CEP 15-02

Monetary News Shocks

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March 2015; revised 7 April 2016

CARLETON ECONOMIC PAPERS

Carleton UNIVERSITY

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Current version: April 7, 2016
Previous version: August 2015

Abstract

We pursue a novel empirical strategy to identify a monetary news shock in the U.S. economy. We use a monetary policy residual, along with other variables in a VAR, and identify a monetary news shock as the linear combination of reduced form innovations that is orthogonal to the current residual and that maximizes the sum of contributions to its forecast error variance over a finite horizon. Real GDP declines in a hump-shaped manner after a positive monetary news shock. This contraction in economic activity is accompanied by a fall in inflation and a rapid increase in the nominal interest rate.

JEL classification: E32, E52, E58

Key words: Monetary News Shocks, Monetary Policy Residual, Federal Funds Rate, Forward Guidance, New Keynesian DSGE Models

1 Introduction

Over at least the past 30 years there has been a significant effort in macroeconomics to identify the causal effects of monetary policy on economic activity by studying the impact of shocks to central bank rate decisions.[1] These shocks are interpreted as unanticipated or 'surprise' deviations of a central bank's policy rate from that implied by a systematic policy rule. Yet recent evidence suggests that at least some of these policy deviations may have not have been surprises after all (see, for example, Gurkaynak et al. (2005), Campbell, Evans, Fisher and Justiniano (2012) and Nakamura and Steinsson (2015)). Rather, a portion of FOMC rate decisions appear to be anticipated by private agents in advance of the formal rate-setting meetings. While a handful of recent papers have examined the impact of anticipated components of rate changes in both asset markets and on economic activity, relatively little work has been done to identify anticipated shocks using the structural vector autoregression (SVAR) framework. The main objective of our paper is to fill this gap. We pursue a novel empirical strategy to identify anticipated monetary policy shocks, or monetary news shocks—that is, pure changes in expectations about the non-systematic component of future monetary policy, orthogonal to non-systematic policy in the present—and determine their effects on the U.S. economy during the Greenspan-Bernanke era of Federal Reserve Chairmanship.

A potential source of monetary news shocks is the practice of forward-guidance through which a central bank provides information about the future course of monetary policy (Rudebusch and Williams (2008), Den Haan (2013), Svensson (2014)). To quantify the impact of forward guidance in dynamic stochastic general equilibrium (DSGE) models, recent work has included anticipated components to the exogenous (non-systematic) portion of the policy rule, similar to the news shock approach of Beaudry and Portier (2006).[2] Moreover, central banks are also currently using this approach for policy analysis.[3] A key contribution of our paper is to provide a benchmark for assessing whether the effects of monetary news shocks in DSGE models are consistent with their empirical

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[1]See Ramey (2016) for a detailed overview.

[2]See, for example, Lasseen and Svensson (2011), Milani and Treadwell (2012), Gomes et al. (2013), Harrison (2014), De Graeve et al. (2014), McKay et al. (2015), Gavin et al. (2015), among others.

[3]See, for example, the 'FRBNY DSGE Model' (Del Negro et al. (2013)) and the 'Chicago FED DSGE model' (Brave et al. (2012)).

counterparts. Our approach is in the spirit of the empirical literature that provided a benchmark for evaluating the effects of surprise monetary policy shocks in theoretical New Keynesian models (e.g. Christiano et al. (1999)).[4] We see our paper as a first-step towards developing this empirical benchmark for monetary news shocks.

While much of the literature has focused specifically on forward-guidance as a source of anticipated shocks to monetary policy, it need not be limited to that. Informal comments by policymakers to the media outside of regularly scheduled meetings, implicit communication implied by changes in the structure or membership of monetary policy operating committees, or public commentary by market observers all have the power to shape private agents' expectations about about future monetary policy rate decisions. Our approach allows for all channels through which changes in expectations may arise, focusing on identifiable pure changes in expectations about non-systematic policy themselves rather than changes in information that emerge through a particular channel.

Focusing on identifying pure changes in expectations about non-systematic monetary policy allows us to exploit empirical advances made in the news shock literature beginning with Beaudry and Portier (2006) that identify the impact of changes in expectations about total factor productivity. Our approach involves first constructing a monetary policy residual that measures deviations from an interest rate rule that tracks the observed federal funds rate well during 1988-2007.[5] Next, we propose a restriction to identify monetary news shocks using the maximum-forecast error variance (MFEV) approach within the SVAR framework similar to Barsky and Sims (2011), Francis et al. (2014), and Ben Zeev and Khan (2015). We identify a monetary news shock as the linear combination of reduced form innovations orthogonal to current policy residual that maximizes the sum of contributions to policy residual's forecast error variance over a finite horizon.

There are at least five advantages of our MFEV approach. First, by imposing orthogonality with respect to the current policy residual we can isolate a pure news component. We define it as

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[4]Other examples on identifying unanticipated monetary shocks include Bernanke and Mihov (1998), Romer and Romer (2004), and Barakchian and Crowe (2013).

[5]The baseline sample start-date is due to the availability of federal funds futures contract data from 1988. We choose 2007 as the sample end-date to first establish results without the influence of a variety of considerations in the post-2007 data (such as the financial crisis and the Great Recession (2008-09), Quantitative Easing, and the zero-lower-bound).

the monetary news shock since it accounts for the largest forecast error variance of the monetary policy residual over a particular horizon. Second, our approach complements event-studies or high-frequency approaches that identify the effects of specific policy announcements, for example, Gurkaynak et al. (2005) by allowing forward guidance to be communicated to private decision makers via all channels available to the FOMC as mentioned above. Third, the approach allows for a general view of forward guidance without taking an explicit view on the level of commitment.[6] Fourth, the MFEV approach to identifying news shocks remains relatively precise even in the presence of measurement errors. Fifth, we partition out systematic from non-systematic monetary policy and focus on changes in expectations about the non-systematic portion of policy. This reduces the chances that we pick up changes in expectations about monetary policy due to the expected future systematic response of the central bank to other non-monetary shocks.

Our paper is related to Campbell, Evans, Fisher and Justiniano (2012) who identify forward guidance shocks at the quarterly data frequency. Like them, we also consider the monetary policy residual based on an interest rate rule as the starting point for identifying monetary news shocks. The identification approaches, however, are completely different. Campbell, Evans, Fisher and Justiniano (2012) use quarterly aggregates of Blue Chip forecasts and interest rate futures prices in an interest rate rule with two lags of the interest rates and measures of the unemployment gap and inflation.[7] The monetary policy residual has both an unanticipated contemporaneous component and a forward guidance component that is anticipated by the public up to four quarters before the change in interest rate. The four quarter ahead forward guidance shock is constructed as a gap between two expectations: the four-period ahead expected interest rate minus the four-period ahead expected interest rate implied by the interest rate rule, given values of the parameters in the rule. They use a Generalized Method of Moments (GMM) approach to estimate the parameters

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[6]Campbell, Evans, Fisher and Justiniano (2012), for example, distinguish between Odyssean and Delphic central bank communications. The former implies public commitment to action whereas the latter do not. Calomiris (2012) has stressed the importance of forward guidance on the information provided to the public about the Federal Reserve's objectives and beliefs about the long-run natural rate of unemployment.

[7]Bruneau (2015) pursues a similar strategy and uses Survey of Professional Forecasters data to infer anticipated monetary shocks, via a Kalman Filtering estimation procedure and examines their effect on durable consumption.

and regress various macroeconomic variables on their news shock series finding that the latter raise interest rates but also raise economic activity and inflation, which is inconsistent with conventional wisdom.

Instead, we construct a policy residual conditional on calibrated values of the interest rate rule parameters, and identify monetary news shocks via the MFEV approach imposing orthogonality with the contemporaneous residual. The information set we consider includes credit spreads and interest rate futures in the VAR. We also impose orthogonality with respect to credit spreads to ensure that monetary news shocks are distinct from credit supply shocks.

Our main empirical findings are as follows. There is a hump-shaped decline in real GDP after a positive monetary news shock. This contractionary effect builds up gradually, reaching its trough after 1.5 years. The effect of monetary news shocks on economic activity is, therefore, quite persistent. The federal funds rate responds very little initially but then rises rapidly in subsequent periods, peaking after 3 quarters. The impact on inflation in the initial periods is muted, although there is a statistically significant fall in inflation that coincides with the trough in output. Interestingly, the responses of output and inflation are in line with conventional views of how monetary policy affects the economy. We find that excess bond premium rises significantly after a positive monetary news shock. This suggests imperfections in the financial markets play an important role in the transmission of monetary news shocks. Recently, Gertler and Karadi (2015) documented that positive monetary surprise shocks increase excess bond premiums. Our findings highlight a key point that both types of monetary shocks are likely to have a common transmission mechanism operating through the financial markets.

Over a one year horizon, monetary news shocks account for about 40% of the forecast variance in the policy residual, indicating that monetary news shocks are strongly present in the data. These shocks account for 40% and 50% of the one-year horizon forecast variance in federal funds futures and the federal funds rate, respectively. This suggests two things. First, that market participants' expectations about the stance of monetary policy are strongly affected by signals about future monetary policy. Second, that an important component of the Fed's determination of the interest

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rate is anticipated in advance. Monetary news shocks also account for about 12% of the forecast error variance in the excess bond premium.

Over the same horizon, monetary news shocks account for about 7% and 10% of the forecast variance in output and inflation, respectively, indicating that they play a relatively small role in short-run output and inflation fluctuations.

Are monetary news shocks related to unanticipated monetary policy shocks and non-monetary shocks? We find that the monetary news shock leads many unanticipated shocks previously identified in the literature by one to two quarters with correlation coefficients in the range of 0.2-0.4. The leading nature of the correlations suggests that a portion of the unanticipated shocks identified previously in the literature have news components. Additionally, monetary news shocks have a near-zero correlation with a variety of monetary and non-monetary shocks. This finding suggests that we are not confounding the effects of monetary news shocks with other types shocks.

Finally, we conduct a Monte Carlo exercise to demonstrate that our procedure correctly identifies monetary news. We apply our method to simulated data from a New Keynesian (NK) model with a Diamond-Mortensen-Pissarides (DMP) type labor market based on the modeling approach taken in the Christiano et al. (2016). The results of our experiment reveal that our method does a good job of identifying the true effects of monetary news shocks as well as the monetary news shock series itself.

In the remainder of the paper we proceed as follows. In Section 2, we construct the monetary policy residual and examine its properties. In Section 3 we introduce our empirical approach and identification methodology, and present the main results. In Section 4 we examine robustness. In Section 5 we present a New Keynesian DSGE model to perform a Monte Carlo exercise and to illustrate the responses of variables to monetary news shocks and contrast them with the estimated responses. Section 6 concludes.

2 Monetary policy residual

Consider a policy rule

$$i_t = g(\Omega_t) + \varepsilon_t, \tag{1}$$

where i_t is the nominal interest rate, Ω_t is the time t information set of the policymaker, $g(\Omega_t)$ is a function of the variables in the information set and denotes the unobserved systematic component of policy, and ε_t is a collection of both unanticipated and news shocks to the interest rate. Specifically, ε_t is given as

$$\varepsilon_t = v_t + n_{t-j}, \tag{2}$$

where v_t denotes the unanticipated shock at time t, and n_{t-j} is a news shock received j periods in advance of period t but that impacts the interest rate process in period t.

We consider a linear and historical rule that approximates $g(\Omega_t)$ in (1) over the sample period given as

$$\tilde{g}(i_{t-1}, \pi_t, u_t) \equiv 0.8 i_{t-1} + 0.2\Big(2 + \pi_t + 0.5\big(\pi_t - 2\big) + 2\big(6 - u_t\big)\Big), \tag{3}$$

where π_t is the inflation rate, and u_t is the unemployment rate.[8] This historical rule and its parameterization is similar to that in Clark (2012). The intercept of 2 captures the normal level of the real interest rate, the inflation target is assumed to be 2 percent, and the long-run normal level of unemployment is 6 percent. A policy rule like (3) with unemployment gap in place of the output gap is appealing for two reasons. First, it is often argued, as in Clark (2012), that unemployment is a better reflection of the maximum employment element of the Federal Open market Committee's dual mandate. Indeed, Campbell, Evans, Fisher and Justiniano (2012) also consider a policy rule with unemployment. Second, in the theoretical literature the output gap is a model-specific notion. It is typically defined as the gap between actual and the unobservable flexible price level of output (see Woodford (2003)). Determining an appropriate empirical counterpart to the output gap implied by theory is thus wrought with issues.[9] In this regard, using the unemployment gap

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[8]Later in the paper we investigate robustness to alternative specifications of the policy rule.

[9]See Galí (2011) for additional discussion.

serves as a practical empirical proxy for the economic 'slack' implied by the output gap of many theoretical models. Galí (2011) uses a New Keynesian model with unemployment to show that one can construct unemployment-based measures of the model-implied output gap. Moreover, he shows that a simple policy rule that responds to price inflation and the unemployment rate can not only approximate the optimal policy rule, but better capture movements in the Federal Funds rate from 1987 to 2008 than a similar policy rule using a more traditional HP-filtered measure of the output gap.

We then define the monetary policy residual, MPR_t, as

$$MPR_t = i_t - \tilde{g}(i_{t-1}, \pi_t, u_t), \tag{4}$$

that is, the residual obtained from netting out the assumed historical policy rule (3) from the nominal interest rate, using time series observations on i_t, π_t and u_t.

2.1 Properties of the Baseline Policy Residual

Figure 1 shows a plot of i_t and the baseline policy rule (3) from 1988Q4 to 2007Q4.[10] As is clear from the figure, the baseline policy rule (3) tracks quite closely the federal funds rates for much of the sample. Figure 2 shows MPR_t over this same period, constructed as in (4) using (3). Evidently, the residual moves systematically in and out of recessions, reflecting the unanticipated and news shocks relative to the historical rule.

3 Identifying monetary news shocks

In this section we present our empirical results for identifying the macroeconomic impact of monetary news shocks. We begin by first providing an overview of our empirical approach and then

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[10]The data were obtained from FRED Database at the Federal Reserve Bank of St. Louis. The effective federal funds rate (FF) is a quarterly average. The inflation rate is measured as annualized percentage of the growth in the GDP deflator (GDPDEF). The unemployment rate is the quarterly average of the civilian unemployment rate (URATE).

present our identification and estimation methodology in detail. Following this, we discuss the impulse responses and forecast error variance decompositions of the identified monetary news shocks. Next, we present robustness of the baseline results along a variety of dimensions.

Our assumptions in the previous section imply that

$$MPR_t \approx v_t + n_{t-j}, \tag{5}$$

which forms the basis of our identifying assumption that MPR_t is well approximated as a stochastic process driven unanticipated and news shocks. We use a VAR-based procedure using quarterly time series on MPR, the Federal Funds Rate (FF), the 6-months out expected interest rate from the Federal Funds Rate futures index (FFF), CPI inflation (CPI), the natural log of real GDP per capita (Y), and the credit spread measure constructed by Gilchrist and Zakrajsek (2012). The estimation period for this baseline specification is 1988Q4-2007Q4[11]

We include two forward-looking variables in the VAR, namely, FFF, and the credit spread index. First, the FFF is a contract for the average daily federal funds rate during the particular month of the contract (in our case, 6 months in the future). We include this futures data to exploit its forward-looking nature to capture variation due to changes in expectations about future monetary policy, much in the same way that the original news shock literature such as Beaudry and Portier (2006) used stock prices to capture unobserved changes in expectations about future productivity.

Second, any empirical analysis that deals with the recent credit-supply-driven period requires the inclusion of an appropriate measure of the credit spread so as to be able to ensure that credit supply shocks are not driving the results. We use the measure constructed by Gilchrist and Zakrajsek (2012), who use micro-level data to construct a credit spread index which they decomposed into a component that captures firm-specific information on expected defaults and a residual component

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[11]As noted in the introduction, we exclude the Great Recession period given its uniqueness and a variety of considerations. Campbell, Evans, Fisher and Justiniano (2012) similarly exclude the Great Recession period from their sample. Later in the paper we investigate robustness to including the Great Recession period.

that they termed as the excess bond premium (EBP).[12] We use the EBP series of Gilchrist and Zakrajsek (2012) as the measure of credit spread; Gilchrist and Zakrajsek (2012) interpret VAR innovations in this series as credit supply shocks.

3.1 Methodology

Let $y_t = [EBP_t \ \ MPR_t \ \ FF_t \ \ 100 - FFF_t \ \ \Delta \log(CPI_t) \ \ \log(Y_t)]'$ be a 6×1 vector of observables.[13] Let the VAR in the observables be given by

$$y_t = F_1 y_{t-1} + F_2 y_{t-2} + ... + F_p y_{t-p} + F_c + e_t \tag{6}$$

where F_i are 6×6 matrices, p denotes the number of lags, F_c is a 6×1 vector of constants, and e_t is the 6×1 vector of reduced-form innovations with variance-covariance matrix Σ. The reduced form moving average representation in the levels of the observables is

$$y_t = B(L)e_t \tag{7}$$

where $B(L)$ is a 6×6 matrix polynomial in the lag operator, L, of moving average coefficients and e_t is the 6×1 vector of reduced-form innovations. We assume that there exists a linear mapping between the reduced-form innovations and structural shocks, ε_t, given as

$$e_t = A\varepsilon_t. \tag{8}$$

Equation (7) and (8) imply a structural moving average representation

$$y_t = C(L)\varepsilon_t, \tag{9}$$

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[12]Gilchrist and Zakrajsek (2012) showed that their spread measure has better predicative power for macroeconomic variables than more standard credit spread measures such as the Baa-Aaa Moody's bond spread.

[13]Note that subtracting the federal funds rate futures index value from 100 yields the 6-months out level of the expected interest rate.

where $C(L) = B(L)A$ and $\varepsilon_t = A^{-1}e_t$. The impact matrix A must satisfy $AA' = \Sigma$. There are, however, an infinite number of impact matrices that solve the system. In particular, for some arbitrary orthogonalization, $\widetilde{A}$ (we choose the convenient Choleski decomposition), the entire space of permissible impact matrices can be written as $\widetilde{A}D$, where D is a 6×6 orthonormal matrix ($D' = D^{-1}$ and $DD' = I$, where I is the identity matrix).

The h step ahead forecast error is

$$y_{t+h} - E_t y_{t+h} = \sum_{\tau=0}^{h} B_\tau \widetilde{A}D\varepsilon_{t+h-\tau}, \tag{10}$$

where B_τ is the matrix of moving average coefficients at horizon τ. The contribution to the forecast error variance of variable i attributable to structural shock j at horizon h is then given as

$$\Omega_{i,j} = \sum_{\tau=0}^{h} B_{i,\tau} \widetilde{A}\gamma\gamma'\widetilde{A}'B'_{i,\tau}, \tag{11}$$

where γ is the jth column of D, $\widetilde{A}\gamma$ is a 6x1 vector corresponding with the jth column of a possible orthogonalization, and $B_{i,\tau}$ represents the ith row of the matrix of moving average coefficients at horizon τ. We index the unanticipated EBP and MPR shocks as 1 and 2, and the monetary news shock as 3 in the ε_t vector. Monetary news shocks identification requires finding the γ which maximizes the sum of contribution to the forecast error variance of MPR over a range of horizons, from 0 to H (the truncation horizon), subject to the restriction that these shocks have no contemporaneous effect on MPR. Formally, this identification strategy requires solving the

following optimization problem

$$\gamma^* = \operatorname*{argmax} \sum_{h=0}^{H} \Omega_{2,3}(h) \quad = \quad \operatorname*{argmax} \sum_{h=0}^{H} \sum_{\tau=0}^{h} B_{2,\tau} \widetilde{A} \gamma \gamma' \widetilde{A}' B'_{2,\tau} \tag{12}$$

$$\text{subject to} \quad \widetilde{A}(1,j) \quad = \quad 0 \ \ \forall j > 1 \tag{13}$$

$$\widetilde{A}(2,j) \quad = \quad 0 \ \ \forall j > 2 \tag{14}$$

$$\gamma(1,1) \quad = \quad 0 \tag{15}$$

$$\gamma(2,1) \quad = \quad 0 \tag{16}$$

$$\gamma'\gamma \quad = \quad 1. \tag{17}$$

The first four constraints impose on the identified news shock to have no contemporaneous effect on EBP and MPR. That is, our news shocks is orthogonal to both unanticipated credit supply shocks as well as unanticipated monetary policy shocks. The fifth restriction that imposes on γ to have unit length ensures that γ is a column vector belonging to an orthonormal matrix. This normalization implies that the identified shocks have unit variance.

We follow the conventional Bayesian approach to estimation and inference by assuming a diffuse normal-inverse Wishart prior distribution for the reduced-form VAR parameters. Specifically, we take 2000 draws from the posterior distribution of reduced form VAR parameters $p(F, \Sigma \mid data)$, where for each draw we solve optimization problem (12); we then use the resulting optimizing γ vector to compute impulse responses to the identified shock.[14] This procedure generates 2000 sets of impulse responses which comprise the posterior distribution of impulse responses to our identified shock.

Our benchmark choice for the truncation horizon is H=5, a horizon that generally corresponds to an operative period of monetary policy of less than two years in the future. We examine the sensitivity of our identification to changes in this horizon length in Section 4. The number of lags we use in the VAR is 5, which is the optimal lag length according to the Akaike Information

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[14]Note that F here represents the stacked $(6 \times (p+1)) \times 6$ reduced form VAR coefficient matrix, i.e., $F = [F_1, ..., F_p, F_c]'$.

Criterion.[15]

3.2 Merits of the MFEV Approach

In this section we highlight two advantages of the MFEV approach described above. First, the MFEV approach is superior to imposing an orthogonalization restriction alone. One may argue, for example, to undertake an approach analogous that of Beaudry and Portier (2006) who identify total factor productivity (TFP) news shocks as innovations in stock prices orthogonalized with respect to TFP. In our setting, this would entail identifying monetary news shocks as innovations in FFF orthogonalized with respect to MPR. While such a restriction would partition out the impact of the unanticipated monetary shock ε_t^0, nothing assures us that monetary news shocks are the only shock not affecting MPR on impact and affecting the FFF. For example, consider a persistent demand shock, such as an expansionary preference shock. In this case, output rises in the present and future, and consistent with the monetary policy rule, the FED systematically raises rates in the present. Moreover, agents expect the FED to systematically raise rates in the future, causing FFF to rise in the present. Since such a scenario involves an innovation to FFF in the present but no change in the policy residual in the present, the Beaudry and Portier (2006) approach would capture such a shock, despite that it involves no policy deviation in the future. Instead, to correctly identify our anticipated monetary deviations and exclude shocks such as the above, we require a method that not only orthogonalizes the news shock with respect to contemporaneous policy deviations, but also isolates shocks that best explain future policy deviations.

Second, the MFEV identification approach is little affected in the presence of measurement errors. If the MPR, for example, inherits a measurement error from unemployment and inflation, this would hurt the identification of the surprise shock (which is not the objective of this paper) but not the news shock. More generally, any other economic shock that affects FFR on impact and is not being fully reflected by our specification of the rule, would reduce the precision of the surprise shock identification but not that of the news shock (or at least to a much lesser extent).

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[15]We have confirmed the robustness of our results to different VAR lag specifications. These results are available upon request from the authors.

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Ben Zeev and Pappa (2015) present formal evidence on this point. They show that the presence of measurement errors invalidates the identification of surprise defense shocks whereas the MFEV identification approach remains unaffected by it.

3.3   Impulse response functions

The solid line in Figure 3 shows the median of the posterior distribution of impulse responses to a positive one standard deviation monetary news shock (that is, an anticipated tightening of policy). The dotted lines are the 84th and 16th percentiles of the posterior distribution.

Following a positive monetary news shock there is an immediate and persistent decline in real GDP. The response is hump-shaped, reaching its trough after about 10 quarters. The impact on inflation in the initial periods is muted, although there is a statistically significant fall in inflation that coincides with the trough in output. The policy residual does not change in the initial period by construction because of the orthogonality restriction. It rises quickly, however, in subsequent periods, peaking after 1 quarter. The FFR which is left unrestricted, changes very little initially, and then peaks after 3 quarters. The fact that the FFR does not move in the initial period when the policy residual also does not move shows that there is no significant systematic effect of monetary policy in the initial period.The FFF rises immediately, consistent with the eventual rise in the federal funds rate itself. Finally, the excess bond premium does not move on impact by construction, after which it significantly rises in a hump-shaped manner. One can interpret this response in terms of a financial accelerator mechanism that takes place in response to the contraction in the economy. Thus, financial market imperfections play an important role in the transmission of monetary news shocks.

3.4   Forecast error variance decompositions

Figure 4 shows the median forecast error variance decompositions along with the 68% posterior probability bands. Over a one year horizon, monetary news shocks account for 42% of the forecast variance in the policy residual, indicating that monetary news shocks are strongly present in the

data. These shocks account for 46% and 42% of the forecast variance in federal funds futures and the federal funds rate. This suggests two things. First, market participants' expectations about the stance of monetary policy are strongly affected by signals about future monetary policy. Second, an important component of the Fed's determination of the interest rate is anticipated.[16] Monetary news shocks account for 8% of the forecast variance in output, indicating that they play a relatively small role in short-run output fluctuations.[17]

3.5   Narrative interpretations of monetary news shocks

Figure 5 shows the estimated news shocks realization series. It is important to note that some of the large realizations of our shock series are consistent with the large realizations of the news shocks series of Campbell, Fisher and Justiniano (2012) and hence their narrative interpretation given in Campbell, Fisher and Justiniano (2012) can also be applied here. Specifically, we obtain a very large realization of 2.1 standard deviations in 1994:Q4 (the second largest positive realization in our sample); as discussed in Campbell, Fisher and Justiniano (2012), this could reflect new information transmitted to the public by the congressional testimony by Greenspan on December 7 which was interpreted by The New York Times the following day as "..an unusually clear signal that the Federal Reserve will continue raising interest rates...".

Moreover, we also obtain two large negative realizations in the first and second quarters of 1995 of −1.1 and −1.6 standard deviations, respectively. The first of these can be explained by Greenspan's testimony before the Senate Banking Committee on February which the Wall Street Journal reported the following day as follows: "Mr. Greenspan warned that inflation may still be a threat but added that the Fed might not raise interest rates again even if inflation starts to rise again...". The second realization of 1995:Q2 can be potentially attributed to the June 20

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[16]Our news shocks explain 66% of the contemporaneous variation in the futures index, stressing that market participants react to new information imminently.

[17]Interestingly, our empirical findings on variance decompositions of output and interest rates are similar to those reported in Campbell, Fisher and Justiniano (2012) that are based on an estimated DSGE model. They find that monetary news shocks account for 9% of the forecast variance of GDP and 55% of the forecast variance of the nominal interest rate.

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Greenspan speech at the Economic Club of New York, in which he stressed that the price stability mandate will be the dominant consideration in the subsequent FOMC meeting and that he felt that inflationary pressures were easing.[18]

3.6   Relation with other monetary shocks

We assess whether our monetary news shock is related to the monetary shocks identified in the literature. We report the cross-correlations of our recovered shock series with the unanticipated monetary shocks identified in Bernanke and Mihov (1998), Christiano et al. (1999), Romer and Romer (2004), Sims and Zha (2006), and Barakchian and Crowe (2013), and the monetary news shocks identified in Campbell, Evans, Fisher and Justiniano (2012). All unanticipated monetary shock series were taken from the data Appendix file of Barakchian and Crowe (2013) which contains updated versions of these shock series that cover the sample period 1989Q4-2007Q4, apart from the Bernanke and Mihov (1998) series which starts in 1990Q1. The monetary news shock of Campbell, Evans, Fisher and Justiniano (2012) covers the sample period of 1996Q1-2007Q2.[19]

Figure 6 shows the median and 16th and 84th percentiles of the cross-correlations between monetary news shocks and all the other monetary shocks. The contemporaneous correlation is approximately zero in nearly all the cases indicating that monetary news shocks are distinct from unanticipated monetary shocks. However, a key property of our monetary news shocks stands out—that it strongly leads all of the considered monetary shocks, including the news shock from Campbell, Evans, Fisher and Justiniano (2012). VARs that produced unanticipated monetary shocks had insufficient information relative to ours, thereby resulting in identified shocks that are

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[18]Since the MPR can be viewed as an unexpected shift in the average preferences of policymakers as measured by the weight they place on the importance of inflation relative to output fluctuations (see Christiano et al. (1999)), this Greenspan speech can be interpreted as providing new information that the FOMC will be putting a larger than average weight on inflation in the near future. This implies an excessively expansionary monetary policy news shock given that inflationary pressures were easing. We emphasize that the two possible qualitative interpretations that we provide here are both examples of informal communications made through channels outside of formal FOMC meetings. That is, they do not constitute the formal notion of forward-guidance statements provided during FOMC meetings.

[19]We thank Christopher Crowe for kindly sharing the data from Barakchian and Crowe (2013) and Jonas Fisher for sharing the news shocks from Campbell, Evans, Fisher and Justiniano (2012).

effectively a combination of both the true unanticipated monetary shock as well as lags of monetary news shocks. Similarly, our monetary news shock significantly leads the Campbell, Evans, Fisher and Justiniano (2012) shock (35% one-lagged correlation) due to the larger information set we use. The information set in Campbell, Evans, Fisher and Justiniano (2012) consists of the observables in their Taylor rule (unemployment, inflation, and the federal funds rate), whereas ours contains additional informationally important variables like credit spreads and the federal funds rate futures index.

4 Robustness

We now explore the robustness of our results along a variety of dimensions. Each exercise is motivated by a specific aspect of the assumptions behind the baseline methodology.

4.1 Alternative monetary policy residuals

Proceeding as before when constructing our baseline residual, we now repeat that exercise, this time using two alternative 'Taylor rules' that have also been considered in the literature. The first is exactly identical to that considered in Clark (2012), i.e., one that uses the lagged unemployment rate rather than the current unemployment rate. We view this as a less appealing specification as it is reasonable to assume that the Fed responds within the quarter to variations in the unemployment rate. That said, it is still an informative exercise to conduct. The results are shown in Figure 7a. Impulse responses for all variables are both qualitatively and quantitatively similar to the benchmark ones, with output and inflation both declining and the news shock continuing to raise interest rates with a delay.

The second residual is based on a policy rule that uses the output gap instead of the unemployment gap. Specifically,

$$\tilde{g}(i_{t-1}, \pi_t, y_t) \equiv 0.8i_{t-1} + 0.2(1.5\pi_t + 0.2y_t) \tag{18}$$

where y_t is the output gap. The policy rule in (18) is similar to the one considered in Rudebusch

(2002). Although our baseline rule has advantages relative to (18) for the reasons discussed in Section 2, it is nevertheless useful to confirm that the baseline results are robust to perturbation of the policy rule. The results are shown in Figure 7b. It is noticeable that impulse responses for all variables are both qualitatively and quantitatively similar to the benchmark ones.

4.2 Including the Great Recession period

We apply our identification method to the sample period of 1988Q4-2012Q4, where the ending date is dictated by the availably of the updated credit spread series of Gilchrist and Zakrajsek (2012). Figure 8 presents the impulse responses obtained from this sub-sample estimation exercise. It is apparent that the main results continue to hold: output and inflation fall and the news shocks raises the interest rate and the futures index after a positive monetary news shock. The excess bond premium increases on impact, and but displays a more significant increase at horizons longer than the baseline case.

4.3 An alternative identification approach based on the Federal Funds Rate futures

As we discussed earlier, analogously to how Beaudry and Portier (2006) identified TFP news shocks as the VAR innovations in stock prices orthogonal to current TFP, one may argue that a suitable identification approach for monetary news shock would use the VAR innovation in the federal funds futures rate orthogonal to MPR and EBP. We view this approach as inferior to the MFEV-based approach because it is likely to capture other economic shocks that are orthogonal to EBP and MPR and which affect the federal funds futures rate. That said, it is worthwhile to implement this alternative identification approach and examine how the results from it fare against our baseline results.

Figure 9 presents the impulse response to a shock to the federal funds futures rate orthogonalized with respect to EBP and MPR. Notably, output now rises on impact, and inflation significantly

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rises. Moreover, the response of MPR is weaker than the one obtained in the benchmark case.[20] As discussed in Section 3.2, one of the disadvantages of imposing *only* the orthogonality condition is that the confounding effect of other demand shocks can still be present. The positive response of output on impact is indicative of this problem, and the positive inflation response which also never turns significantly negative is especially emblematic of this concern. These results, therefore, support the view that using the VAR innovation in the federal funds futures rate for identifying the monetary news shock is inferior to the MFEV-based approach as it is likely to pick up other shocks rather than just the monetary news shock. The difference between the impulse responses to the shock from Figure 9 and the baseline monetary news shock may seem somewhat surprising given that the median correlation between the two identified shocks is 85%. Although high, this correlation still indicates an important wedge that, coupled with the observed differences in the impulse responses to the two shocks, supports the view that using the MFEV-based estimation approach is more suitable for identifying monetary news shocks.

4.4   Relationship with other macroeconomic shocks

We now assess whether our monetary news shock is related to other important identified macroeconomic shocks normally considered to be potentially important sources of business cycle fluctuations: the TFP news shock identified in Barsky and Sims (2011); shock to the real price of oil, Romer and Romer (2010) exogenous tax shock measure, Ramey (2011) defense news shock, the uncertainty shock from Bloom (2009) (that appears in his Figure 1), and the innovation to the U.S. economic policy uncertainty index of Baker et al. (2015).[21] We report the correlation between up to four lags and leads of our identified monetary news shock and these other shocks. The results are presented in Figure 11, where the median and 16th and 84th percentiles of the correlations between up to four

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[20]The federal funds futures rate shock accounts for only 30% of the one-year variation in MPR, compared to 42% accounted for by the MFEV news shock. The results on the forecast error variance contributions of the federal funds futures rate shock are available upon request from the authors.

[21]Apart from the Barsky and Sims (2011) TFP news and unanticipated shock series which were used in their raw form, all other shocks were constructed as the residuals of univariate regressions of each of the five raw shock measures on four own lags.

lags and leads of our news shock and the two technology news socks are shown. The results indicate that the cross-correlations are small, with all median correlations lower than 15% in absolute value. This exercise confirms that our results are not driven by other macroeconomic shocks.

4.5   Longer truncation horizons

Figure 10 shows the median and 84th and 16th percentiles of the posterior distribution of impulse responses to a positive one standard deviation monetary news shock obtained from setting the truncation horizons to H=9, i.e., a two year operative anticipation horizon over which forward guidance is conducted. Our view, which is also consistent with that of Campbell, Evans, Fisher and Justiniano (2012), is that the baseline one year horizon is very much reflective of the horizon over which forward guidance is relevant. Even though the effects on output and inflation are quantitatively weaker than the baseline effects, it is clear from Figure 10 that the longer horizon of two years does not change the qualitative nature of our baseline results.

5   Monte Carlo Exercise

We now illustrate the suitability of our identification method for identifying monetary news shocks by applying our identification procedure on simulated data from a standard small-scale New Keynesian (NK) model with a Diamond-Mortensen-Pissarides (DMP) type labor market based on the modeling approach taken in Christiano et al. (2016).[22] We embed the DMP model into the NK model to introduce unemployment into our setting so as to be able to include it in our monetary policy rule, in accordance with our empirical monetary policy rule. The model we use is quite standard and we present below its main building blocks.

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[22]Christiano et al. (2016) estimate two versions of a full-scale NK-DMP model: One version pursues a variant of the labor market modeling approach taken in Hall and Milgrom (2008), in which real wages are determined by alternating offer bargaining, whereas the other version assumes real wages are determined by Nash bargaining. For simplicity, we abstract from capital and assume Nash wage bargaining; extending our setting to include capital and assume alternating offer bargaining rather than Nash bargaining would not change the Monte Carlo results but would complicate the exposition.

5.1 Model Used for Simulations

Households The model consists of many identical infinitely-lived households, each having a unit measure of workers which are all sent to the labor market where they are either employed or unemployed. The household chooses a sequence of state-contingent consumption and bond levels C_t and B_{t+1} to solve the following problem:

$$\max \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \log(C_t) \tag{19}$$
$$s.t \ P_t C_t + B_{t+1} \leq W_t l_t + (1 - l_t)D + R_b B_t + T_t,$$

where l_t is the share of employed workers belonging to the household (i.e., $1 - l_t$ represents the unemployment rate), D is unemployment benefits, R_t is the gross nominal interest rate, W_t is the nominal wage, and T_t is profits net of taxes.

Final Good Producers There are identical and competitive final good firms that supply to the households the homogeneous final good Y_t, produced using specialized inputs, $Y_t(i)$, with the following technology:

$$Y_t = \left[\int_0^1 (Y_t(i))^{\frac{1}{\lambda}} di \right]^{\lambda}, \tag{20}$$

where $\lambda > 1$. The final goods producers purchased their specialized inputs $Y_t(i)$ from monopolistically competitive retailers at price $P(i)$ and choose $Y_t(i)$ to solve the following profit maximization problem subject to production technology (20):

$$\max_{Y_t(i)} P_t Y_t - \int_0^1 P_t(i) Y_t(i) di, \tag{21}$$

where P_t is the price of the final good. The solution to this problem yields the first order condition

$$Y_t(i) = \left(\frac{P_t}{P_t(i)} \right)^{\frac{\lambda}{\lambda - 1}}. \tag{22}$$

Retailers There is a continuum of retail firms which are monopolists in product markets and competitive in factor markets. Each retail firm produces input good $Y_t(i)$ according to the following technology:

$$Y_t(i) = exp(a_t)h_t(i), \tag{23}$$

where a_t represents technology and $h_t(i)$ is the quantity of an intermediate good purchased by the ith retailer. This good is purchased in competitive markets at the price P_t^h from a *wholesaler*. Nominal price rigidities are introduced into the model via a Calvo (1983) pricing scheme where retail firms can readjust prices with probability $1 - \xi_p$ in each period. The optimal price set by the firm that is allowed to re-optimize its price is obtained from solving the following optimization problem:

$$\max_{\widetilde{P}_t(i)} \mathbb{E}_t \sum_{s=0}^{\infty} \xi^{t+s} \frac{\beta^{t+s} C_t P_t}{C_{t+s} P_{t+s}} \left[P_t(i) - MC_{t+s} \right] Y_{t+s}(i)$$

$$s.t \ \ Y_t(i) = \left(\frac{P_t}{P_t(i)} \right)^{\frac{\lambda}{\lambda-1}}, \tag{24}$$

where $\widetilde{P}_t(i)$ is the newly set price; ξ is the Calvo (1983) probability of being allowed to optimize one's price; $MC = \frac{P_t^h}{exp(a_t)}$ is the firm's nominal marginal cost; $\frac{\beta^{t+s} C_t P_t}{C_{t+s} P_{t+s}}$ is the nominal discount factor for households; and the constraint in (24) is simply the demand curve, (22), obtained as the first order condition of the maximization problem solved by the final good producer.

Wholesalers, Workers and the Labor Market The law of motion for aggregate employment, l_t, is given by:

$$l_t = (\rho + x_t)l_{t-1}, \tag{25}$$

where ρ is the probability that a given firm/worker match continues from one period to the next and therefore ρl_{t-1} denotes the number of workers that were attached to firms in period $t-1$ and remain attached at the start of period t; additionally, x_t represents the hiring rate and thus $x_t l_{t-1}$

is the number of new firm/worker matches at the start of period t.

The number of workers searching for work at the start of period t is $1 - \rho l_{t-1}$, the sum of the number of unemployed workers in period $t - 1$, $1 - l_{t1}$, and the number of workers that separate from firms at the end of $t - 1$, $(1 - \rho)l_{t1}$. Hence, we can define the probability, f_t, that a searching worker meets a firm as given by

$$f_t = \frac{x_t l_{t-1}}{1 - \rho l_{t-1}}. \tag{26}$$

It is assumed there is a large number of identical and competitive wholesalers that produce the intermediate good using only labor, l_t, which has a fixed marginal cost of unity. Wholesalers' output is sold to the retailers for real price, $\vartheta_t = \frac{P_t^h}{P_t}$. At the start of period t, a wholesaler firm that wishes to meet a worker must post a vacancy and pay a fixed cost, κ. Immediately after they meet, the worker and firm agree on a wage via bilateral bargaining (to be discussed below) and begin employment where the match survives into period $t + 1$ with fixed, exogenous probability ρ. The value to a firm of employing a worker at the equilibrium real wage rate, w_t, is denoted by J_t, which satisfies the following recursive relationship:

$$J_t = \vartheta_t - w_t + \rho \mathbb{E}_t m_{t+1} J_{t+1}, \tag{27}$$

where $m_{t+1} = \beta \frac{C_t}{C_{t+1}}$ is the real discount factor. Due to free entry, firm profits must be zero:

$$\kappa = J_t. \tag{28}$$

The value to a worker of being matched with a firm that pays w_t in period t is denoted by V_t:

$$V_t = w_t + \mathbb{E}_t m_{t+1}[\rho V_{t+1} + (1 - \rho)(f_{t+1} V_{t+1} + (1 - f_{t+1}) U_{t+1})], \tag{29}$$

where U_{t+1} is the value of being an unemployed worker in period $t + 1$, for which the recursive

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representation is:

$$U_t = D + \mathbb{E}_t m_{t+1}[f_{t+1}V_{t+1} + (1 - f_{t+1})U_{t+1}]. \tag{30}$$

We now turn to the determination of the real wage. Workers and firms choose a wage that solves a joint optimization problem, taking as given future wages:

$$\max_{w_t} (V_t - U_t)^\eta J_t^{1-\eta}, \tag{31}$$

where η is the share of total surplus given to workers. The first order condition of this problem yields the familiar *Nash sharing rule*:

$$J_t = \frac{1 - \eta}{\eta}(V_t - U_t). \tag{32}$$

Finally, we assume a standard matching function that determines how vacancies are matched with workers:

$$x_t l_{t-1} = \sigma_m (1 - \rho l_{t-1})^\sigma (l_{t-1}v_t)^{1-\sigma}, \tag{33}$$

where $l_{t-1}v_t$ denotes the total number of vacancies and v_t denotes the vacancy rate. Dividing Equation (33) by $1 - \rho l_{t-1}$ obtains

$$f_t = \frac{x_t l_{t-1}}{1 - \rho l_{t-1}} = \sigma_m \left(\frac{l_{t-1}v_t}{1 - \rho l_{t-1}}\right)^{1-\sigma} = \sigma_m \Theta_t^{1-\sigma}, \tag{34}$$

where Θ_t denotes labor market tightness.

Labor and Goods Markets Clearing Total purchases of homogeneous inputs by intermediate good producer amount to

$$h_t = \int_0^1 h_t(i)di, \tag{35}$$

and labor market clearing takes place when $h_t = l_t$. The goods market clearing condition is

$$Y_t = C_t + \kappa x_t l_{t-1}. \tag{36}$$

Monetary Policy To close the model, we assume that monetary policy is conducted in accordance with our empirical policy rule:

$$\frac{R_t}{\bar{R}} = \left(\frac{R_{t-1}}{\bar{R}}\right)^{\alpha} \left[(1 + \pi_t)^{\phi_{-}\pi}(u_t - u_s s)^{\phi_u}\right]^{1-\alpha} exp(\epsilon_t), \tag{37}$$

where $\bar{R}$ is the steady state gross nominal interest rate; $u_t = 1 - l_t$ is the unemployment rate and $u_s s = 1 - l_s s$ is the steady state unemployment rate; and ϵ_t is the monetary policy residual.

Exogenous Stochastic Processes We assume that the monetary policy residual is driven by both unanticipated monetary policy shocks, v_t, as well as monetary news shocks, n_t, as follows:

$$\epsilon_t = \rho_\epsilon \epsilon_{t-1} + v_t + n_{t-1}. \tag{38}$$

The technology variable a_t is assumed to follow a first order AR process with technology shock z_t:

$$a_t = \rho_a a_{t-1} + z_t. \tag{39}$$

As we discuss below, we apply our estimation procedure on five-variable VARs. Hence, we attach two measurement errors to inflation and output so as to avoid stochastic singularity, resulting in our model having a total of five shocks.

Solving and Calibrating the Model We solve the model by approximating the equilibrium conditions around the non-stochastic steady state. The parameterization is summarized in Table 1, which includes the calibration for both the model's structural parameters as well as the standard deviations of the output and inflation measurement errors. We calibrate the parameters in the monetary policy rule in accordance with our empirical rule; the persistence parameter in the policy

residual process, ρ_ϵ, as well as the standard deviations of the monetary surprise and news shocks, were set such that the responses of the policy residual to the monetary shocks matches as closely as possible their data counterparts.[23] The rest of the calibration of the goods and labor markets is fairly standard, where the deep parameters of the labor market were set such that the vacancy filling rate (x_t/v_t) is equal to 0.7 in the steady state; steady state unemployment is 0.06; the replacement ratio (D/w_t) is equal to 0.4 in the steady state; and recruitment costs as a share of output $(x_t l_{t-1}/Y_t)$ are equal to 0.01 in the steady state.

5.2   Simulation Results

We simulate 2000 sets of data with 72 observations each from the above model. For each simulation, we estimate the median impulse response from a Bayesian VAR based on 2000 draws from the posterior distribution of the VAR parameters; we include in the Monte Carlo VAR the same variables that we used in the empirical exercise, where the futures variable is measured by the two period ahead expectation of the nominal interest rate. The only difference in our Monte Carlo exercise relative to the empirical VAR is that we do not include a credit supply shock measure because our theoretical model, for the sake of simplicity, does not contain a financial accelerator element and therefore a credit supply shock can not be formally introduced into the model. We draw all of the model's shocks from the normal distribution.

Figure 12 depicts both the theoretical and estimated median impulse responses averaged over the simulations to a monetary news shock. The theoretical responses are represented by the solid lines and the average estimated median responses over the simulations are depicted by the dashed lines, with the dotted lines depicting the 16th and 84th percentiles of the Monte Carlo distribution of estimated median impulse responses. Overall, results can be viewed as quite encouraging. The identification of the impulse responses is largely unbiased and performs reasonably well.[24]   Ac-

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[23]It is also worth noting that the theoretical contribution of the news shock to the variation in the policy residual, output, and inflation is generally consistent with its data counterpart.

[24]It is important to note that the fairly modest downward bias at the impact horizon is mostly a result of the small sample size at hand; e.g., extending the sample size by 25 years, or 100 observations, largely removes this bias and produces a mean correlation of over 90% between the estimated and true news shock

cordingly, the mean correlation between the estimated news shock series and the true one is 72%, suggesting that the shock series is identified fairly well.

6   Conclusion

We pursue a novel empirical strategy to identify monetary news shocks and determine their effects on the US economy during 1988-2007. Starting with a parameterized policy rule we construct a policy residual. Using the maximum-forecast error variance (MFEV) approach, we then identify a monetary news shock as the linear combination of reduced form innovations that is orthogonal to the current residual and that maximizes the sum of contributions to its forecast error variance over a finite horizon. Real GDP declines in a persistent hump-shaped manner after a positive monetary news shock. This contraction in economic activity is accompanied by a fall in inflation and a rapid increase in the nominal interest rate. These effects of monetary news shocks are qualitatively similar even when the sample is extended to include the Great Recession.

Our results suggest that market participants' expectations about the stance of monetary policy are strongly affected by signals about future monetary policy. Moreover, an important component of the Fed's determination of the interest rate is anticipated in advance. The sharp increase in the excess bond premium after a positive monetary news shock points to the important role of financial frictions in the transmission of this shock.

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series.

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Table 1: **Monte Carlo Experiment: Model Parameterization**.

β	Discount Factor	0.99
λ	Elasticity of Demand for Specialized Intermediate Good.	6
ξ	Price Stickiness Parameter for Intermediate Good Producers	0.75
ρ	Match Survival Rate	0.9
κ	Fixed Cost of Hiring	0.1
η	Workers' Share of Surplus	0.9
σ	Power on Unemployment in Matching Function	0.5
σ_m	Parameter in Matching Function	0.65
ϕ_π	Coefficient on Inflation in Taylor Rule	1.5
ϕ_u	Coefficient on Unemployment in Taylor Rule	-2
α	Taylor Rule Persistence	0.8
ρ_ϵ	Taylor Rule Residual Persistence	0.5
ρ_α	Technology Shock Persistence	0.95
σ_v	Monetary Surprise Shock Standard Deviation	0.0018
σ_n	Monetary News Shock Standard Deviation	0.0016
σ_a	Technology Shock Standard Deviation	0.0045
σ_y	Output Measurement Error Standard Deviation	0.001
σ_π	Inflation Measurement Error Standard Deviation	0.001

Notes: This table reports the values assigned to the parameters of the model used to generate the data in the Monte Carlo experiment of Section 5.

Figure 1: **Federal Funds Rate and the Baseline Policy Rule, 1988Q4-2007Q4.**
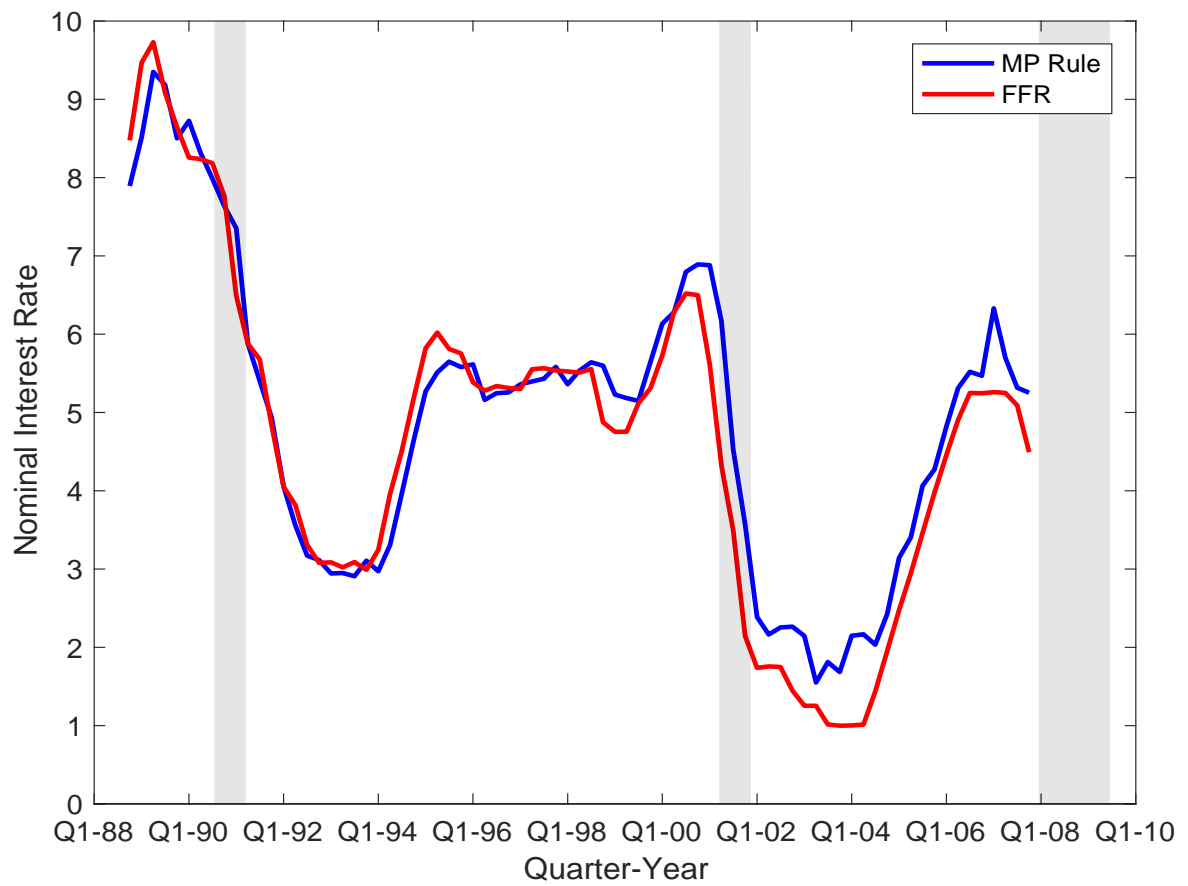
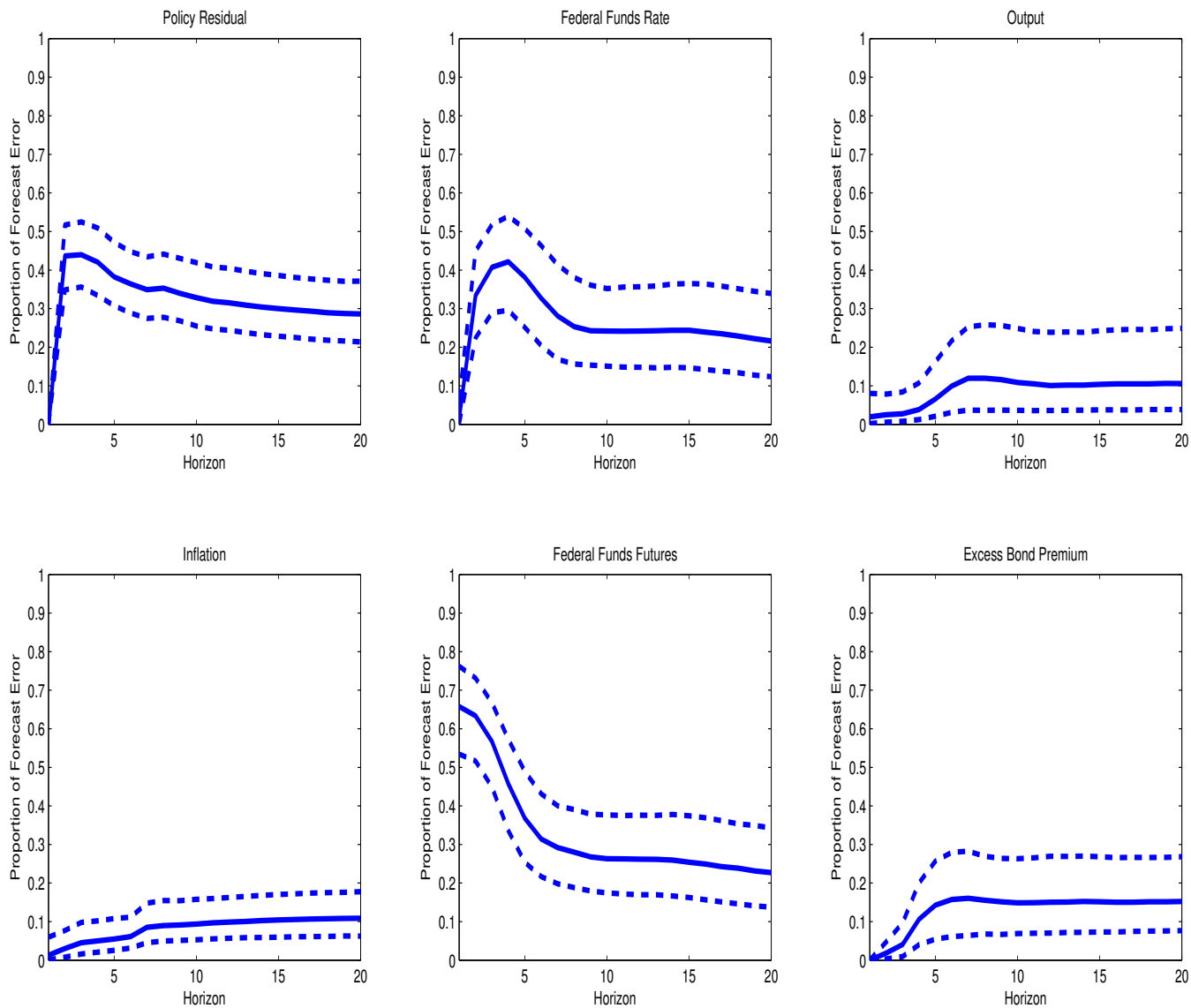Figure 2: **Baseline Policy Residual, 1988Q4-2007Q4**.
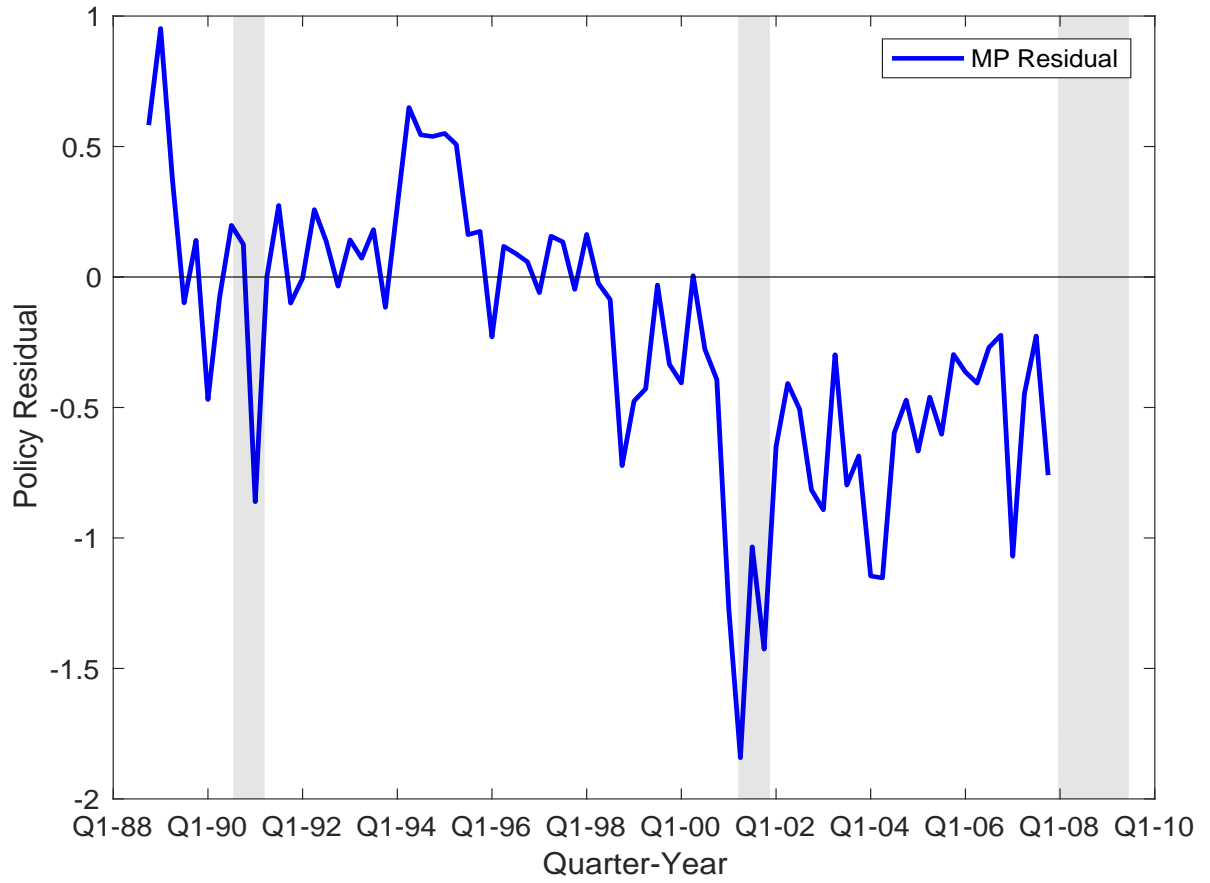
Figure 3: **Impulse Responses to a Monetary News Shock: Baseline**.
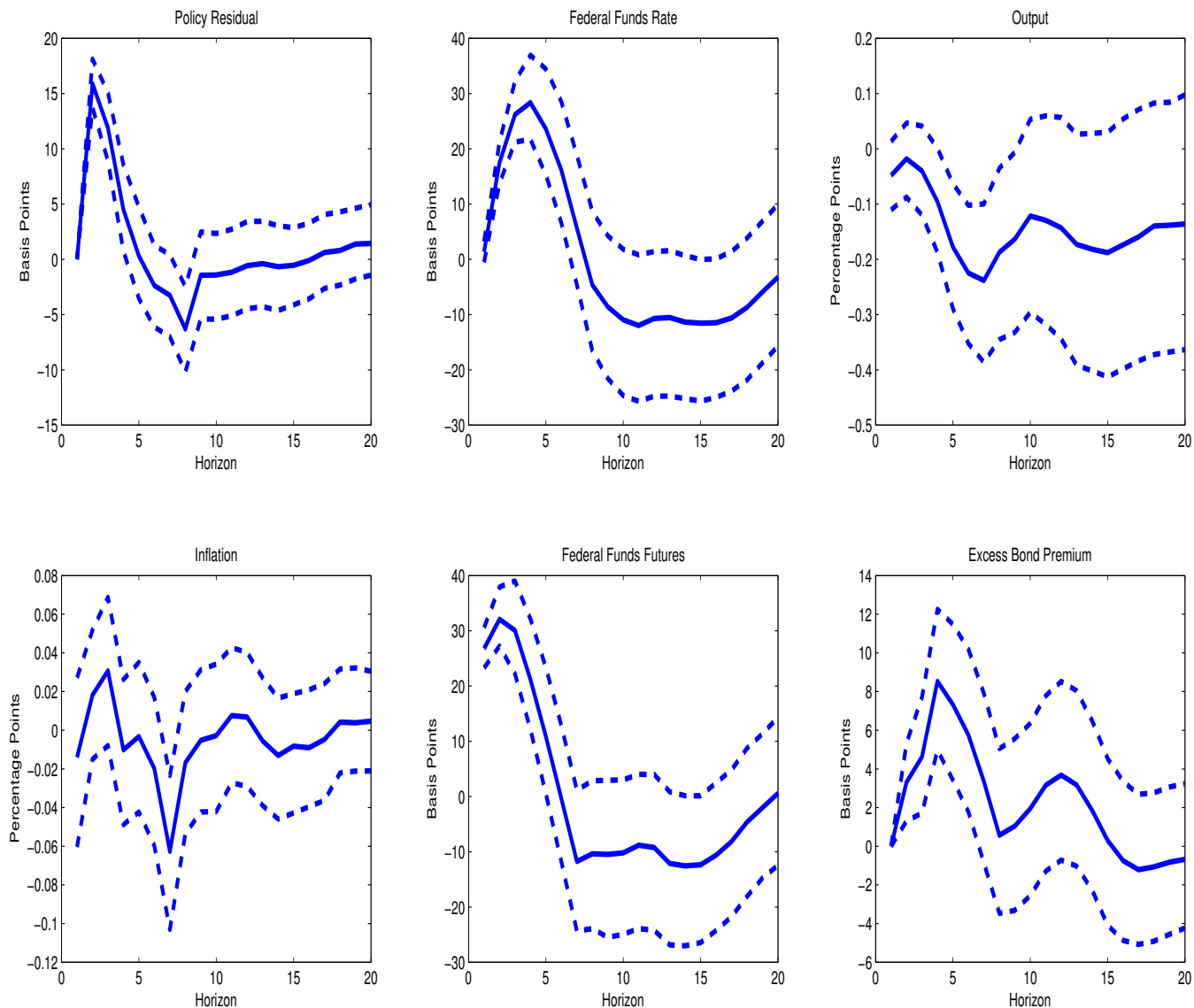
Notes: This figure shows the median and 16th and 84th posterior percentiles of the impulse responses to the monetary news shock from the benchmark VAR.

Figure 4: **Forecast Error Variance Decomposition of Monetary News Shocks: Baseline**.

Notes: This figure shows the median and 16th and 84th posterior percentiles of the contributions of the monetary news shock to the forecast error variance of the variables from the benchmark VAR.

Figure 5: **Monetary News Shock Series from the Baseline VAR**.
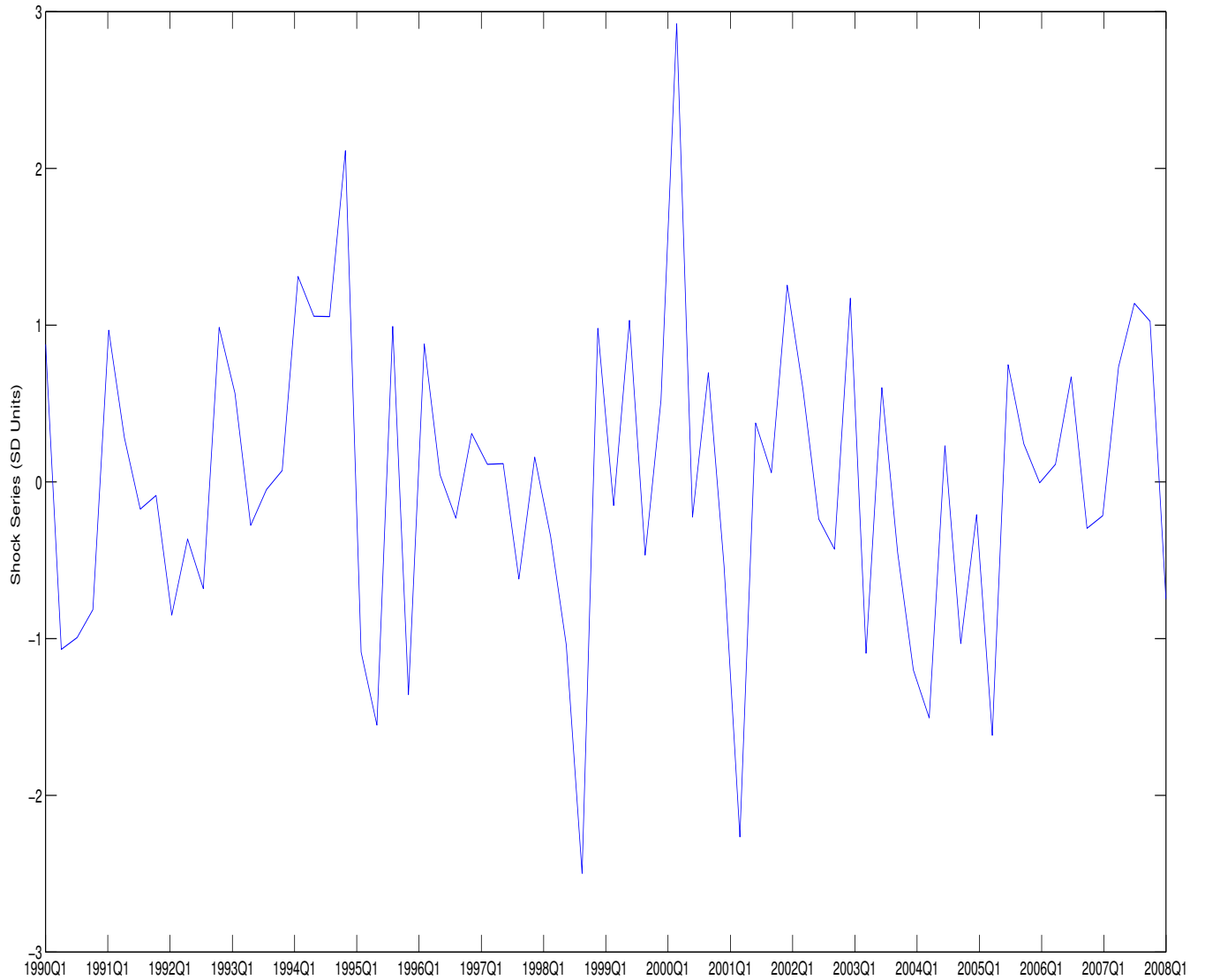
Notes: This figure presents the estimated median shock series from the baseline model. The shock series are in terms of standard deviation units.

Figure 6: **Cross-Correlations between Monetary News Shock and Other Monetary Shocks**.
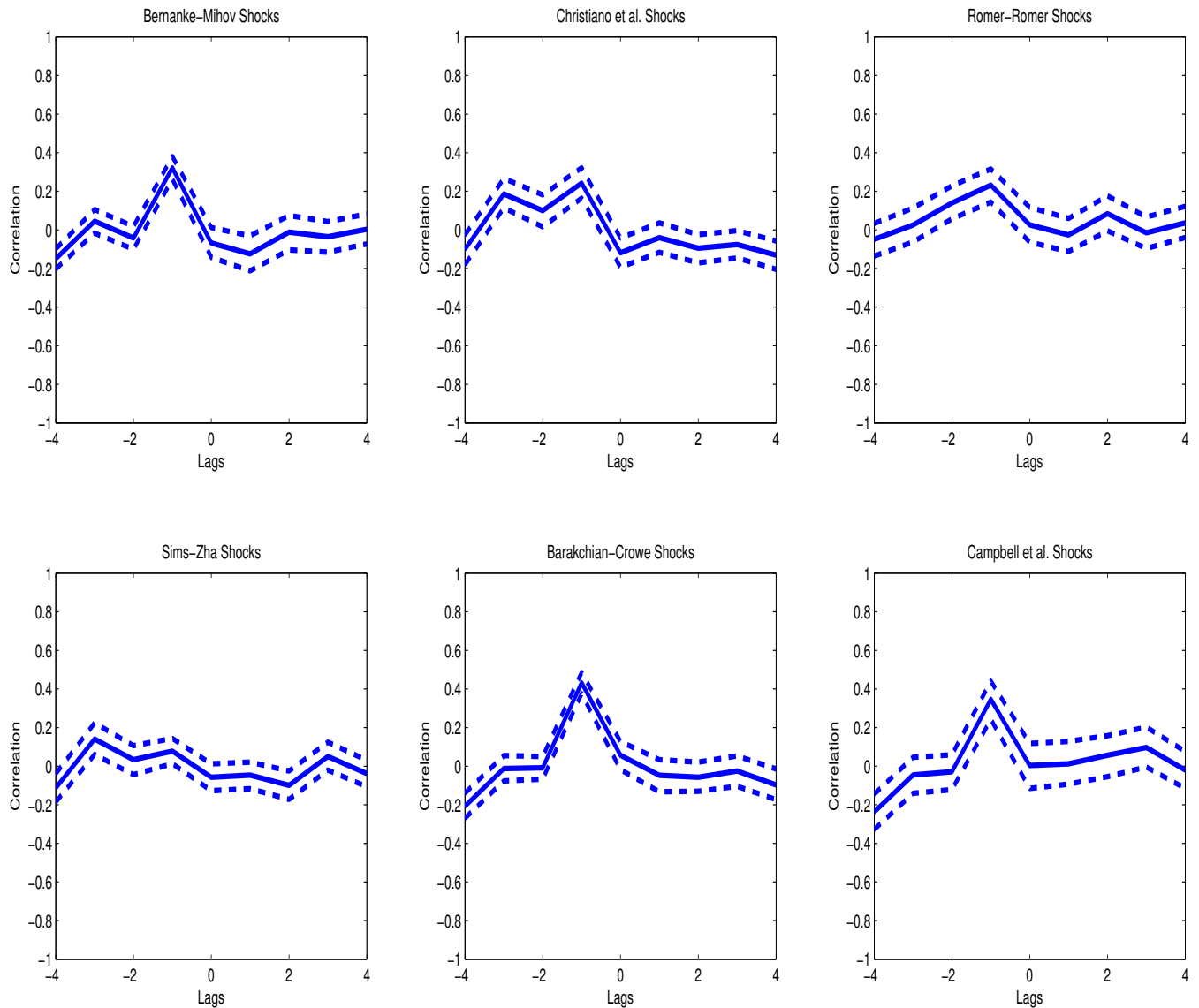
Notes: The solid line is the median cross-correlation and the dashed lines are the 84th and 16th percentiles of the posterior distribution of cross-correlations. The other shocks are the unanticipated monetary shocks identified in Bernanke and Mihov (1998), Christiano et al. (1999), Romer and Romer (2004), Sims and Zha (2006), and Barakchian and Crowe (2013), and the monetary news shocks identified in Campbell, Evans, Fisher and Justiniano (2012). The x-axis gives the lags/leads of the MFEV monetary news shock for which the correlation with the other shocks is computed.

Figure 7: **Robustness to Monetary Policy Rule: (a) Lagged Unemployment Instead of Current Unemployment; (b) Output Gap Instead of Unemployment Gap.**
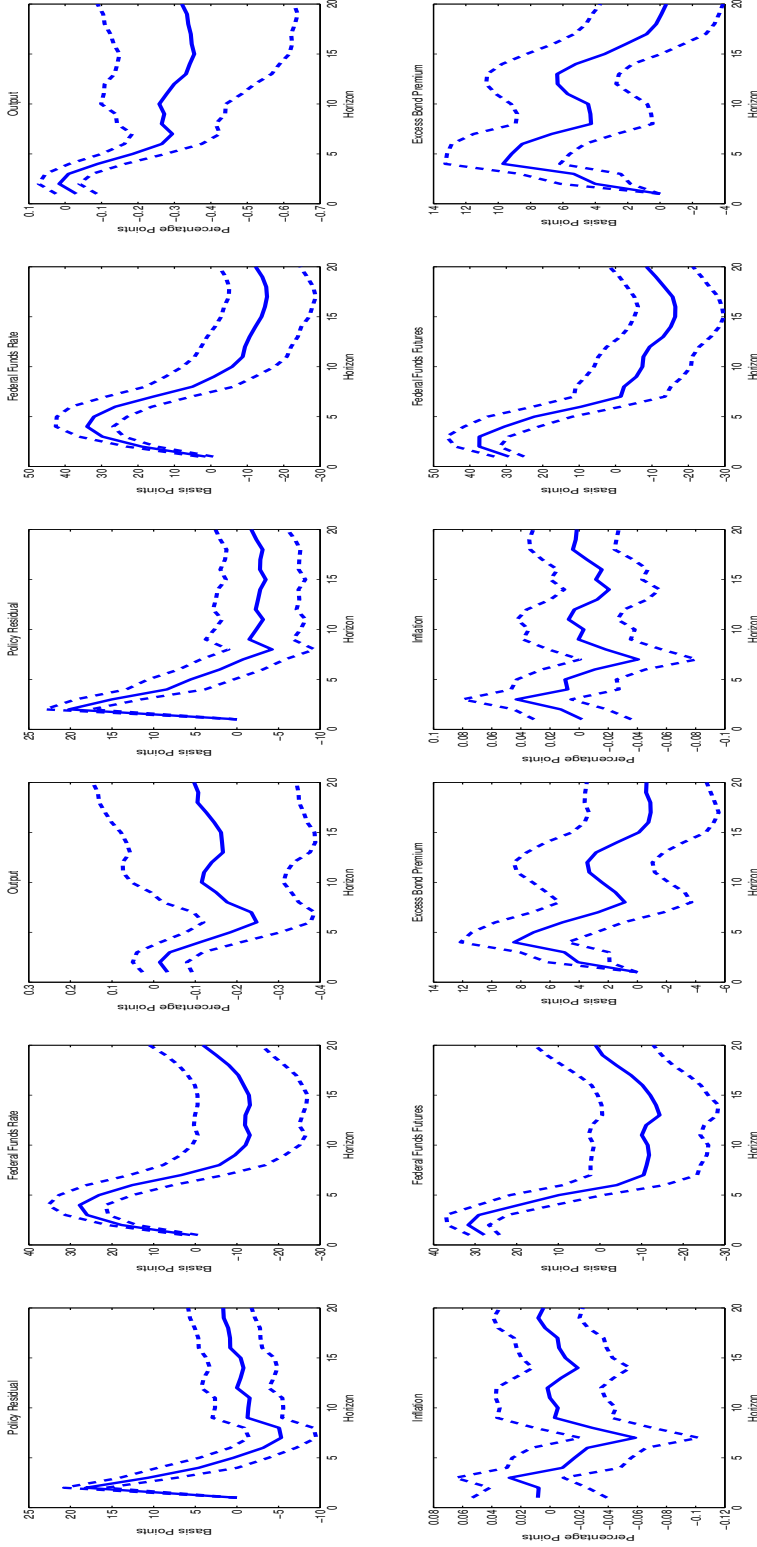
(a) Impulse Responses to a One Standard Deviation Monetary News Shock (H=10).

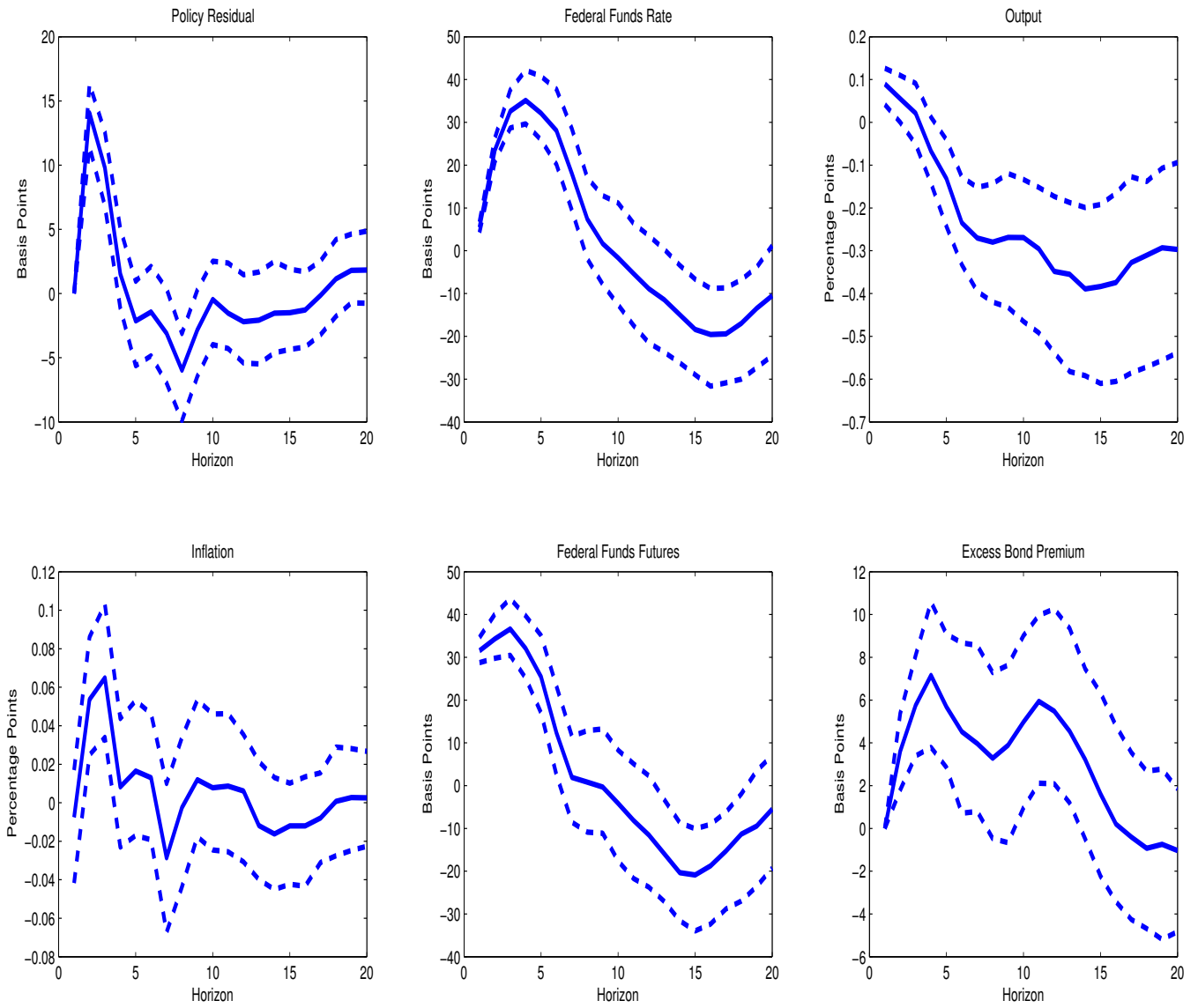**(b) Impulse Responses to a One Standard Deviation Monetary News Shock (H=15).**

Notes: Panel (a): This figure shows the median and 16th and 84th posterior percentiles of the impulse responses to the monetary news shock identified using a monetary policy residual based on a monetary policy rule that includes lagged unemployment instead of current unemployment. Panel (b): This figure shows the median and 16th and 84th posterior percentiles of the impulse responses to the monetary news shock identified using a monetary policy residual based on a monetary policy rule that includes the output gap instead of the unemployment gap.

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Figure 8: **Impulse Responses to a Monetary News Shock: 1988-2012 Sample**.
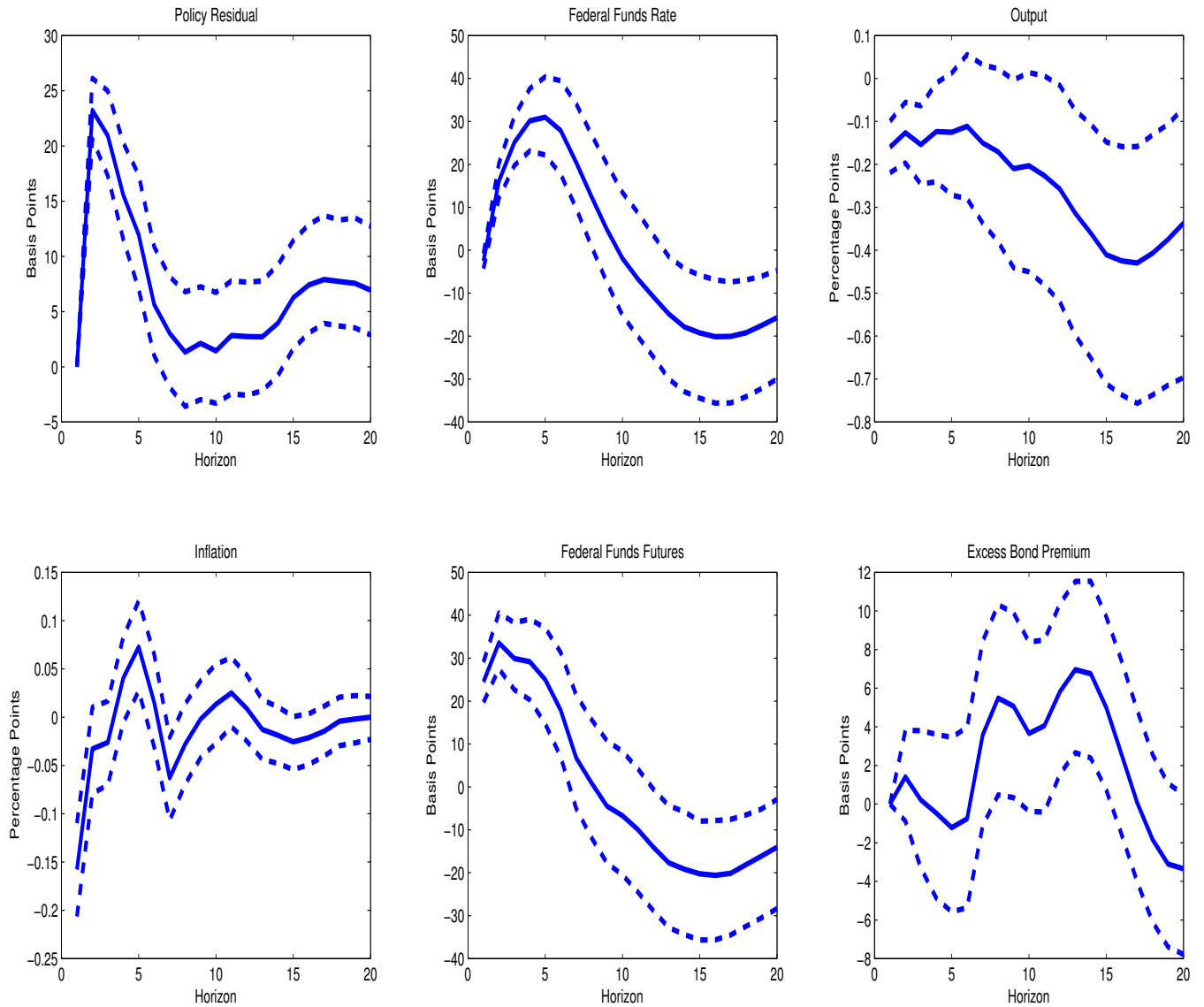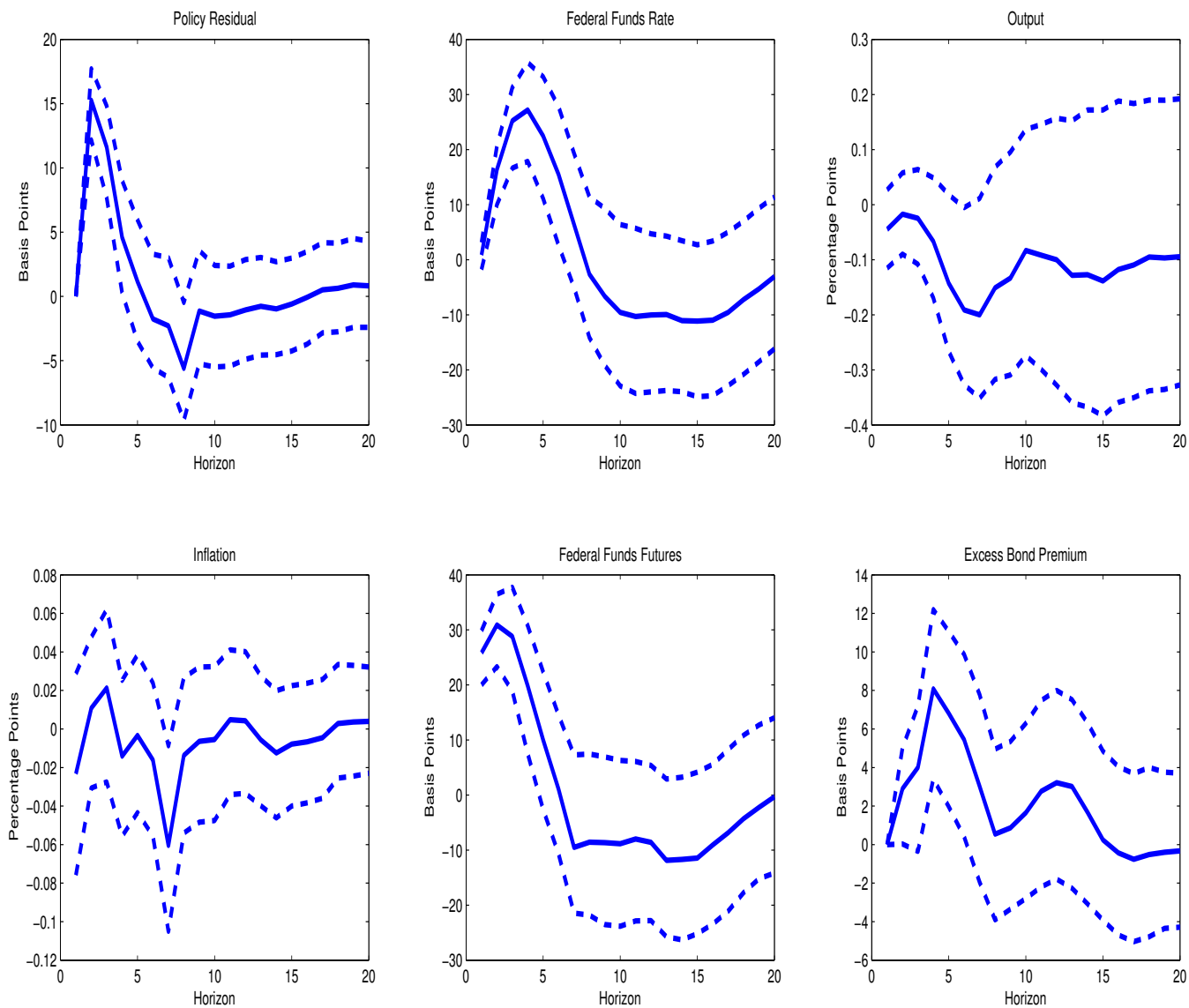

Notes: This figure shows the median and 16th and 84th posterior percentiles of the impulse responses to the monetary news shock obtained from applying our baseline MFEV method on a sub-sample that includes the Great Recession period.

Figure 9: **Impulse Responses to a Monetary News Shock: An Alternative Identification Approach**.

Notes: This figure shows the median and 16th and 84th posterior percentiles of the impulse responses to the monetary news shock obtained as the VAR innovation in the federal funds futures rate orthogonalized with respect to EBP and MPR.

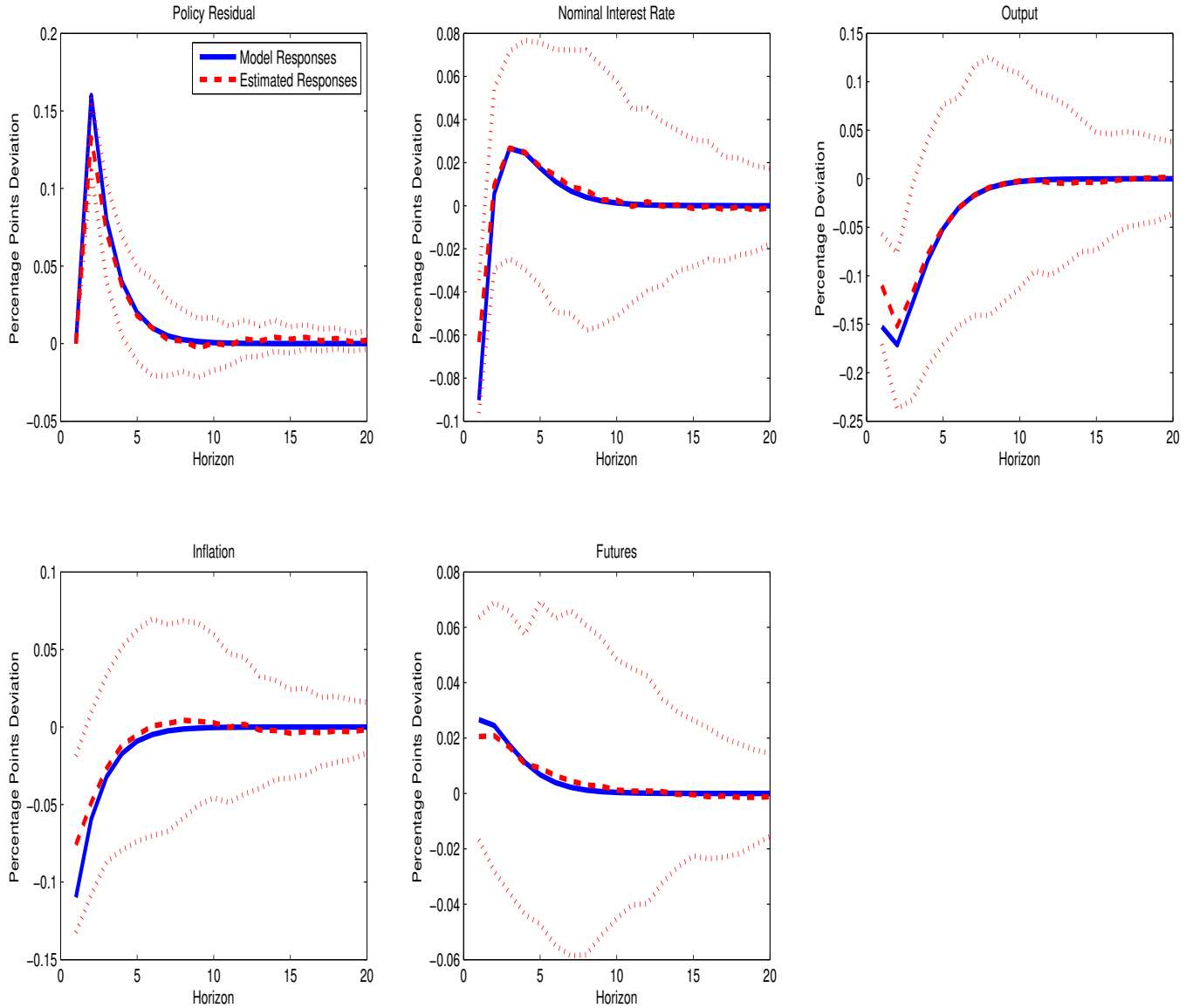Figure 10: **Impulse Responses to a Monetary News Shock: Alternative Truncation Horizon**.

Notes: This figure shows the median and 16th and 84th posterior percentiles of the impulse responses to the monetary news shock identified using a truncation horizon of $H = 9$.

Figure 11: **Cross-Correlations between Monetary News Shock and Other Macroeconomic Shocks**.
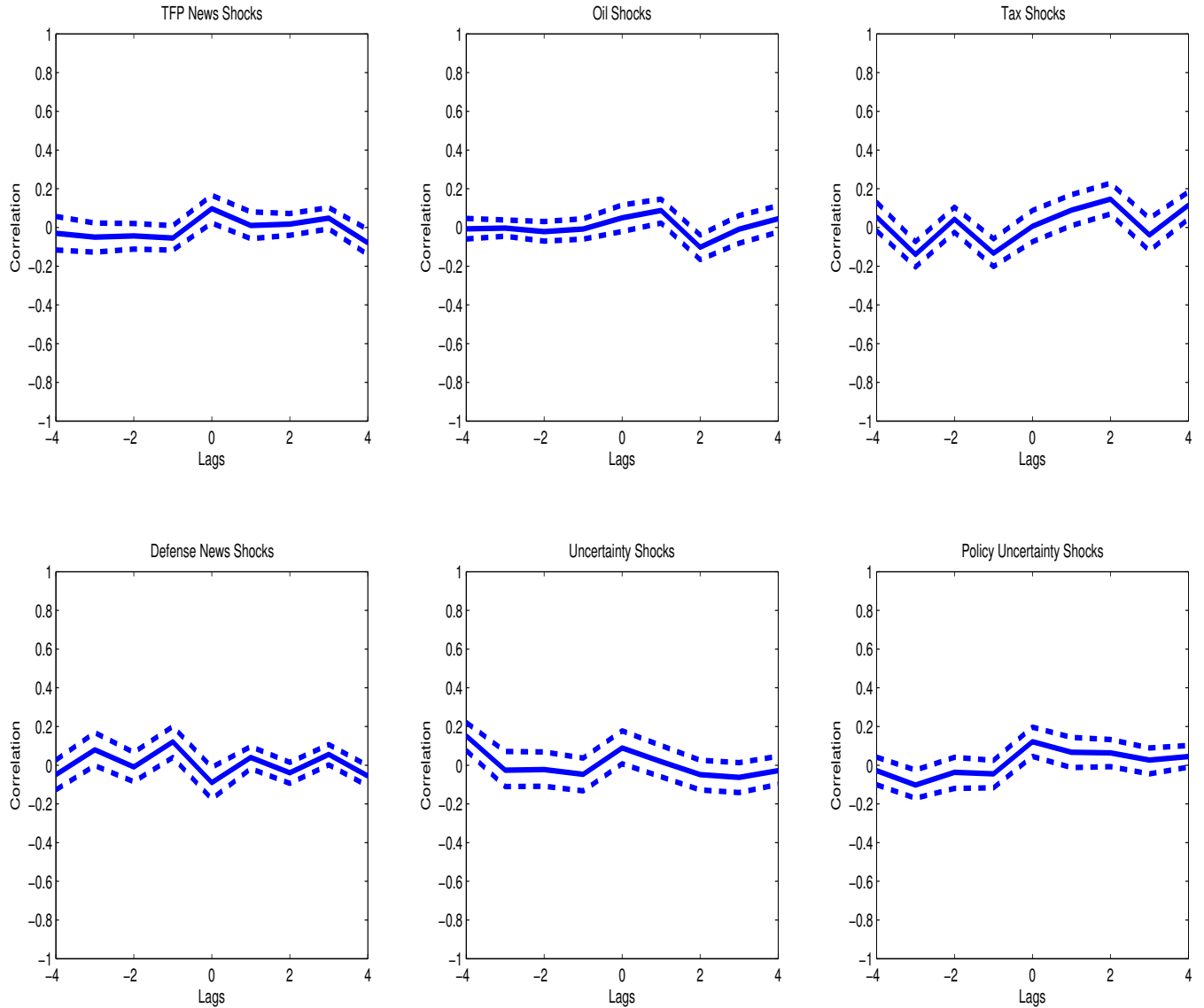
Notes: The solid line is the median cross-correlation and the dashed lines are the 84th and 16th percentiles of the posterior distribution of cross-correlations. The other shocks are the Barsky and Sims (2011) TFP news shock, shock to the real price of oil, Romer and Romer (2010) exogenous tax shock measure, Ramey (2011) defense news shock, the uncertainty shock from Bloom (2009) (that appears in his Figure 1), and the innovation to the U.S. economic policy uncertainty index of Baker et al. (2015). The x-axis gives the lags/leads of the technology news shocks for which the correlation with the MFEV monetary news shock is computed.

Figure 12: **Monte Carlo Evidence: True and Estimated Impulse Responses.**

Notes: The figure is based on 2000 Monte Carlo simulations of the model of Section 5 where in each simulation the impulse responses to the monetary news shock were identified as the median values of impulse responses based on 2000 draws from the posterior distribution of the VAR parameters. The solid line represents the true model impulse responses, the dashed line is the average estimated median impulse response to the defense shock across Monte Carlo replications, and the dotted lines are the 84th and 16th estimated percentiles of the median impulse response.