Monetary Policy Transmission in an Open Economy: New Data and Evidence from the United Kingdom[☆]

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August 2, 2016

Abstract

This paper constructs a new series of monetary policy surprises for the United Kingdom and estimates their effects on macroeconomic and financial variables, employing a high-frequency identification procedure. First, using local projections methods, we find that monetary policy has persistent effects on real interest rates and breakeven inflation. Second, employing our series of surprises as an instrument in a SVAR, we show that monetary policy affects economic activity, prices, the exchange rate, exports, and imports. Finally, we implement a test of overidentifying restrictions, which exploits the availability of the narrative series of monetary policy shocks computed by Cloyne and Huertgen (2014), and find no evidence that either set of shocks contains any 'endogenous' response to macroeconomic variables.

Keywords: Monetary Policy Transmission, External Instrument, High-Frequency Identification, Structural VAR, Local Projections.

JEL Codes: E31, E32, E43, E44, E52, E58.

[☆]We would like to thank Piergiorgio Alessandri, Regis Barnichon, James Cloyne, Andrea Gazzani, Wouter den Haan, Jean Imbs, Oscar Jorda, Sophocles Mavroeidis, Michael McMahon, Silvia Miranda-Agrippino, Emi Nakamura, Evi Pappa, Valerie Ramey, Morten Ravn, Ricardo Reis, Paolo Surico, Jonathan Wright, and seminar participants at the Bank of England, Bank of Italy, BSGE Summer Forum, ECB, European University Institute, and University of Manchester for helpful comments and suggestions. The views expressed in this paper are solely those of the authors and should not be taken to represent those of the Bank of England.

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1 Introduction

The central banks of most industrialized countries use interest rates to stabilize economic activity and inflation. To do this well, they need to know how changes in the policy instrument affect the economy. However, the fact of using these instruments to stabilize the economy makes it difficult to disentangle their effects from their determinants. A number of authors have turned to high-frequency financial market data to identify monetary policy innovations unrelated to the state of the economy(see, for example, Kuttner, 2001, Gurkaynak et al., 2005a, 2007). In essence, they exploit the fact that most of the changes in short-term interest rates futures occurring in a window of a few minutes around policy announcements relate exclusively to monetary policy news.[1]

This paper identifies monetary policy surprises in the United Kingdom (UK) and assesses their effects on financial and macroeconomic variables with data from 1993 to 2015, during which period UK monetary policy was operating under an inflation targeting regime. We contribute to the literature on the effects of monetary policy in three main ways. First, we provide what are to our knowledge the first published estimates of high-frequency monetary policy surprises for the UK,[2] and compare their effects on the macroeconomy with other studies. Second, we employ local projection methods to show that these surprises have persistent effects on financial markets, beyond the day on which they occur. Third, we combine our surprises with overlapping narrative estimates of UK monetary policy innovations (constructed by Cloyne and Huertgen (2014) following Romer and Romer (2004)'s approach) in a test of overidentifying restrictions and find no evidence that either set of 'shocks' is endogenous to the macroeconomy.

We follow closely the methods employed in the literature, applying them as faithfully as possible to the UK, but also implement a number of modifications and extensions.[3] First, the institutional framework for deciding and communicating monetary policy, with separate releases of interest rate decisions and *Inflation Report*, enables us to enlarge the set of events under consideration. Second, we perform estimations at daily frequency employing local projection methods (Jorda, 2005), as well as at monthly frequency using structural Vector Autoregressions (VARs), to demonstrate that these surprises have persistent effects on financial and macroeconomic variables. Unlike VARs, local projection methods do not force our surprises to inherit the average persistence of any disturbance to the interest rate. Third, we enlarge the set of such variables under consideration, considering both daily asset prices and monthly macroeconomic and financial data.

[1] This identification approach based on high frequency data is not new, and dates back to the work by Bagliano and Favero (1999) and Cochrane and Piazzesi (2002). See also Faust et al. (2004) Gurkaynak et al. (2005b), Faust et al. (2007), and Bredin et al. (2009) for other examples.

[2] While writing this paper we became aware of a paper by Miranda-Agrippino (2015) that, using a similar methodology, derives a series of monetary policy surprises for the UK.

[3] See, among others, Nakamura and Steinsson (2013), Barakchian and Crowe (2013), Rogers et al. (2014), Gertler and Karadi (2015), Hanson and Stein (2015), Gilchrist et al. (2015), and Rogers et al. (2015).

Our results show that the monetary surprises we construct have statistically and economically significant effects on interest rates along the nominal spot and forward yield curves. Within this, tightenings tend to raise forward real interest rates and to lower breakeven inflation rate. Employing local projection methods to estimate the impulse responses at trading-day frequency, we find that these effects persist for at least a month after the shock — thus complementing the results of Rogers et al. (2014). Turning to the macroeconomic effects of our surprises, we find that monetary policy tightenings raise unemployment and corporate lending spreads, strengthen the exchange rate, and reduce trade volumes, stock prices and CPI. In sum, our paper advances novel empirical evidence on the monetary transmission mechanism while also confirming some standard results, on monetary non-neutrality and on the credit channel of monetary policy, using a novel data set for the UK. In particular, our effects on financial and macroeconomic variables are comparable to previous studies for the United States and for the UK (for example Nakamura and Steinsson (2013) and Gertler and Karadi (2015); Mountford (2005) and Cloyne and Huertgen (2014)).

The statistical inference that we conduct is reliable under two assumptions: (i) the absence of 'background noise' in our measure of monetary policy surprises and (ii) their exogeneity to developments in the macroeconomy. Since the monetary policy surprises are measured with error, the first assumption should be interpreted as saying that the noise-to-signal ratio is vanishingly small. If this assumption is violated, our OLS parameter estimates will suffer from attenuation bias. The second assumption rules out the possibility that other non-monetary news might affect our monetary policy surprises during the window we consider around policy announcements. If this assumption is violated, the monetary policy surprises can simply measure the central bank's response to its private information about the future evolution of the economy, therefore leading to bias in the estimates.

We test assumption (i) by comparing our OLS estimates (which are only consistent under the assumption that there is no noise in our measured monetary surprises) with the 'identification by heteroskedasticity' estimator proposed by Rigobon (2003) and Rigobon (2003), Rigobon and Sack (2004). We test assumption (ii) by exploiting a series that explicitly controls for the information set of the central bank — the narrative measure of monetary policy innovations of Cloyne and Huertgen (2014) — in a test of overidentifying restrictions. Our results show that both assumptions are satisfied.

The remainder of this paper is structured as follows. Section 2 reviews the framework for setting and communicating monetary policy in the UK and describes how we construct the monetary surprises. Section 3 provides evidence on the impact of monetary policy on 'high frequency' financial variables using local projection methods, and on the absence of background noise in our measure of monetary policy surprises. Section 4 shows their impact on macroeconomic and financial variables in a structural VAR. Section 5 tests the validity of our instrument through overidentifying restrictions. Section 6 concludes.

2 A New Series of Monetary Policy Surprises for the United Kingdom

In this section we derive a new series of monetary policy surprises for the UK, closely following the methodology originally proposed by Kuttner (2001) and Gurkaynak et al. (2005a).

To preview our method, we construct a new dataset using intra-daily data that captures changes in expectations about the monetary policy stance in the UK for every monetary policy "*event*" since operational independence was granted to the Bank of England in 1997. We use the term event to refer to a time at which a policy decision, or change in policy stance by the Monetary Policy Committee of the Bank of England (MPC), was communicated to financial markets. We proxy the changes in expectations about the monetary policy stance by computing the change in interest rate futures (at different maturities) in a thirty-minute window around every monetary policy event. The short time horizon over which these surprises are computed allows us to isolate the monetary policy news from other types of news that can also shift the yield curve.

In what follows we first review the monetary policy framework in the UK and how we compile our set of monetary policy events for the UK. We then describe how we construct the monetary surprises.

Monetary policy events. The UK adopted an inflation target as its nominal anchor in September 1992, following its exit from the Exchange Rate Mechanism of the European Union. To begin with, the Chancellor of the Exchequer (the UK Finance Minister) retained control of the policy instrument (the 'Bank rate') which was adjusted periodically in consultation with the Governor of the Bank of England to meet the inflation target.[4] In May 1997 the Bank of England was given operational independence, i.e. the ability to set monetary policy so as to achieve an inflation target decided by the Government.

Since then, the MPC has held monthly policy deliberations that led to policy announcements and the release of minutes approximately two weeks later. The MPC's view of the economic and financial outlook is also communicated in quarterly *Inflation Reports*, released between the policy decision and the minutes relating to the February, May, August and November MPC meetings. This means that, over the period we study, we have 28 scheduled events of monetary news in each year. However, in our baseline we drop the 12 events each year associated with the publication of meeting minutes, as these often coincided with the release of important macroeconomic data

[4]Bank rate is the rate of interest that the Bank of England pays on reserve balances held by commercial banks.

(specifically, with Labour Market Statistics).[5] This leaves us with 291 monetary policy events: 218 MPC meetings and 73 releases of the Inflation Report.

Since June 1997, the MPC has set Bank rate to achieve its inflation target. A liquid contract based on Bank rate would therefore be the most appropriate contract to compute the surprises. However, and unlike the case of Fed Funds for the US, there is no futures market based on this rate in the UK. Considering the length of the available set of contracts and their market size, the Sterling futures contracts are the most appropriate ones for measuring the expected evolution of interest rates. These contracts are settled based on the 3-month London Interbank Offered Rate (Libor).[6] In particular, in a given year, there are four delivery dates at the end of the following months: March, June, September, and December.[7,8]

Monetary policy surprises. We measure our interest-rate surprises through intra-daily changes in the price of 3-month sterling futures contracts. The price of these contracts is quoted as 100 minus the Libor rate for three-month sterling deposits set on the last trading day of the month in question. So, if investors are risk neutral, the price of a 3-month Sterling future expiring on date h on a given day t is related to expected future interest rates as follows:

$$P_t^h = 100 - \mathbb{E}_t \left[i_h^{(h+90)} \right], \tag{1}$$

where P_t^h denotes the current price for a contract that matures on day h and $\mathbb{E}_t \left[i_h^{(h+90)} \right]$ denotes the expected value (on day t) of the 3-month (i.e., $h + 90$ days) Libor at time h. We define a monetary policy surprise as the change in the price of the 3-month Sterling future in a 30 minutes window around a monetary policy event:

$$s_t = -\left(P_{t,\tau+20}^h - P_{t,\tau-10}^h \right), \tag{2}$$

where t, τ denotes the exact time (in minutes) during day t when a monetary policy event occurred; and P^h denotes the price of a contract that expires on date h. We use the minus in front of the price change to express the surprise such that a positive number means an increase in the expected interest rate implied by P^h. Then, from equations (1) and (2), we can define

[5] Note, however, that we include these events in a robustness exercise and our results are virtually unchanged.

[6] A better, alternative contract would be the Sterling Overnight Indexed Swap (OIS), as suggested by Joyce et al. (2008). This contract is based on the Sterling Overnight Index Average (SONIA), which carries a lower risk premium than Libor. However, OIS contracts at intradaily frequency are available only from 2008 and —since they are traded in OTC markets— the data on intraday transactions are not always available. Appendix A describes in detail all the contracts available for the UK and their characteristics.

[7] For example, on January 1^{st} four contracts are available. These contracts mature at the end March, June, September, and December, respectively. Strictly speaking, there are two additional contracts that expire at the end of January and at the end of February. However, these contracts are very illiquid, therefore in our analysis we only consider the main four contracts mentioned above.

[8] One disadvantage of these contracts compared to the Fed Funds Futures is that the latter has a monthly delivery date and is based on the 30 day average of Fed Funds rate. Appendix provides A more information about these contracts.

the surprise in terms of the expected rate:

$$s_t = \mathbb{E}_{(t,\tau+20)}\left[i_h^{(h+90)}\right] - \mathbb{E}_{(t,\tau-10)}\left[i_h^{(h+90)}\right], \tag{3}$$

where $\mathbb{E}_{(t,\tau+20)}\left[i_h^{(h+90)}\right]$ denotes the expected value of the 3-month Libor at time h, 20 minutes after the monetary policy event that occurred on day t at time τ (i.e., $t, \tau + 20$). We think of these measured surprises, s_t, as noisy signals of the 'true' monetary news $\epsilon_{1,t}^{mp}$ associated with the policy event in question:

$$s_t = \epsilon_{1,t}^{mp} + \eta_t, \tag{4}$$

where the term η_t (which is orthogonal to $\epsilon_{1,t}^{mp}$) represents the noise component of the measurement. We also allow this underlying news (the one that we measure with our surprises) to be, in turn, a subset of the universe of monetary shocks ϵ_t^{mp} that occur within a given period. As purely illustrative example consider:

$$\epsilon_t^{mp} = \epsilon_{1,t}^{mp} + \epsilon_{2,t}^{mp} + \epsilon_{3,t}^{mp}, \tag{5}$$

where — if $\epsilon_{1,t}^{mp}$ are the shocks associated with policy decisions and the Inflation Report — $\{\epsilon_2^{mp}, \epsilon_3^{mp}\}$ could be those shocks associated with speeches by members of the MPC, changes in the membership of the MPC (and the associated change in attitudes towards inflation stabilization), or changes in the mandate of the MPC itself.

Underlying assumptions. The monetary policy surprises s_t can be then used directly in simple regressions to compute consistent estimates of the effect of monetary shocks on financial markets and the economy only if:

(i) the background noise is uncorrelated with developments in the macroeconomy, namely $E[\eta \mid x] = 0$ (where x is the state of the macroeconomy);

(ii) the background noise is negligible, i.e. $E[\eta^2] \simeq 0$.

Further conditions may apply depending on what estimator is being used, and will be examined in detail below. We next discuss how our high-frequency procedure is designed to ensure that these two assumptions are satisfied, and why the procedure might fail.

Starting with assumption (i), the choice of a tight window around the monetary policy event helps to isolate monetary policy news from other types of news. As noted above, we drop those events that coincided with data releases such that no major macroeconomic data releases occurred during our sample windows. One further possibility that would undermine our procedure is that policy events contain significant information about the macroeconomic determinants of monetary policy, as well as news about policy conditional on those determinants. This may be the case if the central bank is perceived to have private information about the outlook for the economy, resulting from privileged access to data or a superior ability to process it. For example, a surprise tightening of monetary policy could therefore be taken to indicate an improvement in the outlook for the macroeconomy. In this case, our monetary surprises will be correlated with non-monetary news about the economy; and the estimated impact of monetary surprises on macroeconomic and financial variables will be biased.

In this regard, Independent Evaluation Office (2015) reviews the accuracy of the Bank of England's macroeconomic forecasts and finds mixed results. *Inter alia*, this study finds that the Bank of England's two-year-ahead forecast errors for GDP, inflation and unemployment were correlated with data available when the forecast was made; and that, at the policy-relevant two-year horizon, private sector forecasts outperformed the Bank of England's. So there is little direct evidence that the MPC's forecasts contain significant incremental information about the determinants of monetary policy. To provide further evidence of the exogeneity of our measured monetary surprises, section 5 conducts a formal test of overidentifying restrictions using another series of monetary innovations that explicitly controls for the Bank of England's private information set. According to the test we cannot reject the null hypothesis that both sets are exogenous to the macroeconomy, against the alternative that at least one of them is not.

We turn now to assumption (ii). The presence of background noise in our measure of monetary surprises (i.e., $E[\eta^2] > 0$ even though $E[\eta \mid x] = 0$, where x is the state of the macroeconomy) can lead to a bias in the parameter estimates depending on whether we are thinking of our surprises as monetary innovations or merely instruments for them. *Prima facie*, the existence of such noise is likely: our surprises are derived from Libor contracts, an interbank rate that can contain significant premia. During the recent crisis, the spread between Libor and the overnight rate SONIA (which carries a much lower risk premium) fluctuated significantly.

Again, the choice of a tight window around the monetary policy event helps in this respect. But we address this concern by also considering an alternative future contract that carries a smaller risk premium (namely, the 3-month forward on the Sterling-US Dollar exchange rate) to which our results are robust. Moreover, in section 3 below, we provide a formal test that the scale of the noise is negligible, justifying their use as direct measures of shocks. And lastly, we provide IV estimates as the baseline when assessing their macroeconomic effects in section 4 and as a robustness check when looking at their effect on financial markets in section 3.

A new series of monetary policy surprises for the UK. Figure 1 displays the series of daily surprises computed using the second front contract of the 3-month Sterling future, i.e. the 3-to-6-month ahead expectation about the 3-month Libor.[9] Our series captures some of the

[9]The data for the monetary policy surprises is available at the authors' web sites. See https://sites.google.com/site/ambropo/CTV_MonPolTransmission_WP_CfM.xls.

main relevant monetary policy events in the period.[10]

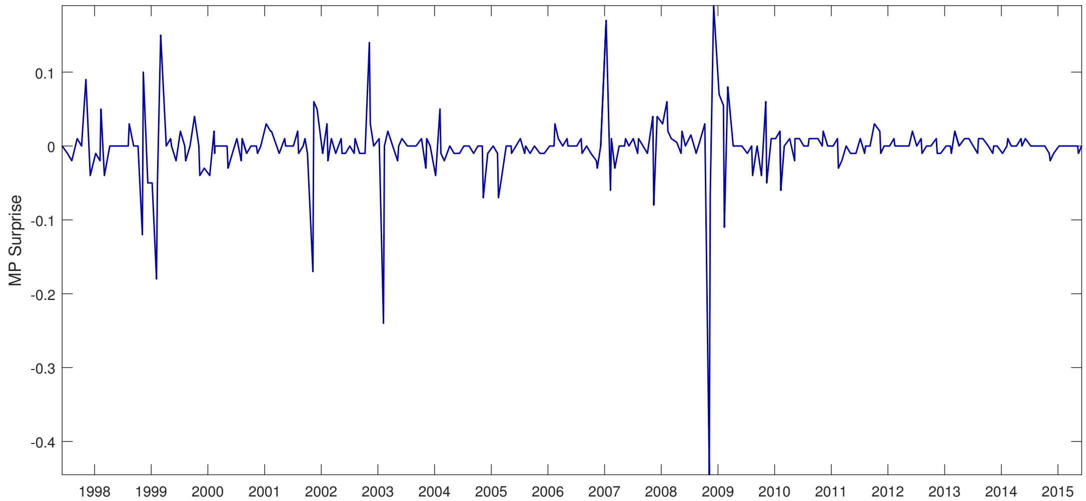

Figure 1 Daily Monetary Policy Surprises. *Note.* Daily monetary policy surprises computed using the second front contract of the 3-month Sterling future, i.e. the 3-to-6-month ahead expectation about the 3-month Libor. The surprises are computed using a 30 minutes window around the identified monetary policy events.

Figure 1 also shows that there is a clear change in the volatility of the monetary policy surprises after March 2009, when Bank rate reached 0.5 percent (i.e., the effective zero lower bound in the case of the UK). This raises the issue of whether short-term future contracts are appropriate to capture monetary policy surprises during the zero lower bound period. For this reason, for the purposes of robustness tests, we also compute monetary policy surprises using the fourth continuous contract of the 3-month Sterling future (i.e., the 9-month to 12-month ahead expectation of the 3-month Libor) and the 3-month forward exchange rate between the British Pound and the US Dollar, a measure that turns out to be highly correlated with more standard measures of monetary news based on the UK yield curve. The UK monetary events in our sample do not overlap with US ones, so this measure can be potentially useful to capture not only conventional monetary policy surprises but also 'unconventional' monetary policy surprises (such as forward guidance and quantitative easing announcements) that became the norm after the Bank rate reached its effective zero lower bound in March 2009.

3 The High Frequency Impact of Monetary Policy Surprises

In this section we provide the details of the methodology that we use to compute the high frequency impact of monetary policy surprises and present the main results. Then, we show

[10] Table D.1 in the Appendix reports the largest surprises identified using this contract and shows that they coincide with important monetary policy events.

that a key assumption underlying our methodology — the effective absence of background noise from our measure of monetary policy surprises — is satisfied and, accordingly, that the statistical inference that we conduct is reliable.

3.1 Methodology and Results

The goal of this section is to estimate the effect of the monetary policy news contained in scheduled MPC announcements and inflation reports on a wide range of 'high frequency' variables, i.e. financial variables that are available at daily frequency. To do that one could estimate the following regression equation:

$$\Delta y_t = \alpha + \beta \Delta i_t + \epsilon_t, \tag{6}$$

where Δy_t denotes the daily variation in a variable of interest (e.g., stock prices, exchange rates, nominal and real interest rates at different maturities) and Δi_t denotes the daily variation in an indicator of the stance of monetary policy, such as a short-term risk-free interest rate. The problem with the estimation of (6) is that Δy_t and Δi_t may be simultaneously responding to news that is not related to monetary policy. As Gurkaynak et al. (2005a) show, these problems are a source of concern not only in monthly or quarterly regressions, but even in daily regressions.

However, as discussed in the previous section, changes in expectations about future interest rates using a tight enough window around monetary events should be dominated by the information about monetary policy. On the assumption that the markets and the central bank have the same information about the determinants of monetary policy, any news that arrives in this short window about how policy is to be set must be about the actions of policy makers given the state of the economy, rather than the state of the economy itself. For econometric purposes, they can therefore be considered as 'exogenous' monetary surprises.[11] It is therefore possible to regress the variable of interest (i) directly on the monetary policy surprise, as it is typically done in the HFI literature; or (ii) on a given policy indicator using the monetary surprises (s_t) as instruments in a 2SLS regression.

It might also be the case that our measured policy surprises are essentially noise, or short-lived disturbances to market interest rates with no persistent effects on monetary or other macroeconomic aggregates. With this in mind, we extend the daily contemporaneous regressions that are typically estimated in the HFI literature by using local projection methods (see Jorda, 2005). Specifically, we estimate the following equation:

$$\Delta y_{t+h} = \alpha + \beta_h s_t + \sum_{j=1}^{J} \gamma_{j,h} \Delta y_{t-j} + \sum_{k=1}^{K} \delta_{k,h} x_{t-k} + \epsilon_t, \tag{7}$$

[11]We test this exogeneity assumption formally in section 5, using an overidentification restriction and the alternative series of monetary policy shocks from Cloyne and Huertgen (2014).

where $h = 0, ..., H$, H is equal to 20 trading days, and x are control variables. The coefficient β_h represents the average impact of a monetary policy surprise on the variable of interest h days after the shock hit. A purely contemporaneous regression would restrict $h = 0$. In our baseline we control for lagged values of the dependent variable, Δy_{t-j} and for lagged value of the policy variable, which in our application is the 1-year nominal gilt yield. We set $J = 5$ and $K = 4$ as suggested by the Akaike information criterion (AIC). If our surprises are exogenous then the inclusion of lags will not affect the probability limit of our estimator $\widehat{\beta_h}$, but will affect its standard error and the value it takes in finite samples.[12]

We estimate the impact of monetary policy surprises on (i) nominal spot and forward gilt yields at different maturities; (ii) forward real gilt yields as measured with the index-linked gilt curve, and (iii) forward breakeven inflation rate at different maturities as implied by the difference between these nominal and real yields. In the list of variables of interest we include stock prices, exchange rates, and financial market spreads. Where daily data are unavailable on certain days for particular series, as it is sometimes the case for the short end of the index-linked gilt curve, we drop those days from our sample for that series only.

In what follows, we describe the results we obtain for each of these sets of variables. Figures 2, 3, and 4 report the point estimates of equation (7), together with 90 and 95 percent confidence intervals. In each case the independent variable is the monetary policy surprise (s_t), while the change in the dependent variable (Δy_{t+h}) is measured over a one-day window at different horizons h. We cumulate IRFs calculated in differences to get cumulative changes in levels over the horizon in question.

Our sample runs from 1997:6 to 2015:5, therefore including both the global financial crisis and its aftermath, i.e. a period where short-term interest rates — the 'typical' monetary policy indicator — reached the zero lower bound. We compute our monetary policy surprises using the second front contract on the 3-month Libor.[13] But our results are robust to (i) using other future contracts at different maturities; and (ii) not including the period where monetary policy was constrained by the zero lower bound.[14]

Nominal interest rates. Figure 2 reports the effects of the monetary policy surprise on nominal gilt yields, both spot rates (upper panel) and forward rates (lower panel). The scale of the monetary policy surprise is arbitrary, so we rescale all impulse response functions (IRFs)

[12] As a robustness check, we run a more conservative specification of equation (7), where we do not include any lags of Δy_t and x_t as a control variable; and a specification where we set $J = 3$ and $K = 2$ as suggested by the Bayesian information criterion (BIC). The results (not reported here, but available from the authors upon request) do not display any significant difference.

[13] We use this future contract because it is the one that displays the higher F-Statistic to explain the daily changes in the 1-year gilt yield.

[14] We report a set of robustness checks in an Online Appendix available on the authors' web sites. See https://sites.google.com/site/ambropo/CTV_MonPolTransmission_OnlineAppendix.pdf.

such that the effect of s_t on the 1-year nominal gilt yield is equal to 25 basis points (upper panel, top left chart), with the standard errors and confidence intervals scaled accordingly.

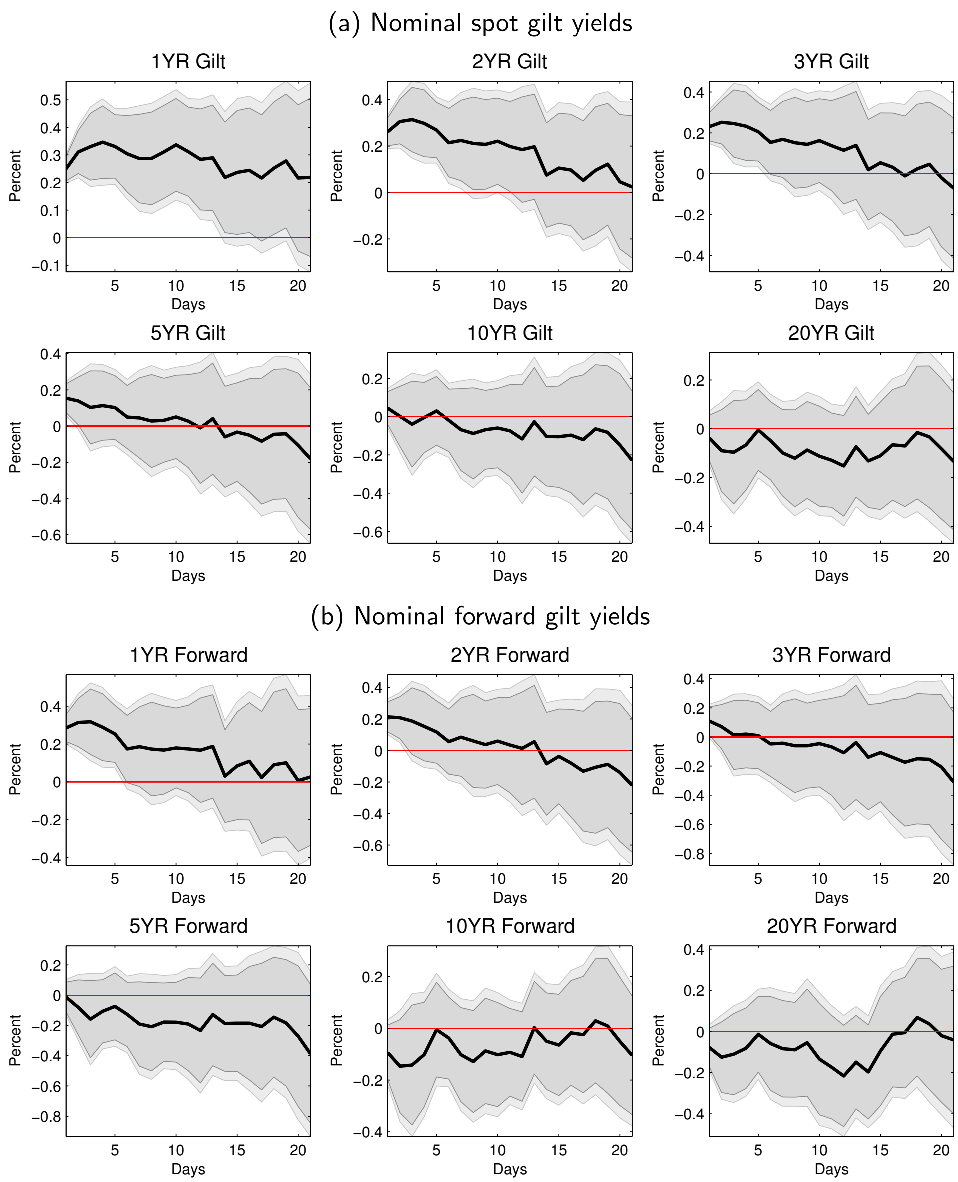

Figure 2 Response of Nominal Interest Rates to the Monetary Policy Surprises. *Note.* Each panel reports the cumulative results from a separate OLS regression as in equation (7). The dependent variable in each regression is the one day change in the variable stated in the panel title. The independent variable is the monetary policy surprise (s_t), computed using the second front contract of the 3-month Sterling future. The sample period is 1997:6 to 2015:5. The solid line and shaded areas report the mean, 90%, and 95% confidence intervals computed using bootstrap with 1,000 replications.

The upper panel shows that the monetary policy surprise has a statistically significant impact on gilt yields. Furthermore, this impact is persistent, statistically different from zero and slightly decreasing over the 20-trading-day horizon under consideration. As we move along the yield curve to longer maturities, the impact of the monetary policy surprise on yields generally becomes weaker and statistically less significant. So, the monetary policy surprises that we measure have an appreciable and persistent impact along a broad swathe of the yield curve.

The bottom panel depicts the decomposition of the impact on the spot yield curve into its effects on forward rates at different horizons. The charts in the lower panel of Figure 2 suggest that the impact on nominal instantaneous forward rates at one-year and two-year horizons is almost identical to that on the one-year spot rate. So the impact of monetary policy surprises on the expected level of short-term interest rates is almost flat for at least two years, suggesting that the markets view policy shocks as highly persistent. Consistent with our findings on the spot curve, as we move further along the forward curve the impact on the instantaneous forward rates falls towards and eventually beyond zero. The effect on 10-year and 20-year instantaneous forward rates is negative but not statistically significant. In summary, our policy surprises have large impact on expected short-term interest rates and then gradually decline at longer horizons.

Real interest rates and inflation. Figure 3 reports the effects of the monetary policy surprise on real forward gilt yields and breakeven inflation rate (left and right panel, respectively), the implied inflation over commensurate horizons obtained from index-linked gilt yields. Index-linked government bonds are not consistently available at short maturities over our sample period, limiting what we can reliably say about the effects of our surprises on the near end of the real government liability curve. So our charts begin with 3-year maturities.

In order to interpret the IRFs correctly, it is necessary to provide some institutional context for the measurement of breakeven inflation through index-linked government securities in the UK. The consumer price index to which UK index-linked gilts are indexed is the General Index of Retail Prices in the UK (RPI).[15] This index includes an estimate of owner-occupier housing costs, which depend, *inter alia*, on mortgage interest rates. The current inflation target of the Bank of England is set in terms of the Consumer Prices Index (CPI), and prior to that in terms of the RPIX, indices that exclude the cost of owner-occupied housing. The target was defined in these terms precisely to avoid the direct effect of the mortgage rates on the target. Our market-based measures of the impulse response of breakeven inflation and real interest rates as would be measured with the CPI will tend to be biased, upwards and downwards respectively, by this effect.

With this in mind, we now turn to describing the estimated effect of monetary policy surprises on real interest rates and breakeven inflation. Panel (a) of Figure 3 shows that real interest rates rise slightly on impact at the three-year maturity, while keeping the other real rates unchanged. This implies that the central bank can have an effect on real interest rates at medium to long

[15]See http://www.dmo.gov.uk/index.aspx?page=gilts/indexlinked for further details.

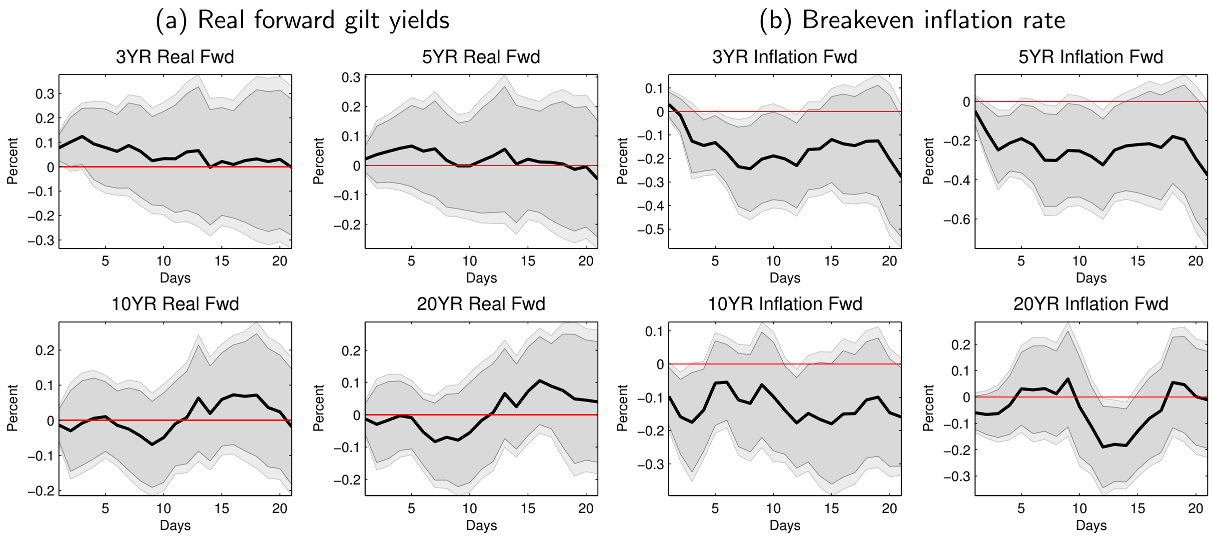

Figure 3 RESPONSE OF REAL INTEREST RATES AND INFLATION TO THE MONETARY POLICY SURPRISES. *Note.* Each panel reports the cumulative results from a separate OLS regression as in equation (7). The dependent variable in each regression is the one day change in the variable stated in the panel title. The independent variable is the monetary policy surprise (s_t), computed using the second front contract of the 3-month Sterling future. The sample period is 1997:6 to 2015:5. The solid line and shaded areas report the mean, 90%, and 95% confidence intervals computed using bootstrap with 1,000 replications.

horizons, a result that echoes the findings in Nakamura and Steinsson (2013) for the United States. Furthermore, this effect is large enough to outweigh the bias imparted by the mechanical effect on RPI explained above.

Panel (b) of Figure 3 shows the effect of the monetary surprises on the breakeven inflation curve. The three, five and ten years rates decline significantly, to the tune of $0.1 - 0.3$ percentage points. Note here that breakeven inflation rates capture both inflation expectations and a risk compensation term. If the risk compensation term responds to monetary policy tightening (similarly to what corporate risk premia do, as we show below), then the impact of the monetary policy surprises on inflation expectations could be larger than the OLS estimates reported above. There is no significant impact at the 20-year forward horizon. Nevertheless, these estimates suggest that monetary policy surprises are able to affect breakeven inflation several years into the future.

FX, equities, and corporate spreads. Figure 4 reports the effects of the monetary policy surprise on some selected exchange rates vis-a-vis Sterling, the FTSE index and a measure of corporate (investment grade) credit spreads.[16] We find that a surprise monetary policy tightening causes the pound Sterling to appreciate vis-a-vis the US Dollar by about 0.5 percent on impact, with the impact rising to a peak of more than 2 per cent. The effect on the exchange rates of Sterling against the euro and the yen is similar, even though slightly smaller in magnitude

[16]Exchange rates are defined such that a rise means an appreciation

and less statistically significant than in the case of the US Dollar. This is reflected in the response of the Exchange Rate Index (ERI), which appreciates by slightly less than 1.5 percent.

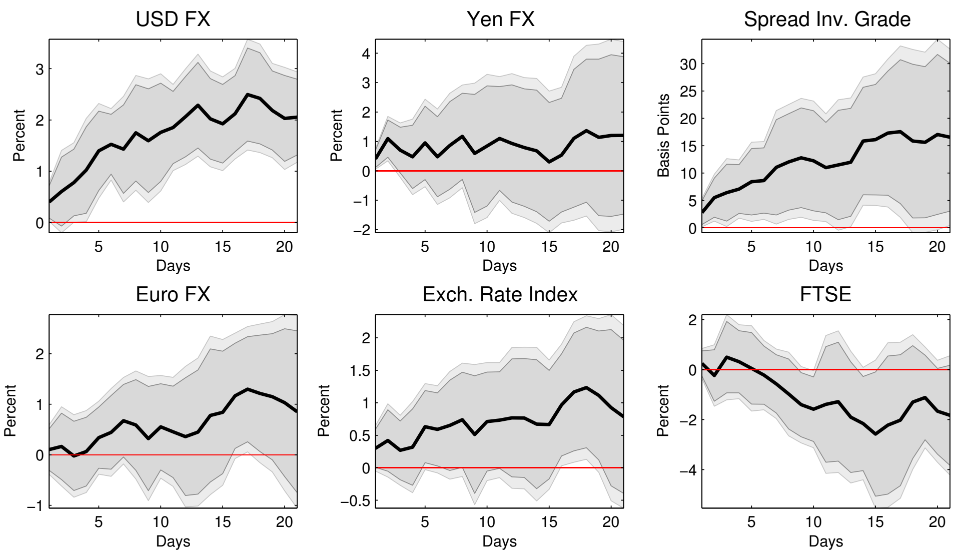

Figure 4 Response of Other Financial Market Variables to the Monetary Policy Surprises. *Note*. Each panel reports the cumulative results from a separate OLS regression as in equation (7). The dependent variable in each regression is the one day change in the variable stated in the panel title. The independent variable is the monetary policy surprise (s_t), computed using the second front contract of the 3-month Sterling future. The sample period is 1997:6 to 2015:5. The solid line and shaded areas report the mean, 90% and 95% confidence intervals computed using bootstrap with 1,000 replications.

The response of equity markets is muted on impact, but the FTSE tends to fall over the horizon considered in our impulse responses getting to a low of about 2 percent one month after the shock hit. Finally, a measure of corporate credit spread (for investment grade firms) tend to increase in the face of a monetary policy tightening, with a maximum impact of about 20 basis points at the end of the horizon considered. The response of the credit spread corroborates one important result put forth by Gertler and Karadi (2015) in their analysis. Specifically, our results point to the presence of a "credit channel" of monetary policy (Bernanke and Gertler, 1995), according to which agency costs create a wedge between the costs of external finance and internal funds.

In summary, our monetary surprises have statistically and economically significant effects on interest rates several years along the yield curve. The response of real and nominal gilt yields is stronger for shorter maturities and null for rates longer than five years. Using index-linked gilts we also identify an effect on breakeven inflation (up to 10 years ahead). These findings are comparable to the ones of Nakamura and Steinsson (2013) for the United States. We also find that a monetary policy surprise induces a significant increase in the corporate spread, in line with the findings of Gertler and Karadi (2015); appreciates the nominal exchange rate vis-a-vis the US Dollar; and generates a fall in equity prices.

Finally, the use of local projection methods at daily frequency allows us to provide some novel evidence on the persistence of these effects over time. Note that, unlike VARs, local projection methods do not force our surprises to inherit the average persistence of any disturbance to the interest rate. In this sense, we see our results as complementary to those of Rogers et al. (2014) for unconventional monetary policy.

3.2 Testing the Extent of Background Noise in the Measured Shocks

As noted above, these impulse responses are estimated with OLS. Implicit in the use of OLS is the assumption that, even if our measure of monetary news s_t does not capture all of the monetary news in a given month or quarter, the choice of a short window implies that it does not contain any background noise. If this assumption is violated, our parameter estimates will suffer from attenuation bias.

Concretely, we can write the measured surprise, s_t, as a function of some underlying monetary news, $\epsilon_{1,t}^{mp}$, and an orthogonal measurement error, η_t, namely:

$$s_t = \epsilon_{1,t}^{mp} + \eta_t \tag{8}$$

Our OLS estimator will only be consistent to the extent that the noise-to-signal ratio is vanishingly small:

$$\begin{aligned} plim\left(\widehat{\beta}_{OLS}^{h}\right) &= \frac{Cov\left(s_t, \Delta y_{t+h}\right)}{Var\left(s_t\right)} = \frac{Cov\left(\epsilon_{1,t}^{mp}, \Delta y_{t+h}\right)}{Var\left(\epsilon_{1,t}^{mp}\right) + Var\left(\eta_t\right)} = \\ &= \frac{Cov\left(\epsilon_{1,t}^{mp}, \Delta y_{t+h}\right)}{Var\left(\epsilon_{1,t}^{mp}\right)} \frac{Var\left(\epsilon_{1,t}^{mp}\right)}{Var\left(\epsilon_{1,t}^{mp}\right) + Var\left(\eta_t\right)}, \end{aligned} \tag{9}$$

where the left-hand term is the effect of pure monetary news on the response variable Δy_t and the right hand term tends towards unity as the noise-to-signal ratio tends towards zero.

We can test this assumption by comparing our OLS estimates, which are only consistent under the assumption that there is no noise in our measured monetary surprises, with the 'identification by heteroskedasticity' estimator (see Rigobon, 2003, Rigobon and Sack, 2004, Nakamura and Steinsson, 2013). This involves collecting data for a control sample of observations during which the variance of background noise is likely be the same, but the variance of monetary news is different. To this end we compile a control group $\{\Delta y_t^c, s_t^c\}$ of movements in the same asset prices during the window 1997:6–2015:5 on the last Wednesday of each month. The heteroskedasticity-based estimator is given by:

$$\widehat{\beta}^h_{RIG} = \frac{Cov\left(s_t, \Delta y_{t+h}\right) - Cov\left(s^c_t, \Delta y^c_{t+h}\right)}{Var\left(s_t\right) - Var\left(s^c_t\right)}. \tag{10}$$

If the difference between this estimator and OLS is small, it follows that the background noise in our measured monetary surprises is small. To conduct inference on this estimator, we follow Nakamura and Steinsson (2013) and construct a test statistic $g\left(\beta^h\right)$ that is zero at the true value of β^h:

$$g\left(\beta^h\right) = \Delta Cov\left(\Delta y_t, s_t\right) - \beta^h \Delta Var\left(s_t\right), \tag{11}$$

where ΔCov and ΔVar denote the difference in sample moments between the treatment and control samples.[17] For a given hypothetical value of β^h we can compute the distribution of $g\left(\beta^h\right)$ with a standard bootstrap procedure. If the hypothesized value of β^h falls within the confidence interval defined by the $\{\alpha/2, 1-\alpha/2\}$ percentiles of the distribution at which $g\left(\beta^H\right) = 0$ we cannot reject the null hypothesis that the OLS estimator is consistent.

We compute the test using the one-year gilt yield as left-hand side variable. We calculate $g\left(\beta^h\right)$ for 10^5 bootstrapped samples over a grid of values of $\beta^h \in [-1, 1]$. Figure 5 plots the median value of $g\left(\beta^h\right)$ as a function of β^h (solid line), together with its 95 percent confidence interval (shaded areas), i.e. where we fixed $\alpha = 5$. The interval for which $g\left(\beta^H\right) = 0$ is defined by $[0.44, 0.75]$, with a median estimate of 0.56. In our baseline results using OLS, we estimate the sensitivity of the change in the one-year gilt yield to our monetary surprise and obtain a coefficient of 0.54 with a standard error of 0.06. That is, our OLS estimate (represented by the dark dot in Figure 5) falls well inside the confidence interval of $[0.44, 0.75]$ and is close to the median estimate of 0.56. We accordingly conclude that the background noise in our measure of monetary news is small enough to be safely ignored, such that estimation and inference based on OLS is reliable.

An alternative possibility when confronted with measurement error of this sort would be to employ Instrumental Variable (IV) techniques, using the measured monetary surprises as an instrument. We can therefore calculate our scaled impulse response as:

$$plim\left(\widehat{\beta}^h_{IV}\right) = \frac{Cov\left(s_t, \Delta y_{t+h}\right)}{Cov\left(s_t, \Delta i_t\right)}, \tag{12}$$

where i_t is the one-year gilt yield. This is not necessary when the measurement error in our monetary policy surprises variable is negligible, and will reduce the precision of our estimates relative to OLS. But it is warranted when estimating a macroeconomic SVAR system to which

[17] As explained in Nakamura and Steinsson (2013), this more sophisticated procedure for inference is necessary when there is a significant probability that the difference in the variance of Δy_t between the treatment and control sample is close to zero. See Fieller (1954) and Staiger et al. (1997).

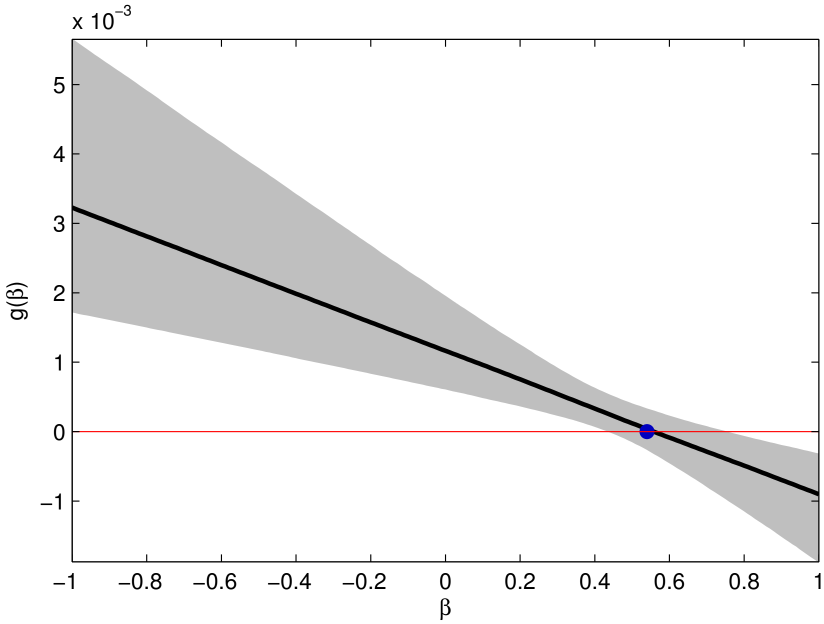

Figure 5 CONFIDENCE INTERVAL FOR $g\left(\beta^h\right)$. *Note.* The solid plots the median value of $g\left(\beta^h\right)$ as a function of β^h; the shaded area plots the 95 percent confidence interval. The interval for which $g\left(\beta^H\right) = 0$ is $[0.44, 0.75]$, with a median estimate of β^h of 0.56. The dark dot plots the sensitivity of the change in the one-year Gilt yield to the monetary surprise s_t obtained with OLS.

to we do not want to add our instrument. It is therefore the approach we take in the following section.

4 The Transmission of Monetary Policy Surprises to the Real Economy

In this section we investigate how of monetary policy surprises transmit to the real economy by estimating a structural VAR for the UK. We first present the empirical model and the procedure we use to identify the 'exogenous' monetary policy innovations. We then report the empirical results.

4.1 The Econometric Framework

The objective of this section is to provide evidence on the transmission of exogenous monetary policy innovations to the UK economy using a structural VAR. Therefore, as it is common in the VAR literature, we need a way to isolate an innovation to the monetary policy indicator that reflects shifts in the monetary policy stance that are not due to a response of central bank to other structural shocks. In order to identify such monetary policy 'shock' we use the external instruments identification approach proposed by Stock and Watson (2012) and Mertens and Ravn (2013), closely following the approach used by Gertler and Karadi (2015) for the case of the United States.

This identification strategy (whose details are reported in Appendix B) uses standard instrumental variable techniques to isolate the variation of the VAR reduced-form residuals that are due to the structural shock of interest. In this way it is possible to identify the contemporaneous response of all endogenous variables in the VAR system to the shock of interest. To obtain the impulse responses at longer horizons, one can then simply iterate the VAR forward.

The variables that we include in our baseline specification are the 1-year gilt yield, our policy indicator; a consumer price index as a measure of prices; the unemployment rate, as a measure of economic activity; the nominal effective exchange rate; stock prices as measure by the FTSE index; a measure of (investment grade) corporate spreads; and imports and exports.[18] We include imports, exports, and the exchange rate in our baseline specification because in a small open economy like the UK, movements in the exchange rate and in the trade balance are crucial determinants of monetary policy transmission. Similarly, we control for global shocks that are potentially important to explain the dynamics of domestic variables. Specifically, we include a global commodity price index and the VIX index in an exogenous block of the VAR. While the inclusion of the exogenous block helps the identification of UK-specific monetary policy surprises and increases the precision of our estimates, all our results are robust to dropping these variables.[19]

The monetary policy surprises that we constructed in section 2 are arguably a measure of monetary policy news that is not correlated with other fundamental disturbances. We can therefore use them to isolate the variation in policy instrument's reduced form residuals that is due exclusively to the monetary policy shock. Since different policy surprises (i.e., computed with different underlying contracts) are available, we choose the one that has the largest F-Statistic in instrumenting the reduced form residuals of the 1-year gilt equation in the VAR. In our case, this is the 2^{nd} front contract of 3-month Sterling future.

Following Sims et al. (1990), we estimate the VAR systems in levels without explicitly modeling the possible cointegration relations among them.[20] We use the BIC information criterion to choose the optimal number of lags, which we set to two. We check that the residuals are not

[18] All variables were Seasonally Adjusted using X13-ARIMA-SEATS program. The consumer price index, nominal effective exchange rate, stock prices, imports and exports enter the VAR in log-levels. Appendix A describes the data sources for all variables.

[19] These robustness exercises, together with the extensive list of other robustness checks mentioned below, are reported in the Online Appendix.

[20] Sims et al. (1990) show that if cointegration among the variables exists, the system's dynamics can be consistently estimated in a VAR in levels.

serially correlated with this specification. Note, however, that the results are robust to different lag specifications.

We estimate the VAR using monthly data for the UK for the period 1993:1–2015:5.[21] We choose the starting point to coincide with the beginning of the inflation targeting regime in the UK. Prior to this, the UK was (i) essentially shadowing the Deutsche Mark and (ii) the target and operating framework for monetary policy were very different. Thus, a sample starting before 1993:1 will likely affected by a structural break. Given the ending point, the data includes the recent crisis and its aftermath, where Bank rate — the 'typical' monetary policy indicator — did not move from the level of 50 basis points reached in 2009. To address this issue, we choose as policy indicator a safe interest rate at longer maturity (i.e., the nominal yield on the 1-year gilt). But we also check the robustness of our results by (i) using longer maturity gilts as a policy indicator; (ii) using the 3-month forward exchange rate between the British Pound and the US Dollar as an instrument; and (iii) excluding the period over which Bank rate did not show any time variation.[22]

Note that our monetary policy surprises are available only for a subsample of the period over which the VAR is estimated, namely from 1997:6 to 2015:5. We choose a longer sample period for the estimation of the VAR so as to estimate with greater precision the lag coefficients and the reduced form residuals. Finally note also that we need to aggregate the daily monetary policy surprises into a monthly series. We do that following the procedure employed by Gertler and Karadi (2015) and we check that results do not change when simply summing the surprises within the month. The time series properties of the monthly surprises are reported in Appendix B, together with their correlation with the monetary policy surprises computed with different contracts.

4.2 Estimation Results

In this section we report two sets of results. Before turning to our full specification, we report the impulse responses from a smaller scale VAR that allows a direct comparison with Gertler and Karadi (2015)'s baseline results.

Comparison with Gertler and Karadi. To allow a comparison with Gertler and Karadi (2015), we estimate a VAR with four variables only: the yield on the 1-year gilt, the CPI, unemployment, and the corporate spread. Figure 6 displays the impulse response function (IRF) to an instrumented increase in the 1-year gilt rate, using as an instrument the 2^{nd} front contract of 3-month Sterling future. Note that the instrument is quite powerful in explaining the behavior of the reduced form residuals of the policy indicator equation. The F-statistic from

[21] For the variables for which data is available at higher frequency, we compute monthly averages.
[22] Again, all these robustness exercises are reported in the Online Appendix.

the first stage is 19.66, well above the relevant threshold of 10 suggested by Stock and Yogo (2002). The R^2 of the first stage regression is 0.08.

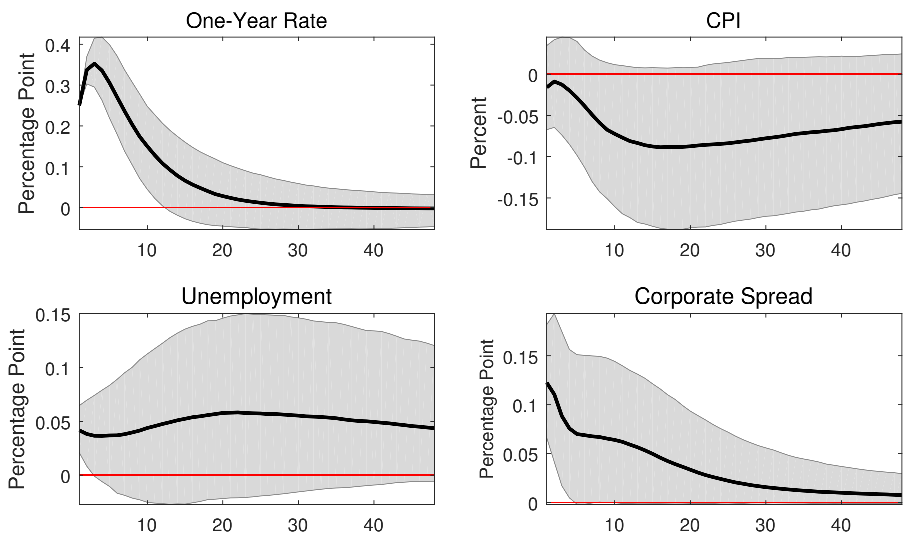

Figure 6 IRFs To A Monetary Policy Shock - Comparison With Gertler And Karadi (2015). *Note.* VAR estimated in log levels, with 2 lags, and a constant over the 1993:1-2015:5 period. The VAR includes as exogenous variables a Commodity Price Index and the VIX index. The 1-year Government Gilt Yield is instrumented using the second front contract of 3-month Sterling future. First stage results: F-Statistic: 19.66 and $R^2 = 0.08$. The solid line and shaded areas report the mean and the 90% confidence intervals computed using wild bootstrap with 1,000 replications.

We normalize the shock so that the 1-year gilt rate increases by 25 basis points. The shock has a persistent effect on the 1-year gilt yield, lasting for about ten months after the shock hit. Consumer prices fall slightly on impact, but the response is not statistically significant. The impact, however, builds up over time and becomes borderline statistically significant (at the 90 percent confidence level) ten months after the shock hit, at a level of about -0.1%. This magnitude is consistent with the evidence reported by Gertler and Karadi (2015) and with the common view that the transmission of monetary policy is slow and gradual. This response also shows that the identification through our external instrument does not suffer from a typical problem that affects VARs identified with short-run restrictions, namely the "Price Puzzle".[23]

The shock induces a small, and statistically significant increase in unemployment, consistent with the contractionary impact of monetary policy shocks. Over the horizon considered in the IRFs, the increase in unemployment reaches a maximum of 0.05%.[24]

[23] In a similar vein to Gertler and Karadi (2015), we show in the Online Appendix that, when the monetary policy shock is identified with short-run restriction (i.e., where policy indicator is not allowed to respond contemporaneously to unemployment and CPI), CPI tends to increase after a monetary policy tightening.

[24] We obtain virtually identical results in an even smaller scale VAR where we drop corporate credit spreads from the list of endogenous variables. Results are reported in the Online Appendix.

Finally, the response of corporate spreads confirms one important result put forth by Gertler and Karadi (2015). Our results show that a monetary policy shock that increases the yield on the 1-year gilt by 25 basis point leads to an increase of corporate spreads by about 12.5 basis points, therefore supporting the view monetary policy operates through a credit channel (Bernanke and Gertler, 1995).[25]

Full specification. We turn now to the full specification where, in addition to the variables included in the previous VAR, we add the nominal effective exchange rate, imports and exports, and the FTSE index. While the result from the first stage regression worsen a little (the F-Statistic is now 13 and the R^2 is 0.06), the inclusion of these variables allows us to consider an important dimension for the transmission of monetary policy shocks in a small open economy like the UK.

Figure 7 reports the IRFs to a monetary policy surprise that increase our policy indicator (the nominal yield on the 1-year gilt) by 25 basis points in this larger VAR. The contractionary monetary policy shock generates a persistent and significant reduction (0.13%) in the CPI that reaches its minimum ten months after the shock. The magnitude of the response is similar to the response in the smaller scale VAR reported in Figure 6.

In line with the daily regressions reported in section 3, the Pound appreciates by about 1% in nominal effective terms and the Investment Grade Corporate Spread increases by about 15 basis points, again confirming the finding of Gertler and Karadi (2015) that the credit channel is a relevant transmission mechanism of monetary policy. The FTSE index declines on impact by 2% in response to the monetary policy tightening.

In principle, the real trade balance can improve or deteriorate in response to a tightening of monetary policy. The exchange rate appreciation tends to switch expenditure towards foreign goods, pushing down on exports and up on imports. On the other hand, the compression in domestic expenditure that a monetary contraction creates will tend to push imports down. In our baseline estimates, we find that a monetary policy shock leads to a 1% decline in exports, which becomes significant four months after the shock. Import volumes fall by a similar amount. Thus, monetary policy does not seem to affect significantly the trade balance in the UK.

The results in Figure 7 are slightly different from the ones of Cloyne and Huertgen (2014), who use data for the UK over the 1975–2007 period. In particular, they find that inflation reacts significantly only 24 months after the shock while in our case the reaction starts at 5 months. This difference is most likely explained by the different sample we use in our analysis, namely the period of Central Bank independence from 1993 to 2015. In a model with time-varying

[25] We estimate the same 4-variable VAR (and the larger-scale VAR below) excluding the extraordinary MPC meetings and the dates of release of the Inflation Report from the set of events. Results are robust and reported in the Online Appendix.

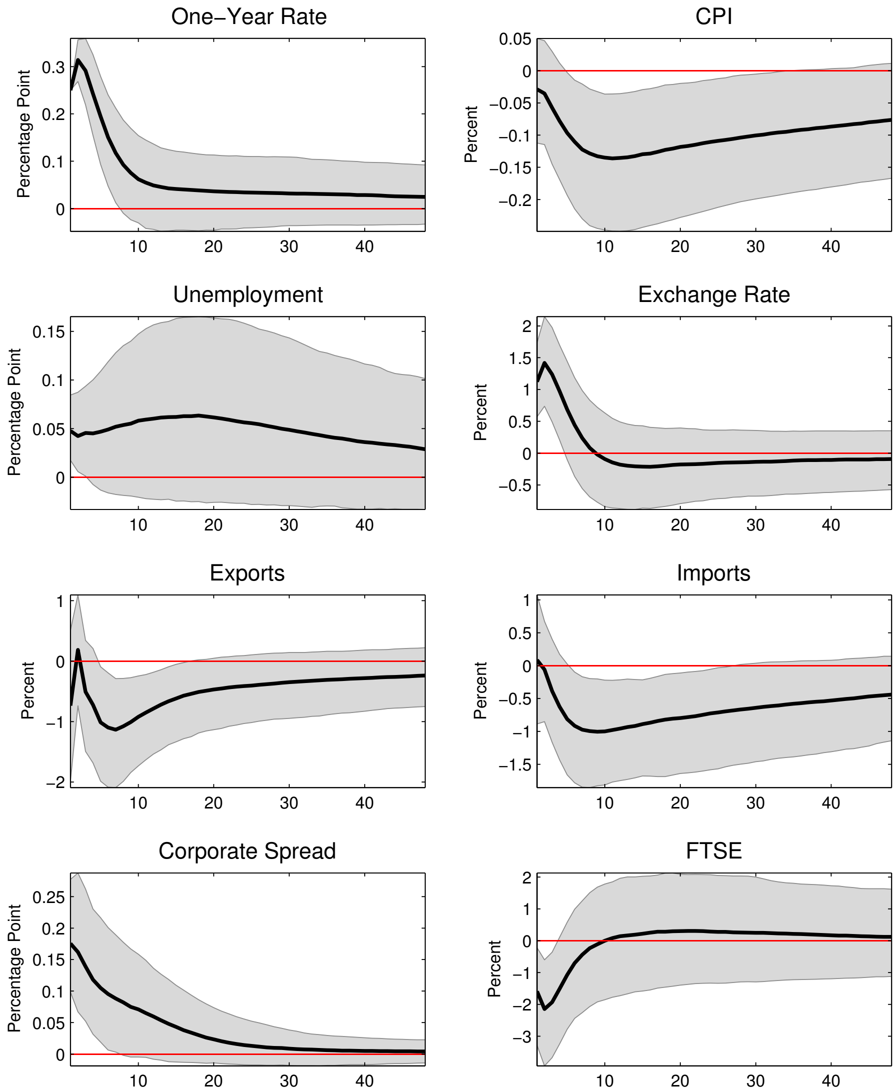

Figure 7 IRFs To A Monetary Policy Shock - Full Specification. *Note.* VAR estimated in log levels, with 2 lags, and a constant over the 1993:1-2015:5 period. The VAR includes as exogenous variables a Commodity Price Index and the VIX index. The 1-year Government Gilt Yield is instrumented using the second front contract of 3-month Sterling future. First stage results: F-Statistic: 13 and $R^2 = 0.06$. The solid line and shaded areas report the mean and the 90% confidence intervals computed using wild bootstrap with 1,000 replications.

coefficients, Ellis et al. (2014) show that monetary policy in the UK has become significantly more effective to affect CPI after 1992, which they claim occurred via the impact of expectations on prices. This accords well with the additional results reported in Cloyne and Huertgen (2014) who show that in the post-1992 period inflation responds slightly faster than in their baseline. Finally, our findings are comparable to Mountford (2005), who identifies the effects of monetary shocks in the UK using sign restrictions, both in terms of magnitudes and timing.

Summing up, the monetary policy surprise has a significant and persistent effect on the macroeconomic variables and affects the economy both through the increase in the Corporate Spread and the appreciation of the pound Sterling. In terms of the effect on activity and inflation, Table 1 contains a detailed comparison with previous findings. Our estimates are broadly within the pack, and if anything somewhat on the small side. However, the difference in the precise measures and orders of integration of activity and inflation employed in the different studies makes an exact comparison difficult. In particular, unemployment is likely to be smoother than industrial production because of labour hoarding in response to temporary shocks, and the large weight of the service and public sectors in total employment, sectors which may respond less than industry to a monetary policy innovation.

Table 1 Summary of Previous Studies on Macroeconomic Effects of Monetary Policy for the US and the UK

Authors	Country	Method	Peak Effects (in %)	
			Output	Prices/Inflation
Bernanke and Mihov, 1998	US	VAR	-0.6 to -1 (GDP)	-0.7 to -1.6 (GDP Defl)
Christiano et al., 1999	US	VAR	-0.7 (GDP)	-0.6 (GDP Defl)
Romer and Romer, 2004	US	Narrative	-1.9 to -4.3 (IP)	-3.6 to -5.9(CPI/PPI)
Uhlig, 2005	US	Sign Rest.	-0.3 (GDP)	-1.0 (GDP Defl)
Bernanke et al., 2005	US	FAVAR	-0.6 (IP)	-0.7 (CPI)
Coibion, 2012	US	Narrative	-1.6 to -4.3 (IP)	-1.8 to -4.2 (CPI Infl)
Barakchian and Crowe, 2013	US	Fed Futures	-0.9 (IP)	-0.1 (CPI)
Gertler and Karadi, 2015	US	Proxy SVAR	-1.0 to -2.0 (IP)	-0.75 to 0.3 (CPI)
Dedola and Lippi, 2005	UK	VAR	-0.5 (IP)	0.2 (CPI)
Mountford, 2005	UK	Sign Rest.	-0.6 (GDP)	-0.15 (GDP Defl)
Ellis et al., 2014	UK	FAVAR	-2.0(IP, 92-05)	-2 (CPI, 92-05)
Cloyne and Huertgen, 2014	UK	Narrative	-0.5 (IP)	-1.0 (CPI Infl)
Cesa-Bianchi et al., 2016	UK	Proxy SVAR	0.2 (Unempl)	-0.5 (CPI)

Note. This table is an update of Table reported in Cloyne and Huertgen (2014). The peak effects correspond to a one percentage point increase in the interest rate. In brackets we include the specific measure of Output and Prices considered in each study. CPI Infl denotes CPI inflation.

5 Tests of Instrument Validity

A key condition for our estimates in the previous sections to be consistent is that our instrument, s_t, is uncorrelated with non-monetary innovations in the system. As argued above, by selecting a short window around policy events and dropping observations containing data releases, we are able to make this so. But it is still possible that the policy decision contains news about the determinants of monetary policy. For example, if the Bank of England has superior information about the state of the economy, changes in monetary policy which comes as a surprise to markets may be systematically correlated with non-monetary developments in the macroeconomy.

In this section we exploit the availability of a complementary (and notionally exogenous) measure of UK monetary policy innovations to check the validity of our instrument with a test of overidentifying restrictions. This alternative measure is the one constructed by Cloyne and Huertgen (2014), following the methodology proposed by Romer and Romer (2004), for the period 1975:1–2007:12. We report it, together with our measure of monetary policy surprises, in Figure 8. The two series overlap on the sample period 1997:6–2007:12. Over this period, the two series are nearly orthogonal, with a contemporaneous correlation of -0.006.[26]

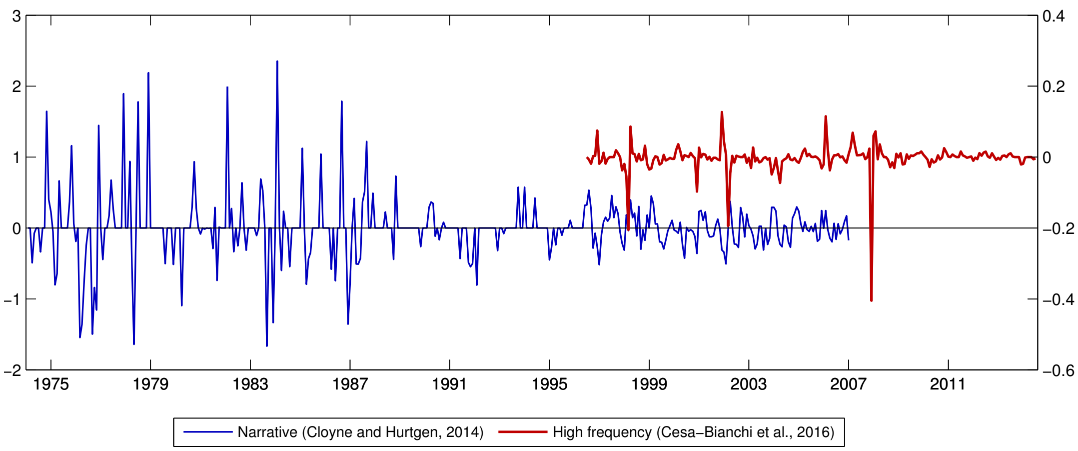

Figure 8 Instruments For Monetary Policy Shock: Narrative And High Frequency Measures. *Note.* The blue line displays Cloyne and Huertgen (2014)'s instrument for monetary policy shocks (left axis). The red line displays the high-frequency instrument developed in this paper (right axis).

To construct the test of overidentifying restrictions we proceed as follows. Consider the structural representation of a VAR (with no constant and only one lag for simplicity of exposition):

$$X_t = F_1 X_{t-1} + B\epsilon_t, \tag{13}$$

where X_t is a $n \times 1$ vector of endogenous variables; ϵ_t is a $n \times 1$ vector of orthogonal structural shocks, with $\Sigma_\epsilon = I_n$; F_1 and B are $n \times n$ matrices of coefficients. The reduced form residuals of the above VAR are given by $u_t = B\epsilon_t$.

Denote by z_1 the series of monetary surprises constructed in this paper (i.e., the surprises s_t) and by z_2 an alternative series of instruments for exogenous monetary shocks, such as those

[26] The contemporaneous correlation between US narrative monetary policy shocks and US monetary policy surprises is also low. We compute it using the data in Tenreyro and Thwaites (2016) (an update of the original Romer and Romer (2004) series up to 2007:12 period) and the original data made available by Gertler and Karadi (2015). The contemporaneous correlation varies between 0.14 and 0.27 depending on the monetary policy surprised used.

in Cloyne and Huertgen (2014). We assume that they are related to the true, full series of monetary shocks ϵ^{mp} as follows:

$$\begin{aligned} \epsilon^{mp} &= \alpha_1 z_1 + \xi_1, \\ \epsilon^{mp} &= \alpha_2 z_2 + \xi_2, \end{aligned} \tag{14}$$

where the ξ_i are orthogonal to z_i for $i = 1, 2$. The idea implicit in this representation is that each instrument captures only a subset of the universe of monetary shocks ϵ_t^{mp} that occur within a given period, while the remaining part is captured by the term ξ_i.[27]

To show how the procedure works, we first re-write the reduced form residuals of the policy rate, CPI, and unemployment equations (u^r, u^{cpi}, and u^u) into two orthogonal components: one due to the monetary policy shock ϵ^{mp} and a second one which is a linear combination of the other structural shocks.[28] Specifically, by letting b_{ij} be the i^{th} and j^{th} element of the B matrix in (13), we have:

$$\begin{aligned} u^r &= b_{11}\epsilon^{mp} + \zeta^r, \\ u^u &= b_{21}\epsilon^{mp} + \zeta^u, \\ u^{cpi} &= b_{31}\epsilon^{mp} + \zeta^{cpi}, \end{aligned} \qquad \text{where} \qquad \begin{aligned} \zeta^r &\equiv \Sigma_{i=2}^n b_{1i}\epsilon^i, \\ \zeta^u &\equiv \Sigma_{i=2}^n b_{2i}\epsilon^i, \\ \zeta^{cpi} &\equiv \Sigma_{i=2}^n b_{3i}\epsilon^i, \end{aligned} \tag{15}$$

where ϵ^i for $i = 2, ..., n$ are the remaining $n - 1$ structural shocks. Now, we can relate the instruments and reduced form residuals to the unobserved components by combining equations (14)-(15):

$$\begin{bmatrix} z_1 \\ z_2 \\ u^r \\ u^u \\ u^{cpi} \end{bmatrix} = \begin{bmatrix} 1/\alpha_1 & -1/\alpha_1 & 0 & 0 & 0 & 0 \\ 1/\alpha_2 & 0 & -1/\alpha_2 & 0 & 0 & 0 \\ b_{11} & 0 & 0 & 1 & 0 & 0 \\ b_{21} & 0 & 0 & 0 & 1 & 0 \\ b_{31} & 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \epsilon^{mp} \\ \xi_1 \\ \xi_2 \\ \zeta^r \\ \zeta^u \\ \zeta^{cpi} \end{bmatrix}. \tag{16}$$

The covariance matrix of the vector of unobservables (whose derivation is reported in Appendix C) gives a vector of 14 parameters to estimate:

$$\theta = \{\alpha_1, \alpha_2, b_{11}, b_{21}, b_{31}, \sigma^2_{\xi_1}, \sigma^2_{\xi_2}, \sigma^2_{\xi_1\xi_2}, \sigma^2_{\zeta^r}, \sigma^2_{\zeta^u}, \sigma^2_{\zeta^{cpi}}, \sigma^2_{\zeta^r\zeta^u}, \sigma^2_{\zeta^r\zeta^{cpi}}, \sigma^2_{\zeta^u\zeta^{cpi}}\},$$

and we observe 15 moments in the covariance matrix of our five observables $\{z_1, z_2, u^r, u^u, u^{cpi}\}$. We estimate the parameters of this system with iterative GMM.[29] Were we to use only our

[27] Note that, consistently with our results in Section 3, this representation assumes that our instruments have no background noise (η). As we shall see below, this representation is also supported empirically, as we find that both z_1 and z_2 are statistically significant in explaining the policy indicator reduced form residual.

[28] Note that, in principle, one could consider up to n reduced form residuals but —as we shall see below— this is not needed to achieve overidentification in our specific application.

[29] We check that the parameters are locally identified at the optimum with the gradient matrix of the moment vector (checking that it is non-singular). We randomize over starting values for the optimization procedure to ensure we have attained a global optimum. Additional details on the GMM procedure we use, together with the

instrument, then our estimates of b_{11}, b_{21}, and b_{31} will be the same as in section 4. The addition of an extra observable will change this to some extent. Our system is overidentified, so the moment conditions are unlikely to hold exactly. But if our restriction holds approximately in the data then the minimized value of our moment conditions will be close to zero. We can accordingly test the null hypothesis that our exclusion restrictions hold with the Hansen-Sargan statistic. The p-value of the resulting test is 0.39, indicating that we do not reject the null of our exclusion restrictions. We accordingly find no evidence that our instruments or those contained in Cloyne and Huertgen (2014) are endogenous.

A complementary and less formal test of the same condition is to estimate the SVAR system using both instruments and see if we get substantially different results. If not, the addition of an extra instrument may help to sharpen these results. Using the overlapping sample period 1997:6–2007:12, we run a simple regression of the reduced form residuals on the two sets of monetary policy surprises, and we find that both series are statistically significant. This suggests that the two series pick up different sources of exogenous monetary innovation.

To make best use of the available data, we need a means of incorporating the non-overlapping sample period. We try two alternative methods of aggregating the two non-overlapping samples. First, we take a simple average of the instruments, having first normalized them to have equal variance, combining them in a unique series that spans the whole second-stage sample period 1993:1–2015:5. The resulting series can therefore be used as an instrument for the reduced form residuals (as in our baseline specification).[30] We do this in light of the fact that both instruments (i) are orthogonal, (ii) explain a significant fraction of the reduced form residual, and (iii) are available over different sample periods that in part do not overlap.

Alternatively, by conduct three separate first-stage regressions — corresponding to one sub-sample in which both instruments are available and two in which only one is — to obtain fitted values of the interest rate. This method has the advantage of not restricting the coefficients on the two (normalized) instruments to be the same, but the disadvantage of running regressions with smaller samples and therefore weaker identification.

Figure 9 displays the IRFs obtained using these new series as an instrument. The F-Statistic of the first stage regression increases (relative to our baseline specification) to 18.1 and the R^2 is 0.07 when using the average of the two instruments (higher than using each instrument separately).[31] The results show that both procedure give very similar results to our baseline.

full system of moment conditions, are reported in Appendix C.

[30] As an alternative (and virtually equivalent) way of combining the two series, we regress the reduced form residuals on each instrument separately and then take an average of their fitted values. The results obtained with these alternative series are robust.

[31] Consistently, the boostrapped responses of macroeconomic and financial variables are similar to our baseline specification, with slightly smaller error bands. See Figure D.1 Appendix D.

Taken together, these results suggest that our instrument is valid and, relatedly, that our results are robust to and somewhat sharpened by combining it with the other available instrument.

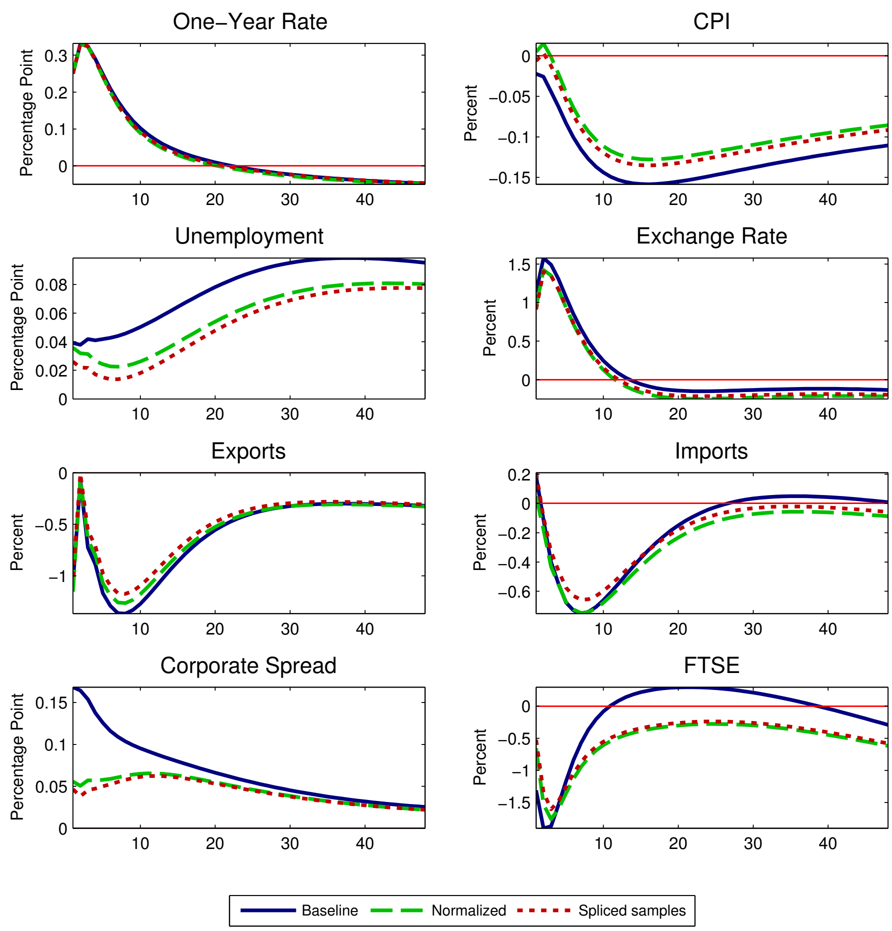

Figure 9 IRFs To A Monetary Policy Shock - Full Specification With Two Instruments - Comparison of Instrument Aggregation Methods. *Note.* VAR estimated in log levels, with 2 lags, and a constant over the 1993:1-2015:5 period. The VAR includes as exogenous variables a Commodity Price Index and the VIX index. The 1-year Government Gilt Yield is instrumented using the second front contract of 3-month Sterling future (blue line). It is combined with Cloyne and Huertgen (2014)'s monetary policy shocks series as a normalized sum (green dashed line) and with 3 subsample regressions (red dotted line).

One final concern is that one (or both) set of instruments is not valid but this is not picked up by the test because it has low power — in which case the test will often not reject a false null hypothesis of invalid instruments. With this in mind, we run our GMM test using the US analogues of our surprises (i.e. Gertler and Karadi, 2015) and those in Cloyne and Huertgen (2014) (i.e. Romer and Romer, 2004). Researchers have separately used alternative tests to claim that both sets of US shocks are forecastable from lagged information, and therefore endogenous to the macroeconomy.[32] We run our GMM-based test above with these two US series and can reject the null that they are both valid instruments at a 5% significance level.[33] This suggests that our test has some power to detect violations of the validity conditions, and accordingly gives some confidence that our non-rejection of the UK instruments is not a false negative.

6 Conclusions

What is the impact of 'exogenous' monetary policy surprises on financial markets and on the macroeconomy? This paper tries to answer this crucial question, using a novel data set for the UK.

To identify exogenous variation in monetary policy, we construct a series of UK monetary policy surprises using the high-frequency methods pioneered by Kuttner (2001) and Gurkaynak et al. (2005a). In line with previous studies, we find evidence of their effect on UK real interest rates and inflation. Applying local projection methods, we are also able to provide some novel evidence on the persistence of these effects in financial markets. We then employ our series of monetary policy surprises as instruments in an structural VAR. A monetary policy tightening generates a persistent and statistically significant reduction in the CPI that reaches its minimum ten months after the shock, appreciates the Pound, and increases corporate spreads. These findings confirm that both the credit and external channels are relevant for the transmission of monetary policy in an open economy like the UK.

The monetary policy surprises that we construct are designed to be exogenous to non-monetary developments in the macroeconomy. But there is still the possibility that policy events contain significant information about the macroeconomic determinants of monetary policy, therefore undermining our procedure. To provide further evidence of the exogeneity of our measured monetary surprises, we exploit the alternative measure of monetary policy innovations constructed by Cloyne and Huertgen (2014) using narrative methods. We propose a new test of overidentifying restrictions and find no evidence that our monetary policy surprises contain any response to macroeconomic variables. This suggests that both series contain complementary information about monetary policy.

Overall, our findings suggest that monetary policy has significant and persistent effects on both financial and macroeconomic variables. This evidence is relevant to improve our under-

[32] See, for example, Gertler and Karadi (2015), Ramey (2016), Campbell et al. (2016), Vicondoa (2016), Miranda-Agrippino (2016).

[33] We first replicated Gertler and Karadi (2015) baseline VAR (i.e., Figure 1 in their paper) using the data available at https://www.aeaweb.org/aej/mac/data/0701/2013-0329_data.zip. We then estimated the same system as in (16) using the residuals of policy rate, industrial production and CPI and the two alternative instruments. The p-value of the Hansen-Sargan test is 0.05, indicating that we do reject the null hypothesis that our exclusion restrictions hold.

standing of the different transmission channels of monetary policy, which has been keenly debated. A key advantage of our series of surprises is that it includes market reaction both to current unexpected changes in policies and to future path of monetary policy related to monetary policy events. Considering that central banks have been relying more on forward guidance, we hope that this new series of surprises, together results presented in this paper, will be useful for the current debate and future research on this area.

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A Appendix. Data

A.1 Events

This section describes the events that we consider to compute the monetary policy surprises. As explained in the main text, we use two main monetary policy events.

Publication of Inflation Report. This report sets out detailed economic analysis and inflation projections on which the Monetary Policy Committee bases its interest rate decisions, and presents an assessment of the prospects for UK inflation. This report is published on a quarterly basis: February, May, August, and November. Dates and times were collected from the Bank of England database and Bloomberg.

Interest rate decisions. The Monetary Policy Committee meets every month to set the interest rate. We use the dates and times in which the decision of the interest rate was announced. This occurs straight after the meeting. This information was collected from the Bank of England database and Bloomberg.

As explained in the main text, we exclude the release of the *Monetary Policy Minutes* since they coincide with the publication of relevant labour market information. This fact would introduce noise in the measurement of the surprise because the reaction of financial markets could be due to the new flow of information about the state of the economy. Note, however, that the main results presented in the paper are robust to adding the *Monetary Policy Minutes* to the set of events.

A.2 High Frequency (Tick-by-Tick) Data

Data about financial contracts was downloaded from Thomson Reuters Tick History Database. All transactions in future markets are recorded with their corresponding time (at the millisecond frequency), price, and volume traded. For our analysis, we use the following contracts.

Sterling Future. These contracts are settled based on the 3-Month London Interbank Offered Rate (LIBOR) and traded at the ICE LIFFE Futures and Options Exchange (LIFFE). In particular, every year there are 4 delivery months: March, June, September, and December plus two serial consecutive months, with the nearest three delivery months being consecutive calendar months. These contracts are traded until the third Wednesday of the delivery month and are cancelled on the next business day.[34] Similar to the Fed Fund Futures, we can extract the expected rate from the price of each contract using the following expression:

$$P_t^h = 100 - \mathbb{E}_t\left[i_h^{(h+90)}\right] \tag{A.1}$$

where P_t^h denotes the current price for a contract that matures on day h and $\mathbb{E}_t\left[i_h^{(h+90)}\right]$ denotes

[34] The following web page https://www.theice.com/products/37650330/Three-Month-Sterling-Short-Sterling-Future contains more information about these contracts.

the expected value of the 3-month (i.e. $h + 90$ days) Libor at h.

Considering the volume traded, we use the continuous synthetic series computed by Thomson Reuters. In particular, we use: *FSScm1*, *FSScm2*, *FSScm3*, and *FSScm4*, which correspond to the first, second, third, and fourth continuous contract respectively. These synthetic series are computed using the underlying contracts at each date. For example: on January 1st, 2000, *FSScm1*, *FSScm2*, *FSScm3* and *FSScm4* track the contracts that expire on March, June, September, and December, respectively. Thus, at every date they capture one-year ahead expectations of the 3-month Libor. However, these continuous series are available since June 1999. In order to complete each series from June 1997, we use the same rolling formula than Thomson Reuters and compute the pricing of each contract using their respective underlying contract. To check for accuracy, we compare our computed series with the ones reported by Thomson Reuters for the period 1999-2000 and they coincide.

Forward FX between Pound and USD. This corresponds to the forward contract based on the expected exchange rate between the Pound and the US Dollar 3 months ahead. Thus, these contracts reflect the expected appreciation/depreciation of the Pound against the US Dollar. Unlike the *Sterling Future*, this contract has a continuous of expiring dates and not just 4 times a year. We use the series under the RIC *GBP3M*, which is available since January 1996 and is very liquid.

Following Gurkaynak et al. (2005a) and Gertler and Karadi (2015), we define a monetary policy surprise as the change in price for each contract between 20 minutes after and 10 minutes before the event (i.e. a 30 minute window).

A.3 Macroeconomic And Financial Data

Daily Data. In our high-frequency section we use the following series:

- *Gilt Yields, Forward Gilt Yields, Real Gilt Yields, Real Forward Gilt Yields*: The data comes from the estimated Government Yield Curves for different types of bonds, which are computed at daily frequency by the Bank of England. We choose representative maturities for the most traded contracts (i.e. with fewer missing values).[35]

- *Expected Inflation*: Implied Inflation for different maturities computed from the estimated real Yield Curve. This series is available at daily frequency and published by the Bank of England.

- *FTSE*: Source: Thomson Reuters Datastream.

- *Euro, Dollar, and Yen bilateral exchange rates and Exchange Rate Index (ERI)*: These series of daily nominal exchange rates are computed by the Bank of England.

[35] The website http://www.bankofengland.co.uk/statistics/pages/yieldcurve/default.aspx contains more information about how the yield curve is estimated.

- *VIX index*: CBOE Volatility Index. Source: FRED Economic Data.
- *Spread IG*: Corporate Spread Investment Grade. This series is computed at a daily frequency by the Bank of England.
- *Spread HY*: Corporate Spread High Yield. This series is computed at a daily frequency by the Bank of England.
- *Spread Libor 6M-Bank Rate*: This variable is computed as the difference between the 6M Libor and the Bank Rate. Source: Bank of England.

Monthly Data. In our VAR analysis we use the following macroeconomic series:

- *One-Year Rate*: One Year Nominal Gilt Yield. Source: Bank of England. We use the monthly average of the daily series.
- *CPI Index*: UK CPI INDEX 00: All items - 2005 = 100. Source: Office for National Statistics, U.K. We seasonally adjust this series using X13-ARIMA-SEATS program.
- *Unemployment Rate*: Unemployment Rate expressed in %. Source: International Financial Statistics (IMF). We seasonally adjust this series using X13-ARIMA-SEATS program.
- *Nominal Exchange Rate*: Nominal Exchange Rate Index. Source: Bank of International Settlements. This index is calculated as geometric weighted averages of bilateral exchange rates. It is available as monthly average and an increase indicates an appreciation.
- *Export and Import*: Volume Indexes (2011=100) at monthly frequency. Source: Office for National Statistics, U.K. We seasonally adjust these series using X13-ARIMA-SEATS program.
- *Corporate Spread*: Investment Grade Corporate Spread Index. Source: Bank of England. This series is available at daily frequency since January-1997. For the VAR, we compute the monthly average of this series. Before 1997, the series was computed as the difference between the Yield on Deventures and the Bank Rate. The former is available at monthly frequency from *Three Centuries of Data* dataset, which is published by the Bank of England.
- *FTSE Index*: Monthly average of FTSE All-Share index. Source: Datastream.
- *Commodity Price Index*: All Commodity Price Index, includes both Fuel and Non-Fuel Price Indices. Source: IMF.
- *VIX index*: Monthly average of the CBOE Volatility Index. Source: FRED Economic Data.

B Appendix. Proxy SVAR

This Appendix describes (i) the proxy SVAR methodology that we use to trace out the impact of monetary policy surprises on the macroeconomy the and (ii) the time series properties of the external instrument, i.e. the monetary policy surprises aggregated at monthly frequency.

B.1 Methodology

Consider the following VAR (with only one lag and no constant or trend for simplicity):

$$Y_t = AY_{t-1} + u_t. \tag{B.1}$$

where Y_t is a $(m \times 1)$ vector of endogenous variables; A is an $(m \times m)$ matrix of coefficients; u_t is a $(m \times 1)$ vector of reduced form residuals with variance-covariance matrix Σ_u. The objective is to recover the structural form of the above VAR, i.e.:

$$Y_t = AY_{t-1} + B\epsilon_t, \tag{B.2}$$

where A and B are $(m \times m)$ matrices of coefficients; and ϵ_t is an $(m \times 1)$ vector of structural residuals with variance-covariance matrix $\Sigma_\varepsilon = I$. Note that the reduced form residuals are a linear combination of the structural residuals. Specifically, $B\epsilon_t = u_t$.

If we partition the vector of endogenous variables Y_t as $(r', X_t')'$ —where r_t is a monetary policy indicator and X_t is the $(m - 1 \times 1)$ vector of remaining endogenous variables— we can re-write the reduced-form VAR as:

$$\begin{bmatrix} r_t \\ X_t \end{bmatrix} = \begin{bmatrix} A_{11} & A_{12} \\ A_{21} & A_{22} \end{bmatrix} \begin{bmatrix} r_{t-1} \\ X_{t-1} \end{bmatrix} + \begin{bmatrix} B_{11} & B_{12} \\ B_{21} & B_{22} \end{bmatrix} \begin{bmatrix} \epsilon_t^{mp} \\ \epsilon_t^{X} \end{bmatrix}, \tag{B.3}$$

where A_{11} and B_{11} are scalars; A_{12} and B_{12} are $(1 \times m - 1)$ vectors; A_{21} and B_{21} are $(m - 1 \times 1)$ vectors; A_{22} and B_{22} are $(m - 1 \times m - 1)$ matrices; and ϵ_t^{mp} and ϵ_t^{X} are the structural residuals associated to monetary policy and the remaining endogenous variables, respectively.

For the sake of argument, let's assume that the structural matrix B is known. Then, we would be able to compute the impulse response to a monetary policy shock. Specifically, the contemporaneous responses of r and X to a unit shock to ϵ_t^{mp} would be given by:

$$\begin{bmatrix} \mathcal{IRF}_0^r \\ \mathcal{IRF}_0^X \end{bmatrix} = \begin{bmatrix} B_{11} \\ B_{21} \end{bmatrix},$$

which, since the model is linear, can be normalized to:

$$\begin{bmatrix} \mathcal{IRF}_0^r \\ \mathcal{IRF}_0^X \end{bmatrix} = \begin{bmatrix} 1 \\ \frac{B_{21}}{B_{11}} \end{bmatrix}. \tag{B.4}$$

Finally, the impulse response functions at longer horizons can be computed as:

$$\mathcal{IRF}_n = A^{n-1} \cdot \mathcal{IRF}_{n-1} \qquad \text{for} \quad n = 2, ..., N. \tag{B.5}$$

Note that if we are interested in computing the impulse responses to the monetary policy shock only we do not need to know all the coefficients of B, but rather only the elements of the first column of B, namely B^1.

We now consider the case of B unknown. To achieve identification, we follow the external instrument identification approach pioneered by Stock and Watson (2012) and Mertens and Ravn (2013). Let u^r and u^X be the OLS estimates of the reduced form residuals in (B.1). Also, let Z_t be a $(z \times 1)$ vector of instrumental variables that satisfy:

$$\begin{aligned} \mathbb{E}\left[\epsilon^{mp} Z_t'\right] &= \phi, \\ \mathbb{E}[\epsilon^X Z_t'] &= 0, \end{aligned}$$

i.e., the instruments are correlated with the monetary policy shock (ϵ^{mp}) but are orthogonal to all the other domestic shocks (the elements of ϵ^X). We can obtain consistent estimates of B^1 from the two-stage least squares regression of u^X on u^r using Z_t as instruments. In other words, since the reduced form residuals of the monetary policy indicator equation (u_t^r) are an imperfect measure of true structural shock (ϵ^{mp}), in the first stage we regress them on the set of instruments (Z_t):

$$u_t^r = \beta Z_t + \xi_t, \tag{B.6}$$

to construct the fitted values $\hat{u}_t^r$. Then we regress the reduced form residuals of the domestic equations (u_t^X) on the fitted values ($\hat{u}_t^r$) to get a consistent estimate of the ratio B_{21}/B_{11}:

$$u_t^X = \frac{B_{21}}{B_{11}} \hat{u}_t^r + \zeta_t, \tag{B.7}$$

where note that $\hat{u}_t^r$ is orthogonal to ζ_t under the assumption that $E[\epsilon^X Z_t'] = 0$.

Finally, we can use the OLS estimates of the matrix A to compute the impulse response functions of all variables to a monetary policy shock using the formula in (B.5).

B.2 Instruments

Figure B.1 reports the monetary policy surprises (aggregated at monthly frequency as described in Section 3) using different contracts, *FSScm1*, *FSScm2*, *FSScm3*, and *FSScm4*, and *GBP3M*. All the monetary policy surprises display a similar behaviour. The largest surprises are concentrated around three events: 1998, 2002 and 2008. Also, the series display higher volatility in the pre-crisis sample, reflecting the fact that monetary policy was constrained by the zero lower bound in the second part of the sample period. The monetary policy surprises, however, display significant variation even in this part of the sample. The similarity between the different

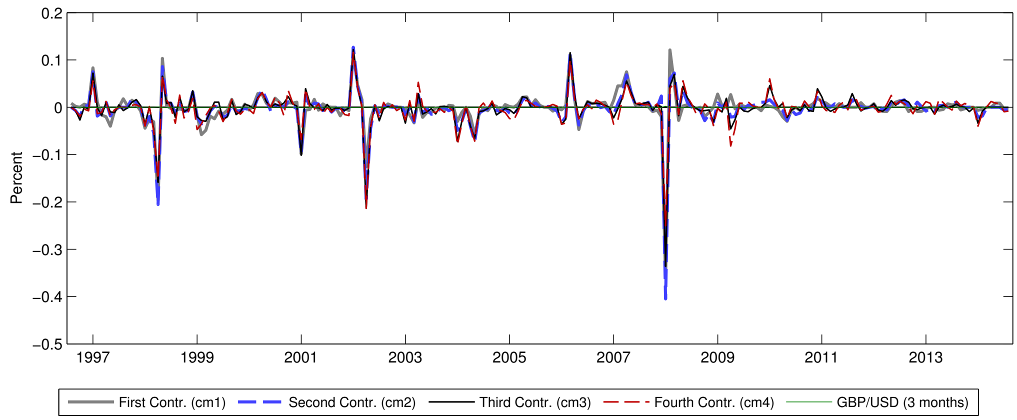

Figure B.1 MONETARY POLICY SURPRISES. *Note.* Each line represent a monetary policy surprise computed with a different contract as explained in Section 3 in the main text.

monetary policy surprises plotted in Figure B.1 is reflected in their correlation. Among all pairs, correlation ranges from a minimum of 0.75 (between *FSScm4* and *GBP3M*) to a maximum of 0.97 (between *FSScm3* and *FSScm4*). The average pairwise correlation (i.e., the average correlation across all pairs) is 0.89. Table B.1 reports the summary statistics for the monetary policy surprises. All series have near-zero mean (between 0 and 0.3 basis points) and a relatively high standard deviation (between 2 and 4 basis points); they are right skewed and display a very high excess kurtosis; and display a small serial correlation that is either positive or negative depending on the monetary policy surprise considered. This is a particularly undesirable feature for a series of arguably 'exogenous' shocks, since any persistence would suggest that the shocks are somewhat predictable. We therefore investigate the statistical significance of those autocorrelation coefficients.

Table B.1 MONETARY POLICY SURPRISES - SUMMARY STATISTICS

	cm1	cm2	cm3	cm4	gbp/usd
Obs	217	217	217	217	217
Mean	0.001	0.003	0.003	0.003	0.000
Max	0.361	0.405	0.336	0.250	0.235
Min	-0.121	-0.127	-0.122	-0.116	-0.049
St. Dev.	0.038	0.041	0.037	0.034	0.020
Auto Corr.	-0.046	-0.032	0.015	0.053	0.003
Skew.	4.359	5.174	4.189	2.911	7.575
Kurt.	44.838	50.643	37.926	23.019	87.779

Note. Summary statistics of the monetary policy surprise (computed with a different contract as explained in Section 3. *Obs* is the number of observations; *Mean* is the sample mean; *Max* is the maximum value; *Min* is the minimum value; *St. Dev.* is teh standard deviation; *Auto Corr.* is teh first lag autocorrelation coefficient; *Skew* is skewness; *Kurt* is kurtosis.

We plot in Figure B.2 the sample autocorrelation function of the monetary policy surprise that we consider in our baseline estimation (*FSScm2*) together with its 95 percent confidence bands (left panel) and its ergodic distribution (right panel). Figure B.2 shows that there is no

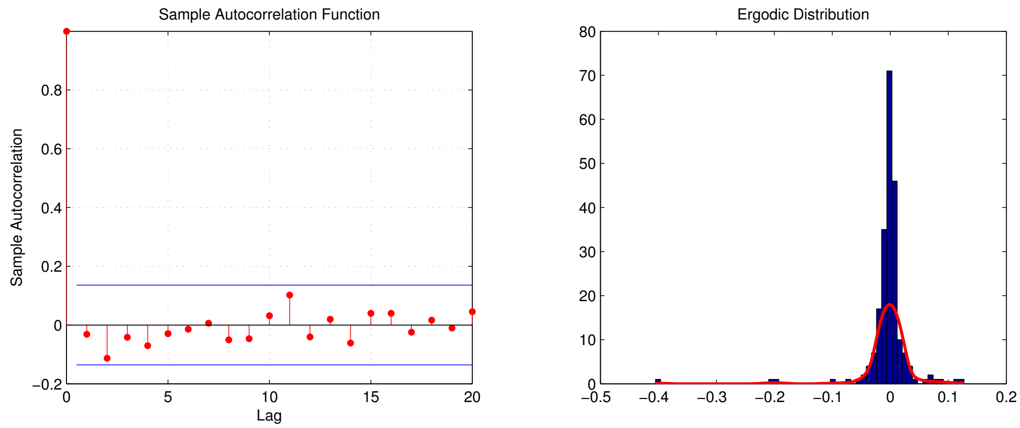

Figure B.2 Monetary Policy Surprise (cm2) - Sample Autocorrelation And Ergodic Distribution. *Note.* The left panel reports the sample autocorrelation function for teh monetary policy surprise compute with the second front contract (cm2), together with 95 percent confidence bands; the right panel plots its ergodic distribution.

statistically significant serial correlation in our series of monetary policy surprises.[36]

Finally, we compare our series of monetary policy surprises with the one constructed by Cloyne and Huertgen (2014). The sample period over which we can compare the two series of monetary policy surprises goes from 1997:6 to 2007:12 — the latest available observation in Cloyne and Huertgen (2014). The two series display quite different behaviour. Indeed the correlation coefficient between the two is extremely low, at -0.006. As shown in Section 5, this somewhat puzzling low correlation simply reflects the fact that our series of shocks and Cloyne and Huertgen (2014)'s capture different information about monetary policy news.

C Appendix. Test of overidentifying restrictions

We start from the relation between the observables and unobservables in equation (16):

$$\begin{bmatrix} z_1 \\ z_2 \\ u^r \\ u^u \\ u^{cpi} \end{bmatrix} = \begin{bmatrix} 1/\alpha_1 & -1/\alpha_1 & 0 & 0 & 0 & 0 \\ 1/\alpha_2 & 0 & -1/\alpha_2 & 0 & 0 & 0 \\ b_{11} & 0 & 0 & 1 & 0 & 0 \\ b_{21} & 0 & 0 & 0 & 1 & 0 \\ b_{31} & 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \epsilon^{mp} \\ \xi_1 \\ \xi_2 \\ \zeta^r \\ \zeta^u \\ \zeta^{cpi} \end{bmatrix},$$

[36]We also checked the monetary policy surprises computed with different contracts. Only the second lag of *FSScm3* and *FSScm4* is statistically significantly correlated with their contemporaneous value at the 95 percent confidence level, while there is no statistically significant association at the 90 percent confidence level.

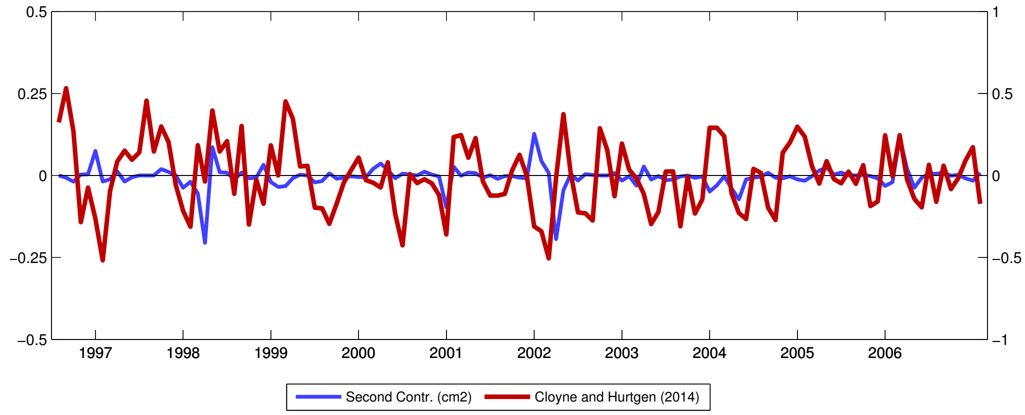

Figure B.3 INSTRUMENTS FOR MONETARY POLICY SHOCK: NARRATIVE AND HIGH FREQUENCY MEASURES (OVERLAPPING SAMPLE PERIOD).*Note.* The red line displays Cloyne and Huertgen (2014)'s instrument for monetary policy shocks (right axis). The blue solid line displays the high-frequency instrument developed in this paper (left axis).

which can be expressed in compact form as:

$$x = \theta \varepsilon.$$

The covariance matrix of the unobservables is given by:

$$Cov \begin{bmatrix} \epsilon^{mp} \\ \xi_1 \\ \xi_2 \\ \zeta^r \\ \zeta^u \\ \zeta^{cpi} \end{bmatrix} = \begin{bmatrix} Var(\epsilon^{mp}) & & & & & \\ Cov(\xi_1, \epsilon^{mp}) & Var(\xi_1) & & & & \\ Cov(\xi_2, \epsilon^{mp}) & Cov(\xi_2, \xi_1) & Var(\xi_2) & & & \\ Cov(\zeta^r, \epsilon^{mp}) & Cov(\zeta^r, \xi_1) & Cov(\zeta^r, \xi_2) & Var(\zeta^r) & & \\ Cov(\zeta^u, \epsilon^{mp}) & Cov(\zeta^u, \xi_1) & Cov(\zeta^u, \xi_2) & Cov(\zeta^u, \zeta^r) & Var(\zeta^u) & \\ Cov(\zeta^{cpi}, \epsilon^{mp}) & Cov(\zeta^{cpi}, \xi_1) & Cov(\zeta^{cpi}, \xi_2) & Cov(\zeta^{cpi}, \zeta^r) & Cov(\zeta^{cpi}, \zeta^u) & Var(\zeta^{cpi}) \end{bmatrix}.$$

Note that:

$$\begin{aligned} Cov(\xi_1, \epsilon^{mp}) &= Cov(\xi_1, z_1 + \xi_1) = \sigma^2_{\xi_1}, \\ Cov(\zeta^r, \epsilon^{mp}) &= Cov(b_{12}\epsilon^2 + ... + b_{1n}\epsilon^n, \epsilon^{mp}) = 0, \\ Cov(\zeta^r, \xi_1) &= Cov(b_{12}\epsilon^2 + ... + b_{1n}\epsilon^n, \xi_1) = 0, \end{aligned}$$

where similar relations hold for other entries of the covariance of unobservables. So, we get:

$$Cov \begin{bmatrix} \epsilon^{mp} \\ \xi_1 \\ \xi_2 \\ \zeta^r \\ \zeta^u \\ \zeta^{cpi} \end{bmatrix} = \begin{bmatrix} 1 & & & & & \\ \sigma^2_{\xi_1} & \sigma^2_{\xi_1} & & & & \\ \sigma^2_{\xi_2} & \sigma^2_{\xi_1 \xi 2} & \sigma^2_{\xi 2} & & & \\ 0 & 0 & 0 & \sigma^2_{\zeta^r} & & \\ 0 & 0 & 0 & \sigma^2_{\zeta^r \zeta^u} & \sigma^2_{\zeta^u} & \\ 0 & 0 & 0 & \sigma^2_{\zeta^r \zeta^{cpi}} & \sigma^2_{\zeta^u \zeta^{cpi}} & \sigma^2_{\zeta^{cpi}} \end{bmatrix}.$$

We have 14 unknowns $\{\alpha_1, \alpha_2, b_{11}, b_{21}, b_{31}, \sigma^2_{\xi_1}, \sigma^2_{\xi_2}, \sigma^2_{\xi_1 \xi_2}, \sigma^2_{\zeta^r}, \sigma^2_{\zeta^u}, \sigma^2_{\zeta^{cpi}}, \sigma^2_{\zeta^r \zeta^u}, \sigma^2_{\zeta^r \zeta^{cpi}}, \sigma^2_{\zeta^u \zeta^{cpi}}\}$. The covariance of the observables gives us 15 moments:

$$Cov \begin{bmatrix} z_1 \\ z_2 \\ u^r \\ u^u \\ u^{cpi} \end{bmatrix} = \begin{bmatrix} Var(z_1) & & & & \\ Cov(z_1, z_2) & Var(z_2) & & & \\ Cov(z_1, u^r) & Cov(z_2, u^r) & Var(u^r) & & \\ Cov(z_1, u^u) & Cov(z_2, u^u) & Cov(u^r, u^u) & Var(u^u) & \\ Cov(z_1, u^{cpi}) & Cov(z_2, u^{cpi}) & Cov(u^r, u^{cpi}) & Cov(u^u, u^{cpi}) & Var(u^{cpi}) \end{bmatrix}.$$

We estimate the parameters of this system with iterative GMM, using Kostas Kyriakoulis' `gmmtbx` toolbox available at https://github.com/tholden/gmmtbx.

We check that the parameters are locally identified at the optimum with the gradient matrix of the moment vector (checking that it is non-singular). We also randomize over starting values for the optimization procedure to ensure we have attained a global optimum. We do that by drawing starting values from a uniform distribution between 0 and 5 with 100 replications. The code quickly converges to a unique minimum.

D Appendix. Additional results

Table D.1 List Of Largest Monetary Policy Surprises

Ranking	Date	Surprise	Event	Description
1	06-Nov-2008	-0.44	MPC rate decision	Bank Rate reduced by 1.5% due to “a sharp slowdown in economic activity”
2	06-Feb-2003	-0.24	MPC rate decision	Bank Rate reduced by 0.25% due to “weaker output than expected”
3	04-Dec-2008	0.19	MPC rate decision	Bank Rate reduced by 1% due to “significant probability of undershooting the inflation target in the medium term”
4	04-Feb-1999	-0.18	MPC rate decision	Bank Rate reduced by 0.5% to “provide a degree of insurance against some of the downward risks” from the international outlook
5	11-Jan-2007	0.17	MPC rate decision	Bank rate increased by 0.25% due to “the world economy was robust, nominal domestic demand was growing strongly and real output growing at least at its potential rate”
6	08-Nov-2001	-0.17	MPC rate decision	Bank rate reduced by 0.50% due to “the prospect of domestic slowdown was largely consequence of the international weakness”

Note. Ranking of the largest monetary policy daily surprises computed using the second front contract of 3-month Sterling future, i.e. the 3-to-6-month ahead expectation about the 3-month Libor.

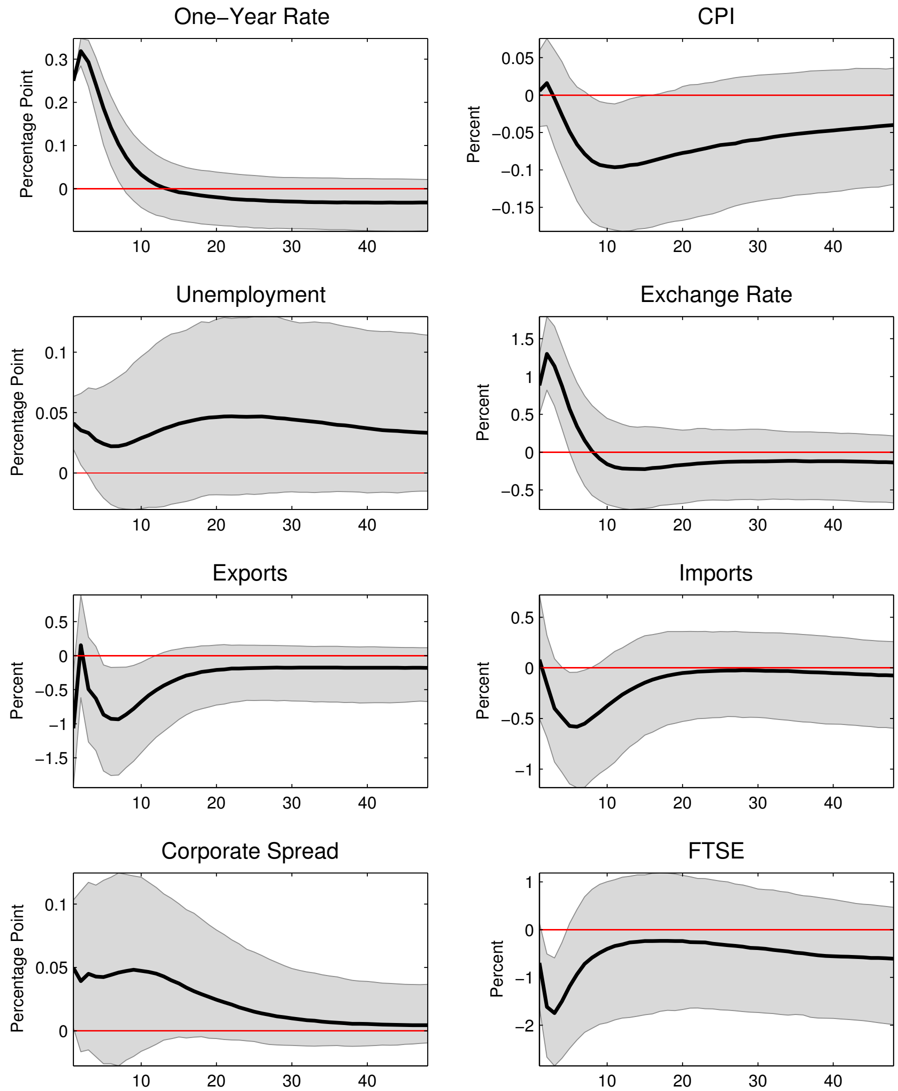

Figure D.1 IRFs To A Monetary Policy Shock - Full Specification With Two Instruments. *Note.* VAR estimated in log levels, with 2 lags, and a constant over the 1993:1-2015:5 period. The VAR includes as exogenous variables a Commodity Price Index and the VIX index. The 1-year Government Gilt Yield is instrumented using the second front contract of 3-month Sterling future combined with Cloyne and Huertgen (2014)'s monetary policy shocks series. First stage results: F-Statistic: 18.65 and $R^2 = 0.07$. The solid line and shaded areas report the mean and the 90% confidence intervals computed using wild bootstrap with 1,000 replications.