Market Confidence and Monetary Policy*

Nina Boyarchenko
Federal Reserve Bank of New York
nina.boyarchenko@ny.frb.org

Valentin Haddad
Princeton University
vhaddad@princeton.edu

Matthew C. Plosser
Federal Reserve Bank of New York
matthew.plosser@ny.frb.org

March 14, 2016

Abstract

We study the ability of monetary policy announcements to influence interest rates at all maturities as well as asset prices more broadly. Using a novel methodology, we discover two distinct monetary policy shocks. The primary factor is a long rate factor, which we find is related to aggregate uncertainty; the second factor is a traditional Fed funds short-rate shock. We show that the two factors have disparate effects on risky asset returns: the long rate is positively related to the aggregate risk premium while the Fed funds rate is negatively related to inflation expectations and positively related to the value premium. Our results highlight that monetary policy announcements influence market confidence even when there is no change in the stance of monetary policy.

*The views expressed in this paper are those of the authors and do not necessarily reflect the position of the Federal Reserve Bank of New York or the Federal Reserve System. The authors thank Peter Hansen, Matthew Yeaton, Sooji Kim, and Sam Stern for valuable research assistance, and Tobias Adrian and Richard Crump for comments on earlier drafts of the paper.

1 Introduction

Modern monetary policy relies heavily on communication as a supplementary policy tool. Indeed, in a presentation at the Federal Reserve's 2001 Jackson Hole conference, Woodford (2001) notes that,

> [S]uccessful monetary policy is not so much a matter of effective control of overnight interest rates... as of affecting...the evolution of market *expectations*...

While central banks often have direct control only over a single interest rate, the mandate to support a well-functioning economy necessitates the ability to influence asset prices and interest rates at all maturities. The recent experience of major central banks has revealed that monetary policy announcements have the potential to impact long rates in a low rate environment (see e.g. Wright, 2012). In this paper, we evaluate whether monetary policy announcements play this role beyond the most recent period of unconventional policy by studying changes in asset prices around scheduled monetary policy announcements.

The key empirical challenge in identifying the transmission of monetary policy shocks to asset prices is that interest rate changes are not exogenous to the state of the economy, since monetary policy is set in response to changes in economic conditions. The previous literature has pursued three approaches to overcoming this endogeneity problem: structural vector autoregressions (VARs) (e.g. Christiano et al., 1999), using changes in interest rates orthogonal to the information contained in internal Fed forecasts (Romer and Romer, 2004), and identification using high frequency changes to interest rates around FOMC announcements (e.g. Kuttner, 2001; Nakamura and Steinsson, 2013).

This paper uses a novel methodology that explicitly allows for multiple monetary policy shocks. We construct the monetary policy shocks as principal components of the variation in yields attributable to policy announcements. This is an extension of the Rigobon (2003) heteroskedasticity approach, where instead of focusing on a single interest rate, we consider the changes to the variance-covariance matrix of the changes to the entire U. S. Treasury nominal yield curve around scheduled Federal Open Market Committee (FOMC) meetings. The identification assumption is that the difference between the covariance matrix in the benchmark periods and the covariance matrix on announcement dates is the covariance matrix of monetary policy shocks. This approach allows us to jointly estimate multiple policy shocks by inferring information from changes to the whole yield curve. Instead, previous studies have focused on a single policy factor related to changes at shorter maturities (see e.g. Gürkaynak et al., 2005a; Nakamura and Steinsson, 2013), and thus confound the short rate factor with an "expected path" factor.

We identify two distinct monetary policy factors. The primary factor, which explains 59 percent of the variation in yields attributable to announcements, is related to level of longer term yields. We find that this factor correlates strongly with measures of uncertainty, suggesting that monetary policy announcements affect longer term risk premia, rather than long-run expectations of the policy rate. The second factor is the more traditional short rate factor, and explains a further 34 percent of the variation in yields attributable to announcements. Importantly, these two factors are distinct in the data, suggesting that movements in the long-rate related to uncertainty are separate from the classic focus on near short rates determined by the stance of traditional monetary policy.

The monetary policy shocks we identify have large and distinct impacts on the returns to other risky assets in the economy. To estimate the correlation of the policy factors with other assets, we compute the differential covariance of the factors with asset returns on announcement days versus non-announcement days. In the inflation-protected Treasury market (TIPS), we find that the first factor affects the entire TIPS yield curve, with a higher correlation between the factor and changes to TIPS yields than that between the factor and nominal yields. This suggests that changes to inflation expectations alone are insufficient to explain the impact of the long-term monetary factor and that the covariance is better explained by changes in the risk premium or in expectations of the future path of real rates. Fed funds shocks, on the other hand, have little impact on real interest rates but do lower inflation expectations. This is consistent with a negative lead-lag relationship between consumption growth and realized inflation (e.g. Stock and Watson, 1999).

We argue that the long-term monetary factor impacts aggregate uncertainty and not necessarily expectations of the future path of interest rates. We find that the long-term factor is significantly positively correlated with changes to various measures of aggregate uncertainty, including the one month option-implied volatility for S&P 500 (VIX), S&P 100 (VXO) and U. S. interest rate swaps (SMOVE). Thus, the long-term monetary policy shock impacts aggregate volatility and should command a risk premium.

Turning next to equity markets, we find that the long rate shock is negatively related to the aggregate market return on announcement days, but positively related on other days. This is consistent with FOMC communications impacting risk premia in the economy by coordinating beliefs about economic uncertainty, as outlined in section 2, as well as decreased firm-level investment after shocks to macroeconomic uncertainty (see e.g. Lucas and Prescott, 1971; Bloom et al., 2007; Bloom, 2009). The Fed funds shock, on the other hand, does not covary significantly with aggregate market returns neither on announcement nor non-announcement days. Instead, the Fed funds shock is negatively related to the "value" portfolio factor (HML) on non-announcement days and positively related on announcement days, whereas the long-rate factor is negatively related in both periods. These results emphasize that the two monetary policy shocks are distinct in their economic impact, as stocks with disparate risk characteristics react differently to the two shocks.

The two policy shocks also have distinct effects on real activity. We find that the long-rate shock is positively correlated with the trade-weighted U. S. exchange rate, commodity risk premia, credit risk premia and physical currency amounts. An unexpected increase in the Fed funds rate, on the other hand, tightens financial conditions for both households and financial corporations and decreasing money supply.

More generally, innovations in the long-rate factor appear to be orthogonal to innovations in the short-rate factor, suggesting that the mechanism by which the FOMC impacts the Fed funds rate is different from the one through which the FOMC impacts the long rate. The strong relation between the long rate factor and aggregate risk premia on announcement days reflects the ability of monetary policy to influence real discount rates in the long term and, as such, should be discussed explicitly as a potential policy tool.

1.1 Related Literature

This paper follows in the long tradition of using an event study approach to identifying monetary policy shocks and their impact on risky asset returns in the economy. Rudebusch (1998) shows that monetary policy shocks identified as surprise changes in the Fed funds rate around FOMC meetings have little correlation with shocks identified from a structural VAR approach. Kuttner (2001) and Bernanke and Kuttner (2005) use the Rudebusch shock identification to measure the reaction of bonds and of the aggregate stock market to policy shocks, respectively. Our paper takes a different approach to identification of monetary policy shocks and, instead of using changes to the price of the front contract in Fed funds futures, uses differential changes to the yield curve on announcement days to identify the shock.

As in our paper, Gürkaynak et al. (2005a) argue for a multi-factor structure for monetary policy surprises. Gürkaynak et al. use unexpected changes in Fed funds rate and Eurodollar futures with maturity up to a year measure a Fed funds shock and a "path" shock. We use, instead, changes to yields on Treasuries of up to ten years in maturity to identify two shocks. We show that our long-rate shock is related to changes in risk premia, rather than the expected path of future monetary policy. In that respect, the approach in this paper is closer to that taken by Hanson and Stein (2015), who measure the response of term premia to changes in the yield on the two year nominal bond on announcement days. Their identification strategy assumes implicitly, however, that the only changes to the yield curve occur around FOMC announcement dates. We avoid making this assumption by using the Rigobon (2003) heteroskedasticity approach. As argued in Rigobon and Sack (2004), this identification approach requires a much weaker set of assumptions to estimate both the policy shock and the response of asset prices to changes in monetary policy than the traditional event study approach.

More recent work uses high frequency changes in interest rates over a 30-minute window surrounding an FOMC announcement to measure monetary policy shocks. Nakamura and Steinsson (2013) perform a heteroskedasticity-corrected version of the high frequency exercise in Gürkaynak et al. (2005a), and use the results to identify parameters in a small-scale New-Keynesian model. Gertler and Karadi (2015) embed the monetary policy shocks identified using this methodology into a structural VAR model of the economy, and find that small shocks to the Fed funds rate can lead to large movements in credit costs, driven both by changes in term premia and credit spreads. Similarly, Gilchrist et al. (2015) use high frequency changes to the yield on the two year Treasury to measure policy surprise in the zero lower bound (ZLB) period. Notably, Gilchrist et al. also use information contained in changes to the yield on the ten year Treasury about policy surprise but concentrate on measuring the response to announcements of unconventional monetary policy. Instead, we show that monetary policy announcements affect long rates even outside the ZLB period. Further, we focus on two day changes because it natural to expect a more gradual adjustment process of longer maturity yields to monetary policy shocks to the extent that these instruments are less liquid than their shorter maturity counterparts. Our results are qualitatively similar if we instead measure changes over the one day interval preceding the FOMC meeting.

2 Motivating Example

In this section, we outline a simple conceptual framework for how news about the economic uncertainty is incorporated into asset prices.[1] To focus on the learning mechanism, we consider a two period model, with the FOMC announcement released at date 0 and consumption occurring at date 1. We assume aggregate consumption at date 1 is log-normally distributed

$$\log C \sim \mathcal{N}\left(\mu_s, \sigma_s^2\right),$$

where $s = 1, 2$ indexes an aggregate state of the economy. If state $s = 1$ is realized, the economy is in an expansionary phase; otherwise, the economy is in a recession. We assume that $\mu_1 \geq \mu_2$, and $\sigma_1 < \sigma_2$, so that recessions are associated with lower expected consumption growth and higher uncertainty about consumption growth. The representative household in

[1] A large literature studies the relation between risk premia, uncertainty and agents beliefs, starting with Miller (1977). In this section, we provide an illustration of the mechanism by which monetary policy announcements may affect market-based measures of uncertainty and risk premia jointly.

the economy evaluates date 1 consumption according to CRRA utility

$$u(C) = \frac{C^{1-\gamma}}{1-\gamma},$$

so that the equilibrium pricing kernel in the economy is given by

$$M = \frac{C^{-\gamma}}{\mathbb{E}\left[C^{-\gamma}\right]} = \frac{\exp\left\{-\gamma \log C\right\}}{\mathbb{E}\left[\exp\left\{-\gamma \log C\right\}\right]}.$$

The representative household starts date 0 with prior beliefs $\vec{\pi}_{-1}$ about the state of the economy, with $\pi_{-1,j} = \mathbb{P}(s=j)$, for $j = 1, 2$. At date 0, the FOMC releases a public signal about the state of the economy y, with the distribution of y conditional on state s realizing given by $f(y|\, s)$. The representative household updates its beliefs about the state of the economy after observing the public signal, so that its posterior belief vector $\vec{\pi}_0$ is given by

$$\pi_{0,j} = \frac{f(y|\, s=j)\, \pi_{-1,j}}{\sum_k f(y|\, s=k)\, \pi_{-1,k}}.$$

We assume that the public signal y is truthful and informative, so that, if the true state of the economy is j, then $\pi_{0,j} > \pi_{-1,j}$.

Given a vector of beliefs $\vec{\pi}_t$, $t = -1, 0$, the time t price of a claim to a share of consumption at date 2 is given by

$$P_t = \mathbb{E}\left[MC|\, \vec{\pi}_t\right] = \sum_{k=1,2} \pi_{t,k} \frac{\exp\left\{(1-\gamma)\mu_k + \frac{(1-\gamma)^2}{2}\sigma_k^2\right\}}{\sum_{j=1,2} \pi_{t,j} \exp\left\{-\gamma\mu_j + \frac{\gamma^2}{2}\sigma_j^2\right\}}.$$

Since $\gamma > 0$, the price of a claim to consumption is decreasing in the uncertainty about consumption growth given the state s. Thus, signals that increase the probability of a recession decrease the price of a claim to consumption, increasing the risk premium. This produces a positive correlation between risk premia in the economy and the representative household's subjective beliefs $\bar{\sigma}_t = \sum_{k=1,2} \pi_{t,k}\sigma_k$ about aggregate uncertainty. Figure 1 illustrates this mechanism: the price P_t of a claim to consumption is lower when the probability of recession is higher (panel 1(a)), and the expected volatility $\bar{\sigma}_t$ is higher (panel 1(b)).

In the remainder of the paper, we analyze the market reaction to monetary policy announcements and provide empirical support for this role for monetary policy. However, we do not take a stand on the source of the signal, y, disclosed by the FOMC. The information might result from superior ability to forecast economic conditions (as in e.g. Romer and Romer, 2004) but also more simply result from superior information about the future stance of monetary policy. A last set of explanations could come from an ability to coordinate decisions (see e.g. Morris and Shin, 2002).

3 Estimation of Monetary Policy Announcement Shocks

To identify the impact of policy announcements we compare changes in yields on announcement days to those on non-announcement days. In this section we describe our methodology, outline the data, and summarize our estimates for announcement day shocks.

3.1 Empirical Model

Our empirical methodology is in the spirit of Rigobon and Sack (2003). We decompose changes in yields into two independent components: a monetary policy surprise component, ε, and a non-policy component, ν. The key identification assumption is that monetary policy surprises occur only on policy announcement dates. Hence, a change in yield of maturity n

at time t can be written as the sum of these two components,

$$\Delta y_t^{(n)} = y_t^{(n)} - y_{t-1}^{(n)} = \nu_t^{(n)} + \varepsilon_t^{(n)}. \tag{1}$$

Given independence between the two components, the variance of the change in yield can be written as the sum of the variance of the policy component, σ_ν, and the non-policy component, σ_ε.[2]

We cannot directly observe the policy component on announcement days because there is also an unobserved non-policy component; however, we can use variances estimates to recover the variance of the policy component. Assuming that policy component is zero on non-announcement days, the announcement and non-announcement day variances are given by:

$$\text{Non-announcement: } \sigma_y^{NA} = \sigma_\nu$$
$$\text{Announcement: } \sigma_y^{A} = \sigma_\nu + \sigma_\varepsilon.$$

If the variances are constant over time, we can subtract the non-announcement variance, σ_y^{NA}, from the announcement variance, σ_y^A, to recover the policy variance, σ_ε.

We can extend this logic to analyze a vector of yields at various maturities. Subtracting the covariance matrix on non-announcement days from that of announcement days results in a policy covariance matrix, Σ_ε. The policy matrix is symmetric and composed of variances on the diagonal and covariances elsewhere.

We use principal component analysis (PCA) to recover the structure of the policy-related

[2] In order to focus on innovations, we assume all variables are demeaned. In unreported analysis we consider a broad array of ways to capture changes in means.

shocks from the policy covariance matrix. PCA identifies linear combinations, or components, of yields and ranks them in order of decreasing importance in explaining overall variance. Based on the contribution of each linear combination to total variance we choose those factors that are most important in explaining the variance of the policy shocks.

While PCA reveals the relevant linear combinations, it cannot reveal the actual level of a factor. The weighted sum of yield changes contains the policy component but we cannot eliminate the non-policy component, $\nu_{i,t}$. Hence, using the PCA weights for component i at time t does not identify the level of the policy shock

$$\tilde{f}_{i,t} = \sum_n \omega_i^{(n)} \Delta y_t^{(n)} = f_{i,t} + \sum_n \omega_i^{(n)} \nu_t^{(n)}. \tag{2}$$

We will account for this when we examine the relation between our policy factors and other assets.

3.2 Data

We construct a daily yield curve from various sources. We obtain zero coupon treasury yields constructed according to Gürkaynak et al. (2007) from the Federal Reserve Board's H.15 data release. These yields are interpolated from closing market bid yields on actively traded Treasury securities in the over-the-counter market. We consider yields at each annual horizon from one year to ten years. For yields of maturity less than one year, we augment the Gürkaynak et al. yield curve with 3-month and 6-month constant maturity Treasury yields from the H.15 data release.The yields at horizons less than a year are available from 1982 onward and the horizons of more than a year from 1961.

The primary policy rate of the Federal Reserve is the overnight Federal Funds rate.

Therefore, in addition to the yield curve, we consider changes in expectations of the Federal Funds target. The Fed Funds rate reflects the average rate at which depository institutions lend funds maintained at the Federal Reserve to other depository institutions overnight. In order to measure changes in market expectations, we use Federal Funds futures contracts traded on the Chicago Mercantile Exchange (CME). Federal Funds futures contracts began trading in 1988. They are written for a specific month and trade at 100 minus the average daily Fed Funds overnight rate for the delivery month. We slightly modify the methodology outlined in Kuttner (2001) to calculate a time-series of daily Fed Funds rate expectations for the front month (i.e. current month).

Our analysis focuses on two-day changes in yields to better capture developments in less liquid, longer horizon bonds. Intra-day evidence, even from shorter maturities, suggests that the market takes time to resolve the impact of FOMC statements (Gürkaynak et al., 2005b). In fact, when we plot the variance of yields over event time, we find there is meaningful variance on the day after announcements which are necessary to capture the response of longer maturity bonds (Figure 3). We ultimately relate yield changes to equity prices – two-day changes also are better for capturing the relation between yields and movements in these stock prices. Our methodology contrasts with higher frequency identification methods (e.g. Nakamura and Steinsson, 2013), which look at intra-day trades but focus solely on shorter, more liquid maturities and classic shocks to changes in the Fed Funds rate.

We compile a list of monetary policy announcement days starting in 1994, the first year that the Fed began to explicitly announce FOMC decisions. Our primary sample runs through 2007 and includes 113 announcement days; post-2007 there are a myriad of atypical policy announcements. Given we are focused on two-day changes in yields, we designate the two-day change following the day of an announcement as "announcement" changes and two-day changes that exclude an announcement as "non-announcement" changes. We exclude two-day changes that fall on announcement days as they likely include some of the impact from a policy shock. Figure 2 illustrates the timing of the "announcement" and "non-announcement" periods.

3.3 Policy Shocks

An implication of our empirical methodology is that yields are more volatile on announcement days versus non-announcement days. If we compare the variance in yields on announcement days to non-announcement days, Figure 4(a), we can see that announcement days in fact are more volatile, consistent with the empirical model. The raw variances of announcement days are above non-announcement days at all maturities from the Fed Funds forward rate to the ten-year treasury yield.

Subtracting the variance on non-announcement days from announcement days we recover the variance of the policy shock, Figure 4(b). The variance of the policy shock is larger than the non-policy shock at the short-end of the yield curve ($<$ 3 months). However, at the longer end the policy component is less volatile than the non-policy component. Hence, the importance of Federal Reserve announcements is preeminent for explaining short rates, but longer rates experience changes that are related to other factors such as the long-term growth of the economy, fiscal policy and the relative attractiveness of dollar denominated investments.

PCA analysis reveals that the variance in the policy component of announcement day yields is primarily explained by two principal components. Figure 5 illustrates the contribution of the first three principal components to the policy variance across maturities. The first component is particularly important for longer maturities and explains almost all the variance at the three and four year horizons. The second component primarily explains the shortest maturities, particularly the Fed Funds surprise (0 maturity), but it also has a modest impact on maturities greater than 6 years. In total, the first two components explain 93% of total policy variance with the first component explaining 59% and the second 34%.

There are several takeaways from examining the loadings for the first two principal components, Figure 6. The first component is strikingly consistent across maturities, but loads very little on the Fed Funds rate. The second component heavily loads on the Fed Funds rate, but falls to zero at the three year maturity before turning slightly negative at longer maturities. The level effect described by the first component is consistent with the the impact on term premia documented by Hanson and Stein (2015), whereas the downward sloping short rate is consistent with the classic tool of monetary policy, the Fed Funds rate, which dissipates at longer maturities in the yield curve. It is important to emphasize these two mechanisms are orthogonal in the data. The policy shock to long rates is not correlated with surprise changes in the short rate. Therefore, the influence of monetary policy must be comprised of more than just surprise changes in the Fed Funds rate.

3.3.1 Robustness

We consider several alternative identification assumptions to verify the robustness of our baseline results. The first considers a different measure for non-policy variance, σ_ν. Non-announcement days may contain some news about Fed policy. FOMC members can deliver public speeches or testify to Congress and these communications might reveal information about future monetary policy announcements. To eliminate the risk of non-announcement day news, we exploit the Federal Reserve Systems communication blackout period. For the seven-days prior to a scheduled meeting there is a strict blackout period during which the

Federal Reserve System and FOMC members do not communicate with the public. We use these dates to construct an alternative measure of policy variance.

Indeed, we find that yield volatility is lower on these days, Appendix Figure 10, especially at the longer maturities. As a result policy variance is larger than under the base specification. When we examine the PCA loadings, we find similar patterns as in our base specification. Figure 8 illustrates the PCA loadings based on non-announcement days (*Baseline*) and those based on blackout days (*Blackout*). The loadings have been rescaled so that they sum to one. The second component is almost exactly the same, whereas the first component is similar, but no longer exhibits elevated weights at the 3 and 6 month maturities.

The second robustness check considers a sub-set of announcement days. Some Fed announcements occur between meetings and may be in response to macroeconomic events. We can construct an alternative estimate of policy variance using the 108 announcements that coincide with scheduled FOMC meetings. The timing of these announcements is predetermined and are plausibly exogenous from other macroeconomic events. Figure 8 also displays component loadings for this set of PCA components (*Scheduled*). The first component closely matches the base specification. The second component trends downward from the short maturities, much like the base case, but slightly favors the 3-month maturity relative to the Fed Funds rate.

Finally, we consider the possibility that macroeconomic news unrelated to the FOMC is released on announcement days. We focus on the release of economic indicators related to the labor market and price levels: non-farm payrolls, the unemployment rate, initial unemployment claims, the produce price index (PPI), and the preliminary ("advance") GDP. If any of these releases falls on an FOMC announcement day we exclude the date from the sample. We drop twelve announcement days during the period of interest. Figure 8 summarizes the component loadings for FOMC only announcements. Dropping the FOMC announcement days that coincide with other macroeconomic announcements does not materially impact the factor loadings.

Thus, the policy factors we uncover are robust to other definitions of both the announcement and the control periods. In unreported results, we also estimated the policy factors using one day changes in yields, rather than two day changes, and found the results to be qualitatively similar. We use the baseline specification, with policy factors identified using two day changes and the full set of FOMC announcement days, in the remainder of our analysis.

4 Characterization of Announcement Shocks

We characterize the impact of the policy shocks estimated in section 3 on the economy by estimating their covariance with other asset prices. We extend our methodology to measure announcement-day covariance between policy shocks and asset prices, such as yields on inflation-linked bonds, equity returns, and market-implied measures of aggregate uncertainty.

4.1 Measuring the Covariance of Shocks With Other Assets

We extend the methodology of section 3 to estimate the covariance between asset returns and policy shocks. As with nominal yields, we allow for non-policy variance on announcement days in addition to the impact of the policy shocks. That is, we assume that returns (or changes in yields) for asset i at time t are the sum of a non-policy component, $\xi_{i,t}$, and the impact of our two policy factors:

$$\text{Non-announcement: } R_{i,t} = \xi_{i,t}$$
$$\text{Announcement: } R_{i,t} = \xi_{i,t} + \beta_{1,i} f_{1,t} + \beta_{2,i} f_{2,t}.$$

The non-policy component is independent of the policy factors but can covary with the non-policy component of nominal yields, so that $\text{cov}(\xi_{i,t}, \nu_t) \neq 0$.

Recall that the PCA weights reveal the policy factor polluted by non-policy shocks (equation (2)). Hence, the covariance between the asset of interest, i, and the observed factor, j, contains a non-policy component as well as the impact of the policy shock,

$$\text{Non-announcement:} \quad \text{cov}\left(R_{i,t}, \tilde{f}_{j,t} | \text{NA}\right) = \text{cov}\left(\xi_{i,t}, \sum_n \omega_j^{(n)} \nu_t^{(n)}\right)$$
$$\text{Announcement:} \quad \text{cov}\left(R_{i,t}, \tilde{f}_{j,t} | \text{A}\right) = \text{cov}\left(\xi_{i,t}, \sum_n \omega_j^{(n)} \nu_t^{(n)}\right) + \beta_{j,i} \text{var}\left(f_{j,t}\right).$$

Assuming the covariance between non-policy shocks is stable over time, we can derive the impact of the policy factor by differencing between the announcement day and non-announcement day estimates.

These covariance can be calculated directly, or operationalized in the form of a regression in which we estimate the differential response of an asset on announcement days versus non-announcement days:

$$R_{i,t} = \alpha_{0,i} + \alpha_{1,i}\mathbb{A}_t + \gamma_{1,i}\Delta y_t^{*,1} + \gamma_{2,i}\Delta y_t^{*,2} + \beta_{1,i}\left(\Delta y_t^{*,1}\mathbb{A}_t\right) + \beta_{2,i}\left(\Delta y_t^{*,2}\mathbb{A}_t\right) + \varepsilon_{i,t}, \tag{3}$$

where $\mathbb{A}$ is a dummy variable equal to one on announcement days and $\Delta y_t^{*,1}$ and $\Delta y_t^{*,2}$ are the two yields spanning the monetary policy shocks. The gamma terms capture the covariance with non-policy shocks and the beta terms identify the variation with the policy shocks.[3]

4.2 Components of Yields

We begin characterizing the two policy shocks by considering whether the changes to the nominal yield curve on announcement days reflect changes in inflation expectations, the future path of the target rate or the risk premium. To disentangle shocks to inflations expectations from the other two effects, we compare the policy factors' covariance with nominal and real yields.

Beginning in 1999, Treasury-Inflation Protected Securities (TIPS) provide a measure of yields absent expectations of future inflation. We obtain zero coupon TIPS yields constructed according to the Gürkaynak et al. (2010) methodology, which interpolates a discount curve from closing market bid yields on actively traded TIPS securities in the over-the-counter secondary market. We focus on yields with maturities between 5 and 10 years as TIPS have historically been issued at 5, 7, and 10 year maturities. If the movement in nominal yields is primarily caused by changes in inflation expectations, we expect covariance with real rates to be close to zero. However, if real rates are further from zero than the corresponding nominal rate, inflation expectations are not the determining factor and the covariance is better explained by changes in the risk premium or the expected path of interest rates.

Figure 7 summarizes the covariance of the factors with nominal and real yields. Similar to the principal component weights, the first factor affects the entire yield curve, absent

[3] We scale the vector of spanning yields $\left[\Delta y_t^{*,1}\ \Delta y_t^{*,2}\right]'$ by the non-announcement covariance matrix on non-announcement dates, and by the announcement covariance matrix on announcement dates, so that the regressors are unit variance and uncorrelated with each other. Therefore, the regression coefficient reveals the difference in covariance between announcement and non-announcement.

the short rate. The covariance between the first factor and real rates is greater than the covariance for the same factor and nominal rates in the 5-10 year maturity range, suggesting that inflation expectations are insufficient to explain the change in rates. The second factor is characterized by a strong covariance with changes in the Fed Funds rate and a negative covariance with longer nominal rates. However the real rates are close to zero, suggesting that at the long end of the yield curve the second factor has little impact on real rates but does lower inflation expectations.[4]

We can estimate the statistical significance of these differences using the regression specification summarized in Equation (3). The interaction between the factors and announcements days reveal the covariance between yields and the policy factors. Table 1 reports the estimated coefficients, together with Newey-West t-statistics. We consider five- and ten-year nominal yields (Columns 1 and 2) and breakeven rates, i.e. inflation expectations, (Columns 3 and 4). Consistent with the figure, the first factor positively covaries with nominal yields at both maturities and the difference is statistically significant at the 1% level. However there is no discernible relation with breakevens in Columns 3 and 4 which suggests that the first factor impacts real yields.

Unlike the first factor, the covariance of the second factor with longer maturities is difficult to distinguish from a change in inflation expectations. Factor two is negatively related with nominal yields and this relation is larger and more significant at the longer maturity (Column 4). The covariance with breakevens is of similar magnitude to those of nominal yields, however the statistical difference from non-announcement covariances is not statistically significant relative to standard levels (the p-values are approximately 14%). The negative relationship between breakevens and factor two is consistent with a classic inverse

[4]See Figure 9 for estimates of the covariance with nominal yields for the full sample periods.

relationship between Fed funds shocks and long-term inflation expectations.

4.3 Aggregate Uncertainty

The absence of a relation between inflation break-even rates and the long-term policy shock suggests that this shock affects either expectations of the future path of interest rates or aggregate uncertainty in the economy and, hence, the risk premia earned by risky assets. From Table 1, we observe that the first policy shock has similar impact on both five- and ten-year nominal yields, making the risk premium channel a more plausible explanation.

The motivating example of section 2 provides a mechanism – beliefs about aggregate uncertainty – through which monetary policy can affect risk premia. Columns 4-7 of Table 2 test this mechanism by estimating Equation 3 via OLS using different measures of changes to aggregate volatility as the dependent variable: one month options-implied volatility of S&P 500 returns (VIX, Column 4), one month options-implied volatility of S&P 100 returns (VXO, Column 5), one month options-implied volatility of interest rate swaps (SMOVE, Column 6), and the Baker et al. (2013) index of political uncertainty (Column 7). For all these measures, we find that the long-term policy shock is positively correlated with changes in volatility on announcement days. Thus, the long-term policy shock leads to increases in aggregate uncertainty on announcement days. The second factor is negatively associated with changes in volatility, but the coefficient is not statistically significant.

4.4 Stock market response

Columns 4-7 of Table 2 show that the long-term policy factor is positively correlated with changes to measures of aggregate uncertainty on announcement days. We now investigate whether this uncertainty commands a risk premium. Time variation in the risk premium can act as a form of policy transmission as the risk premium affects investment choices (see e.g. Lucas and Prescott, 1971; Bloom et al., 2007; Bloom, 2009) and the value of household portfolios (i.e. the "wealth effect") (see e.g. Sandmo, 1970; Grossman and Zhou, 1996; Basak, 1995).

To determine the covariance of our factors with the risk premium, we estimate Equation (3) via OLS using the excess market return as the dependent variable. Column 1 of Table 2 summarizes our findings. Consistent with the results for aggregate uncertainty, we find that the long-term policy factor is negatively correlated with market returns at the 5% significance level. In other words, the equity risk premium is increasing with the policy factor characterized by longer rates. Thus, long rates on announcements days behave in a manner consistent with changing uncertainty which impacts the risk premium for both fixed income (as reflected in yields on both nominal and real Treasury securities) and equity instruments. In contrast, on non-announcement days the long-term factor is *positively* correlated with market returns – emphasizing that monetary policy announcements create a unique risk in equities.

The second factor is also negatively associated with the market return, but the coefficient is not statistically significant.[5]

In addition to the impacting the aggregate risk premium, monetary policy may have a redistributive effect on firms with different risk profiles. To test the redistributive effect of monetary policy, we consider the correlation between HML and SMB factor-mimicking portfolio returns and the two policy shocks. The HML portfolio measures the performance

[5] In contrast, Bernanke and Kuttner (2005) find that the Fed funds shock estimated from changes to the front month Fed funds futures contract around announcement days is significantly, positively related to aggregate market risk premia. Our results show, however, that controlling for non-policy variance and including the long-rate policy factor attenuates the statistical significance of the short-term factor in explaining aggregate market returns.

of value (high book-to-market, H) stocks relative to growth (low book-to-market, L) stocks, while the SMB portfolio measures the performance of small (S) stocks relative to large (B) stocks. These two factors are known to be important in pricing the cross-section of equity returns (see Fama and French, 1993, and the subsequent literature). Columns 2 and 3 of Table 2 summarize the results estimating (3) using HML and SMB returns as the dependent variables, respectively. We find that while neither policy is related in a statistically significant manner to SMB returns on announcement days, the Fed funds shock is significantly positively related to the return on HML on announcement days. That is, an unexpected increase in the Fed funds rate decreases the value premium. This is consistent with the monetary policy target rate having redistributive effects in the economy.

Figure 11 plots the differential covariance between the two policy factors on announcement days for the 25 equity portfolios sorted by size and book-to-market (FF25). Figure 11(a) shows that, consistent with the long-term policy shock impacting the aggregate risk premium, the covariance of the first policy factor on FOMC announcement days with the FF25 portfolio returns is uniformly higher than that on non-announcement days. In contrast, the Fed funds shock has an increase in correlation with the five value portfolios and a decrease in correlation with the five growth portfolios (Figure 11(b)), generating the positive differential covariance with the HML return we saw in Table 2.

5 Real Effects

The previous section measured the responses of financial variables to monetary policy shocks. We now turn to some measures of real activity, and examine the impact of policy shocks on real outcomes.

5.1 Measuring the covariance of shocks with low frequency variables

We now extend the methodology of section 4 to allow the affected variable to be measured at a lower frequency than monetary policy shocks. Suppose that we observe real outcomes m_t every τ days, so that we can only construct τ day changes

$$\Delta_\tau m_t \equiv m_t - m_{t-\tau} = \sum_{s=0}^{\tau-1} \Delta m_{t-j},$$

where we use Δm_t to denote one day changes. If we were able to observe the one day changes directly, similarly to the financial variables in section 4, we would represent Δm_t as the sum of a non-announcement and policy components

$$\begin{aligned} \text{Non-announcement: } \Delta_\tau m_t &= \xi_{m,t} \\ \text{Announcement: } \Delta_\tau m_t &= \xi_{m,t} + \beta_{1,m} f_{1,t} + \beta_{2,m} f_{2,t}. \end{aligned}$$

The non-policy component is independent of the policy factors but can covary with the non-policy component of nominal yields, so that $\text{cov}(\xi_{m,t}, \nu_t) \neq 0$. The τ day change can thus be represented as

$$\Delta_\tau m_t = \sum_{s=0}^{\tau-1} \left(\xi_{m,t-s} + \beta_{1,m} f_{1,t-s} \mathbb{A}_{t-s} + \beta_{2,m} f_{2,t-s} \mathbb{A}_{t-s} \right).$$

We now need to construct the low frequency version of the policy shocks. Using the same expansion of τ day changes of yields into one day yield changes, we can define the low frequency "contaminated" shocks as

$$\tilde{f}_{i,t}^{\tau} = \sum_{n} \omega_i^{(n)} \Delta_\tau y_t^{(n)} = \sum_{s=0}^{\tau-1} \left(f_{i,t-s} \mathbb{A}_{t-s} + \sum_{n} \omega_i^{(n)} \nu_{t-s}^{(n)} \right). \tag{4}$$

The covariance between the low frequency variable and the observed factor j is given by

$$\operatorname{cov}\left(\Delta_\tau m_t, \tilde{f}_{j,t}^{\tau}\right) = \operatorname{cov}\left(\sum_{s=0}^{\tau-1} \xi_{m,t-s}, \sum_{s=0,n}^{\tau-1} \omega_j^{(n)} \nu_{t-s}^{(n)} \right) + \beta_{j,m} \operatorname{var}\left(\sum_{s=0}^{\tau-1} f_{j,t-s} \mathbb{A}_{t-s} \right),$$

where we have allowed the policy factors to be serially correlated. Assuming the covariance between non-policy shocks is stable over time, we can derive the impact of policy shocks by differencing between estimates for periods containing different number of announcement days.

As with the returns, we operationalize estimating the covariance differencing by estimating the differential response of real activity on periods with different numbers of announcements:

$$\begin{aligned} \Delta_\tau m_t = {} & \alpha_{0,m} + \alpha_{1,m} \mathbb{N}_{t,\tau} + \sum_{l=0}^{2} \gamma_{1,m}^{l} \Delta_\tau y_{t-l}^{*,1} + \sum_{l=0}^{2} \gamma_{2,m}^{l} \Delta_\tau y_{t-l}^{*,2} \\ & + \sum_{l=0}^{2} \beta_{1,m}^{l} \left(\Delta_\tau y_{t-l}^{*,1} \mathbb{N}_{t-l,\tau} \right) + \sum_{l=0}^{2} \beta_{2,m}^{l} \left(\Delta_\tau y_{t-l}^{*,2} \mathbb{N}_{t-l,\tau} \right) + \sum_{l=1}^{2} d_l \Delta_\tau m_{t-l} + \varepsilon_{m,t}, \end{aligned} \tag{5}$$

where $\mathbb{N}$ is the indicator variable of the number of announcements over the period $t-\tau$ and t, and $\Delta_\tau y_t^{*,1}$ and $\Delta_\tau y_t^{*,2}$ are the two (re-scaled) yields spanning the low frequency monetary policy shocks. We allow lags of both the policy factors and the dependent variable in the specification to capture the potential lagged response of real activity to monetary policy shocks.

Appendix Table 6 estimates Equation (5) the covariance between Treasury yields and the policy factors using one week changes in yields. While we loose some statistical significance of the estimates, the results are broadly similar to those of Table 1 (which uses the baseline two day changes to identify the covariance). This suggests that the procedure described above performs reasonably well at a weekly frequency.

5.2 Other risky assets

We begin by studying the impact of the policy shocks on other risky assets. Column 1 of Table 3 estimates Equation (3) via OLS using the two day change in the trade-weighted U. S. dollar exchange rate relative to a basket of major currencies as the dependent variable.[6] We find that the long-term policy factor is positively correlated with the exchange rate at the 1% significance level. In other words, the U. S. dollar appreciates against other currencies with the policy factor characterized by longer rates, consistent with the safe-haven currency status of the U. S. dollar.

Columns (2)–(4) of Table 3 estimates Equation (3) via OLS using the returns on the S&P GSCI overall, energy and non-energy spot commodity indices as the dependent variable, respectively.[7] Consistent with the results for aggregate uncertainty and market returns, we find that the long-term factor is negatively correlated with commodity returns at the 5-10% significance level, so that the risk premium paid to commodity investors is increasing with

[6]The trade-weighted U. S. dollar exchange rate is a weighted average of the foreign exchange value of the U. S. dollar against a subset of the broad index currencies that circulate widely outside the country of issue. The major currencies index includes the Euro Area, Canada, Japan, United Kingdom, Switzerland, Australia, and Sweden. The index is higher when the U. S. dollar appreciates relative to other currencies.

[7]The S&P GSCI spot index is a world production-weighted average of prices on liquid commodity futures contracts. The spot index is built to be the equivalent of the S&P 500 equity index for commodities. The overall index consists of 24 futures on physical commodities across five sectors: energy, agriculture, livestock, industrial metals and precious metals. The non-energy subindex includes all futures not included in the energy subindex.

the policy factor characterized by longer rates. In contrast, on non-announcement days, the long-term factor is positively, but not statistically significantly, correlated with commodity returns, highlighting once again that monetary policy announcements create a unique source of risk for risky assets. The short rate factor is positively correlated with commodity returns on announcement days, but the correlation is statistically significant only for the non-energy subindex. That is, an unexpected increase in the Fed funds decreases the risk premium earned by investors in non-energy commodities. A potential explanation may be the signaling effect proposed by Romer and Romer (2004), with market participants interpreting an unexpected increase in the policy rate as a signal for the strength of the economy.

5.3 Credit conditions

Tables 2-3 show that the long-term policy factor is positively correlated with measures in market risk premia on announcement days for fixed income, equity, exchange rate and commodity markets. We now investigate whether this time variation in risk premia affects real outcomes. Table 4 summarizes estimates of Equation (5) via OLS using measures of credit conditions as the dependent variable.

Columns (1)-(4) test the impact of monetary policy shocks on measures of corporate credit conditions. The long-term factor is positively correlated with changes to the commercial paper rate paid by AA-rated non-financial issuers (Column 1), the commercial paper rate paid by AA-rated financial issuers (Column 2) and the yield on the Bank of America (Merrill Lynch) index of U. S. high yield bonds.[8] On announcement days, the long-term factor is even more positively correlated with these three measures of corporate borrowing costs, consistent

[8]The Bank of America (Merrill Lynch) U. S. High Yield Master II index includes U. S. dollar denominated corporate debt issued in the U. S. domestic market with credit rating below investment grade. The securities must have more than one year of remaining maturity, a fixed coupon schedule and a minimum outstanding amount of $100 million.

with the depreciation in response to the long-rate factor experienced by the other risky asset classes above.

The Fed funds shock is also positively correlated with commercial paper interest rates and the yield on high yield bonds. That is, an unexpected increase to the policy rate tightens credit conditions. Indeed, financial conditions as measured by the Chicago Fed National Financial Conditions Index tightens one week after a surprise increase to the Fed funds rate, and continues tightening the following week.[9] In contrast, in other weeks, changes to the NFCI are not statistically significantly correlated with the short rate factor.

Turning to credit conditions for the household sector, Column 5 shows that the spread between the Freddie Mac 30 year fixed rate mortgage interest rate and the 5 year yield narrows in response to the long-rate policy shocks, with the negative correlation becoming more negative in announcement weeks. This negative correlation may be the result of a decrease in the prepayment risk premium component of FRM rates. Columns 6-9 measure the impact of monetary policy shocks on the number of mortgage applications for purchases, refinancing, fixed rate and adjustable rate mortgages, respectively.[10] The number of mortgage applications decreases two weeks after a surprise increase to the Fed funds rate, suggesting that surprise increases in the policy rate have a tightening effect for household credit conditions as well. The number of mortgage applications is also negatively correlated to the lagged long-rate factor, with the correlation on announcement weeks marginally statistically significant at the 10% level for mortgages used to refinance. Combined with the negative

[9]The Chicago Fed National Financial Conditions Index (NFCI) is the first principal component of price and volume measures of financial activity, covering money markets, debt and equity markets, as well as both the traditional and the shadow banking system. The values of the index are standardized; positive values indicate a tightening in financial conditions relative to historical averages.

[10]Mortgage application data is from the weekly mortgage applications survey conducted by the Mortgage Bankers Association. The survey covers roughly 50% of all U. S. residential mortgage applications processed by mortgage banks, commercial banks and thrifts.

effect on the FRM spread, this implies that a surprise increase in the long rate decreases the cost of residential mortgages (relative to similar duration yields) and makes households less likely to refinance their existing mortgages.

5.4 Money aggregates

Table 5 tests whether monetary policy shocks affect money supply. Columns (1)-(3) show that the short-rate policy factor is negatively correlated with money stock (M1), the sum of money stock and wholesale funding (M2), physical currency, and wholesale funding (M2-M1).[11] That is, an unexpected increase in the Fed funds rate decreases the supply of U. S. dollars in the market in excess of the non-announcement week response. This is the traditional channel of monetary policy transmission. Interestingly, the supply of physical currency increases with unexpected changes to the long-rate, suggesting a possible safety premium for physical U. S. currency, which is consistent with the positive response on the trade-weighted exchange rate we found in Table 3.

6 Conclusion

The last twenty years have seen an evolution in how the Federal Reserve communicates monetary policy and the economic outlook. The first major changes happened in 1994, with the FOMC starting to announce its decision on the federal funds rate target in February and adding a description of the state of the economy and the rationale for policy action later in the year. In 1999, the FOMC began releasing a statement each meeting regardless

[11] M1 is the sum of physical U. S. currency in circulation outside of the U. S. Treasury and the Federal Reserve System and demand and checkable deposits. M2 is the sum of M1 and savings and time deposits, and balances in retail money market mutual funds.

of whether policy action was taken, and included its assessment of the "balance of risks" in the economy. More recently, starting in January 2012, statements at every other FOMC meeting incorporate a summary of economic projections of committee members.

In this paper, we demonstrate that throughout this period FOMC announcements impact the prices of risky assets via two distinct channels: directly by announcing changes in Fed funds targets and indirectly by influencing the risk premium. By describing the state of the economy and the rationale for policy action, the FOMC is able to affect economic agents' beliefs about aggregate uncertainty and, thus, the aggregate risk premium. In contrast, the target monetary policy rate impacts expectations about the future path of inflation and performs a redistributive role in the cross-section of equity returns.

The recent experience of monetary policy at the zero lower bound has reignited the debate on the efficacy of monetary policy to impact real outcomes. Our results provide evidence that monetary policy can have an impact – even at the zero lower bound – by acting through the aggregate risk premium. A positive innovation to the long-term policy shock increases aggregate uncertainty, increasing the risk compensation demanded by investors in both bond and equity markets. This channel is distinct from the effects of unconventional monetary policy, as it is present even before the start of the financial crisis in 2007. As the U.S. exits the zero lower bound, it is crucial to revisit the potential role of monetary policy at both the short- and long-end of the yield curve.

References

Baker, S. R., N. Bloom, and S. J. Davis (2013): "Measuring economic policy uncertainty," Research paper n. 13-02, Chicago Booth.

Basak, S. (1995): "A general equilibrium model of portfolio insurance," *Review of Financial Studies*, 8, 1059–1090.

Bernanke, B. S. and K. N. Kuttner (2005): "What Explains the Stock Market's Reaction to Federal Reserve Policy?." *Journal of Finance*, 60, 1221 – 1257.

Bloom, N. (2009): "The impact of uncertainty shocks," *Econometrica*, 77, 623–685.

Bloom, N., S. Bond, and J. Van Reenen (2007): "Uncertainty and investment dynamics," *The Review of Economic Studies*, 74, 391–415.

Christiano, L. J., M. Eichenbaum, and C. L. Evans (1999): "Monetary policy shocks: What have we learned and to what end?" *Handbook of macroeconomics*, 1, 65–148.

Fama, E. F. and K. R. French (1993): "Common risk factors in the returns on stocks and bonds," *Journal of Financial Economics*, 33, 3–56.

Gertler, M. and P. Karadi (2015): "Monetary Policy Surprises, Credit Costs, and Economic Activity," *American Economic Journal: Macroeconomics*, 7, 44–76.

Gilchrist, S., D. López-Salido, and E. Zakrajšek (2015): "Monetary Policy and Real Borrowing Costs at the Zero Lower Bound," *American Economic Journal: Macroeconomics*, 7, 77–109.

Grossman, S. J. and Z. Zhou (1996): "Equilibrium analysis of portfolio insurance," *The Journal of Finance*, 51, 1379–1403.

Gürkaynak, R. S., B. Sack, and E. Swanson (2005a): "The Sensitivity of Long-Term Interest Rates to Economic News: Evidence and Implications for Macroeconomic Models." *American Economic Review*, 95, 425 – 436.

Gürkaynak, R. S., B. Sack, and E. T. Swanson (2005b): "Do actions speak louder than words? The response of asset prices to monetary policy actions and statements." *International Journal of Central Banking*, 1, 55 – 93.

Gürkaynak, R. S., B. Sack, and J. H. Wright (2007): "The U.S. Treasury yield curve: 1961 to the present." *Journal of Monetary Economics*, 54, 2291 – 2304.

Gürkaynak, R. S., B. Sack, and J. H. Wright (2010): "The TIPS yield curve and inflation compensation," *American Economic Journal: Macroeconomics*, 70–92.

Hanson, S. G. and J. C. Stein (2015): "Monetary policy and long-term real rates," *Journal of Financial Economics*, 115, 429–448.

Johannes, M., L. A. Lochstoer, and Y. Mou (2015): "Learning about consumption dynamics," *The Journal of Finance*.

Kuttner, K. N. (2001): "Monetary Policy Surprises and Interest Rates: Evidence from the Fed Funds Futures Market." *Journal of Monetary Economics*, 47, 523 – 544.

Lucas, R. E. and E. C. Prescott (1971): "Investment Under Uncertainty," *Econometrica*, 39, 659–681.

Miller, E. M. (1977): "Risk, uncertainty, and divergence of opinion," *The Journal of Finance*, 32, 1151–1168.

Morris, S. and H. S. Shin (2002): "Social Value of Public Information," *The American Economic Review*, 92, 1521–1534.

Nakamura, E. and J. Steinsson (2013): "High Frequency Identification of Monetary Non-Neutrality," Working paper.

Rigobon, R. (2003): "Identification through Heteroskedasticity." *Review of Economics and Statistics*, 85, 777 – 792.

Rigobon, R. and B. Sack (2003): "Measuring the Reaction of Monetary Policy to the Stock Market." *Quarterly Journal of Economics*, 118, 639 – 669.

——— (2004): "The Impact of Monetary Policy on Asset Prices." *Journal of Monetary Economics*, 51, 1553 – 1575.

Romer, C. D. and D. H. Romer (2004): "A New Measure of Monetary Shocks: Derivation and Implications," *American Economic Review*, 1055–1084.

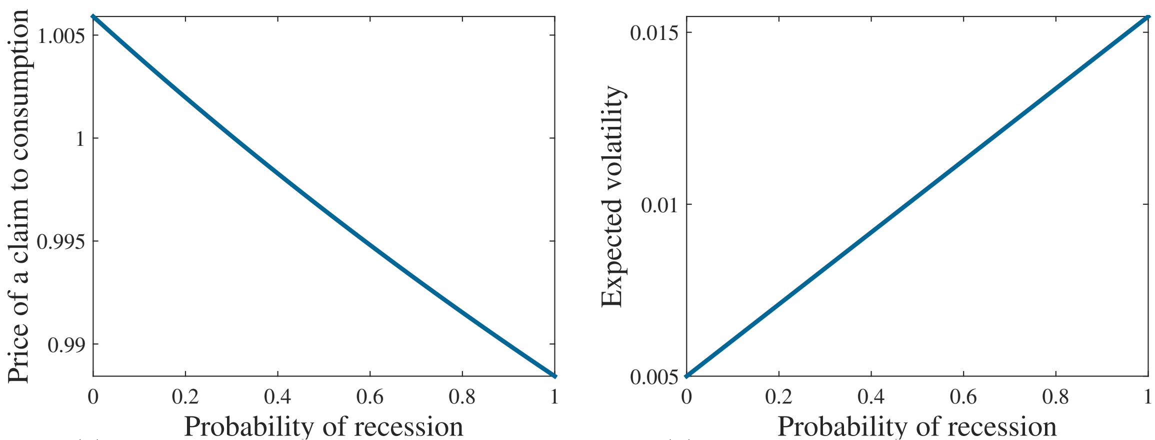
Rudebusch, G. D. (1998): "Do Measures of Monetary Policy in a VAR Make Sense?" *International Economic Review*, 39, pp. 907–931.

Sandmo, A. (1970): "The effect of uncertainty on saving decisions," *The Review of Economic Studies*, 353–360.

Stock, J. H. and M. W. Watson (1999): "Business cycle fluctuations in US macroeconomic time series," *Handbook of Macroeconomics*, 1, 3–64.

Woodford, M. (2001): "Monetary Policy in the Information Economy," in *Symposium on Economic Policy for the Information Economy*, Federal Research Bank of Kansas City, 297–370.

Wright, J. H. (2012): "What does Monetary Policy do to Long-term Interest Rates at the Zero Lower Bound?" *The Economic Journal*, 122, F447–F466.



(a) Price of a claim to consumption

(b) Expected consumption volatility

Figure 1: Stylized equity price and aggregate volatility

Figure 1 illustrates equity price (Figure 1(a)) and expected volatility (Figure 1(b)) as a function of the probability of recession in the conceptual framework of Section 2. The numerical example uses $\mu_1 = .61\%$, $\mu_2 = -.96\%$, $\sigma_1^2 = .25\%^2$, $\sigma_2^2 = 2.39\%^2$ as the parameters of the quarterly consumption growth (source: Johannes et al. (2015)) and $\gamma = 9$ as the coefficient of risk aversion.

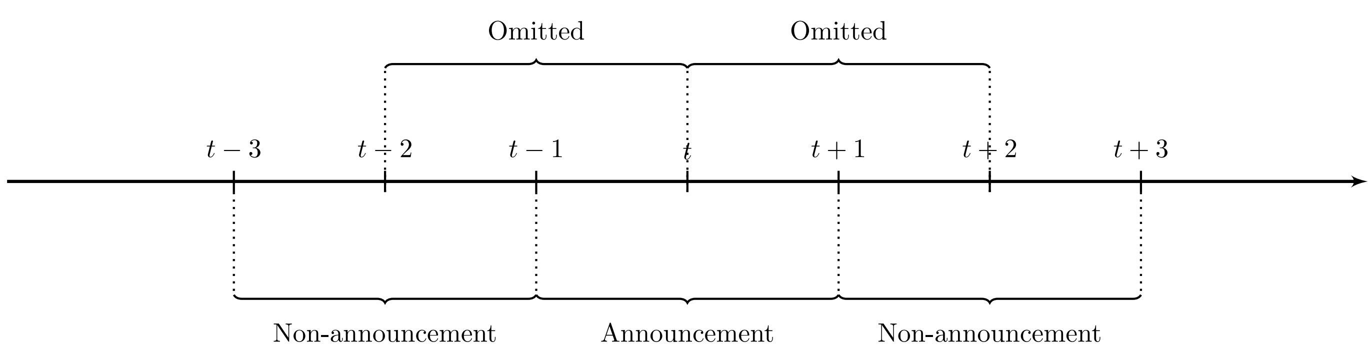


Figure 2: Stylized timeline of shock identification

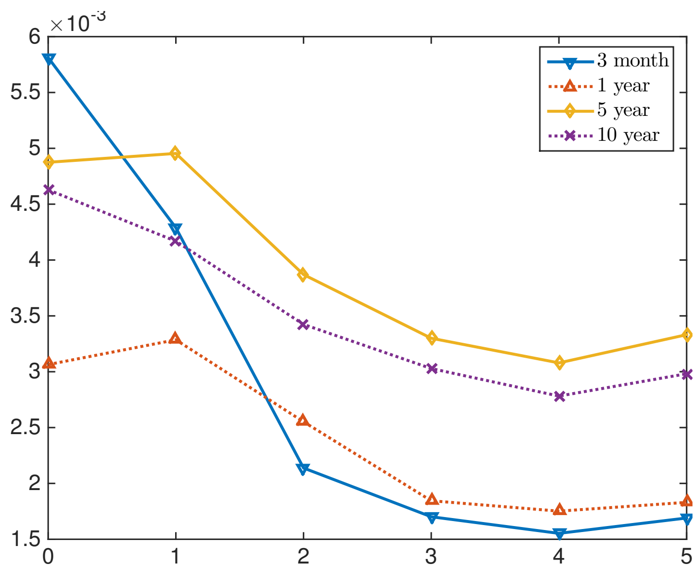
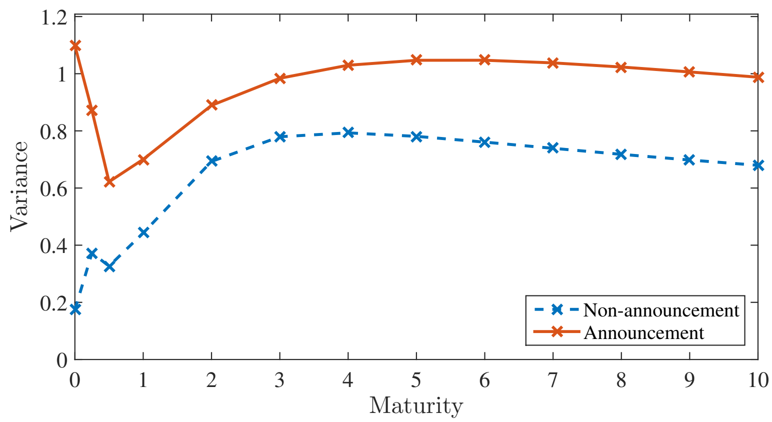
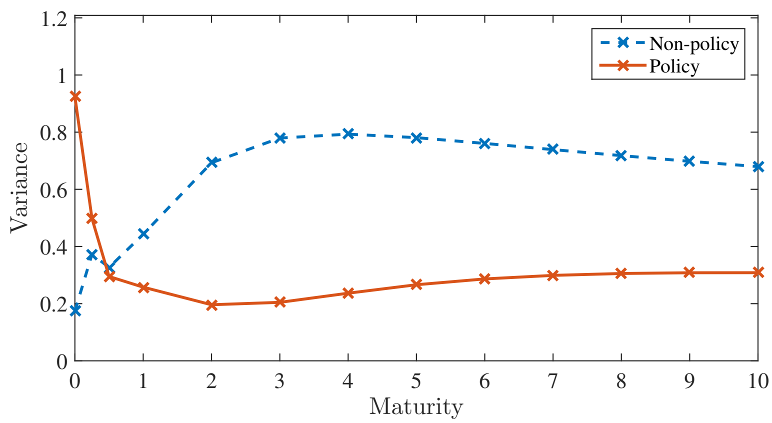


Figure 3: Variance of cumulative one-day changes to the yield curve

Figure 3 illustrates the variance of the cumulative change to the yields of different maturities after an FOMC announcement (date 0). Maturities are in years where the zero maturity yield is the change in the front month Fed Funds future. The sample period runs from 1994-2007 and includes 113 FOMC announcement days.



(a) Announcement vs. Non-Announcement



(b) Policy vs. Non-Policy

Figure 4: Comparing variances in the yield curve

Figure 4 illustrates the variance in yield changes at a range of maturities. Maturities are in years where the zero maturity yield is the change in the front month Fed Funds future. The sample period runs from 1994-2007 and includes 113 FOMC announcement days. Figure 4(a) compares FOMC announcements days versus non-announcement days. Figure 4(b) compares policy variance (announcement less non-announcement) and non-policy variance (announcement variance).

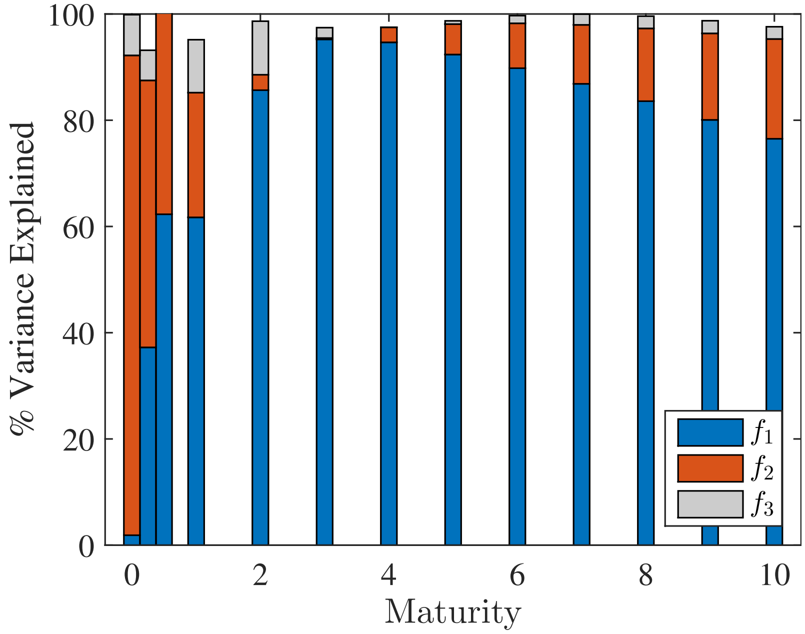


Figure 5: Fraction of explained policy variance by first three principal components

Figure 5 illustrates the policy variance attributable to the first three principal components of the policy matrix at various maturities ranging from the forward Fed Funds rate to 10 years.

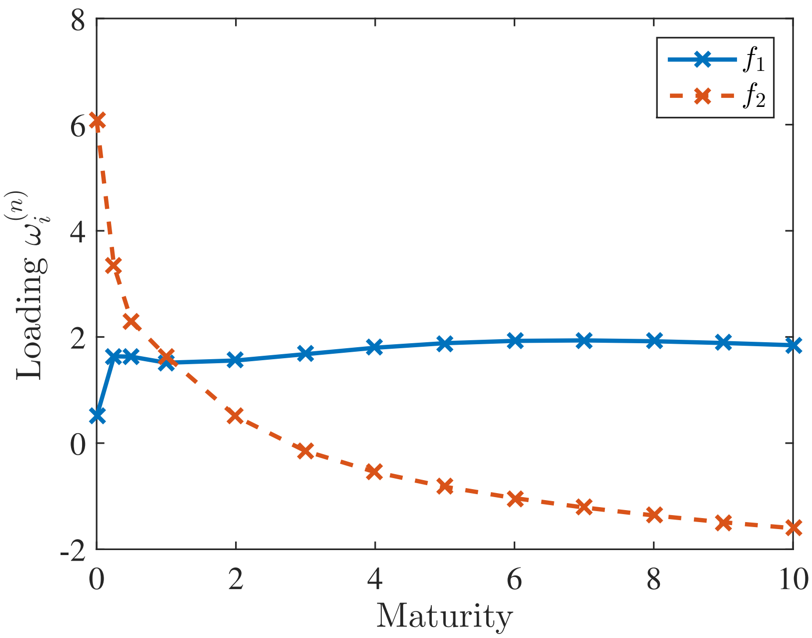


Figure 6: Principal component loading on yields

Figure 6 illustrates the loadings on yields of various maturity for the first two principal components of the policy variance matrix.

$$f_{i,t} = \sum_n \omega_i^{(n)} \Delta y_t^{(n)}$$

Maturities are in years where the forward Fed Funds rate is 0.

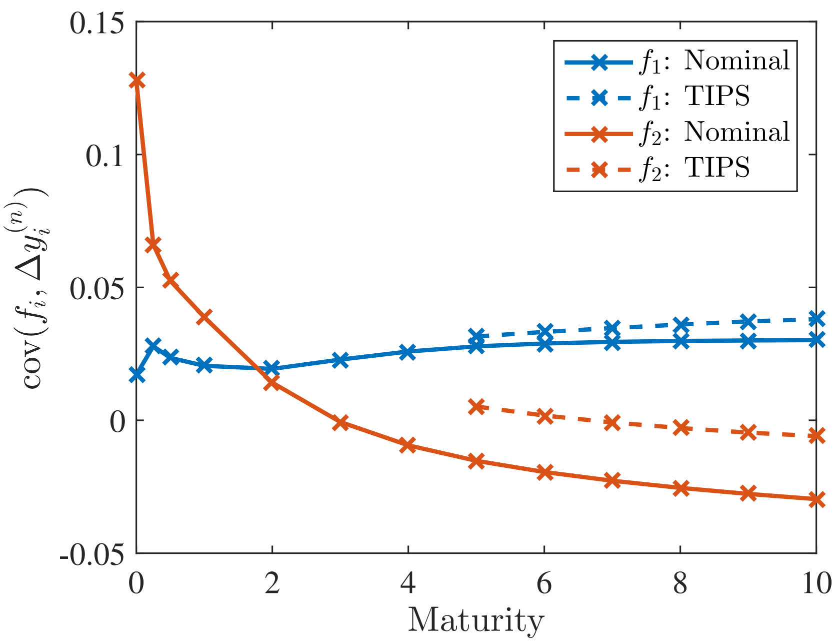
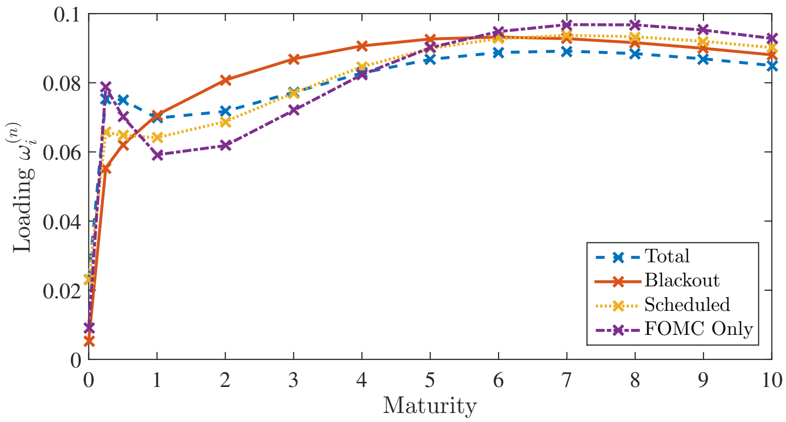
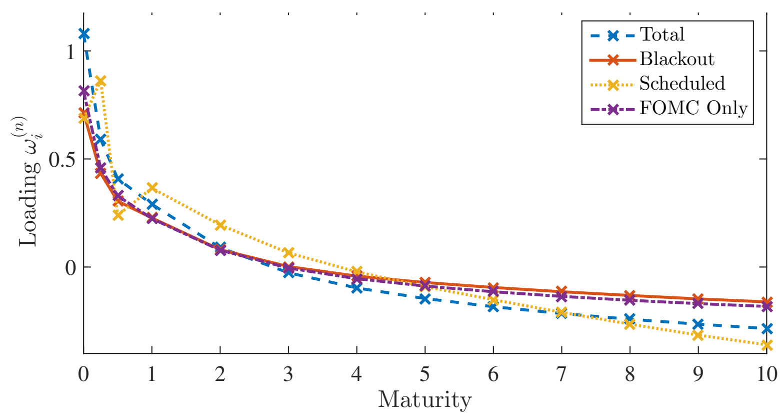


Figure 7: Covariance of yields and policy factors: 1999-2007

Figure 7 illustrates the covariance of the policy factors with changes in yields. Maturities are in years where the forward Fed Funds rate is 0.



(a) f_1



(b) f_2

Figure 8: Comparing principal component loadings across samples

Figure 8 illustrates the loadings on yields of various maturity for the first two principal components of the policy variance matrix.

$$f_{i,t} = \sum_n \omega_i^{(n)} \Delta y_t^{(n)}$$

Loadings in this figure are weighted such that they sum to one. Maturities are in years where the forward Fed Funds rate is 0. Components are derived from three different estimates of the policy covariance matrix. *Total* is the base specification that compares all announcement days vs. all non-announcement days. *Blackout* compares all announcement days to days in the FOMC blackout period. *Scheduled* compares scheduled announcement days to all non-announcement days. Figure 8(a) displays the weights for the first component and Figure 8(b) for the second.

Table 1: Covariance of real yields and breakevens with policy factors

Dep. Var.	(1) 5 year	(2) 10 year	(3) 5 yr BE	(4) 10 yr BE
A	0.00438***	-0.00526**	-0.0105	-0.0117
	(3.48)	(-1.97)	(-1.58)	(-1.50)
f_1	0.144***	0.127***	0.0647***	0.0535***
	(264.07)	(126.02)	(15.59)	(18.95)
f_2	-0.0287***	-0.0323***	-0.0205***	-0.0201***
	(-49.73)	(-26.79)	(-12.30)	(-13.56)
f_1A	0.0561***	0.0554***	0.00211	-0.0130
	(17.26)	(7.24)	(0.13)	(-0.56)
f_2A	-0.00851*	-0.0183***	-0.0159	-0.0192
	(-1.65)	(-6.86)	(-1.41)	(-1.45)
Observations	2132	2132	2132	2132
R-squared	0.99	0.95	0.32	0.31

Table 1 contains coefficient estimates from a regression of changes in Treasury yields on changes in our policy factors. The interaction terms reveal the covariance between monetary policy announcements and the dependent variable. In Columns 1 and 2 the dependent variable is the nominal yield on the five- and ten-year constant maturity Treasury, respectively. In Columns 3 and 4 the dependent variable is the breakeven yield at the five- and ten-year maturity, respectively. The sample period is from 1999-2007. Standard errors in parentheses are calculated over time using Newey-West (5 lags); * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2: Covariance of equity factors and volatility measures with policy factors

Dep. Var.	(1) $R_M - r_f$	(2) R_{HML}	(3) R_{SMB}	(4) Δvix	(5) Δvxo	(6) $\Delta smove$
A	0.310**	-0.0984	0.00893	-0.0396***	-0.0246***	-0.0125**
	(2.21)	(-1.09)	(0.11)	(-5.02)	(-3.02)	(-2.54)
f_1	0.227***	-0.138***	0.180***	-0.00507	-0.00853**	0.0149***
	(2.88)	(-4.13)	(5.30)	(-1.41)	(-2.23)	(5.47)
f_2	0.00873	0.00705	-0.0378**	0.000287	-0.00113	-0.00694***
	(0.34)	(0.50)	(-2.08)	(0.26)	(-1.00)	(-6.72)
f_1A	-0.569**	0.0173	-0.176	0.0306*	0.0330*	0.0218**
	(-2.01)	(0.10)	(-1.05)	(1.70)	(1.85)	(2.39)
f_2A	-0.0775	0.548**	-0.0157	-0.00886	-0.00454	-0.00936*
	(-0.17)	(2.18)	(-0.11)	(-0.64)	(-0.28)	(-1.68)
Observations	3373	3373	3373	3371	3367	3322
R-squared	0.01	0.02	0.02	0.01	0.01	0.05

Table 2 contains coefficient estimates from a regression of equity returns and changes in volatility measures on changes in our policy factors. The interaction terms reveal the covariance between monetary policy announcements and the dependent variable. In Columns 1, 2, and 3 the dependent variable is the excess return on the market, the return on the HML, and the return on SMB, respectively. Column 4 is the change in logs of the VIX, Column 5 is the change in logs of the VXO, Column 6 is the change in logs of the one-year SMOVE, and Column 7 is the change in the economic policy uncertainty from Baker et al. (2013). The sample period is from 1994-2007. Standard errors in parentheses are calculated over time using Newey-West (5 lags); * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 3: Covariance of risky asset retuns with policy factors

	(1)	(2)	(3)	(4)
Dep. Var.	Trade-weighted EX	Commodity Index	Energy Index	Non-Energy Index
A	-0.0216	-0.137	-0.254	0.00454
	(-0.41)	(-0.78)	(-0.97)	(0.05)
f_1	0.105***	0.0620	-0.00625	0.171***
	(4.78)	(0.89)	(-0.06)	(5.08)
f_2	0.00236	-0.0261	-0.0260	-0.0307**
	(0.24)	(-0.78)	(-0.50)	(-2.18)
f_1A	0.293***	-0.728**	-1.004*	-0.286*
	(3.03)	(-2.10)	(-1.85)	(-1.93)
f_2A	-0.00592	0.254	0.308	0.134*
	(-0.11)	(0.85)	(0.72)	(1.76)
Observations	3378	3376	3374	3376
R-squared	0.02	0.00	0.00	0.01

Table 3 contains coefficient estimates from a regression of returns on other risky assets on changes in our policy factors. The interaction terms reveal the covariance between monetary policy announcements and the dependent variable. In Column 1, the dependent variable is the change to the trade-weighted U.S. exchange rate to major currencies. In Columns 2-4, the dependent variable is the change to the S&P GSCI spot index for all commodities, energy commodities, and non-energy commodities, respectively. The sample period is from 1994-2007. Standard errors in parentheses are calculated over time using Newey-West (5 lags); * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 4: Covariance of credit condition measures with policy factors

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Corporate credit conditions					Household credit conditions			
Dep. Var.	AA CP rate (Non.-fin.)	AA CP rate (Fin.)	HY yield	NFCI	FRM spread	Purchases	Refinance	FRM	ARM
A	0.0103**	0.00587	0.00901	0.00237**	-0.00830	1.215	-0.271	0.664	1.257
	(2.27)	(0.93)	(0.95)	(2.05)	(-1.23)	(1.18)	(-0.17)	(0.55)	(0.80)
f_1	0.0189***	0.0165***	0.0200***		-0.186***				
	(7.99)	(6.84)	(4.31)		(-17.08)				
f_2	0.00393***	0.00430***	-0.00555**		0.0355***				
	(2.93)	(2.85)	(-2.42)		(11.01)				
f_1A	0.0402***	0.0340**	0.0290*		-0.0471**				
	(3.43)	(2.56)	(1.68)		(-1.96)				
f_2A	0.0657***	0.0734***	0.00531		0.00769				
	(7.93)	(8.32)	(0.38)		(0.42)				
$L.f_1$				0.000384		-2.919*	-13.30***	-9.113***	-1.059
				(0.28)		(-1.81)	(-5.05)	(-4.41)	(-0.56)
$L.f_2$				-0.000512		-0.248	1.847*	1.182	-0.811
				(-0.96)		(-0.33)	(1.79)	(1.33)	(-0.80)
$L.f_1$A				0.00310		3.501	1.496	1.325	2.797
				(1.00)		(0.88)	(0.25)	(0.27)	(0.69)
$L.f_2$A				0.00645**		2.378	-4.052	-1.065	0.664
				(2.19)		(1.18)	(-0.91)	(-0.34)	(0.32)
$L2.f_1$				0.000975		-1.298	-13.68***	-7.837***	-3.797**
				(0.82)		(-1.00)	(-6.50)	(-4.86)	(-2.53)
$L2.f_2$				-0.000863		-0.0658	4.034***	2.225***	0.231
				(-1.26)		(-0.11)	(3.84)	(2.84)	(0.34)
$L2.f_1$A				-0.00114		-3.767	-11.46*	-7.215	-6.012
				(-0.29)		(-0.92)	(-1.65)	(-1.45)	(-1.12)
$L2.f_2$A				0.00553***		-5.418*	-16.97***	-11.49***	-12.07**
				(3.36)		(-1.84)	(-2.83)	(-2.68)	(-2.15)
Observations	2632	1837	3382	705	707	705	705	705	705
R-squared	0.19	0.17	0.03	0.63	0.63	0.12	0.18	0.15	0.14

Table 4 contains coefficient estimates from a regression of changes to measures of credit conditions on changes in our policy factors. The interaction terms reveal the covariance between monetary policy announcements and the dependent variable at different lags. Columns 1-5 measure credit conditions for the corporate sector. In Columns 1, 2, and 3, the dependent variables is the two day change of the AA CP rate for non-financial borrowers, the two day change of the AA CP rate for financial borrowers, and the two day change of the average yield of a basket of high yield-rated bonds, respectively. Column 4 is the weekly change in the Chicago Fed's National Financial Conditions Index. Columns 5-9 measure credit conditions for the household sector. In Column 6, the dependent variable is the weekly change to the spread between the average fixed rate mortgage rate and 5 year Treasury yield. In Columns 7-10, the dependent variable is the weekly percentage change to the number of mortgage applications for purchases, refinances, fixed rate mortgages and adjustable rate mortgages, respectively. The sample period is from 1994-2007. Regressions in Columns 4 and 6-9 include three lags of the dependent variable. Standard errors in parentheses are calculated over time using Newey-West (5 lags); * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 5: Covariance of money supply measures with policy factors

Dep. Var.	(1) M1	(2) M2	(3) Currency	(4) M2-M1
A	-0.0262	-0.00773	0.0195*	0.00478
	(-0.34)	(-0.35)	(1.66)	(0.22)
f_1	-0.0158	-0.0266	-0.0130	-0.0288
	(-0.23)	(-1.57)	(-0.97)	(-1.15)
f_2	-0.0545**	-0.0205**	-0.00449	-0.00586
	(-2.06)	(-2.21)	(-0.65)	(-0.64)
f_1A	0.0900	0.0235	0.0551*	-0.0205
	(0.36)	(0.36)	(1.90)	(-0.38)
f_2A	-0.946*	-0.334**	-0.0577***	-0.167***
	(-1.66)	(-2.42)	(-2.86)	(-5.21)
$L.f_1$	-0.0960	-0.0256	0.00790	-0.0106
	(-1.03)	(-0.96)	(0.48)	(-0.48)
$L.f_2$	-0.0947	-0.0325**	-0.00956	-0.0148*
	(-1.61)	(-2.06)	(-1.09)	(-1.70)
$L.f_1$A	0.0958	-0.0154	0.0171	-0.0190
	(0.60)	(-0.26)	(0.49)	(-0.39)
$L.f_2$A	0.172	0.0875	-0.00358	-0.00187
	(1.04)	(1.33)	(-0.21)	(-0.08)
$L2.f_1$	0.0600	0.0231	-0.00743	-0.0121
	(0.96)	(1.14)	(-0.62)	(-0.54)
$L2.f_2$	0.0177	0.00122	0.00575	-0.0110
	(0.88)	(0.17)	(0.67)	(-1.43)
$L2.f_1$A	0.112	-0.000117	0.0707*	0.00665
	(0.73)	(-0.00)	(1.69)	(0.10)
$L2.f_2$A	0.123	0.0598	-0.0115	-0.0332
	(1.00)	(1.11)	(-0.42)	(-0.80)
Observations	705	705	705	705
R-squared	0.33	0.25	0.16	0.08

Table 5 contains coefficient estimates from a regression of changes in measures of money supply on changes in our policy factors. The interaction terms reveal the covariance between monetary policy announcements and the dependent variable at different lags. In Columns 1, 2, and 3 the dependent variable is the one week percentage change of M1, M2 and money stock currency, respectively; in column 4, the dependent variable is the one week percentage change in wholesale funding (M2-M1). The sample period is from 1994-2007. Regressions include three lags of the dependent variable. Standard errors in parentheses are calculated over time using Newey-West (5 lags); * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

A Additional Results

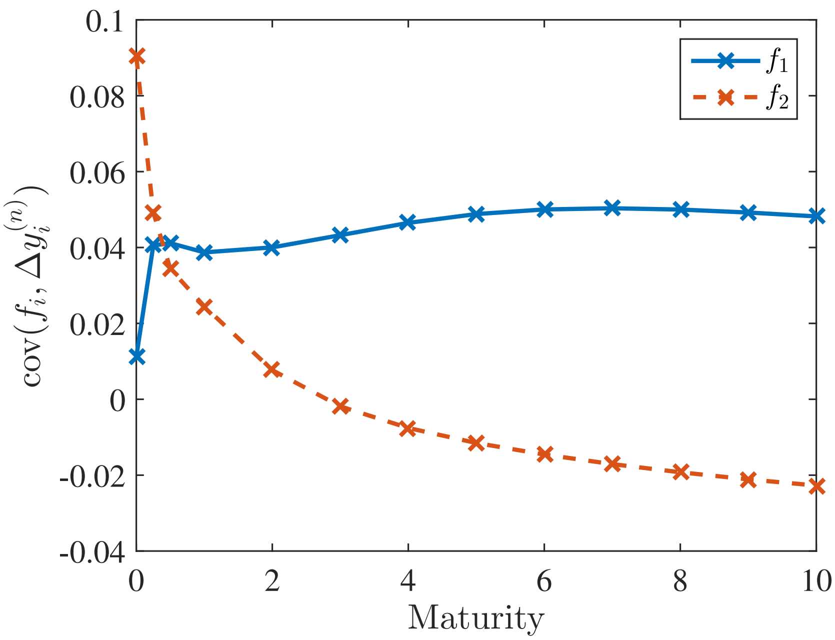
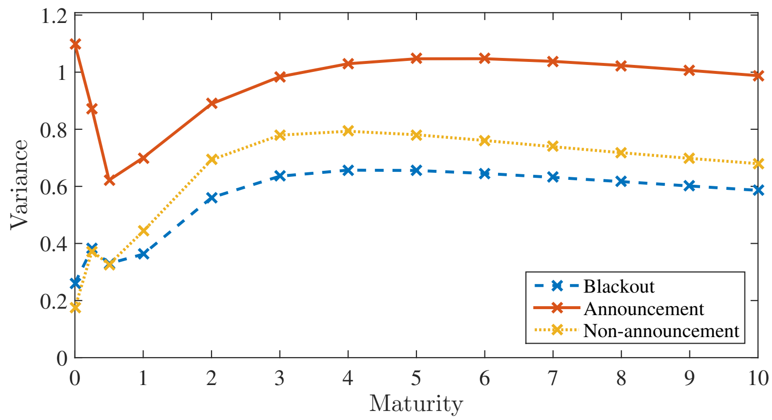
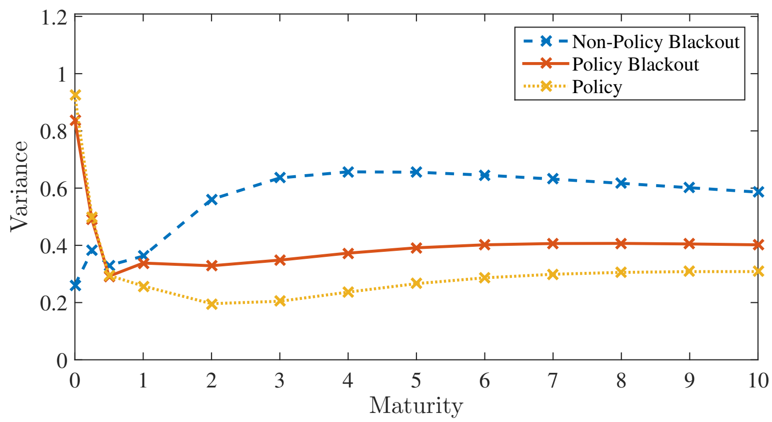


Figure 9: Covariance of yields and policy factors: 1994-2007

Figure 9 illustrates the covariance of the two policy factors with changes in yields. Maturities are in years where the forward Fed Funds rate is 0.



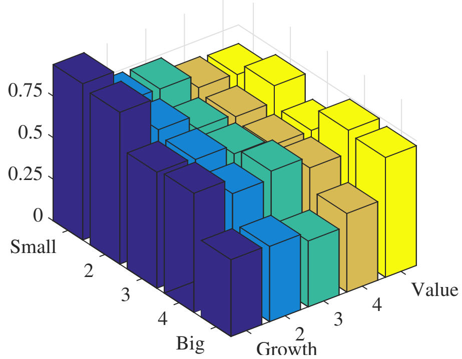
(a) Announcement vs. Non-Announcement



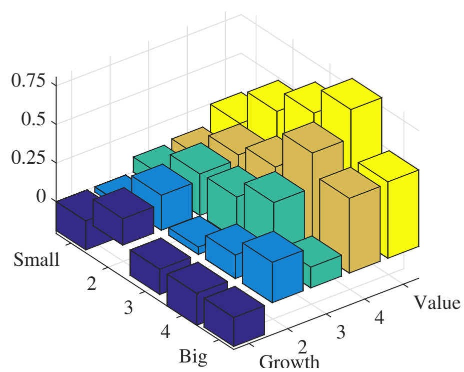
(b) Policy vs. Non-Policy

Figure 10: Comparing variances in the yield curve, including blackout periods

Figure 10 illustrates the variance in yield changes at a range of maturities. Maturities are in years where the zero maturity yield is the change in the front month Fed Funds future. The sample period runs from 1994-2007 and includes 113 FOMC announcement days. Figure 10(a) compares FOMC announcements days versus non-announcement days and versus blackout periods. Figure 10(b) compares policy variance (announcement less non-announcement), blackout policy variance (announcement less blackout) and non-policy (blackout variance).



(a) Factor 1



(b) Factor 2

Figure 11: Covariance of factors with Fama-French 25

Figure 11 illustrates the differential covariance between our factors on announcement days for the FF25 portfolios. Figure 11(a) is the differential covariance of factor 1 on FOMC announcements days versus non-announcement days. Figure 11(b) is the differential covariance of factor 2 on FOMC announcements days versus non-announcement days.

Table 6: Covariance of real yields and breakevens with policy factors at weekly frequency

	(1)	(2)	(3)	(4)
Dep. Var.	5 year	10 year	5 yr BE	10 yr BE
A	0.00301**	-0.00708**	-0.000891	-0.00555
	(2.30)	(-2.29)	(-0.11)	(-0.73)
f_1	0.308***	0.280***	0.0915***	0.0789***
	(203.49)	(85.53)	(7.64)	(8.43)
f_2	-0.0655***	-0.0732***	-0.0312***	-0.0328***
	(-67.93)	(-33.80)	(-6.72)	(-7.86)
f_1A	0.0264***	0.00701	-0.00262	-0.0352*
	(5.33)	(0.96)	(-0.11)	(-1.76)
f_2A	-0.00219	-0.0269***	-0.00868	-0.0267*
	(-0.34)	(-4.76)	(-0.57)	(-1.65)
Observations	707	707	707	707
R-squared	0.99	0.96	0.23	0.24

Table 6 contains coefficient estimates from a regression of changes in Treasury yields on changes in our policy factors at a weekly frequency. The interaction terms reveal the covariance between monetary policy announcements and the dependent variable. In Columns 1 and 2, the dependent variable is the nominal yield on the five- and ten-year constant maturity Treasury, respectively. In Columns 3 and 4 the dependent variable is the breakeven yield at the five- and ten-year maturity, respectively. The sample period is from 1999-2007. Standard errors in parentheses are calculated over time using Newey-West (5 lags); * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.