# Monetary Policy Transmission in the United Kingdom: A High-Frequency Identification Approach ☆

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**Abstract**

This paper investigates the impact of monetary policy shocks on macroeconomic and financial variables in the United Kingdom using a new series of high-frequency monetary policy surprises. Employing our surprises as an instrument in a monthly SVAR over the UK's inflation-targeting period, we show that a monetary policy tightening induces a decline in economic activity and in CPI, an appreciation of the Pound, a reduction in bank credit, and a significant increase in mortgage and corporate bond spreads. UK monetary policy also affects foreign credit spreads, consistent with the extensive presence of large international players in the UK financial intermediation sector. We finally propose a novel test of overidentifying restrictions, which exploits the availability of the narrative series of monetary policy shocks constructed by Cloyne and Hurtgen (2016), and find that our high-frequency monetary policy surprises are not significantly affected by non-monetary news.

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

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

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Online Supplement available at https://sites.google.com/site/ambropo/CTV_MonPolTransmission_OnlineAppendix.pdf

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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 their policy instruments affect the economy. However, the use of these instruments to stabilize the economy makes it difficult to disentangle cause from effect. A strand of the literature has exploited high frequency changes in futures contracts around key monetary events to isolate monetary policy surprises (see, for example, Kuttner, 2001; Gurkaynak et al., 2005a; Gurkaynak et al., 2007). Focusing on a tight window around these events helps to isolate monetary policy news from other types of shocks.

In this paper we revisit the important question of how monetary policy transmits to the broader economy using a high-frequency identification approach for the United Kingdom (UK). We construct what are, to our knowledge, the first series of high-frequency monetary policy surprises for the UK.[1] Following the methodology developed by Stock and Watson (2012) and Mertens and Ravn (2013), we then use these surprises as an external instrument for monetary policy shocks in a structural VAR to identify their causal effects on macroeconomic and financial variables. Finally, exploiting the availability of the narrative series of monetary policy shocks developed by Cloyne and Hurtgen (2016), we propose a new test of overidentifying restrictions to assess the potential information content of the high-frequency monetary policy surprises.

Our approach has two main advantages with respect to previous studies for the UK. First, the high-frequency approach to the identification of monetary policy surprises allows us to analyze simultaneously the dynamic response of both financial and macroeconomic

[1] The high-frequency approach to identification 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. In independently conducted and contemporaneous work to ours Miranda-Agrippino (2016) and Gerko and Rey (2017) also compute series of monetary policy surprises for the UK using a high-frequency approach. Li and Zanetti (2016) compute UK monetary surprises using the change in a market interest rate between two consecutive days instead of *intra-daily* changes in future contracts.

variables without imposing implausible timing restrictions. This is important because financial variables are a crucial determinant of the transmission of monetary policy.[2] Second, since central banks have been increasingly relying on communication to influence markets' beliefs about the evolution of the short-term interest rate, it is important to control for private sector expectations. By focusing on variations in futures contracts around monetary policy events, we capture markets' reaction to the unexpected component of monetary policy.

In line with the existing literature, we document that a tightening of monetary policy has effects on the real economy, raising unemployment and depressing economic activity; as well as appreciating the domestic currency and inducing a significant reduction in CPI. Finally, in line with models with credit frictions (e.g., Bernanke and Gertler, 1995), a monetary policy tightening leads to an increase of both mortgage and corporate bond spreads, leading to a contraction in total credit.

Our analysis also reveals one new finding relative to the existing literature. We show that corporate bond spreads in the US respond significantly to UK monetary policy shocks. This result points to an important dimension of monetary policy that is often overlooked, i.e. the fact that the monetary policy of a relatively small country such as the UK, comprising about 3 percent of world GDP, can affect foreign financial conditions. This finding is not surprising if one takes into account a key feature of the UK financial sector: the extensive presence of large international players such as, for example, foreign pension and mutual funds, as well as global banks. Indeed, using different data and a different methodology to ours, Cerutti et al. (2017) show that UK banking conditions (as summarized by the TED spread) are an important determinant of international capital flows. Moreover, our finding echoes and complements those of Rogers et al. (2018) on the spillover of UK unconventional monetary policy on US government bond rates.

[2] See, for example, Gertler and Karadi (2015) and Caldara and Herbst (2018).

The statistical inference that we conduct is reliable under the assumptions that the monetary policy surprises are exogenous to the developments in the macroeconomy. The short window we consider around policy announcements rules out the possibility that other non-monetary news (such as demand or supply) might affect our monetary policy surprises. When information frictions are present, however, a 'signalling channel' of monetary policy can arise: central bank announcements can simultaneously convey information about monetary policy and the central bank's assessment of the economic outlook (see Romer and Romer, 2000; Melosi, 2017). In this case, market based monetary surprises as the ones constructed in this paper are a combination of two components: (i) the monetary policy news, i.e. the component that we want to isolate; and (ii) a component that captures the release of information from the central bank to financial market participants about the current (or future) state of the economy. The presence of this non-monetary news in our high-frequency surprises would therefore invalidate the assumptions under which we conduct statistical inference, leading to a bias in the estimated impact of monetary policy on financial and macroeconomic variables.

Recent studies on US data have shown that such a signalling component can be sizable in high-frequency market-based surprises around policy announcements by the Federal Reserve.[3] But the presence of this 'information bias' is ultimately an empirical question. Indeed, it depends on the sample period, the financial asset used to compute the surprises, the type of monetary events considered (scheduled vs. unscheduled meetings), etc.[4] In the last part of the paper we provide statistical evidence in support of the assumption that our high-frequency monetary policy surprises are not significantly affected by releases of macroeconomic information. In particular, we test this assumption by exploiting the

[3]See Barakchian and Crowe (2013); Gertler and Karadi (2015); Campbell et al. (2016); Ramey (2016); Nakamura and Steinsson (2018); Miranda-Agrippino and Ricco (2017); Jarocinski and Karadi (2018)

[4]For example, Gertler and Karadi (2015) and Ramey (2016) both show that the high-frequency surprises in the US can be predicted using Fed Staff Greenbook forecasts. However, Caldara and Herbst (2018) show that this result does not hold if one excludes 'unscheduled' meetings from the list of events.

availability of an alternative measure of monetary policy innovations that explicitly controls for the information set of the central bank—the narrative series computed by Cloyne and Hurtgen (2016) following Romer and Romer (2004)'s approach. We combine our monetary surprises with the narrative series of UK monetary policy shocks in a test of overidentifying restrictions. The results of this test suggests that our results are not subject to a statistically significant bias due to the presence of information effects.

The estimated effects on activity are consistent with previous studies for the UK Mountford (2005), Cloyne and Hurtgen (2016), Li and Zanetti (2016), Gerko and Rey (2017); and in line with those of Gertler and Karadi (2015) for the US. Relative to these studies, we provide novel evidence about the transmission of monetary policy in the UK via a domestic credit channel (through domestic credit spreads and bank credit), but also via an international credit channel (through the US corporate bond spreads). Finally, the novel test of overidentifying restrictions, which exploits the narrative series of monetary policy shocks for the UK, sheds light on the content of the identified monetary policy surprises.

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 describes the econometric framework and the identification strategy. Section 4 shows their impact on macroeconomic and financial variables in a structural VAR. Section 5 tests the validity of our instrument through test of overidentifying restrictions. Section 6 concludes.

## 2 Monetary Policy Surprises for the UK

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).[5] Specifically, we construct a new data set using intra-day 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. The term 'event' refers 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.

Since being granted operational independence in June 1997, the MPC has set the Bank Rate to achieve its inflation target.[6] A liquid contract based on the Bank Rate would therefore be the most appropriate contract to compute the surprises. However, and unlike the case of the Fed Funds for the United States, there is no futures market based on this interest 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).[7] In particular, in a given year, there are four delivery dates at the end of the following months: March, June, September, and December.[8,9]

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[5] We would like to thank Georgios Georgiadis, Johannes Grab, and Jonas Jensen for helping us spotting a mistake in the construction of the monetary surprises used in the working paper version of this study.

[6] A short review of the monetary policy framework in the UK is reported in the Appendix A.

[7] 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 is not always available. Appendix A describes in detail all the contracts available for the UK and their characteristics.

[8] 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.

[9] 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 A provides more information

We therefore measure 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:

$$Z_t^{HF} = -\left(P_{t,\tau+20}^h - P_{t,\tau-10}^h\right) = \mathbb{E}_{(t,\tau+20)}\left[i_h^{(h+90)}\right] - \mathbb{E}_{(t,\tau-10)}\left[i_h^{(h+90)}\right], \tag{2}$$

where $t,\tau$ 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 $h$. $\mathbb{E}_{(t,\tau+20)}\left[i_h^{(h+90)}\right]$ 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 $\tau$ (i.e., $t,\tau+20$). 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 $P^h$. The second equality, expressed in terms of expected interest rate, is derived using equation (1). 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.[10]

The largest surprises identified using this contract coincide with important monetary

about these contracts.

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

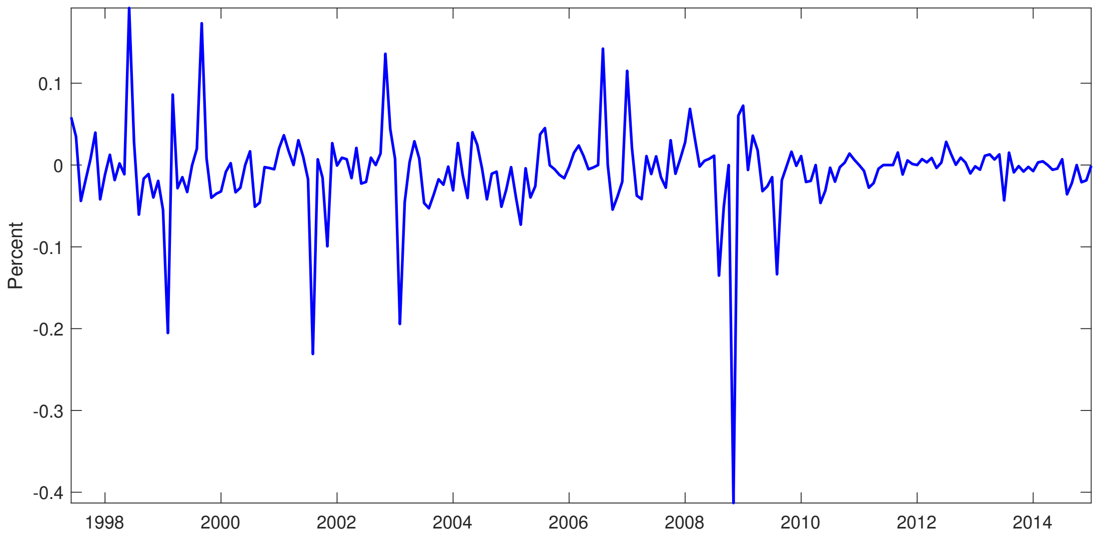


**Figure 1** DAILY MONETARY POLICY SURPRISES. *Note.* Daily monetary policy surprises computed using the second front contract of the 3-month Sterling future, i.e. the 3-to-6-month ahead expectation about the 3-month Libor. The surprises are computed using a 30 minutes window around the identified monetary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

policy events (see Table S1 in the Online Supplement). There is a clear change in the volatility of the monetary policy surprises after March 2009, when Bank Rate reached 0.5 percent (until February 2015 this was considered to be the effective 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 effective 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 is highly correlated with more standard measures of monetary news based on the UK yield curve. Since the UK monetary events in our sample do not overlap with US ones, 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 Estimation and Identification

In order to identify the transmission of monetary policy shocks to the real economy, we use the external instruments identification approach proposed by Stock and Watson (2012) and Mertens and Ravn (2013). This identification strategy uses standard instrumental variable techniques to isolate the variation in the VAR reduced-form residuals that is due to the structural shock of interest.

## 3.1 Proxy-SVAR

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

$$Y_t = AY_{t-1} + u_t, \tag{3}$$

where $Y_t$ is a $(m \times 1)$ vector of endogenous variables; $A$ is an $(m \times m)$ matrix of coefficients; and $u_t$ are an $(m \times 1)$ vector of reduced-form innovations with covariance matrix $\Sigma_{uu'}$. These innovations are a linear combination of the structural shocks:

$$u_t = B\varepsilon_t, \tag{4}$$

where $\varepsilon_t$ is an $(m \times 1)$ vector with variance-covariance matrix $\Sigma_{\varepsilon\varepsilon'} = I$ and $B$ is an $(m \times m)$ matrix that contains the contemporaneous effects of the structural shocks on $Y_t$. Since the structural parameters are not identified from (3), to recover the structural shocks we need to impose some identification restrictions.

If we partition the vector of endogenous variables $Y_t$ as $(r_t', X_t')'$—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 SVAR as:

$$\begin{bmatrix} r_t \\ X_t \end{bmatrix} = \begin{bmatrix} A_{11} & A_{12} \\ A_{21} & A_{22} \end{bmatrix} \begin{bmatrix} r_{t-1} \\ X_{t-1} \end{bmatrix} + \begin{bmatrix} B_{11} & B_{12} \\ B_{21} & B_{22} \end{bmatrix} \begin{bmatrix} \varepsilon_t^r \\ \varepsilon_t^X \end{bmatrix}, \tag{5}$$

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 $\varepsilon_t^r$ and $\varepsilon_t^X$ are the structural shocks associated to monetary policy and the remaining endogenous variables, respectively.

Let $u_t^r$ and $u_t^X$ be the OLS estimates of the partitioned reduced form residuals of equation (3). Also, let $Z_t$ be a $(z \times 1)$ vector of external instruments that satisfies:

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

i.e., the external instruments are correlated with the monetary policy shock ($\varepsilon_t^r$) but are orthogonal to all the other shocks (the elements of $\varepsilon_t^X$). Then, we can obtain a consistent estimate of $B_{21}/B_{11}$ by two-stage least squares. In the first stage, we estimate:

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

to construct the fitted values $\hat{u}_t^r$. Then we regress the reduced form residuals of the remaining equations ($u_t^X$) on the fitted values ($\hat{u}_t^r$):

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

where note that $\hat{u}_t^r$ is orthogonal to $\zeta_t$ under the assumption that $E[\varepsilon^X Z_t'] = 0$.[11] Finally,

[11]See Section 5 for more details about this assumption.

we can use the OLS estimates of the matrix $A$ from equation (3) to compute the impulse response functions of all variables to a monetary policy shock.

## 3.2 Inference

The most popular method for computing the confidence bands in Proxy-SVAR is the wild bootstrap (see, for example, Mertens and Ravn, 2013; Gertler and Karadi, 2015). Jentsch and Lunsford (2019a) and Jentsch and Lunsford (2019b) show that wild bootstrap is, in general, not asymptotically valid for inference about estimators that involve the covariance matrix of VAR innovations identified with external instruments, and may understate the true estimation uncertainty in finite samples. They propose a variant of the moving block bootstrap described in Bruggemann et al. (2016) as an alternative inference approach in proxy-identified SVARs. In our baseline results we report confidence intervals computed using the moving block bootstrap (see Jentsch and Lunsford, 2019b). Following Mertens and Ravn (2019), in the Online Supplement we assess the robustness of our conclusions by reporting confidence intervals based on three alternative methods: the Delta method and the parametric bootstrap (Montiel Olea et al., 2018). For comparison, we also report the results obtained with the wild bootstrap in the Online Supplement.

## 3.3 Baseline Specification and Data

We estimate the VAR using monthly data for the UK for the period 1992:1–2015:1. The starting point coincides with the beginning of the inflation targeting regime in the UK. Immediately prior to this (i) the British Pound was essentially shadowing the Deutsche Mark and (ii) the target and operating framework for monetary policy were very different. Thus, a sample starting before that date will likely be affected by a structural break.[12]

[12] For comparison with Gerko and Rey (2017), the Online Supplement displays the IRFs for the sample 1982:1-2015:1.

In our baseline specification we select a minimal set of endogenous variables that allows us to consider the transmission channels of monetary policy.[13,14] Specifically, we consider the consumer price index as a measure of prices; the unemployment rate, as a measure of economic activity; the nominal effective exchange rate, a key variable for the transmission of monetary policy in an open economy like the UK; and the mortgage and (investment grade) corporate bond spreads as our empirical proxy for credit costs in mortgage and corporate bond markets. Clearly, our data includes the recent crisis and its aftermath, where the Bank Rate—the 'typical' monetary policy indicator until the global financial crisis—did not move from the level of 50 basis points reached in 2009. To address this issue and to capture shifts in expectations about monetary policy, we choose as policy indicator a safe interest rate at slightly longer maturity than is typically considered in monetary VARs, namely the nominal yield on the 1-year gilt.[15]

In order to correctly specify our VAR, there are two additional issues that we need to consider. First, the relatively small scale of our model poses a question of whether all relevant information is included in the specification. It is well known that informationally deficient VARs lead to severe biases in the estimation of the transmission of shocks.[16] Second, Akinci (2013) shows that controlling for global financial conditions in small open economy VARs is important since they constitute an important source of business cycles fluctuations. This is particularly important for an international financial center like the UK where global investors are active in the intermediation of both domestic and foreign credit.

[13] For the variables for which data is available at higher frequency, we compute monthly averages.

[14] In Section 4.2 we report the result from extended specifications, where we augment our baseline specification with additional variables.

[15] We check the robustness of our results by (i) using longer maturity gilts as a policy indicator and (ii) excluding the period over which Bank rate did not show any time variation. Both figures are reported in the Online Supplement.

[16] For example, Bernanke and Mihov (1998) use a factor augmented VAR so as to include all relevant information in their empirical model. Caldara and Herbst (2018) show that the inclusion of credit spreads is crucial for the correct identification of monetary policy shocks, as monetary policy systematically responds to financial conditions.

For these reasons, we add to our baseline specification a measure of US corporate bond spreads, the spread between the yield on Moody's BAA corporate bond index and the yield on US 10-year government bonds. The predictive power of corporate bond spreads for economic activity is well known (see Gilchrist and Zakrajsek, 2012; and references therein). Moreover, as most of (both bank and market-intermediated) international credit is extended in US dollars (see, for example, Bruno and Shin, 2014, 2017), this US measure can also be considered as a global measure of credit market conditions. However, the main conclusions from the analysis do not depend on the inclusion of this variable.[17]

Following Sims et al. (1990), we estimate the VAR systems in levels without explicitly modeling the possible cointegration relations among them.[18] We use the information criteria to choose the optimal number of lags, which we set to two.[19] Moreover, the results are robust to different lag specifications, as detailed in the Online Supplement.

Finally, we employ the series of monetary policy surprises, described in Section 2, as the instrument for the monetary policy shocks. We aggregate the monetary policy surprises at monthly frequency following the procedure developed by Gertler and Karadi (2015).[20] 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 daily change in the 1-year gilt yield. In our case, this is the second front contract of 3-month Sterling future.[21] Our monetary policy surprises are available only for a sub-sample of the period over which the VAR is estimated, namely from 1997:6 to 2015:1. We choose a longer sample period for the estimation of the VAR so as to estimate with greater precision the lag coefficients

[17]The Online Supplement presents the IRFs excluding the US BAA Corporate Spread

[18]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.

[19]According to AIC and BIC the optimal number of lags is 1 while for the HQC is 2. Reduced form residuals of the VAR are not serially correlated only when we consider 2 lags.

[20]The time series properties of the aggregated monetary surprises are reported in the Online Supplement, together with their correlation. Table S2 and Figure S2 show that the monthly series of surprises are not autocorrelated. The Online Supplement also shows that the IRFs remain unchanged if we simply sum of the surprises within a given month.

[21]We get a similar F-statistic using the other front contracts of 3-month Sterling future.

and the reduced form residuals.

As we discuss in more detail below, one 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 (see, for example, Campbell et al., 2012; Campbell et al., 2016; Melosi, 2017; Nakamura and Steinsson, 2018). For example, a surprise tightening of monetary policy could be taken to indicate an improvement in the outlook for the macroeconomy. In this case, our monetary surprises would be correlated with non-monetary news about the economy; and the estimated impact of monetary surprises on macroeconomic and financial variables would be biased. To provide evidence that our surprises are not significantly affected by this issue, in Section 5 we conduct a test of overidentifying restrictions using another series of monetary innovations that explicitly controls for the Bank of England's private information set (Cloyne and Hurtgen, 2016). As we shall show, we cannot reject the null hypothesis that our high-frequency series of monetary policy surprises does not contain non-monetary information.

# 4 The Transmission of Monetary Policy

In this section we present the empirical results of our analysis. We start with the baseline specification described above, which includes standard macroeconomic variables and credit costs in mortgage and corporate bond markets. We then consider four augmented specifications that include a monthly estimate of GDP, a measure of credit quantities, equity prices, and the trade balance, respectively.

## 4.1 Baseline Results

Figure 2 displays the impulse response functions (IRFs) to an instrumented 25 basis points increase in the 1-year gilt rate, using as an instrument the second front contract of 3-month Sterling future. The instrument is powerful in capturing the variation in the reduced form residuals of the policy indicator. The robust F-statistic from the first stage is in fact 40.3, well above the relevant threshold of 10 suggested by Stock and Yogo (2002).[22] The $R^2$ of the first stage regression is 0.12.

We start by analyzing the response of the policy rate, unemployment, and consumer prices. The shock has a persistent effect on the 1-year gilt yield, lasting for about eight months after the shock hits. It induces a statistically significant and delayed increase in unemployment. Over the horizon considered in the IRFs, the increase in unemployment reaches a maximum of 0.10 percentage points. In line with the fall in economic activity, consumer prices fall around 0.1 percent ten months after the shock.

We now turn to credit spreads, and more narrowly to the credit channel of monetary policy. The impulse responses in Figure 2 show that both corporate bond and mortgage spreads increase on impact by about 10 basis points. The response of the mortgage spread is increasing for a few months after the shock, with a peak after about half a year. Credit spreads gradually decline back to their long-run values after about 18 months. These impulse responses provide support to the view monetary policy operates through a domestic credit channel (Bernanke and Gertler, 1995). According to this view, a tightening of monetary policy that reduces the net worth and liquidity of borrowers would increase the effective cost of credit by more than the change in risk-free rates, thus intensifying the effect of the policy action.

But UK monetary policy does not work only through a domestic credit channel. In

[22] Montiel Olea et al. (2018) show that the heteroskedasticity-robust first-stage F-statistic is the measure that can be compared to the Stock and Yogo (2002) critical values.

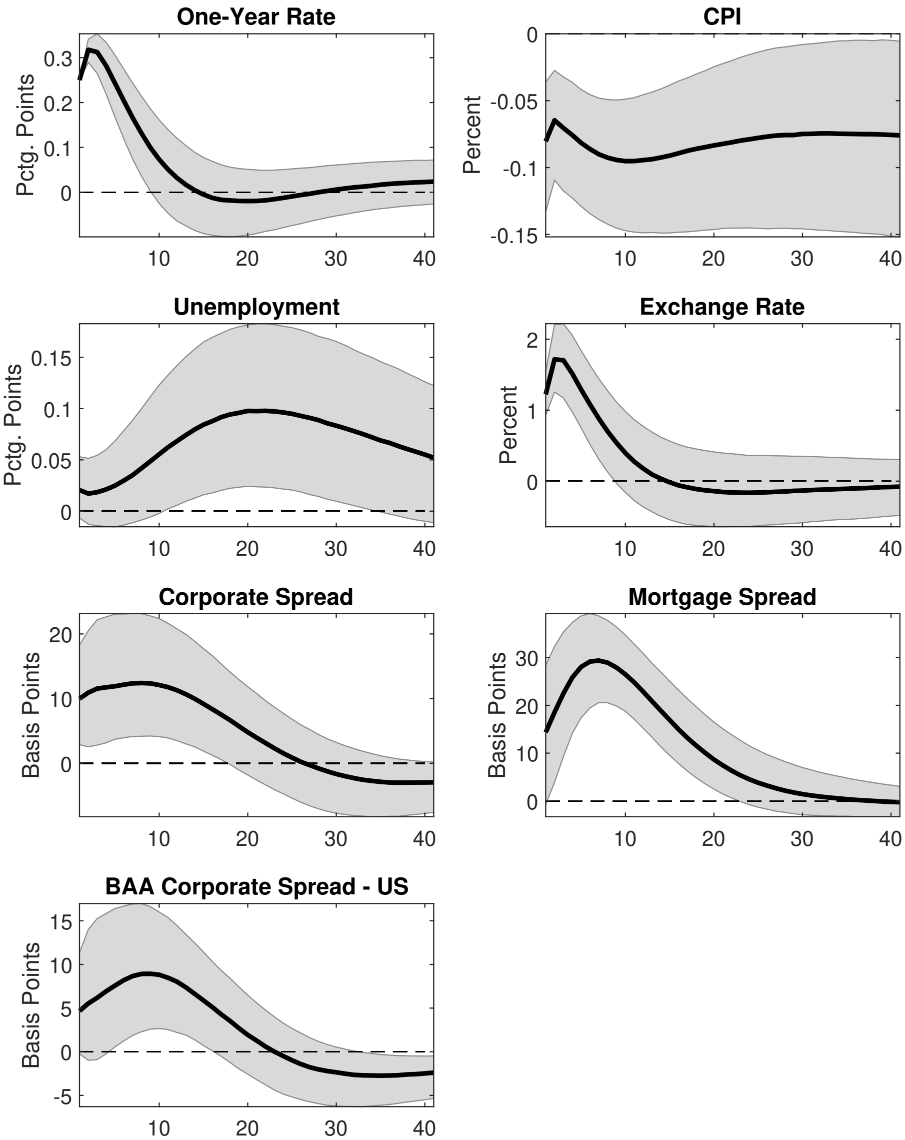


**Figure 2** IRFs To A Monetary Policy Shock - Baseline Specification. *Note*. VAR estimated in log levels, with 2 lags, and a constant over the period 1992:1-2015:1. The 1-year Government Gilt Yield is instrumented using the second front contract of 3-month Sterling future. First stage results: F-Statistic: 40.3 and $R^2 = 0.12$. The solid lines and shaded areas report the mean and the 68% confidence intervals computed using moving block bootstrap with 5,000 replications.

fact, US corporate bond spreads also responds to the UK monetary policy shock, with a statistically significant response of about 8 basis points in the first few months after the shock hit. This new empirical finding on the international financial spillovers of UK monetary policy complements the findings of Gerko and Rey (2017)—who show that UK monetary policy affects the VIX—and challenge the small open economy assumption that is typically made in the study of international transmission of shocks for the UK.[23]

The effect on the US corporate spread is not surprising given extensive presence of large international players in the UK financial intermediation sector (such as foreign pension and mutual funds, as well as global banks, for example). It is well known that monetary policy affects the risk bearing capacity of financial intermediaries (see, for example, Gertler and Karadi, 2015), leading to countercyclical movements in credit spreads. When these financial intermediaries operate across borders, the same effect should be observed in other countries, as we see in our impulse responses. Finally note that our results complement those in Cerutti et al. (2017), who show that banking conditions in the UK affect international capital flows (a key manifestation of the 'global financial cycle'), and those in Rogers et al. (2018), who show that UK monetary policy significantly affects US government bond yields.

## 4.2 Additional Results

In this section, we consider alternative VAR specifications that help us to characterize the transmission of monetary policy in the UK. Specifically, we augment our baseline VAR with a monthly estimate of real GDP; a monthly series of total (real) credit extended by monetary and financial institutions; a stock market index; and the trade balance. For ease of exposition, we report here only the responses of these four variables (see Figure 3), while the full set of impulse responses is reported in the Appendix B. As in our baseline, for each specification we consider a 25 basis points instrumented increase in the 1-year gilt rate, using as an instrument the second front contract of 3-month Sterling future. In all cases, the robust F-statistic from the first stage is significantly larger than the threshold of 10 suggested by Stock and Yogo (2002).

[23]The response of corporate spreads remain unchanged if we add the VIX to our baseline specification. The IRFs are reported in the Online Supplement.

**Real GDP**. We consider first the specification augmented with the monthly estimate of real GDP provided by the National Institute of Social and Economic Research (NIESR), which is computed by combining data on the industrial production and on the other sectors of the economy.[24] As industrial production accounts for only 20 percent of GDP, this is arguably a better and more representative indicator of economic activity in the UK than industrial production. The impulse response of real GDP (reported in Figure 3, top-left panel) shows that the increase in unemployment documented in the previous section is also coupled with a fall in GDP. In particular, real GDP barely moves on impact and then slowly decreases, with a statistically significant peak response of $-0.35$ percent after about 2 years.

**Credit**. If monetary policy affects the risk bearing capacity of the marginal investor pricing corporate bonds and, therefore of the credit supply in bond markets, it may also influence the broader supply of credit (Gilchrist and Zakrajsek, 2012). Then, in the face of an unexpected tightening of monetary policy, we should observe a reduction in total credit extended to the private sector. In our second experiment, we therefore add to the baseline VAR of the previous section a measure of total credit extended by financial institutions to the private non-financial sector. This monthly credit aggregate is deflated with the CPI index. The top-right panel of Figure 3 reports the impulse response of total credit. Similarly to GDP, total credit does not respond significantly on impact, but then slowly and persistently decreases over time. At the end of the horizon considered, the monetary policy shock induces a contraction in total credit of about 1 percent. Thus, the transmission through the credit channel works though rates and total credit.

**Equity prices**. In our third experiment we analyze the response of stock prices by including the FTSE index to our baseline VAR. The bottom-left panel of Figure 3 shows that equity prices contract by more than 0.5 percent, but in a non-significant fashion.

[24] See https://www.niesr.ac.uk/sites/default/files/files/GDP%20Spreadsheets/dp127.PDF for more information on how this index is computed.

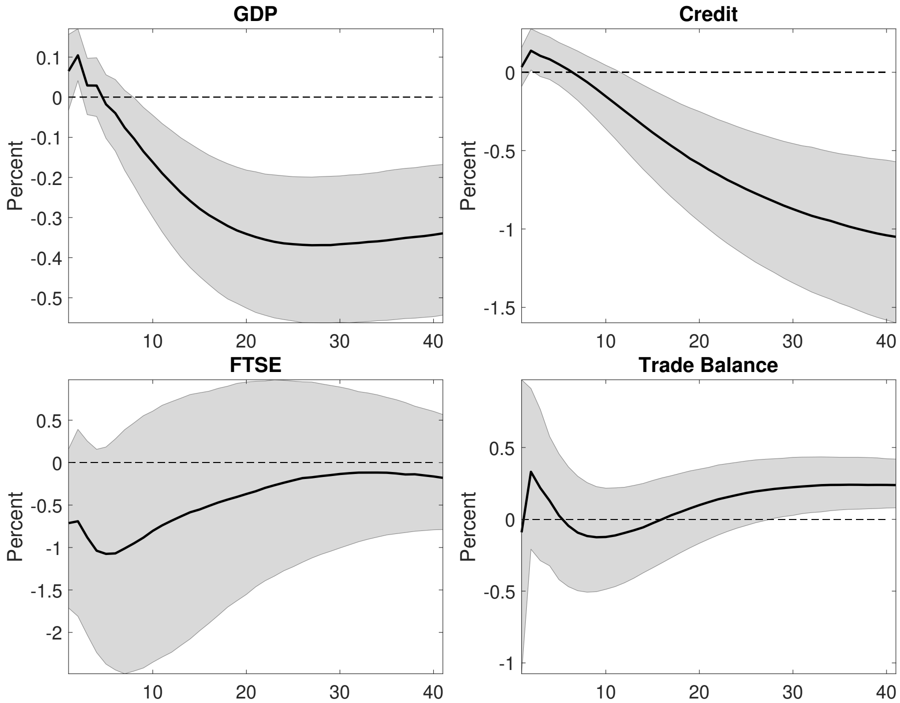


**Figure 3** IRFs To A Monetary Policy Shock - Augmented Specification. *Note.* VAR estimated in log levels, with 2 lags, and a constant over the 1992:1-2015:1 period. The VAR includes the seven variables of the baseline specification plus each of the four variables included in this figure. The 1-year Government Gilt Yield is instrumented using the second front contract of 3-month Sterling future. The F-Statistic of the first stage and the responses of the remaining variables of the system are presented in Appendix B. The solid lines and shaded areas report the mean and the 68% confidence intervals computed using moving block bootstrap with 5,000 replications.

**Trade balance**. Finally, we add the trade balance to our baseline specification to analyze the effects on the foreign sector. Despite the persistent appreciation of the Pound, the monetary policy shock does not induce any significant effect on trade balance. This result is driven by a fall in both exports and imports, which leave the trade balance unchanged.[25]

For robustness, we also estimate the baseline VAR (i) with a set of monetary policy surprises that exclude extraordinary MPC meetings, which are more likely to include information effects (see Nakamura and Steinsson, 2018; Caldara and Herbst, 2018); (ii) with a set of monetary policy surprises that exclude releases of the Inflation Report, which con-

[25] These results are reported in the Online Supplement.

tains forecasts about the evolution of macroeconomic variables. The responses in both specifications, reported in the Online Supplement, remain unchanged. These findings suggest that there is no significant release of information in our monetary policy surprises. In section 5 we address this issue more formally with a statistical test for this hypothesis.

Summing up, monetary policy has a significant and persistent effect on macroeconomic variables and affects the economy both through the increase in the corporate spread and mortgage spread and the decline in total credit. In terms of the effect on activity and inflation, Table 1 contains a detailed comparison with previous findings. Our estimates are broadly in line with these studies for the US and the UK. However, we provide novel evidence on the importance of both corporate and mortgage interest rate spreads and total credit to characterize the effects of monetary policy shocks. Moreover, we also provide novel evidence on the international financial spillovers of UK monetary policy.

# 5 The Role of Information Effects in High-Frequency Monetary Surprises

A key condition for the estimates in the previous sections to be consistent is that our high-frequency instrument, $Z_t^{HF}$, is uncorrelated with the non-monetary innovations in the system ($\varepsilon_t^X$). As argued above, the selection of a short window around policy events should be sufficient to satisfy this condition, as it is unlikely that non-monetary innovations systematically happen in the 30 minutes surrounding the policy announcement. But it is still possible that the unexpected component of policy decisions contains news about the determinants of monetary policy. When information frictions are present, a 'signalling channel' of monetary policy can arise: central bank announcements can simultaneously convey information about monetary policy and the central bank's assessment of the economic outlook (see Romer and Romer, 2000; Melosi, 2017). If this was the case, the high-frequency mon-

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

| Authors | Country | Method | Peak Effects (in %) | |
|---|---|---|---|---|
| | | | Activity | 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 Hurtgen (2016) | UK | Narrative | -0.5 (IP) | -1.0 (CPI Infl) |
| Gerko and Rey (2017) | UK | Proxy-SVAR | -1.8 (IP) | 1.0 (RPIX) |
| This paper | UK | Proxy-SVAR | 0.4 (Unempl)<br>-1.4 (GDP) | -0.4 (CPI) |

*Note.* This table is an update of Table reported in Cloyne and Hurtgen (2016). It shows the response of prices and indicators of economic activity to a one percentage point increase in the interest rate. The peak effects correspond to a one percentage point increase in the interest rate. In brackets we include the specific measure of economic activity and prices considered in each study. CPI Infl denotes CPI inflation.

etary surprises ($Z_t^{HF}$) would be a linear combination of 'true' monetary policy shocks and 'signalling' shocks (also labelled as information shocks in the literature). As the latter are ultimately correlated with the non-monetary shocks the central bank systematically responds to (i.e., $\varepsilon_t^X$), their presence clearly represents a violation of the key exclusion restriction — namely, $\mathbb{E}(\varepsilon_t^X Z_t^{HF\prime}) \neq 0$.

Recent studies on US data have shown that such a signalling component can be sizable in high-frequency market-based surprises around policy announcements by the Federal Reserve.[26] But both the presence and the strength of this signalling component is ultimately

[26] See Barakchian and Crowe (2013); Gertler and Karadi (2015); Ramey (2016); Miranda-Agrippino and Ricco (2017); Jarocinski and Karadi (2018).

an empirical question and depends, among other things, on the sample period, the financial asset used to compute the high-frequency surprises, the type of monetary events considered (scheduled vs. unscheduled MPC meetings), etc. (see, for example, Caldara and Herbst, 2018).

In this section we exploit the availability of a complementary measure of UK monetary policy shocks — the narrative measure constructed by Cloyne and Hurtgen (2016) following the methodology proposed by Romer and Romer (2004), which we label $Z_t^N$ — to test whether our high frequency instrument ($Z_t^{HF}$) is contaminated by information effects with a test of overidentifying restrictions.[27] We exploit one particular feature of the narrative monetary policy shocks, namely that they are constructed as the intended changes in the policy rate purged of the discretionary changes justified by the Central Bank's assessment of current and expected macroeconomic conditions. In other words, $Z_t^N$ explicitly controls for the Central Bank's information set.

We proceed in two steps. First, we employ Cloyne and Hurtgen (2016)'s methodology to extend their original series (which covers the period 1975:1–2007:12) to 2009:2, i.e. the month the policy rate in the United Kingdom hit the Zero Lower Bound. Figure 4 displays our measure of monetary policy surprises together with the narrative series of monetary policy shocks. Over this period, the two series display some slight comovement, with a contemporaneous correlation of 0.16.[28]

Second, we construct a test of overidentifying restrictions. Consider the VAR specifica-

[27] In subsequent work to ours, Angelini and Fanelli (2019) use a similar methodology to test the empirical validity of a proxy-SVAR model for uncertainty shocks.

[28] The contemporaneous correlation between US narrative monetary policy shocks and US monetary policy surprises has a similar magnitude. 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 surprises used.

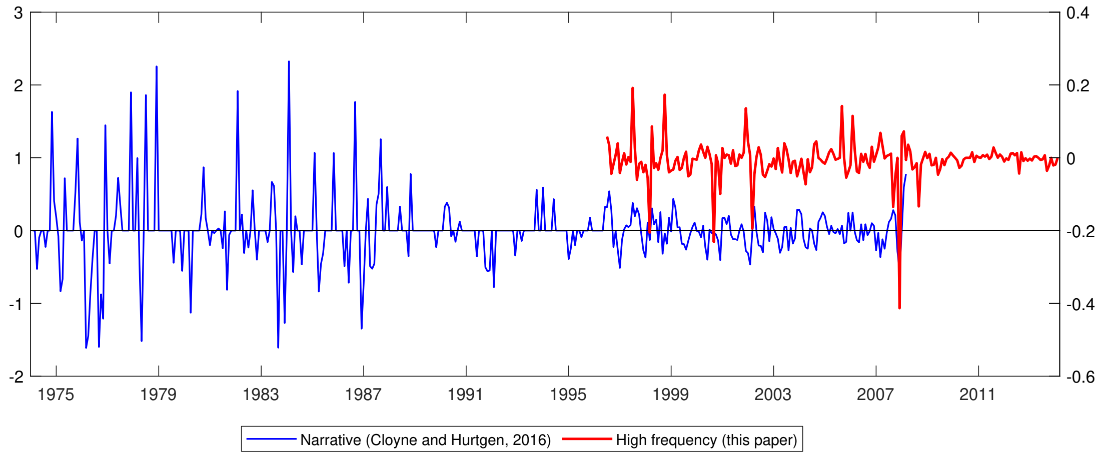


**Figure 4** Instruments For Monetary Policy Shock: Narrative And High-Frequency Measures. *Note.* The blue line displays Cloyne and Hurtgen (2016)'s instrument for monetary policy shocks (left axis). The red line displays the high-frequency instrument developed in this paper (right axis).

tion presented in equation (5). The reduced form residuals can be expressed as:

$$u_t^r = B_{11}\varepsilon_t^r + B_{12}\varepsilon_t^X \tag{8}$$

$$u_t^X = B_{21}\varepsilon_t^r + B_{22}\varepsilon_t^X \tag{9}$$

Then, combining equations (8) and (9), we obtain:

$$u_t^X = B_{21}B_{11}^{-1}u_t^r + \underbrace{\left(B_{22} - B_{21}B_{11}^{-1}B_{12}\right)\varepsilon_t^X}_{\equiv V_t} \tag{10}$$

where $V_t$, the error term of regressing $u_t^X$ on $u_t^r$, is a linear combination of the non-monetary structural shocks. Now, consider the vector of instruments $Z_t = (Z_t^{HF\prime}, Z_t^{N\prime})'$. Differently from the previous Section, $Z_t$ now includes both the high frequency monetary policy surprises and the narrative series of monetary policy shocks. In this case, $B_{21}B_{11}^{-1}$ is overidentified. If the instruments are orthogonal to the non-monetary policy structural shocks (i.e.

$\mathbb{E}\left(\varepsilon_t^X Z_t'\right) = 0$), then the following expression must hold:

$$\mathbb{E}\left(V_t Z_t'\right) = \mathbb{E}\left(\left(\left(B_{22} - B_{21}B_{11}^{-1}B_{12}\right)\varepsilon_t^X\right) Z_t'\right) = 0. \tag{11}$$

We can test that condition (11) holds using a J-Test. Given the above discussion on the properties of $Z_t^N$, it follows that if $Z_t^{HF}$ was a combination of $\varepsilon_t^X$ and $\varepsilon_t^r$ (as it would happen in the case of a signalling channel being at work), we would reject the null hypothesis of $\mathbb{E}\left(V_t Z_t'\right) = 0$.

We proceed as follows. We consider the narrative and high-frequency instruments ($Z_t^N$ and $Z_t^{HF}$) over the period where they overlap, namely from 1997:6 to 2009:2. We then take the reduced-form residuals from our baseline VAR ($u_t^X$ and $u_t^r$), as specified in Section 3, over the same sample. Finally, we use these series to test that condition (11) holds. For each element of $u_t^X$ — where $u_t^X$ includes the reduced form residuals of all variables apart from the policy rate — we cannot reject the null hypothesis that $E\left(V_t Z_t'\right) = 0$. The lowest p-value of the J statistic varies from 0.26 (in the case of the US corporate bond spread residuals) to 0.97 (in the case of the exchange rate residuals), and averages 0.68 across the six set of residuals in our baseline specification. As discussed above, the narrative instrument captures the changes in the policy rate that are not justified by changes in the Central Bank's assessment of current and expected macroeconomic conditions. As such, by construction it explicitly controls for the release of information from monetary policy decisions. The test therefore suggests that our IRFs are not subject to a statistically significant bias due to the presence of information effects.[29]

How good is our procedure at detecting the presence of information effects? To answer this question we exploit the case of the US, for which it is well-known that high-frequency monetary policy surprises include an important signalling component – see for example

[29] Note here that we also tested that condition (11) holds using real GDP as a measure of economic activity. As we do for other variables, we pass the test with a p-value of 0.18, which is similar than for our baseline specification using unemployment.

Nakamura and Steinsson (2018). In this case, therefore, the overidentification test we developed should be able to reject the null that $E\left(V_t Z_t'\right) = 0$, otherwise casting doubts on its power. We proceed as follows. First we estimate a 6-variable VAR as in Gertler and Karadi (2015) (see their Figure 2), including industrial production, CPI, the excess bond premium, mortgage spreads, and corporate spreads for the period 1979:7-2012:6. We then use the narrative series of monetary policy shocks computed by Romer and Romer (2004) and the series of monetary policy surprises of Gertler and Karadi (2015) to test whether condition (11) holds in the case if the US. The lowest p-value of the J statistic is 0.03 (compared to 0.18 in the case of the UK). The test therefore strongly rejects the null hypothesis that $\mathbb{E}\left(V_t Z_t'\right) = 0$, and suggests that US impulse responses based on the high-frequency surprises are significantly biased by the presence of information effects. These results not only are in line with findings in the existing literature, which shows that the non-monetary news can explain a non trivial variance of market based monetary surprises; but can also be interpreted as suggesting that the rejection of the test using UK data is not driven by a problem of low power of the test.[30]

Figure 5 displays the IRFs to a 25 basis points increase in the 1-year gilt yield using the estimated coefficients of the $B$ matrix from equation (10). The responses of the variables are comparable to those of the baseline specification, with the exception of the US corporate spreads that are a bit more delayed than in the baseline specification. Taken together, this suggests that our instrument is not significantly contaminated by informational content and that both series contain complementary information about monetary policy.

Note that the F-statistic of the first-stage regression in the overidentified system is 19.1 and the $R^2$ is 0.15. Although this statistic is higher than 10, it is lower than the one we obtained using our series of high-frequency monetary policy surprises as the only

[30]It is important to note here that the rejection of the null hypothesis $\mathbb{E}\left(V_t Z_t'\right) = 0$ could be driven by an invalidity of any other implicit assumptions imposed on the VAR (e.g., linearity or time constancy of parameters).

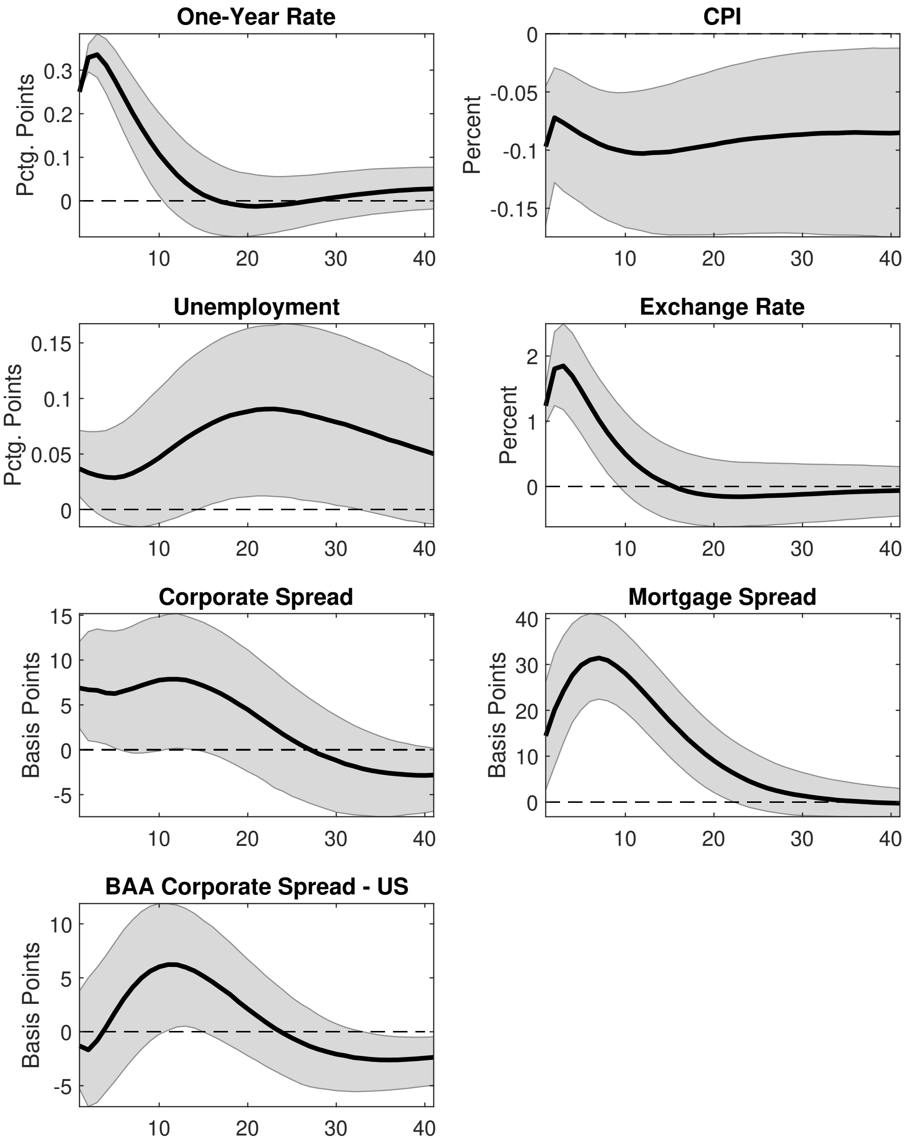


**Figure 5** IRFs To A Monetary Policy Shock - 2 Instruments. *Note*. VAR estimated in log levels, with 2 lags, and a constant over the period 1992:1-2015:1. The IRFs are computed using the first column of B matrix from equation (10). First stage results: F-Statistic: 19.1 and $R^2 = 0.15$. The solid lines and shaded areas report the mean and the 68% confidence intervals computed using moving block bootstrap with 5,000 replications.

instrument (at 40.3).[31] While the exercise reported in this section is informative about the presence of an information bias due to the signalling channel of monetary policy, we think

[31] Skeels and Windmeijer (2018) show that the 5% critical values for the F-stat with 2 instruments lies between 5.83 and 11.57. Thus, in both cases, the F-stat is higher than the critical values.

of our baseline in Section 4.1 as more precisely estimating the impact of monetary policy on macroeconomic and financial variables.

# 6 Conclusions

How does monetary policy transmit to the financial sector and the real economy? Is the credit channel playing a significant role in the monetary transmission? Does the UK monetary policy affect global financial conditions? This paper revisits these crucial questions using a novel series of monetary policy surprises for the UK.

To identify an exogenous variation in monetary policy, we follow the high-frequency methods pioneered by Kuttner (2001) and Gurkaynak et al. (2005a). We then employ the resulting series of monetary policy surprises as an instrument in a structural VAR. A monetary policy tightening induces a decline in economic activity, an appreciation of the Pound, a decline in bank credit, and an increase in both mortgage and corporate bond spreads. Finally, we also find that UK monetary policy also causes an increase in US corporate bond spreads, suggesting that monetary policy conditions in a large financial center like the UK can spillover internationally.

The monetary policy surprises that we construct are designed to be orthogonal to non-monetary developments in the macroeconomy. But policy events may still contain significant information about the macroeconomic determinants of monetary policy, therefore undermining our procedure. To test this hypothesis, we propose a new test of overidentifying restrictions that exploits the availability of the narrative series of monetary policy shocks computed by Cloyne and Hurtgen (2016). We cannot reject the hypothesis that the series of monetary policy surprises do not contain significant information about non-monetary variables. Considering the relatively low correlation between both series, this result suggests that both series contain complementary information about monetary policy.

Overall, our findings suggest that monetary policy has significant and persistent effects on economic activity and asset prices. This evidence is relevant to improve our understanding of the different transmission channels of monetary policy, which has been keenly debated. A key advantage of our new series of monetary surprises is that it includes market reaction to both current unexpected changes in policies and 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 with the 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

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.[32] 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. During our sample period, the MPC's view of the economic and financial outlook was also communicated in quarterly *Inflation Reports*, released between the policy decision and the minutes relating to the February, May, August and November MPC meetings.[33] 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.

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 usually coincided with the release of important macroeconomic data (specifically, with Labour Market Statistics). 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. After removing the minutes we are left with 291 monetary policy events: 218 MPC meetings and 73 releases of the Inflation Report.

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

[33] These arrangements were changed from August 2015, following the publication of the Warsh Review (available at http://www.bankofengland.co.uk/publications/Documents/news/2014/warsh.pdf).

## 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 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

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

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 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.
- *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.
- *Mortgage Spread*: Spliced variable mortgage (M12. Secured and unsecured personal borrowing rates, 1939-2016) rate minus the 5-year government bond yield. Source: Bank of England 'Millenium' Dataset published by the Bank of England.
- *BAA Corporate Spread - US*: the difference between Moody's BAA corporate yield and the yield on 10-Year US Treasury constant maturity. Source: Federal Reserve Economic Data (FRED), St. Louis FED (Code: BAA10Y).

- *GDP*: GDP index at monthly frequency computed by NIESR. This link https://www.niesr.ac.uk/sites/default/files/files/GDP%20Spreadsheets/dp127.PDF contains detailed information on how this series is computed.

- *Real Credit Index*: Lending by Monetary Financial Institutions (in Sterling) adjusted for the impact of securitisations, expressed in real terms using domestic CPI. The series is benchmarked to annual break-adjusted series using Chow-Lin method. Source: Bank of England "Millenium" Dataset.

- *FTSE Index*: Monthly average of FTSE All-Share index. Source: Bank of England "Millenium" Dataset.

- *Trade Balance*: Volume index. Source: Office for National Statistics, U.K. We seasonally adjust the series using X13-ARIMA-SEATS program.

# B Appendix. Specifications included in Section 4.2

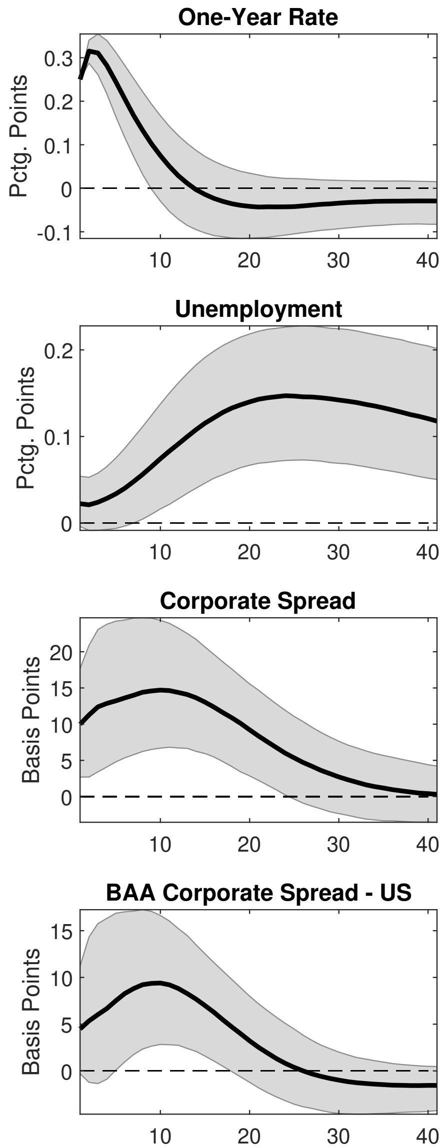


**Figure B1** IRFs To A Monetary Policy Shock - Full Specification. *Note*. VAR estimated in log levels, with 2 lags, and a constant over the 1992:1-2015:1 period. The 1-year Government Gilt Yield is instrumented using the second front contract of 3-month Sterling future. First stage results: F-Statistic: 43.5 and $R^2 = 0.12$. The solid lines and shaded areas report the mean and the 68% confidence intervals computed using moving block bootstrap with 5,000 replications.

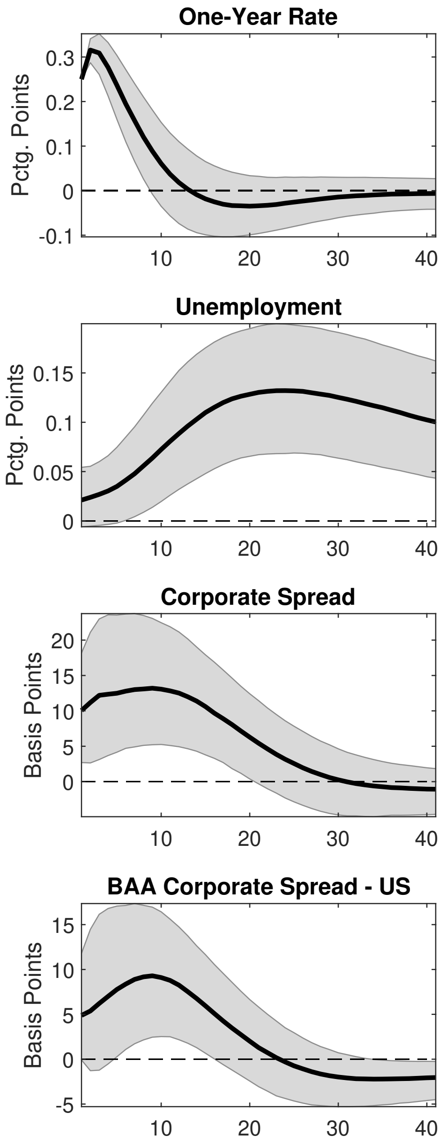


**Figure B2** IRFs To A Monetary Policy Shock - Full Specification. *Note*. VAR estimated in log levels, with 2 lags, and a constant over the 1992:1-2015:1 period. The 1-year Government Gilt Yield is instrumented using the second front contract of 3-month Sterling future. First stage results: F-Statistic: 39.8 and $R^2 = 0.12$. The solid lines and shaded areas report the mean and the 68% confidence intervals computed using moving block bootstrap with 5,000 replications.

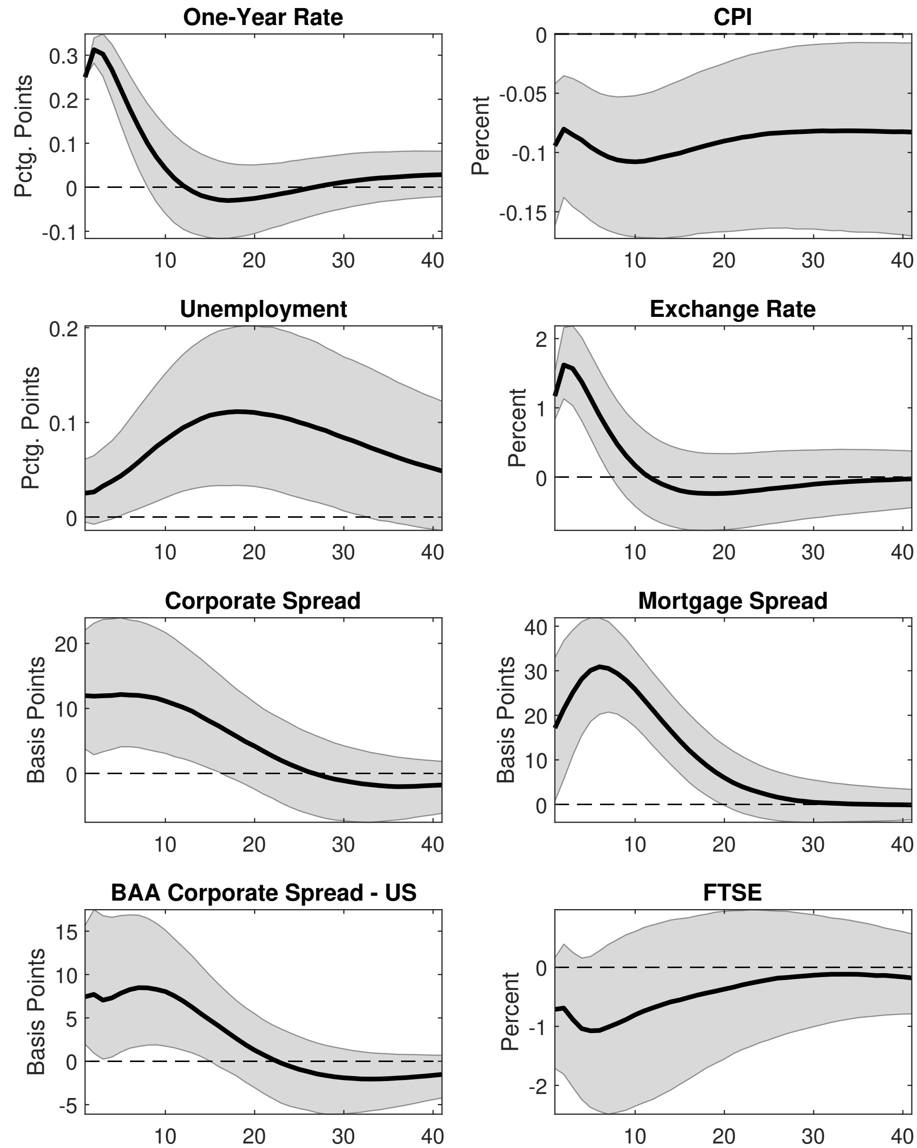


**Figure B3** IRFs To A Monetary Policy Shock - Full Specification. *Note.* VAR estimated in log levels, with 2 lags, and a constant over the 1992:1-2015:1 period. The 1-year Government Gilt Yield is instrumented using the second front contract of 3-month Sterling future. First stage results: F-Statistic: 34.5 and $R^2 = 0.11$. The solid lines and shaded areas report the mean and the 68% confidence intervals computed using moving block bootstrap with 5,000 replications.

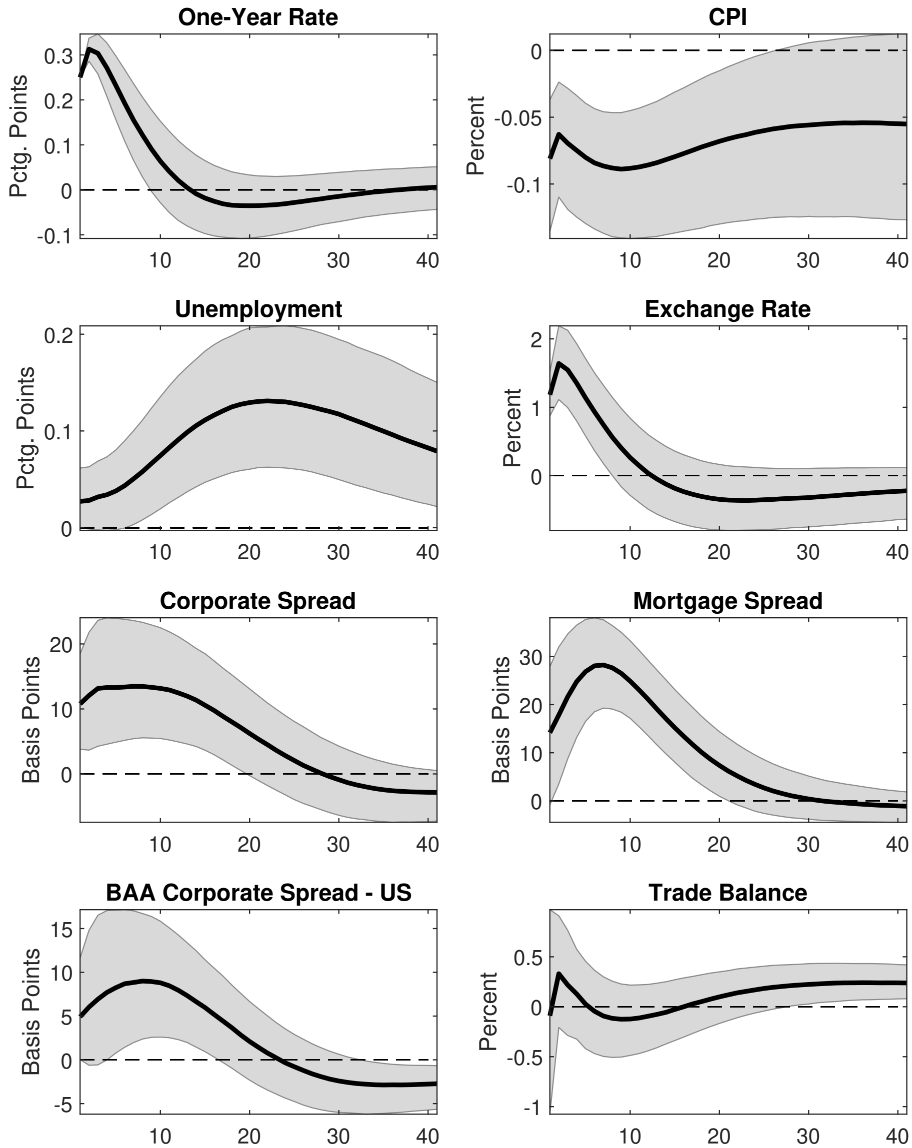


**Figure B4** IRFs To A Monetary Policy Shock - Full Specification. *Note.* VAR estimated in log levels, with 2 lags, and a constant over the 1992:1-2015:1 period. The 1-year Government Gilt Yield is instrumented using the second front contract of 3-month Sterling future. First stage results: F-Statistic: 38.4 and $R^2 = 0.12$. The solid lines and shaded areas report the mean and the 68% confidence intervals computed using moving block bootstrap with 5,000 replications.