

Monetary Policy, Credit Spreads, and Business Cycle Fluctuations

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Abstract

This paper provides new evidence on the transmission of monetary policy shocks to the real economy and to credit markets. The identification of monetary policy shock is based on a high-frequency event study analysis. Two key results emerge from the analysis for the Great Moderation period. First, monetary policy shocks were a key driver of business and credit market conditions, explaining 35% of the forecast error of industrial production and 50% of the excess bond premium. Second, monetary policy shocks explain all movement in the excess bond premium correlated with real activity leaving no role for exogenous disruptions in credit intermediation. Finally, we extend our analysis to include the Great Recession and to examine the role of unconventional monetary policy.

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1 Introduction

Understanding the transmission of monetary policy and its importance to the business cycle is central to modern economics. Starting with [Sims \(1980\)](#), a long literature has assessed the effects of monetary policy using structural Vector Autoregressions (SVARs). While many papers have found identified monetary tightenings generally reduce output,¹ the issue is far from settled, with [Uhlig \(2005\)](#) notably finding that monetary policy has no real effects. Furthermore, the consensus in the literature is that the effects of monetary policy on the real economy have become more muted over time, and in particular during the Great Moderation period.² A more recent strand of the literature has concentrated on assessing the transmission of monetary policy through financial markets, both empirically ([Gertler and Karadi, 2014](#)), and theoretically.³ This paper provides new evidence on the importance of monetary policy shocks for business cycle fluctuations, and shows how the answer to this question crucially hinges on the interaction between monetary policy and credit markets.

In our paper, we use the exogenous instrument SVAR identification to quantify the effects of monetary policy on the real economy. Overall, we find evidence that conventional monetary policy has an important effect on output: in our preferred VAR specification, it explains about 35% percent of the forecast error variance of industrial output at business cycle frequencies, a substantial fraction. Arriving at this conclusion requires explicitly acknowledging the link between measures of credit spreads, which we measure using the excess bond premium (EBP) of [Gilchrist and Zakrajsek \(2012\)](#), and monetary policy. We show first that using this measure—which contains substantial information about future economic activity—to control for credit conditions is crucial to isolating the effects of monetary policy. While this can be seen easily in a univariate framework, jointly accounting for the the relationships between output, credit conditions, and monetary policy requires a SVAR. In our SVAR, we find the surprising result that *conventional monetary policy shocks* explain the bulk of movements in credit conditions at business cycle frequencies. Once these shocks are accounted for, exogenous shocks in credit market explain very little of the portion of EBP that is related to the business cycle, at least over the pre-crisis period we consider as a baseline.

At the heart of our analysis are the shocks identified from high frequency data around FOMC statements. The changes in prices in federal funds rate futures during a narrow window around FOMC statements provides a clean measure of the unexpected component of monetary policy, which we aggregate to monthly frequency. Many studies, some discussed below, have used such a measure of monetary policy. One issue, however, is whether to use all of the events associated with FOMC statements or to use only those that correspond to eight regularly scheduled meetings of the FOMC each year, ignoring an improptu, intermeeting statements. It is common practice to use all of the available shocks, although a few studies exclude the shocks associating with intermeeting

¹See [Bernanke and Blinder \(1992\)](#), [Christiano et al. \(1996\)](#), [Leeper et al. \(1996\)](#), [Leeper and Zha \(2003\)](#), [Romer and Romer \(2004\)](#) and, more recently, [Arias et al. \(2015\)](#).

²See [Hanson \(2004\)](#), [Boivin and Giannoni \(2006\)](#), [Boivin et al. \(2010\)](#), [Castelnuovo and Surico \(2010\)](#)

³Dynamic stochastic general equilibrium models with financial frictions have been pioneered by [Bernanke et al. \(1999\)](#), and [Gertler and Karadi \(2011\)](#) provide a recent application to study the transmission of monetary policy.

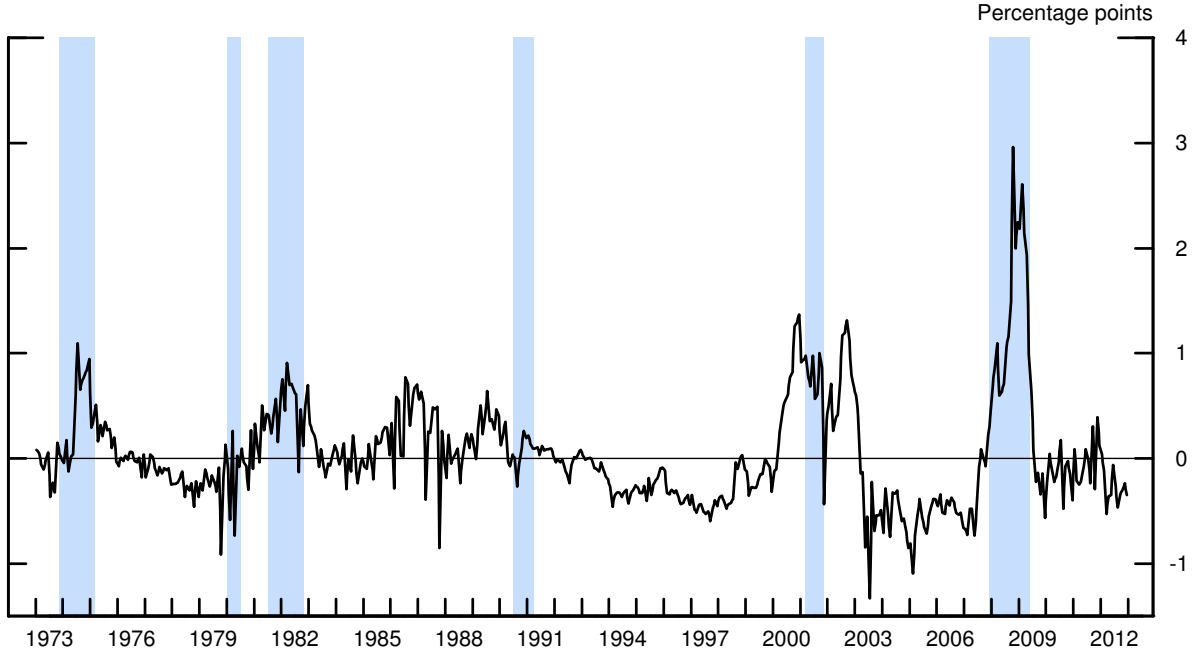
statements, on the assumption that these meetings occur directly in response to other (financial) shocks, making the measure endogeneous.⁴ We explicitly verify this, showing that including intermeeting policy moves—though there are only four in our baseline sample—induces predictability into the aggregated shock series. Including intermeeting moves biases the coefficients associated with monetary policy in least squares regressions and makes inference about the effect of monetary policy extremely difficult.

Our paper follows the literature that uses events studies to examine monetary policy shocks literature pioneered by Kuttner (2001). Other influential studies include Bernanke and Kuttner (2005), Gürkaynak et al. (2005), Campbell et al. (2012), and Gilchrist et al. (2014). The bulk of these studies consider a limited, generally univariate framework for assessing the effects on monetary policy and principally examine asset price movements. In contrast, we are more concerned with isolating with the real effects of monetary policy and quantifying the real importance of these shocks for business cycle fluctuations. Therefore, we use a VAR as our principal framework for analysis. The literature using VARs with event studies is much more sparse, an early example of which is Faust et al. (2004), who use a VAR in an impulse response matching exercise. Our paper is most closely related to Gertler and Karadi (2014) who also study the effects of monetary policy through the lens of an exogenous instrument SVAR using surprises from federal funds rate futures. Indeed, our paper echos many of the same conclusions about the role for credit conditions in transmitting monetary policy. However, there are important differences. Relative to Gertler and Karadi (2014), we show directly that the addition of the EBP to the SVAR leads to a dramatic difference in the response of *all* model variables to the monetary shock, and makes such monetary shocks important drivers of the cycle. In addition, we also document that the monetary shocks are the key driver of credit frictions in normal times, leaving no role for financial shocks to affect the real economy. Finally, in constrast to Gertler and Karadi (2014), we find no role for the monetary policy transmission through term premium effects.

The paper is structured as follows. In Section 2, we describe the features of the event studies data and the credit spread data used in subsequent analysis. In Section 3, we analyze the effects on monetary policy shocks on both industrial output and the EBP using simple univariate regressions. We also provide clear evidence of the endogeneity of shocks derived from intermeeting policy moves. Section 4 presents the exogenous instrument SVAR and the main results of the paper: at business cycle frequencies, conventional monetary policy shocks are a key driver of both the real economy and credit market conditions. In Section 5, we extend the analysis to the period of unconventional monetary policy, starting with the great recession. In Section 6, we consider some robustness exercises around the specification of the VAR and the measures of monetary shocks. Section 7 concludes.

⁴See for instance Nakamura and Steinsson (2013) and Faust et al. (2004).

Figure 1: The Excess Bond Premium



NOTE: Sample period: monthly data from 1973:M1 to 2012:M12. The solid line depicts the estimate of the excess bond premium, an indicator of the tightness of financial conditions (see [Gilchrist and Zakrajsek, 2012](#)). The shaded vertical bars denote the NBER-dated recessions.

2 Data

2.1 Measuring Financial Conditions

We rely on the information contained in corporate credit spreads to measure strains in financial markets. In particular, we use the excess bond premium, an estimate of the extra compensation demanded by bond investors for bearing exposure to U.S. nonfinancial corporate credit risk, above and beyond the compensation for expected losses. As emphasized by [Gilchrist and Zakrajsek \(2012\)](#), the U.S. corporate cash market is served by major financial institutions and fluctuations in the EBP thus capture shifts in the risk attitudes of these institutions and their willingness to bear credit risk and to intermediate credit more generally in global financial markets.⁵

Figure 1 shows this indicator of the effective “risk-bearing capacity” of the financial intermediary sector over the 1973–2012 period. Note that the EBP appears to be a particularly timely indicator of strains in the financial system. As the intensifying downturn in the U.S. subprime mortgage market in the latter half of 2006 led to the emergence of significant strains in term funding markets

⁵This interpretation is also supported by the empirical work of [Adrian et al. \(2010b\)](#) and [Adrian et al. \(2010a\)](#); [Adrian and Shin \(2010\)](#), who show that risk premiums in asset markets are very sensitive to movements in capital and balance sheet conditions of financial intermediaries. Theoretical foundations for such “intermediary” asset pricing theories are developed in the influential work of [He and Krishnamurthy \(2013\)](#) and [Brunnermeier and Sannikov \(2014\)](#).

in the United States and Europe during the summer of 2007, the EBP rose sharply, an increase concomitant with a jump in almost all measures of economic uncertainty.

2.2 Measuring Monetary Policy Shocks

We identify monetary policy shocks applying the high-frequency event study methodology developed in [Kuttner \(2001\)](#). The key aspect to this approach is to measure the unexpected change in the target federal funds rate by calculating the change in (appropriately scaled) current-month federal funds rate futures around a tight window surrounding the release of FOMC statements. [Kuttner \(2001\)](#) uses a daily window, but subsequent studies have shown that even the use of a daily window might not be enough to purge this policy measure from expected (and hence endogenous) movements. Hence, we follow [Gürkaynak et al. \(2005\)](#) and [Gilchrist et al. \(2014\)](#) and use intraday data. In particular, we use a 30-minute window (10 minutes before and 20 minutes after).⁶

Our sample begins in January 1994, the year in which the FOMC started issuing statements immediately after each meeting. We could compute unexpected changes to the target rate using federal funds rate futures from January 1990. But prior to 1994 the FOMC did not issue a statement and changes to the target rate had to be inferred by the size and type of open market operations. [Coibion and Gorodnichenko \(2012\)](#) find an increase in the ability of financial markets and professional forecasters to predict subsequent interest rate changes after 1994, suggesting that improved transparency could have altered the transmission of policy surprises.⁷ But in the Appendix we show that our qualitative results are robust to the inclusion in the sample of the early 90s.

The highest frequency for which the EBP and measures of real activity are available is monthly. Hence, before we can use the shocks series in our analysis, we have to convert it to a monthly frequency. To do this, we follow [Romer and Romer \(2004\)](#) and assign each shock to the month in which the corresponding FOMC meeting occurred. If there are two meetings in a month, we sum the shocks. If there are no meetings in a month, we record the shock as zero for that month. In Sections 3 and 4, where we characterize the effects of conventional monetary policy in normal times, we end the sample in June 2007, three months before the FOMC started to cut interest rates in response to... *the tightening of credit conditions [that] has the potential to intensify the housing correction and to restrain economic growth more generally.*

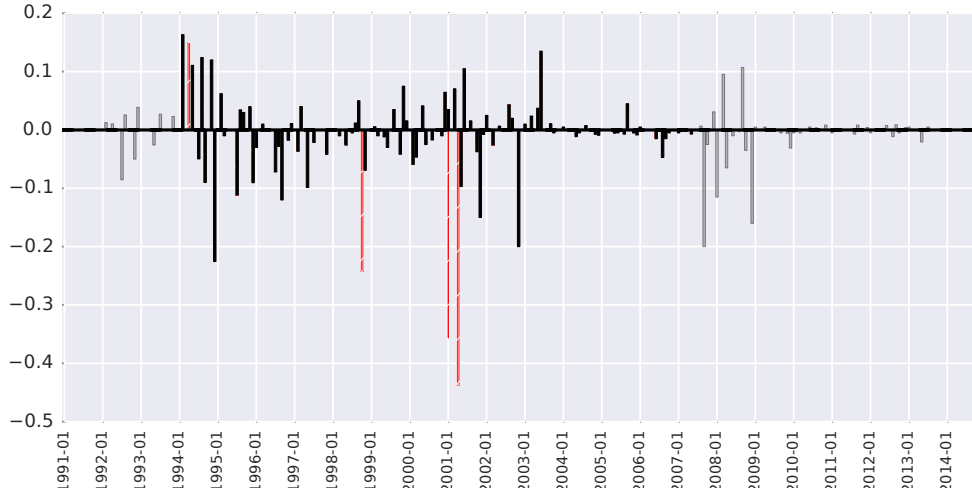
Our sample comprises of 109 announcements associated with regularly-scheduled FOMC meetings, and four announcements associated with intermeeting interest rate moves.⁸ Figure 2 plots the monthly series of unexpected policy changes. The black bars show the aggregated shock series

⁶For the identification of unexpected changes in unconventional monetary policy described below and used in Section 5, we also consider a wide 60-minute window (10 minutes before and 50 minutes after). As argued in [Gilchrist et al. \(2014\)](#), the use of a 60-minute window should allow the market a sufficient amount of time to digest the news contained in announcements associated with unconventional policy measures

⁷Prior to 1994 the FOMC often changed its target for the federal funds rate just hours after the Bureau of Labor Statistics employment report release. But the use of intraday data avoids confounding the truly unexpected change with the reaction of the fed funds rate to the employment report.

⁸As is customary in this kind of analysis, we excluded the announcement made on September 17, 2001, which was made when trading on major stock exchanges resumed after it was temporarily suspended following the 9/11 terrorist attacks.

Figure 2: Unexpected Changes to the Target Federal Funds Rate



NOTE: Sample period: monthly data from 1990:M1 to 2012:M12. The bars depict unexpected movements in the target federal funds rate. The red-dashed bars indicate intermeeting policy moves.

associated with regularly-scheduled meetings for our baseline sample 1994:1-2007:6, with the pre- and post-sample series shown in grey bars. The four intermeeting moves, highlighted in red-dashed bars in Figure 2, occurred on April 18, 1994; October 15, 1998; January 3, 2001; and April 18, 2001. The April 1994 interest rate increase is the second-largest increase in our sample, and the three other meetings represent the three largest cuts. While these four policy actions were largely unexpected, they were taken in response to economic conditions, and particular attention was paid to developments in financial markets. We report below four excerpts from the Minutes (first episode) and the Statements associated with these episodes.

- **April 18, 1994.** Policy change: 25 basis points increase. Unexpected change: 15 basis points increase. *In financial markets, sharp declines in bond and stock prices suggested that speculative excesses had been reduced, and ongoing portfolio realignments probably were shifting long-term financial assets to firmer hands.*
- **October 15, 1998.** Policy change: 25 basis points cut. Unexpected change: 23 basis points cut. *Growing caution by lenders and unsettled conditions in financial markets more generally are likely to be restraining aggregate demand in the future.*
- **January 3, 2001.** Policy change: 50 basis points cut. Unexpected change: 40 basis points cut. *These actions were taken in light of further weakening of sales and production, and in the context of lower consumer confidence, tight conditions in some segments of financial markets, and high energy prices sapping household and business purchasing power.*

- **April 18, 2001.** Policy change: 50 basis points cut. Unexpected change: 43 basis points cut. *Capital investment has continued to soften and the persistent erosion in current and expected profitability, in combination with rising uncertainty about the business outlook, seems poised to dampen capital spending going forward. This potential restraint, together with the possible effects of earlier reductions in equity wealth on consumption and the risk of slower growth abroad, threatens to keep the pace of economic activity unacceptably weak.*

In general, the literature, starting with [Kuttner \(2001\)](#) has considered both shocks associated with pre-scheduled FOMC meetings and those not.⁹ One exception to this is [Nakamura and Steinsson \(2013\)](#) who note that unscheduled meetings may occur in reaction to other shocks and thus be endogenous. We investigate this in detail and show below that the inclusion of these intermeeting moves contaminates the series, badly effecting inference.

As a matter of notation, we refer to the monthly series of aggregated shocks considering only the 109 regular meetings as m_t^{RM} and the monthly series generated by aggregating the four intermeeting moves as m_t^I , so that conventional series of shocks, which considers both types of moves, is given by $m_t^{RM} + m_t^I$.

2.3 Assessing the Monetary Policy Shocks

We investigate whether the inclusion of intermeeting policy moves in our shock series leads to endogeneity. To this end, we estimate a battery of univariate regressions:

$$\epsilon_t^{MP} = \beta_0 + \sum_{i=1}^T \beta_i^x x_{t-i} + \nu_t,$$

where x is in turn the *EBP*, the monthly growth rate of IP, the stock market return, the nominal federal funds rate, a measure of term spread, and a measure of real interest rate.¹⁰ We set ϵ_t^{MP} to either m_t^{RM} , which denotes unexpected policy changes announced at regularly-scheduled FOMC meetings, or to $m_t^{RM} + m_t^I$, where m_t^I denotes intermeeting policy moves.

We report in [Table 1](#) the sum of coefficients expressed in basis points, as well as the F-statistic of the null where each coefficient equals zero. According to the first row of [Table 1](#), m_t^{RM} cannot be predicted by any variable but the 2-year real interest rate. The addition of m_t^I makes monetary shocks predictable by all variables but the stock market returns. A high EBP predicts some of the unexpected cuts in the target rate, and a strong economy predicts some of the unexpected increases in the target rate. These effects are consistent with the narrative from the FOMC Minutes and Statements reported above. Moreover, past values of the nominal fed funds rate predict the shock

⁹See, for example, [Bernanke and Kuttner \(2005\)](#), [Gürkaynak et al. \(2005\)](#), [Campbell et al. \(2012\)](#), [Gilchrist et al. \(2014\)](#), and [Gertler and Karadi \(2014\)](#).

¹⁰The term spread is defined as the difference between the returns on a 10 year and 3 month U.S. Treasury bond. The real interest rate is defined as the 2 year Treasury yield adjusted by the previous month's twelve month inflation, given by PCE prices.

Table 1: Predictability of Monetary Policy Shocks

Predictor	EBP	ΔIP	LMRET	FFR	TS	RR
m_t^{RM}	-0.31 [0.18]	0.30 [0.30]	0.00 [0.99]	-0.21 [1.70]	0.18 [1.72]	-0.26 [3.99***]
$m_t^{RM} + m_t^I$	-2.75 [4.82***]	1.83 [5.12**]	-0.01 [2.56*]	-0.46 [5.73***]	0.87 [1.50]	-0.34 [4.62***]

NOTE: The dependent variable in each specification is ϵ_t^{MP} , a measure of monetary policy shocks. ϵ_t^{MP} equals either m_t^{RM} , which denotes unexpected policy changes announced at regularly-scheduled FOMC meetings, or $m_t^{RM} + m_t^I$, where m_t^I denotes intermeeting policy moves. LMRET = value-weighted total stock market (log) return; TS = 3m/10y term spread; RR = real 2-year Treasury yield. For each regression we report the sum of coefficients $\sum_{i=1}^T \beta_i^x$, where the lag length T is chosen using the Akaike information criterion. We report in bracket the F-statistic of the null where each coefficient β equals zero. The F-statistic is based on HAC standard errors. * $p < .10$, ** $p < .05$, and *** $p < .01$.

series because all four moves were taken to accelerate an ongoing tightening (first episode) or loosening (last three episodes) of the policy stance.

Consistent with this evidence, in the next section we show that an implication of these findings is that the use of m_t^I induces a downward bias in the estimation of the effects of monetary policy shocks on both credit spreads and real activity, while making the coefficients more statistically significant.

3 Monetary Policy, Economic Activity and Financial Conditions

In this section we examine the role of monetary policy on output using the monetary policy shocks in a univariate framework. We show that the endogeneity problem associated with the m_t^I observation contaminates estimates. Using our preferred measure, m_t^{RM} , we show that credit market conditions play a crucial role in helping to isolate the effects of monetary policy on the real economy. Next, we directly examine the relationship between the excess bond premium and monetary policy shocks. We show that there is some evidence that monetary policy plays a direct role in effecting credit market conditions.

3.1 Monetary Policy and Real Activity

As a first pass at estimating the effects of monetary policy on output, we construct simple forecasting regressions along the lines of [Gilchrist and Zakrajsek \(2012\)](#). Specifically, we assess whether the monetary shocks identified using high frequency data affect future industrial production. The regression model is given by:

$$\Delta^h Y_{t+h} = \beta_0 + \sum_{p=1}^P \beta'_{1,p} X_{t-p} + \alpha \epsilon_t^{MP} + \sum_{p=1}^3 \gamma_p EBP_{t-p} + e_{t+h} \quad (1)$$

Y_t , the log of industrial production, is converted in h step average growth rates as using $\Delta^h Y_{t+h} = (Y_{t+h} - Y_{t-1})/(h + 1)$. P lags of a vector of controls X_t is included in the equation, where X_t comprised the monthly growth rate of industrial production, the federal funds rate, and monthly inflation measured by PCE prices. The variable ϵ_t^{MP} is the monetary shock series discussed above, measured both by the full series and a series without intermeeting changes. We also include, in some versions of the regressions, 3 lags of the excess bond premium (EBP_t).¹¹ While the framework of the regression is predictive, since the shocks are (should be) by construction exogenous, estimates of α lend themselves to casual interpretation. As discussed above, we restrict ourselves to the sample from 1994:1 - 2007:6.¹²

Table 2 reports results at the three month, twelve month, and eighteen month forecast horizon. For each horizon, we consider four specifications based on two changes: 1) using the full series ($m_t^{RM} + m_t^I$) or the series purged of intermeeting changes (m_t^{RM}) for ϵ_t^{MP} and 2) including or excluding the lags of EBP_t . The first column of Table 2 reports results for the regression using the full shock series and excluding lagged EBP_t . The OLS point estimates at all three horizons are *positive*, indicating that a monetary tightening is associated with an increase in industrial production—though, at longer horizons this measured effect is not statistically significantly different from zero according to conventional heuristics. The reason for the counterfactual sign of $\hat{\alpha}$ is that the large negative surprises associated with intermeeting moves occurred *in response* to poor credit conditions, associated with lower future output. The next column of the Table reports regression results using the same monetary shock but including 3 lags of EBP as an additional control. Once past credit conditions are included in the regression, the coefficient associated with ϵ_t^{MP} is more-or-less indistinguishable from zero.

The third column shows the results of the regression using only shocks with regularly-scheduled meetings, but not including an measure of past credit conditions. Now, the point estimates of the coefficient α are negative, consistent with the previous evidence discussed. Once we include past EBP , as shown in the regression results in the fourth column, the estimates associated with the monetary policy shock become larger and more precise (though still not “significant” at conventional levels). The inclusion of past credit conditions helps to purge industrial production a large predictable component which can be seen by the adjusted R^2 jumping from 0.11 to 0.35 in the case of the twelve month prediction. Removing this predictable component helps isolate the effects of monetary policy.

These regressions highlight a number of issues. First, using the “entire” shock series (that includes intermeeting policy moves) is problematic in monthly regressions because, as we have seen in Table 1, these intermeeting moves are predictable. This endogeneity biases the coefficients. Controlling for past credit conditions is not sufficient, in this framework, to correct this. Second, using only data from regularly-scheduled meetings indicates that there is some evidence that a monetary tightening leads to a fall in industrial production. Uncovering this result requires purging

¹¹The results are not sensitive to this lag selection.

¹²It should be noted that this set up is quite similar to the local linear projections framework of Jordà (2005).

industrial production of substantial predictability from credit conditions. These results indicate a relationship between output, credit conditions, and monetary policy that is difficult to disentangle. In the next section we look directly at the relationship between EBP and the monetary policy shocks.

3.2 Monetary Policy and Credit Risk

Having seen that controlling for credit market conditions helps isolate the effects of monetary policy, we turn now to the direct relationship between the monetary shock and the excess bond premium. To do this, we estimate regressions of the form,

$$EBP_t = \beta_0 + \sum_{p=1}^P \beta'_1 X_{t-p} + \alpha \epsilon_t^{MP} + e_t \quad (2)$$

Table 3 shows the results of these regressions for both $\epsilon_t^{MP} = m_t^{RM}$ and $\epsilon_t^{MP} = m_t^{RM} + m_t^I$. In addition to considering the two versions of the monetary shock, we also consider two sets of controls: only lags of EBP itself and a broader set including the federal funds rate and industrial production growth. The first and second columns of Table 3 shows the regression results when using both regular meeting and intermeeting-based shocks for the two sets of controls. The coefficient associated with ϵ_t^{MP} in these cases is significant and positive. When we only consider monetary shocks arising after regularly schedule meetings, m_t^{RM} , the magnitude of the coefficients increases, reflecting the simulateneity of EBP and m_t^I . Once cleaned of this, the monetary shocks pass through to the excess bond premium almost one-to-one at the point estimate, although the “significance” of this is moderate. Still, it is instructive to note as we move to the external instruments SVAR.

4 Results from SVARs with Exogenous Instruments

In this section we use SVARs to further explore the interactions between monetary policy, credit market frictions, and the real economy. The methodology we use to implement our identification strategy is based on [Mertens and Ravn \(2013\)](#). We only provide an overview of the methodology, which combines VAR models with exogeneous instruments, and refer the reader to their paper for details. We then describe the results for three alternative VAR specifications which differ in how financial frictions interact with the remaining variables in the model.

4.1 Econometric Methodology

Let us consider the following SVAR:

$$\mathbf{y}'_t \mathbf{A}_0 = \sum_{\ell=1}^p \mathbf{y}'_{t-\ell} \mathbf{A}_\ell + \mathbf{c} + \mathbf{e}'_t \quad \text{for } 1 \leq t \leq T, \quad (3)$$

where $\mathbf{y}_t$ is an $n \times 1$ vector of endogenous variables, $\mathbf{e}_t$ is an $n \times 1$ vector of structural shocks,

$\mathbf{A}_\ell$ is an $n \times n$ matrix of structural parameters for $0 \leq \ell \leq p$ with $\mathbf{A}_0$ invertible, $\mathbf{c}$ is a $1 \times n$ vector of parameters, p is the lag length, and T is the sample size. The vector $\mathbf{e}_t$, conditional on past information and the initial conditions $\mathbf{y}_0, \dots, \mathbf{y}_{1-p}$, is Gaussian with mean zero and covariance matrix $\mathbf{I}_n$ (the $n \times n$ identity matrix). The model described in equation (3) can be written as

$$\mathbf{y}'_t \mathbf{A}_0 = \mathbf{x}'_t \mathbf{A}_+ + \mathbf{e}'_t \quad \text{for } 1 \leq t \leq T, \quad (4)$$

where $\mathbf{A}'_+ = \begin{bmatrix} \mathbf{A}'_1 & \dots & \mathbf{A}'_p & \mathbf{c}' \end{bmatrix}$ and $\mathbf{x}'_t = \begin{bmatrix} \mathbf{y}'_{t-1} & \dots & \mathbf{y}'_{t-p} & 1 \end{bmatrix}$ for $1 \leq t \leq T$. The dimension of $\mathbf{A}_+$ is $m \times n$, where $m = np + 1$. We call $\mathbf{A}_0$ and $\mathbf{A}_+$ the structural parameters. The reduced-form vector autoregression (VAR) model implied by equation (4) is

$$\mathbf{y}'_t = \mathbf{x}'_t \mathbf{B} + \mathbf{u}'_t \quad \text{for } 1 \leq t \leq T,$$

where $\mathbf{B} = \mathbf{A}_+ \mathbf{A}_0^{-1}$, $\mathbf{u}'_t = \mathbf{e}'_t \mathbf{A}_0^{-1}$, and

$$\mathbb{E}[\mathbf{u}_t \mathbf{u}'_t] = \mathbf{\Sigma} = (\mathbf{A}_0 \mathbf{A}_0')^{-1}. \quad (5)$$

For ease of exposition we denote by F the matrix of impact responses $\mathbf{A}_0^{-1}$. Without loss of generality, let us consider the partition $u_t = [u_{1,t}, u_{2,t}]$, where $u_{1,t}$ is a $k \times 1$ vector of policy variables and $u_{2,t}$ is a $n - k \times 1$ vector of macroeconomic variables. We apply a similar partition to $e_t = [e_{1,t}, e_{2,t}]$, where $e_{1,t}$ is the $k \times 1$ vector of policy shocks and $e_{2,t}$ is a $n - k \times 1$ vector of non-policy shocks, and to $F = [F_1, F_2]$. Using this partition, we can write the relationship between reduced-form residuals and structural shocks as follows:

$$u_{1,t} = \eta u_{2,t} + S_1 e_{1,t}, \quad (6)$$

$$u_{2,t} = \xi u_{1,t} + S_2 e_{2,t}, \quad (7)$$

where η , ξ , S_1 and S_2 contain the structural coefficients underlying F . In Section 4.2, $k = 1$ and $u_{1,t}$ includes only the federal funds rate. In Section 5, $k = 2$ and $u_{1,t}$ includes a measure of both short-term and long-term interest rates.

Mertens and Ravn (2013) and Gertler and Karadi (2014) characterize only the response of the endogenous variables to the policy shocks $e_{1,t}$ for which proxies are available:

$$F_1 = \begin{bmatrix} (I - \eta\xi)^{-1} \\ \xi(I - \xi\eta)^{-1} \end{bmatrix} S_1. \quad (8)$$

Instead, we also characterize the effects of the non-policy shocks $e_{2,t}$:

$$F_2 = \begin{bmatrix} (I - \eta\eta)^{-1} \\ \xi(I - \xi\eta)^{-1} \end{bmatrix} S_2. \quad (9)$$

The identification of $e_{2,t}$ is crucial for our analysis because $e_{2,t}$ contains the financial shock

studied in [Gilchrist and Zakrajsek \(2012\)](#). We assume that we observe a $k \times 1$ vector of proxies, denoted by m_t , for the policy shocks $e_{1,t}$. In Section 4.2, $k = 1$ and m_t is the shock series m_t^{RM} described in Section 2.2. In Section 5, $k = 2$ and m_t is augmented with a proxy for unconventional monetary policy constructed by [Gilchrist et al. \(2014\)](#).

We assume that the proxy m_t satisfies

$$\mathbb{E}[m_t e_{1,t}] = \Phi \quad (10)$$

and

$$\mathbb{E}[m_t e_{2,t}] = 0; \quad (11)$$

that is, the proxy m_t has to be correlated with the unobserved monetary policy shock, and uncorrelated with the non-policy shock. From equations (5)-(11) it follows that

$$\Phi F_1 = \mathbb{E}[u_t m_t]. \quad (12)$$

The system in (12), of dimension $n - k$, provides additional identifying restrictions but also depends on the k^2 unknown elements of Φ . Since we do not make any assumptions on Φ other than nonsingularity, equation (12) provides only $(n - k)k$ identification restrictions. These restrictions can be expressed as

$$F_{21} F_{11}^{-1} = \mathbb{E}[u_{1,t} m_t]^{-1} \mathbb{E}[u_{2,t} m_t]. \quad (13)$$

The restrictions in (13) pin down ξ in equation (7). To this this, notice that an estimate of ξ can be obtained by a 2SLS using m_t as instruments. But this is exactly what regression (12) and the rescaling in (13) do. [Mertens and Ravn \(2013\)](#) show that the restrictions in (13), together with the covariance restrictions (5) allow for the identification of η and $S_1 S_1'$, the covariance matrix of the policy shocks $S_1 \epsilon_{1,t}$. It is straightforward to show that these restrictions also identify $S_2 S_2'$, the covariance matrix of the policy shocks $S_2 \epsilon_{2,t}$.

The covariance restrictions are, however, not sufficient to obtain the structural decomposition of the covariance matrices and obtain S_1 and S_2 . We follow [Mertens and Ravn \(2013\)](#) and identify S_1 imposing a Cholesky decomposition of $S_1 S_1'$. Similarly, we identify S_2 imposing a Cholesky decomposition of $S_2 S_2'$.

[Gilchrist and Zakrajsek \(2012\)](#) identify a financial shock using a Cholesky factor with the EBP ordered last in the system. That is, the financial shock has no contemporaneous effects on any variable in the system but the EBP. Since their SVAR includes the federal funds rate, their identification of the financial shock is conditional on the identification of a monetary policy shock. Our identification differs in two assumptions. First, we also condition the identification of the financial shock to the identification of a monetary shock. But the latter is identified through a proxy variable instead of a Cholesky. Second, the identification of the financial shock imposes a Cholesky *within* $u_{2,t}$, where we order the EBP last. These two assumptions imply that the

financial shock *can* affect contemporaneously the other variables in the system. To see this, notice that moves that, in general, $\eta \neq 0$ in equation (6). Hence, a shock that moves the EBP also moves contemporaneously the policy indicators in $u_{1,t}$. In turn, $u_{1,t}$ moves $u_{2,t}$ as in general $\xi \neq 0$ in equation (7). Hence, our assumptions imply that the financial shock has a contemporaneous effect on the other variables in the system *only* through the endogenous response of monetary policy.

Contrary to all papers in the literature using SVAR with exogenous instruments, we use Bayesian techniques to inform the uncertainty of our estimates. Specifically, we use to elicit the posterior of the reduced-form coefficients of the VAR $(\mathbf{B}, \mathbf{\Sigma})$ use standard methods. Given the posterior of the reduced form coefficients, we identify the VAR using restrictions (5) and (13), together with the Cholesky factorization of $S_1 S_1'$ and $S_2 S_2'$.¹³ There exists a unique mapping between the reduced-form parameters $(\mathbf{B}, \mathbf{\Sigma})$ and the structural parameters $(\mathbf{A}_0, \mathbf{A}_+)$, which we can translate into a dogmatic prior for the structural parameters conditional on the reduced-form parameters. As described in Del Negro and Schorfheide (2011), we can draw from the posterior of the structural parameters by drawing from the posterior of the reduced-form parameters and then apply the identification mapping to get draws from the posterior of the structural objects of interest. It should be noted that this approach ignores any uncertainty in the external instruments step. A more coherent Bayesian approach would be to extend the probability model to include the proxy m_t and to conduct joint inference on all the objects simultaneously, which could more completely characterize the uncertainty surrounding the external instrument's relations to the structural shocks. Inference on this extended model is currently ongoing; we view our current approach as an important first pass. The exact algorithm is reported in detail below.

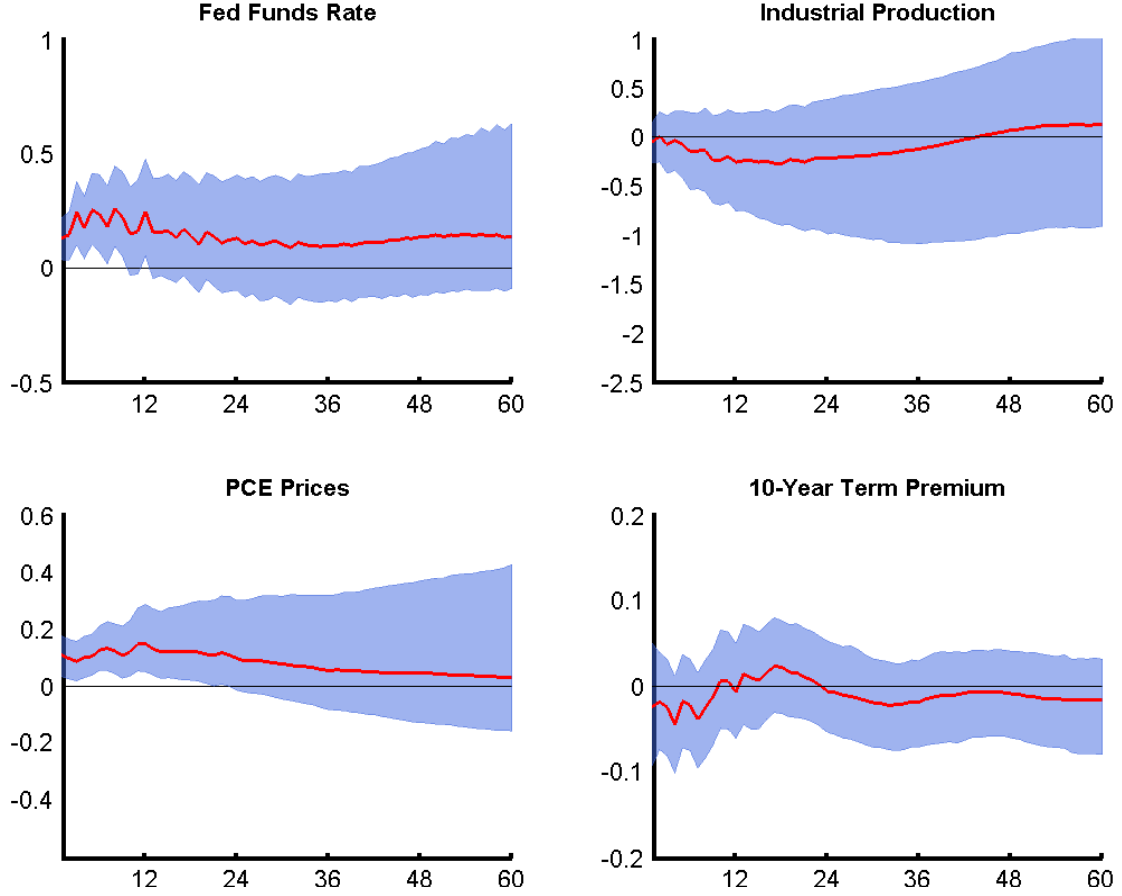
Algorithm 1

1. Draw $(\mathbf{B}, \mathbf{\Sigma})$ from the posterior distribution of the reduced-form parameters.
2. Compute the vector of reduced-form residuals $\mathbf{u}_t$
3. Regress $\mathbf{u}_t$ on m_t and calculate restrictions (13)
4. Combine restrictions (5) and (13) to compute the impact matrix F .
5. Return to Step 1 until the required number of draws from the posterior distribution has been obtained.

The benchmark monthly VAR specification consists of four endogenous variables: (1) the effective nominal federal funds rate; (2) the nominal 10-year Treasury yield; (3) the log of manufacturing industrial production; (4) the log of the PCE price deflator; (5) the EBP. We impose a Normal

¹³Rubio-Ramírez et al. (2010) derive necessary and sufficient conditions for global identification of exactly identified models, which are satisfied by the SVARs studied in this paper.

Figure 3: Macroeconomic Implications of a Monetary Policy Shock
(VAR Estimation without the EBP)



NOTE: The solid line in each panel depicts the median impulse response of the specified variable to a 1 standard deviation shock to the federal funds rate. Shaded bands denote the 90-percent pointwise credible sets.

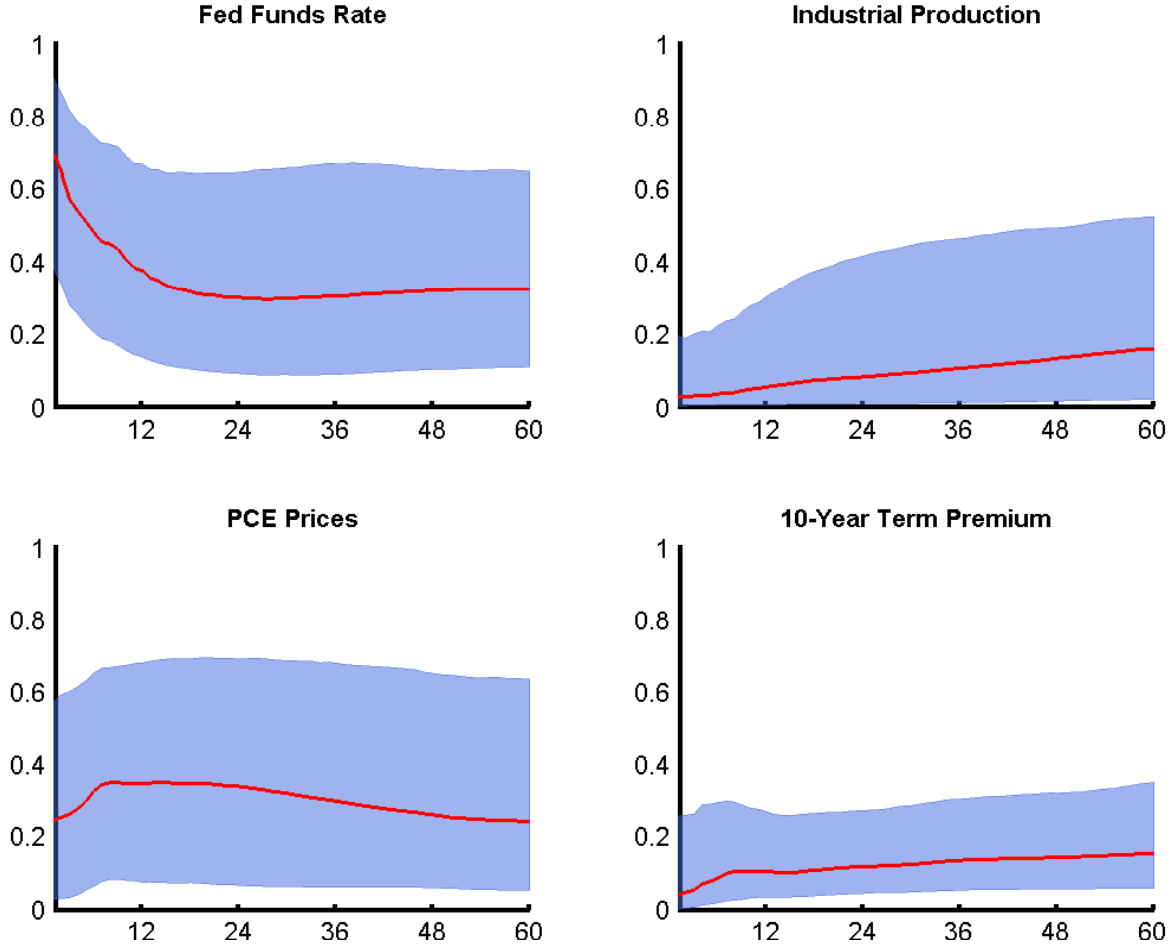
Inverse-Wishart prior on the reduced-form parameters (see Uhlig, 2005). The resulting specification, which includes a constant, is estimated over the 1993:M1–2007:M06 period using twelve lags of the endogenous variables.¹⁴

4.2 Macroeconomic Implications of Monetary and Credit Shocks

We first consider a VAR specification that does not include the EBP. The solid lines in Figure 3 show the median impulse responses of the four endogenous variables to a one standard deviation shock to the federal funds rate, while the shaded bands represent the corresponding 90-percent

¹⁴All the results reported in the paper are based on 10,000 draws from the posterior distribution of the structural parameters (increasing the number of draws had no effect on the reported results).

Figure 4: Forecast Error Variance Decomposition of a Monetary Policy Shock
(VAR Estimation without the EBP)

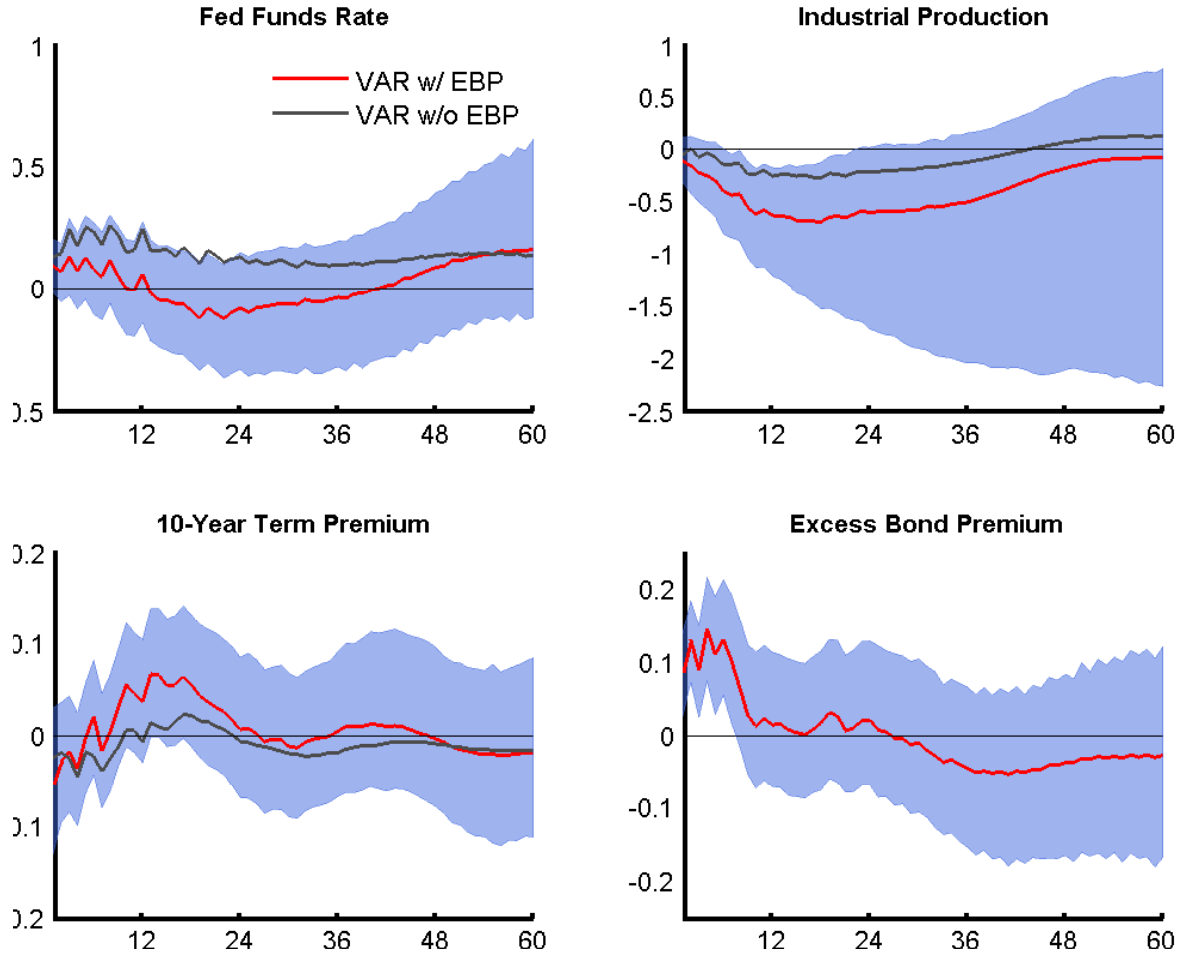


NOTE: The solid line in each panel depicts the median estimate of the portion of the forecast error variance of a specified variable attributable to a 1 standard deviation monetary policy shock. Shaded bands denote the 90-percent pointwise credible sets.

pointwise credible bands. We report the 10-year term premium instead of the response of the 10-year Treasury yield.¹⁵ Overall, the shock leads to a 25 basis point immediate tightening in the monetary policy stance. This is followed by a persistent increase in the federal funds rate which directly translate into movements into the 10-year yield, leaving the term premium unchanged. The rise in the fed funds rate has no meaningful economic or statistical effect on macroeconomic conditions. Finally, the response of prices exhibits a price puzzle, a feature of monetary shocks extensively documented in the literature.

¹⁵The term premium is computed as the difference between the response of the 10-year yield and the cumulative response of the federal funds rate.

Figure 5: Macroeconomic Implications of a Monetary Policy Shock
(VAR Estimation with the EBP)



NOTE: The solid line in each panel depicts the median impulse response of the specified variable to a 1 standard deviation shock to the federal funds rate. Shaded bands denote the 90-percent pointwise credible sets.

The solid lines in Figure 4 show the amount of variation in the same endogenous variables explained by monetary shocks. Results are consistent with the impulse response shown in Figure 3. According to this metric, monetary shocks are a minor source macroeconomic fluctuations—they are estimated to explain around 5 percent of the variation in industrial production 18 months after the policy intervention, the month when the peak response of industrial production is recorded. In addition, such shocks explain little variation in term premia.

We now consider full VAR specification that does include the EBP. The solid red lines in Figure 5 show the median impulse responses of the endogenous variables to a one standard deviation shock to the federal funds rate, while the shaded bands represent the corresponding 90-percent pointwise

credible bands.¹⁶ For comparison, we plot in gray the the median responses from the four equation model.

A one standard deviation shocks leads to an immediate increase in the federal funds rate of about 18 basis points. The shock induces a significant deterioration in macroeconomic and financial conditions. The decline in industrial production reaches its peak approximately 18 months after the shock. The EBP jumps by 12 basis points on impact, and remains above average for about one year. The response of the term premium is both economically and statistically not significant.

The response of the federal funds rate might seem puzzling at first. The initial increase is smaller than in Figure 3 and the dynamic response is not statistically significant. But this response is consistent with the endogenous response of monetary policy to the cycle. The initial increase in rates is quickly reversed due to the negative effect it induces on real activity. It might also be possible that policy also endogenously reacts to the tightening in financial conditions. Hence, a small increase in the policy rate does not necessarily indicate that the HFI proxy is a poor proxy for monetary shocks.

Figure 6 shows that the inclusion of the EBP in the VAR substantially alters the amount of variation in the endogenous variables explained by monetary shocks. Importantly, the monetary shock explains 30% of the FEV of industrial production, and 50% of the FEV in the EBP. Furthermore, the monetary shock explains up to 40% of the variation in the federal funds rate, down from 70% in the specification with the EBP. The first result suggests that monetary policy shocks could be a key driver of the business cycle, as well as the credit cycle. The second result suggests that the systematic component of monetary policy might be a far more important source of fluctuations in the federal funds rate than previously estimated. Furthermore, since this result is obtained including the EBP the federal funds rate equation, the evidence seems to point out that monetary policy reacts to systematically to developments in financial markets beyond the direct effects these have on the real economy.

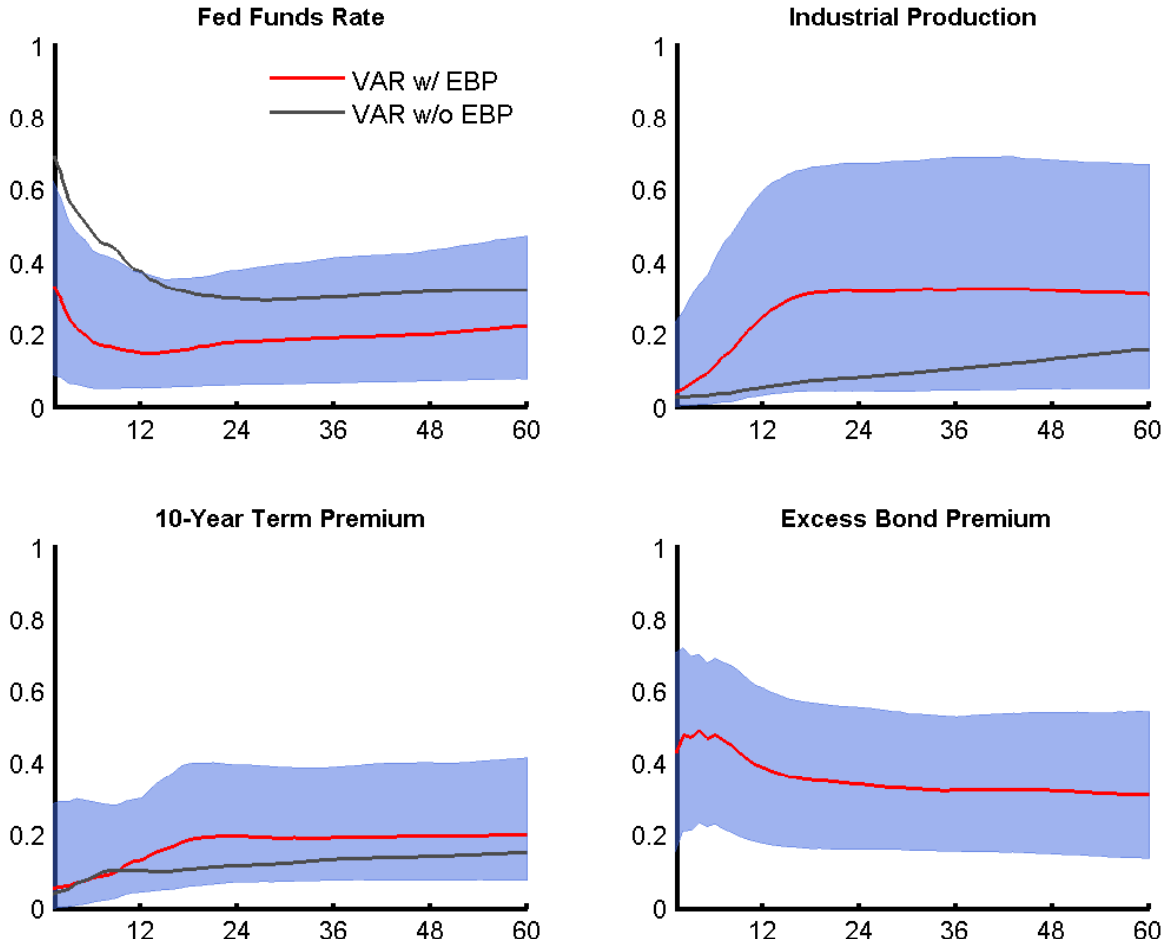
Our last set of results characterize the effects of financial shocks, identified after the monetary shock in our VAR. The solid lines in Figure 7 show the median impulse responses of four endogenous variables to a one standard deviation shock to the EBP. Overall, such an unanticipated jump in the EBP—about 10 basis points—does not induce any response in real activity. Nonetheless, the federal funds rate exhibits a small but persistent drop, adding further evidence that monetary policy might react to developments in credit markets beyond their direct effects on the real economy.

Figure 8 plots the the amount of variation in the endogenous variables explained by the financial shock. Financial innovations do explain a sizeable share of the variability in the EBP, up to 50%, but these movements are uncorrelated to developments in the real economy.

All told, three findings emerge from this section. First, in normal times monetary policy shocks are an important driver of the business and credit market cycles. Second, the federal funds rate seem to respond to movement in credit spreads beyond their direct effects on the economy. Third, monetary policy shocks accounts for movements in the EBP correlated with the business cycle.

¹⁶The response of the price level is similar across specifications, and hence we omit it to save on space.

Figure 6: Forecast Error Variance Decomposition of a Monetary Policy Shock
(VAR Estimation with the EBP)



NOTE: The solid line in each panel depicts the median estimate of the portion of the forecast error variance of a specified variable attributable to a 1 standard deviation monetary policy shock. Shaded bands denote the 90-percent pointwise credible sets.

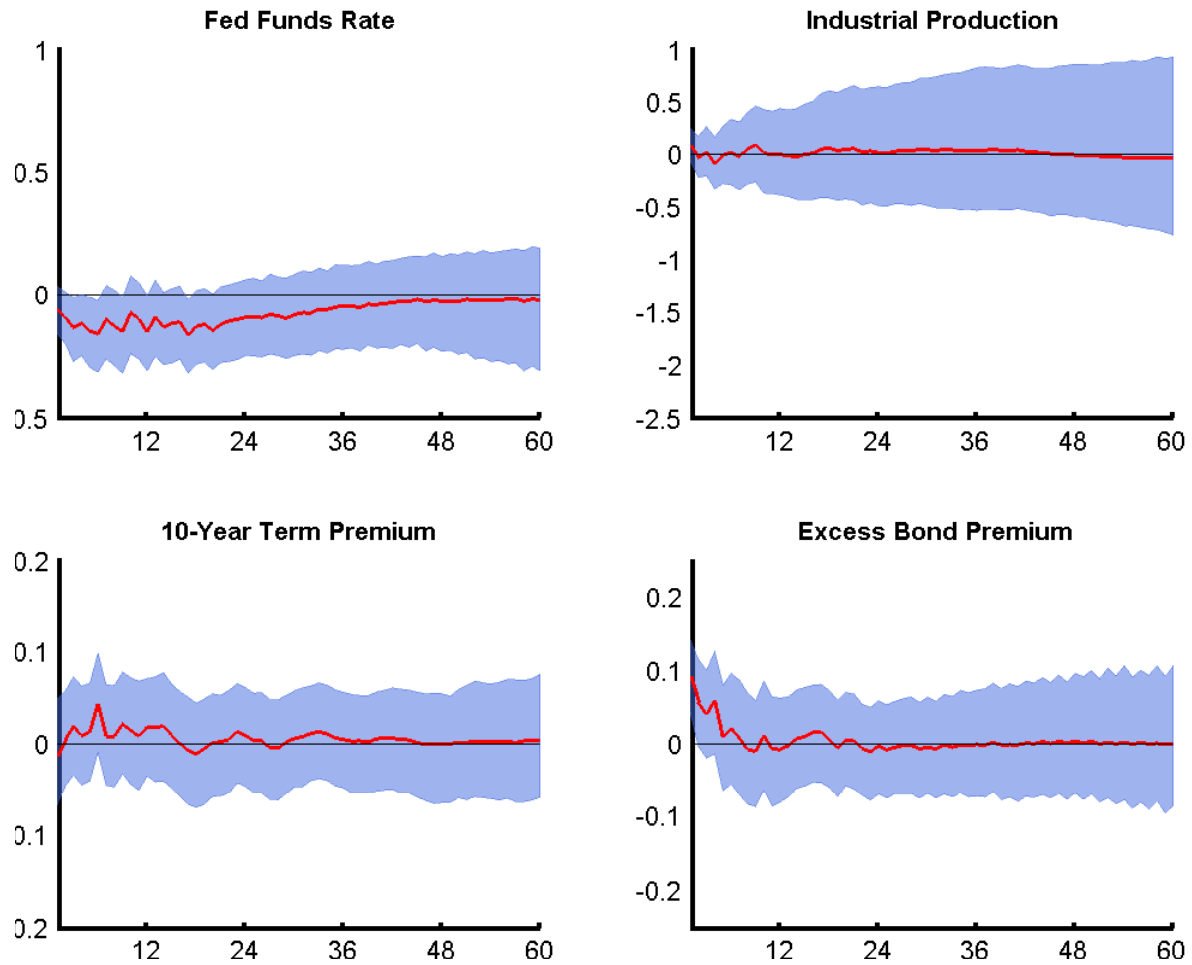
Financial shocks that explain the residual variation in the EBP do not elicit a decline in real activity.

5 The Financial Crisis and Unconventional Monetary Policy

6 Robustness

7 Conclusions

Figure 7: Macroeconomic Implications of a Financial Shock



NOTE: The solid line in each panel depicts the median impulse response of the specified variable to a 1 standard deviation financial shock. Shaded bands denote the 90-percent pointwise credible sets.

Table 2: The Effects of a Monetary Shock on Industrial Production

Horizon $h = 3$ Months				
$\epsilon_t^{MP} = m_t^{RM} + m_t^I$	0.44 [1.93]	0.05 [0.21]	— —	— —
$\epsilon_t^{MP} = m_t^{RM}$	— —	— —	-0.18 [0.45]	-0.55 [1.29]
Lagged EBP	—	X	—	X
Macro Controls	X	X	X	X
Adjusted R^2	0.25	0.40	0.25	0.39
Horizon $h = 12$ Months				
$\epsilon_t^{MP} = m_t^{RM} + m_t^I$	0.22 [1.51]	-0.17 [0.74]	— —	— —
$\epsilon_t^{MP} = m_t^{RM}$	— —	— —	-0.34 [1.13]	-0.72 [1.95]
Lagged EBP	—	X	—	X
Macro Controls	X	X	X	X
Adjusted R^2	0.11	0.34	0.11	0.35
Horizon $h = 18$ Months				
$\epsilon_t^{MP} = m_t^{RM} + m_t^I$	0.09 [0.44]	-0.20 [1.01]	— —	— —
$\epsilon_t^{MP} = m_t^{RM}$	— —	— —	-0.43 [1.14]	-0.53 [1.70]
Lagged EBP	—	X	—	X
Macro Controls	X	X	X	X
Adjusted R^2	0.02	0.14	0.02	0.14

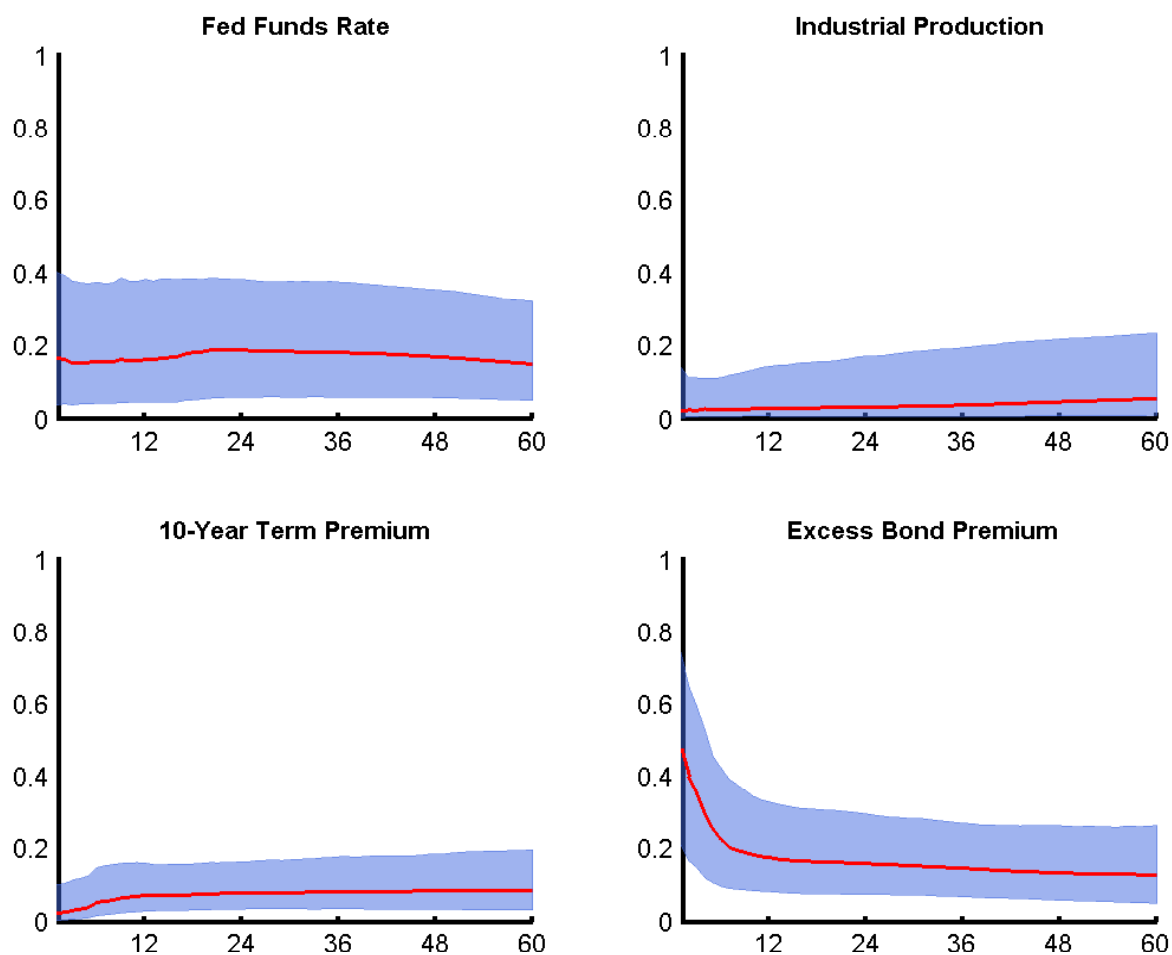
Notes: Sample period: 1994:1-2007:6. Dependent variable is $\Delta^h Y_{t+h}$ where Y_t denotes the log of industrial production (IP) in month t and h is the forecast horizon. The macro controls, X_t , include p lags of $[\Delta IP_{t-1}, R_{t-1}, \Delta PPCI_{t-1}]$, where p is determined by the AIC . Each numerical entry in the table denotes OLS point estimates with absolute asymptotic t -statistics in brackets, associated with HAC errors computed with a uniform weighting kernel for $h-1$ lags, along the lines of [Hodrick \(1992\)](#). An “X” means that the corresponding variables are included the regression.

Table 3: Effect of Monetary Policy Shocks on EBP

	Regression Estimates			
[5pt] $\epsilon_t^{MP} = m_t^{RM} + m_t^I$	0.64 [2.30]	0.68 [2.15]	— —	— —
$\epsilon_t^{MP} = m_t^{RM}$	— —	— —	0.80 [1.46]	0.79 [1.33]
$\sum_{j=1}^p EBP_{t-j}$	0.94 [397.06]	0.95 [308.37]	0.92 [422.18]	0.94 [299.46]
$\sum_{j=1}^p FFR_{t-j}$	— —	0.02 [2.20]	— —	0.02 [1.73]
$\sum_{j=1}^p \Delta \log IP_{t-j}$	— —	0.00 [0.00]	— —	0.00 [0.01]
Adjusted R^2	0.84	0.83	0.84	0.83

Notes: Sample period: 1994:1-2007:6. Dependent variable is EBP_t . The lag length for the controls p is determined by the *AIC*. Each numerical entry in the table denotes OLS point estimates with absolute asymptotic t -statistics / and F-statistics in brackets, associated with HAC standard errors. A $\sum$ term indicates that the entry in the table correspondings the to point estimate and F-statistic associated by the sum of the relevant coefficients.

Figure 8: Forecast Error Variance Decomposition of a Financial Shock



NOTE: The solid line in each panel depicts the median estimate of the portion of the forecast error variance of a specified variable attributable to a 1 standard deviation financial shock. Shaded bands denote the 90-percent pointwise credible sets.

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