BANCA D'ITALIA
EUROSISTEMA

Temi di Discussione

(Working Papers)

Forecaster heterogeneity, surprises and financial markets

by Marcello Pericoli and Giovanni Veronese

July 2015

Number 1020

BANCA D'ITALIA
EUROSISTEMA

Temi di discussione

(Working papers)

Forecaster heterogeneity, surprises and financial markets

by Marcello Pericoli and Giovanni Veronese

Number 1020 - July 2015

ISSN 1594-7939 (print)
ISSN 2281-3950 (online)

Printed by the Printing and Publishing Division of the Bank of Italy

FORECASTER HETEROGENEITY, SURPRISES AND FINANCIAL MARKETS

by Marcello Pericoli* and Giovanni Veronese*

Abstract

We analyze the impact of US macroeconomic surprises and forecaster heterogeneity on the dollar/euro exchange rate and on US and German long-term interest rates from 1999 to 2014. We show that a direct proxy of macroeconomic disagreement, namely the heterogeneity of forecasters' beliefs about upcoming macroeconomic releases, is a factor in explaining daily and intra-day movements. Surprises affect long-term yields and the exchange rate more strongly when forecaster heterogeneity is less pronounced. This result holds for the main US macroeconomic surprises and is robust to the frequency of the data used in the estimation. However, the sensitivity changes with the sample. In this regard, we show that when the same regressions are estimated for a pre-crisis period, a crisis period, and an unconventional monetary policy period, there is time variation in the responses: unconventional monetary policies attenuated the response of the exchange rate to US macroeconomic news, but had little impact on long-term interest rates in the US or Germany. Disagreement among forecasters remains relevant in determining an asymmetric response of these asset prices. The paper demonstrates that in heterogeneity of beliefs is a relevant factor in describing the behavior of asset prices, even at high frequency.

JEL Classification: E44, E52, F31, G14.
Keywords: surprises, forecaster heterogeneity, foreign exchange, long-term interest rates, unconventional monetary policy.

Contents

* Bank of Italy, DG Economics, Statistics and Research

1 Introduction[1]

Scheduled releases of macroeconomic indicators constitute the main source of public information about key macro variables to participants in financial markets. As such, the response of various asset prices to macro data releases has long been studied in order to better understand the link between financial markets and the macroeconomy.

Asset price movements following regularly scheduled macroeconomic announcements provide a unique source of information about the evolution of private sector expectations and about how expectations feed back on the economy. Policy makers and market analysts closely follow these market reactions, and a large and active literature has developed documenting reactions of various markets to macroeconomic news.

This paper investigates how the reaction of exchange rates and long-term yields to macroeconomic news depends on the heterogeneity of beliefs across market participants.

In models featuring interest parity conditions as well as Taylor-rules for monetary policy, macroeconomic news should impact directly interest rates and foreign exchange rates, as they are clearly intertwined with the monetary policy stance. As macroeconomic news buffet these markets, the updated trajectories on inflation and growth will lead to an update also of the expected path of interest rates, thereby affecting long-term yields, as well as the risk premia. So, results can give indications on the dependence of term premia on surprises and the role played by the monetary policy stance in determining such effects.

However, the reaction of asset prices to macroeconomic surprises is not the result of a mechanical rule followed by market participants, but rather it depends on the state of the economy and financial markets. Andersen et al. (2007) show how the impact of macroeconomic news on various asset classes becomes stronger when conditioning on the state of the economy. Beber and Brandt (2010) find that macroeconomic announcements are most important for bond returns and volatilities when they contain bad news during an expansion. More recently, Goldberg and Grisse (2013) show how various measures on the state of the US economy (VIX index, level of interest rates, inflation) interact to determine the impact of macroeconomic news, both on government bond prices and exchange rates.

In this paper we investigate how the heterogeneity among investors, proxied by their beliefs' dispersion, plays a role in determining the intensity and the direction of these responses. We show that heterogeneity, proxied by the dispersion of analysts' forecasts around each macroeconomic release, is an important source of variation in the response of asset prices. We find that the more forecasters disagree ex-ante regarding an upcoming scheduled macroeconomic announcement, the more muted is the price response to the macroeconomic surprise. This finding unveils a novel channel of state dependence in the asset price response to macroeconomic news.

[1]The views expressed are those of the authors and do not necessary represent those of the Bank of Italy. We would like to thank Paul Dozier for providing US intraday data and Alessio Anzuini for the exchange rate intraday data. Jon Frost and participants to the Bank of Italy-ECB-DNB workshop, EFA Conference and Bank of Italy seminar.
Email addresses: marcello.pericoli@bancaditalia.it and giovanni.veronese@bancaditalia.it

Our results suggests that interpreting disagreement as reflecting the market "prior uncertainty" regarding the announcement outcome may be misleading.[2] The role of prior uncertainty in exacerbating price and volume reaction to news, is described in Kim and Verrecchia (1991). In line with a bayesian learning framework the price update is a decreasing function of prior precision as investors revise more their beliefs. Hautsch and Hess (2007) analyze how the strength of the price impact of unanticipated information depends on its "quality", i.e. the relative precision of the macroeconomic news and on the information available before an announcement. In periods when released data is perceived to be more precise, relative to the precision of the prior information, a stronger price reaction should be observed to a given piece of unexpected information. If the disagreement among analysts is taken as a proxy for prior uncertainty our results would appear puzzling at least.

We instead interpret our results as evidence that disagreement on individual macroeconomic releases matter, and that heterogenous beliefs are needed to explain this otherwise puzzling response of asset prices to macroeconomic news. To this end, we reinterpret our results in light of the model of Ottaviani and Sørensen (2015), which predicts that the asset price response to information critically depends on the degree of dispersion of prior beliefs. Their result hinges on wealth effects and on how wealth and beliefs are distributed across market participants. In particular, when beliefs heterogeneity is *more pronounced* the price response to news is shown to be more muted, a prediction in line with what we find in our empirical exercise.

We also document that forecaster heterogeneity regarding several US key macroeconomic indicators is driven not only by variable-specific factors, but also by a sizeable common component. This common component displays some comovement with the VIX index, a popular proxy of "global risk aversion" and overall economic uncertainty.

As a robustness check we extend our analysis to a higher frequency setting, to take into account of the possible reaction of financial markets to other releases of news during the same day. Our findings are robust to this extension, and we continue to find that during phases of low disagreement on the business cycle the impact of macroeconomic news is more pronounced.

During the years covered in our analysis bond markets have been subject to several "anomalies": in the first part of the 2000s, the low level of long-term interest rates raised the perception of a conundrum in US long-term bond yields as they remained low even with an accelerating US business cycle (Greenspan, 2005); in the aftermath of the global financial crisis, movements in bond yields may have been dampened approaching the zero lower bound, especially at shorter maturities. Finally, also more recently the behavior of US long-term interest rates has puzzled many economists, as they did not return to more normal historical levels but, conversely, their general trend has been down since the end of the Great Recession (Hamilton, 2014).

This leads us to investigate whether our findings on the role of agents' disagreement is affected

[2]In their seminal contribution* Andersen et al. (2003) analyze whether uncertainty about the state of the economy, as proxied by the standard deviation of expectations across the individual forecasters, increases following the arrival of bad macroeconomic news in good times.

by the choice of a particular subsample. To this end we apply the methodology of Swanson and Williams (2014a) considering three periods: tranquil (2000-07), crisis (2007-09) and US Quantitative Easing (QE) years (2009-14). In general, our results on the relevance of heterogeneity in beliefs for the response of asset prices continue to hold during tranquil and QE periods. Long-dated US bond yields, while responding to US macroeconomic surprises, also display on average smaller sensitivity during the QE period. Furthermore, we find that during the QE period German long-term yields were more reactive to US macroeconomic surprises, signalling a potential portfolio shift of investors from the US to the German bond market. In contrast, the sensitivity of the USD/EUR exchange rate appears to have changed more radically over time, becoming smaller during the crisis and mostly muted in the QE period. This may suggest that the implementation of non-standard measures by the US monetary authorities has made the exchange rate less reactive to macroeconomic fundamentals as US short-term rates have reached the zero-lower bound.

This result has powerful implications both for theorists and policy makers. In fact, the former can use this result as a theoretical base for linking aggregate uncertainty to individual heterogeneity (or dispersion of beliefs). Several theoretical papers have addressed how introducing a modest amount of heterogeneity and disagreement can help explain several puzzling phenomena for financial markets – Xiong and Yan (2010) for bond prices, and more recently Andrei et al. (2014) for stock prices. Finally, policy makers can use the information content of beliefs' dispersion as an important toolkit for evaluating the effectiveness and the credibility of policy initiatives on markets.

The paper is organized as follows. Section 2 reviews the literature on the impact of surprises on asset prices. Section 3 describes the data. Section 4 presents the results for the case of different disagreement among analysts. Section 5 documents the relationship between disagreement and measures of macro uncertainty and looks at the time-varying sensitivity of asset markets to surprises. Section 6 concludes.

2 Literature review

Starting from the work of Andersen et al. (2003) it is well established that the impact of macroeconomic news on asset prices can be substantial and significant, especially if the time-window to detect the effects is short enough around the announcements. Some variables are found to be more influential than others because of several characteristics. First their outcome is released according to a predefined calendar, analysts are surveyed ahead of each release by the main financial platforms (Bloomberg, Reuters, MMS International) and the actual announcement is associated to a publicly observable and meaningful market median-surprise.[3] Second, and more importantly, markets seem to assign relatively more importance to certain indicators as they reveal in real-time more information on the underlying state of the economy. Not only are they more strongly correlated with the

[3]This has been the case for most US macroeconomic statistics since the late eighties, while it has become common practice only more recently for other countries.

fundamentals, but also their release occurs ahead of important and encompassing variables on the state of the economy (like GDP and inflation). This timeliness makes them also particularly useful to nowcast in real time the state of the economy (Giannone et al., 2008). Finally, they are deemed to be "precise" in the sense of being subject to relatively small volatility and ex-post revisions (Gilbert et al., 2010).

The literature has so far mostly focused on measuring "the impact" of different macroeconomic news, discussing the relevance of high frequency data and assessing which variables are most relevant for the various different asset prices –for a critical review on foreign exchange returns see Neely and Dey (2010). Some authors have investigated the stability over time of the effect of news, testing for parameter constancy and argued that they are mostly stable over time (Faust et al., 2007). Based on this premise, several authors and investment banks have proposed macroeconomic "surprise indices", obtained by re-weighting the individual surprises of the most recent macroeconomic news. The most reknown among these, the *Citi Surprise Index*, based on the work of James and Kasikov (2008), derives the weighting under the assumption of a stable relationship between news releases and asset price responses, suggesting the existence of some sort of mechanical reaction on the part of market participants.

However, the relationship between macroeconomic news and asset prices should not be expected to be stable over time. Time variation may originate in the context of a Taylor-rule type models of policy reaction functions. In a stylized monetary model with zero-lower bound, Swanson and Williams (2014a) show how interest rates are expected to react differently to macroeconomic news as the bound is approached. The prediction from their model is confirmed by the data, with short-term interest rates becoming almost insensitive to macroeconomic news after the QE program in the US, especially at the shorter horizons.

The phase of the business cycle may matter as well, since economic conditions and on investors risk aversion [4]. Among the first papers to investigate the response of asset prices to macroeconomic news at a very high frequency, Andersen et al. (2003) show how U.S. dollar spot exchange rates display conditional mean jumps and document an *asymmetric* reaction to positive and negative surprises. They interpret this result in light of the information processing and price discovery literature. Andersen et al. (2007) extend their earlier analysis to stocks, bonds and foreign exchange markets and find that macroeconomic news indeed produce conditional mean jumps of U.S., German and British stock, bond and foreign exchange markets. Equity markets are found to react *differently to news* depending on the *stage of the business cycle*, which could explain the low correlation between stock and bond returns found when averaged over the cycle. As to bond markets, Beber and Brandt (2010) show that macroeconomic announcements are most important when they contain bad news for bond returns in expansions and, to a lesser extent, good news in contractions.

Ehrmann and Fratzscher (2005) show how macroeconomic news in the United States, Germany

[4] For example, Veronesi (1999) argues that stock prices overreact to bad news in good times and underreact to good news in bad times.

and the euro area impact *daily* US dollar-euro/DEM exchange rate developments (in the period 1993-2003). Not only US macroeconomic news are found to impact more than European news, in light of their greater timeliness, but also the effect of macroeconomic news appears to be magnified in periods of exchange rate volatility, or in the aftermath of a sudden change in the direction of the most recent string of news.

Furthermore, markets may follow with more attention particular macroeconomic variables according to the phase of the business cycle, as well as to the policy reaction function that markets deem likely (Bacchetta and van Wincoop, 2013). For example, in the late 1970s and early 1980s trade balance and unemployment were actively tracked by investors as they feared that US policymakers would respond to deficits or high unemployment with active monetary policies; during the 2000s, the introduction of US policy targets in terms of inflation and output gap has somewhat shifted the focus of markets to more timely and coincident indicators.

More recently Goldberg and Grisse (2013) document a significant time variation in the responses to news of yield curves and exchange rates from January 2000 through August 2011. They resort to *high-frequency data* to show how several indicators of the state of the economy (e.g. the level of policy rates, risk conditions) influence the estimated response.[5] In particular, US bond yields increase in response to "good news", but less so when risk is elevated. They interpret this as reflecting some combination of a market perceived financial stability objective for monetary policy and an influence of risk on the uncertainty associated with the link between macroeconomic announcements and the economic outlook.

Our paper builds upon this literature, and contributes along two dimensions. It explores a new dimension of state dependency, by considering how disagreement among analysts's forecasts can affect the response of asset prices to each specific macroeconomic surprise. [6] To this end, we consider two dimensions of heterogeneity: one specific to each macroeconomic announcement, the other related to a more "common" gauge of macroeconomic disagreement regarding the current state of the economy, extracted from the panel of indicator specific disagreements. Finally, we investigates whether this "state dependence" can be retraced to particular episodes of high/low volatility in financial markets.

2.1 A model with hetoregenous beliefs

Disagreement across forecasters on an upcoming macroeconomic announcement can depend on several factors: difference in information across agents, differential interpretation of the same public signals or difference in priors or models. However, disentangling these sources of disagreement

[5] According to the authors, "Risk conditions matter since they can capture the effects of uncertainty on the information content of news announcements, the interaction of monetary policy and financial stability objectives of central banks, and the effect of news announcements on the risk premium."

[6] Pasquariello and Vega (2007) provide a theoretical model for the role of order flow in US Treasury markets and show how the contemporaneous correlation between order flow and yield changes is *higher* when the disagreement among market participants is *high* and *public announcements* are *noisy*.

can be challenging. Patton and Timmermann (2010) exploit the term structure of forecasts from Consensus Forecasts and provide empirical support that heterogeneity in forecasters's information is not a major driver of observed disagreement, but rather that differences in priors matter the most.

The role of disagreement is explicitly modelled in Ottaviani and Sørensen (2015; henceforth OS), who introduce heterogenous beliefs in a general equilibrium model with learning. In the basic setup, with risk-neutral agents and absence of short sales, the price reaction to a public announcement is found to be decreasing in the degree of heterogeneity of beliefs (Proposition 3 in OS).[7]

In the OS model markets are complete and agents take positions on whether or not a binary event, E, is realized (say the economy will be in an expansion period). There are two state-contingent securities whose payoff is linked to the realization of a binary event $[E, R]$, say expansion and recession.[8] There is a continuum of risk-neutral agents, indexed by i, endowed with an amount of assets w_{i0}, constant across the two states. Agents exchange their assets in a competitive market, but cannot hold a negative amounts of any asset.

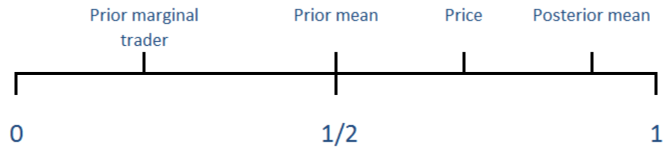
Agents' prior beliefs on the realization of the event E are heterogeneous, q_i, and described by a distribution of initial assets over individuals, $G(q)$, defined over the unit interval. Before trading takes place agents observe a public signal with likelihood ratio L for event E, and update their beliefs according to the Bayes rule. The competitive equilibrium price can be shown to depend on the public signal, as well as on the shape of the distribution of wealth (and hence beliefs) across individuals. OS show that the price underreacts to information (Proposition 2). Underreaction is measured relatively to the one associated to an hypothetical outside observer, whose prior beliefs are those of the average across agents' priors, and who updates his belief in a Bayesian fashion. In the model the equilibrium price does not correspond to the posterior of this "outside observer". Changes in the distribution of wealth across agents (with different prior beliefs) affect the equilibrium and the extent of the underreaction. In particular, more underreaction results when the distribution is less concentrated, that is when agents have greater heterogeneity in beliefs.

To gain intuition on the underlying mechanism, suppose the public signal is in favor of event E, i.e. $L > 1$. Agents will update their beliefs according to Bayes rule, with their posterior odds ratio in favor of E rising one for one with L. Those that were ex-ante optimist, and willing to invest all their wealth in asset E (agents are risk neutral and subject to a trading bound) will continue to do so, but will be able to buy fewer assets E as its price has increased. Similarly the pessimist will be able to buy more assets R, as its price has fallen. This would lead to an insufficient demand for assets E and excess demand for asset R. To balance the market some ex-ante pessimists have to "switch sides", becoming optimist, and seek to sell all their assets R in return for assets E. The effect of information is to move the marginal trader in a direction opposite to the price movement.

[7]The same result, not explicitly discussed in their paper, can be shown to hold also with with no position limits if agents are risk averse, with decreasing absolute risk aversion.

[8]For a more detailed description see the Appendix

Figure 1: Price Underreaction to information



Note: The Figure shows an example where beliefs are ex-ante centered in 0.5 and how the price does not behave like a posterior belief.
The marginal trader has a prior which is more pessimistic than the "average" market prior.

The degree of concentration of beliefs among agents is summarized by a single parameter γ, and we show how this affects the price reaction to the release of a public information signal.

The price elasticity to "information" is described by Figure 1: for a given likelihood ratio ($L = 6$) varying γ leads to different price responses. Greater heterogeneity in beliefs (higher γ), is associated with a smaller price response. This occurs because the marginal trader beliefs are further away from the prior mean and in a direction opposite to the information (more pessimistic). In contrast, when beliefs are very concentrated the marginal trader ex-ante beliefs are closer to the prior mean and accordingly also the price reaction almost parallels the shift in "average" beliefs.

Table 1 –The effect of heterogeneity on the price reaction

Concentration (γ)	Prior mean	Posterior mean	Prior std	Posterior std	Price	Prior marginal trader
0.5	0.5	0.68	0.38	0.35	0.645	0.232
3	0.5	0.84	0.14	0.08	0.793	0.390
10	0.5	0.86	0.04	0.02	0.836	0.459

This result is in stark contrast with the predictions of the learning model by Kim and Verrecchia (1991; henceforth KV), normally used to describe the assets price reaction to public announcements (e.g. Hautsch and Hess (2007) and Gilbert et al. (2010)). In that framework agents heterogeneity is driven by asymmetric information arising from agents observing a combination of private and public signals. The price reaction around a public announcement depends on two components: the precision of the announcement information and the precision of preannouncement information. The precision of preannouncement information reflects the ex-ante uncertainty regarding the risky asset returns, driven by the intrinsic precision on the fundamental and an "average" of agents' private signals precision. Disagreement enters the KV model to the extent that private signals have different precisions. This will lead agents to have have different conditional expectations on the asset payoff. In the KV model however disagreement affects the price in the same direction as overall uncertainty.

Two features determine these different predictions. First, unlike in the OS model, wealth effects are absent in the KV model since agents have CARA utility functions. Second, in the OS model agents' beliefs and the distribution of wealth across agents are explicitly linked, while in the KV model agents are heterogenous in a random fashion, as private information signals are uncorrelated across agents.

3 Data

3.1 Macroeconomic surprises and forecaster disagreement

We use the Bloomberg real-time data on the expectations and realizations of the most relevant U.S. macroeconomic indicators to estimate announcement surprises. The Bloomberg economic calendar contains a survey of economists forecasts for each US scheduled macroeconomic release. Surprises are defined as the difference between the median forecast and the actual announcement. Accordingly, we denote $S_{t,j}$ as the surprise at time t in the macroeconomic indicator j as:

$$S_{t,j} = \frac{R_{t,j} - F_{t,j}}{\hat{\sigma}_{S_j}} \ , \tag{1}$$

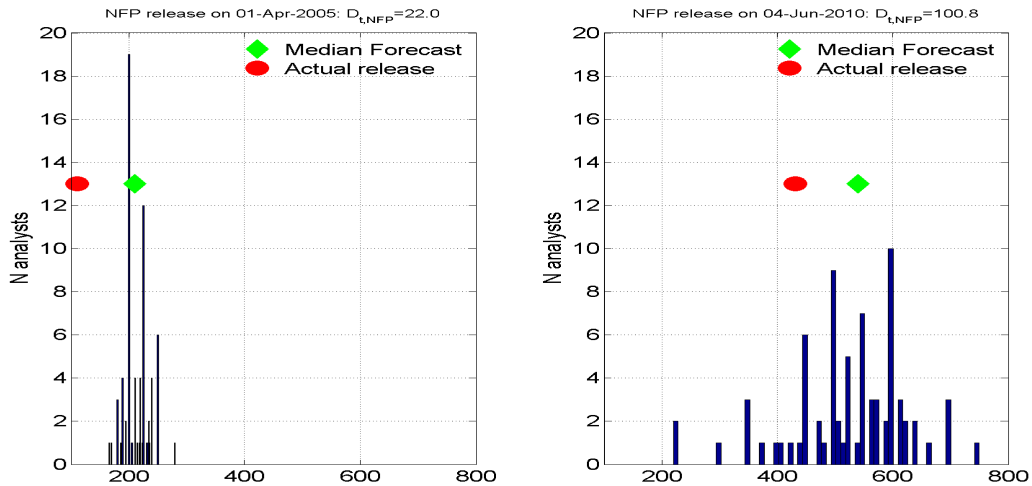
where $R_{t,j}$ is the announced value of the indicator j at the date of release t (in its first estimate), while $F_{t,j}$ is its Bloomberg median forecast.[9] To make surprises comparable across different indicators, they are scaled using their sample standard deviation, $\hat{\sigma}_{S_j}$. This way of proceeding is very common in the literature, see for example Fleming and Remolona (1999) and Gürkaynak et al. (2005).

In this work we also use a relatively unexplored dimension of the Bloomberg survey, consisting of the full set of individual forecasts for each indicator at each release. This dimension is very important as forecasters heterogeneity is found to vary considerably across releases: for example, the US non-farm payrolls releases on 1 April 2005 and 4 June 2010, surprised the median Bloomberg forecast by approximately the same amount, 100k and 109k on the downside respectively, but were associated with a very different distribution of analysts forecasts (see Figure 2). On the first (shown in the left panel), analysts's forecasts were tightly concentrated (with a standard deviation of 22k), while in the second (shown in the right panel), they were very dispersed, with a disagreement more than five-fold. This difference raises the question of whether the response of asset prices to a surprise may depend not only on its magnitude, but also on how forecasters beliefs are positioned ahead of the release. This hypothesis will be explored in Section 4.

We exploit this unique source of information obtained from Bloomberg to construct an announcement specific measure of forecaster heterogeneity. Accordingly, we define disagreement $Disag_{t,j}$

[9] The use of the median as a measure of the market expectations is robust to strategic behavior of some forecasters purposefully trying to the influence the average forecast. Bloomberg continuosly updates the economists forecasts and their median, starting from around two weeks ahead of the scheduled release until the actual announcement. We use the historical estimate of each macroeconomic announcement, even if most of the figures released are subject to revision in the following announcements.

Figure 2: Analysts' expectations on two non-farm payrolls (NFP) releases



Notes: The Figure shows the analysts' forecast of non-farm payrolls on 1 April 2005 (left panel) and 4 June 2010 (right panel); each bar indicates the number of analysts whose forecast is reported on the horizontal axis; the green diamond shows the median forecast implied by the bars and the red dot the actual release. Disagreement was 22k on 1 April 2005 and 100k on 4 June 2010.

among the N forecasters as the standard deviation of their forecasts, namely

$$Disag_{t,j} = \sqrt{\sum_{n=1}^{N} \frac{\left(F_{t,j,n} - \overline{F}_{t,j}\right)^2}{N}} \, , \tag{2}$$

where $F_{t,j,n}$ is the forecast of indicator j released at time t made by forecaster n, while the cross-section average among forecasters $\overline{F}_{t,j} = \frac{1}{N}\sum_{n=1}^{N} F_{t,j,n}$.[10]

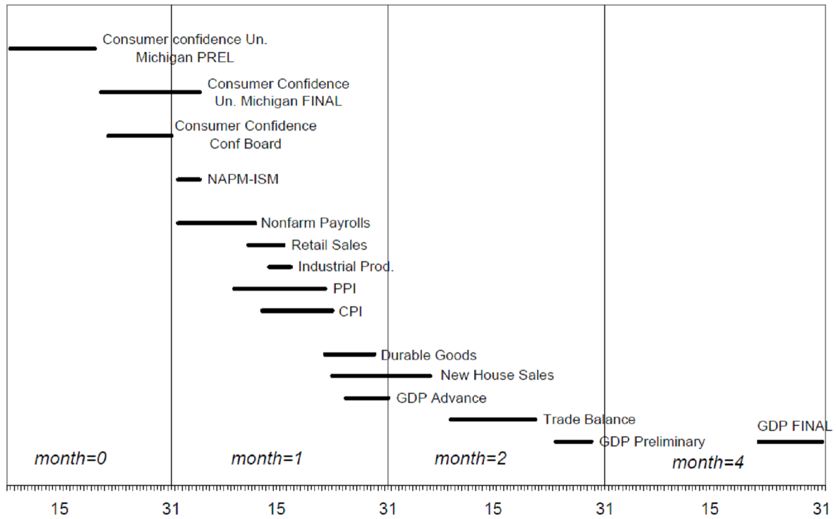
To the extent that each macroeconomic indicator is related to the underlying fundamentals, the corresponding disagreement measured by our proxy is related to the heterogeneity in beliefs on the current state of the economy. Nonetheless, disagreement regarding the current state of the economy is likely to be also an important driver of heterogeneity in the expectations regarding the future evolution as well. This mapping between short and long run disagreement is explored in Andrade et al. (2013), who construct a *term structure* of disagreement among professional forecasters for key macroeconomic variables surveyed in the Federal Reserve Survey of Professional Forecasters.

In our analyis we consider a set of 12 US key macroeconomic indicators. Our choice follows the selection adopted in the previous literature on macroeconomic news (Swanson and Williams, 2014a) and is corroborated by the large number of analysts which routinely post their forecasts on the Bloomberg platform.[11] These include three series on the US labor market (non-farm payrolls,

[10]Other measures of disagreement, such as the interquartile range or the mean absolute deviation have been considered.

[11]The literature on macroeconomic surprises has mostly focused on US scheduled releasese since they are the most

Figure 3: Time-line of the major US macroeconomic data releases



Notes: The Figure shows the days of the month when the release is announced with reference to month 0.

initial jobless claims and unemployment rate), one variable on the inflation rate (CPI core), five indicators on aggregate demand and production (retail sales, advance GDP, core durable consumer goods orders, capacity utilization, trade balance), and three sentiment indicators (ISM/NAPM, Michigan consumer confidence and the Conference Board leading indicator).

The time-line of the US data releases is shown in Figure 3 where each horizontal bar indicates the range of days when data are released with respect to the reference month labeled 0. For example, the ISM-NAPM index is generally released a few days after the reference month, the trade balance is released two months later, while the preliminary consumer confidence of the University of Michigan is published during the same month of reference. Some macroeconomic indicators are more relevant than others for markets, as they are more informative (more related to fundamentals), more timely (released closer to their reference period) and more precise (subject to smaller revisions). For example, the literature has evidenced that the most relevant surprise for markets is the one coming from the release of on non-farm payrolls, which is published the first Friday of the month successive to the reference period.As shown in Table 2, these macroeconomic indicators differ in terms of their volatility and average delay (timeliness). Their relative importance is also reflected by the number of economist disclosing their forecasts on the Bloomberg platform (last column of Table 1). Finally, also their exact time of release is relevant, especially to disentangle correctly their individual impact when using intra-day data.

closely followed by market participants and also span a longer period.

Table 2 –Main US scheduled macroeconomic releases

	frequency	source	release time	σ_R	avg. delay	#analysts
Nonfarm Payrolls	M	BLS	8:30	191.5	5	70
ISM/NAPM	M	ISM	10:00	5.9	2	64
Retail sales ex. autos	M	USCB	8:30	0.7	13	72
GDP (advance)	Q	BEA	8:30	2.1	29	63
Core durable orders	M	USCB	08:30	1.7	26	64
Consumer confidence	M	UnMich	10:00	11.2	-3	59
Initial jobless claims	W	DLS	8:30	75.6	6	36
Capacity utilization	M	Fed	9:15	3.2	16	58
Trade balance	M	USCB	8:30	11.4	44	63
Leading Indicator	M	CB	10:00	0.4	22	50
CPI Core	M	BLS	8:30	0.1	17	67
Unemployment rate	M	BLS	8:30	1.8	5	68

The Table reports descriptive information about the indicators analyzed. Frequency: (W= weekly, M=monthly, Q=quarterly); source: DLS=Department of Labor Statistics, BLS=Bureau of Labor Statistics, BEA=Bureau of Economic Analysis, USCB=US Census Bureau, Fed=Federal Reserve, ISM=Institute for Supply Management, Phil Fed=Philadelphia Federal Reserve, CB=Conference Board, UnMich=U.of Michigan Survey Research Centre; release time: in US Eastern Standard Time which corresponds to GMT-4; σ_R: the standard deviation of the indicator; delay: average delay, in days, with respect to the reference period; #analysts: the average number of analysts providing their forecast to Bloomberg.

When scaled in terms of the reference variable, the standard deviation of the surprise time-series (i.e. the root mean square forecast error of the Bloomberg median) varies widely across indicators (in Table 3, column: $\frac{\sigma_S}{\sigma_R}$): This provides a rough estimate of the *forecastability* of each series, and it ranges from 0.11 for the surprise for Capacity utilization to almost 1 for Core durable goods orders. This pattern is reflected also in the average dispersion surrounding each release, $\overline{Disag}$, when normalized for the original series standard deviation ($\frac{\overline{Disag}}{\sigma_R}$). Interestingly, the cross-sectional dispersion of analysts forecasts is on average one half of the surprise standard deviation.

The actual release often falls outside the support of the distribution of forecasters expectations, suggesting that beyond the difficulty to forecast a particular indicator, there also may be a general tendency to cluster around the consensus median, in order to avoid extreme deviations from the average forecast. As shown in Table 3 (column *Outside*) considerable differences exist across indicators: more than a quarter of the Non-farm payrolls releases fall outside the analysts range of forecasts, compared to less than 5 per cent for the advance GDP growth.

Table 3 –Surprises and disagreement: descriptive statistics

	$\frac{\sigma_S}{\sigma_R}$	$\frac{\overline{Disag}}{\sigma_R}$	Out	R^2_{Disagr}	t_{S-lag}
Nonfarm Payrolls	0.40	0.17	0.28	0.50	6.35
ISM/NAPM	0.34	0.16	0.23	0.24	5.52
Retail sales ex. autos	0.70	0.34	0.16	0.45	5.74
GDP (advance)	0.34	0.21	0.04	0.40	0.02
Core durable orders	0.97	0.40	0.35	0.27	5.11
Consumer confidence	0.34	0.13	0.32	0.21	5.62
Initial jobless claims	0.26	0.11	0.28	0.49	11.84
Capacity utilization	0.11	0.07	0.11	0.30	2.52
Trade balance	0.30	0.11	0.31	0.12	3.35
Leading Indicators	0.42	0.33	0.04	0.08	-0.86
CPI Core	0.97	0.53	0.08	0.21	3.34
Unemployment rate	0.09	0.04	0.19	0.24	5.41

The Table reports descriptive features of the macroeconomic surprises and the corresponding forecaster disagreement. For each indicator, we report: $\frac{\sigma_S}{\sigma_R}$ the standard deviation of the Bloomberg surprise series to the one of the original release; $\frac{\overline{Disag}}{\sigma_R}$: the average cross-sectional dispersion of forecasters $\overline{Disag}$, as a ratio of the original series standard deviation (σ_R); Out: the fraction of times the actual release falls outside the range of forecasters expectations; R^2_{Disagr}: the R^2 from the regression of forecasters cross-sectional dispersion on its lag and the absolute value of the most recent surprise; t_{S-lag} the t-statistics for the past surprise variable.

For most indicators, disagreement displays a significant positive serial correlation. To better characterize this property in Table 3 we report, the R^2 of the regression

$$Disag_{t,j} = \alpha + \beta Disag_{t-1,j} + \gamma |S_{t-1,j}| + u_t$$

of disagreement on its lag, as well as on the size of the most recent absolute value of the surprise. This should provide some indication as to whether disagreement about the upcoming release of an indicator, as proxied by the standard deviation of expectations across the individual forecasters, tends to increase following a particularly large surprise in the earlier release. Overall, we find that, indeed, past surprises amplify the forecaster heterogeneity in the subsequent scheduled release (t_{S-lag}, last column of Table 3).

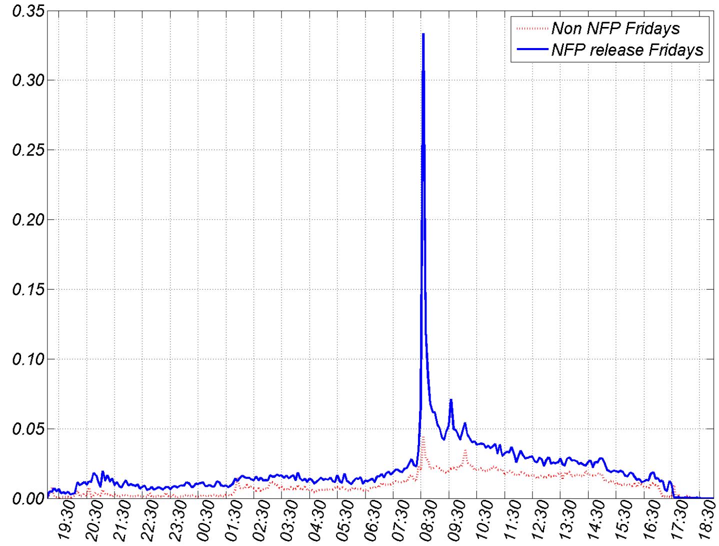
3.2 Asset prices

Throughout the paper we use our novel dataset on US macroeconomic surprises and disagreement in conjunction with the US dollar-euro exchange rate and long term interest rates in the US and in Germany. The data are described in some detail in the Appendix.

To introduce the evidence on the role of disagreement in explaining the asset price response to macroeconomic news, we first report the mean absolute percentage change of the 5-minute 10-year Treasury price in the days (Friday) when the non-farm payrolls are released (Figure 4) – on the time axis we report the Eastern time (EST) hours (which corresponds to four hours before GMT, i.e. 12:30 GMT during EST daylight time, and to five hours before GMT, i.e. to 13:30 GMT, during EST standard time). The absolute change in the 10-year Treasury yield is large in correspondence of the non-farm payrolls release at 8:30, and the change slowly vanishes, suggesting some trading activity in response to the news also in the following hours. In contrast, on the Fridays when no non-farm-payrolls are relased the same prices do not exhibit a marked spike at 8:30EST, and also display a smaller volatility in the rest of the day.

Thus, on average, we confirm the standard finding that indeed the non-farm payroll announcement induces a sizable change in the 10-year Treasury yield. In the next sections we will qualify this result and relate it to the magnitude of the surprise as well to our proxy of analyst heterogeneity in beliefs.

Figure 4: Absolute average impact of non-farm payrolls on 10-year Treasury yield



Note: the Figure shows the absolute percentage change in the first logarithmic difference of the 10-year Treasury price during the days of release of non-farm payrolls; daytime on the horizontal axis. The spike corresponds to 8:30 EST, e.g. 12:30 GMT-4 hours.

4 Results

We evaluate the sensitivity of financial markets to macroeconomic news conditional on the disagreement among forecasters, measured for each individual release. Specifically, we estimate the following equations

$$\Delta y_t = \alpha + \sum_j \beta_j \cdot S_{t,j} + u_t \tag{3}$$

$$\Delta y_t = \alpha + \sum_j \gamma_j \cdot S_{t,j} \cdot I^L_{t,j} + \sum_j \delta_j \cdot S_{t,j} \cdot (1 - I^L_{t,j}) + u_t \tag{4}$$

where Δy_t is the change in the asset yield (first difference of interest rates and first logarithmic difference of the exchange rate), $S_{t,j}$ is the j surprise time series. In specification (3) disagreement is not considered, while in the second specification (4) we distinguish between high and low periods of disagreement, according to a simple binary rule:

$$I^L_{t,i} = \begin{cases} 1, & \text{if } cdf(Disag_{t,i}) < Upper \\ 0, & \text{otherwise} \end{cases} \tag{5}$$

The regime dummies are series specific ($I_{t,j}$), with a common threshold $Upper$: in our estimation we use the $66th$ percentile of the empirical distribution of the disagreement series for the j_{th} indicator (the results do not depend on the particular choice of the percentile). We run the regressions (3) and (4) simultaneously with all the surprise variables, while restricting the estimation on only for days with at least one macroeconomic release. We name the first regression the *baseline* regression.

Results for the US, German 10-year bond yields and for the USD/EUR exchange rate for the baseline regression (3) are shown in the first two columns of Table 4. The top panel of Table 4 shows that US long-term rates are impacted by all news with the exception of capacity utilization, Conference Board leading indicator, CPI core, and unemployment rate. In general, the size of the impact is large even at the 1-day horizon. The top panel also gives a pecking order of news: non-farm payrolls have the largest impact on US long-term yields – e.g. a one per cent unexpected change in non-farm payrolls increases 10-year yields by 4.23 basis points – followed by the ISM index, retail sales, the advance GDP rate of growth, and consumer confidence. Initial jobless claims and trade balance show a significant impact on long-term yields.

The impact of non-farm payrolls, ISM, retail sales and initial claims is also large on German 10-year yields, while the other US surprises do not have a significant impact. Similarly, the USD/EUR exchange rate shows a large sensitivity to surprises upon the release of non-farm payrolls, ISM, and GDP growth – e.g. an unexpected change of 1 per cent in non-farm payrolls determines an appreciation of the US dollar against the euro by 0.30 per cent. All in all, the USD/EUR exchange rate seems reactive to macroeconomic surprises, with a particular attention to the labor market (non-farm payrolls) followed by the hard and soft data from the real economy (GDP, ISM and retail sales).

Results for the two-disagreement regimes regression (4) are shown in the last four columns of Table 4. With low/high disagreement US and German 10-year yields show a larger/smaller sensitivity to surprises; thus, in general we obtain $|\widehat{\gamma}| > \left|\widehat{\beta}\right| > \left|\widehat{\delta}\right|$. More interestingly, we find that some variables turn out to be significant conditional on the disagreement regime; for example, the ISM, consumer confidence, capacity utilization and GDP growth have a large and significant impact on 10-year interest rates during low-disagreement regimes, but not during high-disagreement regimes. In Table 4 we also report the results of the joint hypothesis test of an asymmetric response of asset prices to news across low and high disagreement regimes. We cannot reject the null of no asymmetry for US interest rates, German rates and exchange rates, respectively at the 10%, 20% and 2% level.

In other words, investors react more to scheduled surprises when their beliefs are more concentrated and adjust their portfolios accordingly; on the other hand, an environment with wider dispersion of expectations tends to dampen the impact of macroeconomic surprises on asset prices, perhaps because it increases investors' risk aversion. As pointed out by Carlin et al. (2014), when participants in a market disagree with each other, an investor taking a position based on his unique expectations could face a greater risk of being wrong. This introduces a new form of trading risk (or adverse-selection risk) which is different from the traditional types of market risks that are priced in asset values.

Our results suggests that interpreting disagreement as reflecting the market "prior uncertainty" regarding the announcement outcome may be inappropriate. In models where investors heterogeneous beliefs plays no role, larger uncertainty exacerbates market volatility, as a result of investors revising more rapidly their beliefs, in line with standard Bayesian learning framework where the price update is a decreasing function of prior precision. In particular, Hautsch and Hess (2007), who investigate how release-specific precision measures affect the price impact of the non-farm payrolls release on the US T-bond futures, find that the price impact of more precise information is significantly stronger. In their model the precision of the information is modeled as a decreasing function of prior disagreement, measured by the standard deviation of analysts' forecasts, and an increasing function of the precision of the individual announcement.[12] In their empirical analysis the latter is estimated from the revision process of the non-farm payrolls.

Ottaviani and Sørensen (2015) analyze how asset prices react to information when agents have heterogenous prior beliefs. In their model, albeit specific to binary markets, the asset price response to information critically depends on the degree of dispersion of prior beliefs. In particular, when beliefs heterogeneity is *more pronounced* the price response to news is shown to be more muted, a prediction in line with what we find in our empirical exercise.[13]

Our results also resonate with the recent finding of Goldberg and Grisse (2013), in particular that the level of risk aversion, measured by the VIX index, negatively affects the sensitivity of bond yields and exchange rates to US macroeconomic news. To what extent can we relate our measures based on survey based indicators of disagreement to those of aggregate uncertainty, or "risk"? This is discussed in the next section.

[12] Their learning framework is the following. Let the price, p, depend on a fundamental, say X, $p = \nu E(X)$. Agents have a prior on X which is $X \sim N\left(\mu_F, \frac{1}{\rho_F}\right)$ where ρ_F is the precision of the agents' prior – e.g. the dispersion of beliefs of analysts. Signal on this variable comes as $A = X + \varepsilon$ where $\varepsilon \sim N\left(0, \frac{1}{\rho_\varepsilon}\right)$ and $A \sim N\left(\mu_F, \frac{1}{\rho_\varepsilon + \rho_F}\right)$. The impact of a surprise A on prices is $[(P|A) - E(P)] = \mu_F + (A - \mu_F)\frac{\rho_\varepsilon}{\rho_\varepsilon + \rho_F}$. Adding and subtracting $A - \mu_F$ from the right side yields $[(P|A) - E(P)] = \underbrace{(A - \mu_F)}_{\text{pure surprise}} - \underbrace{(A - \mu_F)\frac{\rho_F}{\rho_\varepsilon + \rho_F}}_{\text{surprise/beliefs interaction}} \underbrace{- \mu_F \frac{\rho_F}{\rho_\varepsilon + \rho_F}}_{\text{pure beliefs}}$. If we let $\mu_F = 0$, the equation reduces to $[(P|A) - E(P)] = A - A\frac{\rho_F}{\rho_\varepsilon + \rho_F}$. This equation can be tested econometrically by defining $(A - \mu_F)$ as the surprise and $\frac{1}{\rho_\varepsilon/\rho_F + 1} = (D/\rho_F + 1)^{-1}$ where $D \equiv \rho_\varepsilon$ is the inverse of the beliefs' dispersion; thus the model is specified by $\Delta p_t = \alpha + \beta \cdot (A - \mu_F)_t + \gamma \cdot (A - \mu_F)_t \cdot D_t + \delta \cdot D_t + u_t$ where by construction $\beta > 0$ and $\gamma, \delta < 0$.

[13] See Proposition 3 of their paper.)

Table 4 – Baseline regression and regression with disagreement regimes

	10-year US Treasury yield					
	Baseline		2-disagreement regimes			
			low disagr.		high disagr.	
Non Farm Payrolls	4.23	(5.03)	6.31	(7.29)	2.23	(1.97)
ISM Manufacturing	2.65	(4.78)	3.54	(5.58)	1.39	(1.55)
Retail sales ex.autos	2.58	(4.72)	2.94	(4.22)	2.18	(2.56)
GDP (advance)	1.85	(1.94)	2.98	(2.20)	0.54	(0.45)
Core durable goods orders	1.60	(3.12)	0.92	(1.25)	2.25	(3.58)
Consumer confidence	1.52	(3.02)	1.80	(3.14)	1.01	(1.14)
Initial claims	-1.48	(-6.01)	-1.37	(-4.08)	-1.55	(-4.52)
Capacity utilization	0.73	(1.28)	1.59	(2.24)	-0.27	(-0.36)
Trade balance	1.01	(2.30)	0.97	(1.84)	1.08	(1.50)
CB leading indicator	0.77	(1.64)	-0.05	(-0.07)	1.07	(1.85)
CPI Core	0.53	(0.86)	0.59	(1.09)	0.61	(0.42)
Unemployment rate	-0.25	(-0.43)	-0.26	(-0.34)	-0.39	(-0.40)
# Observations	1,626					1,626
H_0: $\delta_j = \gamma_j, \forall j$, p-value						0.08
R^2	0.098					0.113
	10-year German Bund yield					
	Baseline		2 disagreement regimes			
			low disagr.		high disagr.	
Non Farm Payrolls	2.12	(3.67)	3.45	(5.78)	0.85	(1.11)
ISM Manufacturing	1.23	(3.33)	1.41	(2.61)	1.03	(2.19)
Retail sales ex.autos	1.40	(3.24)	1.70	(2.92)	1.05	(1.60)
GDP (advance)	0.96	(1.32)	1.32	(1.09)	0.60	(0.76)
Core durable goods orders	0.12	(0.33)	0.09	(0.17)	0.18	(0.43)
Consumer confidence	0.70	(1.86)	0.86	(2.13)	0.40	(0.60)
Initial claims	-0.95	(-4.84)	-1.12	(-4.48)	-0.82	(-2.88)
Capacity utilization	0.58	(1.47)	1.12	(2.48)	-0.05	(-0.07)
Trade balance	-0.11	(-0.31)	0.03	(0.06)	-0.11	(-0.18)
CB leading indicator	0.10	(0.33)	0.20	(0.32)	0.06	(0.17)
CPI Core	0.84	(1.88)	0.60	(1.15)	1.36	(1.67)
Unemployment rate	-0.11	(-0.25)	0.12	(0.20)	-0.47	(-0.74)
# Observations	1,572					1,572
H_0: $\delta_j = \gamma_j, \forall j$, p-value						0.20
R^2	0.051					0.061
	USD/EUR exchange rate					
	Baseline		2 disagreement regimes			
			low disagr.		high disagr.	
Non Farm Payrolls	-0.30	(-5.04)	-0.43	(-5.95)	-0.18	(-2.33)
ISM Manufacturing	-0.15	(-2.78)	-0.10	(-1.34)	-0.22	(-3.40)
Retail sales ex.autos	-0.07	(-1.65)	-0.14	(-2.46)	0.00	(-0.06)
GDP (advance)	-0.25	(-2.07)	-0.12	(-0.82)	-0.41	(-2.66)
Core durable goods orders	-0.03	(-0.88)	-0.04	(-0.83)	-0.03	(-0.52)
Consumer confidence	-0.03	(-0.63)	-0.11	(-1.69)	0.07	(0.89)
Initial claims	0.00	(-0.14)	0.00	(0.15)	-0.02	(-0.44)
Capacity utilization	-0.03	(-0.39)	-0.01	(-0.24)	-0.04	(-0.30)
Trade balance	-0.08	(-1.59)	-0.13	(-2.14)	-0.04	(-0.43)
CB leading indicator	-0.01	(-0.17)	-0.20	(-2.36)	0.06	(0.72)
CPI Core	-0.01	(-0.17)	-0.06	(-0.97)	0.09	(1.44)
Unemployment rate	0.09	(1.79)	0.10	(1.62)	0.09	(1.07)
# Observations	1,628					1,628
H_0: $\delta_j = \gamma_j, \forall j$, p-value						0.02
R^2	0.0268					0.048

The Table reports the coefficients and the White corrected t-statistics (in parenthesis) of regression (3): $\Delta y_t = \alpha + \beta \cdot S_t + u_t$ – baseline regression, and of regression (4): $\Delta y_t = \alpha + \sum_j \gamma_j \cdot S_{t,j} \cdot I^L_{t,j} + \sum_j \delta_j \cdot S_{t,j} \cdot (1 - I^L_{t,j}) + u_t$ – 2 disagreement regimes. H_0 $\delta_j = \gamma_j, \forall j$ is the p-value of the joint test for an asymmetric response to high and low disagreement regimes.

5 Aggregate uncertainty and disagreement

Disagreement may be very different from uncertainty, as shown by Zarnowitz and Lambros (1987). The empirical literature has largely debated over the approximation of uncertainty with disagreement – Bomberger (1996), Giordani and Söderlind (2003) – without establishing a clear and theoretical link between the two. Gürkaynak and Wolfers (2005) argue that for the key US macroeconomic news the relationship between disagreement and uncertainty is rather weak. Similarly, Lahiri and Sheng (2010) argue that the disagreement across forecasters surveyed in the Federal Reserve Survey of professional forecasters is a rather unsatisfactory measure of state uncertainty. Nonetheless, even the widely known measure of macroeconomic uncertainty proposed by Bloom (2009) draws on the same survey to compute the dispersion in the forecasts as proxies for uncertainty about monetary policy and about government budget policies.

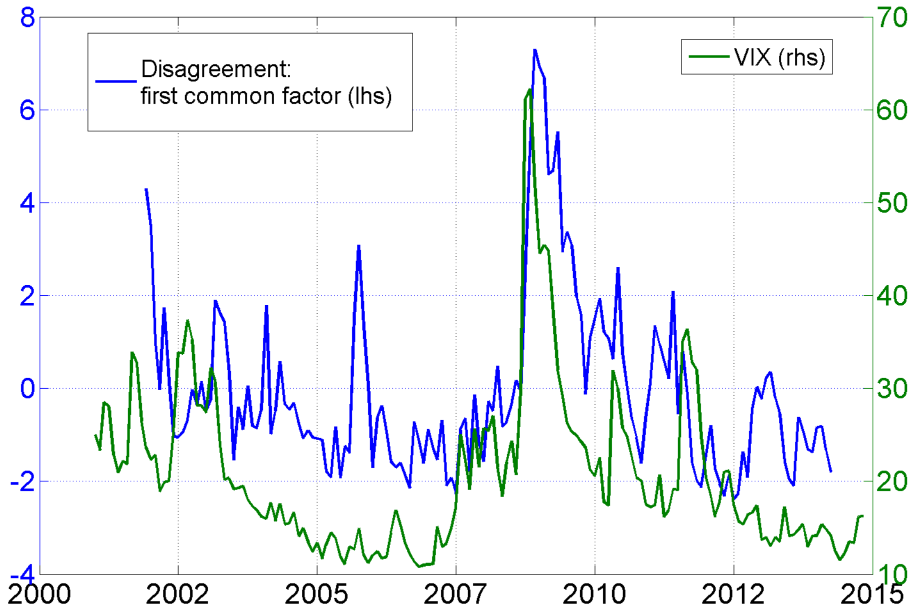
From a theoretical point of view the distinction between disagreement and uncertainty matters, since models featuring some heterogeneity in beliefs can explain several puzzling phenomena for financial markets (Xiong and Yan, 2010). In standard models, without heterogeneous beliefs, agents follow a Bayesian framework to update their expectations about the path of future return on assets. In this setup the financial market sensitivity on day t to a macroeconomic data release depends on the uncertainty of the surprise component of that data release and the day $t-1$ prior variance of the variable being forecast. In general, the theory states that if the market's prior dispersion is very small, then market participants have a great deal of confidence in their expectation of the future short-term interest rate and will respond relatively little to any data release on date t. On the other hand, if the market's date $t-1$ prior dispersion is very large, then markets will respond much more strongly to any news on date t. Swanson and Williams (2014a) and Swanson and Williams (2014b) use a measure of uncertainty, derived from the short-term interest rate option market, to show that it explains well the time-variation in the estimated news response of bond rates.

The empirical findings in the previous Section, that forecaster heterogeneity impacts asset prices –and in a direction opposite to the prediction of the standard model– bears the questions of whether the observed heterogeneity can be traced back to some common factor, or it represents a purely idiosyncratic feature of our dataset. As agents formulate expectations conditional on the current state of the economy, differences in opinion regarding the various data releases should also be mapped into disagreement regarding the expected path of the overall economy.

To answer this question, we exploit the panel of time series on each indicator disagreement and investigate their degree of co-movement. If individual disagreement has a factor structure, the weights can be defined by the eigenvector corresponding to the largest eigenvalue of the covariance matrix of the matrix of individual disagreement.[14] Specifically, we resort to principal components analysis (PCA) on the contemporaneous correlation matrix of the individual disagreement time

[14] A similar assumption is made by Jurado et al. (2013) in constructing a measure of aggregate uncertainty, obtained from aggregating the unforecastable component of a large number of economic indicators

Figure 5: VIX index and disagreement common factor



Note: the Figure shows the VIX index (rhs) and the first PCA factor, $Disag_t^c$, obtained from surprise-specific disagreements (lhs).

series; in particular, we extract the first PCA factor from the cross section of disagreement.[15]. To compute the correlation matrices we need to address two peculiarities of the macroeconomic surprises time series: first, since not all indicators are released on the same day, the dataset is heavily unbalanced at a daily frequency; second, some of the series are not available over the full sample. We apply a procedure similar to Beber and Brandt (2010), filling forward missing data and sub-sampling the data at a lower frequency, so to overcome the local constancy of some of the series. The first common-disagreement PCA factor explains more than 1/3 of the overall variation in the panel: however some series display greater comovement, more so the non-farm payrolls (68%) and the core retail sales series (78%); much less the trade balance and the CPI (21%) and the unemployment rate (31%).

Finally, we then run our 2-regime regressions (4) in Section 4, using this this common disagreement factor, $Disag_t^c$, extracted from the panel, in place of the the regime-specific indicators based on $Disag_{t,j}$. The results from this modified specification continue to confirm our findings of an asymmetric response of asset prices to macroeconomic news, also according to this proxy of economy-wide "disagreement".[16]

[15]A similar approach is followed also in Buraschi and Whelan (2012)

[16]The results are available upon request from the authors.

To illustrate how our measure of disagreement on "current macroeconomic conditions" can be related to more established indicators we plot the first PCA factor with the VIX index (Figure 5). Even if our disagreement factor is extracted from an entirely unrelated information set, the two time series display a striking co-movement, with a sample correlation of 0.63 at a monthly frequency in the period 2001-2014. This finding indicates that, since the VIX index can be interpreted not only as a measure of stock market (economic) uncertainty, but also as reflecting the evolution of global risk aversion (Bekaert et al., 2010), our aggregate measure of disagreement may provide some additional information on the evolution of the latter. This would also be in line with the most recent theoretical literature emphasizing the importance of accounting for divergence of beliefs, which suggests that investors should be compensated for bearing trading risk or the risk due to adverse selection when disagreement arises.

5.1 The time variation of sensitivity

The sensitivity in the response of bond and exchange rates documented in Section 4 could stem from other underlying forces, in particular from changes in the monetary policy environment, which could in turn affect the general disagreement regarding the time profile of interest rates and exchange rates. In this section, we explore whether our results depend on the choice of a particular time period. To this end, we need to allow for some time variation in the coefficients of our regressions. In doing this, we do not follow Faust et al. (2007) in allowing for time variation in all coefficients because in our multivariate framework we would encounter a degrees of freedom problem. Rather, following the approach of Swanson and Williams (2014a), we limit the time variation to that of a "single index" of macroeconomic news, while constraining the relative importance of each indicator to be stable over time. Specifically, we estimate

$$\Delta y_t \quad = \quad \alpha + \sum_{\tau} D_\tau \phi^{L,\tau} \sum_{j} \gamma_j \cdot S_{t,j} \cdot I^L_{t,j} + \sum_{\tau} D_\tau \phi^{H,\tau} \sum_{j} \delta_j \cdot S_{t,j} \cdot (1 - I^L_{t,j}) + u_t \tag{6}$$

with non-linear least squares. Here $\phi^{L,\tau}$ and $\phi^{H,\tau}$ are the coefficients on the time dummies D_τ, defining three subperiods ($\tau = 1, 2, 3$): the tranquil period ($\tau = 1$) – e.g. from January 1999 to July 2007 – the crisis period ($\tau = 2$) – e.g. from August 2007 to March 2009 – and the implementation and the unravelling of the unconventional US monetary policy period $\tau = 3$– from April 2009 to June 2014 – which partially overlaps with the euro-area debt crisis.

In equation (6) we assume that the vector of coefficients γ and δ remain constant across time periods as only the (regime specific) sensitivity to surprises – e.g. ϕ^{L,τ_i} and ϕ^{H,τ_i} – changes.

The coefficients $\phi^{L,\tau}$ and $\phi^{H,\tau}$ provide a measure of the changing sensitivity of asset markets to a single index of US macroeconomic surprises. In other words, they allow to measure the asymmetric response of markets to surprises during different phases of the business cycle (e.g. tranquil vs. crisis period) as well as during the implementation of unconventional monetary policies in the United States. In order to identify the estimates we posit $\phi^{L,1} = \phi^{H,1} \equiv 1$ – e.g. we define the "tranquil period" as the reference one to characterize the impact of surprises. This allows the comparison of the sensitivity across periods. As in Section 4, we perform a joint test of asymmetry of the response, across regimes. When allowing for time variation, the joint test is for the null of $\phi^{H,\tau}\delta_j = \phi^{L,\tau}\gamma_j$, $\forall j$. We find that the null of no asymmetry is rejected for interest rates, during tranquil and QE periods, while it cannot be rejected during the crisis years. In contrast, for exchange rates the null is rejected only during the tranquil period. A synthesis of results is shown in Table 5 while all the estimates are presented in Tables A.1-A.3 in the Appendix.

Table 5 – Time-varying sensitivity of markets to surprises

	10-year US yield			10-year German yield			USD/EUR		
	tranquil	crisis	QE	tranquil	crisis	QE	tranquil	crisis	QE
ϕ^{L,τ_i}	1	0.88***	0.92***	1	0.88***	1.17***	1	0.77***	0.16
ϕ^{H,τ_i}	1	0.98***	0.65*	1	1.24***	-1.06*	1	1.19***	0.36

The Table reports the sensitivity of surprises to US macroeconomic news estimated by equation (6). The sensitivity coefficient is normalized to 1 in the tranquil periods for both the low-disagreement and the high-disagreement regime in order to identify the coefficients. The tranquil period runs from January 1999 to July 2007, the crisis period from August 2007 to March 2009, and the QE period – e.g. the implementation and the unraveling of the unconventional US monetary policy period – from April 2009 to June 2014. ***/**/* indicates that coefficient are significant at the 1/5/10 per cent level. The values in the Table correspond to the coefficients of the last four columns of Tables A.1-A.3.

Results show that 10-year US Treasury yields respond to US macroeconomic surprises but their impact decreases during the QE period. Moreover, there is not much difference between the tranquil and QE period (1 vs. 0.92) in a low-disagreement regime, while the impact is much weaker during QE with high disagreement (1 vs. 0.65). Finally, it turns out, despite not being able to reject (joint) symmetry, that the sensitivity is larger with high disagreement than with low disagreement (0.98 vs. 0.88) in the crisis period, while the order is inverted during the QE period. In contrast, for 10-year German yields there is evidence of a more pronounced time variation, especially during the QE period. Furthermore, the reaction has the opposite sign across the two disagreement regimes, signaling a potential portfolio shift of investors from the US bond to the German bond market especially when disagreement is high.

The sensitivity changes dramatically for the USD/EUR exchange rate. In general the USD/EUR exchange rate does not react to surprises during the QE period while it reacts to them during the crisis and the order goes from low to high disagreement (0.77 vs. 1.19). Even if there is no strong evidence of asymmetry between disagreement regimes during the crisis and QE, the implementation of non-standard measures by the US monetary authorities may have attenuated the exchange rates response to macroeconomic fundamentals.

With reference to exchange rates, our framework allows to evaluate indirectly the scapegoat theory of exchange rates (Bacchetta and van Wincoop, 2013); this theory suggests that market participants may assign time-varying weight to economic fundamentals, which are chosen as scapegoats to explain exchange rate fluctuations. Our results show that exchange rates are less linked to macroeconomic news than bond yields but that this link changes over time being stronger during tranquil periods and, *ceteris paribus*, stronger when disagreement among forecasters is lower.

5.2 Robustness checks

As suggested by Gürkaynak and Wright (2013), the window around the data release for the event type study we have performed should be as small as possible. Indeed, the evidence from Andersen et al. (2003) suggests that the jump in conditional mean following a news announcement occurs within about 10 minutes. To assess the robustness of our results regarding the role of disagreement in the response to the release of macroeconomic news, we also perform the analysis using intraday data at the 5-minute frequency.[17] We consider, for the sake of brevity, the response to a non-farm payrolls surprise, which is the most important piece of macroeconomic news in our set of indicators.[18]

We report the average cumulative impact of the news on non-farm payrolls on the 5-minute returns of the USD/EUR exchange rate (Figure 5, left panel) and 10-year US Treasury (Figure 5, right panel). Results with intra-day data display an even more striking asymmetry between high and low disagreement regimes. What emerges is that, following a positive non-farm payrolls data surprise, bond and USD/EUR returns indeed exhibit a dip which is not offset in the following hours (time axis extends to three hours after the 8.30 a.m. non-farm payrolls announcement).

Furthermore, the average impact of the non-farm payrolls surprise on the intraday 10-year Treasury and the USD/EUR is of the same magnitude as that estimated at daily frequency. A unit (stardardized) positive surprise in the non-farm payrolls is associated with a fall in the 10-year Treasury price by 0.3 per cent in the intraday estimates; this, assuming an average duration for 10-year Treasuries of 7 years, corresponds to a positive change in 10-Treasury yields of around $-0.3 \div 7 \times 100 \simeq 4$ basis points, which is in line with that from the baseline regression 3 on daily data. Similar results hold for the regressions on the the USD/EUR exchange rate: the same non-farm payrolls surprise determines an appreciation of the exchange rate of around 0.30 percentage points. Most of the adjustment occurs in the first hour after the announcement, and persists in the following 2 hours. In general, the impact with low disagreement is twice as large as that with high disagreement (for the 10-year Treasury the impact is three times larger). This confirms that the analysis performed in Section 4 with daily data provides a reliable picture of the impact of

[17] Currently available only for the US Treasury price and the USD/EUR exchange rate.
[18] Similar results are found for the other macroeconomic data releases considered here.

Figure 6: Impact on USD/EUR

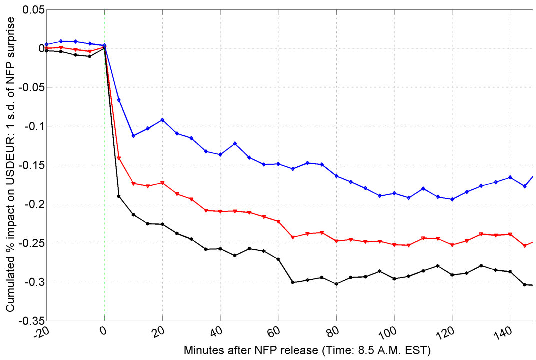
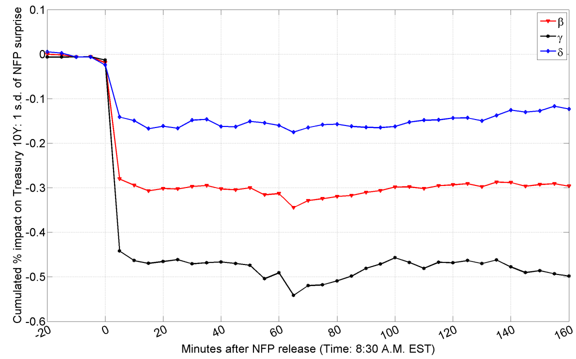


Figure 7: Impact on 10-year Treasury price



The Figure reports the cumulative 5-minutes impact of non-farm payrolls surprises on returns of the USD/EUR exchange rate and 10-year Treasury (basis points) during the first 160 minutes after the non-farm payrolls data release. Estimates from regressions only for the days with a non-farm payrolls release from January 2000 to June 2014. The red line shows the coefficients βs obtained from the regression $(y_{8:30+\tau} - y_{8:30}) = \alpha + \beta \cdot S^{NFP}_{8:30} + u$, where $y$is the logarithm of the USD/EUR exchange rate in the left panel and the level of the 10-year Treasury yield in the right panel, $S^{NFP}_{8:30}$ are the news on non-farm payrolls released at 8:30 EST, and τ is the time after 8:30 EST, which is shown on the horizontal axis; the black and the blue lines report the coefficient γs and δs obtained from the regression $(y_{8:30+\tau} - y_{8:30}) = \alpha + \gamma \cdot (S^{NFP}_{8:30} \cdot I^{NFP,L}_{8:30}) + \delta \cdot (S^{NFP}_{8:30} \cdot I^{NFP,H}_{8:30}) + u$, where $I^{NFP,L}_{8:30}$and $I^{NFP,H}_{8:30}$are the disagreement dummies for low and high disagreement on non-farm payrolls defined in (2).

macroeconomic news on yields and exchange rates, also conditioning on the particular disagreement regime.

6 Conclusion

We analyze the impact of macroeconomic surprises and individual heterogeneity on exchange rates and long-term interest rates in the United States and in Germany from 1999 to 2014. The time span allows the comparison of surprises impact during normal times, during the global financial crisis, and during the implementation of the unconventional measures of monetary policy and forward guidance in the United States and the euro area.

Our work contributes to the recent literature on the impact of macroeconomic surprises on asset prices by introducing a measure of macroeconomic disagreement obtained from the Bloomberg surveys of forecasters; we reconcile the results obtained with data at the daily frequency with those obtained with intra-day data frequency. We show that disagreement on each single macroeconomic variable is made of an idiosyncratic and a common component. The latter explains most of the variation of some particularly informative macroeconomic data, such as non-farm payrolls and retail sales. Moreover, during several periods our measure of economy wide disagreement displays a strong co-movement with the VIX index.

Our contribution is twofold. First, we document how macroeconomic surprises impact differently long-term yield and foreign exchange rates during low-disagreement and high-disagreement regimes.

This result is novel, and unveils a new dimension of state dependence in the response of asset prices to macroeconomic surprises. Our results suggests that interpreting disagreement as reflecting the market "prior uncertainty" regarding the announcement outcome may be inappropriate. They also underscore the relevance of theoretical models which explicitly account for heterogenous beliefs to correctly predict the observed pattern of price responses.

Finally, we estimate the time-varying sensitivity of asset prices to macroeconomic surprises by splitting our sample in a tranquil, crisis, and an unconventional monetary policy period. We show that unconventional monetary policies tended to dampen the response of the USD/EUR to US macroeconomic news, while they left virtually unaffected the response of long-term interest rates. This results may be reconciled with the uncovered interest parity theory as changes in the USD/EUR exchange rate may be constrained when the short-term rates in the two areas are close to the zero-lower bound and short-horizon foreign exchange risk premia cannot change. Conversely, long-term rates and their term premia components continued to respond with the same intensity to the evolution of the macroeconomic outlook even during the crisis and the QE phase. We leave to future research the analysis of the interaction between US and euro-area surprises and their impact on financial markets.

References

Andersen, T. G., Bollerslev, T., Diebold, F. X., & Vega, C. (2003). Micro effects of macro announcements: Real-time price discovery in foreign exchange. *American Economic Review*, (2004/19).

Andersen, T. G., Bollerslev, T., Diebold, F. X., & Vega, C. (2007). Real-time price discovery in global stock, bond and foreign exchange markets. *Journal of International Economics*, 73:251–277.

Andrade, P., Crump, R. K., Eusepi, S., & Moench, E. (2013). Noisy information and fundamental disagreement. Staff Reports 655, Federal Reserve Bank of New York.

Andrei, D., Carlin, B., & Hasler, M. (2014). Model Disagreement and Economic Outlook. NBER Working Papers 20190, National Bureau of Economic Research, Inc.

Bacchetta, P. & van Wincoop, E. (2013). On the unstable relationship between exchange rates and macroeconomic fundamentals. *Journal of International Economics*, 91(1):18 – 26.

Beber, A. & Brandt, M. W. (2010). When it cannot get better or worse: The asymmetric impact of good and bad news onbondreturns in expansions and recessions. *Review of Finance*.

Bekaert, G., Hoerova, M., & Duca, M. L. (2010). Risk, Uncertainty and Monetary Policy. NBER Working Papers 16397, National Bureau of Economic Research, Inc.

Bloom, N. (2009). The impact of uncertainty shocks. *Econometrica*, 77(3):623–685.

Bomberger, W. A. (1996). Disagreement as a measure of uncertainty. *Journal of Money, Credit and Banking*, 28(3):381–92.

Buraschi, A. & Whelan, P. (2012). Term structure models with differences in beliefs.

Carlin, B. I., Longstaff, F. A., & Matoba, K. (2014). Disagreement and asset prices. *Journal of Financial Economics*, 114(2):226 – 238.

Ehrmann, M. & Fratzscher, M. (2005). Exchange rates and fundamentals: new evidence from real-time data. *Journal of International Money and Finance*, 24(2):317–341.

Faust, J., Rogers, J. H., Wang, S.-Y. B., & Wright, J. H. (2007). The high-frequency response of exchange rates and interest rates to macroeconomic announcements. *Journal of Monetary Economics*, 54(4):1051–1068.

Fleming, M. J. & Remolona, E. M. (1999). The term structure of announcement effects. BIS Working Papers 71, Bank for International Settlements.

Giannone, D., Reichlin, L., & Small, D. (2008). Nowcasting: The real-time informational content of macroeconomic data. *Journal of Monetary Economics*, 55(4):665–676.

Gilbert, T., Scotti, C., Strasser, G., & Vega, C. (2010). Why do certain macroeconomic news announcements have a big impact on asset prices. *Applied Econometrics and Forecasting in Macroeconomics and Finance Workshop, Federal Reserve Bank of St. Louis.*

Giordani, P. & Söderlind, P. (2003). Inflation forecast uncertainty. *European Economic Review*, 47(6):1037–1059.

Goldberg, L. S. & Grisse, C. (2013). Time variation in asset price responses to macro announcements. Staff Reports 626, Federal Reserve Bank of New York.

Greenspan, A. (2005). Testimony of Chairman Alan Greenspan. Technical report, Federal Reserve Board's semiannual Monetary Policy Report to the Congress Before the Committee on Banking, Housing, and Urban Affairs, U.S. Senate.

Gürkaynak, R. & Wolfers, J. (2005). Macroeconomic Derivatives: An Initial Analysis of Market-Based Macro Forecasts, Uncertainty, and Risk. In: *NBER International Seminar on Macroeconomics 2005*, NBER Chapters, National Bureau of Economic Research, Inc, pages 11–50. National Bureau of Economic Research, Inc.

Gürkaynak, R. S., Sack, B., & Swanson, E. (2005). The sensitivity of long-term interest rates to economic news: Evidence and implications for macroeconomic models. *American Economic Review*, 95(1):425–436.

Gürkaynak, R. S. & Wright, J. H. (2013). Identification and Inference Using Event Studies. *Manchester School*, 81:48–65.

Hamilton, J. (2014). Bond market conundrum redux. Technical report, Econbrowser.

Hautsch, N. & Hess, D. (2007). Bayesian learning in financial markets: Testing for the relevance of information precision in price discovery. *Journal of Financial and Quantitative Analysis*, 42(01):189–208.

James, J. & Kasikov, K. (2008). Impact of economic data surprises on exchange rates in the inter-dealer market. *Quantitative Finance*, 8.

Jurado, K., Ludvigson, S. C., & Ng, S. (2013). Measuring Uncertainty. NBER Working Papers 19456, National Bureau of Economic Research, Inc.

Kim, O. & Verrecchia, R. E. (1991). Trading volume and price reactions to public announcements. *Journal of Accounting Research*, 29(2):pp. 302–321.

Lahiri, K. & Sheng, X. (2010). Measuring forecast uncertainty by disagreement: The missing link. *Journal of Applied Econometrics*, 25(4):514–538.

Neely, C. J. & Dey, S. R. (2010). A survey of announcement effects on foreign exchange returns. *Review of Federal Reserve St. Louis*, (Sep):417–464.

Ottaviani, M. & Sørensen, P. N. (2015). Price Reaction to Information with Heterogeneous Beliefs and Wealth Effects: Underreaction, Momentum, and Reversal. *American Economic Review*, 105(1):1–34.

Pasquariello, P. & Vega, C. (2007). Informed and Strategic Order Flow in the Bond Markets. *Review of Financial Studies*, 20(6):1975–2019.

Patton, A. J. & Timmermann, A. (2010). Why do forecasters disagree? Lessons from the term structure of cross-sectional dispersion. *Journal of Monetary Economics*, 57(7):803–820.

Swanson, E. T. & Williams, J. C. (2014a). Measuring the effect of the zero lower bound on medium- and longer-term interest rates. *American Economic Review*, forthcoming.

Swanson, E. T. & Williams, J. C. (2014b). Measuring the effect of the zero lower bound on yields and exchange rates in the U.K. and Germany. *Journal of International Economics*, 92, Supplement 1(0):S2 – S21. 36th Annual {NBER} International Seminar on Macroeconomics.

Veronesi, P. (1999). Stock market overreaction to bad news in good times: A rational expectations equilibrium model. *Review of Financial Studies*, 12(5):975–1007.

Xiong, W. & Yan, H. (2010). Heterogeneous Expectations and Bond Markets. *Review of Financial Studies*, 23(4):1433–1466.

Zarnowitz, V. & Lambros, L. A. (1987). Consensus and uncertainty in economic prediction. *Journal of Political Economy*, 95(3):591–621.

Appendix: Tables

Table A.1 – 10-year US Treasury with time variation

	Baseline		2 disagreement regimes				2 disagreement regimes			
			low disagr.		high disagr.		low disagr.		high disagr.	
Non Farm Payrolls	4.29	(7.19)	6.62	(7.10)	2.29	(2.95)	6.63	(6.80)	2.28	(2.92)
ISM Manufacturing	2.62	(4.83)	3.79	(4.95)	1.42	(1.73)	3.78	(4.80)	1.41	(1.67)
Retail sales ex.autos	2.60	(4.42)	3.01	(3.80)	2.32	(2.76)	3.01	(3.76)	2.30	(2.53)
GDP (advance)	1.87	(1.99)	3.12	(2.34)	0.65	(0.45)	3.07	(2.32)	0.58	(0.40)
Core durable goods orders	1.54	(2.78)	1.01	(1.25)	2.34	(2.87)	1.01	(1.25)	2.58	(2.89)
Consumer confidence	1.45	(2.83)	1.86	(2.65)	1.00	(1.21)	1.85	(2.63)	0.91	(1.05)
Initial claims	-1.45	(-5.40)	-1.43	(-3.51)	-1.65	(-4.70)	-1.40	(-3.42)	-1.68	(-4.19)
Capacity utilization	0.68	(1.35)	1.60	(2.14)	-0.23	(-0.28)	1.54	(2.07)	-0.14	(-0.16)
Trade balance	0.92	(1.86)	0.97	(1.32)	1.15	(1.48)	0.95	(1.32)	1.47	(1.65)
CB leading indicator	0.77	(1.58)	-0.02	(-0.02)	1.15	(1.83)	-0.04	(-0.04)	1.21	(1.74)
CPI Core	0.49	(0.94)	0.61	(0.92)	0.68	(0.75)	0.55	(0.82)	0.73	(0.78)
Unemployment rate	-0.26	(-0.55)	-0.29	(-0.42)	-0.49	(-0.62)	-0.25	(-0.37)	-0.48	(-0.54)
ϕ^{crisis}	0.94	(5.49)	0.91		(5.96)		0.88	(4.81)	0.98	(4.93)
ϕ^{QE}	1.14	(5.09)	0.93		(5.33)		0.92	(3.54)	0.65	(1.91)
R^2	0.099				0.114		0.114			
H_0: $\delta_j = \gamma_j, \forall j$					0.027					
H_0: $\delta_j = \gamma_j, \forall j$ (tranquil)							0.003			
H_0: $\phi^{H,crisis}\delta_j = \phi^{L,crisis}\gamma_j, \forall j$							0.133			
H_0: $\phi^{H,QE}\delta_j = \phi^{L,QE}\gamma_j \forall j$							0.004			
# observations	1,626				1,626		1,626			

Columns 1-2 of the Table present estimates of:

$\Delta y_t = \alpha + \sum_\tau D_\tau \phi^\tau \sum_j \beta_j S_{t,j} + u_t$

columns 3-6 present estimates of

$\Delta y_t = \alpha + \sum_\tau D_\tau \phi^\tau \left[\sum_j \gamma_j S_{t,j} I_{t,j} + \sum_j \delta_j S_{t,j}(1 - I_{t,j})\right] + u_t$

columns 7-10 report estimates of

$\Delta y_t = \alpha + \sum_\tau D_\tau \phi^{L,\tau} \sum_j \gamma_j \cdot S_{t,j} \cdot I^L_{t,j} + \sum_\tau D_\tau \phi^{H,\tau} \sum_j \delta_j \cdot S_{t,j} \cdot (1 - I^L_{t,j}) + u_t$

The sensitivity coefficient is normalized to 1 in the tranquil period for both the low-disagreement and the high-disagreement regime in order to identify the coefficients. The tranquil period runs from January 1999 to July 2007, the crisis period from August 2007 to March 2009, and the QE period – e.g. the implementation and the unravelling of the unconventional US monetary policy period – from April 2009 to June 2014. The t-statictics (in parenthesis) are computed with a White-consistent estimator. The values ϕ^{crisis} and ϕ^{QE} in the Table correspond to coefficients of Table 5 in the text. H_0: $\phi^{H,\tau}\delta_j = \phi^{L,\tau}\gamma_j, \forall j$: is the p-value of the joint test for an asymmetric response to high and low disagreement regimes in a given period τ (tranquil, crisis, QE).

Table A.2 – 10-year Bund with time variation

	Baseline		2 disagreement regimes				2 disagreement regimes			
			low disagr.		high disagr.		low disagr.		high disagr.	
Non Farm Payrolls	2.09	(4.75)	3.62	(5.15)	0.81	(1.36)	3.49	(4.83)	0.93	(1.67)
ISM Manufacturing	1.13	(2.87)	1.47	(2.61)	0.97	(1.55)	1.41	(2.54)	0.96	(1.60)
Retail sales ex.autos	1.39	(3.21)	1.73	(2.89)	1.16	(1.74)	1.68	(2.84)	0.70	(1.29)
GDP (advance)	0.85	(1.22)	1.21	(1.21)	0.71	(0.63)	1.21	(1.22)	0.34	(0.36)
Core durable goods orders	0.09	(0.21)	0.10	(0.17)	0.11	(0.17)	0.11	(0.18)	0.48	(0.80)
Consumer confidence	0.61	(1.72)	0.83	(1.62)	0.33	(0.54)	0.86	(1.72)	0.48	(0.87)
Initial claims	-0.88	(-4.22)	-1.13	(-3.56)	-0.87	(-3.21)	-1.11	(-3.49)	-0.41	(-1.65)
Capacity utilization	0.51	(1.44)	1.09	(2.00)	-0.01	(-0.02)	0.90	(1.69)	0.01	(0.02)
Trade balance	-0.18	(-0.52)	-0.02	(-0.04)	-0.51	(-0.86)	0.01	(0.01)	1.24	(1.82)
CB leading indicator	0.14	(0.42)	0.26	(0.37)	0.09	(0.20)	0.24	(0.35)	-0.10	(-0.26)
CPI Core	0.69	(1.83)	0.59	(1.17)	1.26	(1.83)	0.50	(1.02)	1.94	(2.58)
Unemployment rate	-0.10	(-0.31)	0.17	(0.33)	-0.66	(-1.14)	0.16	(0.32)	0.30	(0.58)
ϕ^{crisis}	1.00	(3.76)	0.82		(3.82)		0.88	(3.30)	1.24	(2.31)
ϕ^{QE}	1.36	(3.74)	1.13		(4.17)		1.17	(3.78)	-1.06	(-1.68)
R^2	0.038		0.053				0.054			
H_0: $\delta_j = \gamma_j, \forall j$					0.340					
H_0: $\delta_j = \gamma_j, \forall j$ (tranquil)							0.110			
H_0: $\phi^{H,crisis}\delta_j = \phi^{L,crisis}\gamma_j, \forall j$							0.491			
H_0: $\phi^{H,QE}\delta_j = \phi^{L,QE}\gamma_j, \forall j$							0.020			
# observations	1,626		1,626				1,626			

See note of Table A.1

Table A.3 – USD/EUR with time variation										
	Baseline		2 disagreement regimes				2 disagreement regimes			
			low disagr.		high disagr.		low disagr.		high disagr.	
Non Farm Payrolls	-0.36	(-5.59)	-0.54	(7.10)	-0.19	(-2.53)	-0.56	(-5.36)	-0.16	(-0.16)
ISM Manufacturing	-0.24	(-3.74)	-0.18	(4.95)	-0.23	(-2.73)	-0.20	(-2.34)	-0.20	(-0.20)
Retail sales ex.autos	-0.08	(-1.16)	-0.15	(3.80)	0.01	(0.16)	-0.16	(-1.87)	0.00	(0.00)
GDP (advance)	-0.21	(-1.82)	-0.04	(2.34)	-0.42	(-2.85)	-0.04	(-0.26)	-0.40	(-0.40)
Core durable goods orders	-0.06	(-0.83)	-0.03	(1.25)	-0.04	(-0.42)	-0.03	(-0.36)	-0.02	(-0.02)
Consumer confidence	-0.08	(-1.22)	-0.14	(2.65)	0.08	(0.85)	-0.15	(-1.79)	0.07	(0.07)
Initial claims	0.00	(0.15)	0.01	(-3.51)	-0.02	(-0.45)	0.01	(0.28)	-0.02	(-0.02)
Capacity utilization	-0.02	(-0.28)	-0.05	(2.14)	-0.01	(-0.07)	-0.05	(-0.51)	0.00	(0.00)
Trade balance	-0.18	(-2.43)	-0.22	(1.32)	-0.05	(-0.53)	-0.23	(-2.61)	-0.04	(-0.04)
CB leading indicator	-0.01	(-0.07)	-0.21	(-0.02)	0.12	(1.63)	-0.21	(-1.79)	0.11	(0.11)
CPI Core	0.02	(0.37)	-0.03	(0.92)	0.12	(1.27)	-0.03	(-0.42)	0.09	(0.09)
Unemployment rate	0.16	(2.23)	0.18	(-0.42)	0.06	(0.64)	0.18	(2.24)	0.05	(0.05)
ϕ^{crisis}	0.48	(2.49)	0.87		(4.25)		0.77	(3.42)	1.19	(2.53)
ϕ^{QE}	0.25	(1.14)	0.20		(1.01)		0.16	(0.84)	0.36	(0.58)
R^2	0.038		0.053				0.054			
H_0: $\delta_j = \gamma_j, \forall j$					0.021					
H_0: $\delta_j = \gamma_j, \forall j$ (tranquil)							0.067			
H_0: $\phi^{H,crisis}\delta_j = \phi^{L,crisis}\gamma_j, \forall j$							0.189			
H_0: $\phi^{H,QE}\delta_j = \phi^{L,QE}\gamma_j, \forall j$							0.901			
# observations	1,626		1,626				1,626			

See note of Table A.1

Appendix: Ottaviani and Sørensen model

In the baseline OS model agents take positions on wheter or not a binary event, E,R, is realized (say the economy will be in an expansion period). They trade 2 Arrow-Debreu securities whose payoff is linked to the two possible outcomes. Each asset pays off one unit of cash if the corresponding event occurs (E, R), zero otherwise.

There is a continuum of risk-neutral agents, indexed by i, endowed with a constant (across states) amount of assets in period , w_{i0}. Agents exchange their assets in a competitive market, but cannot hold a negative amounts of any asset.[19]

Agents' prior beliefs on the realization of the event E are heterogeneous, $q_i = Prob_i(E)$, and described by a distribution of initial assets over individuals, $G(q)$ over the interval $[0, 1]$.[20]

Before trading takes place, agents observe a public signal with likelihood ratio L for event E, and update their beliefs according to Bayes' rule. Agent's i posterior odds are revised in a concordant

[19]The aggregate endowment of each asset is 1.

[20]$G(q)$ describes the share of wealth in the hands of agents with prior beliefs $q_i < q$. $G(q)$ is assumed continuous and strictly increasing.

manner using the same L as:

$$\frac{\pi_i}{(1-\pi_i)} \quad = \quad L \times \frac{q_i}{(1-q_i)} \tag{7}$$

When the public information is more favorable to event E, $L > 1$, the distribution of beliefs will also shift to the right, $G(\pi) = G(q|L)$. Since agents are risk neutral they will choose to demand only one kind of asset depending on their beliefs. Those more optimistic, with posterior beliefs π_i greater than the price of asset E, will sell of their initial holding of the asset paying in state R to purchase only assets paying in state E; conversely pessimistic agents, with posterior beliefs π_i lower than the price of asset E, will sell of their initial holding of the asset paying in state E to purchase only assets paying in state R. Demand for the E asset is given the cumulated wealth of agents with posterior beliefs above the price p (who will only hold asset E), divided by the price of the asset. Markets clear when this demand is equal to the aggregate unit endowment of this asset.

The competitive equilibrium price, p, is the unique solution to the equation

$$\begin{aligned} \int_p^1 w_{i,0} dG_i \frac{1}{p} &= \int_0^1 w_{i,0} di = 1 \\ p &= 1 - G\left(\frac{p}{L(1-p)+p}\right) \end{aligned}$$

The price can be associated to the posterior beliefs of a particular agent, the marginal trader. By inverting the Bayesian updating formula, the ex-ante beliefs of this marginal trader can be shown to be $p/[(1-p)L+p]$. This particular agent is the marginal trader, whose beliefs are also an average of the heterogenous subjective prior beliefs of the population. The effect of information is to move the marginal trader prior in a direction opposite to the price movement. The greater the heterogeneity in prior beliefs the more pessimist the marginal trader will have to be ex-ante.

The competitive equilibrium price depends on the public signal, as well as on the shape of the distribution of wealth (and hence beliefs) across individuals. OS show that the price underreacts to information (Proposition 2). Underreaction is measured relatively to the one associated to an hypothetical outside observer, whose prior beliefs are those of the average priors across agents', and who updates his belief in a Bayesian fashion. In the model the equilibrium price does not correspond to the posterior of this "outside observer". Changes in the wealth distribution across agents (with different prior beliefs) affect the equilibrium and the extent of the underreaction. In particular, more underreaction results when the distribution is less concentrated (Proposition 3), that is when agents have greater heterogeneity in beliefs.

To illustrate the model predictions we consider a particular *symmetric* prior distribution $G(q)$, also described in OS: $G(q) = \frac{q^\gamma}{q^\gamma + (1-q)^\gamma}$ where $\gamma > 0$ is a parameter that measures the concentration of beliefs around the average belief $q = 1/2$. The greater γ is, the more concentrated the distribution is. For $\gamma = 1$ beliefs are uniformly distributed, while as $\gamma \to 0$ beliefs become maximally dispersed

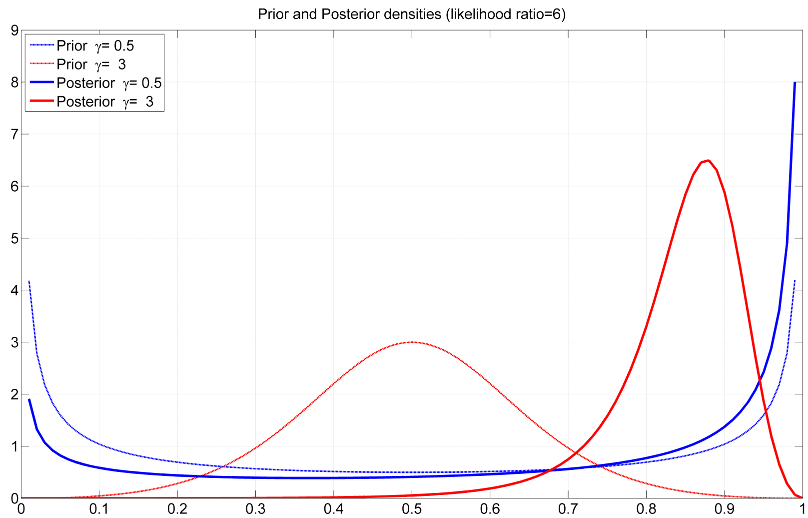


Figure 8: Prior and posterior densities: differing γ, and $L = 6$

on the extremes of $[0, 1]$ (see Figure 8).

To gain intuition on the underlying mechanism, suppose the public signal is in favor of event E, i.e. $L > 1$. Agents will update their beliefs according to Bayes rule, with their posterior odds ratio in favor of E rising one for one with L. Those that were ex-ante optimist, and willing to invest all their wealth in asset E (agents are risk neutral and subject to a trading bound) will continue to do so. However at the higher price they will no longer be able to buy the same amount of assets E. Similarly those ex-ante pessimist will be able to buy more assets R, as its price has fallen. This would lead to an insufficient demand for assets E and excess demand for asset R. To balance the market some ex-ante pessimists have to "switch sides", becoming optimist, and seek to sell all their assets R in return for assets E.

The price elasticity to "information" is described by the parameter γ, and Table 1 in Section 2.1 provides some comparative statics: for a given likelihood ratio ($L = 6$), varying γ leads to different price responses. Greater heterogeneity in beliefs (higher γ), is associated with a smaller price response. This occurs because the marginal trader beliefs are further away from the prior mean and in a direction opposite to the information (more pessimistic). In contrast, when beliefs are very concentrated the marginal trader ex-ante beliefs are closer to the prior mean and accordingly also the price reaction almost parallels the shift in "average" beliefs.

Appendix: data

The US dollar–euro daily exchange rate (USD/EUR) is taken from Thomson Reuters. Long-term yields are the 10-year zero-coupon interest rates computed by the Federal Reserve Board for the US market and by the Bundesbank for the German market. We use zero-coupon yields instead of 10-year benchmark bond yields as the former strip out differences in coupon rates and fluctuations in effective bond duration that arise from changes in the level of interest rates. As a result, zero-coupon yields provide the cleanest measure of the interest rate at any given maturity across countries and over time. This is important in our analysis, because the level of interest rates varies substantially from the early 2000s to the middle of 2014, which would cause the yield-to-maturity on coupon-bearing bonds to fluctuate over time. Note the German interest rate is collected at 13:00 GMT+1 in Frankfurt and then it is not impacted by US data release; so in the empirical analysis we shift forward by one day the German interest rate. Conversely, the USD/EUR is collected at 16:00 GMT and so its level can be impacted by the release of US data. Intraday US Treasury prices are kindly taken from the New York Federal Reserve, the intraday USD/EUR exchange rate from the Olsen database.[21]

Figure 9: Price levels

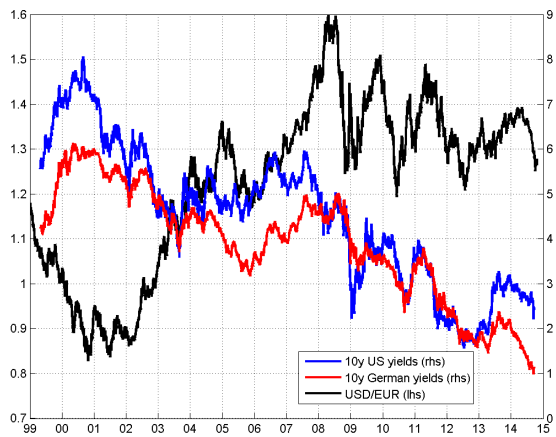
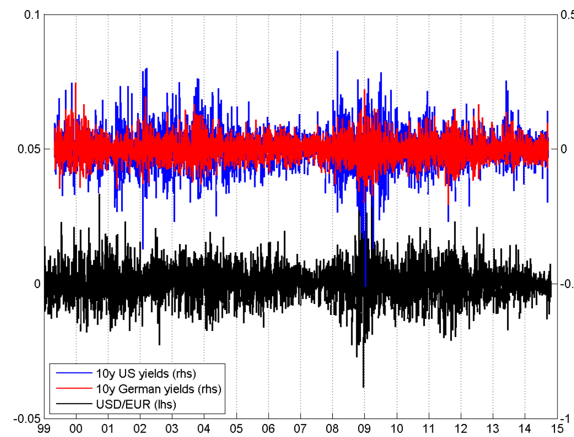


Figure 10: Price changes



The left panel of the Figure shows the daily level of the German 10-year zero-coupon rate, the US 10-year zero-coupon rate and the USD/EUR exchange rate. The right panel shows the first differences of the corresponding interest rates and the logarithmic difference of the USD/EUR exchange rate.

Long-term interest rates in the two countries show a decreasing trend since 1999 (Figures 9). The USD/EUR exchange rate, conversely appreciates from 1999 until 2002 and weakens thereafter, ranging between 1.20 and 1.40. Changes in interest rates are very similar and show some heteroskedasticity and persistence during the financial crisis (Table 2). Statistics show a similar pattern; German and US interest rates have very similar standard deviations, skewness and kurtosis, both in levels and in changes. Moreover, the time series show a very strong short-term persistence in levels – close to 1 for the 1-period lag for interest rates and exchange rates.

[21]Intraday for the German interest rates have are not available yet.

Table 2 – 10-year yields and USD/EUR

	level			change		
	10y Ger. rate	10y US rate	USD/EUR	10y Ger. rate	10y US rate	USD/EUR
mean	3.75	4.22	1.22	0.00	0.00	0.00
std. dev.	1.16	1.24	0.18	0.05	0.06	0.62
skewness	-0.56	-0.29	-0.52	0.16	0.05	0.18
kurtosis	2.38	2.52	2.33	4.57	5.34	5.62
ρ_1	0.87	1.00	1.00	-0.07	0.01	0.01
ρ_5	0.83	0.99	1.00	-0.01	-0.01	0.00
ρ_{20}	0.77	0.98	0.98	-0.02	-0.02	0.00
ρ_{60}	0.75	0.93	0.95	0.01	0.00	-0.02

The Table reports the statistics for the level and the change in daily German 10-year zero-coupon rate, US 10-year zero-coupon rate and USD/EUR exchange rate. The statistics ρ_i show the autocorrelation coefficient with i lagged days.

RECENTLY PUBLISHED "TEMI" (*)

N. 994 – *Trade liberalizations and domestic suppliers: evidence from Chile*, by Andrea Linarello (November 2014).

N. 995 – *Dynasties in professions: the role of rents*, by Sauro Mocetti (November 2014).

N. 996 – *Current account "core-periphery dualism" in the EMU*, by Tatiana Cesaroni and Roberta De Santis (November 2014).

N. 997 – *Macroeconomic effects of simultaneous implementation of reforms after the crisis*, by Andrea Gerali, Alessandro Notarpietro and Massimiliano Pisani (November 2014).

N. 998 – *Changing labour market opportunities for young people in Italy and the role of the family of origin*, by Gabriella Berloffa, Francesca Modena and Paola Villa (January 2015).

N. 999 – *Looking behind mortgage delinquencies*, by Sauro Mocetti and Eliana Viviano (January 2015).

N. 1000 – *Sectoral differences in managers' compensation: insights from a matching model*, by Emanuela Ciapanna, Marco Taboga and Eliana Viviano (January 2015).

N. 1001 – *How does foreign demand activate domestic value added? A comparison among the largest euro-area economies*, by Rita Cappariello and Alberto Felettigh (January 2015).

N. 1002 – *Structural reforms and zero lower bound in a monetary union*, by Andrea Gerali, Alessandro Notarpietro and Massimiliano Pisani (January 2015).

N. 1003 – *You've come a long way, baby. Effects of commuting times on couples' labour supply*, by Francesca Carta and Marta De Philippis (March 2015).

N. 1004 – *Ownership networks and aggregate volatility*, by Lorenzo Burlon (March 2015).

N. 1005 – *Strategy and tactics in public debt manamgement*, by Davide Dottori and Michele Manna (March 2015).

N. 1006 – *Inward foreign direct investment and innovation: evidence from Italian provinces*, by Roberto Antonietti, Raffaello Bronzini and Giulio Cainelli (March 2015).

N. 1007 – *The macroeconomic effects of the sovereign debt crisis in the euro area*, by Stefano Neri and Tiziano Ropele (March 2015).

N. 1008 – *Rethinking the crime reducing effect of education? Mechanisms and evidence from regional divides*, by Ylenia Brilli and Marco Tonello (April 2015).

N. 1009 – *Social capital and the cost of credit: evidence from a crisis*, by Paolo Emilio Mistrulli and Valerio Vacca (April 2015).

N. 1010 – *Every cloud has a silver lining. The sovereign crisis and Italian potential output*, by Andrea Gerali, Alberto Locarno, Alessandro Notarpietro and Massimiliano Pisani (June 2015).

N. 1011 – *Foreign direct investment and firm performance: an empirical analysis of Italian firms*, by Alessandro Borin and Michele Mancini (June 2015).

N. 1012 – *Sovereign debt and reserves with liquidity and productivity crises*, by Flavia Corneli and Emanuele Tarantino (June 2015).

N. 1013 – *Bankruptcy law and bank financing*, by Giacomo Rodano, Nicolas Serrano-Velarde and Emanuele Tarantino (June 2015).

N. 1014 – *Women as 'gold dust': gender diversity in top boards and the performance of Italian banks*, by Silvia Del Prete and Maria Lucia Stefani (June 2015).

N. 1015 – *Inflation, financial conditions and non-standard monetary policy in a monetary union. A model-based evaluation*, by Lorenzo Burlon, Andrea Gerali, Alessandro Notarpietro and Massimiliano Pisani (June 2015).

N. 1016 – *Short term inflation forecasting: the M.E.T.A. approach*, by Giacomo Sbrana, Andrea Silvestrini and Fabrizio Venditti (June 2015).

(*) Requests for copies should be sent to:
Banca d'Italia – Servizio Struttura economica e finanziaria – Divisione Biblioteca e Archivio storico –
Via Nazionale, 91 – 00184 Rome – (fax 0039 06 47922059). They are available on the Internet www.bancaditalia.it.

2012

F. CINGANO and A. ROSOLIA, *People I know: job search and social networks*, Journal of Labor Economics, v. 30, 2, pp. 291-332, **TD No. 600 (September 2006).**

G. GOBBI and R. ZIZZA, *Does the underground economy hold back financial deepening? Evidence from the italian credit market*, Economia Marche, Review of Regional Studies, v. 31, 1, pp. 1-29, **TD No. 646 (November 2006).**

S. MOCETTI, *Educational choices and the selection process before and after compulsory school*, Education Economics, v. 20, 2, pp. 189-209, **TD No. 691 (September 2008).**

P. PINOTTI, M. BIANCHI and P. BUONANNO, *Do immigrants cause crime?*, Journal of the European Economic Association , v. 10, 6, pp. 1318–1347, **TD No. 698 (December 2008).**

M. PERICOLI and M. TABOGA, *Bond risk premia, macroeconomic fundamentals and the exchange rate*, International Review of Economics and Finance, v. 22, 1, pp. 42-65, **TD No. 699 (January 2009).**

F. LIPPI and A. NOBILI, *Oil and the macroeconomy: a quantitative structural analysis,* Journal of European Economic Association, v. 10, 5, pp. 1059-1083, **TD No. 704 (March 2009).**

G. ASCARI and T. ROPELE, *Disinflation in a DSGE perspective: sacrifice ratio or welfare gain ratio?,* Journal of Economic Dynamics and Control, v. 36, 2, pp. 169-182, **TD No. 736 (January 2010).**

S. FEDERICO, *Headquarter intensity and the choice between outsourcing versus integration at home or abroad*, Industrial and Corporate Chang, v. 21, 6, pp. 1337-1358, **TD No. 742 (February 2010).**

I. BUONO and G. LALANNE, *The effect of the Uruguay Round on the intensive and extensive margins of trade,* Journal of International Economics, v. 86, 2, pp. 269-283, **TD No. 743 (February 2010).**

A. BRANDOLINI, S. MAGRI and T. M SMEEDING, *Asset-based measurement of poverty*, In D. J. Besharov and K. A. Couch (eds), Counting the Poor: New Thinking About European Poverty Measures and Lessons for the United States, Oxford and New York: Oxford University Press, **TD No. 755 (March 2010).**

S. GOMES, P. JACQUINOT and M. PISANI, *The EAGLE. A model for policy analysis of macroeconomic interdependence in the euro area,* Economic Modelling, v. 29, 5, pp. 1686-1714, **TD No. 770 (July 2010).**

A. ACCETTURO and G. DE BLASIO, *Policies for local development: an evaluation of Italy's "Patti Territoriali",* Regional Science and Urban Economics, v. 42, 1-2, pp. 15-26, **TD No. 789 (January 2006).**

E. COCOZZA and P. PISELLI, *Testing for east-west contagion in the European banking sector during the financial crisis,* in R. Matoušek; D. Stavárek (eds.), Financial Integration in the European Union, Taylor & Francis, **TD No. 790 (February 2011).**

F. BUSETTI and S. DI SANZO, *Bootstrap LR tests of stationarity, common trends and cointegration,* Journal of Statistical Computation and Simulation, v. 82, 9, pp. 1343-1355, **TD No. 799 (March 2006).**

S. NERI and T. ROPELE, *Imperfect information, real-time data and monetary policy in the Euro area,* The Economic Journal, v. 122, 561, pp. 651-674, **TD No. 802 (March 2011).**

A. ANZUINI and F. FORNARI, *Macroeconomic determinants of carry trade activity,* Review of International Economics, v. 20, 3, pp. 468-488, **TD No. 817 (September 2011).**

M. AFFINITO, *Do interbank customer relationships exist? And how did they function in the crisis? Learning from Italy*, Journal of Banking and Finance, v. 36, 12, pp. 3163-3184, **TD No. 826 (October 2011).**

P. GUERRIERI and F. VERGARA CAFFARELLI, *Trade Openness and International Fragmentation of Production in the European Union: The New Divide?,* Review of International Economics, v. 20, 3, pp. 535-551, **TD No. 855 (February 2012).**

V. DI GIACINTO, G. MICUCCI and P. MONTANARO, *Network effects of public transposrt infrastructure: evidence on Italian regions,* Papers in Regional Science, v. 91, 3, pp. 515-541, **TD No. 869 (July 2012).**

A. FILIPPIN and M. PACCAGNELLA, *Family background, self-confidence and economic outcomes,* Economics of Education Review, v. 31, 5, pp. 824-834, **TD No. 875 (July 2012).**

2013

A. MERCATANTI, *A likelihood-based analysis for relaxing the exclusion restriction in randomized experiments with imperfect compliance*, Australian and New Zealand Journal of Statistics, v. 55, 2, pp. 129-153, **TD No. 683 (August 2008).**

F. CINGANO and P. PINOTTI, *Politicians at work. The private returns and social costs of political connections*, Journal of the European Economic Association, v. 11, 2, pp. 433-465, **TD No. 709 (May 2009).**

F. BUSETTI and J. MARCUCCI, *Comparing forecast accuracy: a Monte Carlo investigation*, International Journal of Forecasting, v. 29, 1, pp. 13-27, **TD No. 723 (September 2009).**

D. DOTTORI, S. I-LING and F. ESTEVAN, *Reshaping the schooling system: The role of immigration*, Journal of Economic Theory, v. 148, 5, pp. 2124-2149, **TD No. 726 (October 2009).**

A. FINICELLI, P. PAGANO and M. SBRACIA, *Ricardian Selection*, Journal of International Economics, v. 89, 1, pp. 96-109, **TD No. 728 (October 2009).**

L. MONTEFORTE and G. MORETTI, *Real-time forecasts of inflation: the role of financial variables*, Journal of Forecasting, v. 32, 1, pp. 51-61, **TD No. 767 (July 2010).**

R. GIORDANO and P. TOMMASINO, *Public-sector efficiency and political culture*, FinanzArchiv, v. 69, 3, pp. 289-316, **TD No. 786 (January 2011).**

E. GAIOTTI, *Credit availablility and investment: lessons from the "Great Recession"*, European Economic Review, v. 59, pp. 212-227, **TD No. 793 (February 2011).**

F. NUCCI and M. RIGGI, *Performance pay and changes in U.S. labor market dynamics*, Journal of Economic Dynamics and Control, v. 37, 12, pp. 2796-2813, **TD No. 800 (March 2011).**

G. CAPPELLETTI, G. GUAZZAROTTI and P. TOMMASINO, *What determines annuity demand at retirement?,* The Geneva Papers on Risk and Insurance – Issues and Practice, pp. 1-26, **TD No. 805 (April 2011).**

A. ACCETTURO e L. INFANTE, *Skills or Culture? An analysis of the decision to work by immigrant women in Italy,* IZA Journal of Migration, v. 2, 2, pp. 1-21, **TD No. 815 (July 2011).**

A. DE SOCIO, *Squeezing liquidity in a "lemons market" or asking liquidity "on tap"*, Journal of Banking and Finance, v. 27, 5, pp. 1340-1358, **TD No. 819 (September 2011).**

S. GOMES, P. JACQUINOT, M. MOHR and M. PISANI, *Structural reforms and macroeconomic performance in the euro area countries: a model-based assessment,* International Finance, v. 16, 1, pp. 23-44, **TD No. 830 (October 2011).**

G. BARONE and G. DE BLASIO, *Electoral rules and voter turnout,* International Review of Law and Economics, v. 36, 1, pp. 25-35, **TD No. 833 (November 2011).**

O. BLANCHARD and M. RIGGI, *Why are the 2000s so different from the 1970s? A structural interpretation of changes in the macroeconomic effects of oil prices*, Journal of the European Economic Association, v. 11, 5, pp. 1032-1052, **TD No. 835 (November 2011).**

R. CRISTADORO and D. MARCONI, *Household savings in China,* in G. Gomel, D. Marconi, I. Musu, B. Quintieri (eds), The Chinese Economy: Recent Trends and Policy Issues, Springer-Verlag, Berlin, **TD No. 838 (November 2011).**

A. ANZUINI, M. J. LOMBARDI and P. PAGANO, *The impact of monetary policy shocks on commodity prices*, International Journal of Central Banking, v. 9, 3, pp. 119-144, **TD No. 851 (February 2012).**

R. GAMBACORTA and M. IANNARIO, *Measuring job satisfaction with CUB models,* Labour, v. 27, 2, pp. 198-224, **TD No. 852 (February 2012).**

G. ASCARI and T. ROPELE, *Disinflation effects in a medium-scale new keynesian model: money supply rule versus interest rate rule,* European Economic Review, v. 61, pp. 77-100, **TD No. 867 (April 2012).**

E. BERETTA and S. DEL PRETE, *Banking consolidation and bank-firm credit relationships: the role of geographical features and relationship characteristics,* Review of Economics and Institutions, v. 4, 3, pp. 1-46, **TD No. 901 (February 2013).**

M. ANDINI, G. DE BLASIO, G. DURANTON and W. STRANGE, *Marshallian labor market pooling: evidence from Italy,* Regional Science and Urban Economics, v. 43, 6, pp.1008-1022, **TD No. 922 (July 2013).**

G. SBRANA and A. SILVESTRINI, *Forecasting aggregate demand: analytical comparison of top-down and bottom-up approaches in a multivariate exponential smoothing framework,* International Journal of Production Economics, v. 146, 1, pp. 185-98, **TD No. 929 (September 2013).**

A. FILIPPIN, C. V, FIORIO and E. VIVIANO, *The effect of tax enforcement on tax morale,* European Journal of Political Economy, v. 32, pp. 320-331, **TD No. 937 (October 2013).**

2014

G. M. TOMAT, *Revisiting poverty and welfare dominance*, Economia pubblica, v. 44, 2, 125-149, **TD No. 651 (December 2007).**

M. TABOGA, *The riskiness of corporate bonds*, Journal of Money, Credit and Banking, v.46, 4, pp. 693-713, **TD No. 730 (October 2009).**

G. MICUCCI and P. ROSSI, *Il ruolo delle tecnologie di prestito nella ristrutturazione dei debiti delle imprese in crisi*, in A. Zazzaro (a cura di), Le banche e il credito alle imprese durante la crisi, Bologna, Il Mulino, **TD No. 763 (June 2010).**

F. D'AMURI, *Gli effetti della legge 133/2008 sulle assenze per malattia nel settore pubblico,* Rivista di politica economica, v. 105, 1, pp. 301-321, **TD No. 787 (January 2011).**

R. BRONZINI and E. IACHINI, *Are incentives for R&D effective? Evidence from a regression discontinuity approach,* American Economic Journal : Economic Policy, v. 6, 4, pp. 100-134, **TD No. 791 (February 2011).**

P. ANGELINI, S. NERI and F. PANETTA, *The interaction between capital requirements and monetary policy*, Journal of Money, Credit and Banking, v. 46, 6, pp. 1073-1112, **TD No. 801 (March 2011).**

M. BRAGA, M. PACCAGNELLA and M. PELLIZZARI, *Evaluating students' evaluations of professors,* Economics of Education Review, v. 41, pp. 71-88, **TD No. 825 (October 2011).**

M. FRANCESE and R. MARZIA, *Is there Room for containing healthcare costs? An analysis of regional spending differentials in Italy,* The European Journal of Health Economics, v. 15, 2, pp. 117-132, **TD No. 828 (October 2011).**

L. GAMBACORTA and P. E. MISTRULLI, *Bank heterogeneity and interest rate setting: what lessons have we learned since Lehman Brothers?,* Journal of Money, Credit and Banking, v. 46, 4, pp. 753-778, **TD No. 829 (October 2011).**

M. PERICOLI, *Real term structure and inflation compensation in the euro area*, International Journal of Central Banking, v. 10, 1, pp. 1-42, **TD No. 841 (January 2012).**

E. GENNARI and G. MESSINA, *How sticky are local expenditures in Italy? Assessing the relevance of the flypaper effect through municipal data,* International Tax and Public Finance, v. 21, 2, pp. 324-344, **TD No. 844 (January 2012).**

V. DI GACINTO, M. GOMELLINI, G. MICUCCI and M. PAGNINI, *Mapping local productivity advantages in Italy: industrial districts, cities or both?*, Journal of Economic Geography, v. 14, pp. 365–394, **TD No. 850 (January 2012).**

A. ACCETTURO, F. MANARESI, S. MOCETTI and E. OLIVIERI, *Don't Stand so close to me: the urban impact of immigration,* Regional Science and Urban Economics, v. 45, pp. 45-56, **TD No. 866 (April 2012).**

M. PORQUEDDU and F. VENDITTI, *Do food commodity prices have asymmetric effects on euro area inflation,* Studies in Nonlinear Dynamics and Econometrics, v. 18, 4, pp. 419-443, **TD No. 878 (September 2012).**

S. FEDERICO, *Industry dynamics and competition from low-wage countries: evidence on Italy*, Oxford Bulletin of Economics and Statistics, v. 76, 3, pp. 389-410, **TD No. 879 (September 2012).**

F. D'AMURI and G. PERI, *Immigration, jobs and employment protection: evidence from Europe before and during the Great Recession,* Journal of the European Economic Association, v. 12, 2, pp. 432-464, **TD No. 886 (October 2012).**

M. TABOGA, *What is a prime bank? A euribor-OIS spread perspective,* International Finance, v. 17, 1, pp. 51-75, **TD No. 895 (January 2013).**

L. GAMBACORTA and F. M. SIGNORETTI, *Should monetary policy lean against the wind? An analysis based on a DSGE model with banking,* Journal of Economic Dynamics and Control, v. 43, pp. 146-74, **TD No. 921 (July 2013).**

M. BARIGOZZI, CONTI A.M. and M. LUCIANI, *Do euro area countries respond asymmetrically to the common monetary policy?,* Oxford Bulletin of Economics and Statistics, v. 76, 5, pp. 693-714, **TD No. 923 (July 2013).**

U. ALBERTAZZI and M. BOTTERO, *Foreign bank lending: evidence from the global financial crisis,* Journal of International Economics, v. 92, 1, pp. 22-35, **TD No. 926 (July 2013).**

R. DE BONIS and A. SILVESTRINI, *The Italian financial cycle: 1861-2011,* Cliometrica, v.8, 3, pp. 301-334, **TD No. 936 (October 2013).**

D. PIANESELLI and A. ZAGHINI, *The cost of firms' debt financing and the global financial crisis,* Finance Research Letters, v. 11, 2, pp. 74-83, **TD No. 950 (February 2014).**

A. ZAGHINI, *Bank bonds: size, systemic relevance and the sovereign,* International Finance, v. 17, 2, pp. 161-183, **TD No. 966 (July 2014).**

S. MAGRI, *Does issuing equity help R&D activity? Evidence from unlisted Italian high-tech manufacturing firms,* Economics of Innovation and New Technology, v. 23, 8, pp. 825-854, **TD No. 978 (October 2014).**

G. BARONE and S. MOCETTI, *Natural disasters, growth and institutions: a tale of two earthquakes,* Journal of Urban Economics, v. 84, pp. 52-66, **TD No. 949 (January 2014).**

2015

G. BULLIGAN, M. MARCELLINO and F. VENDITTI, *Forecasting economic activity with targeted predictors,* International Journal of Forecasting, v. 31, 1, pp. 188-206, **TD No. 847 (February 2012).**

A. CIARLONE, *House price cycles in emerging economies,* Studies in Economics and Finance, v. 32, 1, **TD No. 863 (May 2012).**

G. BARONE and G. NARCISO, *Organized crime and business subsidies: Where does the money go?,* Journal of Urban Economics, v. 86, pp. 98-110, **TD No. 916 (June 2013).**

P. ALESSANDRI and B. NELSON, *Simple banking: profitability and the yield curve,* Journal of Money, Credit and Banking, v. 47, 1, pp. 143-175, **TD No. 945 (January 2014).**

R. AABERGE and A. BRANDOLINI, *Multidimensional poverty and inequality,* in A. B. Atkinson and F. Bourguignon (eds.), Handbook of Income Distribution, Volume 2A, Amsterdam, Elsevier, **TD No. 976 (October 2014).**

M. FRATZSCHER, D. RIMEC, L. SARNOB and G. ZINNA, *The scapegoat theory of exchange rates: the first tests,* Journal of Monetary Economics, v. 70, 1, pp. 1-21, **TD No. 991 (November 2014).**

FORTHCOMING

M. BUGAMELLI, S. FABIANI and E. SETTE, *The age of the dragon: the effect of imports from China on firm-level prices*, Journal of Money, Credit and Banking, **TD No. 737 (January 2010).**

G. DE BLASIO, D. FANTINO and G. PELLEGRINI, *Evaluating the impact of innovation incentives: evidence from an unexpected shortage of funds,* Industrial and Corporate Change, **TD No. 792 (February 2011).**

A. DI CESARE, A. P. STORK and C. DE VRIES, *Risk measures for autocorrelated hedge fund returns,* Journal of Financial Econometrics, **TD No. 831 (October 2011).**

D. FANTINO, A. MORI and D. SCALISE, *Collaboration between firms and universities in Italy: the role of a firm's proximity to top-rated departments,* Rivista Italiana degli economisti, **TD No. 884 (October 2012).**

M. MARCELLINO, M. PORQUEDDU and F. VENDITTI, *Short-Term GDP Forecasting with a mixed frequency dynamic factor model with stochastic volatility,* Journal of Business & Economic Statistics, **TD No. 896 (January 2013).**

M. ANDINI and G. DE BLASIO, *Local development that money cannot buy: Italy's Contratti di Programma,* Journal of Economic Geography, **TD No. 915 (June 2013).**

J. LI and G. ZINNA, *On bank credit risk: sytemic or bank-specific? Evidence from the US and UK,* Journal of Financial and Quantitative Analysis, **TD No. 951 (February 2015).**