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A Unified Measure of Fed Monetary Policy Shocks*

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Abstract

Identification of Fed monetary policy shocks is complex, in light of the distinct policymaking regimes before, during, and after the ZLB period of December 2008 to December 2015. We develop a heteroscedasticity-based partial least squares approach, combined with Fama-MacBeth style cross-section regressions, to identify a US monetary policy shock series that usefully bridges periods of conventional and unconventional policymaking and is effectively devoid of the central bank information effect. Our series has moderately high correlation with the shocks identified by Nakamura and Steinsson (2018), Swanson (2018), and Jarocinski and Karadi (2018), but has crucially important differences. Following both the Nakamura-Steinsson and Jarocinski-Karadi empirical tests, we find scant evidence of the information effect in our measure. We attribute the source of these different findings to our econometric procedure and our use of the full maturity spectrum of interest rate instruments in constructing our measure. We then present evidence confirming an hypothesis in the literature that the information effect can lead to the result that shocks to monetary policy have transmission effects with signs that differ from traditional theory. We find that shocks to series that are devoid of (embody) the information effect display conventionally-signed (perverse) impulse responses of output and inflation. This provides evidence of first-order importance to staff at central banks undertaking quantitative theoretical modeling of the effects of monetary policy.

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1 Introduction

The adoption of unconventional monetary policy tools by the Federal Reserve in the wake of the Great Financial Crisis brought policymaking into new territory and rekindled challenges for research measuring monetary policy shocks and estimating their effects. Much of the new research built on influential work that pre-dated the crisis and used bond market data at daily or intra-daily frequencies (Kuttner (2001), Cochrane and Piazzesi (2002), Rigobon and Sack (2003), Gurkaynak, Sack, and Swanson (2005)). In much of this new work, monetary policy surprises are measured as the change in interest rate futures prices in narrow windows around FOMC announcements (for examples, see Wright (2012), Gertler and Karadi (2015), Nakamura and Steinsson (2018), Rogers, Scotti and Wright (2018), Swanson (2018), and Jarocinski and Karadi (2018)). This represented a departure from traditional approaches to measurement and identification such as the use of orthogonalized innovations to the Federal Funds rate in recursive VARs (Christiano, Eichenbaum and Evans (1996)) or the narrative approach of Romer and Romer (2004). An advantage of the methods developed in the recent papers is that, under certain assumptions, the resulting shock series captures both conventional policymaking, through shocks to the target Fed Funds rate, as well as unconventional policymaking, as reflected in identified shocks to forward guidance (FG) and large-scale asset purchases (LSAPs). The use of narrow time windows around FOMC announcements enhances identification, it is argued, because no other economic news is (routinely) released.

The literature on the central bank private information effect has called into question this assertion, however (Romer and Romer (2000), Campbell et al. (2012 and 2016), Nakamura and Steinsson (2018), Miranda-Agrippino (2016), and Jarocinski and Karadi (2018)). Under this view, the central bank reveals in its meeting day announcements not only pure monetary policy news but also its private information on the state of the economy, its own preferences, or the model it uses to analyze the economy. This in turn causes the private sector to change its outlook for macroeconomic developments. Thus, conventionally-measured monetary policy surprises may be correlated with developments in non-monetary policy economic fundamentals, even in tight windows around central bank announcements. Further confounding identification, these studies document a tendency for private sector expectations to go in the wrong direction. That is, following a contractionary monetary policy surprise, expectations of future GDP growth rise. The empirical presence of the Fed information effect calls into question the central assumption that these surprises are appropriate to identify (pure) monetary policy shocks.

The focus of most of these papers, especially the early ones, is on the transmission to financial markets and expectations. For example, Nakamura and Steinsson (2018) document the effects of their policy news shock on the real interest rate, expected inflation, and expected output growth. Swanson (2018) finds that both forward guidance and LSAP shocks have highly statistically significant effects on a wide variety of assets: Treasuries, corporate bonds, stocks, exchange rates, and options-implied interest rate uncertainty. He also examines the persistence of these shocks, compares magnitudes before and during the ZLB period, and concludes with an appeal to examine the transmission to macroeconomic variables.[1]

In this paper, we develop a heteroscedasticity-based, partial least squares (PLS) approach to identify shocks to US monetary policy, compare our measure to those in the literature, and estimate the macroeconomic transmission effects of shocks.[2] The general idea behind construction of our measure is to use Fama and MacBeth (1973) two-step regressions to estimate the unobservable monetary policy shock. This works initially through the sensitivity of "outcome variables" to FOMC announcements. Specifically, in the first step we run time-series regressions to estimate the sensitivity of interest rates at different maturities to FOMC announcements. This is equivalent to the asset beta in the original Fama-MacBeth method. In order to filter out "background noise", we employ the heteroskedasticity-based estimator of Rigobon and Sack (2003, 2004), implemented with instrumental variables (IV), into this step. In the second step, we regress all outcome variables onto the corresponding estimated sensitivity index from step one, for each time t. In this way, we derive the new monetary policy shock as the series of estimated coefficients from the Fama-MacBeth style second step regressions. The application of this procedure to estimating monetary policy shocks is novel as far as we are aware,[3] and has non-trivial effects on the resulting measure.

Our approach to estimating a monetary policy shock series has a couple of conceptual advantages. One is simplicity. Our procedure has very mild data requirements and is easy to implement econometrically. Compared to the path-breaking work of Romer and Romer (2004), implementing our method involves no need to parse through Federal Reserve transcripts and forecasts. Nor does it require the use of intra-daily data, which is costly to acquire and can have spotty coverage, as in much of the newest research. Thus, a second and related advantage of our method is its greater

[1] "Going forward, there are many important issues that call for further exploration. First and foremost, estimating the effects of forward guidance and LSAPs on macroeconomic variables such as the unemployment rate should be a top priority for future research. After all, the FOMCs stated goal in pursuing these unconventional policies was to boost the economy (pg. 37)."

[2] Wu and Xia (2016) and Jarocinski and Karadi (2018) also focus on transmission to macroeconomic variables, as discussed below.

[3] See Wold (1966, 1975) and Kelly and Pruitt (2013, 2015) for applications to equity returns.

applicability. Our approach can be implemented over longer sample periods and for more countries, for which data requirements often render the process untenable.

We also demonstrate the importance of our series in practice. To begin, we show that our shock series has moderately high correlation with the Nakamura and Steinsson (2018), Swanson (2018), and Jarocinski and Karadi (2108) monetary policy shocks. Focusing on the period surrounding lift-off in December 2015, we show that our shock series reflects the strong forward guidance delivered at the October 2015 FOMC meeting, and thus implies that a contractionary monetary policy shock took place in the meeting *before* the actual interest rate hike, consistent with existing measures. In addition, we show that both the short end and long end of the yield curve respond less to our shock than do medium-horizon maturities like 2-year and 5-year rates, similar to the Swanson forward guidance shock. Moreover, there are many days in which the stock market co-moves positively with our series, consistent with the Jarocinski-Karadi observations that are the focus of their paper.

Similarities with existing measures notwithstanding, we show that there are important differences, beginning with evidence on the Fed information effect. Our investigation includes both testing for the presence of the information effect in the monetary policy shock series and estimating impulse responses from shocks that are purged of the estimated information effect. We follow two prominent approaches in the literature: the Nakamura-Steinsson (2018) expectations-based test and Jarocinski-Karadi (2018) "indirect" test. Using the Nakamura-Steinsson test, we do not find a statistically significant information effect in our new shock series, while we confirm its presence in the series estimated by Nakamura and Steinsson (2018) and Swanson (2018). Second, we test for the information effect "indirectly", following Jarocinski and Karadi (2018), and also find scant evidence of the information effect. These authors examine the high-frequency co-movement of interest rates and stock prices around FOMC announcements. Monetary policy announcements that lead to positive co-movement (within the day) are defined to be those that reveal central bank private information. However, even confining our analysis to observations of our new measure that occur on days with positive co-movement between stock prices and interest rates, we find no evidence of an information effect in the sense of Nakamura and Steinsson (2018).

We reconcile the different findings between our monetary policy shock series and existing measures by pointing to important differences in the construction of the measures. A simple "encompassing" analysis shows that differences in the econometric approach and data used to identify the monetary policy shock series both play a key role. Whereas Nakamura and Steinsson construct their shock series from short-term interest rates up to 2 years, we use the whole yield curve. Inclusion of longer term interest rates is very important, because we find that longer term interest rates display less evidence of an information effect. Compared to NS, our PLS approach extracts a common component from the whole yield curve, and assigns more weight to interest rates that have greater correlation with the policy indicator (the five-year treasury rate in the benchmark case). Because the Fed information effect is essentially non-existent in maturities of five years and longer, the common factor we extract also contains less of an information effect. Jarocinski and Karadi construct their proxy for monetary policy surprises also using only a short rate, the three-month Fed Funds futures rate (FF3). Using their data, we find evidence of the Fed information effect, in the sense of Nakamura-Steinsson, on (JK) information effect days, but as noted above, we do not find it in our measure, even on days of positive co-movement between stock prices and our series.

Finally, we present evidence confirming an hypothesis in the literature that the information effect can give rise to monetary policy shocks having transmission effects with opposite signs from those predicted by traditional theory. Using our series, we find that a positive monetary policy shock leads to significantly negative effects on output and prices, consistent with standard theory. This is true in the full sample and for sub-samples before and during the ZLB. We also find conventional signs using only those of our shocks that occur on Jarocinski-Karadi (JK) information effect days or, equally, only those that occur on non-information effect days. On the other hand, shocks to the alternative measures that embody the information effect produce non-traditional signs. This is especially evident during the ZLB period where output *rises* in response to a positive NS or Swanson monetary policy shock. Similarly, when we use the Jarocinski-Karadi proxy for the monetary policy shock, we replicate their finding that an announcement-day interest rate increase accompanied by a stock price increase leads to significantly higher output and price level, and improvement in financial conditions. However, when we replace their measure with the BRW monetary policy shock we find only minor differences in the impulse responses on information effect days and non-information effect days: with our measure, the responses are always of the conventional sign.

The information effect is an issue of first-order concern to staff at the Federal Reserve and other central banks. Should staff models be constructed to feature the information effect associated with monetary policy announcements? If so, how, what are the appropriate building blocks? Should the impulse responses that the staff's quantitative models attempt to match be of the signs predicted by traditional monetary theory, or of the unconventional signs consistent with the evidence in influential recent papers on the information effect?

In the next section, we describe our econometric approach and the data. In section 2, we display our new series and compare it to existing measures in the literature. In section 3, we test for the presence of the information effect in our Fed monetary policy shock series and alternatives, and reconcile the different findings. In section 4, we confirm the hypothesis in the literature that the information effect can give rise to impulse responses that have signs opposite to those predicted by conventional theory. Section 5 concludes: we provide a US monetary policy shock series that is easy to estimate, that usefully bridges periods of conventional and unconventional policymaking periods, is devoid of the information effect, and that helps substantiate an hypothesis concerning transmission effects to output and inflation.

2 A New Monetary Policy Shock

2.1 Methodology: Fama-MacBeth Meets Rigobon-Sack

We assume that the true monetary policy shock e_t is unobservable. We further assume that the (observable) changes in Treasury yields around FOMC announcement days are driven by a monetary policy shock e_t and nonmonetary policy shock η_t. Our objective is to estimate the former. We normalize the unobserved monetary policy shock to have a one to one relationship with the 5 year Treasury yield,[4]

$$\Delta R_{5,t} = \alpha_0 + e_t + \eta_t \tag{1}$$

where $\Delta R_{5,t}$ is the change in our policy indicator – the 5 year Treasury yield, α_0 is a constant, e_t is the monetary policy shock, and η_t denotes factors unrelated to monetary policy news.[5] We allow η_t to include idiosyncratic noise specific to the 5 year interest rate as well as noise that is common to the entire yield curve.

Our Fama-MacBeth two-step procedure extracts monetary policy shocks e_t from the common component of the outcome variables $\Delta R_{i,t}$. In the first step, we estimate the sensitivity of each outcome variable to monetary policy via time-series regressions. We assume that the outcome of monetary policy decisions is reflected in the movements of zero-coupon yields with maturities of 1

[4] This is motivated by the notion that Fed policy aims to affect interest rates at about this horizon, an assumption that became more appropriate during the ZLB period and is used elsewhere in the literature. We examine (and confirm) robustness of this choice of monetary policy indicator to 2-year and 10-year rates.

[5] This includes factors associated with the Fed information effect, e.g., the market interpreting an FOMC policy announcement as (also) revealing private information it has on the state of the economy, its own preferences for inflation versus output stabilization, etc. The fact that Federal Reserve Board staff construct the index of Industrial Production is one potential source of such private information. Fed staff are situated particularly auspiciously, for example, to ascertain and report to the FOMC in private how noisy is a particular release of the IP series. See Nakamura and Steinsson (2018) for further discussion of "background noise".

year to 30 years. As we demonstrate in section 3, our use of the full maturity structure is important, most notably in producing a shock series that is devoid of the information effect. These outcome variables are also affected by background noise:

$$\Delta R_{i,t} = \alpha_i + \beta_i e_t + \epsilon_{i,t} \tag{2}$$

where $\Delta R_{i,t}$ is the change in the zero-coupon yield with i-year maturity and $\epsilon_{i,t}$ is the idiosyncratic noise for $\Delta R_{i,t}$. We assume the error term $\epsilon_{i,t}$ and the unobserved monetary shock e_t are uncorrelated. Due to our normalization, we can rewrite (2) as,

$$\Delta R_{i,t} = \theta_i + \beta_i \Delta R_{5,t} + \xi_{i,t} \tag{3}$$

where $\xi_{i,t} = -\beta_i \eta_t + \epsilon_{i,t}$ and θ_i is a constant. Recalling that η_t is the error term in the policy indicator ($\Delta R_{5,t}$, see equation (1)), we see that the regressor $\Delta R_{5,t}$ is correlated with the error term $\xi_{i,t}$ due to the component "$-\beta_i \eta_t$". The OLS estimate of β_i is thus biased.

Therefore, although this step could be done using OLS and high-frequency FOMC announcement day data, we instead use the heteroskedasticity-based estimator of Rigobon (2003) and Rigobon and Sack (2003, 2004). This allows us to better filter out background noise. As demonstrated formally in Appendix A1, β_i in (2) can be consistently estimated using instrumental variables (IV). Rewrite (2) as:

$$[\Delta R_{i,t}] = \alpha_i + \beta_i [\Delta R_{5,t}] + \mu_{i,t} \quad i = 1, 2, \cdots, 30 \tag{4}$$

where the independent variable $[\Delta R_{5,t}] = (\Delta R_{5,t}, \Delta R^*_{5,t})'$, $\Delta R_{5,t}$ is the 1-day movement in the policy indicator around the FOMC announcement, and $\Delta R^*_{5,t}$ is the movement with the same event window length but one week before FOMC announcement day.[6] The event window for $[\Delta R_{i,t}]$ corresponds to that of $[\Delta R_{5,t}]$, β_i measures the sensitivity of $\Delta R_{i,t}$ to monetary policy shocks, and $\mu_{i,t}$ is the idiosyncratic noise associated with i only. We exploit the fact that β_i can be estimated using an instrumental variable $\Delta R^{IV}_t = (\Delta R_{5,t}, -\Delta R^*_{5,t})'$ for the independent variable. The underlying assumption is that, on days of FOMC meetings, the variance of the true monetary policy shock increases while that of the background noise remains unchanged. In the estimation, we use a 1-day window, capturing policy surprises between FOMC announcement day (end) and the previous day (end). Because the Fed released no public statements about monetary policy decisions until 1994,

[6] The choice of one week follows Nakamura and Steinsson (2018). We examine (and confirm) robustness to two days before the FOMC announcement day, which is akin to the Rigobon (2003) practice.

we begin estimation of our shock series then.

The second step of our approach, by analogy to Fama and MacBeth, is to recover the aligned monetary policy shock from cross-sectional regressions of $\Delta R_{i,t}$ on the estimated sensitivity index $\hat{\beta}_i$ for each time t,

$$\Delta R_{i,t} = \alpha_i + e_t^{aligned}\hat{\beta}_i + v_{i,t} \quad t = 1, 2, \cdots, T \tag{5}$$

where $e_t^{aligned}$ is the coefficient of interest. This series of T estimated coefficients from the second step regressions is the BRW monetary policy shock series.

2.2 The Data

We collect data on the monetary policy indicator from the Federal Reserve Board public website. As noted above, we examine 2-year, 5-year, and 10-year Treasury rates, with 5-year as benchmark. We also use data on estimated term premia, from Adrian, Crump, and Moench (2013), which are available through the New York Fed website https://www.newyorkfed.org /research/data_indicators/term_premia.html. The policy outcome variables, the zero coupon yields with maturities of 1 to 30 years, are estimated by Gurkaynak, Sack, and Swanson (2005), and available at https://www.federalreserve.gov/pubs/ feds/2006/200628/200628abs.html. To estimate impulse responses, we use monthly industrial production and CPI, both taken from https://fred.stlouisfed.org, the core commodity price index from Thompson Reuters, and the excess bond premium from Gilchrist and Zakrajsek (2012).

2.3 BRW Monetary Policy Shock Series

We display our monetary policy shock series in Figure 1. There are sizable movements before, during, and after the ZLB period. The announcements of QE1, QE2, and QE3, which are marked by navy lines, all generate large expansionary monetary policy shocks. Monetary policy shocks during Operation Twist, denoted by the orange lines, are instead contractionary. We mark with the blue line the FOMC meeting in October 2015, the meeting *preceding* lift-off in December. Zooming in on the last three meetings of 2015, our shock series takes the values -0.080 (September), 0.115 (October), and 0.038 (December). Expectations of a lift-off had been growing throughout the summer and heading into the October meeting. For a variety of reasons, including turmoil in global equity markets, the FOMC decided to keep the target Fed Funds rate unchanged at that meeting but sent a clear signal of a likely rise in December 2015.[7] Our measure indicates that this forward guidance gave rise to a sizable contractionary monetary policy shock in October 2015, one meeting before the actual rate increase. This is consonant with the dynamic pattern of alternative measures that use intra-daily data and estimate separate components of Fed monetary policy shocks. For example, the corresponding values of the policy news shock of Nakamura and Steinsson (2018) are (-0.042, 0.032, 0.016), the forward guidance surprise in Rogers, Scotti, and Wright (2018) are (-0.09, 0.09, 0.03), and in Swanson (2018) (-1.50, 1.67, NA).[8] We analyze this further in the next section.

2.3.1 Comparison with Shocks in the Literature

Moving beyond the issue of plausibility of specific observations around liftoff and QE announcements, we provide in Table 1 a comprehensive comparison of our shock series with well-known measures in the literature: Kuttner (2001), Romer and Romer (2004), Nakamura and Steinsson (2018), Swanson (2018), and Jarocinski and Karadi (2018). The updated R&R shock series, constructed using their same narrative method, runs through the end of 2007. Kuttner (2001) shocks are extracted from changes in Federal Funds futures rates in 30-minute windows around FOMC announcements. Nakamura and Steinsson also examine high-frequency movements around FOMC announcements. Their monetary policy shock is the first principal component of changes in the current month Federal Funds futures rate, the Federal Funds futures rate immediately following the next FOMC meeting, and two, three and four quarter ahead euro dollar futures in the 30-minute event window.[9] Jarocinski and Karadi (2018) use three-month Fed Funds futures (FF3) changes in 30-minute windows around FOMC announcements, while Swanson (2018) separately identifies the effects of forward guidance, large-scale asset purchases, and target Federal Funds rate shocks, also using principal components.[10]

In Table 1 we present the correlation between our measure and the alternatives (figures are available in the online Appendix). As seen in column 1, over the full sample, our shock is reasonably well correlated (around 0.5) with the NS and Swanson shocks, which themselves are relatively large before and during the ZLB. The next two columns decompose the comparison into sub-periods, before and during the ZLB. Before the ZLB, our series is correlated with NS, JK, and the Swanson

[7] As headlined in the Financial Times on October 29, 2015: "Federal Reserve drops warnings on global risks to US economy: Central bank hawkish statement increases chances of December rise in interest rates."

[8] Magnitudes differ due to different normalization choices, especially by Swanson, whose series ends with liftoff.

[9] We obtain these shocks from Nakamura and Steinsson (2018) through 2014m3 (their sample period) and then follow their procedures to update to the present. For this exercise and all of our work using intra-daily data, we obtain the data from the "Event Study" database maintained by Federal Reserve Board staff.

[10] Rogers, Scotti, and Wright (2018) implement an approach similar to Swanson (2018) in computing their three separate components of Fed policy shocks. The series are very highly correlated with those of Swanson, around 0.96.

FG shock at around 0.6. In the final column, we present correlations during the ZLB. The largest correlation, at 0.57, is with the Swanson FG shock. In Figure A7, we display plots of our shock series against the alternatives. Consistent with the correlations above, prior to 2008 our shock series exhibits a similar pattern to the NS, Kuttner, and R&R shocks. After 2008, the alternative series are quite small given that the Fed Funds rate is at zero during the ZLB. In contrast, our new shock series exhibits relatively large movements, consistent with Fed monetary policy being about more than the target FFR. Our shock series is more similar to the FG and LSAP shocks of Swanson.

2.3.2 BRW Series Construction Robustness

We examine several modifications to the construction of the baseline BRW shock series. As previewed above, we consider alternative normalizations of the monetary policy shock series to either the 2-year or the 10-year Treasury rate instead of the 5-year. As seen in Columns 1 and 2 of Table 2, the correlation with our baseline shock series is above 0.97. Thus our approach is robust to different choices of the monetary policy indicator. Our second check is to extend our monetary policy shock series backward to 1969. Before 1994, there was no public announcement of FOMC decisions. Thus, for this earlier period, we use the 1-day policy window between the FOMC announcement day and the following day to capture the policy effect. From the third column of Table 2, we see that the correlation with our BRW shock is over 98%.[11]

Our third modification is to use only zero-coupon yields with 1-, 2-, 5-, 10-, and 30-year maturities, the more commonly-used series, as the outcome variables. The correlation with the baseline shock series, as shown in column 4 of Table 2, is over 0.95. Fourth, we assess robustness to leaving out the QE1 announcement in the alignment process. This announcement, in March 2009, was a sufficiently big event occurring at a time when financial markets were so sluggish that the market response might not represent a typical effect of monetary policy. The new shock series without QE1 is again highly correlated with our baseline series (Column 5). Next, we extend our sample to include all unscheduled FOMC meeting dates since 1995, reconstruct our shock, and find a correlation of 0.9 (Column 6). We then consider using a 2-day event window for both policy indicator and outcome variables. Doing this, we find that the correlation with the baseline shock series is 0.84 (Column 7). We also construct the instrumental variable as the daily movement in the

[11] One feature of our methodology is the need to check the stability of the sensitivities of interest rates with different maturities to monetary policy shocks. Here, we do the rolling sample test for each period of 15 years, expanding the sample size to 1969 - 2017. When we use different monetary policy indicators of 1-, 2-, 5- and 10-year Treasury Rates, the coefficients are not completely stable until early 1990 (figure available in the online Appendix). That is why we start the sample in 1994, when the Fed first released a statement about FOMC policy decisions. The sensitivity index is flat after 1994, indicating stability of our alignment process.

policy indicator *one day* (as opposed to one week) before FOMC announcement day. As presented in Column 8 of Table 2, this alternative shock series has a correlation of 0.99 with the baseline series.

Real-Time BRW Shock. As a final consideration, we construct real-time versions of our series.[12] We use two methods: (1) estimate the first step on the sample up to 2007:12, use the betas from that in the second step regression to compute the aligned monetary policy shock for 2008:1, then roll through the sample one month at a time to construct a real-time shock series after 2008 using these rolling window sensitivity indexes; and (2) estimate the first step regression *only* up through 2007:12 and use the estimated betas from that regression to generate the aligned monetary policy shock series for each observation beginning in 2008:1. The correlations of these two real-time measures with the baseline, "ex-post" BRW shock series are 0.95 and 0.88, respectively (see columns 9 and 10 of Table 2 and the on-line Appendix figures).

2.3.3 Monetary Policy Shocks and the Slope of the Yield Curve

Table 3 provides further evidence on the nature of our shock, with estimates of the effect of it on interest rate spreads at different maturities. Comparisons above suggest that our shock is closely related to forward guidance, which is well captured by movements in 2- or 5-year interest rates. Here we'll consider the 5-year interest rate as benchmark and regress interest rate spreads of different maturities over the 5-year rate on different monetary policy shocks,

$$\Delta SPREAD_{i,t} = \alpha_i + \beta_i e_t + \epsilon_{i,t} \tag{6}$$

where $SPREAD_{i,t}$ is the difference between interest rate with maturity i and the 5-year rate around the FOMC announcement and e_t is, alternatively, the BRW, NS, Swanson, and JK monetary policy shock series.

Column (1) of Table 3 shows the regression results of the 5-year rate itself on the BRW and alternative shock series. The coefficient on BRW is 0.679 and highly statistically significant. The response of the 2-year/5-year interest rate spread -0.113 (Column 4) is significantly negative but close to zero. Thus, the 2-year interest rate responds to our shock in a similar way as does the 5-year rate. Coefficients in regressions for all of the other spreads (6 month and 1 year (Column 2 and 3), 10, and 30 year rates (Column 5 and 6)) are negative and significant, suggesting that

[12] One advantage of using raw surprises as in Kuttner (2001) and JK (2018) is that the resulting shocks are precisely what occurred in real time. Series such as NS (2018), Swanson (2018), and our baseline measure above are (full-sample) estimation-based, do not account for estimation error, and are thus not strictly speaking real-time.

both the short and long end of the yield curve respond to our shock by less than does the 5-year interest rate. Finally, we run same regressions for the NS, JK, and Swanson shock series, as seen in the remaining rows of the table. Our BRW series is similar to Swanson's forward guidance shock series in the sense that both move the 2-year and 5-year interest the most. The NS shock series and Swanson's LSAP shock series capture the movements of the yield curve at the short end and long end, respectively. The JK shock (FF3) affects spreads significantly differently on information effect and non-information effect days, arguably as expected. As seen in the final two rows, FF3 shocks on non-information effect days affect spreads in much the same way as NS shocks, while on information effect days the shock is strongest at the very short end of the yield curve, with zero effect on the 5-year rate itself or the 2-year rate.

3 The Fed Information Effect

Romer and Romer (2000), Nakamura and Steinsson (2018), and Jarocinski and Karadi (2018), among others, advance the hypothesis of a "Fed information effect": monetary policy announcements contain information about central bank forecasts of economic fundamentals. As a by-product, macroeconomic variables such as output and inflation may be influenced not only by the announced policy itself but also by the forecasting information contained in the announcement. The opposite forces from these two sources (the policy and the reaction to it) may cause puzzling impulse responses such as output rising after a contractionary policy shock. Use of even narrow windows around central bank announcements may not alleviate the issue for researchers.[13] In this section, we subject our series to the same tests for the information effect used by Nakamura-Steinsson and Jarocinski-Karadi. We find scant evidence of the information effect in the BRW measure and pinpoint reasons for why our results are different from others.

3.1 A Direct Test and Implications

We begin with the test of Nakamura and Steinsson (2018). We confirm their results for their series and examine robustness to our shock and Swanson's (2018). Specifically, we run regressions of monthly changes in Blue Chip survey expectations of output growth on the monetary policy shock series of that month, and test for the Fed information effect based on the sign of the estimated

[13] Campbell et. al. (2012) also provide evidence of a Fed information effect. Faust, Swanson, and Wright (2004) and Zhang (2019) find no such evidence, however, while Lunsford (2018) argues that in his sample from February 2000 to May 2006 the information effect is present in the first half only.

coefficient.[14] Table 4 reports the results. While the information effect is significant in the measures of Nakamura-Steinsson and Swanson, it is insignificantly different from zero in ours (see the first three columns). For a robustness check, we also find that the two real-time BRW measures are essentially immune from the information effect (the fourth and fifth columns).

In Figure 2, we depict the difference between Fed and Blue Chip forecasts of real GDP growth, a standard proxy for central bank private information used in the literature.[15] Noteworthy are the large negative values around September 11, 2001 and the last quarter of 2008. At these times, the Fed was significantly more bearish on the economy than the private sector.[16] Table 5 reports OLS regressions of the various monetary policy shock series on these forecast differences. The coefficient is positive and significant for the NS and Swanson measures, but insignificantly different from zero in the regression using our series, a regression with an R2 of only 0.02. Once again, the central bank information effect seems barely present in our new series.

3.2 Evidence from an Indirect Approach

Jarocinski and Karadi (2018) construct their information shock by examining the high-frequency co-movement of interest rate and stock price surprises on FOMC announcement days. They argue that when the stock market moves in the same direction as interest rates, the Fed information effect dominates the monetary policy news effect of the announcement. Following Jarocinski and Karadi, we depict in the scatterplot of Figure 3 daily returns on the S&P 500 on FOMC announcement days against the BRW shock (blue dots) as well as the JK surprises – FOMC announcement day high-frequency changes in the third Fed Funds futures contract (in orange). Although the relationship is negative overall, there are clearly many points falling in the first and third quadrants. As emphasized by Jarocinski and Karadi, these are difficult to explain as purely monetary policy shocks. We re-estimate the NS information effect regressions, Blue Chip GDP growth forecast change on the monetary shock, separately on Fed information effect days and non-information days, for both BRW and JK measures. The results are displayed in columns six and seven (BRW) and eight and

[14] In addition, we find robust results running the tests on the NS sub-samples: 1995-2014, 2000-2014, and 2000-2007 (see the online appendix). Extending through 2018 does not alter our conclusions. Also following NS, we exclude from these regressions all observations when FOMC meetings occurred in the first week of the month, as that likely precedes the time the months Blue Chip survey forecast was made.

[15] The series is constructed as follows: (1) prior to December 2013, the average of the first four quarters ahead Greenbook forecasts minus the corresponding Blue Chip forecasts. (2) After January 2014, for which the Greenbook forecasts are not yet publicly available, we use the forecasts from the Fed summary of economic projections (SEP). These are available four times a year: in March, June, September, and December. For the other four FOMC meetings each year, we use the SEP from the previous meeting. We use the current year SEP forecast if the FOMC meeting happens in the first quarter of the year. Otherwise, we use the projection for the following year. We subtract from this the year-ahead Blue Chip forecast.

[16] These were also times when important news events occurred at a higher frequency than the available forecasts.

nine (JK) of Table 4. In regressions with the BRW measure, the point estimates are very small and have no statistical significance. Thus, even during the "Jarocinski-Karadi" information effect days our BRW shock does not display economically or statistically important Fed information effects in the sense of NS. However, the next two columns of Table 4 confirm that the information effect *is* present in the Jarocinski-Karadi data. This naturally sparks the question we address next.

3.3 Why Does Our Shock Series Have Less of a Fed Information Effect?

In order to understand why our monetary policy shock series does not have an information effect in it, we begin by considering the importance of the underlying data and econometric procedure used to construct the series. First, we find that the inclusion of long-term interest rates is important because long-term interest rates are less associated with Fed information effects. Nakamura and Steinsson construct their monetary policy shock from a set of variables that only contains short-term interest rates up to two years. By contrast, we use the whole yield curve to come up with a summary measure of the stance of monetary policy. In Table 6, we report results of the NS information effect regressions–monthly changes in Blue Chip survey expectations of output growth on the 30-minute changes of interest rates–with maturities from 1 day (Fed funds future rate) to 30-year treasury bond yield. This table is similar to Table 4. Columns (2) to (7) correspond to estimated results of changes in the Blue Chip forecasts of GDP on interest rate changes with different maturities. It is clear that as the maturity of interest rates increases, the coefficients become less significant. This indicates that one reason our BRW shock series contains less of a Fed information effect is because we use long term interest rates compared to alternative measures of Fed monetary policy shocks.

Second, we find that the two-step PLS procedure (i.e. Fama-Macbeth) is equally important in reducing the Fed information effect in our shock series. To see this, we input our data into the NS principal components estimation procedure to construct an alternative monetary policy shock series, which we label the "PCA shock". As seen in column 13 of Table 2, the correlation between this shock and our baseline BRW shock is only 0.25. Moreover, estimating the NS information effect regressions with this PCA shock, we find that a positive shock leads to a significant *increase* in the Blue Chip real GDP growth rate forecast in the next quarter, consistent with Fed private information effects being embedded in this alternative series (Table 4, column 12). We conclude that the PCA approach does not remove the Fed information effect even when the underlying data include long-term interest rates.

We conclude our encompassing analysis by inputting high-frequency NS data into our estimation procedure. This includes data in tight windows around FOMC announcements on the expected 3-month eurodollar interest rates with horizons of 2 to 4 quarters, the current month Fed funds futures rate and the Fed funds futures rate immediately following the next FOMC announcement. The "Tight(NS) shock" generated in this way has a correlation of 0.38 with the BRW shock (Table 2, column 14). The information effect regressions of Table 4 indicate that a positive shock to this series is unrelated to changes in the Blue Chip real GDP growth rate forecast (column 11). What happens when we expand the NS data set to include longer horizon maturities? The "Tight shock" is generated with our PLS estimation procedure but with the NS data expanded to further include the expected 3-month eurodollar interest rates with horizons of 1 to 8 quarters and on-the-run Treasury rates of 3 months, 6 months, 2 years, 10 years and 30 years. Using this expanded data increases the correlation with the BRW shock up to 0.50 (Table 2, column 15). Again, the information effect essentially disappears in this Fama-MacBeth aligned shock (Table 4, row 12, Tight(full data)). This confirms the importance of using the Fama-Macbeth procedure in accounting for differences in results concerning the presence of information effects in monetary policy shock series.

The PLS and PCA approaches are similar in the sense of extracting the common component from outcome variables, but the PLS procedure we use assigns weights based on the correlation of outcome variables with the policy indicator (5-year treasury yield).[17] Since the Fed information effect is not present in the 5-year interest rate or interest rates with longer maturities (Table 6), it is to be expected that the common factor we extract also contains less of a Fed information effect. By similar reasoning, because our sample period contains relatively more data after 2007 compared to the existing literature, this will also reduce the degree of Fed information effect in our shock series. We thus conclude that the inclusion of long-term interest rates, a longer post-2007 sample period, and the Fama-MacBeth procedures play important roles in the construction of the BRW shock, and accounts for much of the difference in our findings concerning the information effect.[18]

[17] As pointed out by Kelly and Pruitt (2013, 2015), the PLS forecast asymptotically recovers the latent factor that drives movements in the policy indicator as the number of outcome variables and length of time series both increase.

[18] We also investigated which part of our estimation procedure, IDH or PLS, is more important in isolating the Fed information effect. We constructed an alternative BRW shock series using the Fama-Macbeth two-step procedure *without the use of IDH* but with the same policy indicator and outcome variables as in the baseline. As presented in Table 2 (column labelled OLS), the IDH-free shock is highly correlated with the baseline BRW shock (0.991).

4 Impulse Responses

As noted above, the existing literature has offered the information effect as one reason why the transmission effects of shocks to monetary policy could have signs that differ from those predicted by traditional theory. In this section, we present robust evidence confirming this hypothesis, using the array of monetary policy shock series above to compute impulse responses of output, inflation, and credit conditions. Shocks to series that do not contain the information effect, such as baseline BRW, display conventionally-signed impulse responses while shocks to series that contain the information effect give rise to impulse responses with the opposite signs, especially during the ZLB.

4.1 BRW Shocks

Following Romer and Romer (2004), we place our cumulative shock series in a monthly VAR model to identify the transmission effects of monetary policy shocks. We allow our monetary policy shock to contemporaneously affect all variables: output, inflation, commodity prices and excess bond premium.[19] We include commodity prices in light of the price puzzle (CEE, 1996) and the excess bond premium because of its ability to explain business cycles (Gilchrist and Zakrajsek, 2012) and as an indicator of the price of risk (Creal and Wu, 2016). The variables in our baseline model are thus ordered: cumulative monetary policy shock series, log industrial production, log consumer price index, log commodity price index, and excess bond premium. We use 12 monthly lags.[20]

Figure 4a presents the impulse responses to a contractionary monetary shock using the full sample (1994-2017). Here and throughout the paper we normalize to a 100 basis point positive monetary policy shock on impact. The 68% and 90% standard error confidence intervals, displayed as deep and shallow gray areas respectively, are generated by the bootstrap. Both output and inflation decrease after a contractionary monetary policy shock. The responses reach their troughs after about 10 months. The excess bond premium increases and peaks after about 8 months. These results are conventional, in line with those of Gertler and Karadi (2015), for example.

Figure 4b shows the impulse responses when the model is estimated on the post-2008 sub-sample. The responses are similar. Output and inflation significantly decrease for the first 10 months after a contractionary monetary policy shock, while the excess bond premium increases significantly. Figure 4c shows the responses from the pre-2008 sub-sample. Output decreases immediately and

[19] This also follows Romer and Romer. Our series and theirs are plausibly exogenous, given how they are constructed.

[20] We also examine systems with the 5-year interest rate as an additional variable in the VAR model. These generate similar impulse responses.

reaches its trough about 2 years later. Inflation exhibits a steady downward pattern, and the excess bond premium increases significantly 10 months after the shock. Thus, the impulse responses from a shock to the BRW series in a conventional VAR model are highly stable across sub-periods.

4.2 IRF Robustness with BRW Shocks

In light of standard concerns about potential dynamic mis-specification in VAR models, our first robustness check is to re-estimate using Jorda (2005) local projections.[21] This constructs impulse responses from time-series regressions for each point in time. Appendix Figure A1a presents the impulse responses to a contractionary monetary policy shock using the full sample (1994-2017). After a positive shock, industrial production significantly decreases about 2 months later and reaches its trough after 15 months. Inflation immediately and sharply decreases throughout the 24 months. The excess bond premium responds positively through the first 10 months. Figures A1b and A1c show that results for the sub-periods estimated using local projections are very similar to those of the full sample and hence similar to those estimated from the VAR model.

The next robustness check concerns the term premium. For this purpose, we subtract from the raw interest rates the corresponding term premium on the 5-year Treasury rate and all the zero-coupon yields with 1 to 10-year maturity, as estimated by Adrian, Crump, and Moench (2013). We then reconstruct our monetary policy shock series excluding the term premium. Inserting the cumulative values of that series into the baseline VAR model, we find that the impulse responses are quantitatively identical to the baseline results of Figure 4, although the negative effect on IP is dampened for the first few months (see online Appendix). As shown in column 9 of Table 2, the correlation between the term-premium free shock and our baseline shock is high, 0.79.

4.3 Alternatives: Nakamura-Steinsson, Swanson, and Jarocinski-Karadi

We compare the impulse responses above to those estimated by replacing our shock series with that of, alternately, Nakamura and Steinsson (2018) and Swanson (2018), both of which embody the information effect (Table 4). Nakamura and Steinsson do not directly estimate the effects of their policy news shock on output (nor does Swanson (2018)), but rather focus on the response of expectations of future output growth and real interest rates in a non-VAR framework. These authors also do extensive quantitative modeling and conclude from their estimation of the model that roughly

[21] Again this follows Romer and Romer (2004), who estimate a VAR with cumulative monetary policy shocks and also estimate a version of local projections.

two-thirds of the monetary shock is due to the Fed information effect. Following Gurkaynak, Sack, and Swanson (2005), Swanson (2018) argues that monetary policy has more than one dimension. Changes in the federal funds rate are different from forward guidance announcements, and both of these are different from LSAP announcements, at least in terms of their effects on financial markets. The various shock measures from the Swanson papers thus reflect the effect of, e.g., a 25bp decline in long rates that is carried out through an increase in asset purchases versus one that is accomplished via stimulative forward guidance or a drop in the target rate. Interpretation of the effects of shocks to our series is different but complementary. Our estimates represent the effects of an FOMC meeting day shock that reflects the effect of, e.g., a 25bp decline in the 5-year rate following the words and actions (or inactions) undertaken by the FOMC. Our measure is best thought of telling us the effect of an average 25bp loosening of the 5-year Treasury yield following the FOMC meeting, where this average is in principle a combination of Fed funds rate loosening, some expansionary forward guidance, and some LSAP increases.[22]

Figure 5 presents the results. The sample periods are: full (1994-2015), pre-ZLB (1994-2007) and during the ZLB (2008-2015). For the full and pre-ZLB sub-sample (Figures 5a and b), impulse responses using any of the shocks follow the conventional monetary model. Output and inflation decrease while the excess bond premium increases after a contractionary monetary policy shock. However, during the ZLB sub-sample (Figure 5c), the impulse responses differ across cases. Following a positive shock to the Nakamura-Steinsson measure, both output and inflation *rise* significantly after about 10 months. In response to the shock identified by Swanson (FG plus LSAP), output, inflation and excess bond premium effectively do not change.[23]

To further assess the possible role of Fed private information in accounting for differences in the transmission effects during the ZLB period shown in Figure 5, we replace the original shock series with the residual from the regression of Table 5.[24] This "purged" series represents that component of the raw monetary policy shock that is not accounted for by differences in the Fed-private sector outlook. Impulse responses using the shock series of NS, Swanson, and BRW are reported in Appendix Figure A2a-c, respectively. In the left panels, we depict point estimates and

[22] This can be thought of as a "FRB-US view of the world", in the sense that it mimics how Federal Reserve Board staff analyze monetary policy in their large scale estimated general equilibrium model of the U.S. economy (https://www.federalreserve.gov/econres/us-models-about.htm).

[23] We also perform this exercise with each of the separate Swanson shocks and find similar results. In addition, we estimate impulse responses to identified shocks to the Wu-Xia shadow rate index. During the ZLB, impulse responses are conventional and significant at first, but exhibit the opposite sign at long horizons. Wu and Xia estimate a FAVAR model, different from the basic VAR here, and report conventional responses. For example, they find that expansionary Fed monetary policy shocks raise IP and lower unemployment during the period July 2009 to December 2013, in much the same way that shocks to the effective Fed Funds rate did prior to the ZLB period.

[24] Miranda-Agrippino and Ricco (2017) and Kane, Rogers, and Sun (2018) pursue a similar strategy.

confidence bands from the VARs with the orthogonalized series. In the far right panels are IRFs using the original shock series. The middle column presents the comparison, omitting confidence bands for ease of viewing. For both NS and Swanson purged shocks, the positive responses of output to a contractionary policy shock are diminished compared to IRFs from the raw shocks. Indeed, the responses of shocks to the purged Swanson measure have conventional signs (Figure A2b). With BRW shocks, for which the Fed information effect is insignificant, there are no differences in impulse responses across the two experiments (Figure A2c).

As noted in Section 3, Jarocinski and Karadi (2018) argue that the information effect is empirically important by showing that output, price level, and excess bond premium respond with significantly different signs to a monetary policy shock compared to the shock conditioned on stock prices and interest rates co-moving positively, which they label central bank information shocks. In Figure 6A, we replicate the results of Jarocinski and Karadi (2018) using their monetary policy surprise FF3, while in Figure 6B we re-estimate using our new shock and find quite different results. We depict impulse responses on "non-information effect days", points in the second and fourth quadrants of Figure 3, and on "information effect days", points in the first and third quadrants.[25] In the left (right) panels, we report the point estimates and error bands for the non-information (information) day shocks. In the middle column, we display the point estimate comparison without confidence bands. Consider panel A first. On non-information effect days, the impulse responses exhibit traditional signs. Output and price level fall in response to a monetary contraction, while credit conditions tighten (EBP rises). Impulse responses on information-effect days, the right side column (in blue), produce significantly different results, however, with the transmission effects changing signs. The results are noticeably different when we use our new shock series in place of FF3,however. Transmission effects to output, prices, and credit conditions exhibit conventional signs, irrespective of estimating on information effect days or non-information effect days.[26]

As a final check, we estimate impulse responses from shocks to the various measures constructed in our encompassing analysis of section 3. Results are displayed in the Appendix figures. Responses to the "PCA shock", which embodies the information effect, are unconventional: muted in the full sample and moving in the "wrong" direction during the ZLB period (Figure A4). Impulse responses to a positive "Tight (NS data)" shock, which is devoid of the information effect, look more

[25] We use all available data in the VAR these experiments, and simply set shocks on the other days to zero. This is equivalent to the second of two estimation procedures used by Jarocinski and Karadi, labelled poor man sign restrictions.

[26] Recall from Table 4 that there is little evidence of an information effect, in the sense of NS, in the BRW series even on JK information effect days.

conventional: in the post-2008 sample, the IP and CPI responses are mostly negative, especially at intermediate horizons; the response of EBP is less negative at first and quickly turns positive after a short period of time (Figure A5). Finally, positive shocks to the "Tight (full data)" shock series, also devoid of the information effect (Table 4, row 10), produce impulse responses with *conventional* signs, albeit with some lagged effects compared to those with baseline BRW shocks (Figure A6).

5 Conclusion

We perform a novel application of well-known estimation procedures to derive a US monetary policy shock series that usefully bridges periods of conventional and unconventional policymaking and is effectively devoid of the information effect. Our approach has very mild data requirements and is easy to implement econometrically. The heteroskedasticity-based estimator filters out background noise, while the monetary policy shock is aligned using Fama-MacBeth regressions. We demonstrate the importance of our procedure to the identification of monetary policy shocks through detailed comparison with alternative measures in the literature, including an investigation of the Fed information effect. Overall, using the same testing and "purging" procedures as two prominent approaches in the literature, we find essentially no evidence of an information effect in our new monetary policy shock series.

We then present evidence confirming an hypothesis in the literature that the information effect can lead to monetary policy shocks having transmission effects to output and inflation with signs that differ from those predicted by traditional theory. We find that in response to contractionary shocks to our new measure, output and prices fall significantly, consistent with conventional theory. This result is found in samples both before the ZLB and during the ZLB sub-period with our measure. However, estimating impulse responses to monetary policy shocks that embody the information effect, we find responses that are either zero or positive.

Staff at the Federal Reserve and other central banks want and need to know whether their models should be constructed to feature the information effect. Should the impulse responses associated with monetary policy announcements that the staff's quantitative models attempt to match be of the signs predicted by traditional monetary theory, or of the unconventional signs consistent with the evidence in influential papers like Nakamura-Steinsson and Jarocinski-Karadi?The evidence in this paper, and our unified measure, are useful for guiding these and other exercises in empirical and quantitative theoretical modeling of the effects of Fed monetary policy.

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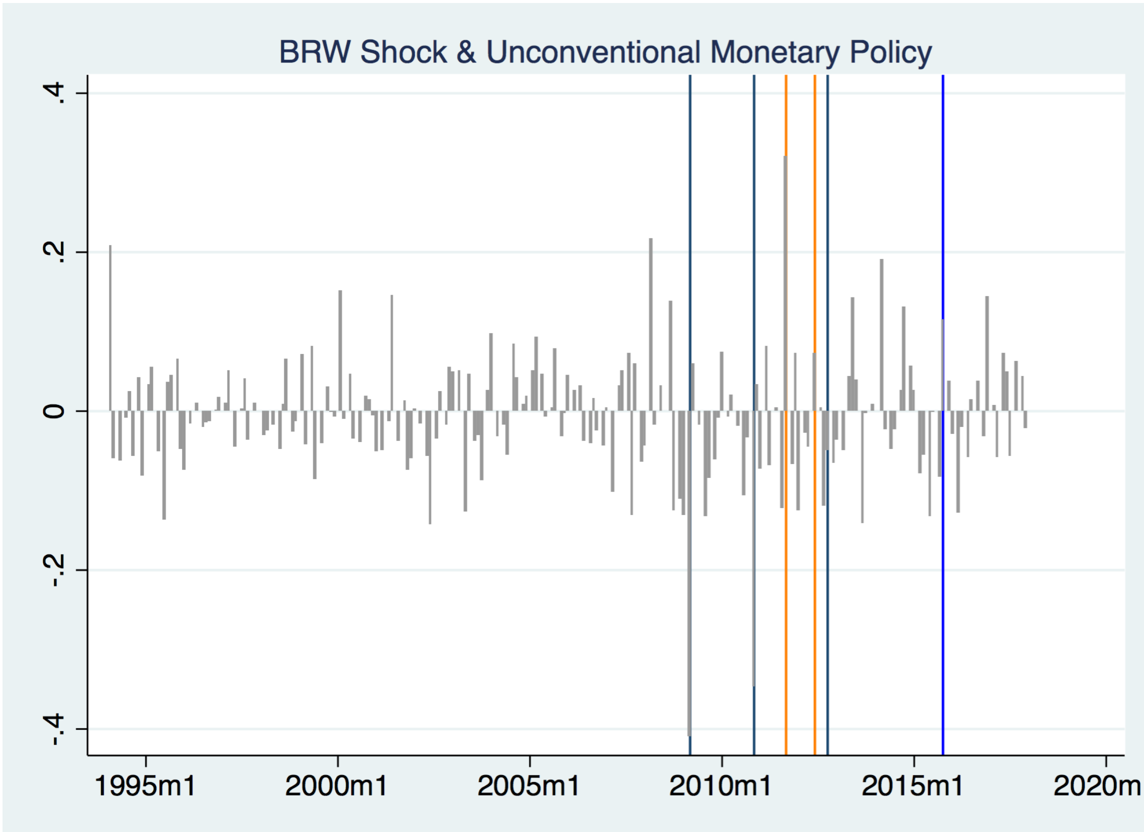
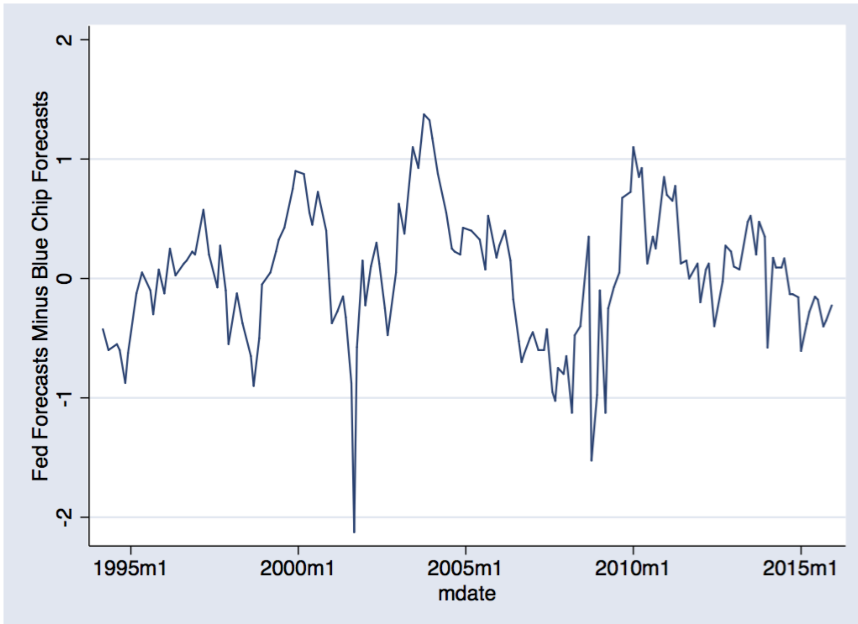
Figure 1: BRW Shock Series Jan 1994 to Dec 2017



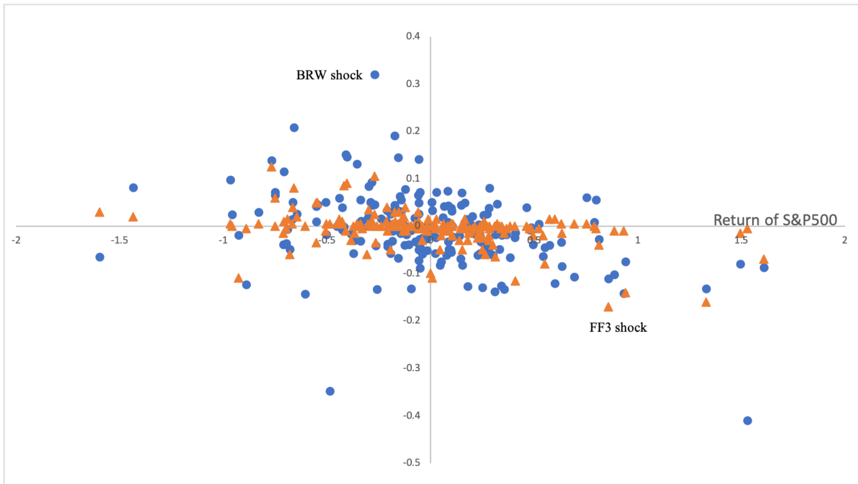
Note: The BRW shock series is estimated from Equations (3) and (4). The navy vertical lines denote announcements of QE1, QE2, and QE3; the orange vertical lines denote the Operation Twist period; and the blue line denotes Oct. 2015, the FOMC meeting prior to liftoff.

Figure 2: GDP Growth Forecasts, Fed Minus Blue Chip



Note: Prior to December 2013, this is the average of the first four quarters ahead Greenbook forecasts less the corresponding Blue Chip forecasts. After January 2014, we use forecasts from the FOMC summary of economic projections (SEP) because the Greenbook data is not yet publicly available. The Fed SEP are available four times per year-in March, June, September, and December. For the other four FOMC meetings, we use the SEP from the previous FOMC meeting. We use the current year SEP forecast for real GDP growth rate if the FOMC meeting happens in the first quarter of the year. Otherwise, we use the next year SEP forecast for real GDP Growth.

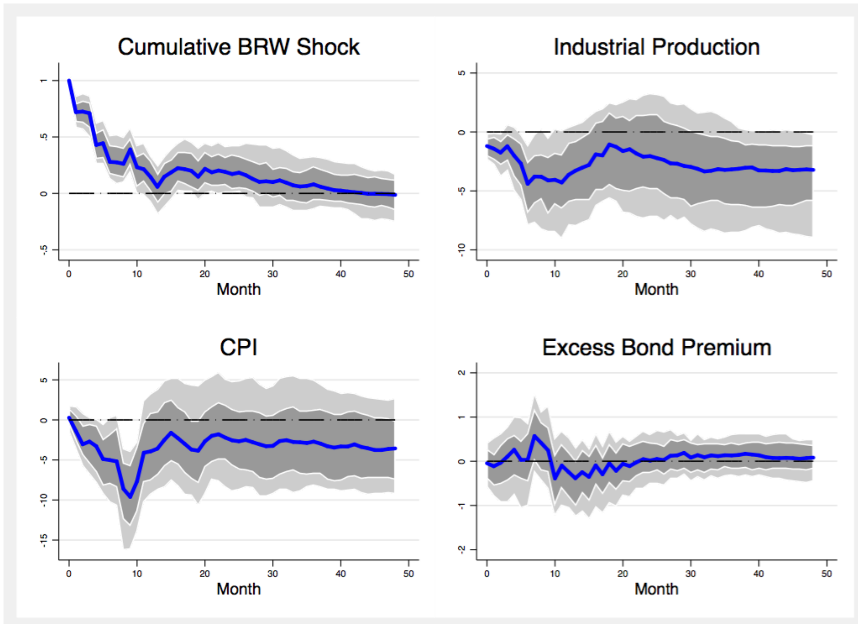
Figure 3: S&P 500, the BRW Shock, and the JK Shock



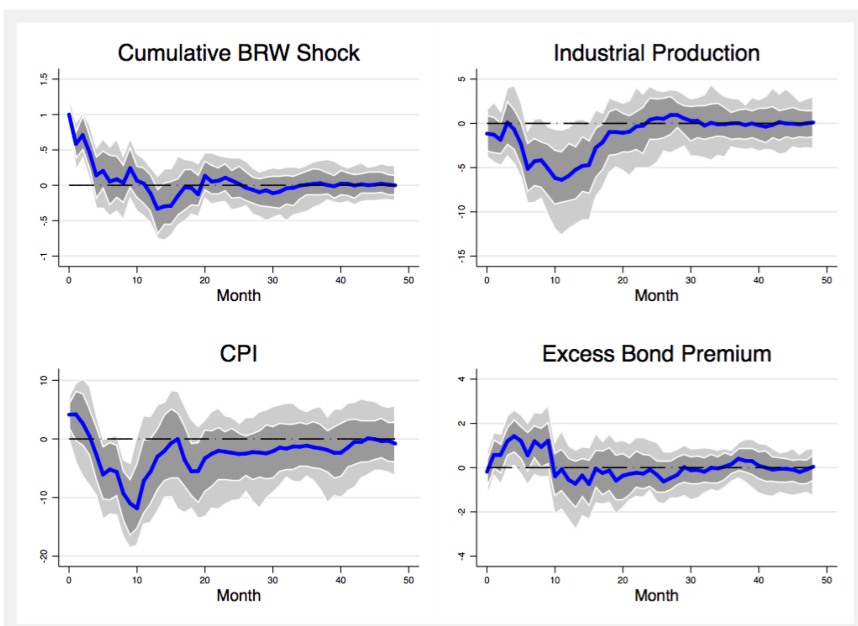
Note: The S&P 500 returns are computed over a 30-minute window around FOMC meeting announcements. The blue dots represent the BRW shocks, and the orange triangles are the surprises of the 3-month federal funds futures that are used by Jarocinski and Karadi (2018).

Figure 4: Baseline SVAR Impulse Responses: BRW Shocks

a. 1994m1-2017m12



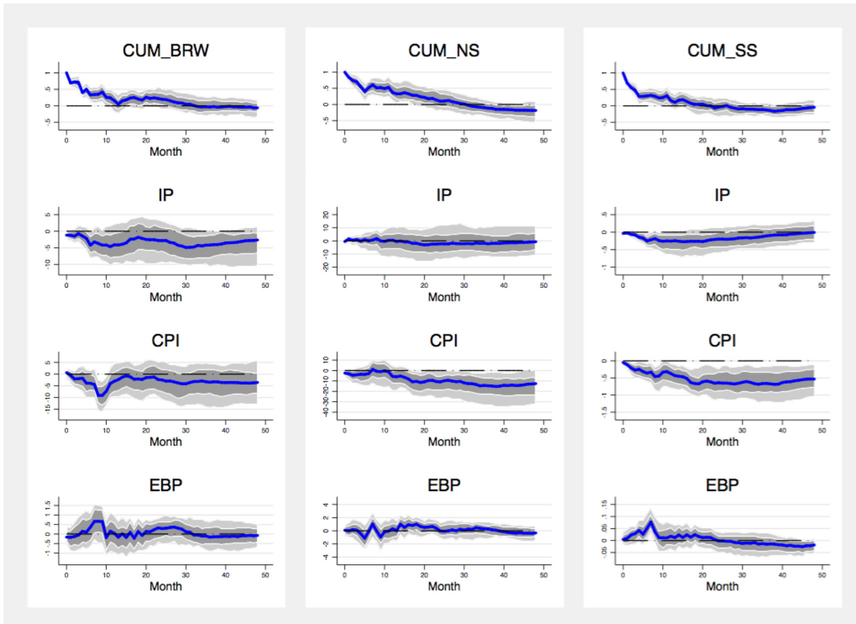
b. 2008m1-2017m12



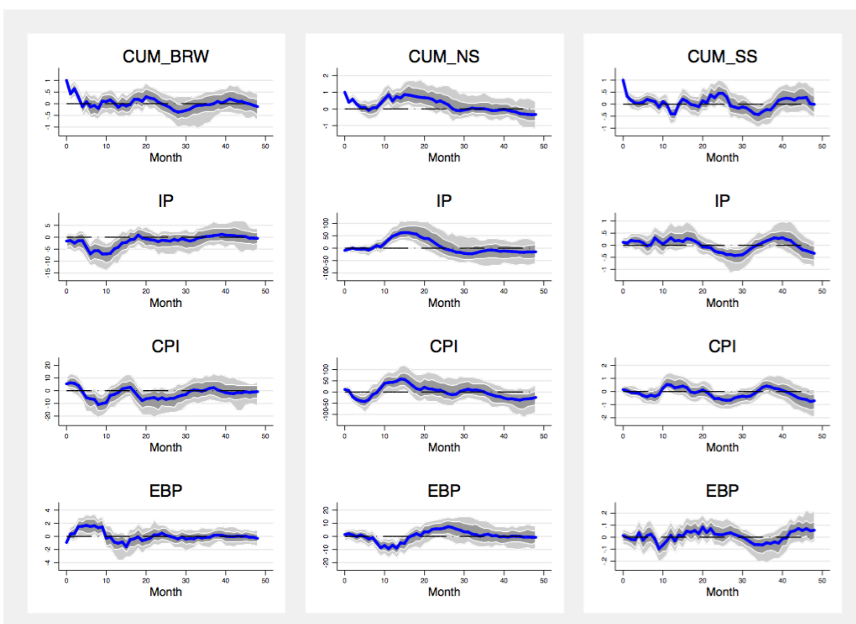
Note: Structural VAR with monthly data, 5 endogenous variables and 12 lags. Variables are ordered as follows: cumulative BRW shock series, log industrial production, log consumer price index (CPI), log commodity prices, and excess bond premium. Graphs show impulse responses estimated over different sample periods to a 100 basis point increase in the cumulative BRW shock series. Deep and shallow gray shaded areas are 68% and 90% confidence intervals produced by bootstrapping 1000 times, respectively.

Figure 5: SVARs with Alternative Shock Series: BRW, NS, and Swanson

a. 1994m1-2015m12



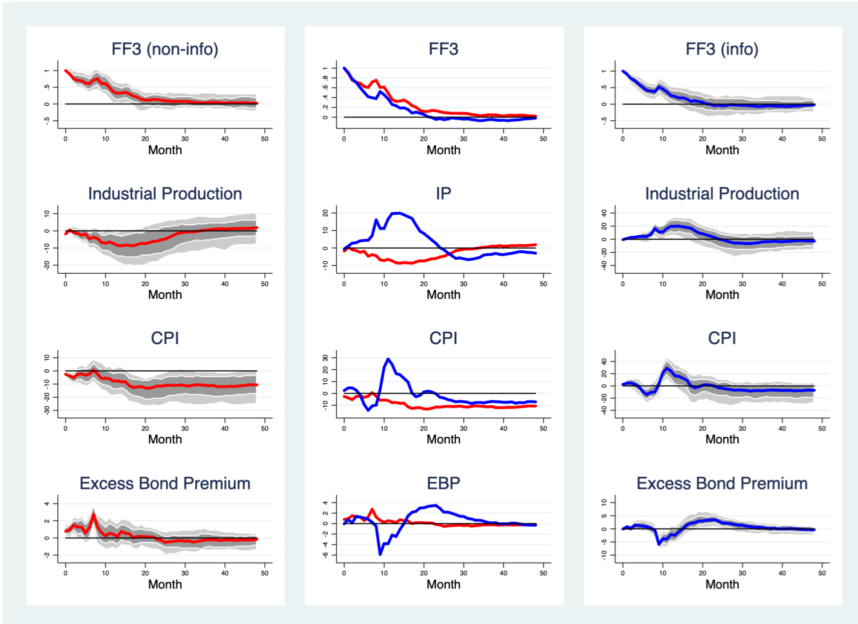
b. 2008m1-2015m12



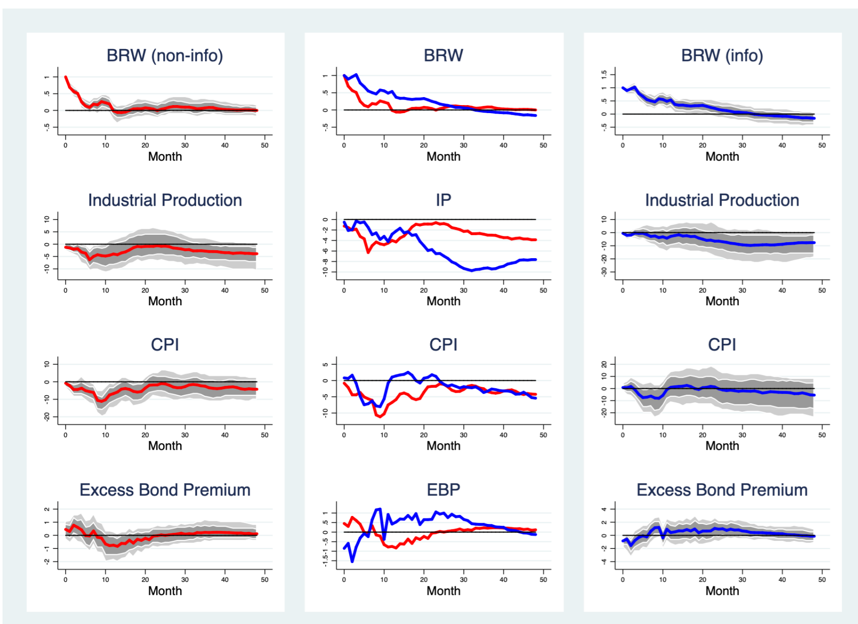
Note: BRW, NS and Swanson refer to cumulative BRW shock series, Nakamura and Steinsson (2018) shock series, and Swanson (2017) shock series, respectively. For these cases, variables are ordered: the cumulative shock series, log industrial production, log consumer price index (CPI), log commodity prices, and excess bond premium. Graphs show impulse response to a 100 basis point increase in the monetary policy indicator series. Deep and shallow gray shaded areas are 68% and 90% confidence intervals produced by bootstrapping 1000 times, respectively.

Figure 6: SVAR on Non-information Days (red) and Information Days (blue)

a. Using the Jarocinski-Karadi FF3 Shock



b. Using BRW Shock



Note: Full sample-period estimation. FF3 is accumulated 3 month federal funds futures rate around the 30-minute FOMC announcement window according to the information day definition in Jarocinski and Karadi (2018). The BRW shock is accumulated in the same way.

Table 1: **Correlation with BRW Shock Series**

	Full Sample	Pre-ZLB	ZLB
NS Shock	0.512	0.653	0.494
SS shock	0.625	0.684	0.532
R&R Shock		0.131	
Kuttner Shock		0.308	
SS_FFR		0.373	
SS_FG	0.492	0.605	0.575
SS_LSAP			0.365
FF3	0.395	0.593	0.336

Note: The benchmark shock is our BRW shock series estimated from Equation (3) and (4). *NS* Shock refers to the policy factor shock of Nakamura and Steinsson (2018), which we update to the present. *SS Shock* refers to the sum of the shock series of the federal funds rate, the forward guidance and the large asset purchases in Swanson (2018). *R&R Shock* refers to the estimated shock series in Romer and Romer (2004). *Kuttner Shock* refers to the 30-minute Fed Funds rate changes around FOMC announcements. *SS_FFR*, *SS_FG*, *SS_LASP* refers to the shock series of the Federal Funds rate, forward guidance and large asset purchases in Swanson (2018). *FF3* is the 30-minute change in 3 month federal funds futures rate around the FOMC announcement. Sample periods are: Full sample 1994m1-2017m12, Pre ZLB 1994m1-2008m12, ZLB 2009m1-2015m12.

Table 2: **Shock Series Robustness: Correlations with Baseline BRW Shock Series**

	N2	N10	BRW69	R5	QE	Unschedule	Day2	IV2	BRW (RT1)	BRW (RT2)	TP	OLS	PCA	Tight (NS)	Tight (Full)
BRW Shock	0.975	0.981	0.983	0.957	0.987	0.903	0.838	0.995	0.954	0.883	0.791	0.992	0.249	0.380	0.500
Observations	191	191	191	191	190	183	191	191	191	191	191	191	190	191	191

Note: *BRW Shock* refers to our BRW shock series estimated from Equation (3) and (4).
N2 refers to the BRW shock series aligned from using the 2-year Treasury Rate as policy indicator.
N10 refers to the BRW shock series aligned from using the 10-year Treasury Rate as policy indicator.
BRW69 refers to our BRW shock series estimated from the sample over 1969m1 to 2017m12.
R5 refers to the BRW shock series aligned using zero-coupon yields with only the 1, 2, 5, 10, 30-year maturities as outcome variables.
QE refers to the BRW shock series excluding the announcement of QE1 in March, 2009.
Unschedule refers to the BRW shock series aligned including all of the unscheduled FOMC meeting dates since 1995.
Day2 refers to the BRW shock series aligned using a 2-day event window around FOMC announcement days.
IV2 refers to the BRW shock series aligned using daily movements in the policy indicator 1-day before FOMC announcement day rather than one week as the instrumental variable.
BRW (RT1) refers to BRW shock series combining rolling sample method post 2008 and original BRW shock before 2008.
BRW (RT2) refers to BRW shock series aligned from sensitivity indexes of pre-2008 subsample.
TP refers to the BRW shock series generated as the baseline approach of Equation (3) and (4) but free of the estimated term premium.
OLS refers to the alternative BRW shock series aligned from the simple Fama-Macbeth method without the IDH procedure.
PCA refers to the shock series generated from extracting the first principal component of our underlying data, i.e., all outcome variables (daily changes of 1 to 30-year zero coupon rate around FOMC meeting).
Tight(NS) refers to the BRW shock series using the data underlying Nakamura and Steinsson (2018), i.e., the 30-minute changes of the current month Fed funds futures rate, the Fed funds futures rate immediately following the next FOMC meeting, and two, three, four quarter ahead euro dollar futures.
Tight(Full) refers to the BRW shock series using the NS data and the 30-minute changes of the 3 month, 6 month, 2 year, 5 year, 10 year, 30 year interest rates around FOMC announcements.

Table 3: **Monetary Policy Shocks and the Slope of the Yield Curve**

	5y	6m - 5y	1y - 5y	2y - 5y	10y - 5y	30y - 5y
BRW	0.679***	-0.432***	-0.351***	-0.113***	-0.232***	-0.782***
	(0.05)	(0.05)	(0.05)	(0.04)	(0.02)	(0.02)
NS	1.102***	-0.211	-0.175	0.076	-0.366***	-0.990***
	(0.14)	(0.14)	(0.12)	(0.08)	(0.05)	(0.11)
SS (FG)	0.508***	-0.350***	-0.284***	-0.0645**	-0.111***	-0.342***
	(0.05)	(0.05)	(0.04)	(0.03)	(0.02)	(0.04)
SS (LSAP)	0.575***	-0.588***	-0.529***	-0.346***	0.0977***	-0.185**
	(0.08)	(0.07)	(0.06)	(0.04)	(0.03)	(0.08)
FF3 (JK info)	-0.292	0.659***	0.472***	0.302**	-0.0124	-0.0773
	(0.18)	(0.19)	(0.16)	(0.12)	(0.06)	(0.14)
FF3 (Non JK info)	0.867***	-0.175	-0.1	0.0563	-0.350***	-0.830***
	(0.16)	(0.15)	(0.13)	(0.08)	(0.06)	(0.13)

Note: Constant term not displayed. Robust standard errors in brackets. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. *5y* refers to the daily change in the 5-year treasury bond yield around the FOMC announcement. *6m-5y, 1y-5y, 2y-5y, 10y-5y, and 30y-5y* refer to the differences between the daily changes in 6 month, 1, 2, 10, and 30 year treasury bond yields around the FOMC announcement and the 5-yr. rate. The (updated) NS Shock is the shock series updated to 2015m12 following the method in Nakamura and Steinsson (2018). The regressions are estimated over each authors' full sample periods. Sample periods are 1994m1-2018m8 for BRW shock series, 1994m1-2015m12 for NS shock series, and 1994m1-2015m11 for Swanson's FG and LSAP shock series.

Table 4: **Fed Information Effect Regressions of Nakamura and Steinsson (2018)**

	BRW	SS	NS	BRW (RT1)	BRW (RT2)	BRW (JK Info)	BRW (JK Ninfo)	FF3 (JK Info)	FF3 (JK Ninfo)	PCA	BRW Tight (NS data)	BRW Tight (full data)
1995-2015	0.01	0.16**	0.76***	-0.01	-0.15	0.39	-0.03	1.06***	0.37	0.53**	-0.15	-0.14
	(0.16)	(0.07)	(0.21)	(0.16)	(0.17)	(0.40)	(0.16)	(0.22)	(0.23)	(0.26)	(0.13)	(0.29)
Obs	134	136	135	134	134	50	84	24	76	133	137	137

Note: Monthly change (current month to next) in Blue Chip survey expectations of output growth over the next 3 quarters regressed on the shock series in that month plus a constant (not displayed). Sample periods are listed at top. Robust standard error in brackets. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. *BRW* refers to our shock series: our data and our PLS estimation procedure. *SS* refers to the sum of the shock series of the Federal Funds rate, forward guidance and large scale asset purchases of Swanson (2018), scaled by 10. *NS* refers to the policy news shocks of Nakamura and Steinsson (2018). *BRW(JK Info)* and *BRW(JK Ninfo)* are two sub-sample regressions focusing only on the Fed information effect days and non-information effect days, as defined by Jarocinski and Karadi (2018). *FF3(JK Info)* and *FF3(JK Ninfo)* are the surprises of the 3-month federal funds futures focusing only on the Fed information effect days and non-information effect days, as used by Jarocinski and Karadi (2018). *PCA* refers to the shock series generated from extracting the first principal component of our underlying data, i.e., all outcome variables (daily changes of 1 to 30-year zero coupon rate around FOMC meeting), scaled by 100. *BRW Tight(NS data)* refers to the BRW shock series computed using PLS with the data in Nakamura and Steinsson (2018), i.e., the 30-minute changes of the current month Fed funds futures rate, the Fed funds futures rate immediately following the next FOMC meeting, and two, three, four quarter ahead euro dollar futures. *BRW Tight(full data)* refers to the BRW shock series using the NS data *plus* the 30-minute changes of the 3 month, 6 month, 2 year, 5 year, 10 year, 30 year interest rates around FOMC announcements. *BRW (RT1)* refers to BRW shock series combining rolling sample method post 2008 and original BRW shock before 2008. *BRW (RT2)* refers to BRW shock series aligned from sensitivity indexes of pre-2008 subsample. *SS (FG), SS (FFR), SS (LSAP)* refers to the shock series of forward guidance, Federal Funds rate and large scale asset purchases of Swanson (2018), respectively, all scaled by 10.

Table 5: **Shock Series Regressed on Fed minus Blue Chip GDP Growth Forecasts**

	(1) NS Shock	(2) Updated NS Shock	(3) BRW Shock	(4) Swanson Shock
Fed - BC	2.00**	1.93***	1.95	0.67**
	(0.77)	(0.70)	(1.53)	(0.31)
Observations	130	150	150	149
R-squared	0.09	0.08	0.02	0.07

Note: Constant term not displayed. Robust standard error in brackets. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. *BRW Shock* refers to our BRW shock series estimated from Equation (3) and (4). *SS Shock* refers to the sum of the shock series of the federal funds rate, the forward guidance and the large asset purchases proposed by Swanson (2018). We scale the SS shock by 100. *NS Shock* refers to the policy factor shocks from Nakamura and Steinsson (2018). The updated NS Shock is the shock series updated to 2015m12 following the method in Nakamura and Steinsson (2018). *Fed - BC* is the difference between Fed and Blue Chip GDP growth Forecasts, constructed as described above. Sample periods are: 1995m1-2014m3, 1994m1-2015m12, 1994m1-2015m12, and 1994m1-2015m11 (Swanson's sample ends just before lift-off).

Table 6: **Fed Information Effect in Interest Rates with Different Maturities**

	Kuttner	6-month	2-yr.	5-yr.	10-yr.	30-yr.
Coef.	0.296***	0.389*	0.368**	0.277	0.308	0.214
	(0.11)	(0.22)	(0.17)	(0.18)	(0.22)	(0.30)
Observations	144	144	144	144	144	144
R-squared	0.04	0.024	0.034	0.017	0.012	0.004

Note: Constant term not displayed. Robust standard error in brackets. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. We regress the monthly change (current month to next) in survey expectations of output growth over the next 3 quarters from Blue Chip Economic Indicators on the shock series in that month. *Kuttner Shock* refers to monetary policy shock of Kuttner(2001). *6 month* refers to the 30-minute change in 6 month treasury note yield around the FOMC announcement. *2, 5, 10, and 30 year* refer to the 30-minute changes in 2, 5, 10, and 30 year treasury bond yields around the FOMC announcement. The sample period is 1994m1-2018m8. Following NS, we exclude the Great Recession period.

Appendix

A1. Implementation of Identification through Heteroskedasticity - IV approach

We assume the monetary policy shock is unobservable. We normalize the shock to have 1-1 relationship with the changes in the 5 year interest rate,

$$\Delta R_{5,t} = \alpha_0 + e_t + \eta_t. \tag{7}$$

The equation of interest is

$$\Delta R_{i,t} = \theta_i + \beta_i \Delta R_{5,t} + \xi_{i,t} \tag{8}$$

where $\xi_{i,t} = -\beta_i \eta_t + \epsilon_{i,t}$, where $\epsilon_{i,t}$ is the idiosyncratic error associated with $\Delta R_{i,t}$, $\epsilon_{i,t}$ is assumed not to correlate with the monetary policy shock e_t, and $\Delta R_{i,t}$ is the change in i year interest rate around FOMC announcements.

For simplicity and without loss of generality, we suppress the subscript i, and demean both $\Delta R_{i,t}$ and $\Delta R_{5,t}$,

$$\Delta R_t = \beta \Delta R_{5,t} + \xi_t. \tag{9}$$

Heteroskedasticity-based estimation – By construction, the regressor $\Delta R_{5,t}$ is correlated with the error term $\xi_{i,t}$ due to the component $-\beta_i \eta_t$. The OLS estimation of β_i is biased due to the errors-in-variables problem.

To deal with this problem, we need to identify two subsamples, which are denoted as M and NM. M is the sample with event windows around FOMC announcements and NM represents the non-monetary windows, which are the corresponding event windows one week before. We also need two assumptions regarding the second moment of the shocks present in the model: on days of FOMC meetings, the variance of the true monetary policy shock increases while that of the background noise remains unchanged.

Assumption 1: $\sigma_e^M > \sigma_e^{NM}$, $\sigma_\eta^M = \sigma_\eta^{NM}, \sigma_\xi^M = \sigma_\xi^{NM}$.

Assumption 2: $E[\eta_t e_t] = E[\xi_t e_t] = 0$.

The implementation is very similar to Rigobon and Sack (2004). Denote the variance covariance matrix of each subsample as

$$\begin{aligned} \Omega^M &= E\left[\left[\Delta R_{5,t}^M \ \Delta R_t^M\right]' * \left[\Delta R_{5,t}^M \ \Delta R_t^M\right]\right] \\ \Omega^{NM} &= E\left[\left[\Delta R_{5,t}^{NM} \ \Delta R_t^{NM}\right]' * \left[\Delta R_{5,t}^{NM} \ \Delta R_t^{NM}\right]\right] \end{aligned} \tag{10}$$

It is clear that

$$\begin{aligned} \Omega^M &= E\begin{bmatrix} \left(\Delta R_{5,t}^M\right)^2 & \Delta R_{5,t}^M \Delta R_t^M \\ \cdot & \left(\Delta R_t^M\right)^2 \end{bmatrix} \\ &= \begin{bmatrix} \left(\sigma_e^M\right)^2 + \left(\sigma_\eta^M\right)^2 & \beta \left(\sigma_e^M\right)^2 \\ \cdot & \beta_1^2 \left(\sigma_e^M\right)^2 + \left(\sigma_\xi^M\right)^2 \end{bmatrix} \end{aligned}$$

The second equality follows from $E[\eta_t e_t] = E[\xi_t e_t] = 0$. Similarly, we can write Ω^{NM} out in terms of σ_η^{NM} and σ_ξ^{NM}.

If we take the difference between these two covariance matrices and let $\left(\sigma_e^M\right)^2 - \left(\sigma_e^{NM}\right)^2 = \lambda$, we have

$$\begin{aligned} \Delta\Omega &= \Omega^M - \Omega^{NM} \\ &= \begin{bmatrix} \lambda & \beta\lambda \\ \cdot & \beta^2\lambda \end{bmatrix} \end{aligned}$$

$$= \lambda \begin{bmatrix} 1 & \beta \\ \cdot & \beta^2 \end{bmatrix}$$

Then, it is clear that β can be estimated as follows,

$$\hat{\beta}_1 = \frac{\Delta\hat{\Omega}_{12}}{\Delta\hat{\Omega}_{11}}$$

Now,

$$\begin{aligned} \hat{\beta}_1 &= \frac{\Delta\hat{\Omega}_{12}}{\Delta\hat{\Omega}_{11}} && (11) \\ &= \frac{cov\left(\Delta R_{5,t}^{M}, \Delta R_t^{M}\right) - cov\left(\Delta R_{5,t}^{NM}, \Delta R_t^{NM}\right)}{var\left(\Delta R_{5,t}^{M}\right) - var\left(\Delta R_{5,t}^{NM}\right)} && (12) \\ &= \frac{E\left[\left(\Delta R_{5,t}^{M}, -\Delta R_{5,t}^{NM}\right)\left(\Delta R_t^{M}, \Delta R_t^{NM}\right)'\right]}{E\left[\left(\Delta R_{5,t}^{M}, -\Delta R_{5,t}^{NM}\right)\left(\Delta R_{5,t}^{M}, \Delta R_{5,t}^{NM}\right)'\right]} && (13) \end{aligned}$$

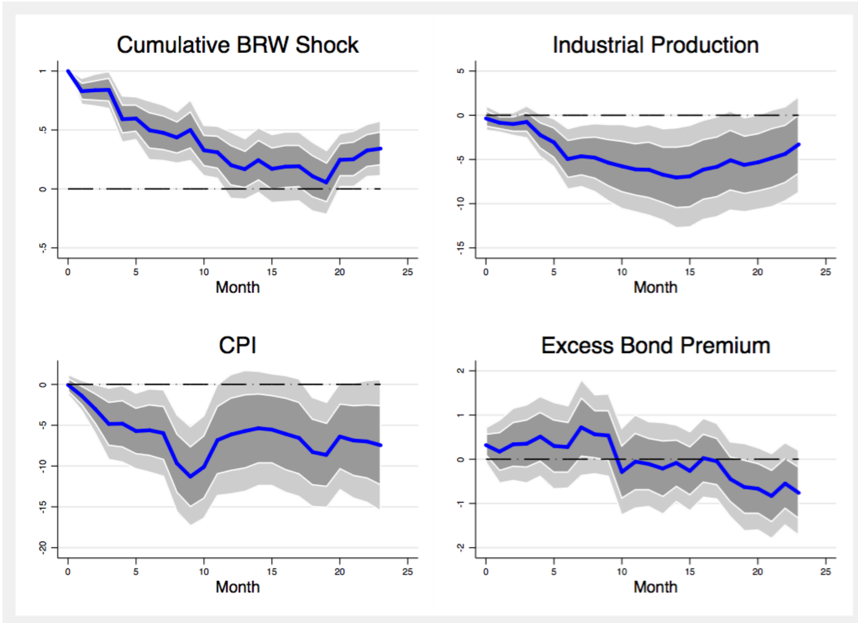
According to (13), we may use an IV approach to implement this estimator. This approach rewrites (8) as:

$$[\Delta R_{i,t}] = \alpha_i + \beta_i[\Delta R_{5,t}] + \mu_{i,t} \quad i = 1, 2, \cdots, 30 \tag{14}$$

where the independent variable $[\Delta R_{5,t}] = (\Delta R_{5,t}^{M}, \Delta R_{5,t}^{NM})'$, the event window of $[\Delta R_{i,t}]$ corresponds to $[\Delta R_{5,t}]$. β_i can be estimated using an instrumental variable $\Delta R_t^{IV} = (\Delta R_{5,t}^{M}, -\Delta R_{5,t}^{NM})'$ for the independent variable. Intuitively, $\left(\Delta R_{5,t}^{M}, -\Delta R_{5,t}^{NM}\right)'$ is able to instrument $\left(\Delta R_{5,t}^{M}, \Delta R_{5,t}^{NM}\right)'$ because, (1) it is clear that they are correlated; (2) $\left(\Delta R_{5,t}^{M}, -\Delta R_{5,t}^{NM}\right)'$ does not correlate with the error terms, which follows directly from Assumption 1 & 2.

Figure A1: BRW Shock Series IRFs using Jorda (2005) Local Projections Method

a. 1994m1-2017m12



b. 2008m1-2017m12

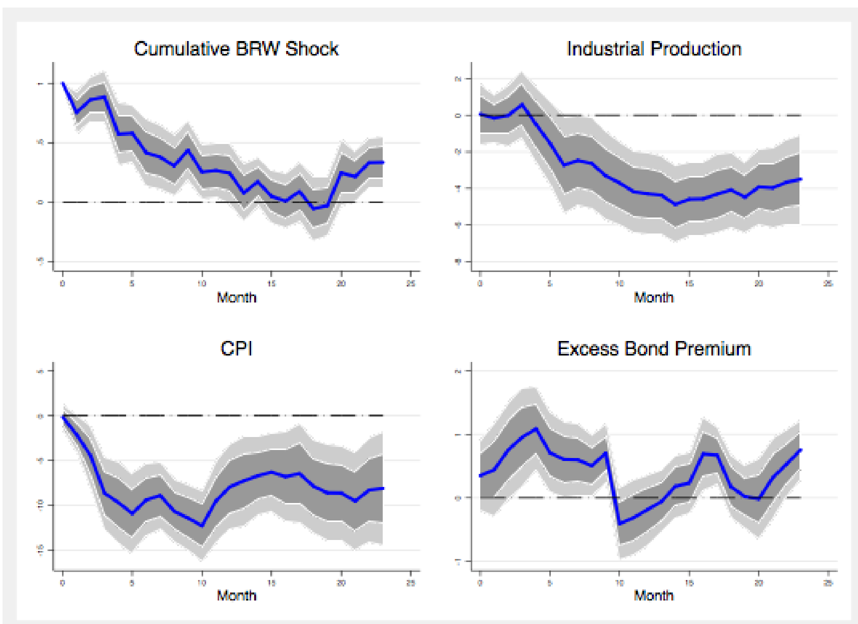
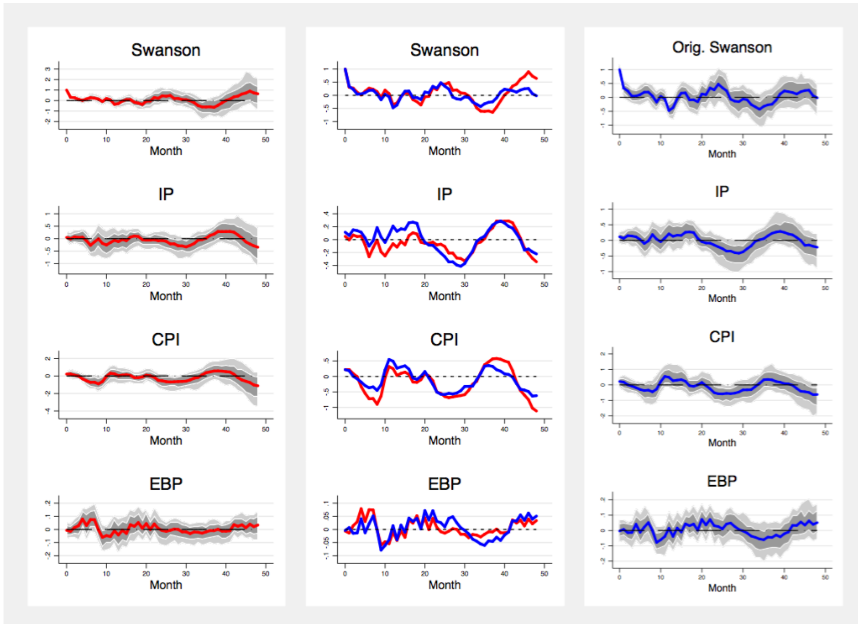


Figure A2: SVARs using shock series purged of the information effect

a. Swanson Shock: Original (blue) versus Purged (red) Shock Series (table 5 residual)



b. N&S Shock: Original (blue) versus Purged (red) Shock Series (table 5 residual)

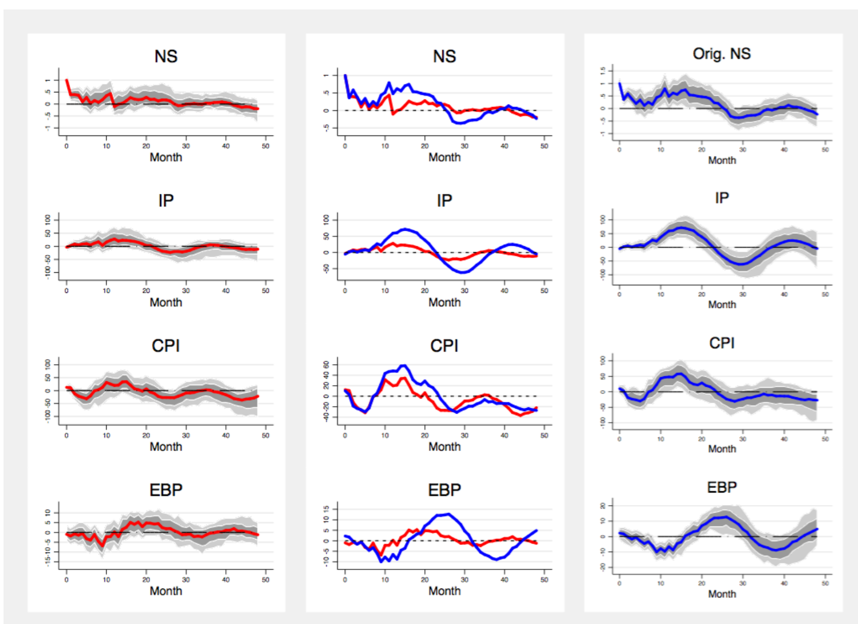
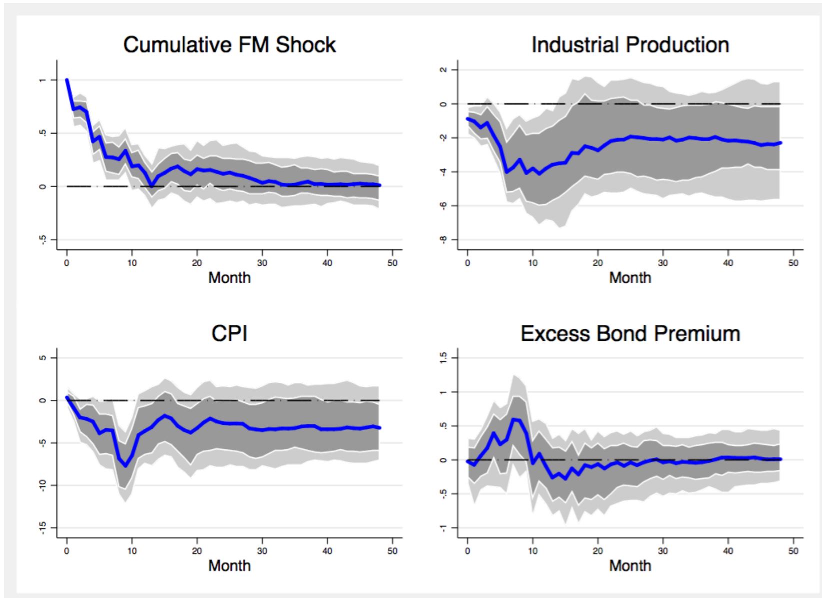
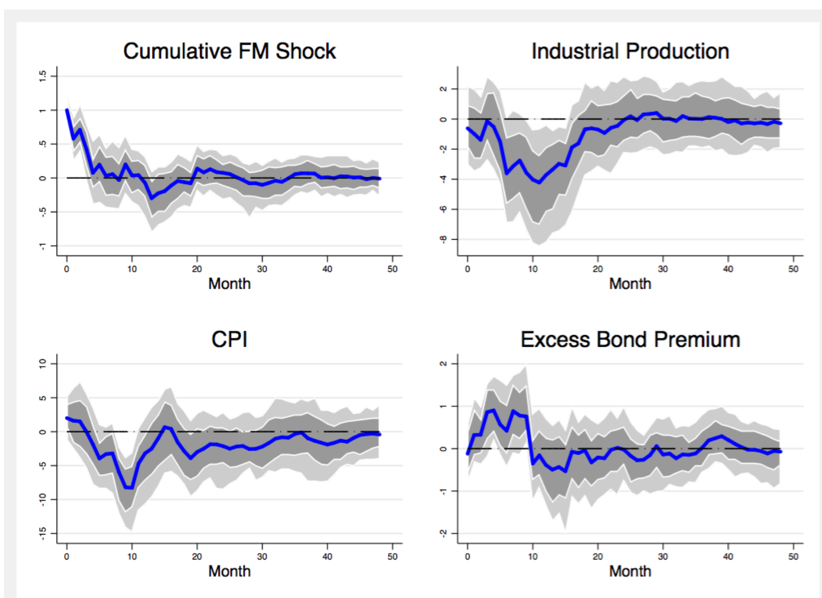


Figure A3: SVAR Impulse Responses with Simple Fama-Macbeth Shock

a. Sample period 1994m1-2017m12



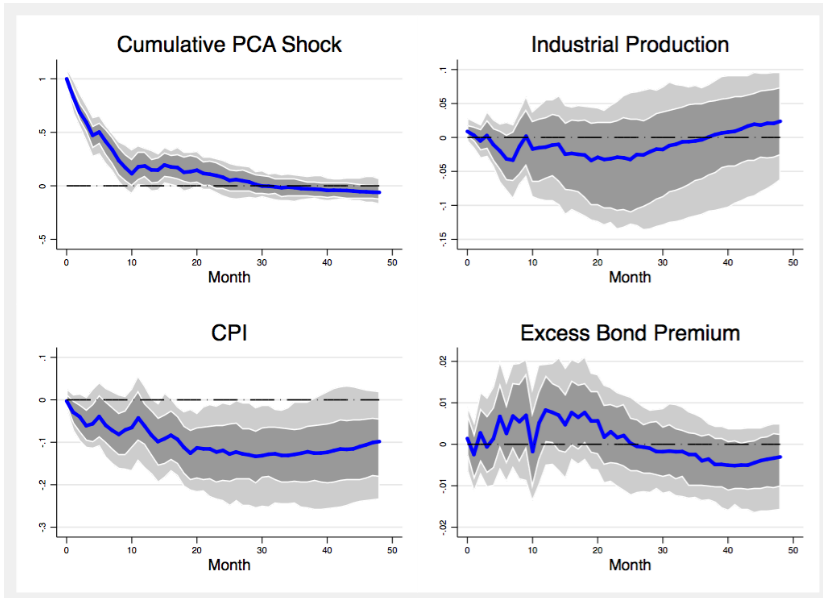
b. Sample period 2008m1-2017m12



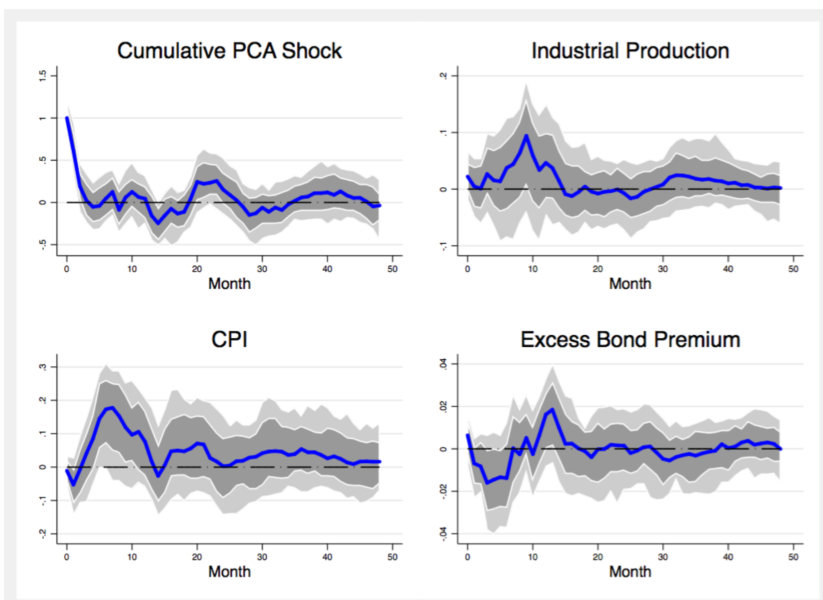
Note: Alternative BRW shock series is aligned from the Fama-Macbeth procedure without IDH. The IRFs are estimated as above.

Figure A4: SVAR Impulse Responses with PCA Shock

a. Sample period 1994m1-2017m12



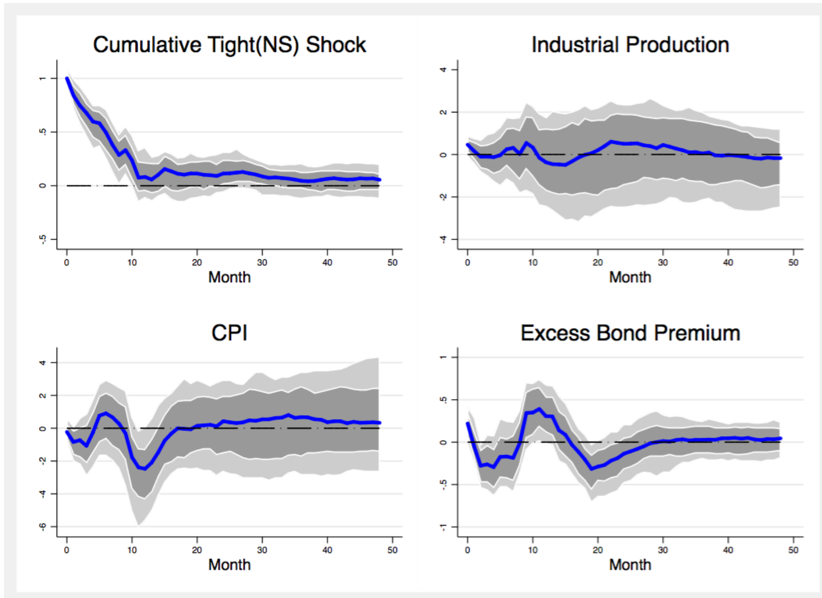
b. Sample period 2008m1-2017m12



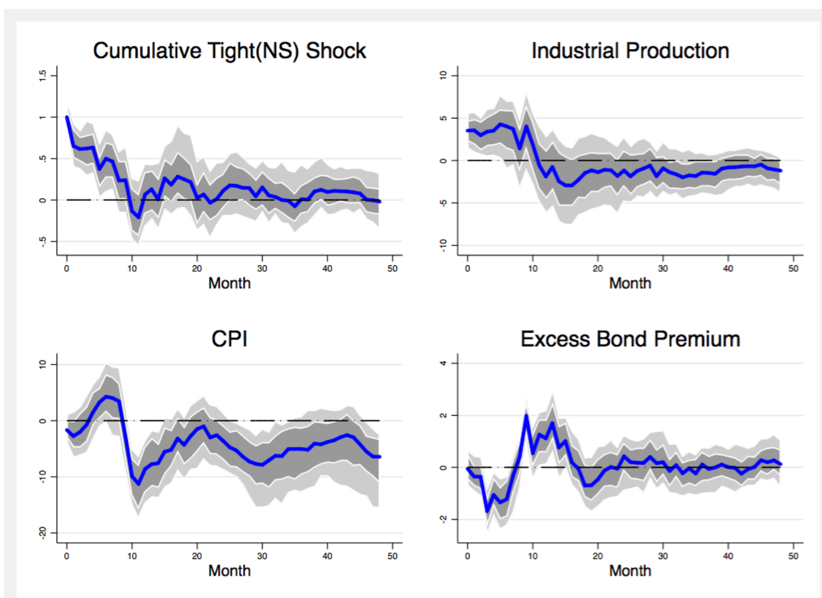
Note: The PCA shock is constructed from applying the Nakamura-Steinsson estimation procedure to our data: extracting the first principal component of all BRW outcome variables (daily changes of 1 to 30-year zero coupon rate around FOMC announcement days). The IRFs are estimated using the same approach as above.

Figure A5: SVAR Impulse Responses with Tight-window(NS data) Shock

a. Sample period 1994m1-2017m12



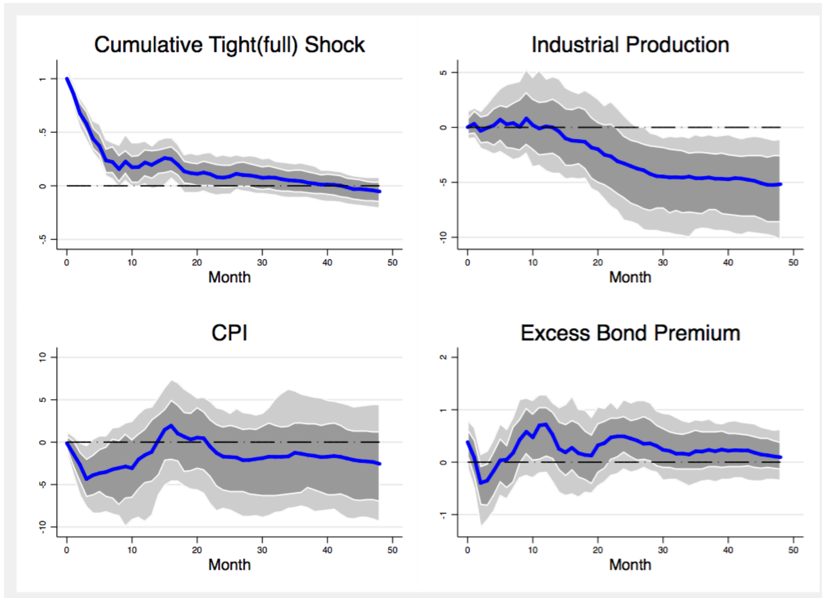
b. Sample period 2008m1-2017m12



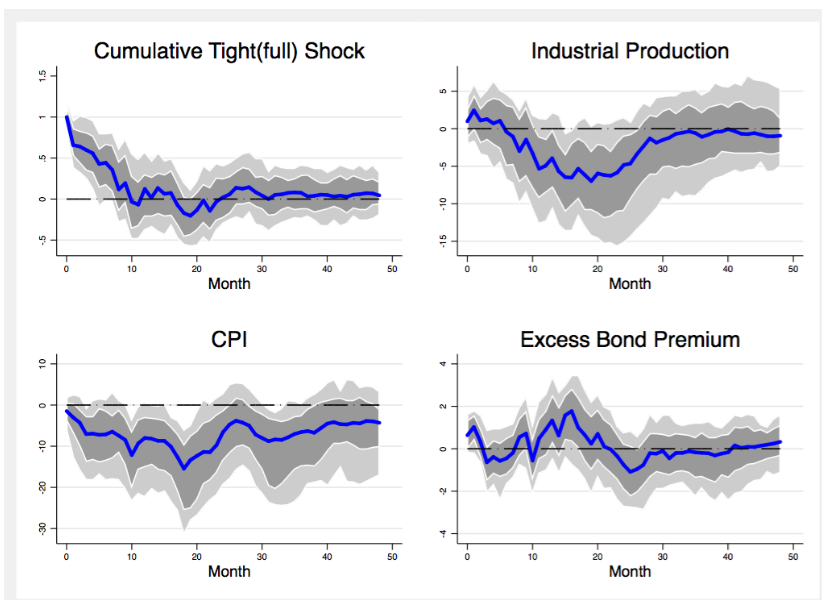
Note: The tight-window(NS data) shock is constructed from using the Nakamura-Steinsson (2018) data with our econometric procedure. The underlying data include the 30-minute changes of the current month Fed funds futures rate, the Fed funds futures rate immediately following the next FOMC meeting, and two, three, four quarter ahead euro dollar futures around the current FOMC announcement. The IRFs are estimated using the same approach as above.

Figure A6: SVAR Impulse Responses with Tight-window(Full data) Shock

a. Sample period 1994m1-2017m12



b. Sample period 2008m1-2017m12



Note: The tight-window shock is constructed using our econometric procedure with the Nakamura-Steinsson (2018) data plus some long term interest rate data.: the 30-minute changes of the current month Fed funds futures rate, the Fed funds futures rate immediately following the next FOMC meeting, the 1-8 quarter ahead euro dollar futures, the 3-, 6 month, and 2-, 5-, 10-, 30-year interest rates around FOMC announcements. IRFs are estimated using the same approach as above.