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The Effect of Monetary Policy Shocks in the United Kingdom: an External Instruments Approach*

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Abstract

This paper uses VAR analysis to identify monetary policy shocks on U.K. data using surprise changes in the policy rate as external instruments and imposing block exogeneity restrictions on domestic variables to estimate parameters from the viewpoint of the domestic economy. The results show large and persistent effects of monetary policy shocks on the domestic economy and point to the critical role of exchange rates and term premia. The analysis resolves important empirical puzzles of traditional recursive identification methods.

JEL Classification: E44, E52, F41.

Keywords: Monetary Policy Transmission, Structural VAR, Small Open Economy, External Instruments Identification.

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1 Introduction

This paper assesses the transmission of monetary policy shocks in the United Kingdom.[1] The analysis uses a vector autoregression model (VAR) that comprises domestic and foreign variables and identifies monetary policy shocks using external instruments in the form of surprise changes in market rates that occur within a two-day window around monetary policy announcements. The estimation imposes block exogeneity restrictions to estimate parameters from the viewpoint of the small open economy. This identification strategy detects large and persistent effects of monetary policy shocks and points to a critical role of exchange rates and term premia for the transmission of monetary policy shocks. The analysis resolves empirical puzzles related to the response of inflation and the exchange rate to monetary policy shocks in recursive identification methods.

Our identification strategy is powerful in addressing two problematic assumptions of recursive identification schemes of monetary policy shocks in small open economies. First, recursive identification methods impose timing restrictions on the effect of monetary policy shocks, restricting the domestic interest rate to have a lagged effect on macroeconomic variables. This restriction is unrealistic for VAR models with financial variables and exchange rates since these variables react immediately to monetary policy shocks.[2] Our identification scheme relaxes this restriction and allows all the variables in the VAR model to respond simultaneously to exogenous changes in monetary policy. The identifying assumption is that surprise changes in market rates within a narrow time window are mainly due to monetary policy shocks and therefore are orthogonal to movements in non-monetary shocks. Second, traditional estimation of VAR models for small open economies does not assume independence of foreign variables from movements in domestic variables, which is a central

[1]The U.K. is an archetypal example of a small open economy since exports and imports equal approximately 30% of GDP, and the size of the overall economy is small compared to the world economy. We believe that our results are therefore informative on the transmission mechanism in small open economies.

[2]See Rigobon and Sack (2004), Faust et al. (2004), and references therein for a discussion of the issue.

assumption in small open economy models.[3] Our identification strategy instead imposes block exogeneity restrictions that isolate the foreign economy from movements in the domestic variables and therefore estimates the VAR model from the viewpoint of the small open economy.

Our study reveals several important results. First, monetary policy shocks have large and persistent effects on output and yields of long maturities. A monetary policy shock that raises the domestic policy rate by 25 basis points reduces output (proxied by industrial production) by approximately 0.25% within one year. We show that the same VAR model estimated with recursive identification methods leads to a smaller decrease in output of 0.17%. Our findings are similar to those in Cloyne and Hurtgen (2016), who identify the effect of a monetary policy shock for the U.K. economy using the narrative approach in Romer and Romer (2004) on a new real-time forecast data. With our identification scheme, the 10-year government bond yield increases by 32 basis points in response to the monetary policy shock whereas it falls by 11 basis points under recursive identification methods. The rise in the 10-year government bond yield in our identification scheme conforms with the theory of the term structure of interest rates that links movements in long yields to a weighted average of present and expected short yields. This finding reveals that rates with long maturities bear information about expected changes in future economic activity that entail an important propagation channel for monetary policy shocks in small open economies. Our analysis therefore provides support for the relevance of long-term yields for movements in economic activity.[4]

Second, we find that a 25 basis point increase in the domestic interest rate significantly decreases the price level by 0.12% within two years. This result sharply contrasts with those

[3]See Gali and Monacelli (2005) for a discussion on the exogeneity assumption of the foreign variables in the context of monetary policy in small open economy models.

[4]For a discussion on the topic, see Estrella and Hardouvelis (1991), Estrella (2005) and the recent survey by Wheelock and Wohar (2009).

from recursive identification methods that instead detect a rise in inflation in response to an unexpected tightening in monetary policy, an empirical phenomenon labeled "price puzzle," originally established by Sims (1992) on U.K. data.[5] We show that the counterfactual increase in inflation detected by recursive identification methods is tightly linked with the behavior of financial markets, as encapsulated by the counterfactual decrease in the 10-year government bonds. The decline in long-yield bonds is consistent with a strong economy and therefore rising inflation. Thus, our analysis shows that empirically plausible movements in the term premia are critical for a realistic response of inflation to changes in monetary policy.

Third, our identification scheme produces large and empirically plausible movements in the exchange rate. The exchange rate appreciates by 1.6% on impact in response to a 25 basis point increase in the domestic interest rate, which then quickly depreciates. The interest rate differential induced by the contractionary domestic monetary policy is offset by expected future depreciation in the domestic currency, a consistent finding with the exchange rate overshooting hypothesis described by Dornbusch (1976). We show that this finding is in sharp contrast with the results from an identical VAR model estimated with a recursive identification method, which finds that a contractionary domestic monetary policy induces an extremely weak appreciation in the domestic currency that becomes statistically insignificant from the second period after the shock.[6]

Our study is related to the realm of research that focuses on the monetary transmission mechanism in the U.K. Ellis et al. (2014) and Mountford (2005) use sign restrictions on structural VAR and FAVAR models to study the transmission mechanism of an array of macroeconomic shocks in the U.K. and find a limited role for monetary policy shocks. Cloyne

[5]See Castelnuovo and Surico (2010) for a recent discussion on the price puzzle. Our results show that identification with external instruments resolves counterfactual dynamics in small open economies.

[6]Eichenbaum and Evans (1995) first detected the counterfactual dynamics of exchange rate movements. See Engel (2014) for a review on the issues and and Benigno et al. (2011) for further discussion.

and Hurtgen (2016) use the narrative approach in Romer and Romer (2004) to identify monetary policy shocks on U.K. data and find similar results to ours. These studies focus on real activity and inflation, and instead they abstract from open economy issues and the inclusion of financial variables, which are central to our investigation. Sims (1992) is the seminal study to show anomalies related to movements in inflation and exchange rates using a recursive identification scheme estimated on U.K. data. Cushman and Zha (1997) show that the prize puzzle disappears using block exogeneity restrictions in a VAR model identified with Cholesky restriction on Canadian data. Similar to these studies, our work addresses important empirical puzzles, but we also focus on a broader set of variables and take advantage of the information in financial and exchange rate markets using a less restrictive identification scheme. Our analysis is also closely related to recent studies by Cesa-Bianchi et al. (2016) and Miranda-Agrippino and Ricco (2017a, b) that identify the effect of monetary policy surprises on the economy using high-frequency data for the U.K. We show that an analysis based on a two-day window around monetary policy announcements produces robust and consistent results compared to monetary policy instruments derived from data with intra-day window.

The analysis also relates to the studies that identify monetary policy shocks using external instruments. Our methodology is similar to Mertens and Ravn (2013, 2014), who use external instruments to investigate the dynamic effects of changes in taxes and tax multipliers. Stock and Watson (2012) use external instruments to estimate the effect of six structural shocks, including monetary policy shocks, to investigate alternative explanations of the Great Recession. Gertler and Karadi (2015) use a similar methodology to estimate the effect of monetary policy shocks and the role of forward guidance. Similar to these studies, our investigation identifies monetary policy shocks using external instruments, but our analysis focuses on the monetary transmission mechanism in a small open economy for a wider set of variables, including financial variables and exchange rates.

The remainder of the paper is organized as follows. Section 2 outlines the VAR model and the identification scheme. Section 3 presents the selection criteria for external instruments and discusses the findings. Section 4 performs robustness analyses. Section 5 concludes.

2 The identification scheme

In this section, we describe the VAR model and the identification scheme based on external instruments.

Consider the structural VAR model:

$$Ax_t = \sum_{j=1}^{p} C(j)x_{t-j} + \varepsilon_t, \tag{1}$$

where x_t is the vector of variables, A, and $C(j)$ with $j \geq 1$ are conformable coefficient matrices and ε_t is a vector of white noise structural shocks. Multiplying each side of equation (1) by A^{-1} yields the reduced-form VAR representation

$$x_t = \sum_{j=1}^{p} B(j)x_{t-j} + v_t, \tag{2}$$

where $B(j) = A^{-1}C(j)$, $v_t = S\varepsilon_t$, $S = A^{-1}$, and p is the number of lags. The variance-covariance matrix of the reduced form representation (2) is $E[v_t, v_t'] = E[SS'] = \Sigma$. Partitioning equation (2) between domestic (y_t) and foreign (y_t^*) variables yields

$$\begin{bmatrix} y_t \\ y_t^* \end{bmatrix} = \sum_{j=1}^{p} \begin{bmatrix} B_{11}(j) & B_{12}(j) \\ B_{21}(j) & B_{22}(j) \end{bmatrix} \begin{bmatrix} y_{t-j} \\ y_{t-j}^* \end{bmatrix} + \begin{bmatrix} u_t \\ u_t^* \end{bmatrix}, \tag{3}$$

where $B_{11}(j)$, $B_{12}(j)$, $B_{21}(j)$ and $B_{22}(j)$ are conformable coefficient matrices associated with domestic and foreign variables, and u_t and u_t^* are the reduced form shocks for domestic and foreign variables, respectively. The reduced form shocks are a function of the structural shocks such that

$$\begin{bmatrix} u_t \\ u_t^* \end{bmatrix} = \begin{bmatrix} S_{11} & S_{12} \\ S_{21} & S_{22} \end{bmatrix} \begin{bmatrix} \epsilon_t \\ \epsilon_t^* \end{bmatrix}, \tag{4}$$

where ϵ_t and ϵ_t^* are vectors of the structural shocks to domestic and foreign variables, respectively.

To estimate the VAR model, we impose block exogeneity restrictions that prevent changes in domestic variables to affect foreign variables while allowing foreign variables to be a source of economic fluctuations for the domestic economy. These restrictions amount to imposing $B_{21}(j) = 0$ for $j = 1, 2, ..., p$ in the reduced form VAR model in equation (3). They enable us to estimate the coefficients in the VAR model from the viewpoint of the small open economy. Our use of block exogeneity restrictions builds on work by Cushman and Zha (1997) that uses similar restrictions to identify monetary policy shocks with standard Cholesky decomposition methods, showing that they improve the identification of monetary policy shocks in open economies. We use least squares estimation of the reduced form (3) to obtain estimates of the coefficients in each matrix $B_{11}(j)$, $B_{12}(j)$, and $B_{22}(j)$ that are necessary to simulate the model and derive the reduced form residuals for the domestic economy (u_t) and the foreign economy (u_t^*).

To identify the effect of domestic monetary policy shocks, $\epsilon_{r,t}$, we partially identify the matrix S. Let s denote the column of S associated with the impact of the structural domestic policy shock $\epsilon_{r,t}$. Partition s as

$$s = \begin{bmatrix} s_1 \\ s_2 \end{bmatrix}, \tag{5}$$

where s_1 is the on-impact response of domestic variables to one unit of domestic monetary shock, $\epsilon_{r,t}$ and s_2 is the on-impact response of foreign variables to $\epsilon_{r,t}$. Note that the block exogeneity restrictions that prevent changes in domestic variables to affect foreign variables implies that $s_2 = 0$.

We estimate s_1 using an instrumental variables approach that follows the methodology developed by Stock and Watson (2012) and Mertens and Ravn (2013). A valid instrumental variable m_t is correlated with the primitive monetary shock $\epsilon_{r,t}$ (i.e. $\mathbb{E}(m_t \epsilon_{r,t}) \neq 0$) and is orthogonal to any other structural shock in the VAR model (i.e. $\mathbb{E}(m_t \epsilon_{k,t}) = 0$, for $k \neq r$). The contemporary response s_1 is estimated by regressing all reduced form non-monetary policy shocks in the domestic block on the reduced form monetary policy shock $u_{r,t}$ using m_t as an instrument. Since m_t is orthogonal to non-monetary shocks, the instrumental estimation yields a consistent estimate for s_1.[7] We normalize s_1 to represent a monetary shock of 25 basis points.

3 Estimation

This section describes the data and discusses the selection of external instruments. It presents and compares the results from our identification method with those from a standard recursive identification scheme.

3.1 Data

We use monthly frequency data for the period January 1994−December 2007 on a variety of economic and financial variables for the U.K. (domestic economy) and the U.S. (foreign economy).[8] To focus on a stable time period, we chose the sample to coincide with the

[7] The instrumental regressions estimate contemporary responses of U.K. variables to a unit of monetary shock, yielding s_1/σ_r, where σ_r is the standard deviation of $\epsilon_{r,t}$.

[8] Our identification method requires changes in realized interest rates to achieve the identification of monetary policy shocks. Since the interest rate remained unchanged in the U.K. in post-2007 period and monetary policy was conducted using alternative instruments, as outlined in Kapetanios et al. (2012), our method fails to identify monetary policy shocks in periods of unconventional monetary policy. Miranda-Agrippino and Ricco (2017a, b) overcome the issue by assuming that the transmission mechanism of monetary policy is invariant across times of normal and unconventional monetary policy. Studies by Rogers et al. (2016), Liu et al. (2017) and Swanson (2017) also consider alternative identification strategies of monetary policy

introduction of inflation targeting and a permanent flexible exchange rate in the U.K., which excludes the period of the financial crisis.[9]

The vector of domestic variables is defined as: $y_t = (g_t, p_t, \pi_t, r_t, \pounds_t, f_t, i_t)'$. It comprises readings for real activity as measured by industrial production (g_t) and the manufacturing Purchasing Managers Index (p_t), inflation (π_t) as measured by changes in the Consumer Price Index index, the domestic policy rate (r_t), and three financial variables that include the exchange rate index for the pound sterling ($\pounds_t$), returns to the FTSE (f_t) and nominal yields to 10-year government bonds (i_t).[10] The vector of foreign variables (y_t^*) comprises the U.S. policy rate (r_t^*).[11] Appendix A provides description and sources of the data.[12]

3.2 Selection of external instruments

We construct external instruments for monetary policy shocks by measuring surprise changes in market interest rates in a short time frame around the Monetary Policy Committee (MPC) policy announcements. The interest rates we consider are: the one-month libor rate (libor-1m), the three-month libor rate (libor-3m), the six-month libor rate (libor-6m) and the

shocks during the post-2007 period of unconventional monetary policy. This topic remains an interesting subject for future research.

[9]Zanetti (2014) and references therein provide a discussion on the key policy changes and time series properties in the U.K. economy. Ikeda et al. (2020) study the change in the effectiveness monetary policy during the financial crisis and the effective lower bound on the short term interest rate.

[10]Industrial production, PMI and the nominal exchange rate index enter are in logs. The other variables are levels of CPI inflation, the policy rate, monthly returns to the FTSE, yields to 10-year government bonds and the Fed funds target rate. A time trend is included in the model. Results remains similar if we use the real exchange rate.

[11]Since the Euro Area accounts for a large share of U.K. trade, we have estimated the model including one-month Euribor to the foreign block and results remain substantially unchanged. The benchmark results are generally robust to the inclusion of additional measures of real activity and financial variables in the foreign bloc. However, the response of inflation becomes insignificant when we include US and Euro Area industrial production or inflation. An appendix that details the findings is available on request to the authors.

[12]We include the the manufacturing Purchasing Managers Index (PMI) as a measure of real activity since it is more timely avaiable to financial markets (it is usually released at the end of the month or in early days of the next month) than industrial production and is closely monitored by financial markets. Koenig (2002) shows that the PMI index is closely linked to the direction of monetary policy changes and therefore it is a powerful measure for identification of monetary policy shocks. Similarly, D'Agostino and Schnatz (2012) document that the PMI index retains strong predictive and forecasting power of economic activity.

twelve-month libor rate (libor-12m). These particular interest rates are chosen since they broadly capture the market's expectations on the policy rate. Important for our analysis, we assume surprise changes in the rates in a short time frame around MPC announcements are orthogonal to non-monetary shocks included in the VAR model. As discussed in Kuttner (2001), the inflow of news in a short time frame that contains monetary policy events is dominated by monetary policy announcements, and policy announcements do not disclose private information of the central bank and do not affect risk premia of the short-term interest rates we consider. Faust et al. (2004) document that surprise changes in the federal funds futures rate due to policy announcement do not contain the Fed's private information on macroeconomic indicators. Similarly, Piazzesi and Swanson (2008) establish that surprise changes in the federal funds futures rate in a short time frame around Fed announcements do not encompass risk premia. Thus, changes in the short-term interest rate before and after monetary policy announcements can be treated as changes in the level of monetary policy.[13] Piazzesi and Cochrane (2002) measure monetary shocks as changes in the one-month Eurodollar rate a day before and a day after Fed policy announcements. They show that monetary policy surprises constructed in this way are powerful estimates of monetary shocks.[14] A number of studies embed the assumption that surprise changes in market interest rates within a short time frame are good measures for monetary shocks.[15]

We measure surprise changes in the interest rates in a two-day window within monetary

[13] In our analysis we make the same assumption. However, further research is needed to establish whether the same result holds in U.K. data.

[14] Piazzesi and Cochrane (2002) find that monetary shocks constructed as surprise changes in market interest rates are robust to the omitted-variable problem, the time-varying parameter problem and the orthogonalization problem that usually affect policy-rule-based or VAR-based monetary shocks.

[15] Other studies follow this approach. Bernanke and Kuttner (2005) study the impact of monetary policy on equity returns using futures rate surprises to measure monetary shocks. Faust et al. (2003) use futures rate surprises to identify the effects of monetary shocks on exchange rates. Gurkaynak et al. (2005b) study responses of long-term interest rates to monetary shocks. Barakchian and Crowe (2013) is the first study that estimates the impact of monetary shocks on output using a measure of the Fed's private information to identify clean monetary shocks measured from surprises changes in market rates. Gorodnichenko and Weber (2016) study responses of return volatilities to monetary shocks of firms with different levels of price stickiness.

policy MPC policy rate announcements. Since libor rates are set and reported at 11:00 a.m., one hour earlier than MPC announcements at 12:00 p.m., we do not expect libor rates to respond to monetary policy on the same day. However, a careful inspection of the data shows that libor rates occasionally change on the day of policy announcements. For example, the six-month libor rate jumped to 5.44% from 5.38% on 12 January 1994, when the Bank of England decided to keep Bank Rate unchanged, and they reverted to 5.38% one day later. In another instance, on 8 February 1994, the Bank of England cut its Bank Rate by 25 basis points. The rate cut was recorded in the six-month libor rate on the day of announcement but not one day later. Dating errors lead to abnormal changes on announcement dates. To deal with the issue of potential dating errors, we follow the approach in Cochrane and Piazzesi (2002) and define unexpected changes in libor rates due to policy announcements as the move from one day before to one day after policy announcements.[16] As a robustness check, we discuss results from instruments measured in a one-day window in section 4.

To address the problem of weak instruments, we statistically test the degree of correlation of the instruments with reduced-form residuals from the monetary policy equation in the VAR model ($u_{r,t}$) using F-statistics from the first-stage regression. Stock et al. (2002) suggest that a value of the F-statistics lower than 10 is evidence of weak instruments. Table 1 shows values of F-statistics from regressing VAR residuals of the monetary policy equation ($u_{r,t}$) on external instruments (m_t). The F-statistics show that all of the four instruments have values higher than 10 and therefore avoid the weak instrument problem. While all four instruments are statistically valid, we need to ensure that rate surprises do not result from *timing* shocks, which reflect a surprise in the *timing* of rate changes rather than a surprise in the expected *level* of future interest rates. As discussed in Bernanke and Kuttner (2005), this distinction is critical for estimating the impact of monetary policy shocks since shocks

[16] In principle, it would be interesting to use intra-day changes in market interest rates around the hour of policy announcements. However, intra-day data are unavailable for the time period we consider.

that alter the level of interest rates in the near future have stronger impact on the economy than shocks that reflect unanticipated timing in the monetary policy shock.

To investigate whether timing shocks substantially affect libor rate surprises, Figure 1 plots surprise changes in libor rates of different maturities. The left panel plots surprise changes in the six-month libor rate against surprise changes in the one-month rate. The right panel plots surprise changes in the three-month libor rate against surprise changes in the one-month rate. The 45 degree line corresponds to a one-for-one response of the six-month libor rate to the one-month libor rate. Observations lying below the 45 degree line in the northeast quadrant and above the southwest quadrant are those associated with a less than one-for-one effect on the surprise changes in the six-month libor rate; those lying along a steeper line had a greater than one-for-one effect. Regressing surprise changes in the six-month libor rate on surprise changes in the one-month libor rate yields an estimated slope of 0.63, which is significantly different from 1 at the 1% level. Similarly, a regression of surprise changes in the three-month libor rate on surprise changes in the one-month libor rate yields an estimated slope of 0.69, which is also significantly different from 1 at the 1% level. This analysis suggests that the generally stronger responses of the one-month libor rate to MPC decisions may capture timing shocks, which have smaller impacts on three- and six-month libor rates. Thus, we choose the six-month rate as our preferred instrument since it avoids the weak instrument problem and is less affected by timing shocks than surprises in the one- and three-month libor rate. Surprise changes in the twelve-month rate are less affected by timing shocks than surprise changes in the six-month libor rate. But the instrument is marginally weak for our estimation as shown in Table 1. In other words, we have selected our preferred instrument, six-month libor rate surprises, to minimize the impact of timing monetary shocks as much as possible. Nevertheless results using the twelve-month libor rates are similar to the baseline results. In section 4, we perform robustness analysis to further investigate the selection of external instruments using principal component analysis.

3.3 Findings

Figure 2 shows impulse responses to a 25-basis-point surprise in domestic monetary policy (thick line) with 68% confidence intervals (thin lines).[17] To appraise the differences between our identification method and standard recursive methods, Figure 3 shows impulse responses to a 25-basis-point surprise in monetary policy using a standard Cholesky identification (thick-dashed line) with 68% confidence bands (thin-dashed lines) alongside our identification (solid line).[18]

The figure shows that a 25-basis-point increase in the domestic interest rate generates a significant decline in industrial production of approximately 0.25% within one year after the shock. Figure 3 shows that the same VAR model estimated with the recursive identification method generates a decline in output of approximately 0.17% within one year after the shock.[19] Thus, the response of output is almost twice as strong in our identification method. Our finding is consistent with the results in Cloyne and Hurtgen (2016), who estimate the effect of monetary policy shocks using the narrative approach in Romer and Romer (2004) and find that industrial production declines approximately 0.5% (i.e. 2.3% in response to a 1% monetary policy shock).

Figure 2 shows that a 25-basis-point increase in the domestic interest rate reduces the

[17] We estimate the model in equation (3) using the ordinary least squares and imposing the block exogeneity restriction, $B_{21}(j) = 0$. We include two lags in the estimation (i.e. $p = 2$). The inclusion of additional lags requires the estimation of a large number of parameters that that proves difficult with the relatively short number of observations. We compute confidence intervals using a wild bootstrap procedure with 1000 replications. For the wild bootstrapping, we randomly draw a vector ι_t of the length of the sample from $\{-1, 1\}$, then multiply $\hat{u}_t$ and m_t element-by-element by ι_t to generate bootstrapped residuals $\hat{u}_t^{bs}$ and the corresponding external shocks m_t^{bs}. Bootstrapped samples are generated using the estimated coefficient matrices $\hat{A}$ and $\hat{u}_t^{bs}$. Then we estimate a VAR model for the bootstrapped sample, and finally implement the identification using bootstrapped instruments m_t^{bs}.

[18] The ordering of the variables in the VAR model is $(g_t, \pi_t, r_t, £_t, f_t, i_t, p_t, r_t^*)$. The results are robust to different orderings of the variables. A companion appendix that details the robustness of the findings is available on request to the authors. See Carlstrom et al. (2009) and Castelnuovo (2012) for a discussion on the issues related to the Cholesky decomposition in closed economies.

[19] A weak response of output to the monetary policy shock is also detected by Mountford (2005), using sign restrictions on a structural VAR model, and Ellis et al. (2014), using a factor-augmented VAR model on U.K. data. They find that GDP falls by approximately 0.2% and 0.6%, respectively, in response to a 25-basis-point surprise in monetary policy.

price level by 0.12% within two years after the shock. The response is significantly different from zero and it is surrounded by a large degree of uncertainty, as captured by the large confidence interval. Mountford (2005) and Ellis et al. (2014) estimate weaker negative responses of the price level of 0.04% and 0.05%, respectively. The magnitude of our results is consistent with Romer and Romer (2004), who show that a 25-basis-point, positive shock reduces CPI (PPI) by 0.9% (1.5%) in 48 months. Cloyne and Hurtgen (2016) find similar results to ours on U.K. data. Figure 3 shows the response of the price level identified with recursive identification methods. The price level increases in response to the shock. This counterfactual response is a pervasive issue in the identification of monetary policy shocks using recursive identification schemes, as originally detected by Sims (1992) and labeled "price puzzle." Our identification scheme based on external instruments resolves this issue and generates a realistic fall in inflation in response to tightening in domestic monetary policy.[20]

Figure 2 shows that an increase of 25 basis points in the domestic interest rate generates an impact appreciation of 1.6% in the exchange rate index (£ERI). The response shows that appreciation in the domestic interest rate relative to the foreign interest rate is followed by a persistent depreciation of the domestic currency. This result is consistent with the exchange rate overshooting hypothesis described by Dornbusch (1976) that predicts an appreciation in the nominal exchange rate in response to an increase in the interest rate is followed by depreciation in the nominal exchange rate.[21] The response of the exchange rate from our identification scheme is larger than estimates from alternative identification methods. For instance, Faust et al. (2003) use a VAR model identified by a response-matching approach and find that the peak appreciation of U.S. dollars against U.K. pounds is less than 1% in

[20]Numerous studies focus on possible solutions to the price puzzle. See Castelnuovo and Surico (2010), Engel (2014) and references therein for a discussion.

[21]This finding is consistent with the results in Benigno and Thoenissen (2003) from a theoretical model calibrated on U.K. data.

reaction to a 25-basis-point shock.[22] Bjørnland (2009) identifies a VAR model using a long-run neutrality restriction of monetary policy on the exchange rate and finds that nominal exchange rates appreciate on impact within the range $0.4 - 1.0\%$ in four countries: Australia, Canada, New Zealand and Sweden. Figure 3 compares the response of the exchange rate index against those estimated with recursive identification methods. The exchange rate increases by approximately 0.4% in the first period after the shock, and it then becomes unresponsive to the monetary policy shock. This is a common counterfactual finding in VAR models identified with recursive schemes. Caglayan et al. (2013) finds a similar puzzling response of exchange rate movements to the monetary policy shock in a VAR model identified with sign restrictions on U.K. data.

Figure 2 shows that an increase of 25 basis points in the domestic interest rate generates a fall of approximately 1.8% on impact in the FTSE index and the index declines by approximately 2.6% within nine months after the shock. The impact response is close to the range of values $0.6 - 1.1\%$ in Bernanke and Kuttner (2005) from an event study on the effect of an increase of 25 basis points in the nominal interest rate on U.S. equities.[23] Figure 3 shows that the response of the stock market returns estimated with recursive identification methods is similar on impact, but it becomes insensitive to the monetary policy tightening five months after the shock.

Figure 2 shows that an increase of 25 basis points in the domestic interest rate increases the yields to 10-year government bonds by 30 basis points. The impact reaction of the 10-year yields is stronger than the reaction of the short-term policy rate, showing that monetary policy has powerful and persistent effects on rates with long maturities, consistent with the findings in Hanson and Stein (2015). However, the strong increase in the 10-year yields is

[22]Note that this study uses interval identification. To make a comparison, we compare the median of their estimates to our median estimates.

[23]Bredin et al. (2007) perform an event study on U.K. data and show that a 25-basis-point monetary shock reduces the stock market returns by 0.2%, quantitatively smaller than Bernanke and Kuttner (2005).

unlinked with expectations of higher inflation (inflation falls in response to the shock), but it is generated by a rise in the term premium, which is consistent with the monetary policy channel described in models with financial frictions.[24] Figure 3 shows that the response of the yields to 10-year government bonds estimated with recursive identification methods is negative and therefore inconsistent with the theory of the term structure of the interest rate that links the increase in the short-term policy rate with a rise in rates of long-term maturities.[25]

4 Robustness analysis

To establish whether our baseline results are robust, we undertake a number of robustness checks on the choice of the external instruments and the specification of the VAR model.

We begin by investigating to what extent our instrument selection affects the results. This robustness check is particularly important since results from the F-statistics in Table 1—based on the threshold of 10 suggested by Stock et al. (2002)—show that all variables are potentially valid instruments. As shown in Figure 1, while the instruments are strongly correlated, there are substantial idiosyncratic components in the instruments as manifested by loosely scattered points along the 45-degree lines.[26] In some cases, libor rates of different maturities move against each other as a response to MPC announcements, as shown by points in the northwest and southeast quadrants. In addition, as discussed in section 2, one particular example is the significantly less than one-for-one effect between one- and three/six-month libor rate surprises. To investigate the issue of idiosyncratic changes in the

[24]See the survey by Brunnermeier et al. (2012) for a recent review on the topic.

[25]Several studies performed on closed economies data are consistent with our identified response in yields with long maturities. Cook and Hahn (1989) and Kuttner (2001) show that long-term nominal yields respond to monetary shocks in the same direction. Hanson and Stein (2015) show that distant *real* forward rates respond strongly to monetary shocks. Gurkaynak et al. (2005a) show distant nominal forward rates respond oppositely to monetary shocks.

[26]Correlation between one- and six-month libor surprises is 0.74, and that between one- and three-month libor surprises is 0.84.

valid instruments, we follow the approach in Gurkaynak et al. (2005a) and use principal component analysis to extract a common "level" factor of monetary shocks derived from market rates of different maturities.[27] We then use the principal component of all the statistically valid instruments as the external instrument to identify monetary policy shocks. Gurkaynak et al. (2005a) point out that the first principal component of surprising changes in market interest rates provides information on the level shock of monetary policy, which shifts market interest rates in the same direction. Since the first principal component is a level factor that moves the one-to-twelve month libor rates in the same way, it is free of idiosyncratic changes of the libor rates, such as timing shocks that mainly affect short-term libor rates and "future policy path shocks," mainly captured by libor rates with longer maturities.

We find that the first principal component is a strong instrument for our identification, with F-statistics from the first stage fit equal to 29.6. It accounts for 86% of variation in the instruments.[28] Figure 4 shows the impulse responses to a 25-basis-point surprise in monetary policy using the first principal component to identify monetary policy shocks (thick-dashed line) with 68% confidence intervals (thin-dashed lines). Responses from our preferred instrument (six-month labor rate) are reported (black line). The figure shows that the impulse responses from the principal component analysis are remarkably close to our baseline results. We interpret this finding as evidence that our preferred instrument captures accurately most of the information on monetary policy changes conveyed by statistically powerful instruments.[29]

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[27]To describe the principal component analysis, denote the set of n valid instruments by Z (with n columns). We calculate the first principal component in three steps. First, we demean Z by subtracting its row means and denote the demeaned matrix by $\bar{Z}$. Second, we calculate the eigenvectors of the covariance matrix of $\bar{Z}$; denote the eigenvectors by V and sort them in descending order of their associated eigenvalues. Third, we derive the first principal component (F_1) by computing $F_1 = \bar{Z}V_1$, where V_1 is the first column of sorted V.

[28]The fraction of variations explained by the jth principal component is calculated as the ratio of the jth eigenvalue (ranked in diminishing order) and the sum of total eigenvalues. The second principal component explains 11% of variations in the four variables.

[29]An appendix that compares principal component analysis on the surprise changes in one- and three-

To ensure results are robust to the specification of our instruments, we use a narrower time window to estimate monetary policy surprises and consider changes in libor rates in a one-day window within MPC announcements. Since libor rates are set and reported before 11:00 a.m., one hour earlier than MPC announcements at 12:00 p.m., we measure monetary shocks as changes in libor rates in the day after announcements. Figure 5 presents the impulse responses identified with one-day monetary shocks using the six-month libor rate (dashed line) alongside the benchmark results. Results are generally similar to the baseline identification, although the responses of inflation and FTSE index are somewhat smaller. As explained in section 3.2, the different responses may be due to measurement errors on the one-day window around the policy announcements, which is reflected by the wider uncertainty around the estimated responses in the VAR model.

Finally, to ensure our identified monetary policy shocks represent accurately the exogenous component of monetary policy and are not accounting for anticipation effects, we follow the methodology in Miranda-Agrippino and Ricco (2017a, b) and regress our identified monetary policy shock on forecast variables from Cloyne and Hurtgen (2016) that capture information available prior to the MPC meetings. The identification of shocks based on the residual of this regression, which can be interpreted as a cleaned proxy for exogenous monetary policy shocks, produce similar impulse response functions to those of our benchmark specification.[30]

month libor rates is available on request to the authors. The analysis shows that the shorter is the maturity of the instrument, the weaker is the estimated impact of monetary shocks on macroeconomic and financial variables. To the limit, impulse responses from the one-month libor rate are closer to the responses from recursive Cholesky identification methods. We interpret this finding as showing that the information conveyed in the external instrument is critical for the identification of monetary policy shocks and rates with very short maturities embed a wide range of information that is not informative on the changes in the level of monetary policy.

[30] An appendix that details the findings is available on request to the authors.

5 Conclusion

This paper assessed the transmission of monetary policy shocks in the U.K. Our identification scheme uses surprise changes in the policy rate within a narrow time window as external instruments and imposes block exogeneity restrictions on domestic variables to estimate parameters from the viewpoint of the domestic economy. The estimation detects large and persistent effects of monetary policy shocks on real activity and long-term yields, pointing to a critical role of the exchange rate and term premia. The analysis addresses important empirical puzzles related to movements in inflation and exchange rates in response to exogenous changes in monetary policy detected by recursive identification schemes.

The analysis offers some interesting directions for future research. The results shows that unexpected changes in monetary policy generate sharp movements in the term premia that are consistent with the theory of the term structure of the interest rate. It would certainly be interesting to develop small open economy models with financial frictions that introduce yields of different maturities to identify important theoretical channels that link movements in term premia to fluctuations in economic activity. This framework will also shed light on what extent and through what mechanisms the inclusion of frictional financial markets enhance the response of exchange rates to monetary policy surprises. These extensions are open to future research.

A Appendix: Data sources

The table below describes the data definition and sources.

		Definition	Source
Endogenous variable			
	g_t	log of U.K. Index of Production (Manufacturing)	Office for National Statistics
		log of U.K. Manufacturing PMI	Markit/CPIS
	π_t	month-over-month CPI inflation of the U.K.	Office for National Statistics
	r_t	Bank Rate of the Bank of England	Bank of England
	$£_t$	log of Sterling effective exchange rate index	Bank of England
	f_t	monthly returns to FTSE ALL SHARE index	FTSE
	i_t	yields to 10-year U.K. government bonds	Bank of England
Exogenous variable			
	r_t^*	Fed funds target rate	Fed
Instrument			
	libor	two-day changes of GPB LIBOR (various maturities) around MPC policy rate announcements	Bank of England

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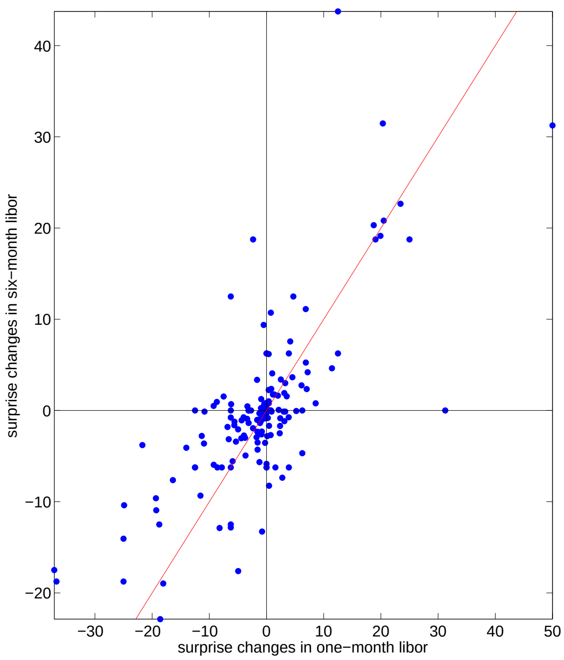
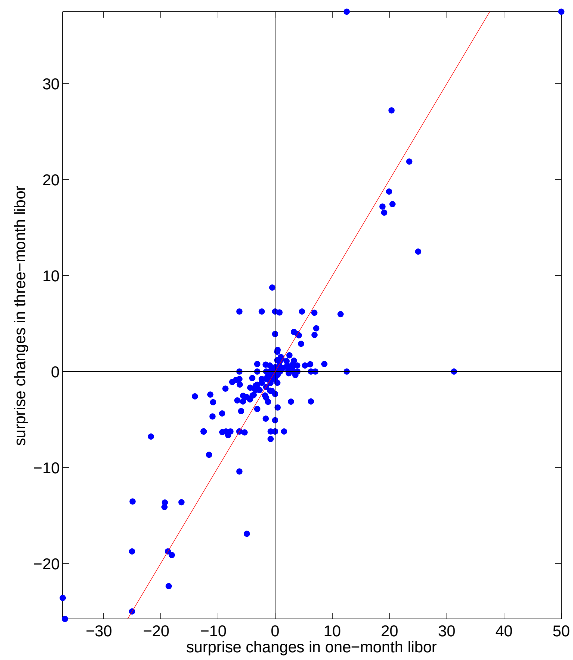
Table 1: Regressing VAR Residuals from the Monetary Equation on External Instruments for Monetary Shocks

	libor-1m	libor-3m	libor-6m	libor-12m
Coefficient	0.45	0.43	0.35	0.27
(s.e.)	(0.06)	(0.08)	(0.08)	(0.08)
Observations	165	165	165	165
R^2	0.23	0.14	0.10	0.07
F	48.7	26.4	19.0	12.3

Notes: This table reports results for regressing residuals from the monetary policy equation in (2) on various external instruments as shown in the columns. Both VAR residuals from the monetary equation and instruments are expressed in percentage points. The sample is monthly, from January 1994 to December 2007.

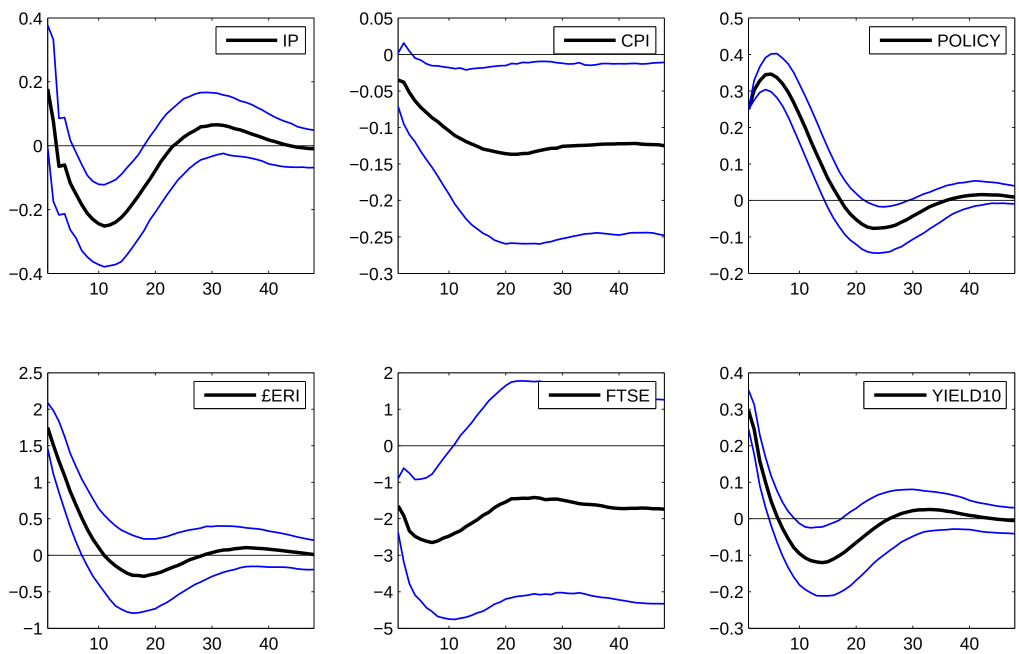
Figure 1: Surprise Changes in Libor Rates





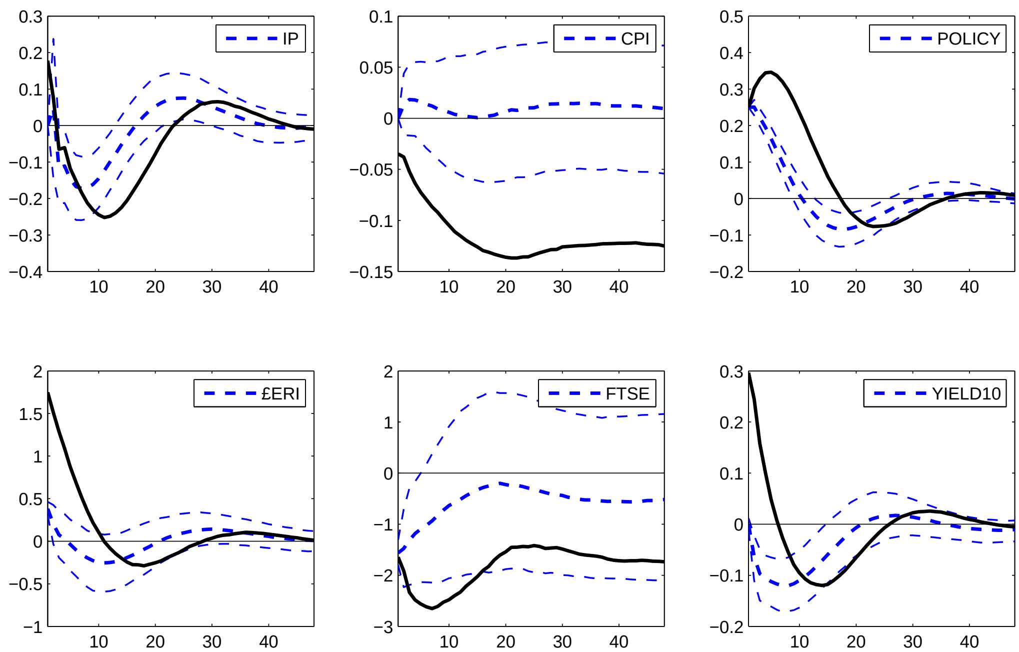
Notes: This graph shows common and idiosyncratic components of surprise changes in the one-, three- and six-month libor rates due to MPC announcement. The left panel plots surprise changes in the six-month libor rate against surprise changes in the one-month rate. The right panel plots surprise changes in the three-month libor rate against surprise changes in the one-month rate. The common component among the rates is shown by the strong correlations, while idiosyncratic components are manifested by loosely scattered points along the 45-degree line.

Figure 2: The Baseline Result



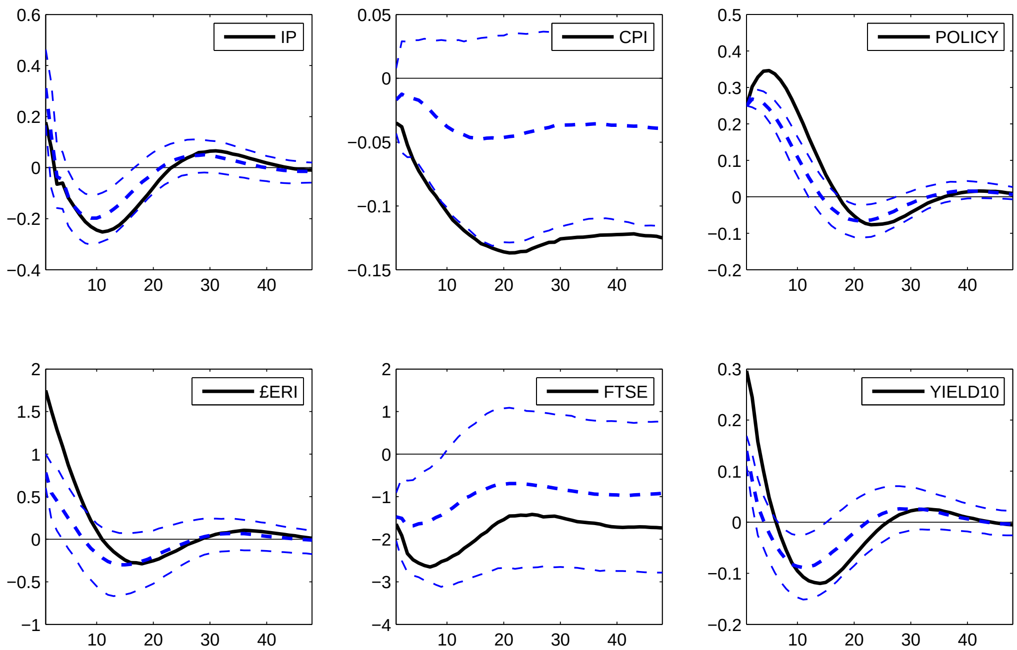
Notes: The VAR has 6 endogenous U.K. variables: log(IP), CPI inflation, the policy rate, log(£ERI), monthly returns to the FTSE and yields to 10-year government bonds; and one exogenous variable: the Fed funds target rate. Two lags of endogenous and exogenous variables are included in the VAR model. Identification is achieved with instruments for monetary policy shocks constructed as surprise changes in the six-month libor rate in a two-day window around MPC announcements. Cumulative responses from inflation and FTSE returns are reported. The 68% confidence interval is generated using a wild bootstrap procedure with 1000 replications. The sample is monthly, from January 1994 to December 2007.

Figure 3: Compare Instrumental Identification with Recursive Identification



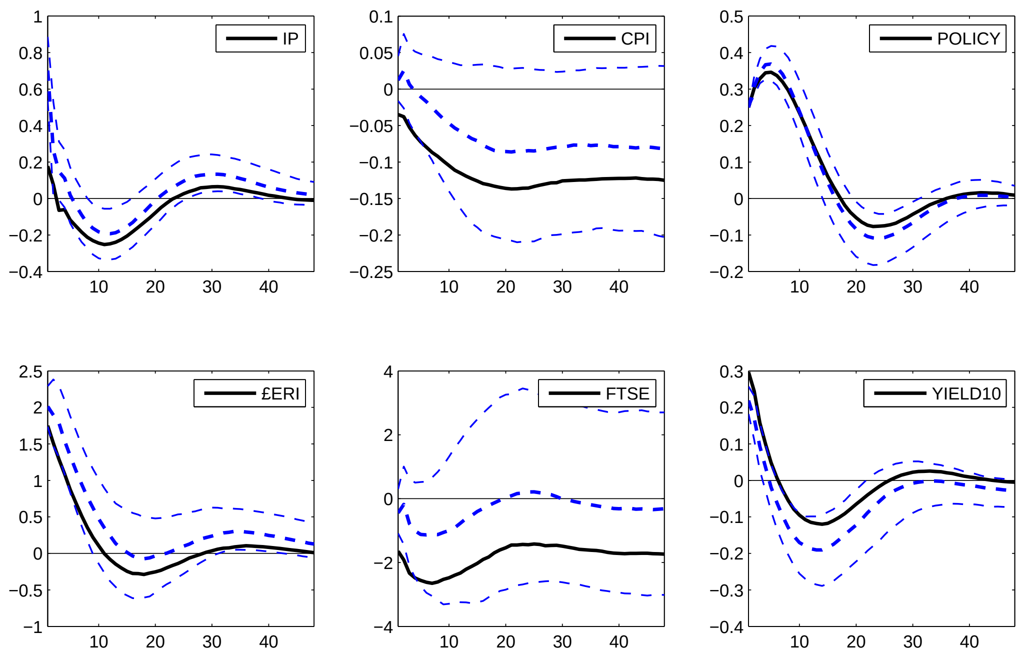
Notes: Dashed lines show median impulse responses and the 68% confidence intervals from recursive identification ordering financial variables after the monetary variable. Impulse responses from our instrumental identification are shown in solid lines.

Figure 4: Identification Using the First Principal Components of Monetary Shocks



Notes: This graph shows median impulse responses and their 68% confidence intervals (dashed lines) for a model identified with the first principal component of surprise changes in the one-, three-, six and twelve-month libor rates. Impulse responses from our baseline instrumental identification are shown in solid lines.

Figure 5: Measuring Monetary Shocks in a One-Day Window



Notes: This graph shows impulse responses identified using monetary shocks measured as changes in the six-month libor rate in a one-day window around MPC announcements (dashed lines). Impulse responses from our baseline instrumental identification are shown in solid lines.