Unsurprising Shocks:
Information, Premia, and the Monetary Transmission

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Abstract

This paper proposes a new method to construct high-frequency futures-based proxies for the identification of monetary policy shocks in structural VARs that conditions on both central banks and market participants information sets. The use of narrow time frames to measure monetary policy surprises is potentially non sufficient to guarantee their exogeneity. Raw monetary "surprises" are, in fact, predictable. These findings are interpreted as suggesting that risk premia and news shocks are likely to be interfering. This results in a violation of the assumptions in Proxy SVARs and induces potentially non-trivial distortions in the estimation of the contemporaneous transmission coefficients. Consequences for the estimation of structural IRFs can be dramatic, both qualitatively and quantitatively. Identification of monetary policy shocks via the orthogonal proxies is shown to correctly retrieve the contemporaneous transmission coefficients even in small, potentially informationally insufficient VARs.

Keywords: Monetary Surprises; Identification; Monetary Policy; Expectations; Information Asymmetries; Event Study; Proxy SVAR.

JEL Classification: E52, E44, G14, C36

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Introduction

The idea of using financial market instruments to extract expectations about the monetary policy path dates 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), it would, however, take a few years for this insight to translate into an effective way to also measure monetary policy shocks.

Building on the work of Sack (2004), Gürkaynak et al. (2005) show how monetary policy surprises – defined as the price revision that follows a monetary policy announcement – can be effectively extracted from intraday quotes of futures on interest rate. To the extent that futures on interest rate embed accurate market-based expectations about the future path of policy rates, if the monetary surprises are computed within a sufficiently narrow window tightly surrounding the policy rate announcement, it should arguably be possible to interpret them as a measure, with error, of the underlying monetary policy shock. Motivated by this assumption, Gertler and Karadi (2015) use a transformation of the daily surprise measures constructed in Gürkaynak et al. (2005) as a *proxy* for the monetary policy shock in a Proxy Structural VAR (see Stock and Watson, 2012; Mertens and Ravn, 2013), and are thus able to identify the effects of unexpected monetary policy actions on a wider set of endogenous observables – among which credit spreads and term premia – for which more standard recursive identification schemes are ill suited, due to the implausibility of the contemporaneous zero restrictions. The availability of intraday data has since spurred a number of diverse applications whereby monetary surprises, extracted from financial market instruments linked more or less directly to a manifestation of the monetary policy stance, have been repeatedly used to quantify the effects of both conventional and unconventional monetary policy shocks. To mention just a few, Hanson and Stein (2015) diagnose 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. Motivated by a similar research question, Nakamura and Steinsson (2013) employ a "policy news shock" – defined as the principal component of monetary surprises calculated using a selection of interest rate futures with maximum maturity of about one year – to show that long-term nominal and real rates respond roughly one to one to monetary policy shocks, and that such changes are mostly due to changes in expectations. Similarly, Swanson (2015) identifies a "forward guidance" and "large-scale asset purchases" dimension 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 Treasury yields, stock prices and exchange rates. 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 et al. (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 area considered.

Two crucial assumptions make the futures-based surprises the ideal candidates for the role of external proxies for the monetary policy shock: (*i*) markets 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 premium; and (*ii*) the information set of the central bank and that of market participants coincide, leading to the equivalence between price updates and monetary policy shocks. Stated differently, these assumptions make it possible to first map all price updates 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 a 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 both risk premia and informational asymmetries are likely to pollute the measurement, thereby casting doubts on the exogeneity of the resulting proxies. If all the relevant information is effectively accounted for in market prices, it is then reasonable to expect that the surprises, being solely a function of the unexpected move in the target rate, be themselves unpredictable. If markets efficiently process all available information, no piece of data available prior to the announcement should be able to predict what the next "surprise" would be. We find that this is not the case. The predictability of monetary surprises – effectively market returns realised over a tiny time interval – suggests that either relevant pieces of information are intentionally ignored, or, more plausibly, that market participants require a risk compensation associated to the uncertainty about the future path of policy that is resolved precisely at the time of the announcement (see e.g. Fama and French, 1989; Fama, 1990, 2013). Likewise, raw surprises are predictable using central bank forecasts. This suggests that market participants are likely to infer the inputs (e.g. forecasts of inflation and output) of the central bank reaction function based on its interest rate decisions and that, therefore, rate announcements could potentially be the conveyor of both monetary and news shocks. If the central bank is adjusting the policy rate in a way that is fully consistent with its own rule and projections, which is typically the case, the fact that markets are nonetheless adjusting their expectations in any direction is relevant in its own right, but should not be interpreted as a measure of the monetary policy shock.

This paper contributes to the discussion on the identification of monetary policy shocks using high-frequency-based external proxies along different dimensions.

(i) While the use of intraday data and narrow time frames around which the surprises are calculated ensures a certain degree of insulation with respect to events other than the monetary policy announcements, it is shown that this is not necessarily sufficient to guarantee exogeneity of the resulting futures-based surprises. Central bank forecasts and past information that is available to market participants well before the date of the announcement are significantly predictive of future surprises. The predictability of monthly surprises makes the assumption concerning the validity of the proxy, impossible to verify in practice, also questionable from a theoretical *ex ante* point of view. This results in a violation of the key identifying assumptions and induces potentially non-trivial distortions in the estimated contemporaneous transmission coefficients. Consequences for the estimation of structural IRFs are shown to be dramatic, both qualitatively and quantitatively.

(ii) The paper thus proposes a new method to construct futures-based monetary policy surprises that conditions on both central banks' and market participants information sets. The composition of the conditioning set is similar to the one in Romer and Romer (2004). Conditioning on a summary measure of the information available to the agents is intended to fulfil the requirement that the proxy be orthogonal to premium effects. Conditioning on central banks forecasts, on the other hand, while not required specifically by the Proxy SVAR identification strategy, is crucial to make sure that what's being captured is in fact the monetary policy shock, and not a more general news shock.

The proposed approach for the construction of the orthogonal proxies has three main advantages: (1) it transforms the proxies *ex ante*, such that they can then be readily used regardless of the composition of the information set in the preferred reduced-form monetary VAR; (2) 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; (3) it includes the minimum set of controls to ensure that the proxies are neither endogenous nor measured around apparent deviations from the policy rule that were in fact responses to either current or future expected economic developments.

(*iii*) Applications to both the US and the UK show that (1) when the raw surprises are used as proxies for the monetary policy shock the identification procedure is invalidated and responses obtained using the corrupted proxies are highly distorted; (2) contrary to the benchmark cases, the use of the orthogonal proxies allows to correctly identify the monetary policy shock, recover the correct contemporaneous transmission coefficients, and deliver economically sound responses of the main output and price variables even in small, potentially informationally insufficient monetary VARs.

The paper is organised as follows. Section 1 discusses the properties that the candidate proxy for the structural shock is required to have for the contemporaneous transmission coefficients to be consistently estimated in a Proxy SVAR framework. A description of the variation in daily futures on monetary announcements days is reported in Section 2. Results on surprises predictability are in Section 3. Section 4 discusses the construction of the orthogonal surprises and impulse responses to monetary shocks identified using these measures as external proxies for the shock. Section 5 concludes.

The identification scheme is reviewed in Appendix A where also proofs are provided. The construction of the raw monetary surprises in both the US and the UK is detailed in Appendix B. Surprises for the UK, both raw and orthogonal, are novel. Details on the theoretical assumptions behind their construction and on the financial instruments from which they are extracted are thus reported. Issues related to the temporal aggregation of daily surprises (US) and to the informational content of surprises calculated around policy rate announcements only, as opposed to a broader collection of policy relevant events (UK) are also discussed. Additional charts are collected in Appendix C at the end of the paper. The full set of tables referring to surprises predictability and robustness checks are in the Online Appendix.

1 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. The identification relies on a set of key identifying assumptions that, being a function of unobserved shocks, are unverifiable in practice. For it to deliver consistent estimates of the transmission coefficients, however, the procedure also requires the candidate proxy to meet certain requirements that are only functions of observables and are thus fully testable. This section discusses the required properties of the external proxy and introduces the notation used in Appendix A, where the identification scheme is reviewed in detail and proofs are provided. Main references are Stock and Watson (2012); Mertens and Ravn (2013).

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

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

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

Suppose one is interested in calculating the responses of y_t to a particular shock in ε_t, call it the monetary policy shock and denote it by $\varepsilon_t^\bullet$. Within the Proxy SVAR framework, the identification of the relevant column $\mathrm{B}^\bullet$ of the $\mathbf{B}$ matrix is obtained via a set of r variables z_t, not in y_t, such that:

$$\begin{aligned} \mathbb{E}[\varepsilon_t^\bullet z_t'] &= \varphi', \\ \mathbb{E}[\varepsilon_t^\circ z_t'] &= 0, \end{aligned} \tag{2}$$

where φ is non-singular. ε_t° denotes structural shocks other than the one of interest. If one or more variables z_t can be found such that these conditions are satisfied, then it is possible to identify $\mathrm{B}^\bullet$ up to scale and sign using only moments of observables:

$$\mathbb{E}[u_t z_t'] = \mathbb{E}[\mathbf{B}\varepsilon_t z_t'] = \begin{bmatrix} \mathrm{B}^\bullet & \mathrm{B}^\circ \end{bmatrix} \begin{bmatrix} \mathbb{E}[\varepsilon_t^\bullet z_t'] \\ \mathbb{E}[\varepsilon_t^\circ z_t'] \end{bmatrix} = \mathrm{B}^\bullet \varphi'. \tag{3}$$

The conditions in (2) 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 proxy variables need to be relevant and exogenous with respect to the *unobserved* shocks. What this implies, in practice, is that the relevance of the proxy can only be assessed *after* the system is identified and a realisation of the structural shock $\varepsilon_t^\bullet$ is estimated; moreover, when – as it is the case in this example – only partial identification is achieved, exogeneity cannot be tested, and it is therefore crucial to make sure that the proxy variables are constructed in such a way that makes exogeneity a reasonable assumption to make.

An equivalent way of addressing the identification of $B^{\bullet}$ that also allows to neatly single out the observable characteristics that the external proxy z_t is required to have is to cast the problem in a measurement error framework.

Consider the following error-in-variables (EIV) model, where the true relationship of interest is

$$y_t = \mathbf{A}^{\star}\mathcal{Y}_t^{\star} + w_t, \tag{4}$$

where $\mathcal{Y}_t^{\star} \equiv [\mathcal{Y}_t', \; \varepsilon_t^{\bullet\prime}]'$ and $\mathcal{Y}_t$ is the $[np \times 1]$ vector collecting the lags of y_t. $\mathbf{A}^{\star} \equiv [\mathbf{A} \; B^{\bullet}]$, $\mathbf{A} \equiv [A_1, \ldots, A_p]$. Rather than $\mathcal{Y}_t^{\star}$, the researcher observes $\mathcal{Y}_t^{+}$ where

$$\mathcal{Y}_t^{+} \equiv [\mathcal{Y}_t', \; z_t']' = \Psi\mathcal{Y}_t^{\star} + \zeta_t. \tag{5}$$

z_t is a *proxy* for the unobserved "regressor" $\varepsilon_t^{\bullet}$ or, equivalently,

$$z_t = \Phi\varepsilon_t^{\bullet} + \nu_t, \tag{6}$$

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 $\mathbb{E}[\nu_t\nu_{\tau}'] = 0$, $\forall \tau \neq t$; Φ is non-singular. Therefore, z_t is effectively a scaled version of the shock up to a random measurement error.

The researcher thus estimates

$$y_t = \mathbf{C}\mathcal{Y}_t^{+} + \eta_t, \tag{7}$$

in lieu of the true model in (4). Because $\mathcal{Y}_t^{\star}$ is measured with error – (5), projecting y_t onto $\mathcal{Y}_t^{+}$ yields a biased estimate of $\mathbf{C}$; in particular, if $\hat{\mathbf{C}}$ denotes the least squares estimates of $\mathbf{C}$, and η_t and ζ_t are normally distributed, $\hat{\mathbf{C}} = \mathbf{C}\Lambda$, where Λ is the reliability matrix of $\mathcal{Y}_t^{+}$ (see equation A.7, Bowden and Turkington, 1984; Gleser, 1992).

Combining (7) with (5), we have that $\mathbf{A}^{\star} = \mathbf{C}\Psi$, therefore, the transmission coefficients in $B^{\bullet}$ can be recovered as a function of the parameters in the EIV system using

$\mathbf{A}^{\star} = \hat{\mathbf{C}}\Lambda^{-1}\Psi$, which, if $\mathbb{E}[z_t, \mathcal{Y}_t'] = 0$, reduces to (3).[1]

Whilst the identifying assumptions in (2) cannot be verified, the identification scheme based on the use of a proxy variable for the structural shock of interest, relies on a number of assumptions that only involve observables and are thus fully testable. As just discussed, if the external instrument is assumed to be as in (6), the procedure described in (3) delivers a consistent estimate of the transmission coefficients only if the instrument z_t is uncorrelated with the lagged endogenous included in the VAR, that is

$$\mathbb{E}[z_t \mathcal{Y}_t'] = 0. \tag{8}$$

Furthermore, (6) implies that

$$\mathbb{E}[z_t z_\tau'] = 0, \qquad \tau \neq t; \tag{9}$$

$$\mathbb{E}[z_t | \Omega_{t-1}] = 0, \tag{10}$$

where Ω_t denotes the information set at any time t. The conditions in (9) and (10) further require that, just like the shock itself, the proxy variable should not be forecastable given lagged information, be such 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 dynamic of z_t. The intuition here is that if this is not the case, then there is no real reason why one wouldn't 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). In fact, an equivalent way of estimating the transmission coefficients would be to include the proxy variable in the set of endogenous observables and identify the monetary policy shock by ordering it first in a standard Cholesky triangularisation.

The orthogonality assumption in (8) can be relaxed if the estimation of the contem-

[1] See proof in Appendix A.

poraneous transmission coefficients is achieved with a two-step procedure, rather than within the system in (7). 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 (3). It is important to notice, however, that the variance of z_t enters both the measures of instruments reliability Λ – equation (A.7), and the F-statistic customarily used to test the joint significance of the coefficients estimated in the regression of u_t onto z_t, that is, in the second stage of the procedure just described. When (9) does not hold, the presence of autocorrelation artificially increases both statistics and thus leads to overstating the effective relevance of the instrument.

Overall, successful identification of the contemporaneous transmission coefficients is ultimately a question of both specifying the VAR correctly, and singling out a proxy that can be reasonably thought of as being solely a function of the structural shock of interest.

2 A Closer Look at High-Frequency Responses

The following set of charts illustrate the daily variation in intraday quotes of interest-rate futures customarily used to measure monetary surprises, focussing on a selection of policy-relevant events. The contract used for the UK is the next expiring Short Sterling (SS) future – or front contract – that, depending on the relative market liquidity, can be either the one expiring within the current month [M0] or within the next month [M1]; these contracts embed expectations about the policy rate up to an horizon of about three months.[2] 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]. Results are robust to the use of the fourth FF contract which has a maturity of three months, roughly matching the horizons in SS.

To aid with the interpretation of the results, I use the median of the responses to the Bloomberg Survey of Economists (BSE) to proxy for markets' expectations about the

[2]More details on Short Sterling futures are provided in Section ?? of Appendix B.

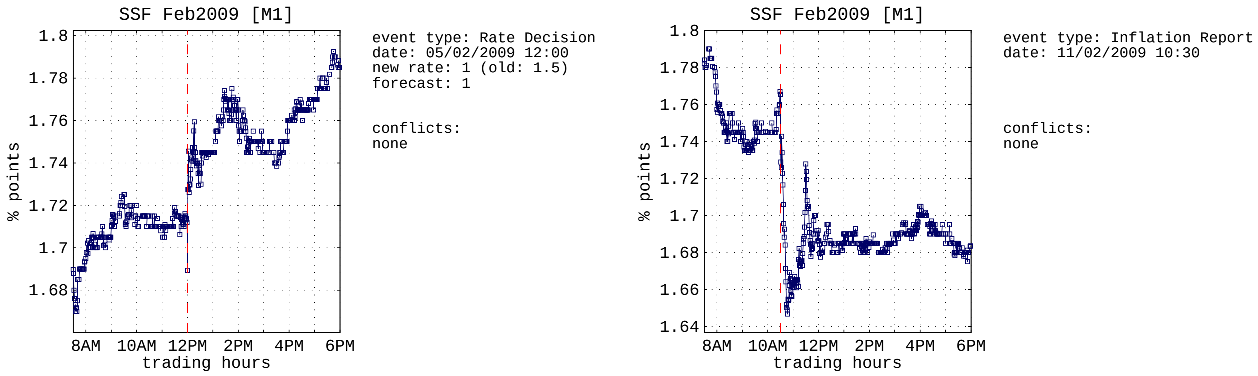

FIGURE 1: February 2009: The decision meeting of the MPC on the 5th [LEFT] is followed by the release of the Inflation Report on the 11th [RIGHT] where the Governor announces the MPC is ready to implement QE. The announcement reverts the reaction that followed the the 50bp rate reduction perceived on impact as unsatisfactory by market participants. *Source:* Bloomberg and Thomson Reuters Tick History Database, author's calculations.

policy rate;[3] moreover, all the episodes discussed in this section are selected among those for which there are no registered conflicting contemporaneous data releases. All times reported are London times.

The charts in Figure 1 show how the full impact of a single monetary action is shaped once more information about the events that motivate the actions are revealed and how market participants react not just to the decision itself, but also to the information about the state of the economy that they likely infer from central banks actions.

On February 5th, 2009, the Bank of England 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, markets reacted to the decision by *raising* futures rates (left panel of Figure 1). One possible explanation is that some players in the market were attaching some probability to an even larger move. However, an at least equally likely explanation is that the move can be at least in part explained by an increase in the risk premium fuelled by perceived news of deteriorating economic (and financial) conditions.[4] This particular MPC meeting followed the release, on January 23,

[3] 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.

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

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%.[5] The full scale of the MPC strategy was to be revealed shortly after the meeting when, 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 obvious during the press conference that the MPC was ready to launch its Quantitative Easing program. The announcement this time induced a visible contraction of futures quotes that fully reverted to the level at which they were prior to the interest rate decision; the press conference revealed 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*".[6]

A more striking picture is in Figure 2. All four episodes refer to days in which the Bank of England and the Fed decided not to change the level of the target interest rates.

In the top row, the MPC maintained the 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. Here the 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 median Bloomberg forecast reveals that market participants expected both the MPC and the FOMC not to move the policy rate in all instances,

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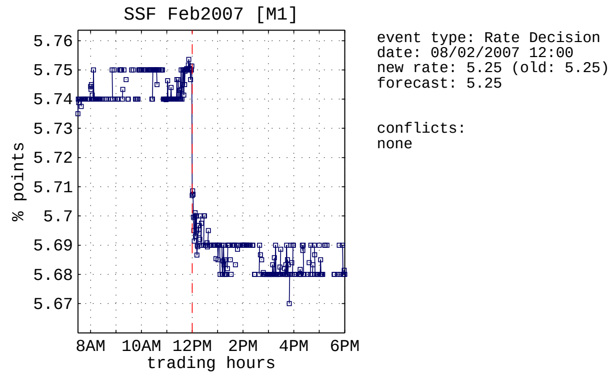
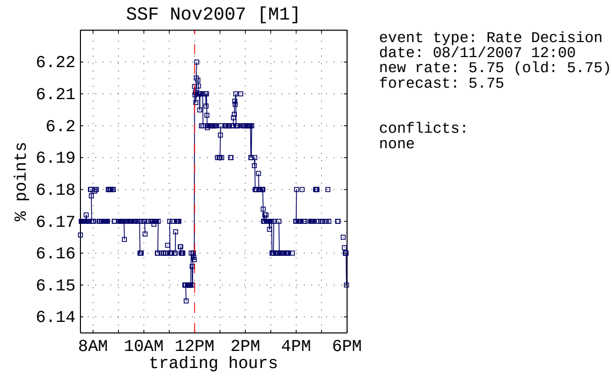
[5]Other significant data releases on the day of the MPC decision were the Halifax House Price Index for January at -17.20% on a 3month/year basis at 9:00 AM, US Jobless Claims relative to January at +38K compared to December at 1:30 PM, and US Factory Orders for December 2008 at -3.9% month-on-month,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

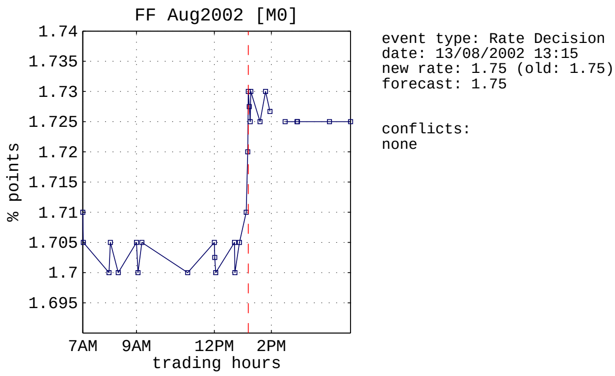
[6]Inflation Report Press Conference, Opening Remarks by the Governor, 11 February 2009: http://www.bankofengland.co.uk/publications/Pages/inflationreport/ir0901.aspx



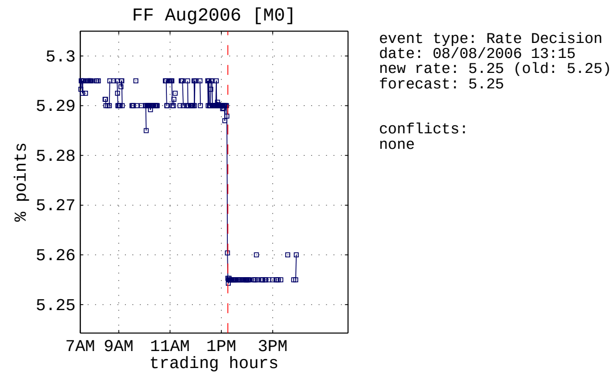
(A) MPC: fully anticipated no-change policy decision.



(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, the reaction is of Short Sterling futures around MPC decisions to maintain the Bank Rate at the current level. The bottom row shows intraday movements in Federal Fund futures on FOMC announcement days where again the decision resulted in the Target Fed Fund Rate being confirmed at the previous level. In all four instances, median Bloomberg forecasts indicate that the decisions were fully anticipated by markets. *Source:* Bloomberg and Thomson Reuters Tick History Database, author's calculations.

which makes these four moves fully anticipated by markets. 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 are then markets reacting at all? Being these four fully anticipated no-change moves, and holding the assumptions discussed in the introduction, one should reasonably expect that following any event as the ones in Figure 2, markets would have no need of adjusting. Furthermore, and more strikingly, Figure 2 shows how not only markets can and do react to fully anticipated moves, but they can also do so in opposite, seemingly inconsistent ways. While this is hard to reconcile with a framework in which markets information sets are aligned with that of the central bank and prices only adjust following revision in expectation triggered by an unexpected policy rate change,

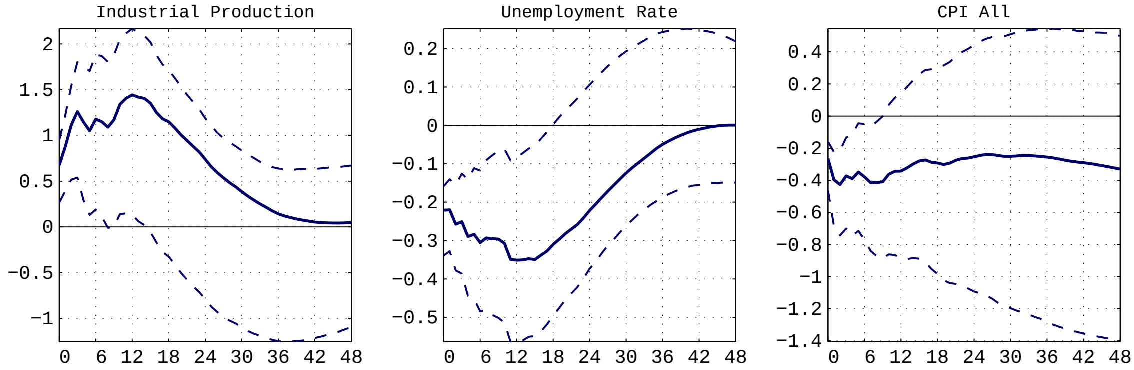


FIGURE 3: Impulse response functions to a shock inducing a 100bp increase in the policy rate identified using the raw FF4-based proxy as an external instrument. VAR(12) estimated in levels over 1969:1 - 2014:12. The monetary policy variable is the 1-year rate. Dashed lines limit 90% confidence bands obtained using 10,000 bootstrap replications. The full set of IRFs is in Figure 6.

this behaviour can be rationalised by allowing markets 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 seem to suggest that monetary surprises are potentially a contemporaneous function of more than just the monetary policy shock. If one is willing to accept this interpretation, then it is easy to see that if the VAR in use does not properly account for future expectations and premia by including them in the set of endogenous variables, raw monetary surprises 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 4 and depict responses to a monetary policy shock identified using the raw monetary surprise computed using the fourth Federal Fund future (FF4) as an external proxy (z_t in Section 1). The selection of the variables 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; crucially, however, this specification of the VAR intentionally leaves out the Excess Bond Premium of Gilchrist and Zakrajšek (2012). The shock is normalised 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. We interpret these anomalous responses as reflecting the extent to which confounding the shocks can induce distortions in the estimates of the contemporaneous transmission coefficients. In particular, we postulate that the reaction of both output and unemployment can be partly rationalised as the effect of a news shock, that is, an increase in the policy rate signals that the central bank is forecasting improved economic conditions ahead. Conversely, interpreting the sign of the effect of 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).

3 Predictable Surprises

This section addresses the concerns in Section 2 more formally. In what follows, raw US surprises are those in Gürkaynak et al. (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 future (SS1). To assess the behaviour of market participants around policy relevant events other than the rate announcements, UK raw surprises are also computed to account for the release of the minutes of the meetings (SS1M), and of the quarterly Inflation Report (SS1MIR). The reader is referred to Appendix B for a throughout description of the raw surprises and the financial instruments used for their construction.

Tables 1 and 2 collect results from predictive regressions where the raw surprises are projected onto different sets of variables that are intended to summarise the information set of both market participants and the central bank. Formally, the tables report the adjusted R^2 and the F statistics for the regression

$$y_t - y_{t-\Delta t} = \kappa_{\mathrm{c}} + \kappa_{\mathrm{x}} \mathrm{X}_{t-1} + \epsilon_t, \tag{11}$$

where $y_t - y_{t-\Delta t}$ is the raw monetary surprise and X_{t-1} is a set of observables whose realisation is known before the announcement (i.e. $\Delta t < 1$). Full regression outputs are collected in the online appendix, that also reports robustness checks. The regressions are estimated in-sample and at monthly frequency. The length of the measurement window (Δt) is typically equal to 30 minutes, with the exception of the (SS1MIR) case, for which 90 minutes are allowed.[7] When X_{t-1} contains either data or factors, these enter the specification with a month lag, conversely, when predictability is tested against collections of forecasts, these are aligned such that the compilation of the forecast always precedes the monetary surprise.

It should be stressed that these results are produced in support of the claims made in Section 2 concerning the possibility that raw monetary surprises are not just a function of the underlying monetary policy shock, and that time varying risk premia (proxied by lagged observables) and news about future developments (proxied by central bank projections) are significantly informative of future surprises. We therefore 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.

[7]See Appendix B for details.

	$MP1_t$		$FF4_t$		$ED2_t$		$ED3_t$		$ED4_t$	
	R^2	F	R^2	F	R^2	F	R^2	F	R^2	F
Macro-Financial Factors	0.044	2.243**	0.079	3.298***	0.097	3.871***	0.070	3.004**	0.056	2.579**
Lagged Observables										
ISM Composite	0.043	13.16***	0.072	21.82***	0.094	28.78***	0.076	23.03***	0.063	19.13***
Consumer Sentiment	0.000	1.56	0.008	3.13*	0.018	5.93**	0.020	6.46**	0.016	5.22**
Effective Fed Funds Rate	0.018	5.83**	0.046	13.80***	0.078	23.52***	0.063	18.96***	0.052	15.75***
3M T-bill Spread	0.089	27.05***	0.053	15.88***	0.027	8.38***	0.020	6.57**	0.014	4.67**
AAA-FFR Spread	0.010	3.69*	0.000	0.85	0.000	0.02	0.000	0.06	0.000	0.09
1Y T-Bond FFR Spread	0.054	16.19***	0.035	10.70***	0.021	6.73**	0.017	5.67**	0.013	4.44**
S&P 500 Composite	0.009	3.52*	0.017	5.72**	0.032	9.92***	0.040	12.12***	0.034	10.25***
CBOE VIX	0.018	5.82**	0.019	6.14**	0.056	16.82***	0.057	17.23***	0.052	15.71***
Greenbook Forecasts										
Output	0.054	2.844**	0.089	4.154**	0.128	5.774***	0.113	5.141***	0.074	3.571***
Inflation	0.000	1.059	0.000	1.009	0.000	0.594	0.000	0.469	0.000	0.207
Unemployment	0.027	1.888*	0.031	2.033*	0.104	4.746***	0.065	3.267**	0.039	2.310**
Greenbook Forecasts Revisions										
Output	0.086	4.831**	0.102	5.576***	0.135	7.346***	0.117	6.387***	0.090	4.987***
Inflation	0.000	0.531	0.000	0.483	0.000	0.580	0.000	0.410	0.000	0.311
Unemployment	0.024	2.002*	0.066	3.848*	0.065	3.812*	0.048	3.047**	0.045	2.916**

TABLE 1: Sufficient information in US-based raw monetary policy surprises. The table reports adjusted R^2 and F statistics for the null $H_0 : \kappa_x = 0$ in (11) 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 Fund future (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.

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 in McCracken and Ng (2015).[8] Surprises are predictable by past information, summarised by the macro-financial factors. Individual t-statistics (not reported) are significant at least at 5% level for three out of the ten factors and for all the raw surprises. The joint null (reported in the top row of the table) is rejected at 1% level in almost all cases. One concern with regressing on factors is that they are estimated on the last available vintage of data, that thus include 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 they are extracted was not entirely visible to market participants 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 difference with the exception of surveys and spreads. Both surveys and financial data, not subject to revisions and relative to the month prior to the announcement, are significantly predictive of future monthly surprises.[9] These results complement the findings in Piazzesi and Swanson (2008) and suggest that monthly raw monetary surprises seem to be significantly contaminated by time variation in risk premia.[10]

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 a year ahead. Greenbook forecasts are aligned to match the FOMC meeting they refer to. Results in the table confirm that forecasts and successive

[8] 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 variance 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 initialised with static principal components and OLS for the state space parameters. Prior to estimation, all variables are opportunely transformed to achieve stationarity.

[9] A more comprehensive set of regression results is reported in the Online Appendix. Results on predictability survive if the surprises are computed using only scheduled FOMC meetings. 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.

[10] 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.

	$SS1_t$		$SS1M_t$		$SS1MIR_t$	
	R^2	F	R^2	F	R^2	F
Macro-Financial Factors	0.021	2.164*	0.014	1.770	0.029	0.038**
Lagged Observables						
PMI Composite	0.000	0.72	0.000	0.24	0.000	0.80
CPI All Items	0.032	7.93***	0.026	6.62**	0.031	7.74***
Consumer Confidence	0.000	0.000	0.000	0.05	0.000	0.760
Bank Rate	0.004	1.92	0.004	1.74	0.000	0.84
FTSE All Share	0.020	5.64**	0.017	4.74**	0.027	6.88***
3M LIBOR	0.029	7.27***	0.028	7.31***	0.022	5.70**
3M LIBOR-OIS Spread	0.144	29.45***	0.144	29.44***	0.154	31.70***
1Y Gilt Spread	0.050	12.66***	0.050	12.35***	0.052	12.66***
Official Reserves	0.027	7.06***	0.023	6.03***	0.028	7.06***
Inflation Report Forecasts						
Output	0.000	0.051	0.000	0.069	0.000	0.048
Inflation	0.000	0.033	0.000	0.028	0.000	0.059
Unemployment	0.000	0.047	0.000	0.156	0.000	0.177
IR Forecasts Revisions						
Output	0.000	0.035	0.000	0.038	0.000	0.028
Inflation	0.000	0.114	0.000	0.041	0.000	0.050
Unemployment	0.000	0.031	0.000	0.512	0.000	0.399

TABLE 2: Sufficient information in UK-based raw monetary policy surprises. The table reports adjusted R^2 and F statistics for the null $\mathrm{H}_0 : \kappa_{\mathrm{x}} = 0$ in (11) 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.

forecast revisions for output and unemployment are highly informative for all the raw monetary surprises considered.[11]

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 variables.[12] As it was the case for the US, there is evidence that monthly raw surprises extracted from Short Sterling futures are also

[11]Predictability with respect to Greenbook forecasts is also noted in Gertler and Karadi (2015) and Ramey (2015).

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

contaminated by time-varying risk premia. On the other hand, however we do not find evidence of predictability with respect to the forecasts and forecasts revisions contained in the Inflation Report. It should be noted, however, that the quarterly availability of the Report makes estimates more uncertain with respect to the US case.

4 Conditional Monetary Policy Surprises and Shock Identification

The results collected in the previous suggest that the raw monthly monetary surprises cannot be safely used as proxies for the monetary policy shock unconditionally. As shown, the mere fact of narrowing down the measurement window to a short span surrounding the time of the announcement does not guarantee that the raw surprises thus computed are in fact a clean measure of the underlying monetary policy shock. Within the Proxy SVAR framework, successful identification is ultimately a combination of both using a valid external proxy, and correctly specifying the VAR, that is, ensuring that the information included in the set of the endogenous variables is sufficient, in the sense of Forni and Gambetti (2014). If, however, small-scale VARs are the basis for the analysis, information deficiency is a non-negligible risk that must be mitigated by correcting the proxies in a way that ensures that what is being captured are in fact exogenous, unexpected policy changes.

4.1 Orthogonal Proxies

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 variables intended to capture both central banks' private information, and a summary measure of the information available to the agents. The latter component of the conditioning set is intended to clean the dependence on time-varying risk premia. The necessity of conditioning on central banks forecasts, on the other hand, while not required specifically by the Proxy SVAR identification strategy, is crucial to make sure that what's being captured is in fact the monetary policy shock, and not a more general news shock which results from market participants trying to infer the central bank's projections from their decision on the target policy rate. If the central bank is adjusting the policy rate in a way that is fully consistent with its own rule and projections, which is typically the case, the fact that markets are nonetheless adjusting their expectations in any direction is relevant in its own right, but should not be interpreted as a measure of the monetary policy shock. As shown in Miranda-Agrippino and Ricco (2015), this procedure is also consistent with a framework in which monetary policy responses are calculated in presence of informational asymmetries/frictions; in this case, markets are not assumed to be perfectly incorporating information, and thus monetary surprises are rather interpreted as the result of inattentive market participants only partially acquiring the information being released through central bank's communication.

The proposed approach for the construction of the orthogonal proxies has three main advantages: (*i*) it transforms the proxies *ex ante*, such that they can then be readily used regardless of 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 the minimum set of controls to ensure that the proxies are effectively capturing surprises orthogonal to the all the available information and that result from policy decisions that are not taken in response to either current or future economic developments.

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 a year ahead and of current unemployment – as a proxy for central bank's private information and (b) the lagged bank rate and the observed change in the target rate markets respond to, as a proxy for markets' information. The composition of the conditioning set resembles the one in Romer and Romer (2004) and ensures that only unsystematic policy changes are used, and that, to the extent that monetary policy typically moves in response to changes in macroeconomic and financial conditions, past target rates are a

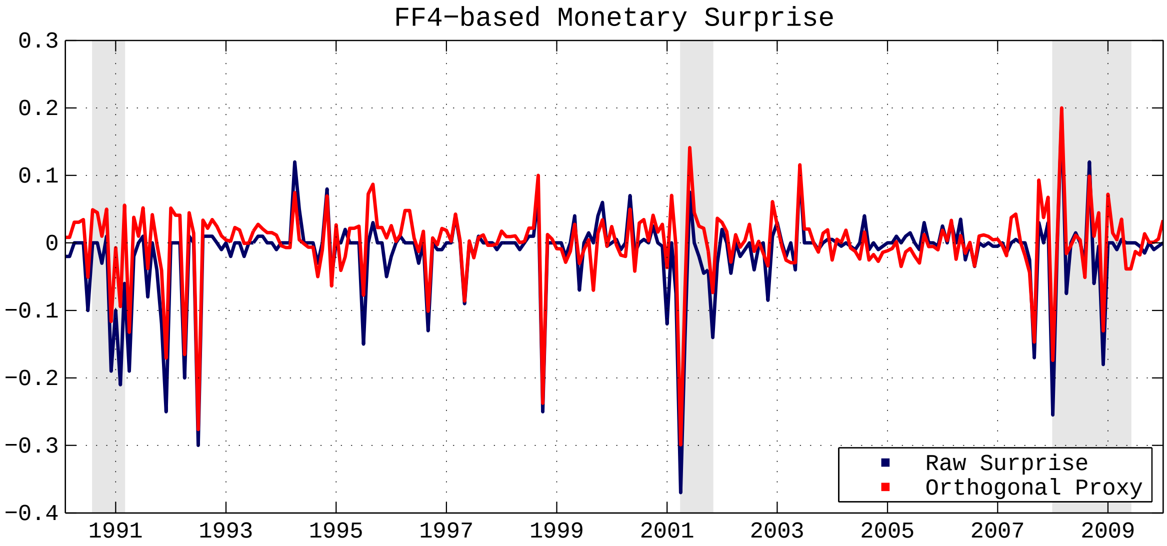


FIGURE 4: Raw and orthogonal monetary policy surprises at monthly frequency. The chart plots the raw surprise extracted from the fourth Federal Funds future (FF4 – blue line) and the surprise orthogonal to both central bank's and market participants' information sets (FF4* – red line). Shaded areas denote NBER recessions. See main text for details.

sufficient measure of the state of the economy. 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.[13] The raw (FF4) and orthogonal (FF4*) monthly surprises extracted from the fourth Federal Funds future are plotted in Figure 4.

As discussed in Appendix B, measuring responses to a monetary policy shock in the UK using high-frequency futures data presents some difficulties primarily linked to the fact that no financial contract with a sufficiently long history is directly linked to the Bank Rate. A further complication arises from the fact that, contrary to the FOMC, the Bank of England's MPC do not inform their judgement on official forecasts that are updated prior to all scheduled meetings, that is, there is no equivalent of the Greenbook forecasts to proxy for the central bank's private information. To overcome these issues, the conditioning set over which the raw monetary surprises are calculated is rather composed by (a) forecasts and forecast revisions for output and inflation for the previous and the current quarter and up to a year ahead and for current unemployment extracted

[13]Specifically, MP1: F =0.684 (p-value 0.738); FF4: 1.035 (0.417); ED2: 1.184 (0.306); ED3: 1.118 (0.352); ED4: 1.231 (0.276).

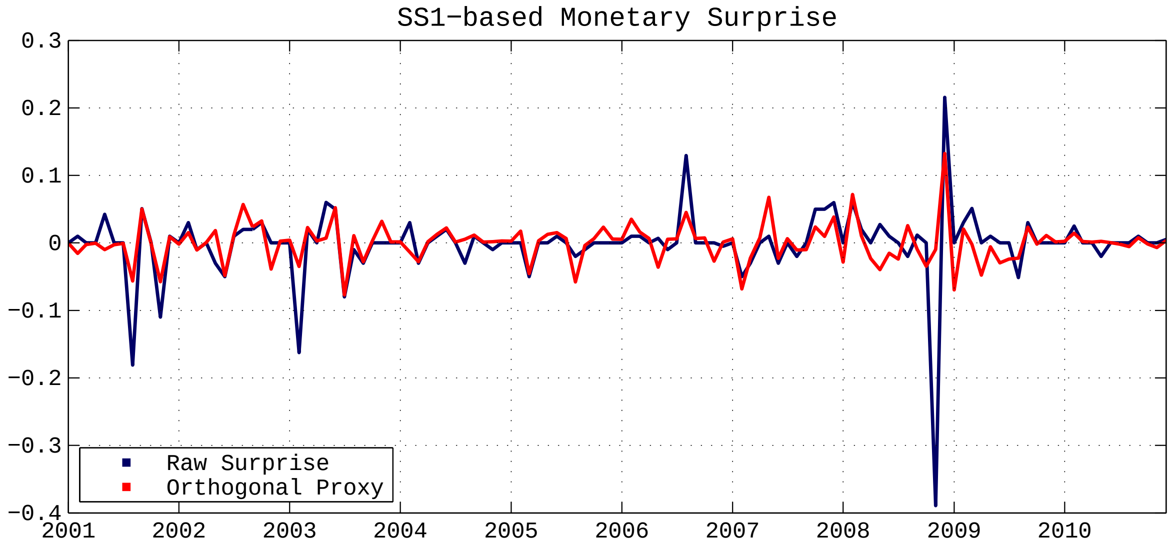


FIGURE 5: Raw and orthogonal monetary policy surprises at monthly frequency. The chart plots the raw surprise extracted from the first Short Sterling future (SS1 – blue line) and the surprise orthogonal to both central bank's and market participants' information sets (SS1⋆ – red line). See main text for details.

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 rate markets respond to, as a proxy for markets' information. The use of Inflation Report forecasts to proxy for the Bank of England's private information 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. The inclusion of the LIBOR-OIS spread is motivated by the decomposition in equation (B.9). Being linked to the Sterling-based LIBOR rate, 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 (see Figure B.2), it needs to be controlled for when extracting revisions in expectations about the policy rate.[14] The raw UK monetary surprise used is the one computed around rate announcements only (i.e. blue line in Figure B.3). The orthogonal surprise SS1⋆ is plotted in Figure 5 against its raw counterpart SS1. It is worth noticing that the largest peak in the raw surprises disappears in the

[14]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, I use the daily spread instead.

orthogonal series, in support to the claim that not all price movements contemporaneous to policy announcements are necessarily a reaction to a monetary policy shock only. In fact, the peak coincides with the sudden increase in the Libor-OIS 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.

4.2 Identification of Monetary Policy Shocks

US I finally 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 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 (2015), 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, IRFs to a monetary policy shock identified using a recursive Cholesky scheme with the effective federal funds rate 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.[15] Responses are normalised 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 Effective Fed Fund Rate ordered last (CHOL). Dark blue lines are obtained

[15]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.

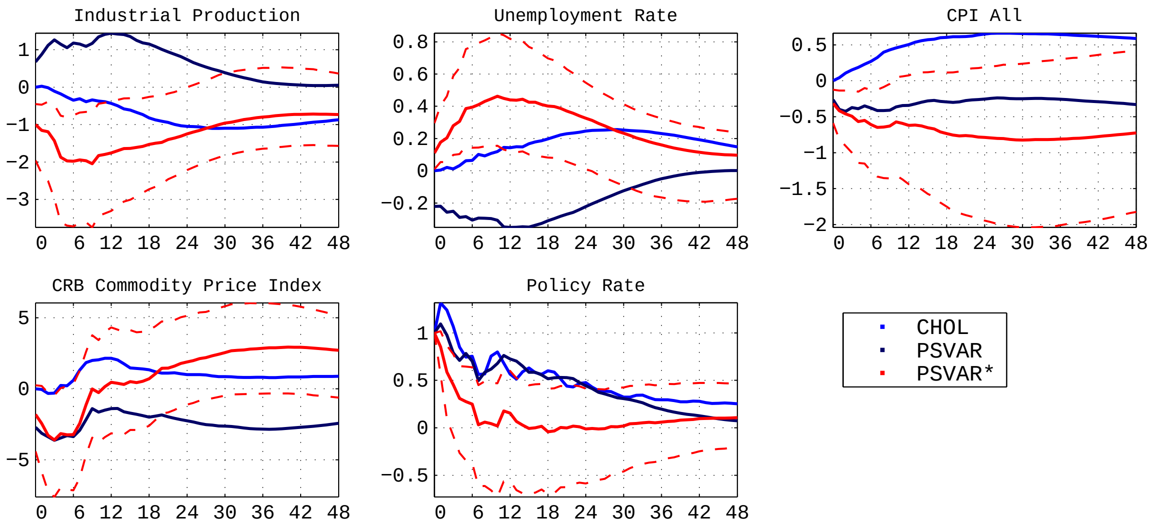


FIGURE 6: The chart compares impulse responses to a monetary policy shock obtained estimating a VAR(12) over the sample 1969:1 - 2014:12 and using different identification schemes. Light blue lines are for the recursive identification scheme with the Effective Fed Fund Rate ordered last (CHOL). Dark 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.

when the shock is identified using the raw FF4-based surprise (PSVAR). 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 PSVAR lie outside the confidence bands of PSVAR⋆ in almost all cases. The issues highlighted for the raw weighted 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 an equally small VAR where, however, they include the excess bond premium (EBP) of Gilchrist and Zakrajšek (2012). Other than a good (in-sample) predictor of real activity, the EBP is constructed using micro-level data on corporate spreads with average maturity of about 7 years. This is likely to be at least partially capturing forecasts about future realisations that "clean" the VAR residuals and thus still deliver responses of the expected sign.[16] On the 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 – Figures C.1 and C.2.

UK The quality of the conditional and unconditional monetary SS1-based surprises is evaluated in their ability to recover consistent responses to monetary policy shocks in a Proxy SVAR for the 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 are the log of industrial production, the LFS unemployment rate and the log of RPI.[17] The VAR is estimated in levels with 12 lags over the period 1979:1 to 2014:12, responses are again normalised such that the policy rate increases by 1% on impact. The identification of the contemporaneous transmission coefficients uses the full length of the orthogonal SS1*, that is 2001:1 - 2009:12.[18]

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 the 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*. 90% bootstrapped confidence bands are obtained

[16] 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 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.

[17] The Bank Rate and 1-year government bond rate are from the Bank of England; prices, output and unemployment data are from the Office of National Statistics.

[18] While IR forecasts are released at quarterly frequency and with no significant lag, and thus their timely availability is not a concern, I end the sample in 2009:12 to avoid introducing distortions caused by the Bank Rate reaching the zero (effective) lower bound in 2009. The lower time bound to the construction of the orthogonal surprise is constrained by the availability of the LIBOR-OIS spread.

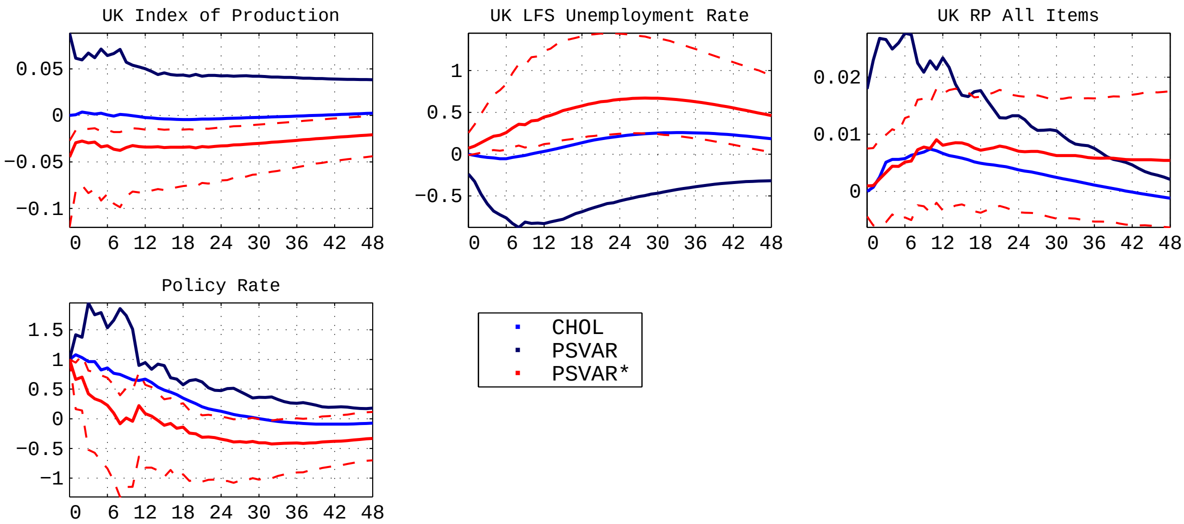


FIGURE 7: The chart compares impulse responses to a monetary policy shock obtained estimating a VAR(12) over the sample 1979:1 - 2014:12 and using different identification schemes. Light blue lines are for the recursive identification scheme with the Bank Rate ordered last (CHOL). Dark blue lines are 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.

with 10,000 replications for the PSVAR* case. Responses in Figure 7 confirm the extent to which responses can be biased 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 it was the case in 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 C.3, 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.

5 Concluding Remarks

Recent advances in the identification of monetary policy shocks have proposed the use of market-based 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 working assumption made throughout is that, to the extent that futures on interest rate provide accurate measures of market-based expectations of future policy rates, if the surprises are computed within a sufficiently narrow window tightly surrounding the monetary announcement, the change in intraday interest rate futures around these times can be regarded as a measure, with error, of the underlying monetary policy shock. Two crucial assumptions make the futures-based surprises the ideal candidates for the role of external proxies for the monetary policy shock: (*i*) markets 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 premium; and (*ii*) the information set of the central bank and that of market participants coincide, leading to the equivalence between price updates and monetary policy shocks. Stated differently, these assumptions make it possible to first map all price updates 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 a 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 both risk premia and informational asymmetries are likely to pollute the measurement, thereby casting doubts on the exogeneity of the resulting proxies. Raw monthly monetary "surprises" are shown to be predictable using both private central bank's forecasts and past information that was available to market participants well before the announcements. Building on the predictability of the raw surprises, this paper proposes a new set of proxies that are orthogonal to both central banks' and market participants' information sets, and are thus better candidates for the task of capturing only the unexpected monetary policy decisions. The latter component of the conditioning set is intended clean the surprises of time-varying risk premia considerations and, more generally, to make them exogenous with respect to the history of the endogenous variables that might be included in the VAR, but without having to know *a priori* what the researcher's preferred VAR model would ultimately be. The necessity of conditioning on central banks forecasts, on the other hand, while not required specifically by the Proxy SVAR identification strategy, is crucial to make sure that what's being captured is in fact the monetary policy shock, and not a more general news shock which results from market participants trying to infer the central bank's projections from their decision on the target policy rate.

Results on 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, less reliant on the composition of the VAR information set and invariant to sample splits.

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A Proxy SVAR Identification

Using the notation introduced in Section 1, let y_t be an n-dimensional vector of endogenous observables whose dynamic is described by the following system of equations:[19]

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

where $\mathbf{B}^{-1}$ and $\mathcal{A}_i$, $i = 1, \ldots, p$, are square matrices of structural coefficients and ε_t is an n-dimensional vector of structural shocks such that $\mathbb{E}[\varepsilon_t] = 0$, $\mathbb{E}[\varepsilon_t \varepsilon_t'] = \mathbb{I}_n$ and $\mathbb{E}[\varepsilon_t \varepsilon_\tau'] = 0$, $\forall \tau \neq t$. 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 combination of the structural shocks:

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

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 realisations of the structural shocks:

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

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 restriction such that identification can be achieved. A prime example entails assuming that $\mathbf{B}$ is lower triangular and equal to the Cholesky factor of Σ_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,

[19] The content of this Appendix draws heavily from Mertens and Ravn (2013); Montiel Olea et al. (2012).

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

$$u_t = \mathbf{B}\varepsilon_t = \begin{bmatrix} \mathrm{B}^\bullet & \mathrm{B}^\circ \end{bmatrix} \begin{bmatrix} \varepsilon_t^\bullet \\ \varepsilon_t^\circ \end{bmatrix}, \tag{A.5}$$

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

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

$$\begin{aligned} \mathbb{E}[\varepsilon_t^\bullet z_t'] &= \varphi', \\ \mathbb{E}[\varepsilon_t^\circ z_t'] &= 0, \end{aligned} \tag{2}$$

where φ is non-singular. If a variable z_t can be found such that the validity conditions in (2) are satisfied, then it is possible to identify $\mathrm{B}^\bullet$ up to scale and sign:

$$\mathbb{E}[u_t z_t'] = \mathbb{E}[\mathbf{B}\varepsilon_t z_t'] = \begin{bmatrix} \mathrm{B}^\bullet & \mathrm{B}^\circ \end{bmatrix} \begin{bmatrix} \mathbb{E}[\varepsilon_t^\bullet z_t'] \\ \mathbb{E}[\varepsilon_t^\circ z_t'] \end{bmatrix} = \mathrm{B}^\bullet \varphi', \tag{3}$$

implying that further normalisation is needed to back out the elements in $\mathrm{B}^\bullet$. Montiel Olea et al. (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 $\mathrm{B}^\bullet$ equal to 1. In what follows, let v_i denote the i-th element of any column vector V, and use the subscript notation $\backslash i$ to denote the sub-vector of V not containing v_i, such that $[\mathrm{v}_1, \mathrm{v}_{\backslash 1}']'$ is a partition of V. With $\mathrm{B}^\bullet = [1, \mathrm{b}_{\backslash 1}^{\bullet\prime}]'$, and using the relation established in (3)

$$\begin{bmatrix} \mathrm{b}_{\backslash 1}^\bullet \mathbb{E}[u_{1,t} z_t'] \\ \mathbb{E}[u_{\backslash 1,t} z_t'] \end{bmatrix} = \begin{bmatrix} \mathrm{b}_{\backslash 1}^\bullet \varphi' \\ \mathrm{b}_{\backslash 1}^\bullet \varphi' \end{bmatrix},$$

which, rearranging terms, is equivalent to writing:

$$\mathrm{b}_{\backslash 1}^\bullet \mathbb{E}[u_{1,t} z_t'] = \mathbb{E}[u_{\backslash 1,t} z_t']. \tag{A.6}$$

Equation (A.6) establishes that, given the normalisation discussed above, the elements of $\mathrm{B}^\bullet$ can be estimated using moments of observables; in particular, if z_t only contains one proxy variable, $\mathrm{b}_{\backslash 1}^\bullet = \mathbb{E}[u_{\backslash 1,t} z_t'] / \mathbb{E}[u_{1,t} z_t']$, that is, it is equal to the ratio between the coefficients of the regression of the reduced-form VAR innovations onto the instrument.[20]

A.1 The contemporaneous transmission coefficients in the EIV framework

Let the true model be:

$$y_t = \mathbf{A}^{\star}\mathcal{Y}_t^{\star} + w_t, \tag{4}$$

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

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

$$z_t = \Phi\varepsilon_t^{\bullet} + \nu_t, \tag{6}$$

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

$$y_t = \mathbf{C}\mathcal{Y}_t^{+} + \eta_t, \tag{7}$$

where

$$\mathcal{Y}_t^{+} \equiv [\mathcal{Y}_t',\ z_t']' = \Psi\mathcal{Y}_t^{\star} + \zeta_t. \tag{5}$$

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

$$\Lambda = [\Sigma_{\mathcal{Y}^{+}}]^{-1}[\Sigma_{\mathcal{Y}^{+}} - \Sigma_{\zeta}] \tag{A.7}$$

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.

[20] An alternative formulation is discussed in Mertens and Ravn (2013), where $\mathrm{b}_{\setminus 1}^{\bullet}$ in (A.6) is replaced with $\tilde{\mathrm{b}}_{\setminus 1}^{\bullet} \equiv [\mathrm{b}_1^{\bullet}]^{-1}\mathrm{b}_{\setminus 1}^{\bullet}$ and thus the ratio between the coefficients of the regressions of the VAR innovations onto the instrument delivers a scaled version of $\mathrm{B}^{\bullet}$. The unscaled $\mathrm{B}^{\bullet}$ is then recovered by noting that:

$$\mathrm{b}_1^{\bullet} = \sqrt{\Sigma_{1,1} - (\Sigma_{\setminus 1,1} - \tilde{\mathrm{b}}_{\setminus 1}^{\bullet}\Sigma_{1,1})'\Gamma^{-1}(\Sigma_{\setminus 1,1} - \tilde{\mathrm{b}}_{\setminus 1}^{\bullet}\Sigma_{1,1})},$$

where $\Gamma = \tilde{\mathrm{b}}_{\setminus 1}^{\bullet}\Sigma_{1,1}\tilde{\mathrm{b}}_{\setminus 1}^{\bullet\prime} - (\Sigma_{\setminus 1,1}\tilde{\mathrm{b}}_{\setminus 1}^{\bullet} + \tilde{\mathrm{b}}_{\setminus 1}^{\bullet}\Sigma_{\setminus 1,1}') + \Sigma_{\setminus 1,\setminus 1}$ and $\Sigma_{i,j}$ are appropriate partitions of Σ_u.

The coefficients in $\mathbf{A}^\star$, and thus $\mathrm{B}^\bullet$, can be recovered using $\mathbf{A}^\star = \hat{\mathbf{C}}\Lambda^{-1}\Psi$. In what follows I will show that a necessary condition for this procedure to deliver the coefficients in $\mathrm{B}^\bullet$, is that the proxy z_t be orthogonal to the history of y_t included in the VAR, that is, $\mathbb{E}[z_t\mathcal{Y}_t'] = 0$, as claimed in Section 1.

Using (A.7), and OLS estimates of $\mathbf{C}$ from (7):

$$\begin{aligned} \mathbf{A}^{\star\prime} = [\mathbf{A}\ \mathrm{B}^\bullet]' &= \Psi'\Lambda^{-1}\hat{\mathbf{C}}' \\ &= \Psi'\left[\Sigma_{\mathcal{Y}^+}^{-1}\left[\Sigma_{\mathcal{Y}^+} - \Sigma_\zeta\right]\right]^{-1}\Sigma_{\mathcal{Y}^+}^{-1}\Sigma_{\mathcal{Y}^+y}. \end{aligned} \tag{A.8}$$

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

$$\Psi' = \begin{bmatrix} \mathbb{I}_{np} & \mathbf{0} \\ \mathbf{0} & \Phi' \end{bmatrix} \quad \text{and} \quad \Sigma_{\mathcal{Y}^+} = \begin{bmatrix} \Sigma_{\mathcal{Y}} & \mathbf{0} \\ \mathbf{0} & \Sigma_z \end{bmatrix}, \tag{A.9}$$

where $\mathbf{0}$ denotes matrices of zeros of suitable dimensions. Equation (5) can thus be rewritten as:

$$\mathcal{Y}_t^+ \equiv \begin{bmatrix} \mathcal{Y}_t \\ z_t \end{bmatrix} = \begin{bmatrix} \mathbb{I}_{np} & \mathbf{0} \\ \mathbf{0} & \Phi \end{bmatrix}\begin{bmatrix} \mathcal{Y}_t \\ \varepsilon_t^\bullet \end{bmatrix} + \begin{bmatrix} \mathbf{0} \\ \nu_t \end{bmatrix}. \tag{5'}$$

After some algebra, plugging (A.9) into (A.8) yields:

$$\mathbf{A}^{\star\prime} = \begin{bmatrix} \mathbb{I}_{np} & \mathbf{0} \\ \mathbf{0} & \Phi' \end{bmatrix}\left[\begin{pmatrix} \Sigma_{\mathcal{Y}}^{-1} & \mathbf{0} \\ \mathbf{0} & \Sigma_z^{-1} \end{pmatrix}\begin{pmatrix} \Sigma_{\mathcal{Y}} & \mathbf{0} \\ \mathbf{0} & \Sigma_z - \Sigma_\nu \end{pmatrix}\right]^{-1}\left[\begin{pmatrix} \Sigma_{\mathcal{Y}}^{-1} & \mathbf{0} \\ \mathbf{0} & \Sigma_z^{-1} \end{pmatrix}\begin{pmatrix} \Sigma_{\mathcal{Y}y} \\ \Sigma_{zy} \end{pmatrix}\right]. \tag{A.10}$$

Due to the block diagonal structure of the elements in (A.10), 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:

$$\begin{aligned} \mathrm{B}^{\bullet\prime} &= \Phi'\left[\Sigma_z^{-1}[\Sigma_z - \Sigma_\nu]\right]^{-1}\Sigma_z^{-1}\Sigma_{zy} \\ &= \Phi'[\Sigma_z - \Sigma_\nu]^{-1}\Sigma_{zy} \\ &= \Phi'[\Phi\Phi']^{-1}\Sigma_{zy} = \Phi^{-1}\Sigma_{zy}, \end{aligned} \tag{A.11}$$

which is equivalent to (3).

B Monetary Policy Surprises from Financial Markets Instruments

B.1 US Raw Monetary Surprises

Sack (2004) discusses the technical procedure for the extraction of policy expectations from both Federal Funds and Eurodollar futures that are shown to be accurate predictors of the policy rate in Gürkaynak et al. (2006). Let $P_{FF}^{(\iota)}$ and $P_{ED}^{(\iota)}$ denote respectively the price of the FF and ED expiring on day ι of a given month m, and let N be the number of days in m, then:

$$P_{FF}^{(\iota)} = 100 - \frac{1}{N}\sum_{i=1}^{N} r_i; \tag{B.1}$$

$$P_{ED}^{(\iota)} = 100 - \$\text{lib}_{\iota}^{(\iota+90)}; \tag{B.2}$$

where r is the effective fed fund rate and $\$\text{lib}_{\iota}^{(\iota+90)}$ is the 3-month US Dollar-based Libor fixing on day ι. When expressed in rates at any time t, the equations above transform as follows:

$$FF_t^{(\iota)} = \mathbb{E}_t\left[\frac{1}{N}\sum_{i=1}^{N} r_i\right] + \xi_{FF,t}^{(\iota)}; \tag{B.3}$$

$$\begin{aligned} ED_t^{(\iota)} &= \mathbb{E}_t\left[\$\text{lib}_{\iota}^{(\iota+90)}\right] + \xi_{ED,t}^{(\iota)} \\ &= \mathbb{E}_t\left[\bar{\mathbf{r}}_{\iota}^{\iota+90}\right] + \mathbb{E}_t\left[\$\text{lib}_{\iota}^{(\iota+90)} - \bar{\mathbf{r}}_{\iota}^{\iota+90}\right] + \xi_{ED,t}^{(\iota)}. \end{aligned} \tag{B.4}$$

Where $\bar{\mathbf{r}}_{\iota}^{\iota+90}$ denotes the average rates over the 90 days (3 months) starting from day ι, i.e. $\bar{\mathbf{r}}_{\iota}^{\iota+90} \equiv \frac{1}{90}\sum_{i=1}^{90} r_{\iota+i}$. While the link between FF and r is direct, when dealing with EDs an additional step in which expectations about future Libor fixings are translated into expectations about the policy rates is required. The terms $\xi_{\cdot,t}^{(\iota)}$ denote (possibly time-varying) term/risk premia in both equations. In (B.4), the ED rate is expressed as a function of three terms: (i) the expectation of the short-term rate over the three-month period starting from the expiration of the contract – ι; (ii) 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 (iii) a residual risk premium which encompasses everything which is not explicitly associated to either (i) or (ii).

The construction of monetary surprises in the US is discussed in Kuttner (2001) for futures referring to the current month and daily data, and in Gürkaynak et al. (2005) for futures covering maturities which go out about 3.5 quarters and intraday quotes. Federal Fund futures (FF) 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 future just before ($-\Delta t$) the FOMC meeting, and r_t is the EFFR:

$$FF^{(0)}_{t-\Delta t} = \frac{n}{N}\mathbb{E}_{t-\Delta t}[r_{\tau\leq t}] + \frac{N-n}{N}\mathbb{E}_{t-\Delta t}[r_{\tau\geq t}] + \xi^{(0)}_{FF,t-\Delta t}. \tag{B.5}$$

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 $\xi^{(0)}_{FF,t-\Delta t}$ 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 contract if certain timing criteria are met (see Gürkaynak, 2005).

If $\tilde{r}_t$ denotes the target rate (i.e. policy rate) and $r_t = \tilde{r}_t + \epsilon_t$, where ϵ_t is some targeting error which is assumed to be unchanged within the Δt time frame, the *raw* monetary policy surprise – $mps^{(0)}_t$ – can be computed as:

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

Gürkaynak et al. (2005) assume that the latter term in the equation above is zero, *de facto* implying that it takes longer than the Δt 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 et al. (2005) to the month in which the corresponding meeting was scheduled to happen (red line).[21] 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.

[21]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.

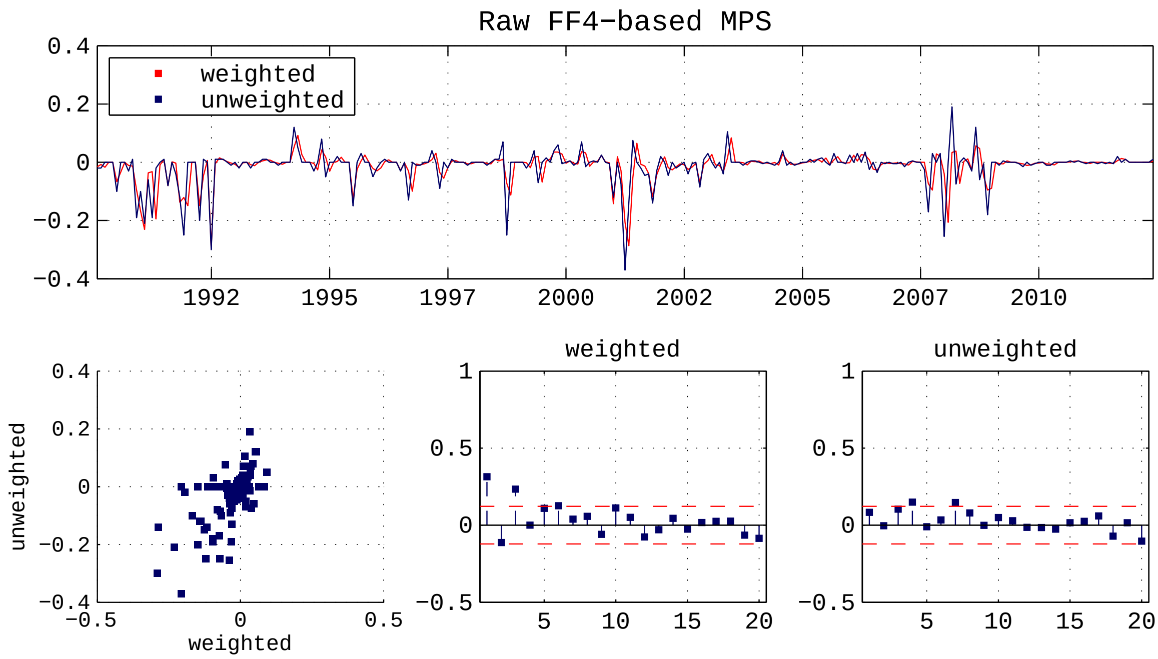

FIGURE B.1: Raw FF4-based monetary surprises at monthly frequency. The weighted series is from Gertler and Karadi (2015), while the unweighted surprise is constructed as the sum of daily surprises in Gürkaynak et al. (2005). In the bottom panel, from left to right, the different information content in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

The weighted series exhibits some degree of autocorrelation, also noted in Ramey (2015). 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 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 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 the Bank Rate, to which, however, no financial market instrument is directly linked. The closest alternative are Overnight Indexed Swap (OIS) rates. 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 the Bank Rate closely, furthermore, since these contracts are constructed in a way that minimises credit risk, the implied path of SONIA rates 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.[22] 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 based on the (USD) 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 P_{SS}^{ι} denote the price of a Short Sterling future expiring on day ι, we have that

$$P_{SS}^{(\iota)} = 100 - \text{£lib}_{\iota}^{(\iota+90)}, \tag{B.7}$$

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

$$\begin{aligned} SS_t^{(\iota)} &= \mathbb{E}_t\left[\text{£lib}_{\iota}^{(\iota+90)}\right] + \xi_{SS,t}^{(\iota)}, \\ &= \mathbb{E}_t\left[\bar{\mathbf{r}}_{\iota}^{\iota+90}\right] + \mathbb{E}_t\left[\text{£lib}_{\iota}^{(\iota+90)} - \bar{\mathbf{r}}_{\iota}^{\iota+90}\right] + \xi_{SS,t}^{(\iota)}, \end{aligned} \tag{B.8}$$

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

The rates involved in (B.8) and a detail on the time variation of the LIBOR-OIS (LOIS)

[22]The quality of market-based policy path forecasts, including those derived from SS contracts, is discussed in Joyce et al. (2008). The exercise is similar in spirit to Gürkaynak et al. (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.

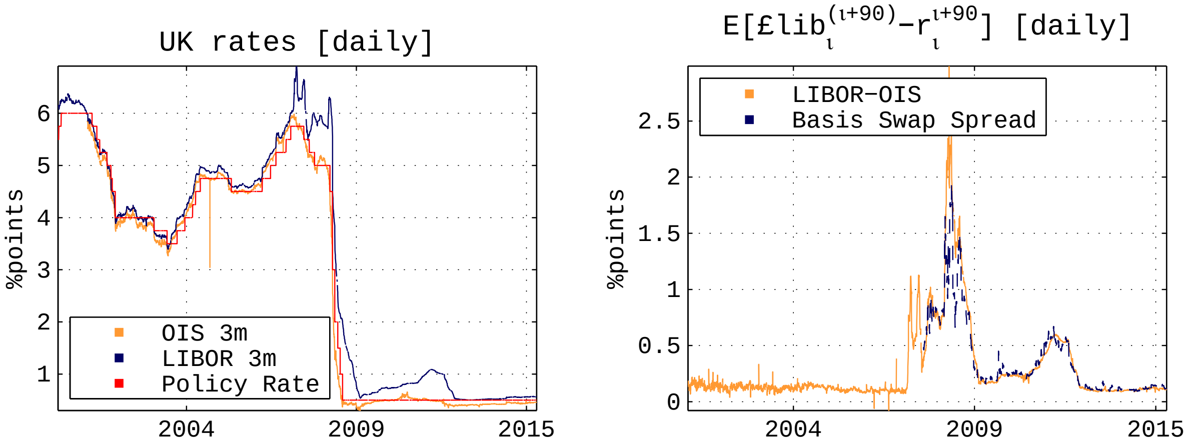

FIGURE B.2: [LEFT] Relevant interest rates for Short Sterling futures rates decomposition. [RIGHT] LIBOR-OIS (LOIS) spreads obtained as the difference between the 3-month Sterling LIBOR and the 3-month OIS curve, and from basis swaps (front contract, Basis Swap Spread). All rates are at daily frequency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

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, LIBOR rates, 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 considered as one of the key measures of risk premium, the LOIS spread drew relatively little attention in the years preceding the onset of the 2007 financial 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. Starting from 2008, however, raising doubts about financial institutions’ solvency and concerns relative to market liquidity induced a rise in LIBOR rates which made the spread jump at unprecedented levels. As the LOIS 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 directly 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 $\iota = 0$ in $\mathbb{E}_t\left[\text{£lib}_\iota^{(\iota+90)} - \bar{\mathbf{r}}_\iota^{\iota+90}\right]$.

Let $BS_t^{(\iota)}$ 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:

$$\begin{aligned} mps_t^{(\iota)} &= \left[SS_t^{(\iota)} - SS_{t-\Delta t}^{(\iota)}\right] - \left[BS_t^{(\iota)} - BS_{t-\Delta t}^{(\iota)}\right], \\ &= \left[\mathbb{E}_t\left[\bar{\mathbf{r}}_\iota^{\iota+90}\right] - \mathbb{E}_{t-\Delta t}\left[\bar{\mathbf{r}}_\iota^{\iota+90}\right]\right] + \left[\xi_t^{(\iota)} - \xi_{t-\Delta t}^{(\iota)}\right]. \end{aligned} \tag{B.9}$$

Figure B.3 plots the monthly surprises in the first Short Sterling future 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.[23] To construct the raw monetary surprise, at any date in the sample I use the next expiring SS future, or front contract (SS1). Because liquidity in these markets tends to vanish when the expiration date approaches, if the MPC date falls in the vicinity of the expiry date, I switch to the next contract. 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 time windows tightly surrounding the policy event.[24] For the construction of the monthly surprise I 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 IR are contemporaneous to a number of relevant data releases that are also likely to substantially influence markets.[25] 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 the Labour Force Survey data and statistics on Money and Lending activities

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

[24] I reconstruct the historical set of policy rate decisions dates and times and the decision that resulted from the committee meetings using Bloomberg. The raw monetary surprises are computed by measuring changes in the first Short Sterling future contract rates within a narrow 30-minute window surrounding the event. A different strategy is adopted in case of the release of the Inflation Report: due to the press conference associated to the release lasting a full hour, I allow some more flexibility in this case by employing a 90-minute window. Raw intraday data are from Thomson Reuters Tick History Database.

[25] 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).

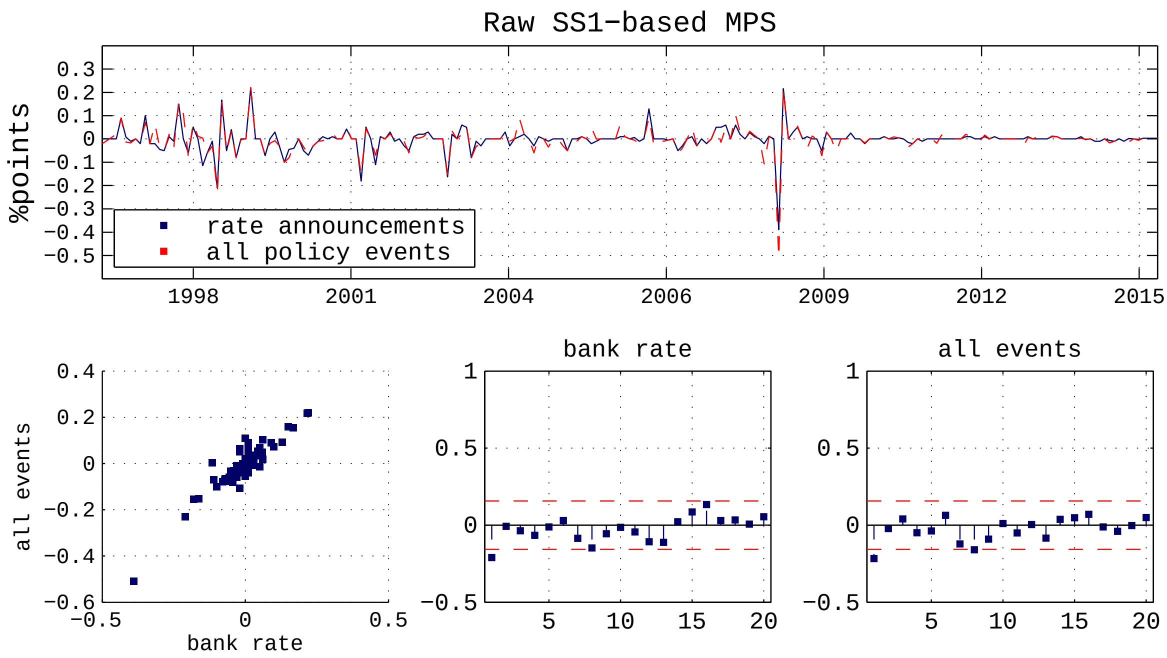

FIGURE B.3: Raw SS1-based monetary surprise at monthly frequency. Responses are reported around Bank Rate announcements only (blue line) and when also minutes and releases of the Inflation Report are taken into account (red dotted line). All surprises control for data releases contemporaneous to the 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.

and, in some instances, even GDP figures. To account for these interferences, in all cases I 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 et al. (2014).

The top panel and the bottom left subplot of Figure B.3 reveal that while there are 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 Bank of England'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 zero lower bound period (post 2009) is removed from the analysis.

C Additional Charts

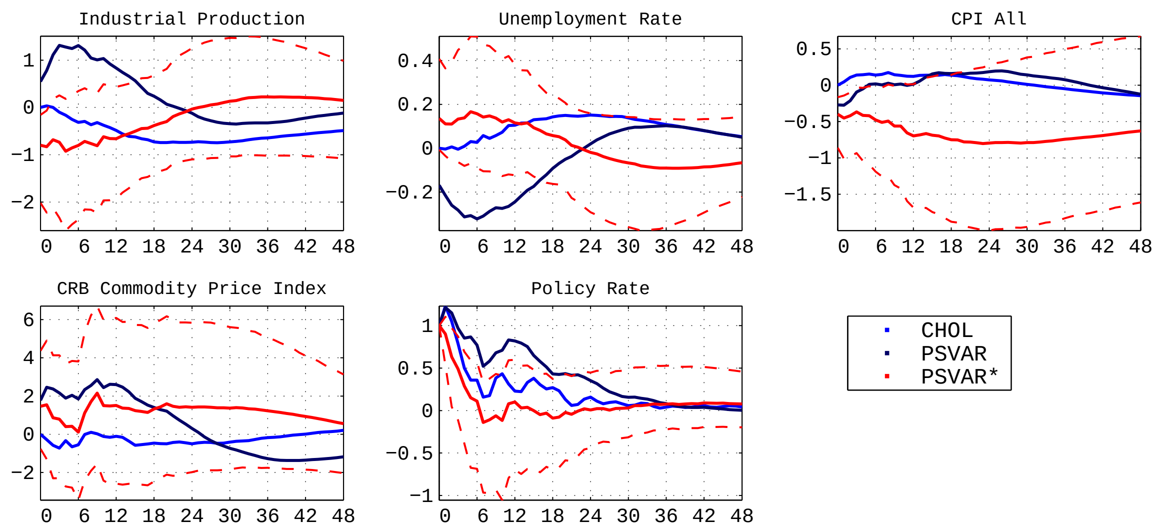

Figure C.1: The chart compares impulse responses to a monetary policy shock obtained estimating a VAR(12) over the sample 1969:1 - 2007:12 and using different identification schemes. Light blue lines are for the recursive identification scheme with the Effective Fed Fund Rate ordered last (CHOL). Dark blue lines are obtained when the shock is identified using the weighted 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.

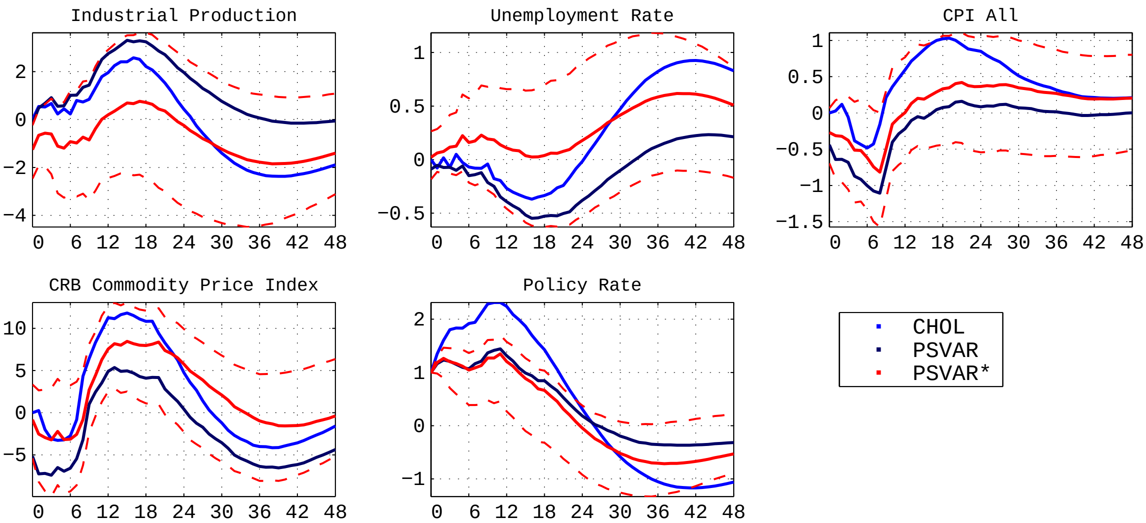

Figure C.2: The chart compares impulse responses to a monetary policy shock obtained estimating a VAR(12) over the sample 1990:1 - 2014:12 and using different identification schemes. Light blue lines are for the recursive identification scheme with the Effective Fed Fund Rate ordered last (CHOL). Dark blue lines are obtained when the shock is identified using the weighted 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.

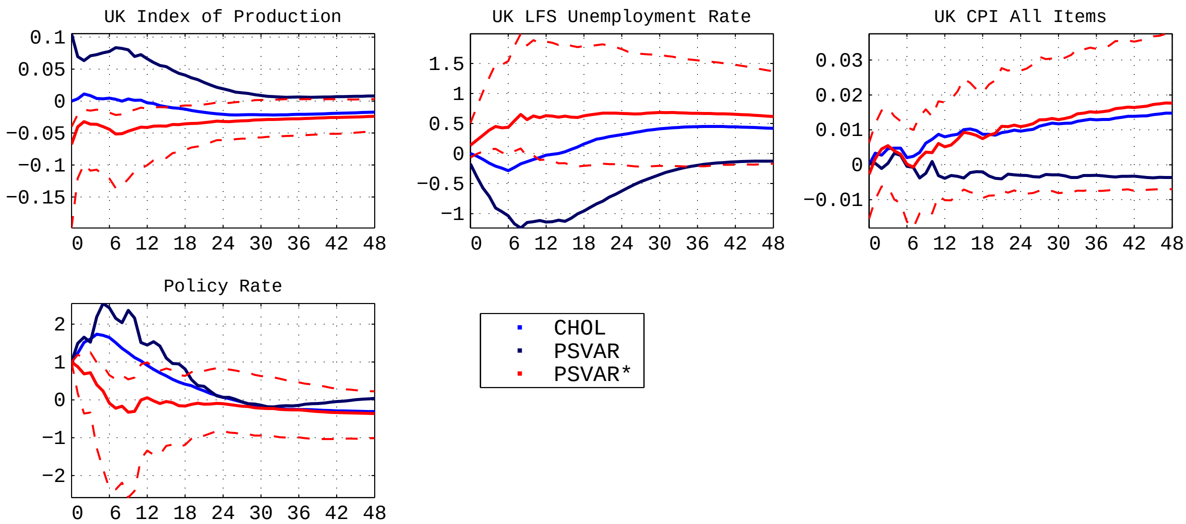

FIGURE C.3: The chart compares impulse responses to a monetary policy shock obtained estimating a VAR(12) over the sample 1990:1 - 2014:12 and using different identification schemes. Light blue lines are for the recursive identification scheme with the Bank Rate ordered last (CHOL). Dark blue lines are 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.