significant at a 5% confidence level. If no lags are found to be significant, then the selected lag length is zero. In this way we expect to remove any serial correlation in the residuals.

[34] The critical values for ADF tests are obtained from MacKinnon [1996], whereas the reported critical values for the LM test statistic of the KPSS test are based upon the asymptotic results presented in Kwiatkowski, Phillips, Schmidt and Shin [1992].

[35] The same unit root tests performed on first-differenced series support the stationarity hypothesis. Detailed test results are available from the authors upon request.

sequential modified likelihood ratio test as discussed in Lütkepohl (2005, p. 143), the final prediction error and the most widely used information criteria (Akaike, Schwarz, Hannan-Quinn). There is no ambiguity in the results and two lags are retained ($p = 2$). Table C2 presents the outcome of the VAR lag order selection tests for households. In this case, the information criteria suggest three lags. However, given the moderate sample size, we favour a parsimonious specification which involves the minimum number of parameters to estimate, thus also in this case we choose $p = 2$.

We next run a battery of cointegration tests, whose purpose is to determine whether the three non-stationary series (investment, value added or disposable income, user cost of capital) are cointegrated. Given that the series appear to have a trend, we assume a linear trend in the level data (unrestricted constant in the VECM model) and in the cointegrating relations. Results are shown in Table C3 (firms) and in Table C4 (households). The hypothesis that $\rho = 0$ is rejected by the trace test as well as by the maximum eigenvalue test, while the hypothesis that $\rho = 1$ cannot be rejected. This implies the existence of a single stationary relationship (and two common trends) both for firms and households, at the 5% confidence level. Therefore, we conclude that investment, value added (or disposable income) and the user cost of capital share one cointegrating relationship. The estimation of this equation for firms and households will be the object of the next two sections.

4.4 Results for firms' investment

In this section we set up a dynamic investment model for non-financial corporations. As a first step, we estimate a cointegrated VAR model represented in VECM form as in equation (8). Therefore, the two main long-run determinants of investment here considered are real value added and the real user cost of capital. Additional exogenous variables aimed at tracking the short-run investment behaviour will be introduced later on in this section. Estimation is conducted at a quarterly frequency over the 1995–2016 period, employing the maximum likelihood method (Johansen [1995]).

Table 10 presents the estimates of the VECM model without exogenous variables. In all the tables the focus will be on the first column, which refers to the investment equation. The upper part of the table displays estimates of the exactly identified cointegrating vector, after the normalization procedure suggested by Johansen [1995]. The lower part provides estimates of the speed of adjustment parameters and of the short-run dynamics.

Table 10: Firms – VECM estimates

	Coint. Equation		
$\ln(I_{t-1})$	1.0000		
$\ln(Y_{t-1})$	-1.4311***		
	[-6.1360]		
r_{t-1}	0.2126***		
	[5.2426]		
$Trend$	0.0047***		
	[8.1628]		
$Const.$	4.1476		
Error Correction:	$\Delta(\ln(I_t))$	$\Delta(\ln(Y_t))$	$\Delta(r_t)$
Speed of adj.	-0.1203***	-0.0519***	-0.0835
	[-5.5194]	[-4.4455]	[-0.5953]
$\Delta(\ln(I_{t-1}))$	0.4109***	0.0559	-0.0179
	[2.9760]	[0.7567]	[-0.0202]
$\Delta(\ln(Y_{t-1}))$	0.4337	0.4206**	0.9576
	[1.2673]	[2.2956]	[0.4346]
$\Delta(r_{t-1})$	0.0446**	0.0273**	0.5245***
	[2.0582]	[2.3588]	[3.7627]
$Const.$	0.0007	0.0010*	-0.0074
	[0.6633]	[1.6416]	[-1.0306]
Adj. R-squared	0.7284	0.6368	0.3048
Log likelihood		730.1695	
Schwarz criterion		-17.6610	

Notes: t-statistics are in [ ]. $***$, $**$ and $*$ denote statistical significance at 1, 5 and 10%, respectively.

Focusing on the investment equation, the estimated specification is:

$$\begin{aligned} \Delta\ln(I_t) &= 0.001 + 0.411\Delta\ln(I_{t-1}) + 0.434\Delta\ln(Y_{t-1}) + 0.045\Delta r_{t-1} \\ &- 0.120(\ln(I_{t-1}) - 1.431\ln(Y_{t-1}) + 0.213 r_{t-1} + 4.148 + 0.005t) \end{aligned} \tag{11}$$

Overall, the identified cointegrating vector confirms the long-run relationship between firms' investment and its two standard determinants (value added and user cost). In particular, results indicate that the long-run investment elasticity with respect to value added is highly significant and above unity. At the same time, our estimate of the negative real user cost elasticity lies in the range of the values found in the empirical literature for business investment (e.g. between -0.17 estimated for Italy by Bacchini et al. [2017] and -0.44 estimated for the UK by Ellis and Price [2004]). One plausible reason why firms' aggregate investment is not very responsive to the cost of capital is due to aggregation bias, due to the fact that different components of aggregate investment react differently to the cost of capital. In particular, Bacchini et al. [2017] find a significant negative long-run relationship between investment and the real user cost of capital for the non-ICT components of capital accumulation and a non significant association for the ICT item.[36]

Turning to the lower panel of Table 10, the speed of adjustment coefficient in the investment equation is both significant and negative (-0.12), suggesting the in each period investment adjusts partially to its long-run equilibrium level (as do the other two endogenous variables), again in accordance with the flexible neoclassical theory. Lastly, most of the short-run coefficients have significant coefficients and expected signs, especially in the investment and value added equations. The significant and positive role of lagged changes in investment in the first column could point, for example, to the lumpiness of investment. Overall, the investment equation performs well in-sample, as indicated by the adjusted

[36] As in the cointegration tests discussed in the previous section, a linear trend is also included in the cointegrating relationship, given that some of the variables – notably the user cost of capital – contain deterministic trends. Interest rates on bank loans indeed dramatically declined in the years prior to the adoption of the euro, i.e., the first years of our sample, recording a marked downward trend. Estimating our VECM model with demeaned and detrended variables yields similar results to those reported in Table 10, except for the point estimate of the trend coefficient, which becomes statistically not different from zero (results available upon request).

R-squared (which is above 70 per cent).

We next enrich the VECM model by including the additional weakly exogenous variables in order to better describe the short-run dynamics of investment. Consistent with the VECM(1) formulation, initially we consider a specification containing both contemporaneous and one-quarter lagged exogenous variables. Then, we test the significance of the corresponding coefficients and we keep only those statistically different from zero at conventional levels. The final specification in Table 11 includes lagged first differences of the business climate, lagged first differences of uncertainty and contemporaneous values of indebtedness (its 4-term moving average is considered). Focusing on the investment equation, the estimated relationship is:[37]

$$\begin{aligned}\Delta\ln(I_t) &= 0.034 + 0.350\Delta\ln(I_{t-1}) + 0.389\Delta\ln(Y_{t-1}) + 0.046\Delta r_{t-1} \\ &- 0.116(\ln(I_{t-1}) - 1.784\ln(Y_{t-1}) + 0.133r_{t-1} + 7.061 + 0.002t) \\ &- 0.134\Delta(Uncertainty)_{t-1} + 0.039\Delta(Climate)_{t-1} - 0.049MA(Debt\_to\_GDP_t)\end{aligned} \tag{12}$$

There are no notable differences in the estimated coefficients relative to the previous table, except that related to an increase in the income elasticity and a decline in the real user cost of capital elasticity. Turning to the weakly exogenous variables, both higher uncertainty and a deteriorated business climate are significantly associated with lower investment, a result in line with theoretical predictions and with the existing empirical findings for Italy (Busetti, Giordano and Zevi [2016]). The link between investment and indebtedness is also significant and negative, as expected.[38]

[37] Business climate is reported in percent (i.e., it is multiplied by 100) such that the corresponding estimated coefficient has a magnitude that is comparable to the other variables' coefficients.

[38] Results, presented in Table D1 in Annex D, are unchanged if we use the ratio of debt to its sum with equity as a measure of leverage instead of the baseline measure of debt–to–GDP used herein.

Table 11: Firms – VECM estimates with weakly exogenous variables (1)

	Coint. Equation		
$\ln(I_{t-1})$	1.0000		
$\ln(Y_{t-1})$	-1.7838***		
	[-7.8362]		
r_{t-1}	0.1325***		
	[3.4167]		
$Trend$	0.0020*		
	[1.7128]		
$Const.$	7.0613		
Error Correction:	$\Delta(\ln(I_t))$	$\Delta(\ln(Y_t))$	$\Delta(r_t)$
Speed of adj.	-0.1160***	-0.0462***	0.2276
	[-3.5496]	[-2.6270]	[1.1212]
$\Delta(\ln(I_{t-1}))$	0.3501**	0.0451	0.1839
	[2.5544]	[0.6119]	[0.2160]
$\Delta(\ln(Y_{t-1}))$	0.3889	0.4253**	2.2868
	[1.1528]	[2.3451]	[1.0917]
$\Delta(r_{t-1})$	0.0463**	0.0254**	0.4155***
	[2.1697]	[2.2128]	[3.1374]
$\Delta(Uncertainty)_{t-1}$	-0.1337**	-0.0588*	-0.3536
	[-2.1857]	[-1.7862]	[-0.9306]
$\Delta(Climate)_{t-1}$	0.0393**	0.0266***	0.3910***
	[2.1313]	[2.6856]	[3.4118]
$MA(Debt\_to\_GDP_t)$	-0.0494***	-0.0163***	0.1238*
	[-4.4337]	[-2.7212]	[1.7900]
$Const.$	0.0342***	0.0119***	-0.0956*
	[4.3129]	[2.7959]	[-1.9431]
Adj. R-squared	0.7574	0.6728	0.4224
Log likelihood		745.0876	
Schwarz criterion		-17.5409	

Notes: t-statistics are in [ ]. $***$, $**$ and $*$ denote statistical significance at 1, 5 and 10%, respectively. $MA(Debt\_to\_GDP_t)$ is a 4-term moving average of the debt-to-GDP ratio.

Table 12: Firms – VECM estimates with weakly exogenous variables (2)

	Coint. Equation		
$\ln(I_{t-1})$	1.0000		
$\ln(Y_{t-1})$	-1.2349***		
	[-4.3259]		
r_{t-1}	0.1557***		
	[3.1542]		
$Trend$	0.0034***		
	[4.4485]		
$Const.$	2.8436		
Error Correction:	$\Delta(\ln(I_t))$	$\Delta(\ln(Y_t))$	$\Delta(r_t)$
Speed of adj.	-0.0861***	-0.0350***	0.1744
	[-4.0242]	[-3.0798]	[1.2901]
$\Delta(\ln(I_{t-1}))$	0.4376***	0.0658	-0.0872
	[3.3984]	[0.9617]	[-0.1072]
$\Delta(\ln(Y_{t-1}))$	0.2586	0.3627**	2.6717
	[0.7752]	[2.0472]	[1.2673]
$\Delta(r_{t-1})$	0.0557**	0.0318***	0.3970***
	[2.5692]	[2.7650]	[2.9005]
$\Delta(Uncertainty)_{t-1}$	-0.1079*	-0.0490	-0.4261
	[-1.7914]	[-1.5324]	[-1.1195]
$\Delta(Climate)_{t-1}$	0.0327*	0.0241**	0.4074***
	[1.7888]	[2.4787]	[3.5383]
$Creditconstraints_t$	-0.1079***	-0.0426**	0.2847
	[-3.3870]	[-2.5189]	[1.4139]
$Const.$	0.0079***	0.0038***	-0.0305*
	[3.1142]	[2.8058]	[-1.9123]
Adj. R-squared	0.7659	0.6920	0.4227
Log likelihood		747.9253	
Schwarz criterion		-17.6136	

Notes: t-statistics are in [ ]. $***$, $**$ and $*$ denote statistical significance at 1, 5 and 10%, respectively.

In Table 12 we replace indebtedness with the credit constraints variable described in Section 3. The latter is statistically significant and displays the expected negative sign. All previously described results hold and the fit of the investment equation further improves.

Finally, to take into account the potential impact of the higher exit rate of Italian firms during the sovereign debt crisis, documented in the microeconomic literature and discussed in Section 2, we also attempted to control for step or one-off dummies in the years 2010–2011. Results do not support the evidence of a structural break in those years, thereby confirming that with our macroeconomic dataset we are unable to account for firm turnover and its impact on investment in this period. We therefore postpone the analysis of this issue to future research, necessarily to be conducted on firm-level data.

4.5 Results for households' investment

In this section, we turn our attention to households' investment. Table 13 summarises the estimates of the cointegrated VECM model as in equation (8), with no additional, weakly exogenous variable.

Table 13: Households – VECM estimates

	Coint. Equation		
$\ln(I_{t-1})$	1.0000		
$\ln(Yd_{t-1})$	-2.3887***		
	[-21.5668]		
r_{t-1}	0.0852***		
	[3.8825]		
Trend	0.0044 ***		
	[14.1878]		
Const.	13.0670		
Error Correction:	$\Delta(\ln(I_t))$	$\Delta(\ln(Yd_t))$	$\Delta(r_t)$
Speed of adj.	-0.1220***	-0.0218	0.8590***
	[-5.0237]	[-1.2810]	[5.2132]
$\Delta(\ln(I_{t-1}))$	0.7551***	0.1727***	0.0628
	[9.7934]	[3.1903]	[0.1202]
$\Delta(\ln(Yd_{t-1}))$	-0.3967**	0.1177	6.6050***
	[-2.2906]	[0.9682]	[5.6228]
$\Delta(r_{t-1})$	0.0399***	0.0327***	-0.1293
	[2.7510]	[3.2114]	[-1.3149]
Const.	0.0004	0.0011*	-0.0192***
	[0.4092]	[1.7136]	[-3.1852]
Adj. R-squared	0.7332	0.3556	0.3897
Log likelihood		708.2334	
Schwarz criterion		-17.0986	

Notes: t-statistics are in [ ]. $***$, $**$ and $*$ denote statistical significance at 1, 5 and 10%, respectively.

Focusing on the investment equation, the estimated specification is:

$$\begin{aligned} \Delta\ln(I_t) &= 0.0004 + 0.755\Delta\ln(I_{t-1}) - 0.397\Delta\ln(Yd_{t-1}) + 0.040\Delta r_{t-1} \\ &- 0.122(\ln(I_{t-1}) - 2.389\ln(Yd_{t-1}) + 0.085r_{t-1} + 13.067 + 0.004t) \end{aligned} \quad (13)$$

The long-run investment elasticity to real disposable income is significant and higher than 2, a value which appears to be larger than what usually found in the empirical literature (see, for instance, Arestis and González [2014]).[39] The elasticity of investment

[39]When considering another proxy of households' demand for investment, i.e. real gross savings, we

to the user cost of capital is statistically significant at conventional levels and negative (-0.085), as predicted by theory and in line with similar estimates recently provided by Loberto and Zollino [2016] for housing investment in Italy. Furthermore, in the investment equation results underscore a negative and significant speed of adjustment coefficient to the equilibrium value, supporting a valid error-correction representation.

As a second step we examine the short-run influence of weakly exogenous variables (namely, uncertainty, confidence and indebtedness) on households' investment (Table 14). Focusing on the investment equation, the estimated relationship is:

$$\begin{aligned} \Delta \ln(I_t) &= 0.010 + 0.764\Delta \ln(I_{t-1}) - 0.422\Delta \ln(Y_{t-1}) + 0.035\Delta r_{t-1} \\ &- 0.128(\ln(I_{t-1}) - 2.537 \ln(Y_{t-1}) + 0.090 r_{t-1} + 14.294 + 0.003t) \\ &+ 0.010\Delta(Uncertainty)_{t-1} + 4.075\Delta(Confidence)_{t-1} - 0.012MA(Debt\_to\_income_t) \end{aligned} \quad (14)$$

The previously estimated long-run relationship is confirmed in this augmented specification, as is the speed of adjustment coefficient. Similarly to firms, initially we consider a specification containing both contemporaneous and one-quarter lagged weakly exogenous variables. Then, for the sake of parsimonious modelling, we keep only those coefficients that are statistically different from zero at conventional levels. In the short run, higher (lagged) confidence is positively associated with households' investment. Conversely, the coefficient attached to uncertainty turns out to be not significant. This may be due to its high correlation with the confidence variable, as mentioned in Section 3.2. Lastly, higher indebtedness is significantly associated with lower investment.

obtain an estimate of the long-run investment elasticity that is highly significant and close to 1. In this case, the estimate of the elasticity to the user cost of capital is also significant and higher in absolute value (-0.189). One possible explanation of why we obtain such a high elasticity to income is that we are considering current, as opposed to permanent, income, the latter being more appropriate to model households' investment demand but more difficult to estimate. On this, see, for example, Duca, Muellbauer and Murphy [2012], which too finds an income elasticity of over 2 in the case of US housing capital stock when current income is employed (and around 1 when permanent income is included).

Table 14: Households – VECM estimates with weakly exogenous variables

	Coint. Equation		
$\ln(I_{t-1})$	1.0000		
$\ln(Y_{t-1})$	-2.5365***		
	[-14.1030]		
r_{t-1}	0.0895***		
	[3.7001]		
$Trend$	0.0032**		
	[2.3118]		
$Const.$	14.2944		
Error Correction:	$\Delta(\ln(I_t))$	$\Delta(\ln(Y_t))$	$\Delta(r_t)$
Speed of adj.	-0.1276***	-0.0090	0.6382***
	[-3.9792]	[-0.4088]	[2.9390]
$\Delta(\ln(I_{t-1}))$	0.7639***	0.1691***	0.3890
	[9.0467]	[2.9314]	[0.6802]
$\Delta(\ln(Y_{t-1}))$	-0.4216**	0.1344	6.5951***
	[-2.3128]	[1.0789]	[5.3420]
$\Delta(r_{t-1})$	0.0349**	0.0314***	-0.1397
	[2.3164]	[3.0499]	[-1.3696]
$\Delta(Uncertainty)_{t-1}$	0.0096	-0.0234	0.1667
	[0.2672]	[-0.9476]	[0.6817]
$\Delta(Confidence)_{t-1}$	0.0407*	0.0327	0.0372
	[1.6570]	[1.4758]	[0.1692]
$MA(Debt\_to\_income_t)$	-0.0200***	-0.0056	0.2245***
	[-2.7164]	[-1.1119]	[4.5033]
$Const.$	0.0098***	0.0036	-0.1263***
	[2.7059]	[1.4706]	[-5.1369]
Adj. R-squared	0.7327	0.3891	0.3905
Log likelihood		714.4894	
Schwarz criterion		-16.7563	

Notes: t-statistics are in [ ]. $***$, $**$ and $*$ denote statistical significance at 1, 5 and 10%, respectively. $MA(Debt\_to\_income_t)$ is a 4-term moving average of the debt-to-income ratio.

4.6 The role of financing constraints in explaining firms' and households' investment

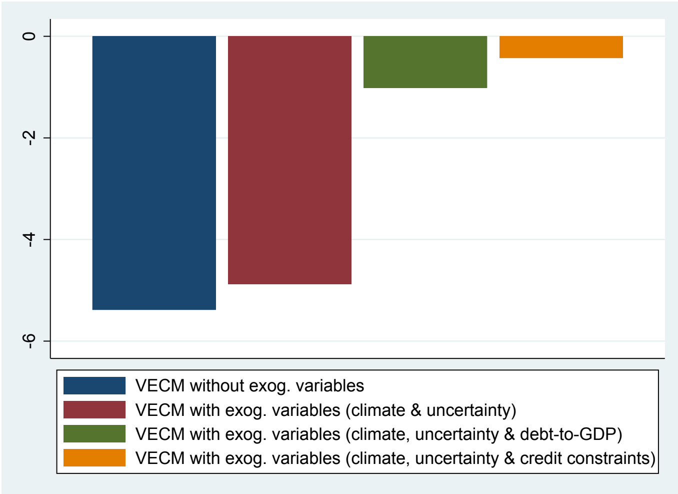
Thus far, we have provided empirical evidence that, amongst other results, firms' indebtedness or credit constraints, on the one hand, and households' indebtedness, on the other hand, have been significantly and negatively associated with firms and households' investment respectively since 1995. In this section, we test how much the inclusion of these financial variables matter in explaining (in-sample) investment activity in Italy, in particular during the recent double-dip recession, when Italy's gross fixed capital accumulation process diverged significantly from that of the other main euro-area countries (recall Figure 1).

As in Barkbu et al. [2015], we analyse the residuals of the various models estimated herein. We term "unexplained investment growth gap" the cumulative sum of residuals, i.e the cumulative difference between the predicted and the actual quarter-on-quarter growth rates of real investment in the crisis years 2009-2012.

The comparison of different models allows us to better understand what the drivers of this shortfall are. Starting with firms, in Figure 7 we first plot the gap resulting from the flexible neoclassical model, which is based on estimation results in Table 10. The second bar is based on a version of Table 11, available upon request, which includes only the two non-financial weakly exogenous variables (business climate and uncertainty). The third gap is based on the model in Table 11, which also includes indebtedness, whereas the fourth bar refers to the model in Table 12, which alternatively includes financing constraints. The chart is compelling: firms' investment growth gap is indeed significant and negative, in that after the eruption of the global financial crisis, investment dynamics in Italy were systematically lower than the model-predicted growth rates, but the measured shortfall is significantly more contained when either proxy of financing constraints is included, yet in particular when using the survey-based measure.

Figure 7: Non-financial corporations: Unexplained investment growth gap (2008:Q4–2012:Q4)

(percentage points)

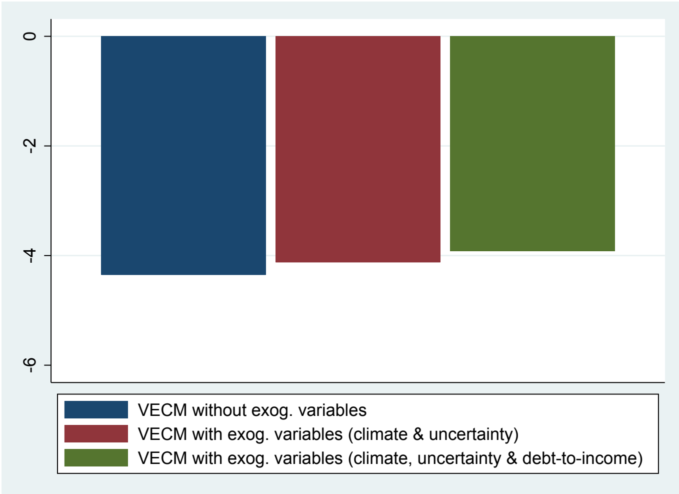


Source: Authors' estimates.

In Figure 8 we repeat the same analysis for households, with the difference that we only have a indebtedness measure of financial constraints. Also in this case, households' investment gap is negative and the shortfall appears to be smaller when exogenous variables are included. However, the improvement across models is less remarkable compared to that measured for non-financial corporations, possibly because of the lower indebtedness of households than firms in Italy.

Figure 8: Households: Unexplained investment growth gap (2008:Q4–2012:Q4)

(percentage points)



Source: Authors' estimates.

5 Conclusions

The total-economy investment rate in Italy currently stands at its lowest levels since 1995, the first year for which official national account series are available; taking a longer historical perspective, it has never been so low since 1950. Although gross fixed capital formation fell substantially in most advanced economies after the outbreak of the global financial crisis, Italy stands out as having being harder hit and for recovering more gradually.

Following a resurging literature, this paper examines the main macroeconomic determinants affecting capital accumulation in Italy and, in particular, the role of financial constraints. The value added of the paper with respect to the existing macroeconomic studies is threefold. First, we rely on institutional sector national accounts, which enable us to disaggregate investment of non-financial corporations and of households. Second, we adopt a vector error correction approach, which mimics the neoclassical theory of aggregate investment, as put forward by Hall and Jorgenson [1967], and at the same time disentangles long-run and short-run determinants of investment. Third, using financial account data we construct a macro measure of financial constraints (the degree of indebtedness) which, at least for firms, we find to be significantly correlated in dynamics with a survey-based measure of credit constraints, thereby bridging the macro-micro gap.

For both institutional sectors we find that the relationship between investment, on the one side, and the real user cost of capital and output, on the other side, is in line with the neoclassical theory. As a second step, we integrate the model with selected weakly exogenous (relative to the cointegrating relationship) variables with the aim of better describing the short-run dynamics of investment in Italy. We find that investment growth is positively correlated with improvements in sentiment; furthermore, there is evidence of a negative association between uncertainty and firms' investment at least. Turning to financing constraints, firms' and households' investment are significantly dampened by high indebtedness. Our results for firms are further confirmed when replacing indebtedness with our survey-based measure of credit rationing. In addition, based on our sample, we show that investment models not accounting for financing constraints are unable to satisfactorily explain Italy's recent and exceptional investment downturn, especially in connection with firms' capital accumulation.

In the interest of extending the validity of these findings, the empirical framework presented here for Italy could be applied to other countries for which the necessary data (in particular, financial accounts) are available, within an integrated and consistent statistical framework that allows for international comparability. Existing studies have indeed shown that, for example, the contribution of financial variables to real fluctuations is heterogeneous across countries (Chirinko, de Haan and Sterken [2008]; Hubrich et al. [2013]). It would therefore be interesting to assess whether in Italy they matter to a different extent relative to other comparable advanced economies, thereby potentially accounting for Italy's sharper contraction and subsequent slower recovery in investment. We leave this empirical issue to be tested in future research.

Appendix A: Alternative proxies of economic uncertainty

All proxies of uncertainty are intrinsically imperfect. Indeed, in order to measure economic uncertainty correctly, ideally one would like to know the subjective probability distributions over future events from firms and households, but this information is generally unavailable (ECB [2016]). The measure used in this paper captures the dispersion in economic agents' expectations. The main limitation of this measure is that expectations may be differing across agents, but "certain" (i.e., with probability close to one) for the individual agents expressing them. However, we believe that our proxy is more suited than other possible indicators for the purpose of our analysis. First, it is the only possible measure that we could construct separately for both firms and households, moreover at a quarterly frequency. Second, our uncertainty proxy does not rely on financial asset prices, thereby allowing us to distinguish empirically between uncertainty and financing constraints (on this see, for example, Caldara et al. [2016]). Third, we find that the shortcomings of alternative proxies are larger than the drawbacks of our survey-based indicator, given the aim of our analysis.

The most prominent, alternative measures of uncertainty put forward by the empirical literature are: a) the dispersion in GDP growth forecasts of professional analysts, as in Zarnowitz and Lambros [1987] and Bomberger [1996]; b) economic policy uncertainty, measured as the newspaper article count of the words "uncertainty" and similar, as proposed by Baker, Bloom and Davis [2016]; c) actual or implied stock market volatility, in accordance with Bloom [2009].

The first measure assumes that increasingly diverging opinions about the economic outlook among professional forecasters imply that it is ever more uncertain to make economic forecasts and therefore that economic uncertainty is on the rise. As discussed in ECB [2016], the level and fluctuations in this dispersion measure may, however, also be traced back to other factors, such as differences in forecast techniques and in information sets. Moreover, analysts may keep their projections unchanged or revise them all in the same direction, while actual uncertainty may be changing significantly. In a similar vein, Jurado, Ludvigson and Ng [2015] propose an economic uncertainty index based on the implied forecast errors for real economic activity derived from a factor model that employs many economic and financial series, which suffers from similar limitations.

The second measure captures a broader dimension of uncertainty than economic uncertainty, including political uncertainty. The selection of newspapers (two per country) might, however, not be representative of the media coverage in a specific country. Moreover, these data for Italy start in 1997, thereby not covering the whole time period under analysis in this paper. Finally, it is a slow-moving indicator and has been found to capture long-run uncertainty (Barrero, Bloom and Wright [2017]); therefore it is not appropriate to capture short-term dynamics.

The third measure proxies a canonical measure of uncertainty in the financial markets but does not necessarily mirror actual developments in economic uncertainty surrounding economic activity, especially in a country such as Italy where listed companies are relatively few. A rise in stock market volatility may be related to changes in risk aversion or market sentiment, which can be independent of a firm's fundamentals. Moreover, as emphasized by Bekaert, Hoerova and Lo Duca [2013], the risk component of stock market data is highly countercyclical, due to the countercyclical nature of contractual and informational frictions associated with financial shocks.

Appendix B: Unit root test results

Table B1: Firms – ADF unit root test on $\ln(I_t)$

Null Hypothesis: $\ln(I_t)$ has a unit root			
		t-Statistic	Prob.*
Augmented Dickey-Fuller test		-1.8794	0.6553
Test critical values:	1% level	-4.0851	
	5% level	-3.4709	
	10% level	-3.1625	

* MacKinnon [1996] one-sided p-values.

Table B2: Firms – KPSS unit root test on $\ln(I_t)$

Null Hypothesis: $\ln(I_t)$ is stationary
Exogenous: Constant, Linear Trend
Bandwidth: 6 (Newey-West automatic) using Bartlett kernel

		LM-Stat.
Kwiatkowski-Phillips-Schmidt-Shin test statistic		0.2885
Asymptotic critical values*	1% level	0.2160
	5% level	0.1460
	10% level	0.1190

* Kwiatkowski-Phillips-Schmidt-Shin (1992, Table 1).

Table B3: Firms – ADF unit root test on $\ln(Y_t)$

Null Hypothesis: $\ln(Y_t)$ has a unit root			
		t-Statistic	Prob.*
Augmented Dickey-Fuller test		-2.7598	0.2165
Test critical values:	1% level	-4.0800	
	5% level	-3.4685	
	10% level	-3.1611	

* MacKinnon [1996] one-sided p-values.

Table B4: Firms – KPSS unit root test on $\ln(Y_t)$

Null Hypothesis: $\ln(Y_t)$ is stationary
Exogenous: Constant, Linear Trend
Bandwidth: 6 (Newey-West automatic) using Bartlett kernel

		LM-Stat.
Kwiatkowski-Phillips-Schmidt-Shin test statistic		0.2708
Asymptotic critical values*	1% level	0.2160
	5% level	0.1460
	10% level	0.1190

* Kwiatkowski-Phillips-Schmidt-Shin (1992, Table 1).

Table B5: Firms – ADF unit root test on r_t

Null Hypothesis: r_t has a unit root

		t-Statistic	Prob.*
Augmented Dickey-Fuller test		-3.3469	0.0658
Test critical values:	1% level	-4.0696	
	5% level	-3.4635	
	10% level	-3.1582	

* MacKinnon [1996] one-sided p-values.

Table B6: Firms – KPSS unit root test on r_t

Null Hypothesis: r_t is stationary
Exogenous: Constant, Linear Trend
Bandwidth: 6 (Newey-West automatic) using Bartlett kernel

		LM-Stat.
Kwiatkowski-Phillips-Schmidt-Shin test statistic		0.2146
Asymptotic critical values*	1% level	0.2160
	5% level	0.1460
	10% level	0.1190

* Kwiatkowski-Phillips-Schmidt-Shin (1992, Table 1).

Table B7: Households – ADF unit root test on $\ln(I_t)$

Null Hypothesis: $\ln(I_t)$ has a unit root			
		t-Statistic	Prob.*
Augmented Dickey-Fuller test		-1.0258	0.9337
Test critical values:	1% level	-4.0851	
	5% level	-3.4709	
	10% level	-3.1625	

* MacKinnon [1996] one-sided p-values.

Table B8: Households – KPSS unit root test on $\ln(I_t)$

Null Hypothesis: $\ln(I_t)$ is stationary
Exogenous: Constant, Linear Trend
Bandwidth: 6 (Newey-West automatic) using Bartlett kernel

		LM-Stat.
Kwiatkowski-Phillips-Schmidt-Shin test statistic		0.3046
Asymptotic critical values*	1% level	0.2160
	5% level	0.1460
	10% level	0.1190

* Kwiatkowski-Phillips-Schmidt-Shin (1992, Table 1).

Table B9: Households – ADF unit root test on $\ln(Yd_t)$

Null Hypothesis: $\ln(Yd_t)$ has a unit root			
		t-Statistic	Prob.*
Augmented Dickey-Fuller test		-1.9481	0.6201
Test critical values:	1% level	-4.0753	
	5% level	-3.4662	
	10% level	-3.1598	

* MacKinnon [1996] one-sided p-values.

Table B10: Households – KPSS unit root test on $\ln(Yd_t)$

Null Hypothesis: $\ln(Yd_t)$ is stationary		
Exogenous: Constant, Linear Trend		
Bandwidth: 6 (Newey-West automatic) using Bartlett kernel		
		LM-Stat.
Kwiatkowski-Phillips-Schmidt-Shin test statistic		0.2604
Asymptotic critical values*	1% level	0.2160
	5% level	0.1460
	10% level	0.1190

* Kwiatkowski-Phillips-Schmidt-Shin (1992, Table 1).

Table B11: Households – ADF unit root test on r_t

Null Hypothesis: r_t has a unit root			
		t-Statistic	Prob.*
Augmented Dickey-Fuller test		-2.2501	0.4556
Test critical values:	1% level	-4.0784	
	5% level	-3.4677	
	10% level	-3.1606	

* MacKinnon [1996] one-sided p-values.

Table B12: Households – KPSS unit root test on r_t

Null Hypothesis: r_t is stationary		
Exogenous: Constant, Linear Trend		
Bandwidth: 6 (Newey-West automatic) using Bartlett kernel		
		LM-Stat.
Kwiatkowski-Phillips-Schmidt-Shin test statistic		0.1975
Asymptotic critical values*	1% level	0.2160
	5% level	0.1460
	10% level	0.1190

* Kwiatkowski-Phillips-Schmidt-Shin (1992, Table 1).

Appendix C: VAR order selection and cointegration rank tests

Table C1: Firms – VAR Lag Order Selection

Lag	LogLik	LR.$^{(1)}$	FPE$^{(2)}$	AIC$^{(3)}$	SC$^{(4)}$	HQ$^{(5)}$
0	274.8617	NA	1.57e-07	-7.1543	-7.0623	-7.1175
1	660.9510	731.5378	7.69e-12	-17.0777	-16.7096	-16.9306
2	713.1676	94.8142*	2.47e-12*	-18.2149*	-17.5709*	-17.9576*
3	715.8615	4.6789	2.92e-12	-18.0490	-17.1290	-17.6813
4	721.3434	9.0885	3.23e-12	-17.9564	-16.7604	-17.4784

(1) Sequential modified LR test statistic (each test at 5% level)
(2) Final prediction error
(3) Akaike information criterion
(4) Schwarz information criterion
(5) Hannan-Quinn information criterion
* Indicates lag order selected by the criterion.

Table C2: Households – VAR Lag Order Selection

Lag	LogLik	LR.$^{(1)}$	FPE$^{(2)}$	AIC$^{(3)}$	SC$^{(4)}$	HQ$^{(5)}$
0	239.1099	NA	4.02e-07	-6.2134	-6.1214	-6.1767
1	626.7054	734.3915	1.89e-11	-16.1765	-15.8085	-16.0294
2	681.1158	98.7979	5.74e-12	-17.3715	-16.7275	-17.1141
3	708.9787	48.3934*	3.50e-12*	-17.8679*	-16.9478*	-17.5002*
4	712.0073	5.0212	4.12e-12	-17.7107	-16.5147	-17.2327

(1) Sequential modified LR test statistic (each test at 5% level)
(2) Final prediction error
(3) Akaike information criterion
(4) Schwarz information criterion
(5) Hannan-Quinn information criterion
* Indicates lag order selected by the criterion.

Table C3: Firms – Cointegration Rank Test test

Trend assumption: Linear deterministic trend (restricted)
Lags interval (in first differences): 1 to 1

Trace

Hypothesized No. of CE(s)	Eigenvalue	Trace(1) Statistic	0.05 Critical Value	Prob.**
None*	0.3038	46.8780	42.9153	0.0191
At most 1	0.1500	18.6370	25.8721	0.3027
At most 2	0.0735	5.9585	12.5180	0.4656

Maximum Eigenvalue

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen(2) Statistic	0.05 Critical Value	Prob.**
None*	0.3038	28.2410	25.8232	0.0236
At most 1	0.1500	12.6785	19.3870	0.3547
At most 2	0.0735	5.9585	12.5180	0.4656

(1) Trace test indicates 1 cointegrating equation(s) at the 0.05 level
(2) Max-eigenvalue test indicates 1 cointegrating equation(s) at the 0.05 level
* Denotes rejection of the hypothesis at the 0.05 level
** MacKinnon-Haug-Michelis (1999) p-values.

Table C4: Households – Cointegration Rank Test test

Trend assumption: Linear deterministic trend (restricted)
Lags interval (in first differences): 1 to 1

Trace

Hypothesized No. of CE(s)	Eigenvalue	Trace[(1)] Statistic	0.05 Critical Value	Prob.**
None*	0.4521	64.0658	42.9153	0.0001
At most 1	0.1272	17.1399	25.8721	0.4046
At most 2	0.0803	6.5254	12.5180	0.3967

Maximum Eigenvalue

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen[(2)] Statistic	0.05 Critical Value	Prob.**
None*	0.4521	46.9259	25.8232	0.0000
At most 1	0.1272	10.6145	19.3870	0.5535
At most 2	0.0803	6.5254	12.5180	0.3967

(1) Trace test indicates 1 cointegrating equation(s) at the 0.10 level
(2) Max-eigenvalue test indicates 1 cointegrating equation(s) at the 0.05 level
* Denotes rejection of the hypothesis at the 0.05 level
** MacKinnon-Haug-Michelis (1999) p-values.

Appendix D: Additional tables

Table D1: Firms – VECM estimates with weakly exogenous variables (3)

	Coint. Equation		
$\ln(I_{t-1})$	1.0000		
$\ln(Y_{t-1})$	-1.0471***		
	[-3.5190]		
r_{t-1}	0.1442***		
	[3.0860]		
$Trend$	0.0034***		
	[4.8430]		
$Const.$	1.4568		
Error Correction:	$\Delta(\ln(I_t))$	$\Delta(\ln(Y_t))$	$\Delta(r_t)$
Speed of adj.	-0.0903***	-0.0366***	0.2069*
	[-4.4953]	[-3.4570]	[1.6432]
$\Delta(\ln(I_{t-1}))$	0.3780***	0.0398	0.01792
	[2.8630]	[0.5725]	[0.0217]
$\Delta(\ln(Y_{t-1}))$	0.3664	0.4026**	2.4227*
	[1.1089]	[2.3124]	[1.1694]
$\Delta(r_{t-1})$	0.0428**	0.0269**	0.4254***
	[2.0632]	[2.4595]	[3.2724]
$\Delta(Uncertainty)_{t-1}$	-0.1210**	-0.0541*	-0.4013
	[-2.0031]	[-1.7014]	[-1.0596]
$\Delta(Climate)_{t-1}$	0.0300*	0.0200**	0.4100***
	[1.8312]	[2.5278]	[3.5511]
$Leverage_{t-1}$	-0.0789***	-0.0332**	0.1721
	[-3.1563]	[-2.5227]	[1.0977]
$Const.$	0.0079***	-0.0332***	0.1721*
	[-3.1563]	[-2.5227]	[1.0977]
Adj. R-squared	0.7612	0.6907	0.4204
Log likelihood		747.0534	
Schwarz criterion		-17.5913	

Notes: t-statistics are in [ ]. ***, ** and * denote statistical significance at 1, 5 and 10%, respectively.

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