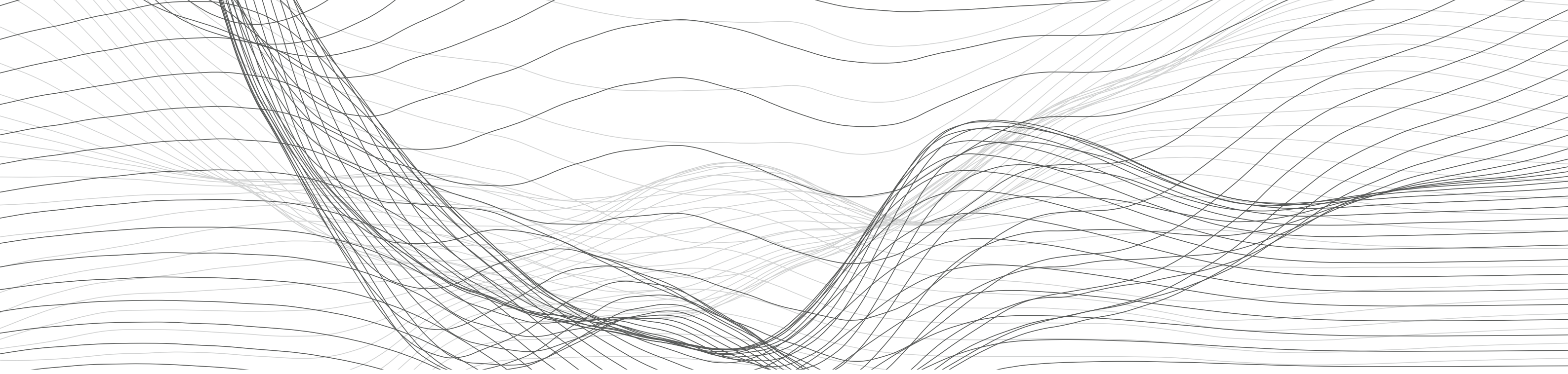

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Maxime Phillot, Samuel Reynard

SNB Working Papers
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ISSN 1660-7716 (printed version)
ISSN 1660-7724 (online version)

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P.O. Box, CH-8022 Zurich

Monetary Policy Financial Transmission and Treasury Liquidity Premia[*]

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August 2021

Abstract

We quantify the effects of monetary policy shocks on the yield curve through their impact on Treasury liquidity premia. When the Fed raises interest rates, the spread between less-liquid assets and Treasuries of the same maturity and risk increases, as the liquidity value of holding Treasuries increases when the aggregate volume of banks' customer deposits decreases. The longer the maturity is, the smaller—but still significant—the increase in the liquidity premium is, as longer-term Treasuries are less liquid. Due to this change in liquidity premia, the spread between a 10-year Treasury bond and a 3-month T-bill yield increases by approximately 5 basis points for a one-percentage-point increase in the policy rate, i.e., the Treasury yield curve steepens, ceteris paribus.

JEL Classification: E52; E43; E41

Keywords: Treasury liquidity premia; Monetary policy; Yield curve; Deposit channel

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[*] We are grateful to Petra Gerlach, Mark Watson, an anonymous referee as well as Swiss National Bank seminar and AEA 2021 meeting participants for their useful comments. We are thankful to Adrian et al. (2017), and Jarociński and Karadi (2020) for providing us with their data. The views, opinions, findings, and conclusions or recommendations expressed in this paper are strictly those of the authors. They do not necessarily reflect the views of the Swiss National Bank (SNB). The SNB takes no responsibility for any errors or omissions in, or for the correctness of, the information contained in this paper.

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1 Introduction

The monetarist literature has emphasized the role of deposit fluctuations as proxies for various substitution effects of monetary policy when many asset prices matter for aggregate demand (Nelson, 2003). This paper identifies and quantifies the macroeconomic dynamic effects of substitutions between short- and long-term Treasuries due to changes in liquidity premia, which occur because of bank customer deposits' response to monetary policy interest rate movements. We thus identify and quantify monetary policy effects on the yield curve through relative liquidity premia.

Nagel (2016) documents, using static regressions, a positive relationship between the short-term Treasury liquidity premium and the federal funds rate. When the federal funds rate increases, the spread between nonliquid short-term assets and T-bills increases. Drechsler et al. (2018) present a monetary policy transmission channel through the banking system or deposits channel that can explain Nagel's empirical findings. As the opportunity cost of holding money increases, agents want to hold less money and more Treasuries to get liquidity services. The spread increases as the liquidity value of holding short-term T-bills increases when the aggregate amount of banks' customer deposits decreases.

We use a macro SVAR model to quantify the effects of well-identified monetary policy interest rate shocks on the yield curve due to changes in liquidity premia. Our SVAR includes, as in Gertler and Karadi (2015), the excess bond premium. In addition, we include and assess the evolution of Treasury liquidity premia for different maturities along the yield curve.

Our methodology is based on Gertler and Karadi (2015) in that we follow the approach given by Gürkaynak et al. (2005) by using the first principal component of monetary policy surprises occurring on FOMC announcement days corresponding to exogenous changes in expectations about current monetary policy.

After aggregating the resulting so-called target factor, we use it as an external instrument (Stock and Watson, 2012) to identify our monthly SVAR between 1991 and 2019. We finally compute the IRFs of the economy and liquidity premia to a monetary policy shock and infer their significance using a recursive-design wild bootstrap procedure (Gonçalves and Kilian, 2004; Mertens and Ravn, 2013).

We show that the monetary effects on the yield curve via liquidity premia vary across maturities. When the Fed raises interest rates, the liquidity premium of longer-term Treasuries significantly increases, but by less than that of shorter T-bills, as they are less liquid. A higher liquidity premium means a relatively lower Treasury yield, as investors are willing to forego yield in exchange for liquidity. As the liquidity premium rises more for short- than for long-term Treasuries, short-term yields relatively decrease, leading to a steepening of the slope of the yield curve not related to policy expectations or a risk premium but reflecting liquidity premia reactions to monetary policy. The overall effect on the yield curve then depends on other factors such as whether the interest rate increase is expected to be temporary or not.

Thus, when measured with Treasury yields, monetary policy interest rate actions have a relatively greater effect on long-term rates than on short-term rates, as the rise in short-term yields is dampened by the higher liquidity premium when the policy interest rate increases, and vice versa. Thus, a decline in deposits is an indicator of a relatively higher long-term real rate occurring via liquidity premia term structure changes.

Our results shed light on the expectation hypothesis, as monetary policy affects the term structure through liquidity premia. Moreover, Treasury liquidity premia fluctuations have recently been used to understand various issues such as real equilibrium interest rate movements (Bok et al., 2018; Ferreira and Shousha, 2020) and exchange rate forecasting (Engel and Wu, 2018). Our results contribute to the quantification of the effect of monetary policy on such fluctuations.

Section 2 discusses the conceptual framework. Empirical results are presented in section 3. Finally, section 4 concludes.

2 The Deposit Channel and Treasury Liquidity

According to the deposit channel presented by Drechsler et al. (2018, DSS hereafter), the effect of monetary policy on the Treasury liquidity premium arises from two facts. First, interest rates offered by commercial banks on customers' deposits adjust slowly and only partially to changes in monetary policy rates. Second, households and firms adjust their deposit holdings to changes in opportunity costs, reflecting traditional money demand motives. Changes in the volume of deposits, and thus of the total liquidity in the financial system which consists of money and other liquid assets, affect the liquidity value of Treasuries.

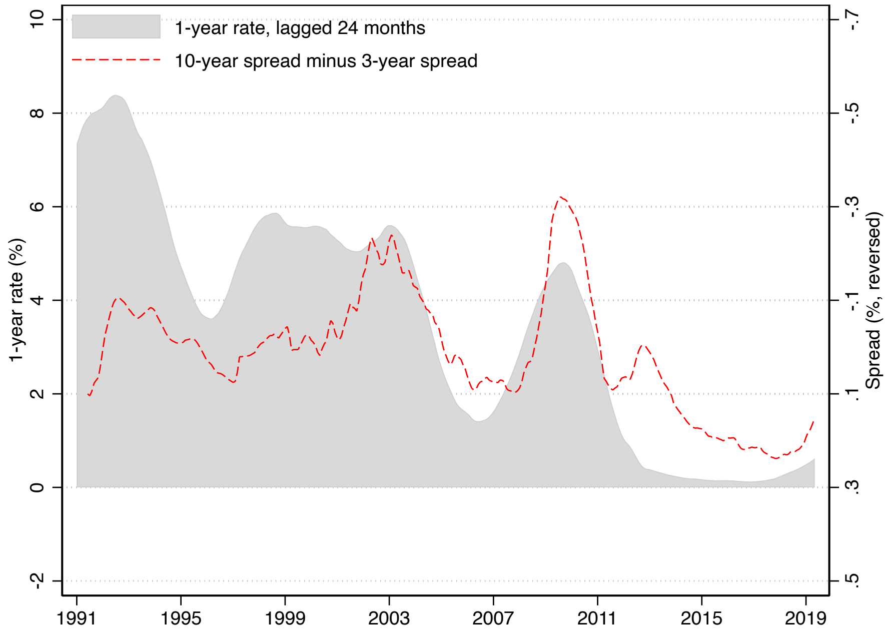
DSS present a model of banks' pricing behavior, where market power causes deposit interest rate spreads to increase with increases in the federal funds rate, as deposit rates adjust only partially to federal funds rate increases. Bank customers respond to this change in opportunity cost, reflecting money demand motives. As the aggregate amount of customer deposits decreases with an increase in the federal funds rate, the liquidity value of holding T-bills increases. The spread between illiquid bond and T-bill yields thus increases. In DSS, deposits are modeled as providing liquidity services in the utility function. DSS show that this deposit channel affects lending and thus monetary policy transmission.

Our empirical results, presented in the next section, show that liquidity premia of Treasuries of different maturities react differently to a monetary policy shock. The longer the maturity is, the less the liquidity premium increases with the federal funds rate. The spread between a 10-year Treasury bond and a 3-month T-bill declines by approximately 5 basis points for a one percentage point increase in the policy rate.

Figure 1 displays the opposite of the difference between the 10-year and 3-month liquidity premia together with the 1-year nominal interest rate that we interpret as the monetary policy rate. Both variables are 24-month moving averages. As shown by this graph, the liquidity premium we study has fluctuated by approximately 60 basis points since the start of the 1990s.

Thus, there is a need to model the process of obtaining deposits from liquid assets such as Treasuries, like in Reynard and Schabert (2010), as Treasuries provide liquidity services because they can easily be converted into deposits to make payments. Such a model should lead to different liquidity premia for different maturities, as we find empirically.

Figure 1: Federal Funds Rate & Liquidity Premium Differences



Notes: The shaded area denotes the 1-year Treasury rate (left scale), while the dashed red line denotes the opposite of the spread between the 10-year liquidity premium and the 3-year liquidity premium (right scale) lagged by 24 months. All variables are 24-month moving averages.

3 The Identified Effects of Monetary Policy on Liquidity Premia

The method used in this paper proceeds over two steps and is largely based on Gertler and Karadi (2015).

First, we use the high-frequency identification (HFI) proposed by Faust et al. (2004) and follow the approach of Gürkaynak et al. (2005) by using the first principal component of monetary policy surprises occurring on Federal Open Market Committee (FOMC) announcement days corresponding to exogenous changes in expectations regarding current monetary policy.

Second, we aggregate the resulting target factor and use it as an external instrument (Stock and Watson, 2012) to identify our monthly structural vector autoregressive process (SVAR). We then compute the impulse response functions (IRFs) of the economy and liquidity premia to a monetary policy shock and infer their significance using a recursive-design wild bootstrap procedure (Gonçalves and Kilian, 2004; Mertens and Ravn, 2013).

3.1 HFI of Monetary Policy Shocks

The main purpose of the HFI scheme is to observe changes in an outcome around the time of shocks within a time window narrow enough to ensure that the changes in the given outcome are caused by shocks and nothing else.

The idea originates from Kuttner (2001), who estimates the effect of changes in the Federal Reserve's policy on various interest rates by separating anticipated from unanticipated changes in the target rate using daily data on federal funds futures. Accordingly, Faust et al. (2004) measure from futures daily data the impact of these unexpected changes in monetary policy on the expected path of interest rates and identify a VAR requiring that the response of the federal funds rate to monetary policy shocks matches that in the data.

Gürkaynak et al. (2005) extend the methodology and argue that two factors underly the response of futures prices to monetary policy. The argument relies on the observation that there have been monetary policy announcements associated with no change in the *target* rate itself but with changes in communication over the future *path* of monetary policy causing futures prices to change.[1,2]

In particular, the authors define a set of five monetary policy surprises mp^j for $j = 1, ..., 5$ computed using 1-, 2-, 3- and 4-month federal funds futures contracts and 6, 9 and 12-month Eurodollar futures contracts. The authors interpret as a change in expectations regarding current and future monetary policy the change in the interest rate implied by the futures contracts on a FOMC announcement day within a 30-minute window around the time of announcement.

The first two surprises, mp^1 and mp^2, are calculated using the federal funds futures and respectively reflect the rate expected to prevail until the next FOMC meeting and that expected to prevail thereafter.[3] Because there are eight scheduled meetings per year, meetings occur on average every six to seven weeks. However, on a day on which a meeting takes place, the maturity of the futures due in the month corresponding to the following scheduled meeting ranges from two to three months. One therefore must compute mp^2 always considering the actual number of months separating one meeting from the next. The last three surprises mp^j for $j = 3, 4$, and 5 are computed as the daily returns on FOMC days of the 6-, 9- and 12-month Eurodollar futures, respectively.

Finally, the authors extract the two principal components stemming from

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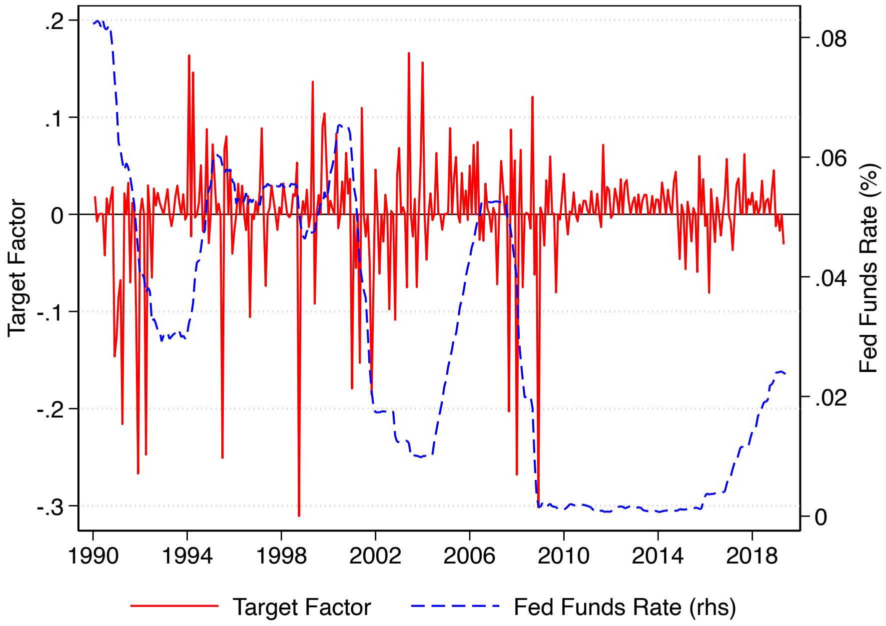
[1] This has been especially true since 2007, when the zero lower bound compelled the Federal Reserve to conduct its monetary policy using so-called forward guidance.

[2] See Appendix A for further details about the HFI methodology.

[3] By construction of the federal funds futures contracts, which pay off according to the average effective federal funds rate prevailing over the agreed-upon month, the contract partly reflects the rate realized thus far in that month and the expected rate to prevail until the end of the period thereof. Measuring the suprise in monetary policy associated with an FOMC meeting therefore requires some adjustments. We provide details regarding these adjustments in Appendix A.

Figure 2: Cumulated Series of Monetary Policy Shocks and Federal Funds Rate



Notes: The figure plots the federal funds rate (dashed line, centered around zero) and the cumulated series of shocks to expectations about monetary policy around the time of FOMC announcement days identified as the target factor (solid line). The latter is the instrument we use for the 1-year rate in the monthly SVAR.

this set of five monetary policy surprises, rotate them to give them a structural interpretation and label them the target factor and path factor. In this paper, however, because we are interested in the effects of unexpected changes in the stance of *current* monetary policy on Treasury liquidity premia, we focus on the target factor. This choice is also motivated by Jarociński and Karadi (2020), who show in their online appendix that the target factor performs marginally better as an instrument for the zero lower bound period.

The last step before estimating our monthly SVAR requires that we aggregate the daily series of monetary policy surprises into a monthly series of shocks that will ultimately serve our specification as an instrument. As in Jarociński and Karadi (2020), we use the sum of the daily surprises occurring in each month.

Figure 2 plots the series stemming from this aggregation of the target factor (solid red line) on the left scale together with the federal funds rate (red dashed line) on the right scale. By construction, the factor series cumulates monetary policy surprises that relate to the target of the central bank, which explains why the effective funds rate appears to be driven by the monetary policy surprises. Recall that these surprises were identified to ensure that they were exogenously driven by changes in the expectations about the future target rate.

3.2 SVAR With External Instrument

Let us consider the following p-th order structural vector autoregressive (SVAR) model with k endogenous variables:

$$AY_t = \sum_{s=1}^{p} \psi_s Y_{t-s} + \varepsilon_t, \tag{1}$$

where Y_t is a $(k \times 1)$ vector of endogenous variables, A and the ψ_ss are $(k \times k)$ matrices of coefficients, and ε_t is a $(k \times 1)$ vector of structural innovations such that $E[\varepsilon_t \varepsilon_t'] = I_k$ and $E[\varepsilon_t \varepsilon_s'] = 0_k$ for all $t \neq s$.

We estimate the following reduced-form of (1):

$$Y_t = \sum_{s=1}^{p} \phi_s Y_{t-s} + u_t, \tag{2}$$

where $\phi_s = A^{-1}\psi_s$, and u_t is a $(k \times 1)$ vector of reduced form disturbances such that $E[u_t u_t'] = \Sigma$ and $E[u_t u_s'] = 0_k$ for all $t \neq s$. We have:

$$u_t = B\varepsilon_t. \tag{3}$$

To retrieve the structural impulse response functions (IRFs) implied by (1) from the estimates of (2), we need to identify B. Suppose we are only interested in the response of the system to the structural innovations of variable y_t^m (the monetary policy rate, in our case). Without a loss of generality, assume that this variable is placed first in Y_t such that the vector of structural innovations can be partitioned as follows:

$$\varepsilon_t = (\varepsilon_t^m \ \varepsilon_t^{\bullet 2} \ \cdots \ \varepsilon_t^{\bullet k})' = (\varepsilon_t^m \ \varepsilon_t^{\bullet})'. \tag{4}$$

Likewise, the relationship between reduced-form residuals and structural innovations can be partitioned as follows:

$$B = (s^m \ s^{\bullet}). \tag{5}$$

To compute the structural IRFs, we thus need to identify only the first column s^m. Suppose further that we have a set of instruments Z_t that fulfills the following two conditions:

$$\begin{aligned} E(Z_t \varepsilon_t^{m\prime}) &= \alpha && (6) \\ E(Z_t \varepsilon_t^{\bullet\prime}) &= 0_{k-1}. && (7) \end{aligned}$$

These conditions are similar to those of a standard instrumental variable, i.e., the relevance and the exogeneity conditions. Under these conditions, we can identify s^m with a two-stage procedure summarized as follows.

I. First Stage: Estimate (2) and obtain the reduced-form residuals:

$$(\hat{u}_t^m \ \hat{u}_t^\bullet) = Y_t - \sum_{s=1}^{p} \hat{\phi}_s Y_{t-s}. \tag{8}$$

Regress the reduced-form residuals that stem from the equation of the variable of interest on the set of instruments:

$$\hat{u}_t^m = \gamma Z_t + \xi_t, \tag{9}$$

and obtain the fitted values $\tilde{u}_t^m = \hat{\gamma} Z_t$.

II. Second Stage: Regress the reduced-form residuals $\hat{u}_t^\bullet$ stemming from the equations of the $k-1$ other variables $y_t^\bullet$ on the fitted values $\tilde{u}_t^m$ separately:

$$\begin{pmatrix} \hat{u}_t^{\bullet 2} \\ \hat{u}_t^{\bullet 3} \\ \vdots \\ \hat{u}_t^{\bullet k} \end{pmatrix} = \begin{pmatrix} \tilde{u}_t^m & 0 & \cdots & 0 \\ 0 & \tilde{u}_t^m & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & \tilde{u}_t^m \end{pmatrix} \begin{pmatrix} \beta^{\bullet 2} \\ \beta^{\bullet 3} \\ \vdots \\ \beta^{\bullet k} \end{pmatrix} + \begin{pmatrix} \eta_t^{\bullet 2} \\ \eta_t^{\bullet 3} \\ \vdots \\ \eta_t^{\bullet k} \end{pmatrix} \tag{10}$$

Column s^m can finally be established using the following:

$$\kappa^{-1} s^m = (1 \ \hat{\beta}^{\bullet 2} \ \cdots \ \hat{\beta}^{\bullet k})', \tag{11}$$

where κ is a scaling factor identified up to a sign convention whose closed solution can be found in Gertler and Karadi (2015, footnote 4).

System (1) can be finally estimated using the p reduced-form $\hat{\phi}_s$s, the $((k-1) \times 1)$ vector $\hat{\beta}^\bullet$ and κ, by imposing the corresponding constraints on the first column of B.

Notwithstanding, we are ultimately interested in computing the IRFs resulting from VAR(p). Using lag operator L defined such that $L^p\eta_t = \eta_{t-p}$, Equation (2) is equivalent to the following:

$$(I_k - L\phi_1 - \cdots - L^p\phi_p)Y_t = u_t. \tag{12}$$

With $\Phi(L) = (I_k - L\phi_1 - \cdots - L^p\phi_p)$, and provided that the VAR in (2) is stable, we can obtain its infinite-order vector moving average representation as follows:

$$Y_t = \Phi(L)^{-1}u_{t-i} = \sum_{i=0}^{\infty} \Gamma_i u_{t-i}, \tag{13}$$

where $\Gamma_0 = I_k$ and $\Gamma_i = \sum_{s=1}^{i} \Gamma_{i-s}\phi_s$ for $i = 1, 2, \ldots$. The notation included in (13) is convenient, as it enables us to see that matrices $\Gamma_i = \partial Y_{t+i}/\partial u_t'$ are the IRFs. Indeed, the j, k entry of Γ_i is the response of the j-th element of Y_t after i periods to a one-time unit shock to the k-th element of u_t. Because we identify the first column of B in the previous section, we can obtain a causal interpretation of the effect of the orthogonalized shock of interest on the whole system from $\partial Y_{t+i}/\partial s^m \varepsilon_t^m$.

Finally, to account for potential conditional heteroskedasticity and avoid any generated regressor problem, we use a recursive-design wild bootstrap procedure to compute confidence intervals (CIs) for the IRFs.

The idea is to draw T independent observations $\{\nu_t\}_{t=1,\ldots,T}$ of a random variable ν_t such that

$$\nu_t = \begin{cases} +1 & \text{with probability } 1/2 \\ -1 & \text{with probability } 1/2 \end{cases} \tag{14}$$

and to recursively generate a pseudoseries Y_t^* according to:

$$Y_t^* = \sum_{s=1}^{p} \hat{\phi}_s Y_{t-s}^* + \hat{u}_t \nu_t, \tag{15}$$

where $\hat{\phi}_s$ and $\hat{u}_t$ have been obtained after estimating (2).

Using the pseudoseries of instruments $Z_t^* = Z_t \nu_t$, one can reestimate the SVAR described above N times. The α-level CIs are then simply the $(\alpha/2)$-th and $(1-\alpha/2)$-th percentiles of the resulting distribution of bootstrapped IRFs.

3.3 Data

We take the 1-year T-Bill rate as the policy rate and instrument it using the target factor previously defined.[4] To characterize the response of the economy to a monetary policy shock, we include in our SVAR the log of the consumer price index and the log of the industrial production index as in Gertler and Karadi (2015). We add the excess bond premium (EBP) (Gilchrist and Zakrajšek, 2012) to match their specification and to account for the so-called credit channel of monetary policy, and we add the log of bank customers' checkable and savings deposits to account for the deposit channel. We also include several measures of liquidity premia.

Our dataset is monthly and runs from June 1991 to May 2019. Note that Gertler and Karadi (2015) use a shorter sample period to identify the contemporaneous response of their system to a monetary policy shock than the one on which they impose the resulting constraints. The authors therefore

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[4] We thank Jarociński and Karadi (2020) for providing us with the updated series of the target factor.

assume that the instrumental subsample is a representative characterization of the way in which surprises in monetary policy affect the economy. We do not need to make this assumption because Refcorp bonds were first issued in 1991.

Data sources.—Table 1 describes the source, frequency and timespan of our data.

Table 1: Data Sources, Timespans and Transformations

Variable	Source	Timespan	Frequency
a. Interest Rates			
Effective Federal Funds Rate	`fred.stlouisfed.org`	1954.07–2019.06	Monthly
3-Month General Collateral Repurchase Agreement	Bloomberg	1991.06–2019.06	Daily
Moody's Seasoned AAA Corporate Bond Yield	`fred.stlouisfed.org`	1954.07–2019.06	Monthly
Z-Year Treasury Constant Maturity Rate*	`fred.stlouisfed.org`	1954.07–2019.06	Monthly
Z-year Refcorp bond yield*	Bloomberg	1990.02–2019.06	Daily
10-year On-the-run/Off-the-run spread	Adrian et al. (2017)	1991.6–2019.06	Monthly
c. Economic Variables			
Consumer Price Index	`fred.stlouisfed.org`	1954.07–2019.06	Monthly
Industrial Production Index	`fred.stlouisfed.org`	1954.07–2019.06	Monthly
Total Checkable Deposits	`fred.stlouisfed.org`	1959.01–2019.06	Monthly
Total Savings Deposits	`fred.stlouisfed.org`	1975.1–2019.06	Weekly
Consumer Price Index	`fred.stlouisfed.org`	1954.07–2019.06	Monthly
Equity Bond Premium	Zakrajsek et al. (2016)	1973.01–2019.06	Monthly
c. Instrument			
Target Factor (30-minute window)	Jarociński and Karadi (2020)	1990.1–2019.06	Daily

Notes: At maturities: $^*Z = \{0.25, 0.5, 1, 2, 5, 10\}$.

Liquidity premia are defined as the difference in yield between two equally risky assets whose liquidity differs. Please note that we take monthly averages of higher frequency data when necessary. We use four measures of liquidity:

1. Our main measures of liquidity are the spreads between Resolution Funding Corporation (Refcorp) bond yields and Treasury zero-coupon bond yields. As argued by Longstaff (2002), Refcorp bonds are unique because their principal is fully collateralized by Treasury bonds. Thus, Refcorp bonds carry the same credit risk as Treasury bonds. Since Treasury bonds are more liquid, comparing their prices to those of Refcorp bonds serves as an ideal way of capturing liquidity premia.

2. Following Krishnamurthy and Vissing-Jorgensen (2012), we make use of the difference between Moody's Seasoned AAA Corporate Bond Yield (mnemonic `AAA`) and the 10-Year Treasury Constant Maturity Rate (`GS10`), which are both available on the Federal Reserve Bank of St. Louis' FRED database.

3. Alternatively, as suggested by Nagel (2016), we make use of the spread between the 3-Month General Collateral Repurchase Agreement (GC repo) Rate and the 3-Month Treasury Constant Maturity Rate. The former is taken from Bloomberg (mnemonic `USRGCGC ICUS Curncy`) and is computed as the midpoint between the bid and ask rates, and the latter is obtained from FRED (`GS3M`).

4. Another measure of the liquidity premium is the spread between on-the-run and off-the-run Treasury securities. The securities we use have a 10-year maturity. See Adrian et al. (2017) for additional details.

Economic variables entering the VAR are monthly and come from the Federal Reserve Bank of St. Louis' FRED database. These include the 1-year

Treasury yield (mnemonic `GS1`), the Consumer Price Index (`CPIAUCSL`), the Industrial Production Index (`INDPRO`), the total checkable deposits (`TCDSL`), and the total savings deposits (`SAVINGSL`). The equity bond premium is derived from Zakrajsek et al. (2016).

The updated series of the first principal component of monetary policy surprises for a 30-minute window around the time of FOMC announcements was kindly provided to us by Jarociński and Karadi (2020) based on an updated dataset of that used in Gürkaynak et al. (2005).

3.4 Results

Figure 3 plots the estimated structural IRFs (black solid lines) together with the respective 95 percent bootstrapped CIs (black dashed lines). Each subplot therefore shows the response of the abovementionned variable to a one-standard deviation monetary policy shock (i.e., a monetary policy tightening). The red dotted lines show the IRFs stemming from the well-known recursive identification, which places economic variables first (log CPI, log IP, log deposits), and financial variables second (1-year rate, the EBP, and liquidity premia). We included the latter to assess the robustness and the advantage of the identification through external instruments.

A review of the IRFs obtained through the instrumental approach reveals that a one-standard deviation positive shock to the target factor generates a response of the 1-year rate of approximately 20 basis points that disapears within two years.[5] This monetary tightening triggers a response from the

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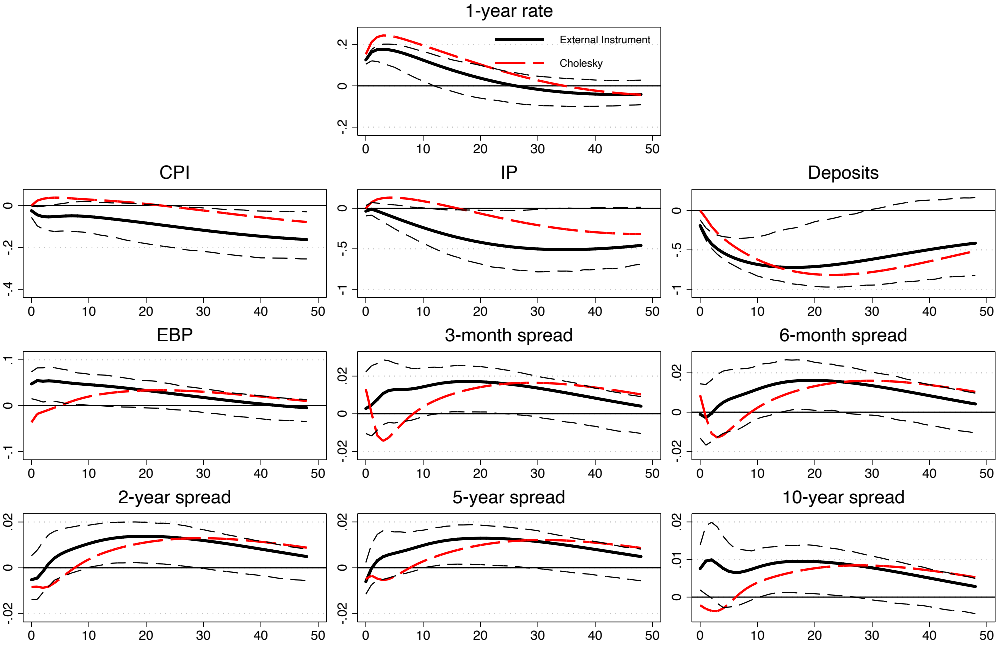
[5] Recall that the target factor is rescaled so as to match units with the first monetary policy surprise, mp_t^1.

Figure 3: IRFs to Monetary Policy Shocks



Impulse: 1-year rate. **Instrument:** Target Factor. **First Stage:** F = 17.61, R^2 = 5.04, N = 334.
CIs: Recursive wild bootstrap with 1000 replications, 0.95 level. **Set Up:** 2 lags, 1991.6 - 2019.5.
Cholesky Order: 1y rate, CPI, IP, Deposits, EBP, 3m spread, 6m spread, 2y spread, 5y spread, 10y spread.

Notes: Each subgraph plots the IRF of the variable mentioned above to a one standard deviation surprise monetary policy tightening (solid black line) together with the CIs surrounding it (dashed black lines) and the Cholesky-identified IRF (red dashed line). See below the figure for more details.

economy consistent with theory: i) a significant and delayed decline in the CPI level (approximately 20 bp) and ii) a significant decrease in output (proxied by industrial production) within a year following the shock peaking after around two years (approximately 50 bp.).

Regarding log deposits, the excess bond premium and liquidity premia, the SVAR estimation corroborates the mechanisms behind both the deposit and credit channels of monetary policy. First, the significant increase in the EBP (approximately 50 bp), echoing Gertler and Karadi (2015), provides evidence that monetary policy tightening deteriorates general credit conditions for up to one year following the shock. Second, the significant long-lasting decrease in log deposits (approximately 70 bp) coupled with the significant increase in the liquidity premium at all maturities supports the mechanism theorized in Drechsler et al. (2018).

When the monetary policy interest rate increases, rates paid on banks' customer deposits do not adjust proportionally, leading to a decrease in aggregate banks' customer deposits: thus, the liquidity value of Treasuries increases. Most importantly, the response of the liquidity premia across maturities is characterized by a decreasing but significant relationship. Indeed, the longer the maturity is, the less the premium increases, as longer-term Treasuries are less liquid and are discounted more heavily when interest rates rise. The responses of the liquidity premia are significant 1 to 2 years after the policy shock. The response of the short-term (3-month) spread is approximately one-tenth of the policy shock, whereas the response of the long-term (10-year) spread is approximately one-fifth of the shock.

We obtain a similar overall result when we consider the Cholesky-identified IRFs. It is worth noting that the shock under study is no longer identified using the instrument, as we rely on the ad hoc assumptions that a one-standard deviation positive shock to the federal funds rate triggers a contemporaneous response of the financial variables included in the SVAR (the EBP and the premia) but that it only affects the economy (the CPI and industrial production) with a lag.

One noticeable strength of the instrumental approach is that it eradicates the well-documented price puzzle. Furthermore, it produces a significant response for all the spreads (which would likely be nonsignificant otherwise).

3.5 Robustness

Excluding the 2008 financial crisis.—As in Gertler and Karadi (2015), we estimate the same model as in section 3.4 while excluding the 2008 financial crisis. In particular, we exclude the period between 2008.7 and 2009.6. The results are quantitatively similar but somewhat less significant due to the loss of observations (12 months plus lags).

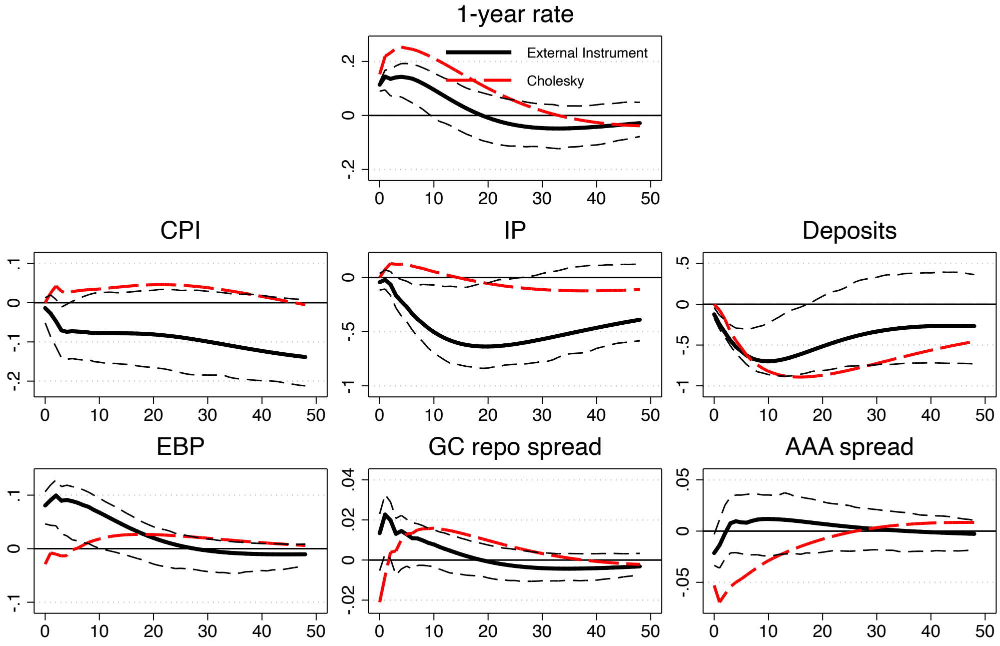
Nonetheless, the chore implications of our benchmark specification provided in section 3.4 remain valid and the IRFs are significant at the 90 percent level.

Using other measures of the liquidity premium.—Additionally, we estimate the same model as in section 3.4 using different measures of liquidity premia.

First, as a measure of short-term liquidity, we make use of the spread between the 3-Month General Collateral Repurchase Agreement (GC repo) Rate and the 3-Month Treasury Constant Maturity Rate as suggested by Nagel (2016). For the measure of liquidity for longer maturities, we follow Krishnamurthy and Vissing-Jorgensen (2012) and use the difference between Moody's Seasoned AAA Corporate Bond Yield and the 10-Year Treasury Constant Maturity Rate.

The results of this alternative specification are shown in Figure 4. The responses of the IP, the CPI, deposits and the EBP remain significant and robust. The GC repo spread increases significantly (approximately 2 bp) on impact, while the AAA spread significantly decreases (approximately 2 bp). Hence, this approach leads to a similar pattern according to which long-maturity liquidity premia react less to monetary policy shocks.

Figure 4: IRFs to Monetary Policy Shocks, GC Repo Spread and AAA Spread



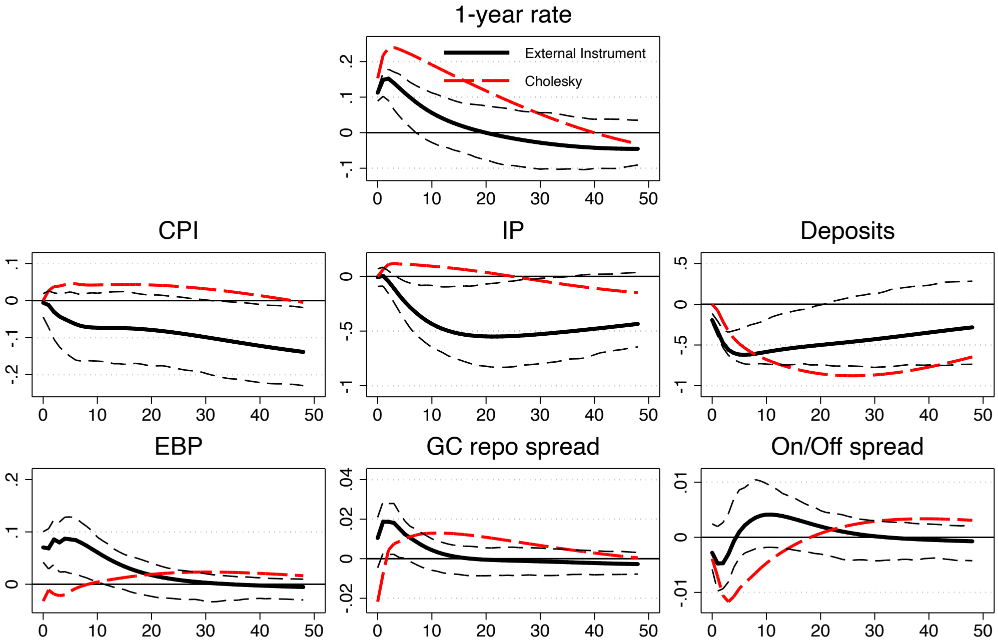
Impulse: 1-year rate. **Instrument:** Target Factor. **First Stage:** F = 11.81, R^2 = 3.3, N = 348.
CIs: Recursive wild bootstrap with 400 replications, 0.95 level. **Set Up:** 4 lags, 1990.2 - 2019.5.
Cholesky Order: 1y rate, CPI, IP, Deposits, EBP, GC repo spread, AAA spread.

Notes: Each subgraph plots the IRF of the variable mentioned above to a one standard deviation surprise monetary policy tightening (solid black line) together with the CIs surrounding it (dashed black lines) and the Cholesky-identified IRF (red dashed line). See below the figure for more details.

Another common measure of the liquidity premium is the spread between on-the-run and off-the-run Treasury securities (see Adrian et al. (2017) for more details). The securities we use have a 10-year maturity, so we include them in the SVAR instead of the AAA spread.

Figure 5 shows the results of this specification. While the GC repo spread significantly increases on impact (approximately 2 bp), the on/off spread somewhat decreases but remains insignificant following the monetary policy shock, which reinforces the view that longer-term liquidity premia react relatively less to monetary policy.

Figure 5: IRFs to Monetary Policy Shocks, GC Repo Spread and On/Off Spread



Impulse: 1-year rate. **Instrument:** Target Factor. **First Stage:** F = 12, $R^2 = 3.5$, N = 334.
CIs: Recursive wild bootstrap with 400 replications, 0.95 level. **Set Up:** 3 lags, 1991.6 - 2019.5.
Cholesky Order: 1y rate, CPI, IP, Deposits, EBP, GC repo spread, On/Off spread.

Notes: Each subgraph plots the IRF of the variable mentioned above to a one standard deviation surprise monetary policy tightening (solid black line) together with the CIs surrounding it (dashed black lines) and the Cholesky-identified IRF (red dashed line). See below the figure for more details.

4 Conclusion

We have estimated the macrodynamic effects of monetary policy on the yield curve through changes in liquidity premia. When the Fed raises interest rates, the spread between less-liquid assets and Treasuries of the same maturity and risk increases, which is significant 1 to 2 years after the policy shock. The longer the maturity is, the lesser—but still significant—the increase in the spread is.

Our empirical results point to the need to explicitly model the process of obtaining deposits from liquid assets such as Treasuries to account for the different liquidity premia at different maturities. Moreover, our results should lead to a better understanding of the expectation hypothesis, which accounts for the fact that monetary policy affects the term structure through liquidity premia, and of real equilibrium interest rates and exchange rate fluctuations, which are substantially influenced by liquidity premia fluctuations according to recent research.

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A Methodology

In this appendix, we provide additional details regarding the methodology. We expose how one can formally extract changes in expectations about future monetary policy using futures data and how one can rotate the resulting factors into interpretable dimensions of the Fed's conduct through its monetary policy.

Identifying Expectations Shocks.—Let us denote by $ff^1_{t-\Delta t}$ as the settlement rate implied by the federal funds rate futures contract expiring within the month, Δt days before a scheduled FOMC announcement. By construction of the futures contracts, which pay off according to the average effective federal funds rate prevailing over the agreed-upon month, $ff^1_{t-\Delta t}$ partly reflects the rate realized thus far in the given month, r_0, and the expected rate to prevail until the end of the period thereof, r_1. Accordingly, by denoting d^1 as the day of the month on which an FOMC meeting will take place and D^1 as the number of days in that same month, we have the following:

$$ff^1_{t-\Delta t} = \frac{d^1}{D^1} r_0 + \frac{D^1 - d^1}{D^1} E_{t-\Delta t}[r_1] + \rho^1_{t-\Delta t}, \tag{16}$$

where ρ^1 accounts for any (risk, term or liquidity) premium present in the contract. Assuming that there is no systematic change in premium ρ^1 within an FOMC announcement day, the surprise associated with a change in the federal funds target rate, $E_t[r_1] - E_{t-1}[r_1]$, is

$$mp^1_t = \frac{D^1}{D^1 - d^1}(ff^1_t - ff^1_{t-1}). \tag{17}$$

Similarly, one can measure the change in expectation, mp^2_t, regarding the rate that will prevail after the next FOMC meeting, r_2, by examining the futures of corresponding maturity. Because there are eight scheduled meetings per year, the next meeting arises within the next two months.[6] By denoting m as the number of months separating the current meeting from the next, it follows that:

$$ff^{1+m}_{t-\Delta t} = \frac{d^{1+m}}{D^{1+m}} E_{t-\Delta t}[r_1] + \frac{D^{1+m} - d^{1+m}}{D^{1+m}} E_{t-\Delta t}[r_2] + \rho^{1+m}_{t-\Delta t}, \qquad (18)$$

where the superscript i in $ff^i_{t-\Delta t}$ indicates the number of months from $t-\Delta t$ within which the futures due date occurs. The change in expectation is therefore characterized by:

$$mp^2_t = \frac{D^{1+m}}{D^{1+m} - d^{1+m}} \left[(ff^{1+m}_t - ff^{1+m}_{t-1}) - \frac{d^{1+m}}{D^{1+m}} mp^1_t \right]. \qquad (19)$$

There are two particular cases one needs to account for. First, when a meeting happens late in a month, the weight given to the surprise is relatively high. To prevent from the potential noise in the data from affecting our measurement, when a meeting occurs within the last seven days of a month, we take the unweighted change in the next month's futures price as the monetary policy surprise. Second, for meetings taking place on the first day of a month, we make use of the unweighted price difference between the federal funds futures rate due in the month of the meeting and that due in the previous month.

Finally, for the remaining contracts, namely, the 6-, 9- and 12-month Eurodollar futures (with a price denoted by ed^i_t), one can directly take the

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[6] NOte that Gürkaynak et al. (2005) assume unscheduled meetings to be expected as happening with zero probability.

daily return as the surprise itself due to their spot settlement nature. Thus, for $j = 4, 5, 6$ and $i = 6, 9, 12$ respectively, we have

$$mp^j = ed_t^i - ed_{t-1}^i. \tag{20}$$

Extracting the target and the path factor.—Let X be a $(T \times n)$ matrix whose entries correspond to the above-defined monetary policy surprises mp_t^j for $j = 1, ..., n$ and $t = 1, ..., T$, that is, the surprise component of the daily change in federal funds futures and Eurodollar futures rates solely associated with FOMC announcements.[7]

Let us assume X to be generated by the following factor model:

$$X = F\Lambda + \nu, \tag{21}$$

where F is a $(T \times \ell)$ matrix of $(\ell < n)$ unobserved factors, Λ is a $(\ell \times n)$ matrix of factor loadings, and ν is a matrix of orthogonal disturbances. Gürkaynak et al. (2005) show that the response of futures prices is sufficiently characterized by two factors (i.e., $\ell = 2$). We therefore estimate $F = \{F_{1t}, F_{2t}\}_{t=1,...,T}$ through principal-component analysis.

Because the two factors yielded by this method are chosen to maximize the share of explained variance, they lack structural interpretation. Gürkaynak et al. (2005) rotate F_1 and F_2 to obtain Z_1 and Z_2. Namely, we define:

$$Z = FU, \tag{22}$$

such that U is a (2×2) orthogonal matrix with Z_2 being associated, on

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[7] We normalize the columns of X so that they have zero mean and unit variance.

average, with no change in the federal funds futures rate for the current month. To recover an interpretation as to the magnitude of these factors, we rescale Z_1 (Z_2) to match its units with mp^1 (mp^4).

According to Gürkaynak et al. (2005), this rotation allows us to see Z_1 and Z_2 as the *target* factor and the *path* factor, respectively. This is the case because Z_1 is defined such as to drive surprises in the current target rate on FOMC announcement days, while Z_2 reflects everything (unrelated to the federal funds target rate) that causes changes to expectations of future monetary policy. Next, we describe the way in which the target factor serves as an external instrument for the estimation of our structural VAR.

Rotation of the Factors.—Gürkaynak et al. (2005) rotate F_1 and F_2 to obtain Z_1 and Z_2. Namely, they define

$$Z = FU, \tag{23}$$

where

$$U = \begin{bmatrix} u_{11} & u_{12} \\ u_{21} & u_{22} \end{bmatrix}', \tag{24}$$

such that U is a (2×2) orthogonal matrix, with Z_2 being associated, on average, with no change in the federal funds futures rate for the current month. The orthogonality between Z_1 and Z_2 requires that:

$$E(Z_1 Z_2) = u_{11}u_{12} + u_{21}u_{22} = 0. \tag{25}$$

Then, because:

$$F_1 = \frac{u_{22}Z_1 - u_{12}Z_2}{u_{11}u_{22} - u_{12}u_{21}}, \tag{26}$$

$$F_2 = \frac{u_{21}Z_1 - u_{11}Z_2}{u_{12}u_{21} - u_{11}u_{22}}, \tag{27}$$

one can assume that Z_2 has no impact on mp^1 by imposing the final restriction:

$$\lambda_2 u_{11} - \lambda_1 u_{12} = 0, \tag{28}$$

where λ_1 and λ_2 are the loadings on mp^1 of F_1 and F_2, respectively. To recover an interpretation as to the magnitude of these factors, one can rescale Z_1 (Z_2) to match its units with mp^1 (mp^4). Rotation matrix U is obtained by solving the last four equations.

In the words of Gürkaynak et al. (2005), this rotation allows us to see Z_1 and Z_2 as the *target* factor and the *path* factor, respectively. This is the case because Z_1 is defined such as to drive surprises in the current target rate on FOMC announcement days, while Z_2 reflects everything (unrelated to the federal funds target rate) that causes changes to expectations of future monetary policy.

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