BANK OF ENGLAND

Staff Working Paper No. 626

Unsurprising shocks: information, premia, and the monetary transmission

Silvia Miranda-Agrippino

November 2016

Staff Working Papers describe research in progress by the author(s) and are published to elicit comments and to further debate. Any views expressed are solely those of the author(s) and so cannot be taken to represent those of the Bank of England or to state Bank of England policy. This paper should therefore not be reported as representing the views of the Bank of England or members of the Monetary Policy Committee, Financial Policy Committee or Prudential Regulation Authority Board.

BANK OF ENGLAND

Staff Working Paper No. 626
Unsurprising shocks: information, premia, and the monetary transmission

Silvia Miranda-Agrippino[1]

Abstract

Central banks' decisions are a function of forecasts of macroeconomic fundamentals. Because private sector forecasts are not bound to be equal to central banks' forecasts, what markets label as unexpected may or may not be unanticipated by the central bank. Monetary surprises can thus incorporate anticipatory effects if market participants fail to correctly account for the systematic component of policy when they are surprised by an interest rate decision. Depending on how market participants perceive the policy decision, their economic projections and the associated risk compensations move in opposite directions, and do so at the time of the announcement. Hence, and regardless of the width of the measurement window, monetary surprises capture more than just the monetary policy shock, and their use as external instruments for identification is potentially compromised. Monetary 'surprises' are dependent on, and shown to be predictable by both central banks' forecasts and past information available to market participants prior to the announcement. A New-Keynesian framework sketches the intuition. The resulting distortions in the estimated impulse response functions can be dramatic, both qualitatively and quantitatively. A new set of monetary surprises, free of anticipatory effects and unpredictable by past information, are shown to retrieve transmission coefficients to a monetary policy shock consistent with macroeconomic theory even in informationally deficient VARs.

Key words: Monetary surprises, identification with external instruments, monetary policy, expectations, information asymmetries, event study, proxy SVAR.

JEL classification: E52, E44, G14, C36.

(1) Bank of England. Email: silvia.miranda-agrippino@bankofengland.co.uk. Web: www.silviamirandaagrippino.com

I am indebted to Raffaella Giacomini, Alejandro Justiniano, Ricardo Reis and Giovanni Ricco for many insightful discussions and for taking time to comment on earlier versions of this paper. I am grateful to Refet Gürkaynak for graciously sharing the updated daily US monetary surprises and to Filippo Ferroni, Elena Gerko, Francisco Gomes, Marek Jarociński, Iryna Kaminska, Roland Meeks, Andrew Meldrum, Lubos Pastor, Michele Piffer, Valerie Ramey, Hélène Rey, Matt Roberts-Sklar, Paolo Surico, colleagues at the Macro-Financial Analysis Division of the Bank of England, seminars and conference participants at the Bank of England, London Business School, the University of Bonn, the University of Surrey, the Graduate Institute in Geneva, the 2016 Annual Meetings of the Royal Economic Society, the 2016 Meeting of the International Association of Applied Econometrics, and the 24th CEPR European Summer Symposium in International Macroeconomics (ESSIM) for helpful suggestions. I also thank Angelo De Marco for answering all my Python-related questions. The views expressed in this paper are those of the author and do not necessarily reflect those of the Bank of England, the Monetary Policy Committee, the Financial Policy Committee or the Prudential Regulation Authority. This paper was previously circulated with the title 'Unsurprising shocks: measuring responses to monetary announcements using high-frequency data'.

Online Appendix available at www.bankofengland.co.uk/research/Documents/workingpapers/2016/swp626appendix.pdf.

Information on the Bank's working paper series can be found at
www.bankofengland.co.uk/research/Pages/workingpapers/default.aspx

Publications Team, Bank of England, Threadneedle Street, London, EC2R 8AH
Telephone +44 (0)20 7601 4030 Fax +44 (0)20 7601 3298 email publications@bankofengland.co.uk

© **Bank of England 2016**
ISSN 1749-9135 (on-line)

1 Introduction

Recent advances in empirical monetary economics have proposed to use monetary surprises as external instruments for monetary policy shocks to achieve identification of the contemporaneous transmission coefficients in structural vector autoregressions (VARs). Monetary policy surprises are typically computed as the price updates in futures on interest rates that follow central banks' communication of the interest rate decisions. The argument in favour of their use is that, to the extent that these prices provide sensible measures of expectations about future policy rates, if the surprises are computed within a sufficiently narrow window around the announcement, they can then be regarded as a measure, with error, of the underlying monetary policy shock.

Two crucial assumptions make monetary surprises the ideal candidate for the job: (*i*) markets efficiently incorporate all the relevant information as it becomes available, and it takes longer than the measurement window for the monetary policy shock to modify the risk premium in these contracts, and (*ii*) the set of economic forecasts on the basis of which central banks' decisions are taken, and those of market participants coincide, leading to the equivalence between price updates and monetary policy shocks. These assumptions make it possible to first map all price changes into revisions in market-implied expectations about the policy rate and, second, to effectively interpret these announcement-triggered revisions as *the* monetary policy shock, up to scale and a random measurement error. This makes the surprises valid external instruments for the identification of the contemporaneous transmission coefficients.

This paper produces evidence that challenges both these assumptions and argues that under general conditions, and independent from the length of the measurement window, monetary surprises capture more than just the monetary policy shock. Because private sector forecasts are not bound to be, and are generally not equal to central banks' forecasts, what markets label as unexpected may or may not be unanticipated by the central bank. In other words, monetary surprises can incorporate anticipatory effects if market participants are not able to correctly account for the systematic component of policy when they are surprised by a policy announcement. The dependence of monetary sur-

prises on forecasts of macroeconomic fundamentals also makes the risk premium change within the measurement window. We document a new stylized fact, namely, that other than being dependent on central banks' forecasts, as also noted in Gertler and Karadi (2015) and Ramey (2016), high-frequency monetary surprises are predictable by public data whose release *predates* the announcement. We interpret these results as indicating also the presence of a time-varying risk premium that is bound to change at the time of the announcement, because of the partial resolution of uncertainty about the future path of policy, and of macroeconomic conditions more generally, that is triggered by the policy decision. This casts doubts on the exogeneity of the monetary surprises as external instruments for the monetary policy shock. The resulting violation of the key identifying assumptions induces potentially non-trivial distortions in the estimated contemporaneous transmission coefficients. Consequences for the estimation of structural impulse response functions (IRFs) can be dramatic, both qualitatively and quantitatively.

A contractionary monetary policy shock that materializes as an increase in the policy rate induces a contraction in output and prices. In such a scenario, forecasts are likely to be revised downward, and a counter-cyclical risk premium to increase following the announcement. However, an increase in the policy rate may just as well be a signal that the central bank is anticipating buoyant times, along with inflationary pressures. If this is the case, forecasts will be revised upward and the premium will contract as result. Depending on whether market participants see the interest rate move as the result of a monetary policy shock, or as part of the systematic reaction of policy to the economic outlook, their economic forecasts and the premium will change in opposite directions, inevitably altering the signal in the monetary surprises. In the first case the two effects reinforce one another. In the second case, they have opposite effects on the futures price reaction to the announcement. As a consequence, we show that futures prices can rise in response to expected monetary easing, and that they can be revised in opposite directions following seemingly equivalent and expected no-change announcements (Section 3).

Under the assumption that market participants know the central bank's reaction function, the monetary policy surprise extracted from a futures contract that pays the

nominal interest rate i_t at the end of the current month can be expressed as (Section 2)

$$\begin{aligned} mps_t &\equiv p_t - p_{t-\Delta t} \\ &= f\left(\widehat{\Omega}^{CB}_{t|t} - \widehat{\Omega}^{M}_{t|t}\right) + \zeta\left(\widehat{\Omega}^{CB}_{t|t} - \widehat{\Omega}^{M}_{t|t}\right) + e_t, \end{aligned} \tag{1}$$

where e_t is the monetary policy shock, $f(\cdot)$ denotes the monetary policy rule, and ζ is the time-varying risk premium. $\widehat{\Omega}^{CB}_{t|t}$ and $\widehat{\Omega}^{M}_{t|t}$ denote the nowcast of macroeconomic fundamentals of the central bank and of market participants respectively. Equation (1) is central to our argument and establishes that a necessary condition for the monetary surprise to be a contemporaneous function of the monetary policy shock only, and thus be a valid external instrument for its identification, is that the central bank and market participants share the same forecasts about current and future macroeconomic fundamentals. This assumption has been challenged in a number of papers, starting with the seminal contribution of Romer and Romer (2000). Note that for (1) to hold it suffices that the central bank and market participants only differ in the forecasting model they choose to employ, everything else being equally known to both.

As a result, monetary 'surprises' depend on and are predictable by both central banks' forecasts and past, public information that was available to market participants well before the announcement. The predictability of monetary surprises – effectively market returns realized over a tiny time interval – can be naturally interpreted as indicating the presence of a time-varying risk premium (see e.g. Fama and French, 1989; Fama, 1990, 2013).

Building on the predictability of monetary surprises, this paper proposes to construct conditional futures-based surprises to be used for the identification of monetary policy shocks by projecting the raw surprises onto a set of central banks' forecasts and forecasts revisions of the key variables that are likely to enter the central bank's reaction function. The composition of the conditioning set, similar to the one in Romer and Romer (2004), also ensures that systematic reactions to interest rate changes of either sign are duly controlled for. Moreover, if the futures contract is on an interest rate other than the overnight (e.g. LIBOR), we include a correction term that takes into account the

discrepancy between the two. The resulting market-based proxies, free of anticipatory effects by construction, are shown to be uncorrelated with summary measures of the information available to the public, suggesting that the procedure effectively removes the time-varying risk premium that otherwise plagues market surprises. Projection of the conditional surprises onto a set of lagged macro-financial factors produces F-statistics well below critical values, suggesting that the conditioning set is sufficient to remove the dependence of monetary surprises on past information. The orthogonal proxies proposed in this paper are thus better candidates for the task of capturing only the unanticipated monetary policy shock. The importance of purging anticipatory effects from monetary surprises stems from these causing opposite responses of key macroeconomic variables such as output, unemployment and prices, when compared to a monetary policy shock. Campbell, Evans, Fisher and Justiniano (2012); Nakamura and Steinsson (2013); Campbell, Fisher, Justiniano and Melosi (2016) document compelling evidence in the context of forward guidance using survey-based forecasts.

Applications to both the US and the UK show that the use of the orthogonal proxies, contrary to the raw monetary surprises, allows to retrieve economically sound responses of the main output and price variables even in small, potentially informationally insufficient monetary VARs.

This paper extends the work of Barakchian and Crowe (2013), who are the first to discuss the assumption of equivalence between private sector forecasts and central banks' forecasts in the identification of monetary policy shocks using daily surprises in futures markets.

Early uses of financial market instruments to extract expectations about the path of short-term interest rates date back at least to the early nineties (see e.g. Cook and Hahn, 1989; Svensson, 1994; Soderlind and Svensson, 1997; Kuttner, 2001; Cochrane and Piazzesi, 2002; Piazzesi, 2002). Rudebusch (1998) was the first to suggest the inclusion of futures on interest rates in monetary VARs to overcome the potentially misspecified reaction function implicitly estimated in these models. Estimates of the unexpected component of policy have become more sophisticated with the availability of high-frequency financial data (Sack, 2004; Gürkaynak, 2005; Gürkaynak, Sack and Swanson, 2005). Gertler

and Karadi (2015) are the first to use monetary surprises as external instruments for the monetary policy shock in a Proxy Structural VAR (Stock and Watson, 2012; Mertens and Ravn, 2013). The availability of potentially clean measures of monetary shocks has since spurred a number of diverse applications whereby monetary surprises extracted from financial market instruments have been used to quantify the effects of both conventional and unconventional monetary policy shocks. To mention just a few, Hanson and Stein (2015) find large responses of long-term real rates to monetary policy shocks and explore the transmission of monetary policy to real term premia using intraday changes in the two-year nominal yield. Nakamura and Steinsson (2013) employ a 'policy news shock' – defined as the first principal component of monetary surprises calculated using a selection of interest rate futures – to show that long-term nominal and real rates respond roughly one to one to monetary policy shocks. Similarly, Swanson (2015) identifies 'forward guidance' and 'large-scale asset purchases' dimensions of monetary policy shocks at the zero lower bound using principal components of a selection of futures on short-term interest rates and long-term government bond yields, and employs them to study the effects of unconventional monetary policy on asset prices. Glick and Leduc (2015) use monetary surprises in federal funds futures and a collection of Treasury rate futures at longer maturities to study the effects of conventional and unconventional monetary policy on the dollar. Finally, Rogers, Scotti and Wright (2014) measure the pass-through of unconventional monetary policy implemented by four different central banks on asset prices by using monetary surprises calculated from long-term government bond yields in each of the monetary areas considered.

Miranda-Agrippino and Ricco (2016) relax some of the assumptions in this paper and use Bayesian Local Projection (BLP) to study the transmission of monetary policy shocks in the presence of informational frictions – including the challenges to identification posed by central bank signalling and by the slow and imperfect absorption of information by market participants. A transformation of monetary surprises is there used to construct an identification strategy that is robust to non-nested information sets of the central bank and private agents.

The paper is organized as follows. The theoretical framework is presented in Sec-

tion 2. Section 3 provides some illustrative evidence on the intraday adjustments of futures prices on monetary announcements days, and on the potential distortions caused to the estimated IRFs when raw monetary surprises are used to identify monetary policy shocks. Section 4 discusses the properties that the candidate instrument for the structural shock is required to have for the contemporaneous transmission coefficients to be consistently estimated in a Proxy SVAR; results on surprises predictability are in Section 5. Section 6 discusses the construction of the orthogonal surprises and illustrates impulse responses to monetary shocks identified using these measures as external proxies for the shock. Section 7 concludes. Additional results and technical details on the construction of the monetary surprises are reported in the Appendixes at the end of the paper and in the Online Appendix available at http://www.bankofengland.co.uk/research/Documents/workingpapers/2016/swp626appendix.pdf.

2 Theoretical Background

In this section we introduce a simple illustrative framework to provide the intuition behind the informational content of monetary surprises. The model we refer to is a workhorse three-equation New-Keynesian model (see Woodford, 2003; Galí, 2008, for textbook treatment).

Consider an economy in which the behaviour of households and firms is described by the following two equations:

$$x_t = x_{t+1|t} - \sigma\left(i_t - \pi_{t+1|t} - r_t^n\right), \tag{2}$$

$$\pi_t = \beta\pi_{t+1|t} + \kappa x_t. \tag{3}$$

Equation (2) is obtained from the linearized Euler equation and expresses the current output gap x_t as a function of the expected output gap $x_{t+1|t} \equiv \mathbb{E}_t[x_{t+1}]$ and of future expected deviations of the real interest rate from its natural rate r_t^n. The output gap is defined as the difference between the actual level of output and its 'natural' rate that would prevail with fully flexible prices. Within this simple model, the natural rates of

interest and of output are both functions of exogenous shocks to technology and preferences. One could think of richer frameworks where the natural interest rate is also a function of other shocks, such as to households borrowing constraints, or to the financial sector, without altering the essence of the argument discussed below. The parameter σ denotes the intertemporal elasticity of substitution.

The behaviour of inflation is regulated by the Phillips curve – Equation (3), that expresses current inflation as a function of future expected inflation, and of the current output gap. The parameter κ regulates the size of the response of inflation to changes in the output gap.

The central bank is assumed to set the interest rate according to the following simple rule:

$$r_t \equiv i_t - \pi_{t+1|t} = r_t^n + e_t, \tag{4}$$

therefore, the monetary authority chooses the real interest rate in such a way to track the natural rate of interest, with deviations from the rule denoted by e_t.[1]

Solving equations (2) and (3) forward one obtains

$$x_t = -\sigma \sum_{j=0}^{\infty} \left(r_{t+j|t} - r_{t+j|t}^n \right), \tag{5}$$

$$\pi_t = \kappa \sum_{j=0}^{\infty} \beta^j x_{t+j|t}. \tag{6}$$

Absent any monetary policy shock – i.e. if $e_t = 0$, the real interest rate equals the natural rate, and both the output gap and expected inflation are equal to zero. Conversely, a monetary policy tightening (loosening) will result in the real rate being larger (smaller) than the natural rate, a contraction (expansion) in economic activity, and a decline (rise) in inflation.

Within this framework, agents form expectations by projecting on current realizations

[1] For the sake of building the intuition, we choose to adopt the simple framework in Andrade and Ferroni (2016), however, the rule can be extended to include a Taylor principle as in Nakamura and Steinsson (2013).

of the shocks, of which current macroeconomic fundamentals are a contemporaneous function. The expected value of the level of the nominal interest rate just before the announcement can be expressed as

$$i_{t|t-\Delta t} = r^n_{t|t-\Delta t} + e_{t|t-\Delta t}, \tag{7}$$

where we assume that agents know that the central bank will revert to its rule following any shock, from which it follows that $\pi_{t+1|t-\Delta t} = 0$. The announcement is scheduled in the interval $(t - \Delta t, t)$. The forecast about the natural rate r^n_t can be further expressed as

$$r^n_{t|t-\Delta t} = \Theta \Omega_t, \tag{8}$$

whereΩ_t is the vector collecting the current realizations of macroeconomic fundamentals, and Θ is a non-linear function of primitive model parameters and denotes the coefficients of the projection.

In reality, however, the current value of macroeconomic fundamentals is not known in real time and must be estimated. Equation (8) thus transforms into

$$r^n_{t|t-\Delta t} = \Theta \widehat{\Omega}_{t|t}, \tag{9}$$

where $\widehat{\Omega}_{t|t}$ denotes the nowcast ofΩ_t. It is assumed that in the time interval Δt no news relative to macro fundamentals are released to the public, that is, the monetary announcement is the only event in the measurement window. In the absence of competing data releases, and conditional on the forecasting model being unchanged, $\widehat{\Omega}_{t|t} = \widehat{\Omega}_{t|t-\Delta t}$.

Consider the price of a futures contract on the nominal interest rate that pays the rate prevailing at some future date $t + h$

$$p^{(h)}_t = \mathbb{E}_t[i_{t+h}] + \zeta^{(h)}_t, \tag{10}$$

where $\zeta^{(h)}_t$ denotes the risk premium that may be present in the contract. Equation (10) expresses the price of the futures contract as a function of the expected future nominal

rate i_t plus a possibly time-varying risk compensation that investors require to hold such a contract to maturity.

Using (9) and (10) one can express the price just before the announcement as

$$p^{(h)}_{t-\Delta t} = i_{t+h|t-\Delta t}(\ \widehat{\Omega}_{t|t}\ ) + \zeta^{(h)}(\ \widehat{\Omega}_{t|t}\ ), \tag{11}$$

where the dependence on the economic forecasts is made explicit. In Equation (11) the time-variation in the risk premium is derived from the dependence of the premium on either realized or expected macroeconomic fundamentals.

Without loss of generality, consider the futures contract expiring at the end of the current month, i.e. the front contract. Assume that market participants have access to the same pool of public data as the central bank, and that they know the reaction function of monetary authority. The price that investors attach to such a contract just before the relevant monetary policy announcement is equal to

$$\begin{aligned} p_{t-\Delta t} &= i^{M}_{t|t-\Delta t} + \zeta(\ \widehat{\Omega}^{M}_{t|t}\ ), \\ &= f(\ \widehat{\Omega}^{M}_{t|t}\ ) + \zeta(\ \widehat{\Omega}^{M}_{t|t}\ ). \end{aligned} \tag{11$'$}$$

$i^{M}_{t|t-\Delta t}$ is the expected policy decision. Given market participants' forecasts about$\Omega_{\ t}$, and the central bank's reaction function f, what investors expect the interest rate to be after the announcement is equal to $f(\ \widehat{\Omega}^{M}_{t|t}\ )$. Conditional on the same set of forecasts, the risk premium equals $\zeta(\ \widehat{\Omega}^{M}_{t|t}\ )$.

After the policy decision is revealed, the futures price is updated accordingly

$$p_t = \underbrace{f(\ \widehat{\Omega}^{CB}_{t|t}\ ) + e_t}_{i_t} + \zeta(\ \widehat{\Omega}^{CB}_{t|t}\ ). \tag{12}$$

The new policy rate is a function of the central bank's forecast $\widehat{\Omega}^{CB}_{t|t}$ and of a possibly non-zero shock e_t. Consequently, the newly demanded risk premium is also revised. The risk compensation is associated to the uncertainty about the future path of policy, and of

macroeconomic conditions more generally; if the forecast forΩ_t changes, the risk premium that investors demand will reflect that change.

Monetary policy surprises around announcements are computed as the price update that follows the communication of the interest rate decision, that is, $mps_t \equiv p_t - p_{t-\Delta t}$.[2] It is then clear that, all else being equal, the fact that the economic forecasts of the central bank may not coincide with those of the private sector makes the surprises a contemporaneous function of more than just the monetary policy shock. In fact, for the price update to be mapped into the monetary policy shock it has to be the case that $\widehat{\Omega}^{CB}_{t|t} = \widehat{\Omega}^{M}_{t|t}$; in general, the monetary surprise will otherwise be equal to

$$\begin{aligned} mps_t &= p_t - p_{t-\Delta t} \\ &= f\left(\widehat{\Omega}^{CB}_{t|t} - \widehat{\Omega}^{M}_{t|t}\right) + \zeta\left(\widehat{\Omega}^{CB}_{t|t} - \widehat{\Omega}^{M}_{t|t}\right) + e_t. \end{aligned} \tag{1}$$

Monetary surprises can therefore incorporate anticipatory effects. The importance of purging these effects from monetary surprises is crucial for the analysis of the effects of e_t. If the central bank is raising the policy rate because it anticipates higher inflation or growth above potential, failing to account for the anticipation will result in misleadingly attributing the cause of higher growth and inflation to the higher interest rate.

Using Blue Chip forecasts, Nakamura and Steinsson (2013) show that a monetary contraction naturally sees agents revise their expectations about the short-term nominal interest rate upward. The same data on real output growth and inflation, however, reveal that the same contraction also results in a significant upward revision for growth forecasts, up to about a year into the future. Expected inflation measured using the GDP deflator reacts with the expected sign, but the adjustments are not significant. Similar evidence is reported in Campbell, Evans, Fisher and Justiniano (2012) and Campbell, Fisher, Justiniano and Melosi (2016). Campbell et al. (2012) regress Blue Chip forecast revisions for unemployment and CPI inflation on market-based monetary surprises constructed extending the work of Gürkaynak, Sack and Swanson (2005). Private forecasts of current and future unemployment are revised downward following a monetary

[2]Specific details on futures on interest rates and their use in the construction of monetary surprises are in Appendix B.

tightening. Similarly, an unexpected rise in the policy rate sees professional forecasters expect higher inflation, albeit the evidence in this case is weaker. Campbell et al. (2012) argue that these puzzling responses are due to the fact that the central bank and the public are not equally well informed about macroeconomic fundamentals, that is, their forecasts differ. The policy decisions transfer knowledge about the central bank's forecasts in what are labelled as Delphic announcements, triggering private sector forecasts revisions of the 'wrong' sign. Campbell et al. (2016) explore this hypothesis further, and find that the counterintuitive sign of the responses of both output and unemployment is indeed explained by the asymmetry of beliefs entertained by the central bank (measured by Greenbook forecasts) and the public (summarized by the Blue Chip Economic Indicators).

3 A Closer Look at Announcement-Triggered Price Revisions

This section collects some illustrative evidence on the movements observed in interest rate futures on a selection of policy-relevant dates. The contracts used are those typically employed as a basis for the construction of monetary surprises.

The contract used for the UK is the next expiring short sterling (SS) future – or front contract – that can be either the one expiring within the current month [M0] or within the next month [M1] depending on the relative market liquidity; these contracts embed expectations about the policy rate up to a horizon of about three months.[3] For the US, the reference contract is the next expiring federal funds (FF) future; this is typically the one expiring within the current month [M0] unless the policy announcement is made at the end of the month, in which case the focus shifts to the second contract [M1]. Charts displaying the variation in the fourth FF contract, which has a maturity of three months, roughly matching the horizons in the SS discussed here, are nearly identical.

To aid with the interpretation of the charts, intraday futures variations are com-

[3]The market for futures on interest rates tends to be very thin in the days that immediately precede their expiry date, for this reason we switch to the next expiring one when the number of trades for the contract expiring in the current month is low. More details on short sterling futures are provided in Appendix B.

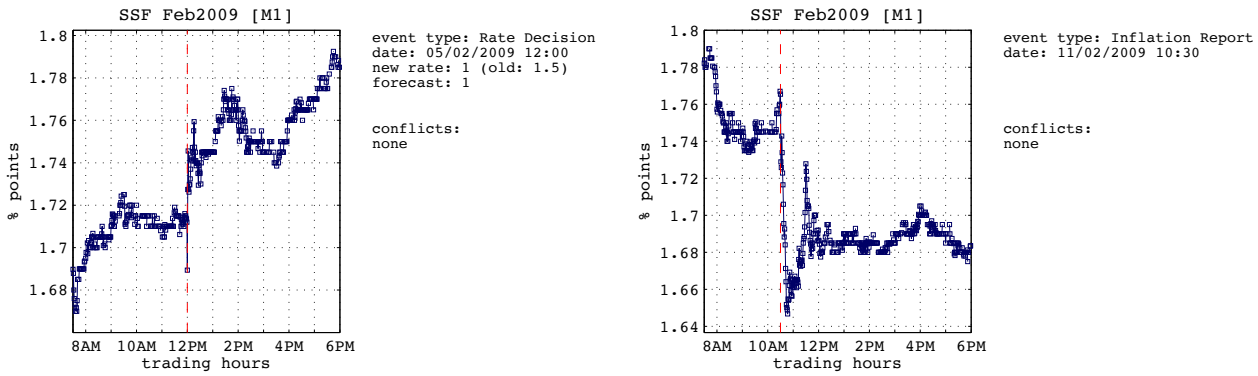


FIGURE 1: February 2009: The decision meeting of the MPC on the 5th [LEFT] is followed by the release of the Inflation Report on the 11th [RIGHT]. In each subplot, forecasts refer to the median of the Bloomberg Survey of Economists. Conflicts refer to major data releases scheduled within the hour surrounding the policy decision, marked with a vertical red dashed line. *Source:* Bloomberg and Thomson Reuters Tick History Database, author's calculations.

pared with expectations about the policy rate embedded in the median responses to the Bloomberg Survey of Economists (BSE).[4] To avoid interference of competing events that may contribute to alter expectations about both macroeconomic fundamentals and the upcoming interest rate decision, all the episodes discussed in this section are selected among those for which there are no conflicting contemporaneous data releases.

The charts in Figure 1 show how futures prices can raise in response to an expected monetary easing and how the counterintuitive price update can be rationalized against the background of the evolving economic environment, and the dependence on central bank's forecasts that provide context to the decision itself, released the week after the announcement.

On February 5th, 2009, the Bank of England's Monetary Policy Committee (MPC) voted to lower the policy rate by 50 basis points, to 1%. While the median forecast suggests that the move was largely anticipated, futures rates *rose* following the announcement (left panel of Figure 1). One possible explanation is that some players in the market were attaching some probability to an even larger cut in the policy rate. However, an equally plausible argument is that the move can be at least in part explained by an increase in the

[4]Survey-based expectations for all market-sensitive data are collected and published by Bloomberg over the two-week period immediately preceding all relevant data releases. Survey participants can contribute their forecasts up to 24 hours prior to the release itself and their views are collected for a variety of macroeconomic data releases, including the policy rate.

risk premium prompted by the stream of news of deteriorating economic (and financial) conditions that were hitting the markets, and reflected in the sizeable degree of uncertainty that surrounded the expected outcome of this policy decision.[5,6] This particular MPC meeting followed the release, on January 23, of the advance figure for real GDP growth relative to 2008 Q4, showing a contraction of 1.5% on a quarter-on-quarter basis, which had surprised market participants and institutional forecasters alike: the median Bloomberg forecast was at -1.2% on the day before the release, while the most recent World Economic Outlook, released on November 6th 2008, had it at a mere -0.5%; the IMF, however, were to release a new issue of the WEO only five days later, on January 28, where the estimate was revised downward to -1.8%.[7]

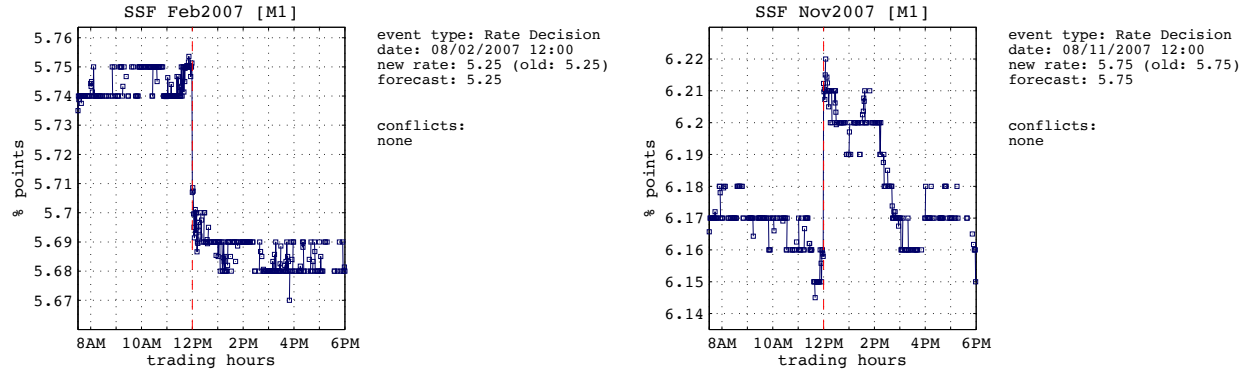
On the 11th of February, the Bank of England published its quarterly Inflation Report (IR). During the Opening Remarks at the start of the IR press release, at 10:30 AM, the then Governor King announced that the UK economy was in a deep recession and, more importantly, it became clear during the press conference that the MPC was likely to introduce further easing. The announcement this time induced a visible fall of futures quotes that fully reverted to the level at which they were prior to the interest rate decision. During the press conference the Governor stated that "*three weeks ago, the UK Government announced a five-point plan to restore the flow of lending. One of the five points is the creation of an asset purchase facility operated by the Bank of England and aimed at increasing the availability of corporate credit. The Bank of England will open its facility to make purchases later this week*", and that "*at its February meeting the Committee judged that an immediate reduction in Bank Rate of 0.5 percentage points to 1% was warranted. Given its remit to keep inflation on track to meet the 2% target in the medium term, the projections published by the Committee today imply that further easing in monetary policy might well be required.*"[8]

[5]Survey-based forecasts ranged from 0.5 to 1.5%, the median and average forecast were equal to 1%.

[6]A similarly puzzling reaction to the easing announcement was registered in the currency markets, where sterling rose following the announcement. *Source:* Bloomberg and Financial Times, Friday February 6, 2009.

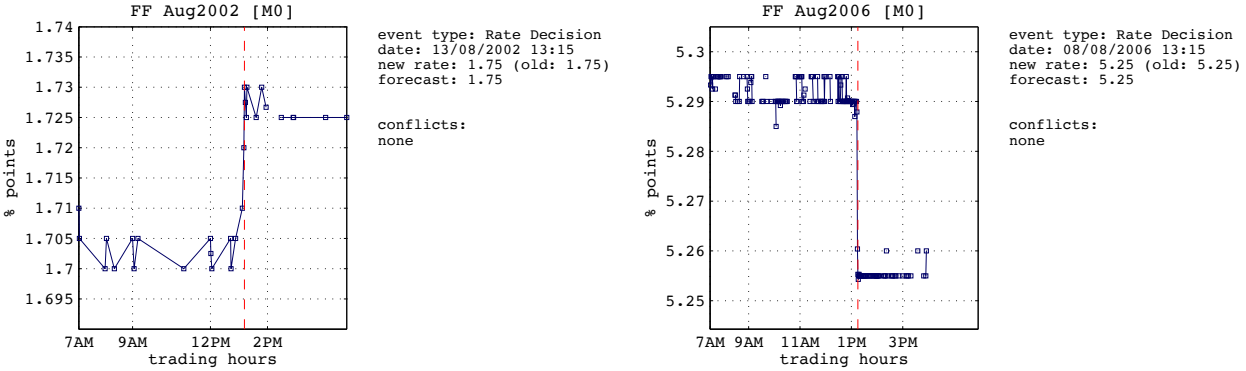
[7]Other significant data releases on the day of the MPC decision were the Halifax House Price Index for January at -17.20% on a 3month over year basis at 9:00 AM, US Jobless Claims relative to January at +38K compared to December at 1:30 PM, and US Factory Orders for December 2008 at -3.9% month-on-month, released at 3:30 PM.

[8]Both quotes are extracted from the opening remarks by the Governor delivered at the start of the press conference for the publication of the Inflation Report of the 11 February 2009, and available at



(A) MPC: fully anticipated no-change policy decision.

(B) MPC: fully anticipated no-change policy decision.



(C) FOMC: fully anticipated no-change policy decision.

(D) FOMC: fully anticipated no-change policy decision.

FIGURE 2: Fully anticipated no-change events triggering opposite reactions. In the first row, short sterling futures around MPC decisions to maintain Bank Rate at the previous level. In the bottom row, federal fund futures on FOMC announcement days where the level of the target fed funds rate was not changed. Forecasts refer to the median of the Bloomberg Survey of Economists. Conflicts refer to major data releases scheduled within the hour surrounding the policy decision, marked with a vertical red dashed line. *Source:* Bloomberg and Thomson Reuters Tick History Database, author's calculations.

A more striking picture is in Figure 2. All four episodes refer to days in which the Bank of England and the Federal Reserve decided not to change the level of the policy rate.

In the top row, the Bank of England's MPC maintained Bank Rate at the previous level, both on February 8, 2007 and on November 8 of the same year, at 5.25 and 5.5% respectively. The same is true for the charts in the bottom row. The Fed's Federal Open Market Committee (FOMC) agreed not to change the target interest rate both on August 13, 2002 and on August 8, 2006, leaving it at 1.75 and 5.25% respectively. The

http://www.bankofengland.co.uk/publications/Pages/inflationreport/ir0901.aspx

median Bloomberg forecasts reveal that market participants expected both the MPC and the FOMC not to move the policy rate in all instances, which makes these four moves largely anticipated.[9] Recall also that in none of the four selected cases other relevant data releases were scheduled in the hour surrounding the central bank decision. Why then are market prices reacting at all?

The salient feature of Figure 2 is that it shows how not only markets can and do react to fully anticipated moves, but they can also do so in different directions. While this is hard to reconcile with a framework in which investors and the central bank share the same forecasts and prices only adjust following revision in expectation triggered by an unexpected policy rate change, it can be rationalized by allowing market participants and the central bank to entertain different beliefs about the state of the economy, and the premium to change within the measurement window.

The episodes in Figures 1 and 2 provide suggestive evidence in support of the decomposition in Equation (1), where the monetary surprises are expressed as potentially a contemporaneous function of more than just the monetary policy shock. If one is willing to accept this interpretation, it is then easy to see that if the VAR in use does not properly account for future expectations and premia (e.g. by including them in the set of endogenous variables), proxying for monetary policy shocks using futures-based price revisions can produce IRFs that are highly distorted. Figure 3 illustrates the point.

The IRFs in Figure 3 are an excerpt of those reported in Section 6 (Figure 6), and depict responses to a *contractionary* monetary policy shock identified using the average monetary surprise computed using the fourth federal funds future (FF4) as an external instrument.[10] The set of variables included in the VAR is the one in Coibion (2012), and the VAR is estimated in levels over the sample 1969-2014 using 12 lags. The identification is borrowed from Gertler and Karadi (2015) and uses the 1-year rate as the monetary policy (endogenous) variable and the average FF4-based surprise as a proxy for the shock. Contrary to Gertler and Karadi (2015), however, the specification of the VAR intentionally leaves out the Excess Bond Premium of Gilchrist and Zakrajšek (2012). Note here

[9] Forecast ranges for the four episodes are (clockwise from top left panel): 5.25 - 5.5%, 5.5 - .75%, 5.25 - 5.5%, 1.25 - 1.75%.

[10] The identification scheme is discussed in the next section and further detailed in Appendix A. See Section 6 for additional discussion on the VAR specification and the role of the variables included.

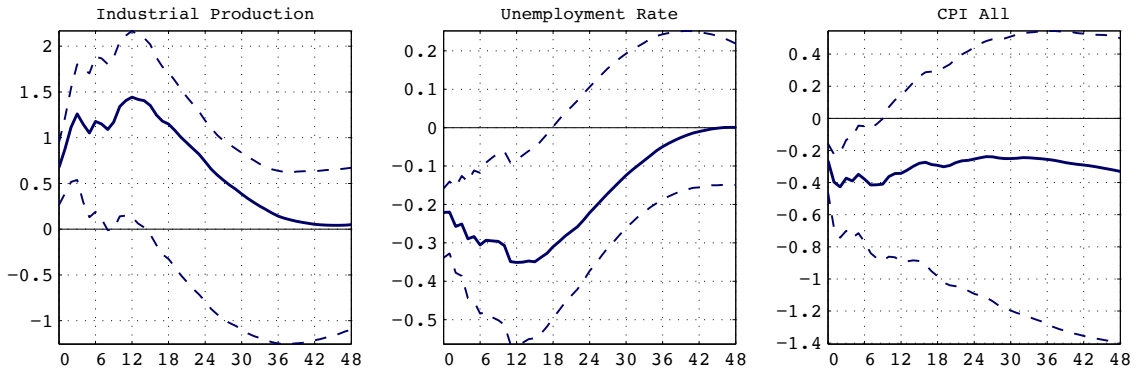


FIGURE 3: Distortion in contemporaneous transmission coefficients resulting from the use of raw monetary surprises as external proxies for the monetary policy shock. IRFs to a shock inducing a 100bp increase in the policy rate identified using the raw FF4-based proxy as an external instrument. VAR(12) estimated in levels over 1969:1 - 2014:12. The monetary policy variable is the 1-year rate. Dashed lines show 90% confidence bands obtained using 10,000 bootstrap replications. The full set of IRFs is in Figure 6.

that the composition of the VAR should be of little practical relevance, were the chosen external instrument truly a sole function of the monetary policy shock. The shock is normalized to induce a 1% *increase* in the policy rate.

According to the figure, a contractionary monetary policy shock induces a significant and persistent increase in output and an equally sizeable reduction in unemployment, while prices slightly contract. The sign of these responses is reminiscent of the responses of survey-based forecast revisions that were discussed in Section 2, suggesting that similar mechanisms may well be at play. We therefore interpret these anomalous responses as reflecting the extent to which confounding the shocks can induce distortions in the estimates of the contemporaneous transmission coefficients. An increase in the policy rate may be signalling that the central bank is forecasting improved economic conditions ahead, hence explaining the sign of the responses. Conversely, interpreting the sign of the effect of change in the risk premium is less obvious. If premia are assumed to be countercyclical (see e.g. Campbell and Cochrane, 1999) a monetary contraction could likely induce an increase in risk aversion, leading to an amplified effect on output and prices. However, this need not necessarily be the case (De Paoli and Zabczyk, 2012).[11]

[11]The positive responses of output and unemployment are in this case amplified by the use of the average monthly markets surprise (i.e. the one in Gertler and Karadi, 2015). When the monthly sum of daily surprises is used instead, that is, no weighting is performed on the daily monetary surprises as is the case for example in Stock and Watson (2012) and Caldara and Herbst (2015), the expansionary effects of the signalling channel and the contractionary effects induced by the monetary policy shock

4 Proxies for Structural Shocks

Proxy SVARs achieve identification of the contemporaneous transmission coefficients that express reduced-form VAR innovations as a combination of the structural shocks, by using external proxies – not included in the set of endogenous variables – as instruments for the latent shocks (Stock and Watson, 2012; Mertens and Ravn, 2013; Montiel-Olea, Stock and Watson, 2016). This section briefly reviews the identification scheme that is then fully developed in Appendix A. A discussion on the desired properties of the external instrument is reported at the end of the section.

Let y_t be an n-dimensional vector of endogenous observables whose responses to the structural shocks in e_t are given by

$$y_t = [\, A(L) \,]^{-1} u_t = \mathcal{C}(L)\mathbf{B}e_t, \tag{13}$$

where $\mathcal{C}(L)\mathbf{B}$ are the structural impulse response functions. u_t are the reduced-form innovations with $u_t = \mathbf{B}e_t$ and $\mathcal{C}(L) = [\, A(L) \,]^{-1} \equiv [\, \mathbb{I}_n - A_1 L - \ldots - A_p L^p \,]^{-1}$, where A_i, $i = 1, \ldots p$, are the matrices containing the reduced-form autoregressive coefficients. $\mathbf{B}$ collects the contemporaneous transmission coefficients. The structural shocks are such that $\mathbb{E}[\, e_t \,] = 0$, $\mathbb{E}[\, e_t e_t' \,] = \mathbb{I}_n$ and $\mathbb{E}[\, e_t e_\tau' \,] = 0 \;\; \forall \tau \neq t$.

Suppose one is interested in calculating the responses of y_t to a particular shock in e_t, call it the monetary policy shock, and denote it by $e_t^\bullet$. The identification of the relevant column $\mathbf{b}^\bullet$ of $\mathbf{B}$ that links the reduced-form innovations to $e_t^\bullet$ is achieved via a set of variables z_t, not in y_t, such that

$$\mathbb{E}[\, e_t^\bullet z_t' \,] = \phi', \tag{14}$$

$$\mathbb{E}[\, e_t^\circ z_t' \,] = 0, \tag{15}$$

and ϕ is non-singular. e_t° denotes structural shocks other than the one of interest. If

balance out, resulting in muted responses at all horizons for both output and unemployment in the same VAR used here. IRFs for this case are not reported but available upon request. Further details on the weighting scheme are in Appendix B.

one or more variables z_t can be found such that these conditions are satisfied, then it is possible to identify $\mathbf{b}^\bullet$ up to scale and sign using only moments of observables.

Conditions (14) and (15) are the key identifying assumptions, and require that the proxy variables are correlated with the structural shock of interest, and that they are not correlated with all the other shocks. While these requirements resemble the standard conditions for external instruments' validity, it is important to notice that here the instruments need to be relevant and exogenous with respect to the *unobserved* structural shocks.

An equivalent way of addressing the identification of $\mathbf{b}^\bullet$ is to cast the problem in a measurement error framework where the structural shock of interest is treated as an unobserved regressor, and the external instrument is explicitly modelled as a proxy variable

$$z_t = \Phi e_t^\bullet + \nu_t, \tag{16}$$

where ν_t is an i.i.d. measurement error and Φ is non-singular. In this case, all the relevant model parameters (i.e. $A(L)$ and $\mathbf{B}$) are jointly estimated in an error-in-variable (EIV) system where z_t is effectively treated as a scaled version of the shock up to a random error.

Whilst in general there is no formal way to verify that (14) and (15) hold, the identification via external instruments also relies on a number of other requirements that only involve observables and are thus fully testable. The estimation of the EIV system delivers a consistent estimate of the transmission coefficients only if the instrument is uncorrelated with the lagged endogenous variables included in the VAR, that is

$$\mathbb{E}[\, z_t \mathcal{Y}_t' \,] = 0, \tag{17}$$

with $\mathcal{Y}_t \equiv [y_{t-1}', \ldots, y_{t-p}']'$.[12] Furthermore, (16) implies that just like the shock itself, the proxy variable should not be forecastable given lagged information relative to own lags or lags of any other variable, regardless of whether it is included in y_t or not. These conditions resemble the informational sufficiency requirement on the observables included in any structural VAR, and call for the absence of any endogenous variation in the

[12] See Appendix A.

dynamics of z_t. The intuition here is that if this is not the case, then there is no reason why one would not want to include z_t in the set of endogenous observables y_t and let it act as an instrument for itself (see e.g. Bagliano and Favero, 1999; Barakchian and Crowe, 2013). In fact, an equivalent way of estimating the transmission coefficients is to include z_t in the set of endogenous observables and identify the monetary policy shock by ordering it first in a standard Cholesky triangularization.[13]

The orthogonality requirement in (17) can be relaxed if the estimation of the contemporaneous transmission coefficients is achieved with a two-step procedure, rather than within the EIV system. In this case, the VAR is estimated in the first stage, and then the reduced-form residuals u_t, orthogonal to $\mathcal{Y}_t$ by construction, are projected onto the instruments to estimate the coefficients in $\mathbf{b}^{\bullet}$. If, however, $\mathbb{E}[\, z_t \mathrm{X}'_{t-1} \,] \neq 0$, where X_{t-1} is a set of variables omitted from the VAR specification, but such that $\mathbb{E}[\, u_t \mathrm{X}'_{t-1} \,] \neq 0$, the two-step procedure will be misspecified, resulting in potentially severely biased estimates of the parameters in $\mathbf{b}^{\bullet}$. The discussion in the reminder of the paper, related to the predictability of the monetary surprises, will technically fall within this context.

Overall, empirically, a successful identification of the contemporaneous transmission coefficients is ultimately a question of both specifying the VAR correctly, and singling out a reasonably valid proxy. In the optimal case in which the instruments are truly a sole function of the structural shock of interest, the composition of the VAR should be of little practical relevance, and IRFs should be invariant to the number and type of variables included in the VAR. In practice, the evidence collected in this paper suggests this is not necessarily the case. Intuitively, if doubts arise about the effective exogeneity of the chosen instrument, one way to mitigate the distortions on the estimated contemporaneous transmission coefficients is to enrich the information set of the VAR itself to produce 'cleaner' innovations.

[13] The variance of z_t enters both the measures of instruments' reliability Λ – Equation (A.12), and the F-statistic customarily used to test the joint significance of the coefficients estimated in the regression of u_t onto z_t. The presence of autocorrelation can artificially increase both statistics leading to overstating the effective relevance of the instrument.

5 Predictable Surprises

This section addresses the concerns discussed in the previous sections more formally. In what follows, raw US surprises are those in Gürkaynak, Sack and Swanson (2005), extended until 2012. Namely, surprises are extracted from the first (MP1) and fourth (FF4) federal funds futures, and from the second (ED2), third (ED3) and fourth (ED4) Eurodollar futures. UK surprises are novel, and constructed using the next expiring short sterling futures (SS1). To assess the behaviour of market participants around policy-relevant events other than the rate announcements, UK raw surprises are also computed on extended sets of dates that add to the rate decision the release of the minutes of the MPC meetings (SS1M), and of the quarterly Inflation Report (SS1MIR). Because the latter events are often contemporaneous to major economic data releases that are also market movers, we control for all data releases which are scheduled within the measurement window. The reader is referred to Appendix B for a thorough description of the raw surprises and their time series properties, and of the financial instruments used for their construction.[14]

Tables 1 and 2 collect results from predictive regressions where the raw surprises are projected onto different sets of variables that are intended to summarize the information set of both market participants and the central bank. These regressions are motivated by the decomposition in Equation (1). Results show that monthly monetary surprises respond significantly to central banks' forecasts and forecasts revisions about output, inflation, and unemployment, in support of the view that central bank's forecasts do enter the specification. As discussed, investors' risk compensation is also likely to change within the measurement window as a consequence of the dependence on economic forecasts, implying that the surprises may be contaminated by a time-varying risk premium. As is standard in the finance literature, we test for the presence of time-varying risk premia by regressing the surprises (i.e. intraday returns) on a collection of macroeconomic and financial observables known to market participants prior to the announcement itself, and show that these are highly significant in explaining future surprises. Once the anticipa-

[14] Cesa-Bianchi, Thwaites and Vicondoa (2016) also use high-frequency data to construct proxies for monetary policy shocks in the UK; their proxies roughly correspond to the raw surprise calculated around all policy events constructed here (SS1MIR) and further discussed in Appendix B.

tory effects are controlled for, as is the case in the orthogonal surprises constructed in the next section, data released prior to the announcement lose their significance, as one would expect (see Section 2).

In the language of Section 4, here we test for $\mathbb{E}[\, z_t \mathrm{X}'_{t-1} \,] = 0$, where X_{t-1} is a collection of variables likely to be in the information set of either or both the central bank and market participants at the time of the monetary announcement.[15]

Formally, the tables report the adjusted R^2 and the F statistics for the regression

$$\underbrace{p_t - p_{t-\Delta t}}_{mps_t} = \kappa_{\mathrm{c}} + \kappa_{\mathrm{x}} \mathrm{X}_{\mathrm{t-1}} + \epsilon_{\mathrm{t}}, \tag{18}$$

where mps_t is the monetary policy surprise and X_{t-1} is a set of observables whose realization is known before the announcement (i.e.$\Delta\ t < 1$). Full regression outputs are collected in the Online Appendix, which also reports robustness checks, including those where the specification in (18) is augmented with the lagged monthly surprise and, for the US, those where the sample starts after 1994 to avoid the discontinuity introduced by the release of a statement accompanying the decision of the FOMC. The regressions are estimated in-sample and at monthly frequency. The length of the measurement window (Δt) is equal to 30 minutes, with the exception of the broad UK-based surprises that also cover the release of the minutes and of the Inflation Report (i.e. the SS1MIR case). When the IR is the relevant policy event, we set$\Delta\ t$ equal to 90 minutes to account for the duration of the IR press conference. When X_{t-1} contains either realized economic variables or estimated factors, these enter the specification with a month's lag. When predictability is tested against collections of forecasts, these are aligned such that the compilation of the forecast always precedes the monetary surprise.

The top row of Table 1 reports predictability results relative to a set of ten lagged macroeconomic and financial factors estimated from the 134 US monthly series assembled

[15] We abstract from concerns related to the design of trading strategies and out-of-sample predictability of monetary surprises that, while relevant in their own right, go beyond the scope of the present analysis.

in McCracken and Ng (2015).[16] Surprises are predictable by past information, summarized by the lagged macro-financial factors. Because raw surprises are effectively market returns realized over a tiny time span, significant predictability with respect to lagged observables and factors can naturally be interpreted as suggesting that the surprises are contaminated by a time-varying risk premium. Individual t-statistics (not reported, see Online Appendix) are significant at least at the 5% level for three out of the ten factors and for all the raw surprises. The joint null of no predictability (reported in the top row of the table) is rejected at the 1% level in almost all cases.

One concern with regressing on these factors is that they are estimated on the last available vintage of data, that thus includes revisions that occurred after the surprise was measured. Moreover, due to the sometimes significant delay with which data are released, the information set from which the factors are extracted was not entirely visible at the time of the announcements, even if factors are lagged one month. In order to assess the predictability of surprises using data that were effectively available at the time of the announcement, the middle panel of Table 1 reports results of individual regressions on a subset of the variables used for the factors extraction. Lagged observables are taken in first differences with the exception of surveys and spreads. Both surveys and financial data, which are not subject to revisions and relative to the month prior to the announcement, are significantly predictive of future monthly surprises. These results complement the findings in Piazzesi and Swanson (2008) and suggest that monthly monetary surprises seem to be significantly contaminated by time variation in risk premia.[17,18]

[16] Factors are obtained by estimating a Dynamic Factor Model (Forni et al., 2000; Stock and Watson, 2002) with VAR(1) dynamics and diagonal idiosyncratic variance. Maximum likelihood estimates of the factors, their variances and model parameters are obtained using the EM algorithm and Kalman filter for the DFM cast in state space form, and iterating until convergence. The algorithm is initialized with static principal components and least squares estimates for the state space parameters. Prior to estimation, all variables are opportunely transformed to achieve stationarity.

[17] Results on predictability survive for samples starting after 1994 and if the surprises are computed using only scheduled FOMC meetings (reported in the Online Appendix). The dates of the unscheduled FOMC meetings, taken from Lucca and Moench (2015), are April 18, 1994, October 15, 1998, January 3, 2001, April 18, 2001, September 17, 2001, January 21, 2008 and October 7, 2008.

[18] Piazzesi and Swanson (2008) regress the daily surprises in Kuttner (2001) on Treasury yield spreads over the sample 1994-2005 and fail to reject the null of no-predictability at daily frequency.

	$MP1_t$		$FF4_t$		$ED2_t$		$ED3_t$		$ED4_t$	
	R^2	F	R^2	F	R^2	F	R^2	F	R^2	F
Macro-Financial Factors	0.042	2.040**	0.078	2.994***	0.095	3.502***	0.068	2.739***	0.055	2.379***
Lagged Observables										
ISM Composite	0.0394	10.71***	0.0682	18.36***	0.0964	26.28***	0.0802	21.66***	0.0673	18.11***
Consumer Sentiment	0.0103	3.46*	0.0221	6.35**	0.0374	10.22***	0.035	9.70***	0.0273	7.65***
Effective Fed Funds Rate	0.016	4.88***	0.0437	11.78***	0.076	20.40***	0.0618	16.56***	0.0517	13.87***
3M T-bill FFR Spread	0.0839	22.72***	0.0464	12.53***	0.0233	6.64**	0.019	5.50**	0.0123	3.95**
AAA-FFR Spread	0.006	2.36	0.003	0.23	0.003	0.25	0.003	0.23	0.003	0.24
1Y T-Bond FFR Spread	0.0549	14.76***	0.036	9.84***	0.0225	6.47**	0.0181	5.37**	0.0134	4.23**
S&P 500 Composite	0.0108	3.59*	0.0203	5.89**	0.0404	10.95***	0.049	13.22***	0.0416	11.25***
CBOE VIX	0.0214	6.19**	0.0239	6.79***	0.0682	18.33***	0.0685	18.42***	0.063	16.93***
Greenbook Forecasts										
Output	0.056	3.825**	0.083	5.301***	0.118	7.379***	0.092	5.832***	0.057	3.856**
Inflation	0.007	1.359	0.015	1.712	0.007	1.329	0.015	1.747	0.005	1.237
Unemployment	0.038	2.863**	0.048	3.414**	0.098	6.156***	0.056	3.827**	0.034	2.69**
Greenbook Forecasts Revisions										
Output	0.071	5.557***	0.078	6.019***	0.11	8.391***	0.079	6.114***	0.047	3.941***
Inflation	0.004	1.218	-0.009	0.468	-0.006	0.650	0.010	1.589	0.015	1.904
Unemployment	0.042	3.619**	0.088	6.763***	0.082	6.284***	0.051	4.230**	0.039	3.432**

TABLE 1: Predictability of US raw surprises. The table reports adjusted R^2 and F-statistics for the null $H_0 : \kappa_x = 0$ in (18) estimated at monthly frequency over the sample 1990:1 - 2012:6 (1990:1 - 2009:12 for Greenbook forecasts). Variables in X_{t-1} are listed in the first column. The ten factors are extracted from the set of 134 monthly macroeconomic and financial variables in McCracken and Ng (2015). Lagged observables are taken in first difference with the exception of surveys and spreads. ***, ** and * denote significance at 1, 5 and 10% level respectively. The raw monetary surprises are extracted from the first and fourth fed funds futures (MP1 and FF4 and the second, third and fourth Eurodollar future (ED2, ED3, ED4). Monthly raw surprises are obtained as the sum of the daily series in Gürkaynak et al. (2005). See main text for details. Full regression output is reported in the Online Appendix.

The bottom panel of Table 1 reports predictability results relative to Greenbook forecasts and forecast revisions for output, inflation and unemployment for the previous and current quarter and up to three quarters ahead. Greenbook forecasts are aligned to match the FOMC meeting they refer to. Results in the table confirm that forecasts and successive forecast revisions for output and unemployment are highly informative for all the monetary surprises considered, as noted also in Gertler and Karadi (2015) and Ramey (2016).

Results referring to UK-based surprises are in Table 2, where the same data transformations adopted for the case of the US are used. The five monthly factors are extracted from a set of 47 macroeconomic and financial indicators selected to be a UK counterpart of the set in McCracken and Ng (2015).[19] As was the case for the US, there is evidence that monthly surprises extracted from short sterling futures are also contaminated by time-varying risk premia. On the other hand, the evidence of predictability with respect to the forecasts and forecasts revisions contained in the Inflation Report (IR) is more mixed.[20] F statistics reported in Table 2 refer to the case in which forecasts and revisions are all included in the same regression, however, specifications where these are alternatively included turn out to be more inconclusive. In particular, while F statistics are still above critical levels, individual significance is less obvious, potentially due to the forecasts being highly correlated among them. Moreover, the quarterly availability of the Report and the shorter estimation sample imply that these estimates are based on a smaller number of observations compared to the US case, which makes them necessarily more uncertain. UK raw surprises are available only since June 1997, a date chosen to coincide with the first decision meeting after the Bank of England's MPC was granted operational independence for setting monetary policy. Complementary evidence is reported in Figure B.3 in Appendix B, where the SS1 and SS1MIR series are plotted. As shown, expanding the set of policy events to include the minutes and the IR does not seem to alter the overall informational content of the SS1-based monthly surprise series. We take

[19] The complete list of data and the transformations applied prior to the factor extraction are reported in the Online Appendix.

[20] Inflation Report forecasts are obtained by conditioning on a market-based path for the interest rate. This conditioning is not a cause of concern in the present case, however, since it is made on market data which are realized prior to the compilation of the forecast itself.

	$SS1_t$		$SS1M_t$		$SS1MIR_t$	
	R^2	F	R^2	F	R^2	F
Macro-Financial Factors	0.044	2.390**	0.045	2.417**	0.044	2.395**
Lagged Observables						
PMI Composite	0.001	0.7	0.003	0.33	0.001	0.80
CPI All Items	0.0322	7.99***	0.0333	8.23***	0.0298	7.46***
Consumer Confidence	0.000	0.000	0.005	0.01	0.001	0.740
Bank Rate	0.007	2.47	0.0097	3.06*	0.001	1.17
FTSE All Share	0.019	4.95**	0.016	4.40**	0.025	6.34**
3M LIBOR	0.031	7.69***	0.0351	8.63***	0.024	6.13**
3M T-bill Spread	0.094	22.75***	0.102	24.86***	0.108	26.29***
1Y Gilt Spread	0.061	14.71***	0.065	15.61***	0.058	13.93***
Official Reserves	0.025	6.42**	0.025	6.29**	0.028	7.03***
IR Forecasts and Revisions						
Output	0.098	1.938*	0.121	0.121**	0.195	3.088**
Inflation	0.131	2.297**	0.161	2.658**	0.165	2.702**
Unemployment	0.132	2.316**	0.113	2.094**	0.192	3.048***

TABLE 2: Predictability of UK raw surprises. The table reports adjusted R^2 and F statistics for the null $H_0 : \kappa_x = 0$ in (18) estimated at monthly frequency over the sample 1997:1 - 2014:12. Variables in X_{t-1} are listed in the first column. The five macro-financial factors are extracted from a set of 47 monthly macroeconomic and financial variables. Lagged observables are taken in first difference with the exception of surveys and spreads. ***, ** and * denote significance at 1, 5 and 10% level respectively. The raw monetary surprises are extracted from the first short sterling future and computed around rate announcement only (SS1), rate decision and release of the minutes (SS1M), rate decision, release of the minutes and of the Inflation Report (SS1MIR). All raw surprise series control for contemporaneous data release. See Appendix B for details on UK-based raw surprises. Full regression output is reported in the Online Appendix.

this as evidence of the fact that monetary surprises are a function of the central bank's assessment of current and future economic conditions, despite the lack of significance of the individual coefficients of some of the IR forecasts.

6 Orthogonal Monetary Surprises and Shock Identification

Consistent with the predictions made in Section 2, the results collected in the previous section suggest that monthly monetary surprises extracted from futures on interest rates cannot be safely used as external instruments for the monetary policy shock unconditionally. The mere fact of narrowing down the measurement window to a short time

span surrounding the time of the announcement does not guarantee that the computed surprises are a clean measure of the underlying monetary policy shock.

6.1 Orthogonal Monetary Surprises

To construct conditional futures-based surprises to be used for the identification of monetary policy shocks, we propose to project the raw surprises onto a set of forecasts and forecast revisions of the key variables that are likely to enter the central bank's reaction function. The composition of the conditioning set is motivated by the decomposition in Equation (1) and similar to the one in Romer and Romer (2004). The orthogonal monetary policy surprises ($mps_t^\star$) are defined as the residuals of the following regression estimated at monthly frequency:

$$mps_t = \mu + \underbrace{\alpha i_{t-1} + \beta \Delta i_t}_{\text{rate decisions}} + \underbrace{\sum_{j=-1}^{3} \gamma_j \, \widehat{\Omega}^{CB}_{t|q+j}}_{\text{level forecasts}} + \underbrace{\sum_{j=-1}^{2} \delta_j \left[\, \widehat{\Omega}^{CB}_{t|q+j} - \widehat{\Omega}^{CB}_{t-1|q+j} \, \right]}_{\text{forecast revisions}} + mps_t^\star \, . \quad (19)$$

To proxy for the information included in the central bank's reaction function at the time of the announcement, we use staff forecasts produced ahead of policy meetings for output, inflation, and unemployment. Forecasts horizons considered are the previous and current quarter, and up to three quarters ahead. We include in the conditioning set both the level forecasts and forecast revisions between consecutive forecast dates. Depending on the release schedule of the variables of interest, these forecasts are substituted with actually released data whenever they become available. As in Romer and Romer (2004), we only include unemployment nowcasts into the conditioning set, due to the strong correlation between output growth and unemployment. Other components of the conditioning set are the lagged level of the target interest rate and the rate decision itself. This is included to control for any systematic response of market participants to policy decisions of either sign. Lastly, if the future contract is on an interest rate other than the overnight (e.g. LIBOR), we augment (19) with a correction term that takes into account the discrepancy between the two.

The proposed approach for the construction of the orthogonal surprises has three main advantages: (i) it transforms the proxies ex ante, such that they are less dependent

on the composition of the information set in the preferred reduced-form monetary VAR; (*ii*) the variables that enter the conditioning set are either unrevised or have a trackable revision history, meaning that the conditioning can be carefully done to ensure that the different information sets are properly aligned at all times; (*iii*) it includes a minimum set of controls to ensure that the proxies are effectively capturing surprises orthogonal to all the available information, and that result from policy decisions that are not taken in response to either current or future economic developments. Projection of the orthogonal surprises on the same set of macro-financial factors used in Section 5 produces F statistics well below critical values, suggesting that the conditioning set is sufficient to remove the dependence of monetary surprises on past information.

For the US, the conditioning set contains (a) Greenbook forecasts and forecast revisions for output and inflation for the previous and the current quarter and up to three quarters ahead, and of current unemployment; and (b) the lagged federal funds rate and the observed change in the target interest rate. Regressions of the orthogonal proxies on the same set of ten lagged factors extracted in the previous section produce F statistics all below critical levels; specifically, MP1: F =0.775 (p-value 0.653); FF4: 1.162 (0.318); ED2: 1.498 (0.141); ED3: 1.212 (0.284); ED4: 1.069 (0.387). The raw (FF4) and orthogonal (FF4*) monthly surprises extracted from the fourth federal funds future are plotted in Figure 4 for the period 1990-2009. The upper time bound to the construction of the orthogonal surprises is constrained by the 5-year publication lag of the Greenbook forecasts and more generally motivated by the fed funds rate reaching the zero lower bound in 2009.

Measuring responses to a monetary policy shock in the UK using high-frequency futures data presents some difficulties, primarily related to the fact that no financial contract with a sufficiently long history is directly linked to Bank Rate.[21] A further complication in the present context arises from the fact that, contrary to the case of the US, over the sample considered the Bank of England's MPC meets twelve times a year, while official forecasts are updated once a quarter. The conditioning set over which the orthogonal monetary surprises are calculated is in this case composed by (a) forecasts and

[21] See Appendix B.

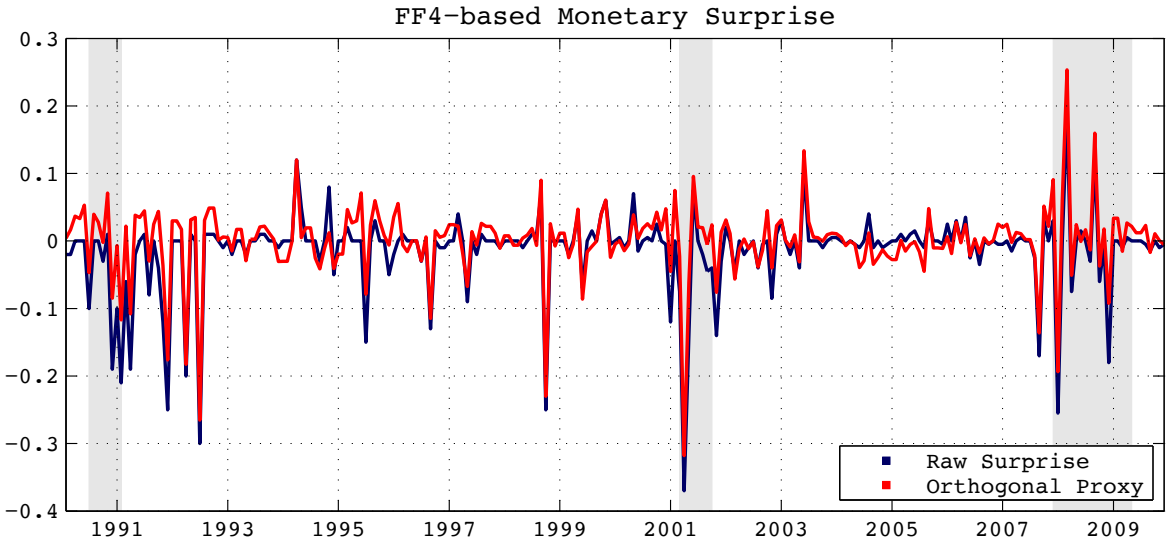


FIGURE 4: Raw (FF4 – blue line) and orthogonal (FF4* – red line) monetary surprises for the US at monthly frequency. Both sets of surprises are extracted from the fourth federal funds future. Shaded areas denote NBER recessions. See main text for details.

forecast revisions for output and inflation for the previous and the current quarter and up to three quarters ahead, and for current unemployment, extracted from the quarterly Inflation Report, and (b) the lagged Bank Rate, the lagged level of the LIBOR-OIS spread, and the observed change in the target interest rate. The use of Inflation Report forecasts is also used in Cloyne and Hürtgen (2014) to construct a narrative account of UK monetary policy decisions not taken in response to current and forecast macroeconomic conditions in the spirit of Romer and Romer (2004). The inclusion of the LIBOR-OIS spread is intended to partially offset the fact that the contracts used to extract the surprises are not a direct function of the interest rate set by the MPC. Being linked to the sterling LIBOR, the raw surprises in short sterling futures are rather a measure of the expected change in the 3-month interbank rate and, to the extent that the relation between the two rates is neither zero, nor constant, it needs to be controlled for when extracting revisions in expectations about the policy rate.[22] The raw UK monetary surprise used is the one computed around rate announcements only. The orthogonal surprise SS1* is plotted in Figure 5 against its raw counterpart SS1 for the period 2001-2015. While IR

[22]See Figure B.2. Ideally, one would want the correction for the LIBOR-OIS spread to happen at the time of computing the surprises at intraday frequency; however, due to unavailability of intraday swap quotes for the selected period, the daily spread is used instead.

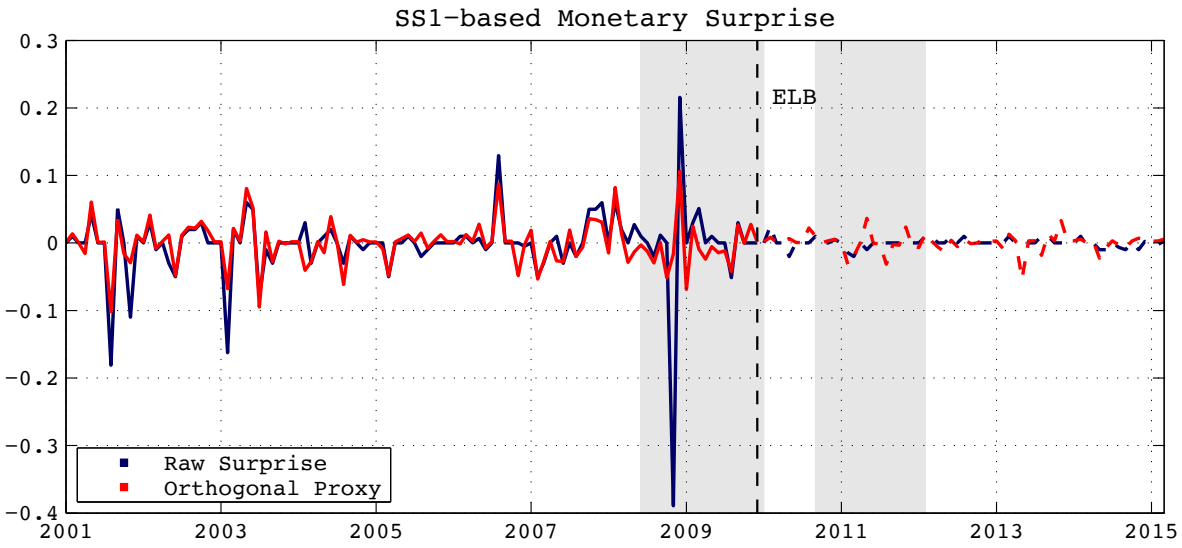


FIGURE 5: Raw (SS1 – blue line) and orthogonal (SS1* – red line) monetary surprises for the UK at monthly frequency. Both sets of surprises are extracted from the first short sterling future. Shaded areas denote Economic Cycle Research Institute (ECRI) recessions. The vertical dotted line denotes start of the Effective Lower Bound (ELB) zone. See main text for details.

forecasts are released at quarterly frequency and with no significant lag, and thus their timely availability is not a concern, we end the benchmark sample for the identification in 2009 to avoid introducing potential distortions caused by Bank Rate reaching its effective lower bound (ELB). The orthogonal surprise calculated over the benchmark sample only is plotted in Appendix C. The start date for the construction of the orthogonal surprise is instead constrained by the availability of the LIBOR-OIS spread. It is worth noticing that the largest peak in the raw surprise disappears in the orthogonal series, in support to the claim that not all price movements contemporaneous to policy announcements are necessarily a reaction to monetary policy shocks only. In fact, the peak coincides with the sharp forecast revisions to growth and unemployment at the onset of the 2009 recession and the sudden increase in the LIBOR-OIS spread that occurred in late 2008, and that was signalling increased fears of insolvency and concerns related to credit availability which had arguably little to do with the monetary policy decision.

6.2 Identification of Monetary Policy Shocks

In the remainder of this section we illustrate the implications of the orthogonalization proposed above for the identification of monetary policy shocks in small, potentially informationally insufficient VARs.

US We test the implications for monetary shock identification using the FF4 and FF4* series as external instruments in a Proxy SVAR where the monetary policy variable is the end-of-month 1-year government bond rate. The identification is borrowed from Gertler and Karadi (2015) and is intended to capture both conventional and unconventional monetary policy that were likely to affect interest rates at medium maturities during the zero lower bound period. Other endogenous variables include the log of industrial production, unemployment rate, the log of CPI and the CRB commodity price index. All variables are taken from the St. Louis FRED Database, with the exception of the commodity price index, distributed by the Commodity Research Bureau. The composition of the set is the same as in Coibion (2012) and Ramey (2016), and it is intentionally kept small to let the differences between the different identifications stand out. For the sake of completeness and comparability with results in these papers, impulse response functions (IRFs) to a monetary policy shock identified using a recursive Cholesky scheme with the effective federal funds rate (EFFR) replacing the 1-year rate and ordered last are also reported. The VAR is estimated in levels with 12 lags over the period 1969:1 - 2014:12. The identification of the contemporaneous transmission coefficients uses the full length of the orthogonal FF4*, that is 1990:1 - 2009:12. Responses are normalized such that the policy rate increases on impact by 1%. Results are in Figure 6. Light blue lines are for the recursive identification scheme with the EFFR ordered last (CHOL). Dark blue lines are obtained when the shock is identified using the FF4-based surprise (PSVAR) of Gertler and Karadi (2015); these are the IRFs plotted in Figure 3. Red lines are responses obtained when the orthogonal FF4* surprise series is used instead – PSVAR*. 90% bootstrapped confidence bands are obtained with 10,000 replications for the PSVAR* case; the wild bootstrap of Gonçalves and Kilian (2004) is used.

Differences between the three identifications are stark. IRFs from both CHOL and

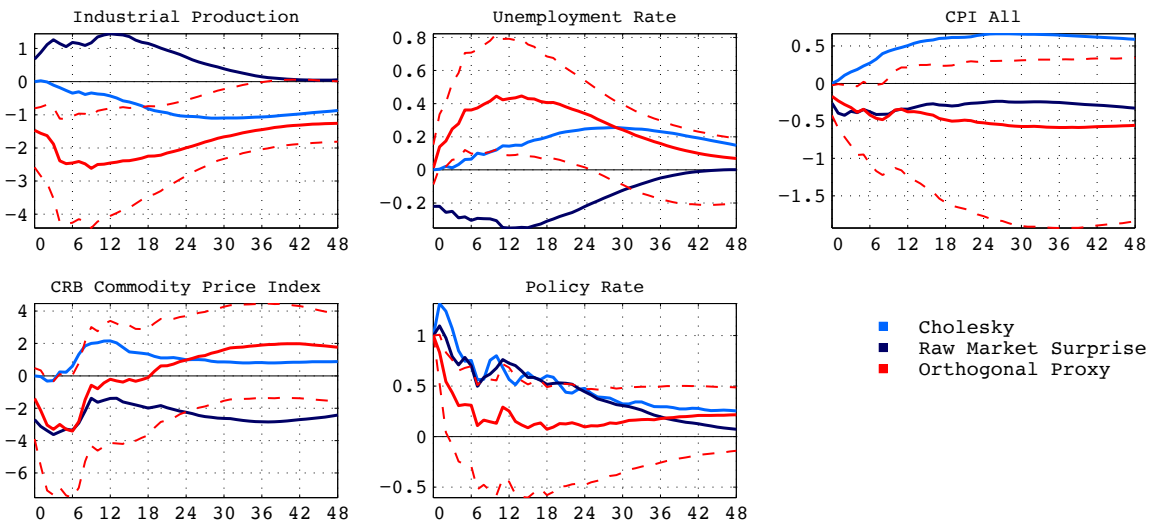


FIGURE 6: The chart compares impulse responses to a monetary policy shock obtained estimating a VAR(12) over the sample 1969:1 - 2014:12 and using different identification schemes. Light blue lines are for the recursive identification scheme with the effective fed funds rate ordered last (CHOL). Dark blue lines are obtained when the shock is identified using the raw FF4-based surprise in a Proxy SVAR with the 1-year rate as the monetary policy variable (PSVAR). Red lines are responses obtained when the conditional, orthogonal surprises are used instead – PSVAR*. Red dotted lines limit 90% bootstrapped confidence bands obtained with 10,000 replications for the PSVAR* case. All shocks are normalised to induce a 1% increase in the policy rate. See main text for details.

PSVAR lie outside the confidence bands of PSVAR* in almost all cases, and particularly so for the nearer horizons. The issues highlighted for the raw FF4 measure, coupled with a small, presumably informationally deficient VAR, deliver distorted and counterintuitive responses for both industrial output and unemployment. Gertler and Karadi (2015) use the raw weighted FF4 measure to identify effects of the monetary policy shock in a similarly small VAR where, however, they also include the excess bond premium (EBP) of Gilchrist and Zakrajšek (2012). Other than being a good predictor of real activity, the EBP is constructed using micro-level data on corporate spreads with average maturity of about 7 years. The long maturity of spreads involved in the calculation of the EBP is likely to be at least partially capturing also forecasts about future realizations that 'clean' the VAR residuals and thus still deliver responses of the expected sign.[23] On the

[23] As noted, successful identification of the shocks in a Proxy SVAR depends both on the quality of the proxy and on the correct specification of the VAR. The importance of the inclusion of the Excess Bond Premium for the identification of the monetary policy shock in otherwise informationally deficient VARs is also discussed in Caldara and Herbst (2015). The authors find that monetary policy shocks are

other hand, PSVAR* responses are less reliant on the composition of the information set in the VAR. Although necessarily less precise, PSVAR* responses are robust to sample splits as shown in Figures C.1a and C.1b in Appendix C.

UK To stress the importance of using orthogonal surprises, we again rely on a small-scale monetary VAR where the raw SS1 and the orthogonal SS1* are used as external instruments, and the monetary policy variable is the end-of-month 1-year government bond rate. Other endogenous variables are the log of industrial production, the LFS (Labour Force Survey) unemployment rate and the log of the retail price index (RPI). Data for Bank Rate and the 1-year government bond rate are from the Bank of England; prices, output and unemployment data are from the Office of National Statistics. The VAR is estimated in levels with 12 lags over the period 1979:1 to 2014:12; responses are again normalized such that the policy rate increases by 1% on impact. The identification of the contemporaneous transmission coefficients uses the orthogonal SS1* in the pre-ELB sample, that is 2001:1 - 2009:12. Responses obtained using the orthogonal SS1* extended to include the ELB period are essentially unaltered, and reported in Appendix C.[24]

Responses to a monetary policy shock in the UK are in Figure 7. As before, light blue lines are for the recursive identification scheme where Bank Rate is ordered last (CHOL). Dark blue lines are obtained when the shock is identified using the raw SS1-based surprise (PSVAR). Red lines are responses obtained when the conditional, orthogonal SS1* surprise series is used instead – PSVAR*. Responses in Figure 7 confirm the extent to which the estimates of the contemporaneous transmission coefficients can be distorted when raw surprises are used to proxy for the monetary policy shock. Again, CHOL and PSVAR responses lie outside the PSVAR* confidence bands throughout most of the horizons, and particularly so on impact. Moreover, as was the case for the US, the spurious information included in the raw SS1 produces responses for output, unemployment and prices that are hard, if not impossible, to reconcile with economic theory. The responses in Figure

important drivers of the EBP at business cycle frequencies and that once these shocks are accounted for, exogenous credit shocks explain a smaller portion of the residual forecast error variance of the EBP and industrial production.

[24] A further backward extension to June 1997 (not reported) is obtained by assuming that the LIBOR-OIS spread is constant and equal to its pre-crisis average over the period 1997:6 - 2000:12. IRFs in this case are qualitatively the same, but estimated with significantly greater uncertainty.

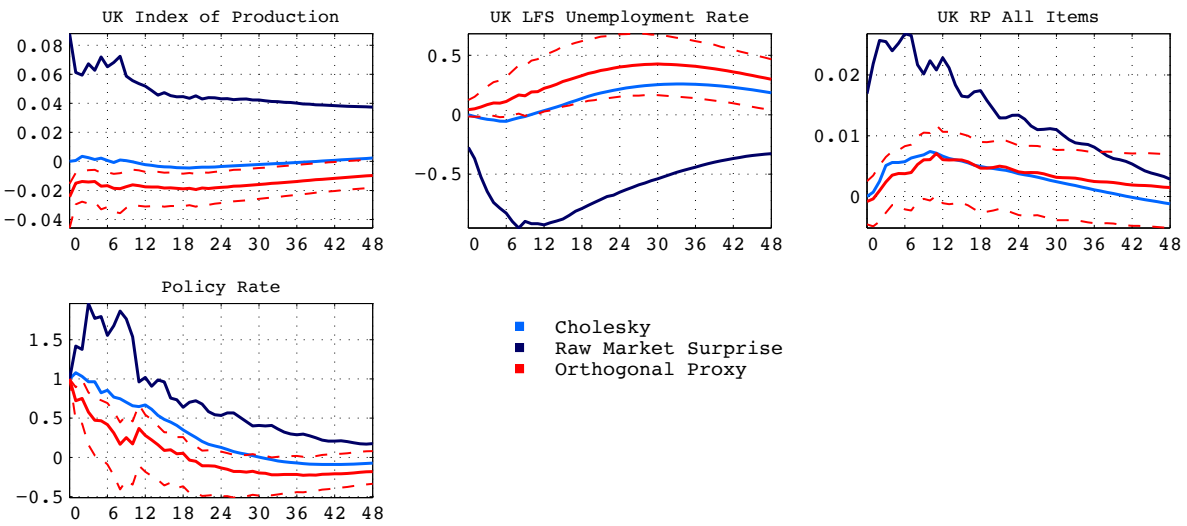


FIGURE 7: The chart compares impulse responses to a monetary policy shock obtained estimating a VAR(12) over the sample 1979:1 - 2014:12 and using different identification schemes. Light blue lines are for the recursive identification scheme with Bank Rate ordered last (CHOL). Dark blue lines are obtained when the shock is identified using the raw SS1-based surprise in a Proxy SVAR with the 1-year rate as the monetary policy variable (PSVAR). Red lines are responses obtained when the conditional, orthogonal surprises are used instead – PSVAR*. Red dotted lines limit 90% bootstrapped confidence bands obtained with 10,000 replications for the PSVAR* case. All shocks are normalised to induce a 1% increase in the policy rate. See main text for details.

C.2a, obtained when the RPI is replaced with the consumer price index and the VAR is estimated from 1990:1 to 2014:12, show that again the identification is robust to sample splits, and the composition of the VAR information set.

7 Concluding Remarks

Recent advances in the identification of monetary policy shocks have proposed the use of futures-based monetary surprises as external instruments in Proxy SVARs to back out the contemporaneous transmission coefficients that link the structural shock of interest to the reduced-form VAR innovations. The rationale for their use is that, to the extent that interest rate futures provide accurate market-based expectations of future policy rates, if the surprises are computed within a sufficiently narrow window around the announcement, the price update around these times can be regarded as a measure, with error, of the underlying monetary policy shock.

Two crucial assumptions are needed to make the futures-based surprises valid external instruments for the monetary policy shocks: (*i*) market prices efficiently incorporate all the relevant available information as it comes along and it takes longer than the measurement window for the monetary policy shock to modify the risk premium in these contracts; and (*ii*) the set of economic forecasts on which the central bank builds its decision, and those of market participants coincide, leading to the equivalence between price updates and monetary policy shocks.

This paper produces evidence that challenges both these assumptions and argues that under general conditions, and independent from the length of the measurement window, monetary surprises capture more than just the monetary policy shock. Because private sector forecasts are not bound to be, and are generally not equal to central banks' forecasts, what markets label as unexpected may or may not be unanticipated by the central bank. In other words, monetary surprises can incorporate anticipatory effects. The dependence of monetary surprises on forecasts of macroeconomic fundamentals also makes the risk premium change within the measurement window. This casts doubts on the exogeneity of the monetary surprises as external instruments for the monetary policy shock. As a result, raw monthly monetary 'surprises' are predictable by both private central bank's forecasts and data whose release predates the announcement.

Building on the predictability of monetary surprises, this paper proposes a new set of proxies that are orthogonal to central banks' forecasts and forecasts revisions of the key variables that are likely to enter the central bank's reaction function. The resulting market-based proxies, free of anticipatory effects by construction, are shown to be uncorrelated with summary measures of information available to the public. The orthogonal proxies proposed in this paper are thus better candidates for the task of capturing only the unanticipated monetary policy decisions.

Results for both the US and the UK show that while raw monetary surprises fail to correctly identify the shocks in standard monetary VARs, and produce counterintuitive responses for critical variables such as output and prices, impulse responses to a monetary policy shock identified using the orthogonal proxies are shown to be in line with economic theory and less reliant on the composition of the VAR information set and the sample

considered even in small, potentially informationally insufficient VARs.

References

Altavilla, Carlo, Domenico Giannone, and Michele Modugno (2014) "Low Frequency Effects of Macroeconomic News on Government Bond Yields," Working Papers ECARES ECARES 2014-34, ULB – Universite Libre de Bruxelles.

Andrade, Philippe and Filippo Ferroni (2016) "Were the ECB announcements Odyssean or Delphic?" mimeo, Banque de France.

Bagliano, Fabio C. and Carlo A. Favero (1999) "Information from financial markets and VAR measures of monetary policy," *European Economic Review*, Vol. 43, No. 4-6, pp. 825–837, April.

Barakchian, S. Mahdi and Christopher Crowe (2013) "Monetary policy matters: Evidence from new shocks data," *Journal of Monetary Economics*, Vol. 60, No. 8, pp. 950–966.

Bowden, Roger J. and Darrell A. Turkington (1984) *Instrumental Variables*, Econometrics Society Monographs in Quantitative Economics: Cambridge University Press.

Caldara, Dario and Edward Herbst (2015) "Monetary Policy, Credit Spreads, and Business Cycle Fluctuations," mimeo, Board of Governors of the Federal Reserve System (U.S.).

Campbell, John Y. and John Cochrane (1999) "Force of Habit: A Consumption-Based Explanation of Aggregate Stock Market Behavior," *Journal of Political Economy*, Vol. 107, No. 2, pp. 205–251, April.

Campbell, Jeffrey R., Charles L. Evans, Jonas D.M. Fisher, and Alejandro Justiniano (2012) "Macroeconomic Effects of Federal Reserve Forward Guidance," *Brookings Papers on Economic Activity*, Vol. 44, No. 1 (Spring), pp. 1–80.

Campbell, Jeffrey R., Jonas D. M. Fisher, Alejandro Justiniano, and Leonardo Melosi (2016) "Forward Guidance and Macroeconomic Outcomes Since the Financial Crisis," in *NBER Macroeconomics Annual 2016, Volume 31*: National Bureau of Economic Research, Inc.

Cesa-Bianchi, Ambrogio, Gregory Thwaites, and Alejandro Vicondoa (2016) "Monetary Policy Transmission in an Open Economy: New Data and Evidence from the United Kingdom," Staff Working Paper 615, Bank of England.

Cloyne, James and Patrick Hürtgen (2014) "The macroeconomic effects of monetary policy: a new measure for the United Kingdom," Bank of England working papers 493, Bank of England.

Cochrane, John H. and Monika Piazzesi (2002) "The Fed and Interest Rates: A High-Frequency Identification," NBER Working Papers 8839, National Bureau of Economic Research, Inc.

Coibion, Olivier (2012) "Are the Effects of Monetary Policy Shocks Big or Small?," *American Economic Journal: Macroeconomics*, Vol. 4, No. 2, pp. 1–32, April.

Cook, Timothy and Thomas Hahn (1989) "The effect of changes in the federal funds rate target on market interest rates in the 1970s," *Journal of Monetary Economics*, Vol. 24, No. 3, pp. 331 – 351.

De Paoli, Bianca and Pawel Zabczyk (2012) "Why Do Risk Premia Vary Over Time? A Theoretical Investigation Under Habit Formation," *Macroeconomic Dynamics*, Vol. 16, No. S2, pp. 252–266, September.

Fama, Eugene F (1990) " Stock Returns, Expected Returns, and Real Activity," *Journal of Finance*, Vol. 45, No. 4, pp. 1089–1108, September.

Fama, Eugene F. (2013) "Two Pillars of Asset Pricing," December. Nobel Prize Lecture.

Fama, Eugene F. and Kenneth R. French (1989) "Business conditions and expected returns on stocks and bonds," *Journal of Financial Economics*, Vol. 25, No. 1, pp. 23 – 49.

Forni, Mario, Marc Hallin, Marco Lippi, and Lucrezia Reichlin (2000) "The Generalized Dynamic-Factor Model: Identification and Estimation," *Review of Economics and Statistics*, Vol. 82, No. 4, pp. 540–554.

Galí, Jordi (2008) *Monetary Policy, Inflation, and the Business Cycle: An Introduction to the New Keynesian Framework*: Princeton University Press.

Gertler, Mark and Peter Karadi (2015) "Monetary Policy Surprises, Credit Costs, and Economic Activity," *American Economic Journal: Macroeconomics*, Vol. 7, No. 1, pp. 44–76.

Gilchrist, Simon and Egon Zakrajšek (2012) "Credit Spreads and Business Cycle Fluctuations," *American Economic Review*, Vol. 102, No. 4, pp. 1692–1720.

Gleser, Leon Jay (1992) "The Importance of Assessing Measurement Reliability in Multivariate Regression," *Journal of the American Statistical Association*, Vol. 87, No. 419, pp. pp. 696–707.

Glick, Reuven and Sylvain Leduc (2015) "Unconventional monetary policy and the dollar: conventional signs, unconventional magnitudes," Working Paper Series 2015-18, Federal Reserve Bank of San Francisco.

Gonçalves, Silvia and Lutz Kilian (2004) "Bootstrapping autoregressions with conditional heteroskedasticity of unknown form," *Journal of Econometrics*, Vol. 123, No. 1, pp. 89–120, November.

Gürkaynak, Refet S (2005) "Using federal funds futures contracts for monetary policy analysis," Finance and Economics Discussion Series 2005-29, Board of Governors of the Federal Reserve System (U.S.).

Gürkaynak, Refet S, Brian Sack, and Eric T. Swanson (2005) "Do Actions Speak Louder Than Words? The Response of Asset Prices to Monetary Policy Actions and Statements," *International Journal of Central Banking*, Vol. 1, No. 1, May.

——— (2006) "Market-based Measures of Monetary Policy Expectations," Working Paper Series 2006-04, Federal Reserve Bank of San Francisco.

Hanson, Samuel G. and Jeremy C. Stein (2015) "Monetary policy and long-term real rates," *Journal of Financial Economics*, Vol. 115, No. 3, pp. 429–448.

Joyce, Michael, Jonathan Relleen, and Steffen Sorensen (2008) "Measuring Monetary Policy Expectations from Financial Market Instruments," Working Paper Series 0978, European Central Bank.

Kuttner, Kenneth N. (2001) "Monetary policy surprises and interest rates: Evidence from the Fed funds futures market," *Journal of Monetary Economics*, Vol. 47, No. 3, pp. 523–544, June.

Lucca, David O. and Emanuel Moench (2015) "The Pre-FOMC Announcement Drift," *The Journal of Finance*, Vol. 70, No. 1, pp. 329–371.

McCracken, Michael W. and Serena Ng (2015) "FRED-MD: A Monthly Database for Macroeconomic Research," Working Papers 2015-12, Federal Reserve Bank of St. Louis.

Mertens, Karel and Morten O. Ravn (2013) "The Dynamic Effects of Personal and Corporate Income Tax Changes in the United States," *American Economic Review*, Vol. 103, No. 4, pp. 1212–47.

Miranda-Agrippino, Silvia and Giovanni Ricco (2016) "The Transmission of Monetary Policy Shocks." University of Warwick.

Montiel-Olea, José Luis, James Stock, and Mark W. Watson (2012) "Inference in Structural VARs with External Instruments." Harvard University.

Montiel-Olea, José Luis, James H. Stock, and Mark W. Watson (2016) "Uniform Inference in SVARS Identified with External Instruments." New York University.

Nakamura, Emi and Jón Steinsson (2013) "High Frequency Identification of Monetary Non-Neutrality," NBER Working Papers 19260, National Bureau of Economic Research, Inc.

Piazzesi, Monika (2002) "The Fed and Interest Rates - A High-Frequency Identification," *American Economic Review*, Vol. 92, No. 2, pp. 90–95, May.

Piazzesi, Monika and Eric T. Swanson (2008) "Futures Prices as Risk-Adjusted Forecasts of Monetary Policy," *Journal of Monetary Economics*, Vol. 55, No. 4, pp. 677–691, May.

Ramey, Valerie A. (2016) "Macroeconomic Shocks and Their Propagation," Working Paper 21978, National Bureau of Economic Research.

Rogers, John H., Chiara Scotti, and Jonathan H. Wright (2014) "Evaluating Asset-Market Effects of Unconventional Monetary Policy: A Cross-Country Comparison," International Finance Discussion Papers 1101, Board of Governors of the Federal Reserve System (U.S.).

Romer, Christina D. and David H. Romer (2000) "Federal Reserve Information and the Behavior of Interest Rates," *American Economic Review*, Vol. 90, No. 3, pp. 429–457.

——— (2004) "A New Measure of Monetary Shocks: Derivation and Implications," *American Economic Review*, Vol. 94, No. 4, pp. 1055–1084.

Rudebusch, Glenn D. (1998) "Do Measures of Monetary Policy in a Var Make Sense?" *International Economic Review*, Vol. 39, No. 4, pp. 907–931.

Sack, Brian (2004) "Extracting the Expected Path of Monetary Policy From Futures Rates," *Journal of Futures Markets*, Vol. 24, No. 8, pp. 733–754.

Scotti, Chiara (2013) "Surprise and uncertainty indexes: real-time aggregation of real-activity macro surprises," International Finance Discussion Papers 1093, Board of Governors of the Federal Reserve System (U.S.).

Soderlind, Paul and Lars Svensson (1997) "New techniques to extract market expectations from financial instruments," *Journal of Monetary Economics*, Vol. 40, No. 2, pp. 383–429, October.

Stock, James H. and Mark W. Watson (2002) "Macroeconomic Forecasting Using Diffusion Indexes," *Journal of Business & Economic Statistics*, Vol. 20, No. 2, pp. 147–162.

——— (2012) "Disentangling the Channels of the 2007-09 Recession," *Brookings Papers on Economic Activity*, Vol. 44, No. 1, pp. 81–156.

Svensson, Lars E.O. (1994) "Estimating and Interpreting Forward Interest Rates: Sweden 1992 - 1994," NBER Working Papers 4871, National Bureau of Economic Research, Inc.

Swanson, Eric T. (2015) "Measuring the Effects of Unconventional Monetary Policy on Asset Prices," NBER Working Papers 21816, National Bureau of Economic Research, Inc.

Woodford, Michael (2003) *Interest and Prices: Foundations of a Theory of Monetary Policy*: Princeton University Press.

A Identification with External Instruments

Using the notation introduced in Section 4, let y_t be an n-dimensional vector of endogenous observables whose dynamic is described by the following system of equations:[25]

$$\mathbf{B}^{-1}y_t = \mathcal{A}_1 y_{t-1} + \ldots + \mathcal{A}_p y_{t-p} + e_t, \tag{A.1}$$

where $\mathbf{B}^{-1}$ and $\mathcal{A}_i$, $i = 1, \ldots, p$, are square matrices of structural coefficients and e_t is an n-dimensional vector of structural shocks such that $\mathbb{E}[\, e_t \,] = 0$, $\mathbb{E}[\, e_t e_t' \,] = \mathbb{I}_n$ and $\mathbb{E}[\, e_t e_\tau' \,] = 0$, $\forall \tau \neq t$. Deterministic terms are allowed to enter (A.1) but are omitted in what follows for notational brevity.

The reduced-form version of the SVAR in (A.1) reads

$$A(L)y_t = u_t, \tag{A.2}$$

where $A(L) \equiv [\, \mathbb{I}_n - A_1 L - \ldots - A_p L^p \,]$, $A_i \equiv \mathbf{B}\mathcal{A}_i$, $i = 1, \ldots p$, and the reduced-form VAR innovations are linear combinations of the structural shocks

$$u_t \equiv \mathbf{B}e_t, \tag{A.3}$$

with

$$\mathbb{E}[\, u_t u_t' \,] = \mathbf{B}\mathbf{B}' = \Sigma_u. \tag{A.4}$$

If $A(L)$ is invertible, y_t can be expressed as an infinite sum of present and past realizations of the structural shocks

$$y_t = [\, A(L) \,]^{-1} u_t = \mathcal{C}(L)\mathbf{B}e_t, \tag{13}$$

where $\mathcal{C}(L)\mathbf{B}$ are the structural impulse response functions. While the coefficients in $\mathcal{C}(L)$ are easily estimated as a function of the reduced-form autoregressive parameters, recovering the elements of $\mathbf{B}$ typically requires imposing a set of identifying restrictions such that identification can be achieved. A prime example entails assuming that $\mathbf{B}$ is lower triangular and equal to the Cholesky factor ofΣ_u; the resulting $n(n-1)/2$ contemporaneous restrictions grant exact identification of the system in (A.4).

Within the Proxy SVAR framework, on the other hand, the relevant columns of the $\mathbf{B}$ matrix are identified using an external instrument (or *proxy*), not included in the VAR,

[25] The content of this Appendix draws heavily from Montiel-Olea, Stock and Watson (2012); Mertens and Ravn (2013).

that can be thought of as a measure – possibly with error – of the structural shock (Stock and Watson, 2012; Mertens and Ravn, 2013). Without loss of generality, suppose that the shock of interest – call it the monetary policy shock, $e_t^\bullet$ – is ordered first in the vector e_t, such that $\mathbf{B}$ can be partitioned as follows:

$$u_t = \mathbf{B}e_t, \tag{A.5}$$

$$\begin{matrix}[1\times 1]\\ [(n-1)\times 1]\end{matrix} \left(\begin{array}{c} u_t^\bullet \\ \hdashline u_t^\circ \end{array}\right) = \left(\begin{array}{c:c} \mathbf{b}^\bullet & \mathbf{b}^\circ \end{array}\right)\left(\begin{array}{c} e_t^\bullet \\ \hdashline e_t^\circ \end{array}\right), \tag{A.6}$$

where $\mathbf{b}^\bullet$ denotes the first column vector of $\mathbf{B}$, $\mathbf{b}^\circ$ is of dimension $[n \times (n-1)]$ and e_t° collects the remaining shocks.[26]

Suppose there exists a set of variables z_t, not in y_t, such that:

$$\mathbb{E}[\, e_t^\bullet z_t' \,] = \phi', \tag{14}$$

$$\mathbb{E}[\, e_t^\circ z_t' \,] = 0, \tag{15}$$

where ϕ is non-singular. If a variable z_t can be found such that the validity conditions in (14) and (15) are satisfied, then it is possible to identify $\mathbf{b}^\bullet$ up to scale and sign:

$$\mathbb{E}[\, u_t z_t' \,] = \mathbb{E}[\, \mathbf{B}e_t z_t' \,] = \left(\begin{array}{cc} \mathbf{b}^\bullet & \mathbf{b}^\circ \end{array}\right)\left(\begin{array}{c} \mathbb{E}[\, e_t^\bullet z_t' \,] \\ \mathbb{E}[\, e_t^\circ z_t' \,] \end{array}\right) = \mathbf{b}^\bullet \phi', \tag{A.7}$$

implying that further normalization is needed to back out the elements in $\mathbf{b}^\bullet$.

Montiel-Olea, Stock and Watson (2012) assume that a unit positive increase in the shock induces a unit positive increase in the first variable; this translates into setting the first element of $\mathbf{b}^\bullet$ equal to 1. With $\mathbf{b}^\bullet = (\,1,\ \mathrm{b}^{\circ\prime}\,)'$, and using the relation established in (A.7)

$$\left(\begin{array}{c} \mathrm{b}^\circ \mathbb{E}[\, u_t^\bullet z_t' \,] \\ \mathbb{E}[u_t^\circ z_t' \,] \end{array}\right) = \left(\begin{array}{c} \mathrm{b}^\circ \phi' \\ \mathrm{b}^\circ \phi' \end{array}\right),$$

which, rearranging terms, is equivalent to writing

$$\mathrm{b}^\circ\, \mathbb{E}[\, u_t^\bullet z_t' \,] = \mathbb{E}[\, u_t^\circ z_t' \,]. \tag{A.8}$$

[26] $\mathbf{B} \equiv \left(\begin{array}{c:c} \mathbf{b}^\bullet & \mathbf{b}^\circ \end{array}\right)$, where $\mathbf{b}^\bullet$ is the column vector containing the coefficients that link the reduced-form residuals to $e_t^\bullet$. $\mathbf{b}^\bullet$ is further partitioned such that $\mathbf{B} = \left(\begin{array}{c:c} \mathrm{b}^\bullet & \\ \mathrm{b}^\circ & \mathbf{b}^\circ \end{array}\right)$, where $\mathrm{b}^\bullet$ is the coefficient that links the first entry in u_t to $e_t^\bullet$.

Equation (A.8) establishes that, given the normalization discussed above, the elements of $\mathbf{b}^{\bullet}$ can be estimated using moments of observables; in particular, if z_t only contains one proxy variable, $\mathrm{b}^{\circ} = \mathbb{E}[\, u_t^{\circ} z_t' \,]/\mathbb{E}[\, u_t^{\bullet} z_t' \,]$, that is, it is equal to the ratio between the coefficients of the regression of the reduced-form VAR innovations onto the instrument.[27]

A.1 The contemporaneous transmission coefficients in the EIV framework

Let the true model be:

$$y_t = \mathbf{A}^{\star} \mathcal{Y}_t^{\star} + w_t, \tag{A.9}$$

where $\mathbf{A}^{\star} \equiv (\, \mathbf{A} \;\; \mathbf{b}^{\bullet} \,)$, $\mathbf{A} \equiv (\, A_1, \ldots, A_p \,)$. $\mathcal{Y}_t^{\star} \equiv (\, \mathcal{Y}_t', \;\; e_t^{\bullet\prime} \,)'$, where $\mathcal{Y}_t \equiv [y_{t-1}', \ldots, y_{t-p}']'$ is only partially observable, as it contains the latent structural shock of interest – $e_t^{\bullet}$. The relevant contemporaneous transmission coefficients are collected in the column vector $\mathbf{b}^{\bullet}$.

Given a proxy z_t for $e_t^{\bullet}$ such that

$$z_t = \Phi e_t^{\bullet} + \nu_t, \tag{16}$$

where ν_t is an i.i.d. measurement error with $\mathbb{E}[\, \nu_t \,] = 0$, $\mathbb{E}[\, \nu_t \nu_t' \,] = \Sigma_{\nu}$, and $\mathbb{E}[\, \nu_t \nu_{\tau}' \,] = 0$, $\forall \tau \neq t$ and Φ is non-singular, the researcher estimates

$$y_t = \mathbf{C} \mathcal{Y}_t^{+} + \eta_t, \tag{A.10}$$

where

$$\mathcal{Y}_t^{+} \equiv (\, \mathcal{Y}_t', \;\; z_t' \,)' = \Psi \mathcal{Y}_t^{\star} + \varsigma_t. \tag{A.11}$$

[27] An alternative formulation is discussed in Mertens and Ravn (2013), where b° in (A.8) is replaced with $\tilde{\mathrm{b}}^{\circ} \equiv [\, \mathrm{b}^{\bullet} \,]^{-1} \mathrm{b}^{\circ}$ and thus the ratio between the coefficients of the regressions of the VAR innovations onto the instrument delivers a scaled version of $\mathbf{b}^{\bullet}$. The unscaled $\mathbf{b}^{\bullet}$ is then recovered by noting that:

$$\mathrm{b}^{\bullet} = \sqrt{\Sigma_{11} - (\Sigma_{n1} - \tilde{\mathrm{b}}^{\circ} \Sigma_{11})' \Gamma^{-1} (\Sigma_{n1} - \tilde{\mathrm{b}}^{\circ} \Sigma_{11})},$$

where $\Gamma = \tilde{\mathrm{b}}^{\circ} \Sigma_{11} \tilde{\mathrm{b}}^{\circ\prime} - (\Sigma_{n1} \tilde{\mathrm{b}}^{\circ} + \tilde{\mathrm{b}}^{\circ} \Sigma_{n1}') + \Sigma_{11}$ and Σ_{ij} are appropriate partitions of Σ_u

$$\mathbb{E}[\, u_t u_t' \,] = \Sigma_u = \left(\begin{array}{c|c} \underset{[1\times 1]}{\Sigma_{11}} & \underset{[1\times(n-1)]}{\Sigma_{1n}} \\ \hline \underset{[(n-1)\times 1]}{\Sigma_{n1}} & \underset{[(n-1)\times(n-1)]}{\Sigma_{nn}} \end{array} \right).$$

Because $\mathcal{Y}_t^\star$ is measured with error, the OLS estimate of $\mathbf{C}$ is biased. In particular, if $\widehat{\mathbf{C}}$ denotes the least squares estimates of $\mathbf{C}$, and η_t and ς_t are normally distributed, $\widehat{\mathbf{C}} = \mathbf{C}\Lambda$, where

$$\Lambda = [\,\Sigma_{\mathcal{Y}^+}\,]^{-1}[\,\Sigma_{\mathcal{Y}^+} - \Sigma_\varsigma\,] \tag{A.12}$$

is the reliability matrix of $\mathcal{Y}_t^+$ (Bowden and Turkington, 1984; Gleser, 1992).Σ_x denotes $\mathbb{E}[x_t x_t']$ for any x_t.

The coefficients in $\mathbf{A}^\star$, and thus $\mathbf{b}^\bullet$, can be recovered using $\mathbf{A}^\star = \widehat{\mathbf{C}}\Lambda^{-1}\Psi$. A necessary condition for this procedure to deliver the coefficients in $\mathbf{b}^\bullet$, is that the proxy z_t be orthogonal to the history of y_t included in the VAR, that is, $\mathbb{E}[\,z_t\mathcal{Y}_t'\,] = 0$.

Using (A.12), and OLS estimates of $\mathbf{C}$ from (A.10)

$$\begin{aligned}\mathbf{A}^{\star\prime} = (\,\mathbf{A}\ \mathbf{b}^\bullet\,)' &= \Psi'\Lambda^{-1}\widehat{\mathbf{C}}' \\ &= \Psi'\left[\,\Sigma_{\mathcal{Y}^+}^{-1}\ [\,\Sigma_{\mathcal{Y}^+} - \Sigma_\varsigma\,]\,\right]^{-1}\Sigma_{\mathcal{Y}^+}^{-1}\Sigma_{\mathcal{Y}^+ y}.\end{aligned} \tag{A.13}$$

If $\mathbb{E}[\,z_t\mathcal{Y}_t'\,] = 0$,

$$\Psi' = \begin{pmatrix} \mathbb{I}_{np} & \mathbf{0} \\ \mathbf{0} & \Phi' \end{pmatrix} \quad \text{and}\Sigma_{\mathcal{Y}^+} = \begin{pmatrix} \Sigma_{\mathcal{Y}} & \mathbf{0} \\ \mathbf{0} & \Sigma_z \end{pmatrix}, \tag{A.14}$$

where $\mathbf{0}$ denotes matrices of zeros of suitable dimensions. Equation (A.11) can thus be rewritten as

$$\mathcal{Y}_t^+ \equiv \begin{pmatrix} \mathcal{Y}_t \\ z_t \end{pmatrix} = \begin{pmatrix} \mathbb{I}_{np} & \mathbf{0} \\ \mathbf{0} & \Phi \end{pmatrix}\begin{pmatrix} \mathcal{Y}_t \\ e_t^\bullet \end{pmatrix} + \begin{pmatrix} \mathbf{0} \\ \nu_t \end{pmatrix}. \tag{A.11$'$}$$

After some algebra, plugging (A.14) into (A.13) yields

$$\mathbf{A}^{\star\prime} = \begin{pmatrix} \mathbb{I}_{np} & \mathbf{0} \\ \mathbf{0} & \Phi' \end{pmatrix}\left[\begin{pmatrix} \Sigma_{\mathcal{Y}}^{-1} & \mathbf{0} \\ \mathbf{0} & \Sigma_z^{-1} \end{pmatrix}\begin{pmatrix} \Sigma_{\mathcal{Y}} & \mathbf{0} \\ \mathbf{0} & \Sigma_z - \Sigma_\nu \end{pmatrix}\right]^{-1}\left[\begin{pmatrix} \Sigma_{\mathcal{Y}}^{-1} & \mathbf{0} \\ \mathbf{0} & \Sigma_z^{-1} \end{pmatrix}\begin{pmatrix} \Sigma_{\mathcal{Y}y} \\ \Sigma_{zy} \end{pmatrix}\right]. \tag{A.15}$$

Due to the block diagonal structure of the elements in (A.15), the components of $\mathbf{A}^\star$ can be solved for separately. It is easily seen that the first np equations deliver the least squares estimates of the VAR autoregressive coefficients, that is, the elements in

$(A_1, \ldots, A_p)'$. The remaining conditions produce the parameters of interest:

$$\begin{aligned}\mathbf{b}^{\bullet\prime} &= \Phi' \left[\Sigma_z^{-1} \left[\Sigma_z - \Sigma_\nu \right] \right]^{-1} \Sigma_z^{-1} \Sigma_{zy} \\ &= \Phi' \left[\Sigma_z - \Sigma_\nu \right]^{-1} \Sigma_{zy} \\ &= \Phi' \left[\Phi \Phi' \right]^{-1} \Sigma_{zy} = \Phi^{-1} \Sigma_{zy}, \end{aligned} \tag{A.16}$$

which is equivalent to (A.7).

B Monetary Policy Surprises from Futures on Interest Rates

B.1 US Raw Monetary Surprises

Sack (2004) discusses the technical procedure for the extraction of policy expectations from both Federal Funds (FF) and Eurodollar (ED) futures that are shown to be accurate predictors of the policy rate in Gürkaynak, Sack and Swanson (2006). Let $ff^{(h)}$ and $ed^{(h)}$ denote respectively the price of the FF and ED expiring on day h of a given month with N days then:

$$ff^{(h)} = 100 - \frac{1}{N}\sum_{j=1}^{N} i_j, \tag{B.1}$$

$$ed^{(h)} = 100 - \$\text{lib}_h^{(h+90)}, \tag{B.2}$$

where i_t is the effective fed fund rate and $\$\text{lib}_h^{(h+90)}$ is the 3-month US dollar LIBOR fixing on day h. When expressed in rates at any time t, the equations above transform as follows:

$$p_{t,FF}^{(h)} = \mathbb{E}_t\left(\frac{1}{N}\sum_{j=1}^{N} i_j\right) + \zeta_{FF,t}^{(h)}, \tag{B.3}$$

$$\begin{aligned} p_{t,ED}^{(h)} &= \mathbb{E}_t\left[\ \$\text{lib}_h^{(h+90)}\ \right] + \zeta_{ED,t}^{(h)} \\ &= \mathbb{E}_t\left[\ \bar{\mathbf{i}}_h^{h+90}\ \right] + \mathbb{E}_t\left[\ \$\text{lib}_h^{(h+90)} - \bar{\mathbf{i}}_h^{h+90}\ \right] + \zeta_{ED,t}^{(h)}. \end{aligned} \tag{B.4}$$

$\bar{\mathbf{i}}_h^{h+90}$ denotes the average rates over the 90 days (3 months) starting from day h, i.e. $\bar{\mathbf{i}}_h^{h+90} \equiv \frac{1}{90}\sum_{j=1}^{90} i_{h+j}$. While the link between FF and i_t is direct, when dealing with EDs an additional step in which expectations about future LIBOR fixings are translated into expectations about the policy rate is required. The terms $\zeta_{.,t}^{(h)}$ denote (possibly time-varying) term/risk premia in both equations. In (B.4), the ED rate is expressed as a function of three terms: (a) the expectation of the short-term rate over the three-month period starting from the expiration of the contract – h; (b) a term reflecting 'basis risk', that is, the compensation that investors require for lending to an institution over a 3-month period rather than on an overnight basis; and (c) a residual risk premium which encompasses everything which is not explicitly associated to either (a) or (b).

Kuttner (2001) constructs monetary surprises in the US using daily data on federal funds futures expiring in the current month. Gürkaynak (2005) and Gürkaynak et al.

(2005) use futures covering maturities which go out about 3.5 quarters and intraday quotes. Federal funds futures settle based on the average effective federal funds rate (EFFR) calculated over the relevant expiry month, therefore, if $ff^{(0)}_{t-\Delta t}$ denotes the current month futures just before $(-\Delta t)$ the FOMC meeting, and i_t is the EFFR:

$$ff^{(0)}_{t-\Delta t} = \frac{n}{N}\mathbb{E}_{t-\Delta t}[\, i_{\tau \le t}\,] + \frac{N-n}{N}\mathbb{E}_{t-\Delta t}[\, i_{\tau \ge t}\,] + \zeta^{(0)}_{FF,t-\Delta t}. \tag{B.5}$$

In the equation above, N is the number of days in the month and n is the day of the FOMC meeting, t the time of the announcement, and $\zeta^{(0)}_{FF,t-\Delta t}$ a risk or term premium that may be present in the contract. The scaling is such that it avoids overweighting when the FOMC meets at the end of the month by using the next month's contract if certain timing criteria are met (see Gürkaynak, 2005). The monetary policy surprise – $mps^{(0)}_t$ – can be computed as:

$$\begin{aligned} mps^{(0)}_t &= \frac{N}{N-n}\left[\, ff^{(0)}_t - ff^{(0)}_{t-\Delta t}\,\right] \\ &= \left[\mathbb{E}_t[\, i_{\tau \ge t}\,] - \mathbb{E}_{t-\Delta t}[\, i_{\tau \ge t}]\right] + \left[\, \zeta^{(0)}_{FF,t} - \zeta^{(0)}_{FF,t-\Delta t}\,\right]. \end{aligned} \tag{B.6}$$

Gürkaynak et al. (2005) assume that the latter term in the equation above is zero, de facto implying that it takes longer than theΔ t time frame for the announcement to modify the premium. The surprises that relate to announcements further ahead in the future are derived in an equivalent way using futures that refer to the month in which the relevant FOMC announcement is scheduled to happen.

The raw monetary surprise extracted from the fourth Fed Fund future (FF4) and aggregated at monthly frequency is plotted in Figure B.1. The top panel of the chart reports the monthly average surprise in Gertler and Karadi (2015) (blue line) and the raw series that assigns each daily surprise in Gürkaynak, Sack and Swanson (2005) to the month in which the corresponding meeting was scheduled to happen (red line).[28] The bottom row of the chart reports (from left to right) the scatter plot of the two monthly measures and the partial autocorrelation function of the weighted and unweighted monthly surprises respectively. The weighted series exhibits some degree of autocorrelation, also noted in Ramey (2016). The weighting procedure can be summarised in two steps: (1) for each day of the month, the surprise is equal to the sum of surprises in FOMC days within the past month; (2) for each month, the surprise is equal to the average of the daily series in the previous step. The procedure induces a significant time-dependence in the

[28]The procedure follows Romer and Romer (2004); if there is more than one FOMC meeting in the same month, the monthly surprise is equal to the sum of the surprises registered in that month.

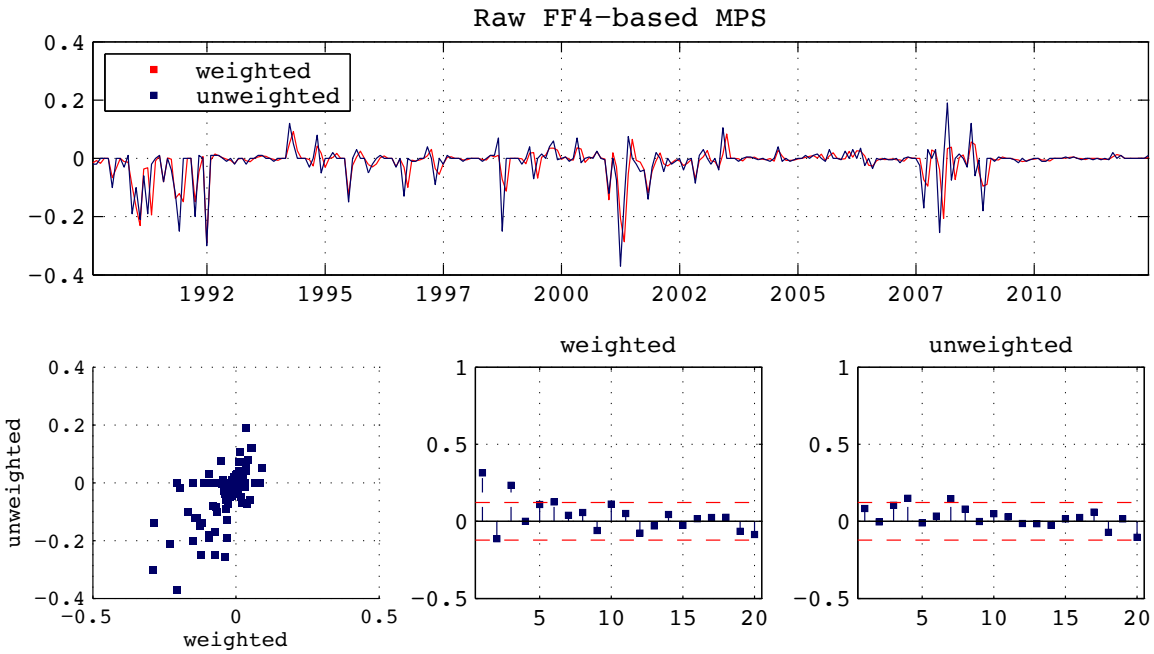


FIGURE B.1: Raw FF4-based monetary surprises at monthly frequency. The weighted series is from Gertler and Karadi (2015), while the unweighted surprise is constructed as the sum of daily surprises in Gürkaynak et al. (2005). In the bottom panel, from left to right, the different information content in the two series and their partial autocorrelation functions.

monthly series. To see this, note that the autocorrelation is only marginally significant when monthly surprises are just the sum of daily movements (unweighted series). A more serious concern, however, is in the alignment of the two series, visible in the top panel of the chart. The weighting of daily surprises shifts the monthly surprise series forward; this implies that also the alignment with the information set (and thus the residuals) of the VAR is distorted. As a result, we use the unweighted monthly surprises as the basis for our analysis.

B.2 UK Raw Monetary Surprises

The case for the UK differs form the US in some non-trivial ways. The Bank of England implements the Monetary Policy Committee's (MPC) decisions by adjusting the level of Bank Rate, to which no financial market instrument is directly linked. The closest alternative is Overnight Indexed Swap (OIS) contracts. In these contracts, the parties agree to exchange fixed interest rate payments against payments based on the Sterling Overnight Index Average (SONIA); because the level of credit risk in overnight transactions is typically very low, SONIA rates track Bank Rate closely, furthermore, since these contracts are constructed in a way that minimises credit risk, the implied path

of SONIA rates at short horizons should also be relatively free of material risk premia. The contracts, however, are only available for a limited time span and, until the years immediately preceding the global financial crisis, seldom traded at maturities beyond 6 months. The next best alternative is to use short sterling (SS) futures contracts, whose forecasting performance is only slightly inferior to OIS rates.[29] These contracts settle based on the 3-month interbank (GBP) LIBOR rate rather than on overnight rates, but are exchange-traded and available for a much longer history.

Because Eurodollar (ED) futures also settle on the (US dollar) LIBOR rather than on the effective fed funds rate, they are the natural starting point to work out policy expectations in the UK. Building on the decomposition in Sack (2004) – equation (B.4), let $ss^{(h)}$ denote the price of a Short Sterling future expiring on day h, we have that

$$ss^{(h)} = 100 - £\text{lib}_h^{(h+90)}, \tag{B.7}$$

where $£\text{lib}_h^{(h+90)}$ is the 3-month sterling LIBOR fixing on day h. Following the same logic in (B.4), the rate at time t can then be expressed as

$$\begin{aligned} p_{t,ss}^{(h)} &= \mathbb{E}_t\left[£\text{lib}_h^{(h+90)}\right] + \zeta_{SS,t}^{(h)}, \\ &= \mathbb{E}_t\left[\bar{\mathbf{i}}_h^{h+90}\right] + \mathbb{E}_t\left[£\text{lib}_h^{(h+90)} - \bar{\mathbf{i}}_h^{h+90}\right] + \zeta_{SS,t}^{(h)}, \end{aligned} \tag{B.8}$$

where it is assumed that the overnight rate i_t is equivalent to the policy rate up to a negligible additive error. $\bar{\mathbf{i}}_h^{h+90}$ denotes the average overnight rate over the 90 days (3 months) starting from day h, i.e. $\bar{\mathbf{i}}_h^{h+90} \equiv \frac{1}{90}\sum_{j=1}^{90} i_{h+j}$.

The rates involved in (B.8) and a detail on the time variation of the LIBOR-OIS spread are in Figure B.2 for the sample 01/01/2000 - 31/05/2015. The overnight rate is the one that most closely tracks the policy rate over the whole sample considered, LIBORs, on the other hand, typically lie above the policy/overnight rates reflecting the risk involved in lending at further away maturities. While it is now widely regarded as one of the key measures of credit risk premia, the LIBOR-OIS spread – e.g. the second term in Equation (B.8), drew relatively little attention in the years preceding the onset of the 2007 finan-

[29] The quality of market-based policy path forecasts, including those derived from SS contracts, is discussed in Joyce, Relleen and Sorensen (2008). The exercise is similar in spirit to Gürkaynak, Sack and Swanson (2006), but in this case also yield curves are added to the horserace. The two zero-coupon yield curves used in the analysis are the ones estimated and published by the Bank of England; the Government Liability Curve (GLC), derived from UK government bonds ('gilts') and general collateral repo rates, and the Bank Liability Curve (BLC), based instead on LIBOR interest rates, short sterling futures, Forward Rates Agreements and LIBOR-based interest rates swaps. Since yield curves are estimated and published at daily frequency, we discard them from the subsequent analysis.

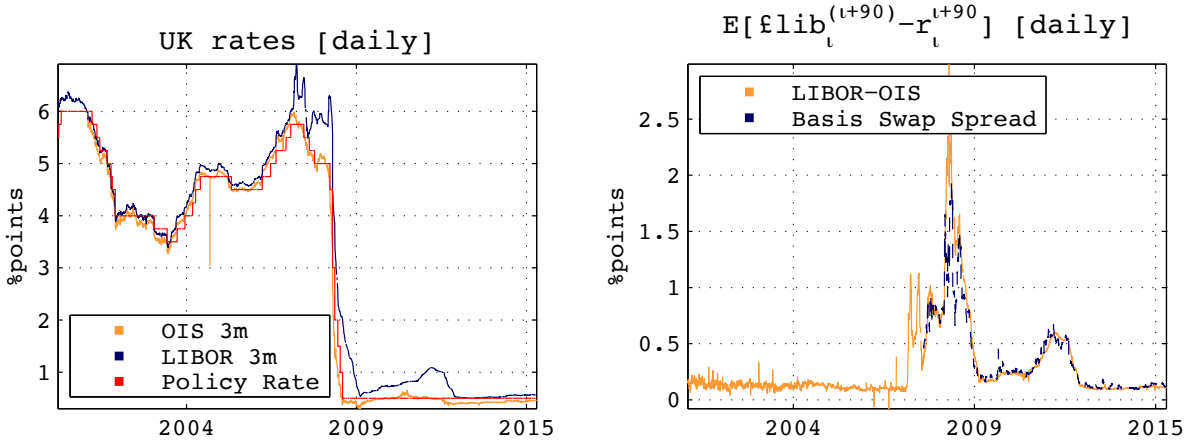


FIGURE B.2: [LEFT] Relevant interest rates for short sterling futures rates decomposition. [RIGHT] LIBOR-OIS spreads obtained as the difference between the 3-month sterling LIBOR and the 3-month OIS curve, and from basis swaps (front contract, basis swap spread). All rates are at daily frequency over the sample 01/01/2000 - 31/05/2015. See equation (B.8) for details. *Source:* Bloomberg, author calculations.

cial crisis: its level remained very low (around 11 basis points) and substantially flat for years, reflecting the belief that the level of credit risk involved in the financial system was not only very small, but also constant over that period. Starting from 2008, however, doubts about financial institutions' solvency and concerns related to market liquidity induced a rise in LIBORs which made the spread jump to unprecedented levels. As the LIBOR-OIS spread moved away from its long-run average, basis swaps involving expected risk at different maturities started being traded and thus, from that date, expectations about future spreads can be read from the swap quotes. In the absence of such contracts, that is, prior to 2008, the actual difference between the 3-month sterling LIBOR and the 3-month OIS curve can be used to compute the expected spread; this is equivalent to setting $h = 0$ in $\mathbb{E}_t\left[£\text{lib}_h^{(h+90)} - \bar{\mathbf{i}}_h^{h+90}\right]$.

Let $p_{t,BS}^{(h)}$ denote the basis swap quotes matching the expectation components in (B.8) at any time t, and let the relevant policy announcement happen within the time interval $[t-\Delta t, t]$, such that Δt denotes the width of the time window around which the response is measured. In the absence of any conflicting event, the raw unconditional monetary policy surprise is thus given by:

$$\begin{aligned} mps_t^{(h)} &= \left[p_{t,SS}^{(h)} - p_{t-\Delta t,SS}^{(h)}\right] - \left[p_{t,BS}^{(h)} - p_{t-\Delta t,BS}^{(h)}\right], \\ &= \left[\mathbb{E}_t\left[\bar{\mathbf{i}}_h^{h+90}\right] - \mathbb{E}_{t-\Delta t}\left[\bar{\mathbf{i}}_h^{h+90}\right]\right] + \left[\zeta_t^{(h)} - \zeta_{t-\Delta t}^{(h)}\right]. \end{aligned} \tag{B.9}$$

Figure B.3 plots the monthly surprises in the first short sterling futures from June 1997 to 2015. The starting date is chosen to coincide with the first decision meeting after the MPC independence. SS delivery dates are such that the first three contracts expire towards the end of three consecutive months, the first of which is the current one.[30] To construct the raw monetary surprise, at any date in the sample we use the next expiring SS futures, or front contract (SS1). Because liquidity in these markets tends to become very thin when the expiration date approaches, if the MPC date falls in the vicinity of the expiry date, the next contract is used instead. The top panel of the chart compares monthly surprises measured around announcement only (blue line) and all policy-relevant events in the same month, that is, the release of the minutes and of the Inflation Report (red dotted line). Surprises are computed in narrow 30-minute windows tightly surrounding the policy event. The historical set of policy rate decisions dates and times, and the decision that resulted from the committee meetings are reconstructed using Bloomberg. A different strategy is adopted in case of the release of the Inflation Report: due to the press conference associated with the release lasting a full hour, more flexibility is allowed in this case by employing a 90-minute window. Raw intraday data are from Thomson Reuters Tick History Database. For the construction of the monthly surprise we again follow Romer and Romer (2004) and assign each surprise to the month of the corresponding announcement.

In a non-negligible number of instances within the sample considered, some of the policy-relevant events around which the surprises are computed are contemporaneous to major macroeconomic data release. While the Bank Rate decision is typically released to the public at 12:00 noon, when no other data releases are scheduled, the release of the minutes and of the Inflation Report (IR) are contemporaneous to a number of relevant data releases that are also likely to substantially influence markets.[31] This is particularly true for the release of the minutes of the MPC meetings, the date and time of which often coincide with the release of labour market data and statistics on money and lending activities and, in some instances, GDP figures. To account for these interferences, in all cases we control for (standardised) data news falling within the time window around which the surprise is measured. Data news are computed as the difference between the released value and the median nowcast of the Bloomberg Survey of Economists as in Scotti (2013) and Altavilla, Giannone and Modugno (2014).

The top panel and the bottom left subplot of Figure B.3 reveal that while there are

[30] https://www.theice.com/products/37650330/Three-Month-Sterling-Short-Sterling-Future

[31] In the summer of 2015 the Bank of England adopted a different release schedule whereby the rate announcement and the minutes of the meeting are released simultaneously to the public at 12 noon. When the IR is also due for release, it is added to the block (e.g. "super Thursday" of August 6th, 2015).

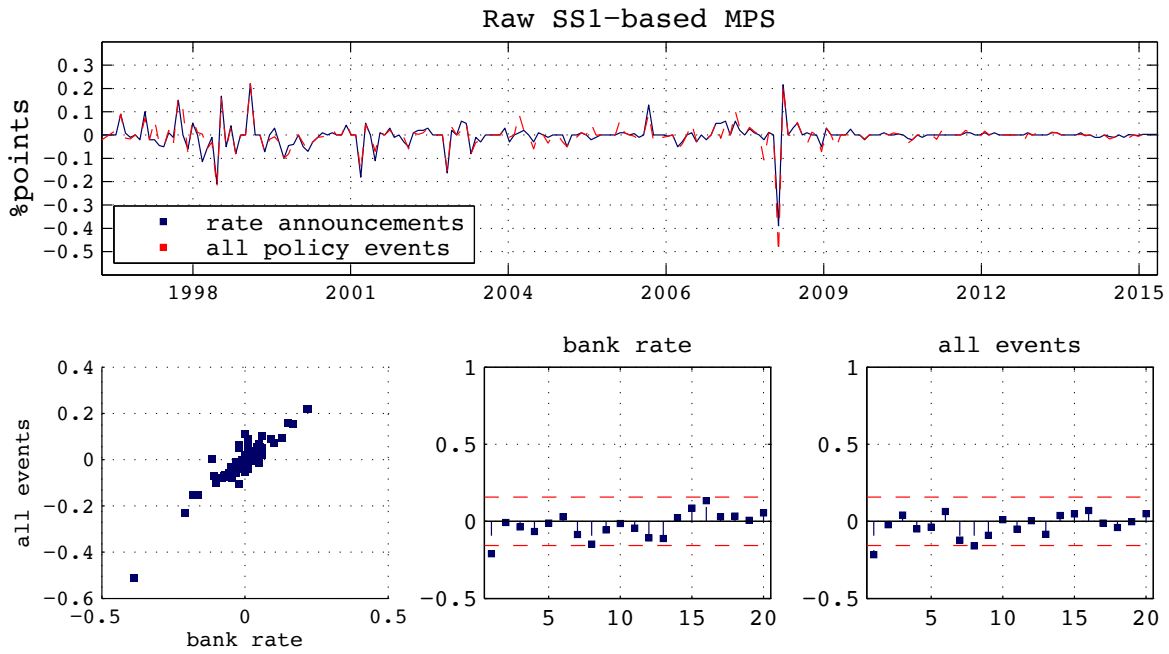
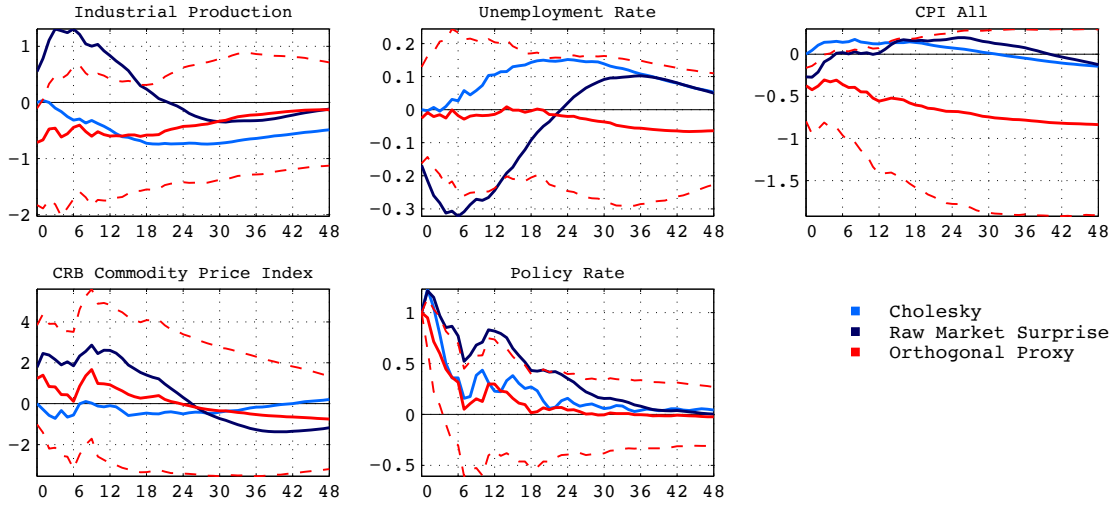


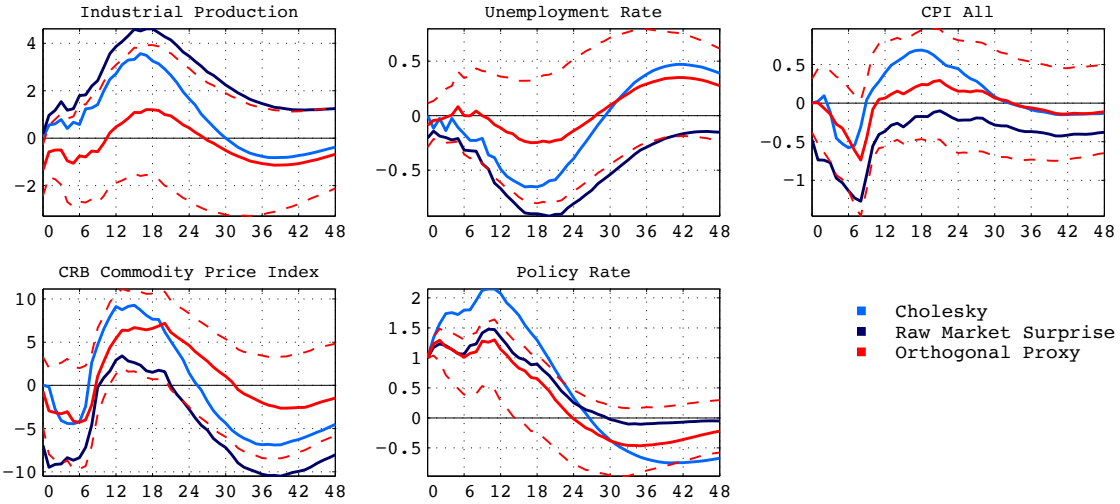
FIGURE B.3: Raw SS1-based monetary surprise at monthly frequency. Surprises are computed around Bank Rate announcements only (SS1 - blue line) and when also minutes and releases of the Inflation Report are taken into account (SS1MIR - red dotted line). All surprises control for data releases contemporaneous to the policy events in the sample considered. In the bottom panel, from left to right, the different information content in the two series and their partial autocorrelation functions.

some differences between the two series, expanding the set of policy events to include the minutes and the IR does not seem to modify substantially the overall information content of the monthly surprise series. We take this as evidence of the fact that on the day of the rate decisions, market participants infer what the MPC's assessment for current and future economic outlook is likely to be, and interpret the policy decision accordingly. Contrary to the US, raw UK-based monthly surprises display some (negative) autocorrelation even if no weighting scheme is adopted in their construction. The presence of autocorrelation in the first lag persists also if the effective lower bound period (post March 2009) is removed from the analysis.

C Additional Charts

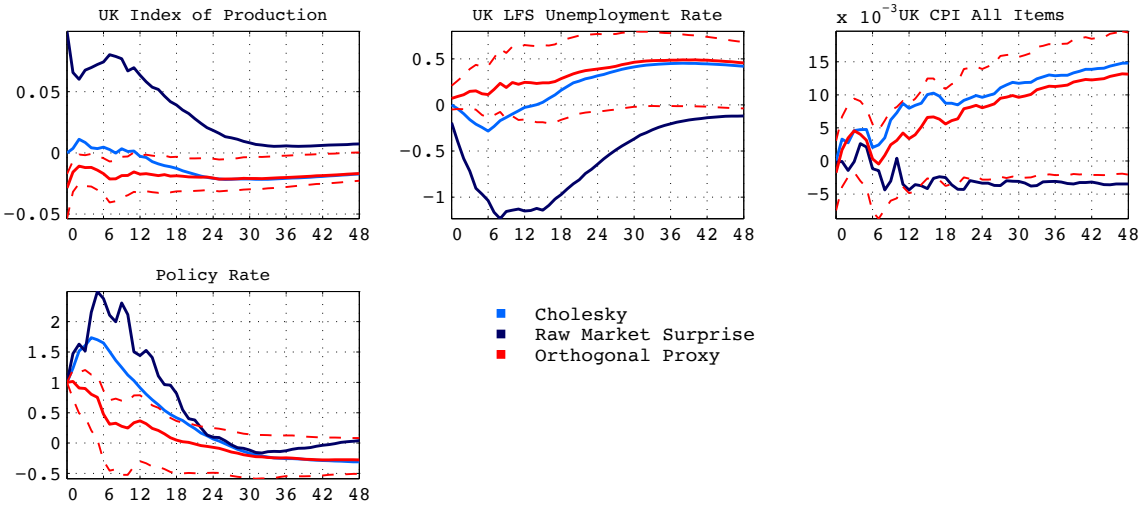


(A) US - VAR(12). Estimation sample 1969:1 - 2007:12, identification sample 1990:1 - 2007:12.

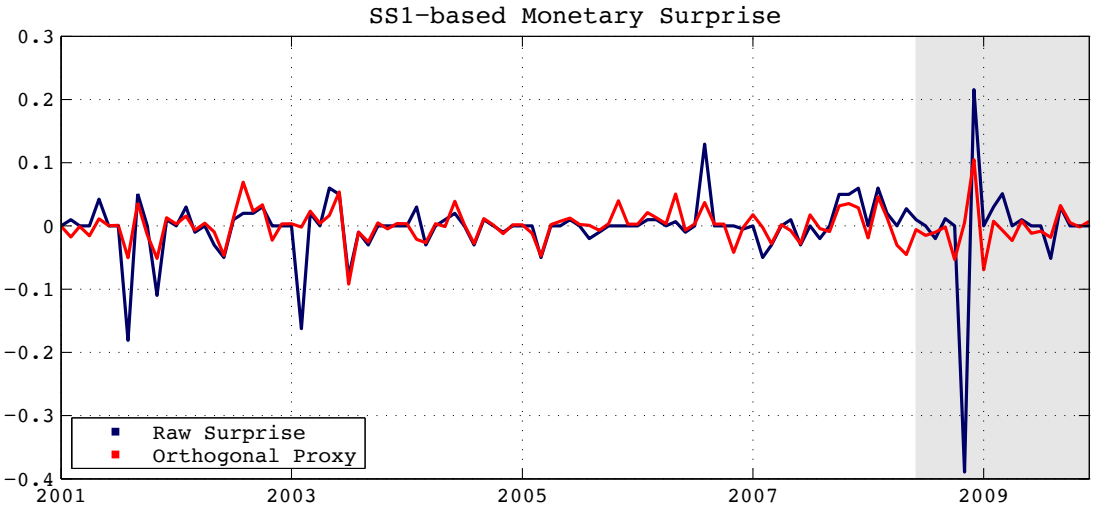


(B) US - VAR(12) Estimation sample 1990:1 - 2012:12, identification sample 1990:1 - 2009:12.

FIGURE C.1: US - Alternative estimation and identification samples. Recursive identification (light blue) vs identification with external instruments based on the weighted raw FF4 (dark blue) and on the orthogonal FF4* monetary surprise (red - PSVAR*). 90% bootstrapped confidence bands are obtained with 10,000 replications for the PSVAR* case. Shocks are normalized to induce a 1% increase in the policy rate. See main text for details.

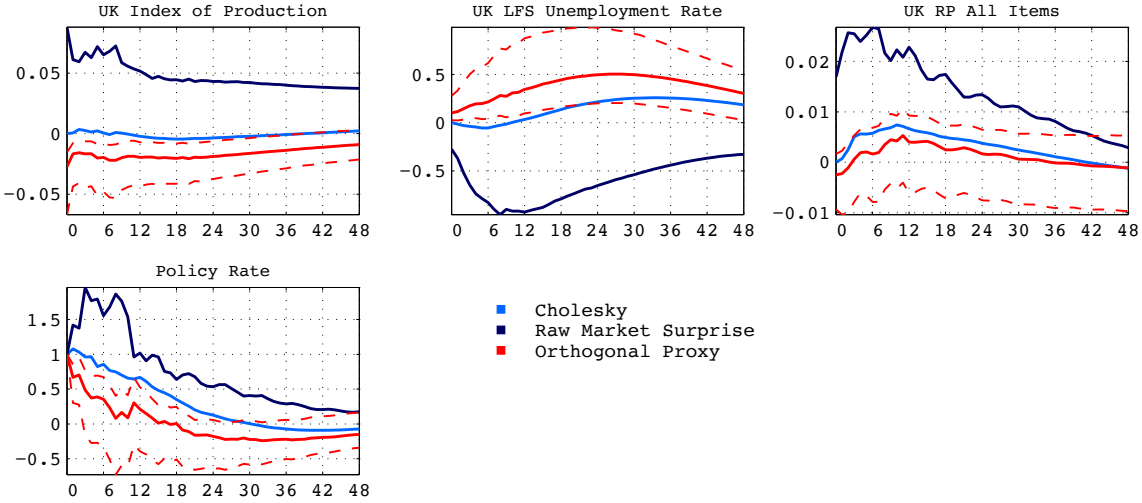


(A) UK - VAR(12). Estimation sample 1990:1 - 2014:12, identification sample 2001:1 - 2009:12. Recursive identification (light blue) vs identification with external instruments based on the weighted raw SS1 (dark blue) and on the orthogonal SS1* monetary surprise (red - PSVAR*). 90% bootstrapped confidence bands are obtained with 10,000 replications for the PSVAR* case. Shocks are normalized to induce a 1% increase in the policy rate.

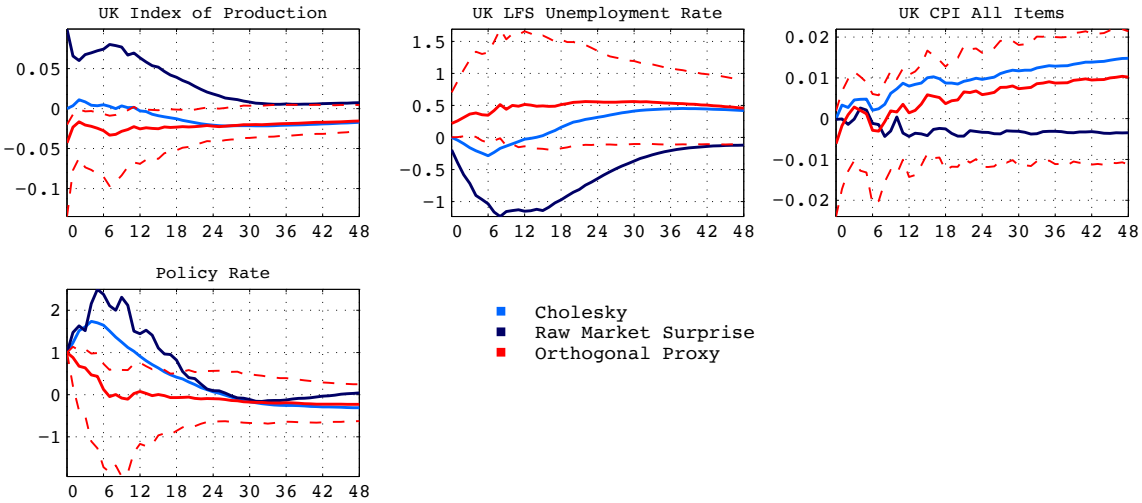


(B) UK - Benchmark Sample 2001:2009. Raw (SS1 – blue line) and orthogonal (SS1* – red line) monetary surprises at monthly frequency. Both sets of surprises are extracted from the first short sterling future. Shaded areas denote Economic Cycle Research Institute (ECRI) recessions.

FIGURE C.2: See main text for details.



(A) UK - VAR(12). Estimation sample 1979:1 - 2014:12, identification sample 2001:1 - 2015:3.



(B) UK - VAR(12) Estimation sample 1990:1 - 2014:12, identification sample 2001:1 - 2015:3.

FIGURE C.3: US - Alternative identification samples. Recursive identification (light blue) vs identification with external instruments based on the weighted raw SS1 (dark blue) and on the orthogonal SS1* monetary surprise (red - PSVAR*). 90% bootstrapped confidence bands are obtained with 10,000 replications for the PSVAR* case. Shocks are normalized to induce a 1% increase in the policy rate. See main text for details.