Monetary Policy Spillovers and the Global Secular Decline in Long-Term Interest Rates*

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Abstract

The global secular decline in long-term interest rates since the 1990s can be largely linked to the cumulative spillover effects of U.S. monetary policy. U.S. monetary policy news, reflected in yield changes during the three-day windows around U.S. monetary policy announcements, can account for the bulk of the decline in long-term government bond yields across advanced economies over the past three decades. In contrast, monetary policy announcements by other advanced economy central banks have not produced lasting effects on long-term yields domestically or abroad. The effect works mainly through the expected path of future short-term interest rates rather than through term premia or exchange rate adjustments. It appears to be driven by a "Fed response to news" channel as the cumulative yield declines of global long-term rates are concentrated in FOMC windows preceded by US macro news announcements.

Keywords: Monetary policy, bond yields, interest rate trends, global financial cycle.

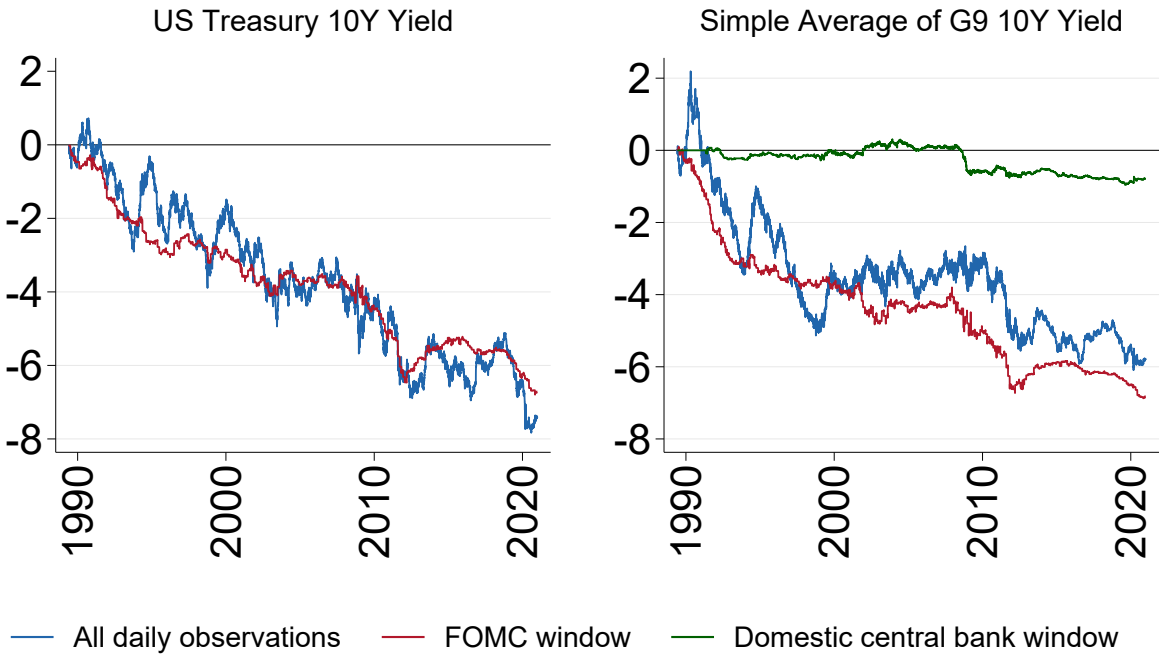
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1 Introduction

Long-term interest rates in advanced economies have co-moved closely over recent decades. In particular, they displayed a common secular decline since the 1990s, which is often attributed to global long-term structural factors such as falling productivity growth, excess savings, demographic decline, and reduced capital investment opportunities (see Bernanke (2005), Carvalho et al. (2016), Summers (2014).)

Figure 1: Cumulative yield changes in the U.S. and other countries (%).



Notes. "All daily observations" is the sum of all daily yield changes over the sample period. "FOMC window" is the sum of daily yield changes within the 3-day FOMC windows. "Domestic central bank window" is the sum of daily yield changes within the 3-day announcement window of each country's domestic central bank.

In this paper, we present new evidence suggesting that cumulative international spillovers of U.S. monetary policy decisions may have been an important factor driving the global decline in long-term rates. Our analysis shows that only U.S. FOMC announcement windows exhibit a strong downward shift in yields and a significant pass-through to international interest rates, whereas other countries' central bank windows

have limited effects on long-term rates both domestically and internationally. Figure 1 summarizes this key result, plotting the cumulative daily changes in 10-year government bond yields in the United States and the simple average of yields in the remaining G9 currencies. In both panels, 10-year yields have trended downward by about 6 percentage points over the past 30 years. As suggested by Hillenbrand (2025), the left panel shows that the secular decline in U.S. government bond yields is closely associated with the cumulative yield changes that occur only during the three-day windows surrounding U.S. Federal Open Market Committee (FOMC) monetary policy announcement dates (FOMC windows). More importantly, the right panel shows that the cumulative changes of G9 countries' government bond yields are also well explained by yield movements during the FOMC windows—indeed, even more so than by those during the three-day windows surrounding their own domestic central bank monetary policy announcements.

The further analysis in the paper explores the key mechanisms underlying this result. Specifically, we analyze the strong downward shift in yields and the significant pass-through to international interest rates during the FOMC windows in two steps. (1) We examine the pass-through mechanism from the FOMC windows to foreign yield changes. (2) We investigate the factors driving yield declines during the FOMC windows.

We first explore the international spillovers of U.S. monetary policy using various decompositions. Significant spillovers from U.S. yields to other countries' yields are documented during FOMC announcement windows through an interest parity decomposition, a term-structure decomposition, and a Granger causality test for the policy rate.

The baseline Mundell–Fleming model suggests that a country can maintain monetary autonomy when its exchange rate is flexible to adjust in response to foreign shocks. Using an uncovered interest parity (UIP) decomposition that allows for deviations from UIP, the analysis shows that, during the FOMC windows, changes in U.S. long-term interest rates have a persistent and nearly one-to-one passthrough to other countries' interest rates. Exchange rates, however, do not respond systematically to U.S. interest rate movements. This finding contradicts the predictions of the Mundell–Fleming model and suggests that global long-term interest rates are influenced by U.S.-centric common factors—a trend-level analogue of the global financial cycle, as suggested by Rey (2013).

In the next step of our assessment of the mechanisms underlying the international U.S. monetary policy spillovers, we investigate whether they work primarily through expected short-term rates or term premia.

To this end, we decompose daily yield changes into expected short-rate and term-premium components using the term structure model of Bauer and Rudebusch (2020). The results show that international yield movements during FOMC windows are largely driven by revisions in the path of expected short-term rates, while the contribution of term-premium changes is limited. This evidence suggests that U.S. monetary policy announcements trigger persistent adjustments in market expectations about future policy rates—not only in the United States but also across other advanced economies.

Given the major variation in expected short-term rates in the wake of U.S. monetary policy announcements, we further corroborate the international transmission of U.S. monetary policy by examining the directional link between policy rates across countries. Specifically, we analyze the Granger causality relationship between the U.S. policy rate and those of other advanced economies. The results reveal a clear asymmetry: changes in the U.S. policy rate systematically Granger-cause movements in foreign policy rates, while no foreign central bank's policy rate Granger-causes the U.S. rate. This asymmetry underscores the dominant role of U.S. monetary policy in shaping global interest rate dynamics and reinforces the interpretation that FOMC meetings serve as pivotal events for the global yield curve.

The pass-through analysis naturally raises the question of why the declines in U.S. and global long-term yields are so strongly concentrated in FOMC announcement windows. To shed light on this concentration, we examine whether the pattern reflects systematic interactions between macroeconomic news and subsequent monetary policy decisions.

We find that the downward shifts in yields occur predominantly during a subset of FOMC meetings that are preceded by major U.S. macroeconomic announcements—namely, CPI and GDP releases. We refer to these as macro-news FOMC windows. When we partition FOMC meetings into those preceded by macro releases and those that are not, we find that nearly all of the cumulative yield decline—over 95 percent in the case of U.S. 10-year Treasuries—originates from macro-news FOMC windows.

The decline that occurs during the "macro news FOMC windows" is closely associated with Bauer and Swanson (2023a)'s "Fed response to news" component of high-frequency monetary policy surprises—that is, the portion of monetary policy surprises that is ex post explainable by macro variables known before the shock. In contrast, we find that several alternative measures of information effects—including macro announcement days themselves, FOMC announcements associated with the "dot plot," and Fed Chair speeches—cannot account for the secular decline in long-term interest rates.

If long-run macroeconomic conditions alone determined long-term interest rates, regardless of central bank decisions, then the macro announcements themselves should be the most informative about the long-run trend. Conversely, if the decline were purely driven by releases of the central bank's private information about underlying macroeconomic conditions, the meetings with economic projections and the Fed Chair's speeches should be more informative about the long-run trend. The empirical pattern we document thus suggests that it is the specific nexus between updates of macroeconomic information and the subsequent FOMC actions or interpretations of the news that contributes to the decline.

Literature review. Since the pioneering studies of Kuttner (2001) and Gürkaynak et al. (2005), a large empirical literature has employed high-frequency financial data to identify monetary policy shocks and trace their effects on asset prices and the real economy. Subsequent contributions such as Bernanke and Kuttner (2005) and Gertler and Karadi (2015) established the macroeconomic importance of these shocks in shaping output and inflation dynamics. More recent work—exemplified by Hanson and Stein (2015) and Nakamura and Steinsson (2018)—provides compelling evidence that monetary policy is far from neutral. Refinements in the identification and interpretation of policy shocks have been advanced by Bauer and Swanson (2023b), Bauer and Swanson (2023a), Jarociński and Karadi (2020), Altavilla et al. (2019), Bu et al. (2021), Miranda-Agrippino and Ricco (2021) Acosta (2023), and Boehm and Kroner (2024). Swanson (2021) apply the technique to studies a broader set of monetary policy, including forward guidance and quantitative easing. Collectively, this literature has substantially deepened our understanding of how monetary policy decisions—and the information they convey—transmit through financial markets and the global economy.

Our paper is closely related to the work of Hillenbrand (2025) and Alam (2023). Hillenbrand (2025) demonstrate that the three-day window surrounding FOMC monetary policy announcement dates accounts for much of the secular decline in long-term U.S. interest rates. Alam (2023) focus on the pre-FOMC drift and the FOMC equity premium during macro news announcements and document a similar secular decline in these FOMC windows. We make three key contributions to this literature. First, we examine the monetary policy announcement effects across multiple central banks and show that other central banks do not produce comparable patterns. Second, we demonstrate that the underlying mechanism is consistent with a "Fed responses to news" interpretation and directly link the FOMC window decline to the corresponding monetary policy surprises.[1] Third, we offer several decompositions of the relationship between monetary

[1]Relatedly, a growing literature explains yield movements through fiscal events. Studies include Treasury issuances (Lou et al. (2022)), budget announcements (Gomez Cram et al. (2025)), and Treasury auctions (Somogyi et al. (2025)). We comple-

policy announcements and the secular decline in interest rates. For example, we show that the yield decline is primarily associated with FOMC windows preceded by macro news and changes in expected short-term rates.

This paper further contributes to the extensive literature on monetary policy spillovers. Much of the existing research focuses on the impact of U.S. policy on real economic variables. Early studies, such as Calvo et al. (1993) and Mackowiak (2007), examine the spillover effects on emerging economies, while Kim and Roubini (2001) and Kim (2007) emphasize the effects on advanced economies. The global financial cycle literature, pioneered by Rey (2013), identifies U.S. monetary policy shocks as a critical global factor driving risky asset prices worldwide. This work has been expanded by important studies including Albagli et al. (2019), Dedola et al. (2017), Gilchrist et al. (2019), Miranda-Agrippino and Rey (2020), and Brusa et al. (2020), which focus on the Federal Reserve. Boehm and Kroner (2025) highlights the role of macro news announcements in driving the global financial cycle. Additionally, Gerko and Rey (2017), Jarociński (2022), Miranda-Agrippino and Nenova (2022), and Kearns et al. (2023) highlight the central roles of the Federal Reserve, the European Central Bank, and the Bank of England. Jordà et al. (2024) document that the effects of monetary policy can be very long-lasting, persisting for decades. The literature has focused on examining how identified shocks drive short-term fluctuations in asset prices. This paper introduces a novel perspective by investigating the association between monetary policy and long-term trends in safe asset prices—a phenomenon we term the "global safety trend". Importantly, we show that the yield changes during the FOMC windows can almost entirely account for the secular changes in world long-term interest rates.

The literature examining the secular decline in the natural rate of interest has evolved from initial estimation efforts to broader explorations of its structural determinants. Foundational empirical frameworks developed by Laubach and Williams (2003), Holston et al. (2017), Hamilton et al. (2016), and Del Negro et al. (2017) established that natural rates have trended downward across advanced economies since the 1990s. More recent work extends beyond measurement to investigate the underlying mechanisms shaping this decline.

A growing strand of research emphasizes that monetary and financial factors themselves may exert long-term influences on r^* Borio et al. (2017), McKay and Wieland (2021), and Rungcharoenkitkul and Winkler

ment this literature by focusing on the interaction between macro news and monetary policy announcements, as well as their implications for cross-country yields.

(2022) demonstrate that persistent monetary accommodation, financial frictions, and policy non-linearities can anchor equilibrium real rates at lower levels than would be implied by fundamentals alone. These findings challenge the traditional view that r^* is purely exogenous to monetary policy, opening up the possibility of feedback between policy regimes and structural interest rate trends. Another key driver arises from changes in productivity and technological progress. Faster productivity growth raises the marginal product of capital, increasing equilibrium rates as firms demand more investment (Cesa-Bianchi et al. (2022); Mankiw (2022)). The prolonged slowdown in productivity growth observed across advanced economies has therefore been a central factor behind the sustained decline in r^*. Financial globalization further shapes equilibrium interest rates through both capital flows and asset supply. The scarcity of safe assets—particularly U.S. Treasuries—has increased their valuation and lowered real yields (Bernanke (2005); Caballero et al. (2008); Del Negro et al. (2019); Bárány et al. (2023); Krishnamurthy and Vissing-Jorgensen (2012)). At the same time, the growing integration of emerging markets into global capital markets offers alternative investment destinations, exerting upward pressure on r^* in advanced economies (Obstfeld (2023)). Finally, demographic trends exert profound and persistent effects on saving behavior and potential growth. Changes in fertility, longevity, and dependency ratios influence aggregate saving for retirement and the balance between desired saving and investment, thereby shifting the natural rate over time (Carvalho et al. (2016); Gagnon et al. (2021)).

Taken together, this literature underscores that the secular decline in natural rates reflects an interplay of structural, financial, and policy-driven forces. Against this backdrop, the present study contributes new empirical evidence highlighting the role of U.S. monetary policy transmission in shaping the long-run trajectory of global equilibrium interest rates.

The remainder of the paper is organized as follows. Section 2 describes the sovereign yield data and central bank event dates. Section 3 presents our baseline empirical results and compares the international effects of various central bank announcements on international sovereign yields. Section 4 investigates how U.S. monetary policy announcements spill over to interest rates in other countries. Section 5 investigates the channels through which FOMC announcements may have driven the secular declines in the U.S. long-term yield which then spilled over internationally. Section 6 conducts robustness checks examining the statistical significance of cumulative effects of FOMC announcements. Section 7 concludes.

Table 1: Sources of daily sovereign yields.

Country	Abbreviation	Source	Start
Australia	AUD	Bloomberg	Apr 16, 1991
Canada	CAD	Bloomberg	Jun 25, 1991
Switzerland	CHF	Bloomberg	Feb 25, 1994
Germany	EUR	Bloomberg	Oct 03, 1991
U.K.	GBP	Bank of England	Jan 02, 1979
Japan	JPY	Bloomberg	Sep 30, 1992
Norway	NOK	Bloomberg (after 1998) & `Investing.com`	Feb 18, 1994
New Zealand	NZD	Bloomberg	Mar 9, 1992
Sweden	SEK	Bloomberg	Feb 25, 1994
U.S.A.	USD	Federal Reserve Board (GSW)	Jun 5, 1989

Notes: The plots start from the first FOMC date (June 5, 1989) or the first date the yields become available, whichever is later. GSW refers to Gürkaynak et al. (2007). The starting dates are those for the 10-year yield series.

2 Data

Daily Sovereign Yields. Our main source of daily sovereign yields is Bloomberg, following Du et al. (2018).[2] Our analysis focuses on the 10-year yield, but our findings extend to other maturities. Moreover, we rely on the entire yield curve data to estimate changes in the expected interest rates and term premia during monetary policy announcement windows. The Federal Reserve Board and the Bank of England provide daily sovereign yields on their websites with starting dates earlier than the Bloomberg data. In this case, we use the yield curve data provided by the central bank. Table 1 summarizes the sources of daily sovereign yields and the respective starting dates. For each country, the Bloomberg yield curve dataset contains maturities of 3 months, 6 months, and annual maturities of 1, 2, 3, 4, 5, 7, and 10 years. Gürkaynak et al. (2007) (GSW) covers maturities from 1 year to 30 years with yearly increments. We augment the GSW yield curve data with 3-month and 6-month interest rate data from the FRED. For the U.S., we select the same maturities as Bloomberg to be consistent with other countries. For Norway, while the Bloomberg dataset begins in 1998 and lacks observations between November 2012 and April 2014, `Investing.com` provides daily Norwegian 10-year yield data dating back to February 18, 1994. We complement the Bloomberg data with this source to create a longer and more complete time series. The sample ends on Aug 21, 2023.

FOMC Announcement Dates. The Federal Open Market Committee (FOMC) makes U.S. monetary

[2]We focus on the G10 currencies, including the Australian Dollar, Canadian Dollar, Swiss Franc, Euro, British Pound, Japanese Yen, Norwegian Krone, Swedish Krona, and U.S. Dollar. We use German bond yields to represent the Euro area. Our analysis centers on the 10-year bond yield, the longest maturity widely available across currencies.

policy decisions. Since 1981, the FOMC has typically held eight scheduled meetings per year. Most monetary policy decisions since 1994 have been made during these *scheduled* meetings, while a few were made during *unscheduled* meetings. In contrast, unscheduled meetings accounted for a large fraction of changes in the federal funds rate before 1994. Some of the unscheduled meetings were not followed by immediate policy actions or a statement. The public learned about these meetings with a significant time lag. These meetings are excluded from the list of FOMC announcement dates.

In line with Hillenbrand (2025), our selection of FOMC announcement dates corresponds to the time when the public receives information about the meetings. Before 1994, changes in monetary policy were typically disclosed to the market one day after the meeting through open market operations. Therefore, for dates before 1994, we rely on the dates that the market associated with a monetary policy change, as identified by Kuttner (2001, 2003). Our first FOMC announcement window was in June 1989. After 1994, monetary policy decisions were predominantly made during scheduled FOMC meetings, with the Fed consistently releasing a statement if the federal funds rate changed. Consequently, we utilize the dates when these statements were released. In Appendix A, we list the FOMC announcement dates.

Other Central Banks' Monetary Policy Announcements. Whenever possible, we collect monetary policy announcement dates from the respective central bank's website. We supplement the calendar with records from `Investing.com` and verify that the latter coincides with the official central bank calendar when the two samples overlap. Most announcement series begin in 2000, with the latest starting date being February 3, 2000, for the Bank of Canada.

Fed Chair Speeches. We obtain the dates of the Federal Reserve Board Chair speeches from the Federal Reserve Board's official website. The website documents the dates of speeches by the Federal Reserve Board Chair, Vice Chair, and other officials from 1996 onwards.

3 Monetary Policy Announcements and Bond Yield Dynamics

In this section, we examine the behavior of 10-year government bond yields across G10 economies, distinguishing between windows around U.S. monetary policy announcements and those tied to domestic central banks. The key finding is that the long-run dynamics of G10 interest rates are largely accounted for by yield movements surrounding U.S. monetary policy decisions.

We examine the dynamics of the 10-year government bond yield during different event windows. The sample period is divided into three parts: US central bank monetary policy announcement windows (FOMC windows), domestic central bank monetary policy announcement windows (domestic CB windows) and non-monetary policy announcement windows (non-MP windows). We follow Hillenbrand (2025) to define monetary announcement windows as the days $T-1, T, T+1$ surrounding each monetary policy announcement date T, with non-monetary policy announcement windows encompassing the remaining days. For each event window, we construct the cumulative change in the 10-year government bond yield by summing daily yield movements. To preserve a longer sample, the analysis primarily relies on nominal yields.

Formally, we define the cumulative change series as

$$\nabla y_t^W = \sum_{s=t_0+1}^{t} (y_s - y_{s-1})\, \mathbf{1}_{s \in W}, \tag{1}$$

where t and s index daily observations, t_0 denotes the starting date of the sample, y_s is the log n-year Treasury zero-coupon yield at date s, and $\mathbf{1}_{s \in W}$ is an indicator function for the event set W. The set $W \in \{\text{FOMC, domestic CB, nonMP}\}$ corresponds to either FOMC policy monetary policy announcement window, the domestic central bank announcement window, or the complementary set of non-announcement days. The domestic cental bank announcement windows that overlap with FOMC windows are treated as FOMC windows, so the partition is mutually exclusive and exhaustive. The total yield change from the start of the sample can be decomposed into the contributions of the three windows:

$$y_t - y_0 = \nabla y_t^{AllDates} = \nabla y_t^{FOMC} + \nabla y_t^{domesticCB} + \nabla y_t^{nonMP}. \tag{2}$$

For simplicity, we normalize yields by subtracting their initial value so that $\nabla y_t^{AllDates}$ and y_t can be treated equivalently.

Quantitatively, we evaluate the model fit using a pseudo R^2 measure. The measure is computed as the ratio of the squared changes within an event window W relative to the sum of the squared changes within each of the three event windows:

$$pseudo\ R^{2\{W\}} \equiv \frac{\sum_{t=1}^{T} \left(\nabla y_t^W\right)^2}{\sum_{W' \in \{\text{FOMC, domestic CB, nonMP}\}} \left[\sum_{t=1}^{T} \left(\nabla y_t^{W'}\right)^2\right]} \tag{3}$$

The pseudo R^2 associated with the three event windows sum up to 100% for each country. Moreover, it is bounded between zero and one, and can be interpreted as analogous to the conventional R^2. The numerator represents the total variation that occurred within the window of interest, while the denominator captures the total variation across all windows that partition the sample space. For example, a larger pseudo $R^{2\{FOMC\}}$ indicates that greater variations occur during FOMC windows and the yield change is almost zero during other times, suggesting that the FOMC windows better explain the dynamics of the observed yields.

3.1 FOMC and Other Central Banks' Announcements

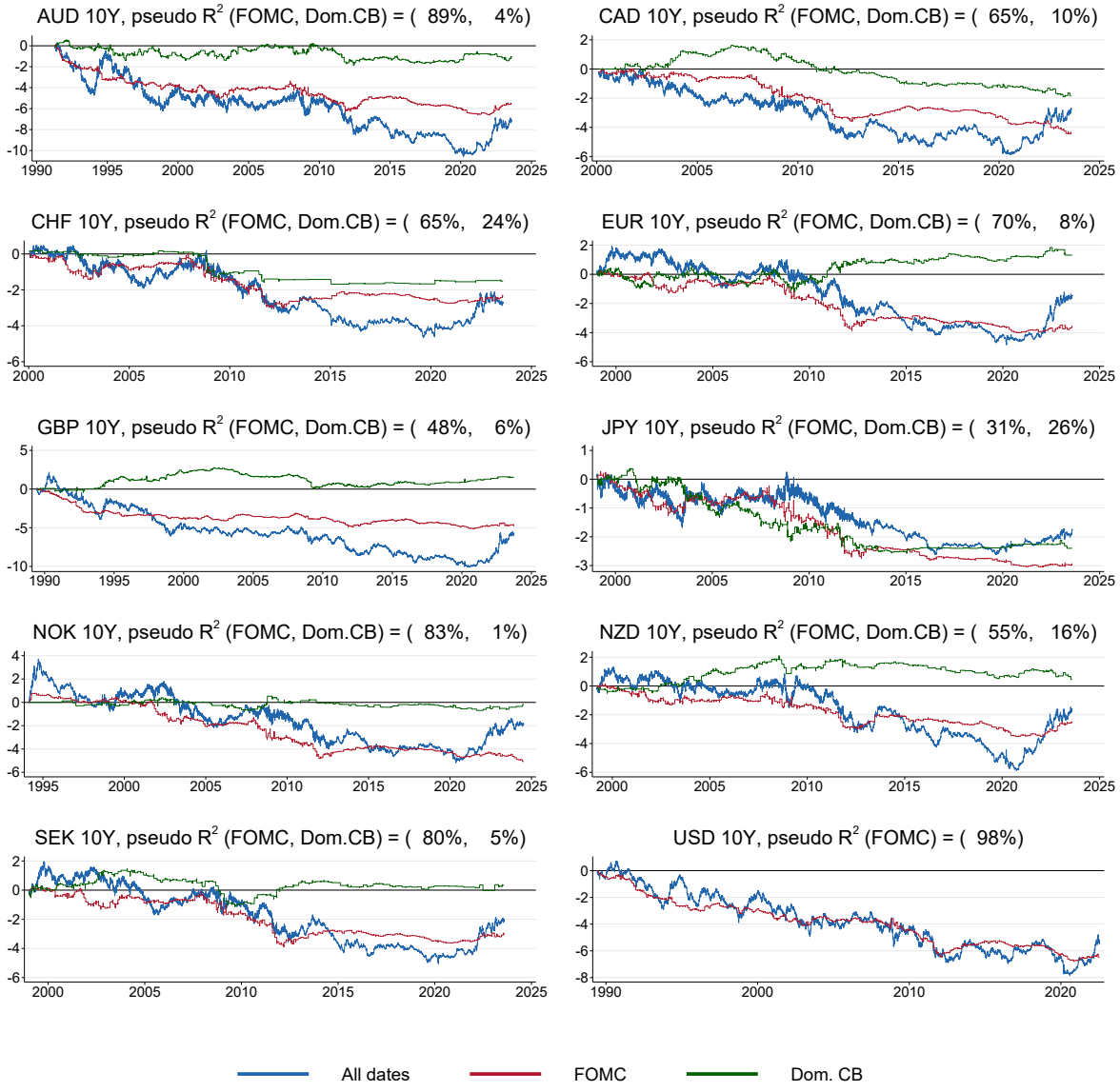
In Figure 2, we plot the cumulative sum of daily changes in 10-year government bond yields (blue lines), along with the corresponding cumulative changes during the FOMC window (red lines) and the domestic central bank window (green lines) for the G10 countries. The titles of the subfigures report the pseudo R^2 measures for the FOMC and domestic central bank windows.

To ensure a fair comparison between the FOMC window and the domestic window, we begin each country's sample when the domestic central bank's monetary announcement dates first become available in the data archive. Despite the slightly different start dates across countries, the first general observation from the blue lines in the figure is a consistent decline in the 10-year government bond rate. The secular decline starts at the beginning of the sample, with the largest drop occurring in Australia, reaching about –10% just before 2020. The average trough across countries is approximately –6%.

Focusing on the United States, and consistent with Hillenbrand (2025), the cumulative decline in the 10-year yield during FOMC windows closely mirrors the observed data, reaching a trough of about -6%. Importantly, the figure shows not only an endpoint-to-endpoint alignment but also that the path of the FOMC-window decline tracks the overall cumulative decline throughout the sample. The pseudo R^2 is 98%, indicating an exceptionally close fit.

Strikingly, the FOMC window also captures the overall decline in other countries' interest rates quite

Figure 2: Cumulative changes in nominal 10-year government bond yields during FOMC and other central banks' announcement windows.



Notes. Unit of the y-axis is percentage p.a.. "All dates" refers to the actual yield change relative to the initial value, "FOMC" is the cumulative yield change during FOMC windows, and "Dom. CB" is the cumulative yield change during domestic central bank announcement windows.

well. Take Australia as an example: the Australian 10-year yield shows a persistent decline starting in 1991, reaching about –6.5% by the end of the sample. The decline during FOMC windows is similarly persistent, amounting to roughly –6% by the end of the sample. By contrast, the decline during the Reserve Bank of Australia's announcement windows is much more limited, with a maximum cumulative drop of about –2% around 2012. Over time, its effect is largely negligible.

The case of Australia is not unique; the decline in its 10-year government rate, which is closely tracked by the FOMC window and whose domestic window decline is mostly flat, is similarly observed in Canada, Switzerland, Germany (EUR), Norway, New Zealand, and Sweden. In two cases, Germany and New Zealand, the cumulative changes in yield during the domestic central bank window are actually positive for an extended period of time, which is the opposite of the observed 10-year government rate pattern in the corresponding country.

There is one exception, which is that the FOMC window does not universally dominate the Bank of Japan. In Japan, the secular decline in interest rates is about –2% over the sample period. However, the cumulative declines during both FOMC windows and Bank of Japan announcement windows are similar, with the latter marginally tracking the observed yields more closely in the later part of the sample. This is perhaps unsurprising, given the yield curve control and various quantitative easing policies implemented in Japan over the past 30 years.

Overall, the FOMC series fit the actual yields quite well. The most impressive one is the U.S. 10-year yield. The pseudo R^2 indicates that the FOMC windows explain 98% of the total variation in the U.S. 10-year yield. Additionally, the FOMC windows demonstrate impressive explanatory power for yields in Australia, Canada, Switzerland, Germany, Norway, New Zealand, Sweden and accounting for two-thirds of the observed total variations in non-U.S. G10 10 year yield. On the other hand, the average pseudo R^2 for non-US domestic central bank is only 11%.

A key reason why the U.S. FOMC window accounts for a large share of the global yield decline is the persistent decrease in U.S. yields during FOMC windows, coupled with significant spillovers to other countries. This pattern is not observed for other central banks: either there is no persistent domestic yield decline during their announcement windows, or any such decline does not generate meaningful spillovers across borders.

We illustrate this point using two regression analyses. In Panel I of Table 2, we examine whether domestic

yields systematically decline during each central bank's own announcement window. Specifically, we estimate the following time-series regression for each country:

$$\Delta_3 y_{c,t} = \alpha_c + \beta_c \mathbf{1}_{c,t} + u_{c,t} \tag{4}$$

where $\Delta_3 y_{c,t}$ is the 3-day change in the 10-year yield in country c, and $\mathbf{1}_{c,t}$ is an indicator for whether t belongs to the monetary policy announcement window of country c. That is, $\mathbf{1}_{c,T-1} = \mathbf{1}_{c,T} = \mathbf{1}_{c,T+1} = 1$ if there is a monetary announcement of country c on date T. This regression is estimated using daily data from 2000 onward, the period for which we have complete monetary policy announcement dates for all countries in our sample.

The coefficient β_c indicates whether there is a systematic drift during central bank announcement dates relative to non-announcement dates. The results are consistent with the visual patterns shown in Figure 2. Not all central banks experience a systematic decline in domestic long-term yields during their announcement windows. Only the announcement dates of the Bank of Canada, the Bank of Japan, and the Norges Bank exhibit statistically significant declines in their domestic long-term yields akin to those observed during Fed announcements.

However, as shown in Panel II, only the Fed generates systematic spillovers to the yields of other countries. To test this, we estimate the following panel regression:

$$\Delta_3 y_{i,t} = \alpha_i + \sum_c \beta_c \mathbf{1}_{c,t} + u_{i,t}, \tag{5}$$

where $\Delta_3 y_{i,t}$ is the 3-day change in the 10-year yield for country i. We include a central bank monetary announcement indicator for *each* central bank and control for country-fixed effects. Standard errors are clustered at the event level.

Although some non-U.S. central bank windows are associated with domestic yield declines (panel I), only the Fed's announcement window has a statistically significant impact on yields across other countries (panel II). On average, the FOMC window is associated with an additional 0.87 basis point decline in foreign 10-year yields, relative to non-FOMC days. In contrast, none of the other central bank announcement windows have statistically significant effects on foreign yields. A joint test of the null hypothesis that all non-Fed central bank windows have zero effect on foreign yields produces weak evidence against the null (F=1.05 with a

Table 2: Average three-day yield changes during or outside monetary policy announcement windows.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
	I: $\Delta_3 y_{c,t} = \alpha_c + \beta_c \mathbf{1}_{c,t} + u_{c,t}$, country-by-country									
β_c	-0.29	-0.69**	-0.23	0.13	-0.02	-0.56**	-0.94***	0.34	0.40	-1.06***
se	(0.36)	(0.32)	(0.53)	(0.34)	(0.31)	(0.23)	(0.36)	(0.37)	(0.40)	(0.28)
	II: $\Delta_3 y_{i,t} = \alpha + \sum_c \beta_c \mathbf{1}_{c,t} + u_{i,t}$, panel regression									
β_c	-0.36	-0.05	0.66	-0.08	0.08	-0.39	-0.39	0.62*	0.55	-0.87**
se	(0.31)	(0.34)	(0.57)	(0.35)	(0.33)	(0.33)	(0.39)	(0.37)	(0.42)	(0.42)
	F-statistic for all non-US central bank coefficients being jointly zero: $F = 1.05$, $p = 0.39$.									

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The units are basis points per annum. Panel I reports the results from the panel regression $\Delta_3 y_{i,t} = \alpha + \sum_c \beta_c \mathbf{1}_{c,t} + u_{i,t}$, where $\Delta_3 y_{i,t}$ is the 3-day change in the 10-year yield in country, and $\mathbf{1}_{c,t}$ is an indicator for whether t belongs to the announcement window of country c. Each column reports $\hat{\beta}_c$ and se($\hat{\beta}_c$) of announcement window of country c. The last row reports the F-statistic and p-value for all non-U.S. central banks jointly having zero coefficients. The standard errors are clustered at the event level. Panel II reports the domestic effect of each country's monetary policy announcements from the country-by-country time series regression $\Delta_3 y_{c,t} = \alpha_c + \beta_c \mathbf{1}_{c,t} + u_{c,t}$.

p-value of 0.39).

We conclude this section with a 10x10 matrix analysis to understand how each central bank announcement window fits each country's yield in the sample. We repeat the exercise described by Equation (2) for each central banks' announcement dates, and then we investigate whether the announcement windows can explain interest rate trends internationally.

To facilitate comparison across different monetary policy announcement windows and different country yields, we use the root mean square error (RMSE) to summarize how well the cumulative sum of daily yield changes within each central bank's monetary policy announcement windows can fit the observed 10-year yield data. The RMSE for the cumulative sum of daily yield changes of country i within an event window W is defined as:

$$RMSE_i = \sqrt{\frac{1}{T}\sum_t \left(\nabla y_{i,t} - \nabla y_{i,t}^W\right)^2}. \tag{6}$$

Table 3 reports the RMSE for each country's monetary policy windows (each column) when fitting the 10-year yield of the corresponding row country. The table shows that, across all central bank announcement windows, the cumulative yield change during the FOMC window provides the best fit among all countries. The average RMSE for the FOMC column is 1.3%, while the RMSE for the other columns ranges from 2.5% to 5%.

Table 3: Fitting 10-year bond yields with the cumulative announcement effects of central banks.

Yields \ CB	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
AUD	3.50	3.42	3.63	2.88	3.56	4.33	2.01	5.29	2.59	2.11
CAD	4.19	2.39	2.01	2.68	2.62	5.05	2.94	5.13	2.52	1.29
CHF	2.79	2.34	2.13	3.40	2.65	3.67	2.27	4.05	3.14	0.99
EUR	5.29	3.57	3.72	4.00	3.98	4.63	2.85	5.32	3.73	1.32
GBP	1.89	1.11	2.11	2.53	2.74	1.81	3.36	6.08	2.73	1.11
JPY	0.97	0.62	0.8	1.22	0.95	1.99	1.31	2.47	1.31	0.64
NOK	7.23	5.30	2.58	2.30	1.77	3.63	2.72	5.51	3.61	0.90
NZD	4.49	1.73	2.92	3.28	3.96	3.92	2.98	5.14	2.59	1.50
SEK	5.05	3.58	4.23	3.45	3.10	4.67	3.31	5.63	4.20	1.52
USD	4.91	4.91	4.42	5.19	4.64	5.09	2.73	4.59	5.47	1.41

Notes. The table reports $RMSE = \sqrt{\frac{1}{T}\sum_t \left(y_t - \nabla y_t^{MP}\right)^2}$, where y_t is the actual level of the yield of the row country, and ∇y_t^{MP} is the cumulative changes in y_t during three-day windows bracketing the announcement of the column central bank. The sample is post-2000. The unit is percentage points per annum.

Overall, the main message of this section is that a significant portion of the trend decline in global long-term interest rates occurred during the FOMC windows. By contrast, monetary policy announcements of other central banks did not exert a similarly persistent effect on domestic or other countries' long-term interest rates.

4 International Spillovers of FOMC Announcements

In the previous section, we demonstrated that the U.S. monetary policy announcement window has a uniquely strong explanatory power for yield variations across countries. In this section, we delve deeper into the transmission of these spillovers during the FOMC windows in three steps. First, using an adjusted uncovered interest parity identity, we show that even after accounting for exchange rate movements, foreign long-term interest rates move almost one-for-one with U.S. long-term rates. Second, we show that during the FOMC window, the vast majority of yield changes arise from movements in expected short-term rates, with only limited changes in term premia. Finally, we show that the U.S. policy short rate Granger-causes other countries' policy short rates.

4.1 Spillover Decomposition via Interest Parity

We utilize an uncovered interest parity (UIP) decomposition during FOMC windows to analyze the transmission mechanism. For any non-U.S. bond return, denoted as i_t^*, we can write the UIP equation as:

$$i_t^* = i_t + (\mathbf{E}_t[s_{t+1}] - s_t) + \alpha_t \tag{7}$$

where s is the foreign currency per U.S. dollar (an increase is a U.S. dollar appreciation), i is the return on a U.S. bond, and i^* is the return on a foreign bond, α is the excess return of investing in a foreign currency, or equivalently, the UIP deviation. Therefore, when the foreign interest rate adjusts, it must be associated with a change in the U.S. interest rate, exchange rate adjustment, or a deviation of the UIP equation.

Since we are interested in the daily changes in 10-year bond yields, we take the first difference of the UIP equation and relate it to the holding-period return of 10-year government bonds over a daily horizon. Furthermore, due to the lack of exchange rate expectation data at a daily frequency, we follow the literature and focus on the ex-post exchange rate change:

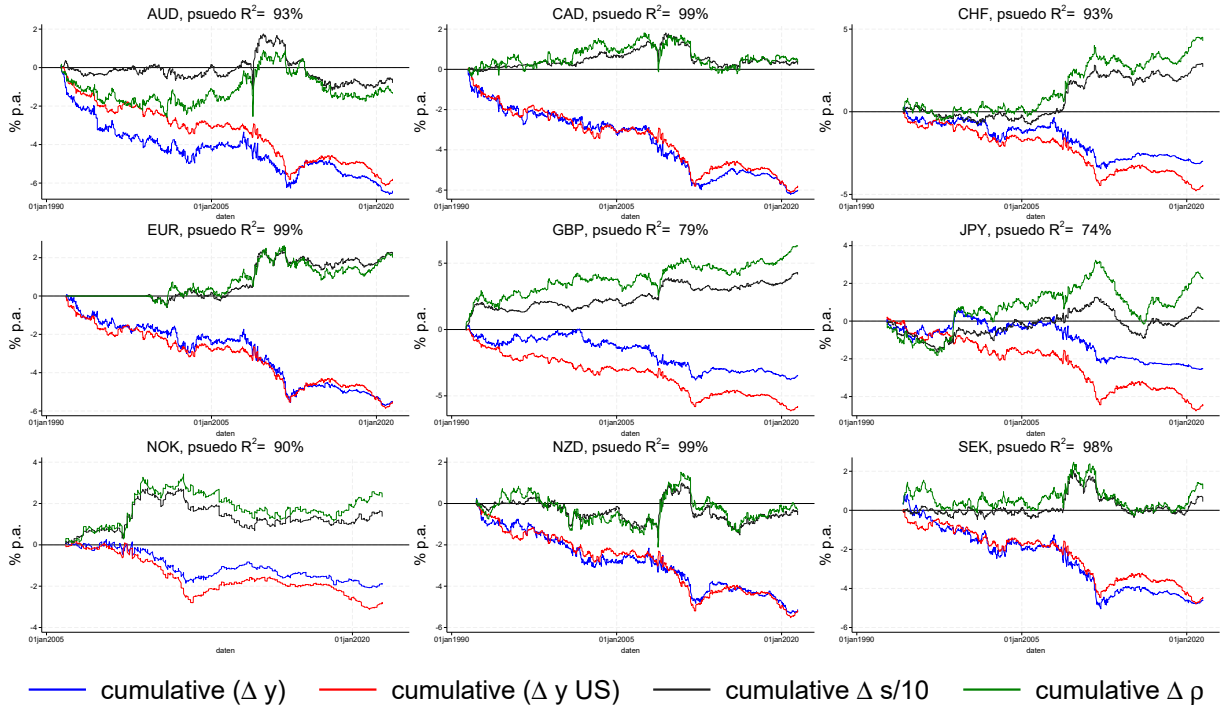
$$\textit{one day holding return}_t^{(10)*} = \textit{one day holding return}_t^{(10)} + \Delta_1(s_{t+1}) + \Delta_1\alpha_t, \tag{8}$$

where Δ_1 denotes the daily-change operator. We ignore the change in maturity over the daily holding period, so the daily holding period return is approximated by $\textit{one day holding return}_t^{(10)} \approx p_{t+1}^{(10)} - p_t^{(10)} = -10\Delta_1 y_t^{(10)}$. Plugging this term back to equation (8) and scaling all terms by 1/10, we get an identity in changes:

$$\Delta_1 y_{t+1}^{(10)*} = \Delta_1 y_{t+1}^{(10)} + \Delta_1 s_{t+1}/10 - \underbrace{\Delta_1\alpha_t/10}_{\equiv\Delta_1\rho_t}, \tag{9}$$

To the extent that the Mundell-Fleming framework is correct and all these G10 currencies operate under a floating exchange rate regime, we should expect a change in the 10-year rate in the U.S., $\Delta_1 y_t^{(10)}$, to be offset by a corresponding change in the exchange rate, leaving the change in the foreign rate, $\Delta_1 y_t^{(10)*}$, and the excess return, $\Delta_1\rho_t$, unchanged.

Figure 3: UIP decomposition within the FOMC windows (Equation (9))



Notes: The daily UIP condition is $\Delta_1 y_{t+1}^{(10)*} = \Delta_1 y_{t+1}^{(10)} + \Delta_1 s_{t+1}/10 - \Delta_1 \rho_t$. The figure plots the cumulative sums of daily changes in $\Delta_1 y_{t+1}^{(10)*}$ (blue), $\Delta_1 y_{t+1}^{(10)}$ (red), $\Delta_1 s_{t+1}/10$ (black), and $\Delta_1 \rho_t$ (green) during FOMC windows. The interest rates are the 10-year government bond yields. *pseudo* $R^2 \equiv \dfrac{\sum_{t=1}^{T}\left(\nabla y_t^{\text{US yield change}}\right)^2}{\sum_{W' \in \{\text{US yield change, exchange change, excess return}\}}\left[\sum_{t=1}^{T}\left(\nabla y_t^{W'}\right)^2\right]}$

In Figure 3, we cumulatively sum each term in equation (9) during FOMC windows. In all of these plots, one can see that the cumulative changes in U.S. bond yields and foreign bond yields track each other almost perfectly during the FOMC windows. Additionally, the cumulative $\Delta_1 s_{t+1}/10$ and cumulative $\Delta_1 \alpha_t/10$ also closely track each other, indicating that they largely cancel each other out in equation (9) at any time t, resulting in a direct passthrough of the U.S. 10 year yield to the foreign 10-year yield. The pseudo R^2 of the figures reports how well the cumulative 10-year U.S. yield change can explain the cumulative 10-year yield change in each foreign country.[3] The minimum pseudo R^2 is 74%, with seven out of nine countries exceeding

[3]The foreign yield change can be partition to three difference component (US yield change, exchange rate variation and excess return). We can define a pseudo R^2 to measure how well the cumulative foreign yield change, $\Delta_1 y_{t+1}^{(10)*}$, is

90%.

Papers by Gabaix and Maggiori (2015), Itskhoki and Mukhin (2021), as well as Devereux and Engel (2002), have emphasized the role of UIP deviations in explaining exchange rate movements. The findings on exchange rates here are broadly consistent with the exchange rate determination literature, but we provide evidence focusing on the FOMC window and its implications for interest rate passthrough. Contrary to the Mundell-Fleming prediction, we observe that the exchange rate does not adjust to monetary policy actions in the U.S. during the FOMC window. Instead, foreign interest rates strongly respond to monetary actions, resulting in an almost one-to-one response to U.S. interest rates.

In the Appendix D, we conduct a similar exercise using a covered interest parity (CIP) decomposition. The results are largely consistent with the UIP decomposition.

4.2 Spillover Decomposition with a Term Structure Model

We have shown that the cross-country passthrough is primarily transmitted to long-term yields rather than through exchange rate movements. In this subsection, we decompose long-term yields into the expected path of short-term interest rates and the term premium, and show that the main variation originates from changes in expected short-term rates.

To decompose long-term interest rates into their term premium and expected future short-term interest rates, we adopt the "observed shifting endpoint" (OSE) affine term structure model developed by Bauer and Rudebusch (2020). The model is particularly appropriate for our analysis because it allows for the possibility of a time-varying trend r_t^* rather than assuming the standard assumption of a constant r^*.[4]

We closely follow the model developed by Bauer and Rudebusch (2020) and estimate the term structure model on a country-by-country basis. Details of the model and estimation procedures are deferred to Appendix B. The model is estimated using the first five principal components of each country's yield curve as the state variable. We approximate the "observed shifting endpoint", denoted τ_t, as the first principal

explained by the cumulative US yield change, $\Delta_1 y_{t+1}^{(10)}$, during the FOMC windows. The exact definition is *pseudo* $R^2 \equiv \frac{\sum_{t=1}^{T}\left(\nabla y_t^{\text{US yield change}}\right)^2}{\sum_{W' \in \{\text{US yield change, exchange change, excess return}\}}\left[\sum_{t=1}^{T}\left(\nabla y_t^{W'}\right)^2\right]}$.

[4] Bauer and Rudebusch (2020) show that once an otherwise standard affine term structure model incorporates the cointegration trend as a state variable, the long-term behaviors of expected interest rates and term premia are opposite to those implied by canonical models. The reason is that canonical term structure models assume stationary state processes, so expected interest rates at long horizons must converge to the unconditional mean.

component of the cumulative daily changes in domestic and U.S. zero-coupon yields across maturities during the three-day FOMC windows.

We use the OSE model to decompose daily 10-year yields into risk-neutral rates and term premia, then compute the sum of their daily changes during the FOMC windows. Figure 4 presents the contributions of the risk-neutral and term premium components to the cumulative yield change during the FOMC windows, ∇y_t^{FOMC}. The series "FOMC window, yield" replicates the FOMC-window cumulative yield change series in Figure 2. The series "FOMC window, RNY" and "FOMC window, TP" denote the cumulative change in the risk-neutral yield (RNY) and term premium (TP) during FOMC windows. Clearly, the risk-neutral component explains most of the cumulative changes in the FOMC windows for all country's interest rates, except Japan. The risk-neutral component for Japan has remained nearly constant, as the Japanese short-term rate has been stuck at zero for most of our sample period. In contrast, the term premium component for all countries, except Japan, shows limited variation during the FOMC windows, thus explaining only a small portion of the variation in the cumulative responses of long-term interest rates to FOMC announcements.

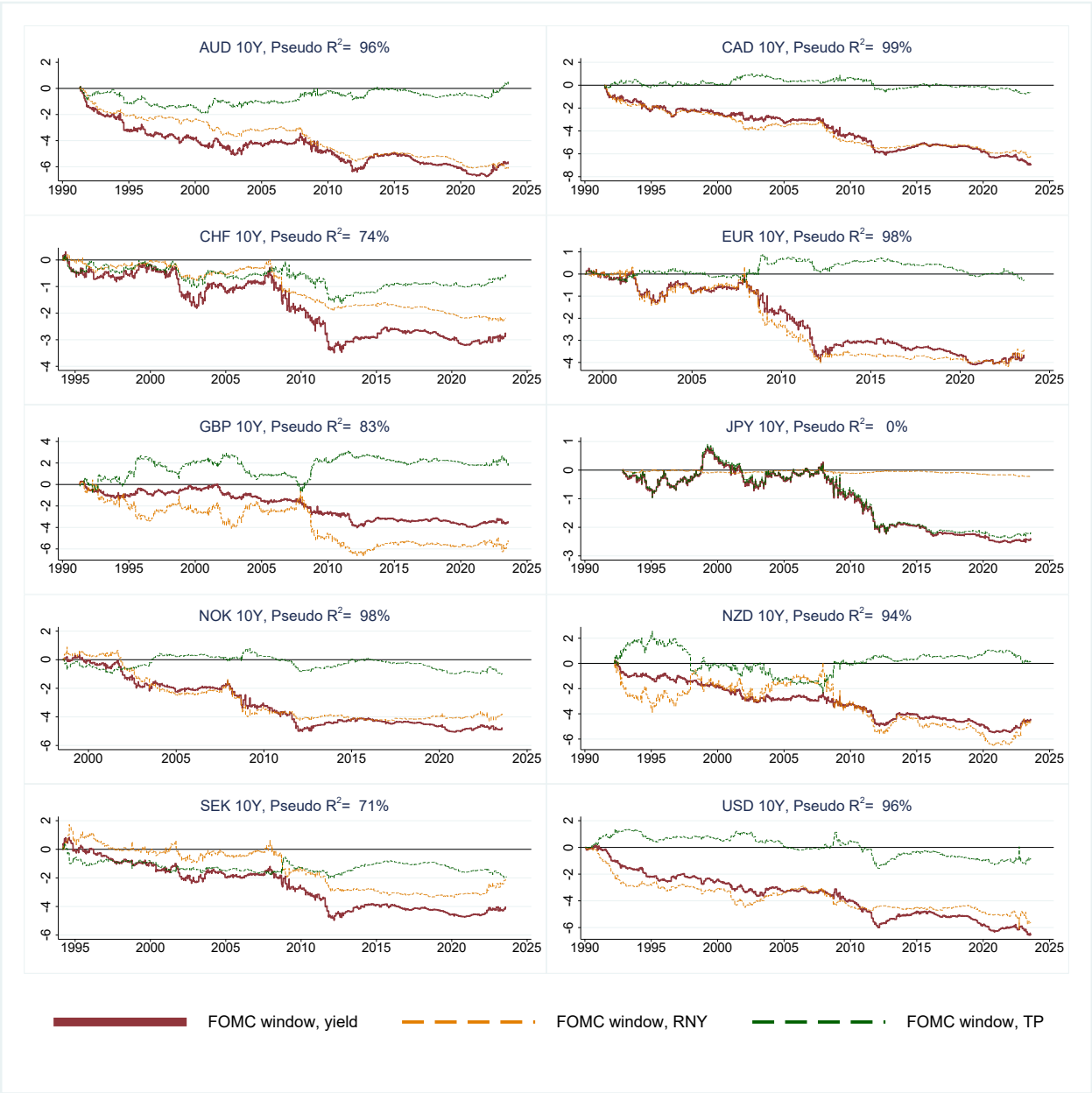
To quantitatively evaluate how much of the observed cumulative yield changes during FOMC windows can be explained by risk-neutral yields, we redefine the pseudo R^2 as

$$pseudoR^2 \equiv 1 - \frac{\sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC}\right)^2}{\sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{TP,FOMC}\right)^2 + \sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC}\right)^2}. \tag{10}$$

Since $\nabla y_t^{FOMC} = \left(\nabla y_t^{FOMC} - \nabla y_t^{TP,FOMC}\right) + \left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC}\right)$, the pseudo R^2 is analogous to R^2 measuring how well the risk-neutral yield fits the actual yield during the FOMC windows. Except for Japan, the risk-neutral yields account for over 70% of the cumulative changes in the 10-year yields, and the number exceeds 90% for six out of the ten currencies.

This exercise sheds light on the specific spillover mechanisms at play. During the FOMC windows, the secular decline in long-term interest rates in the U.S. is largely driven by the expected short-term rate component. This decline in U.S. rates also passes through to other countries, highlighting the interconnectedness of global financial markets. These results are consistent with the findings of Hanson and Stein (2015) in the U.S. context, as well as Albagli et al. (2019), who examine the contemporaneous pass through from U.S. monetary policy to foreign countries. We demonstrate that the short rate pass through is a key factor in

Figure 4: FOMC window: risk-neutral yield (RNY) vs. term premia (TP).



Notes. Unit of the y-axis is percentage p.a.. The figure decomposes the FOMC-window cumulative change in the 10-year yield into cumulative changes in the risk-neutral yield (RNY) and the term premium (TP) components. The pseudo R^2 measures the fraction of FOMC-window cumulative yield changes explained by the risk-neutral component. The figure shows the decomposition of the U.S. yield, and the table reports the results for other countries. $pseudoR^2 \equiv 1 - \frac{\sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC}\right)^2}{\sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{TP,FOMC}\right)^2 + \sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC}\right)^2}$

Table 4: FOMC windows and term structure decomposition.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
I: Share of FOMC window yield change explained by risk-neutral yield (RNY), equation (11).										
pseudo R^2 (%)	96	99	74	99	83	0	98	94	71	96
II: Share of overall risk-neutral yield changes explained by risk-neutral yield changes in FOMC windows.										
pseudo R^2 (%)	98	100	74	90	87	2	95	98	94	90
III: Share of overall term premium changes explained by term premium changes in FOMC windows.										
pseudo R^2 (%)	27	4	56	5	33	28	26	37	39	32

explaining the secular decline in interest rates, underscoring the dominant role of U.S. monetary policy in shaping global interest rate trends.

4.2.1 Risk-Neutral Yields and Term Premia Between FOMC Meetings

We have demonstrated that, conditional on FOMC windows, sovereign yield dynamics are predominantly driven by changes in risk-neutral yields and are largely unrelated to term premium dynamics. Is the muted term premium result driven by the estimation, such as a small overall estimated term premium?

To answer this, we construct the cumulative sum of daily changes in these two components during FOMC windows and on non-FOMC dates. We summarize the results in Table 4 and present the figures of the cumulative changes in Appendix B.5. In Panel II of Table 4, we report the pseudo R^2 for explaining the actual risk-neutral yield series using its changes over FOMC windows. Overall, FOMC windows account for over 70

FOMC windows have much smaller impacts on term premia than on risk-neutral yields. Panel III of Table 4 presents the pseudo R^2 for explaining the term premium using its dynamics during FOMC windows. On average, FOMC windows explain only about 30% of the total variation in term premia. For example, the pseudo R^2 for Germany suggests that FOMC windows account for just 5% of the total variation in the 10-year term premium. Since FOMC and non-FOMC windows together constitute the whole sample, this implies that most of the term premium variation occurs outside FOMC windows. The appendix figure shows that term premia in all countries have fluctuated substantially over the past three decades; however, these fluctuations are not concentrated in FOMC windows.

4.3 Spillover of U.S. policy rates

We have just shown that the spillover is mostly explained by changes in expected short-term rates. By definition, FOMC windows are the windows in which short-term policy rate decisions are made in the U.S.. In this subsection, we provide evidence that U.S. monetary policy exerts significant spillover effects on the policy rates of other countries, whereas the reverse does not hold.

To formally test the direction of spillovers from the U.S. to foreign economies, we estimate a series of monthly bivariate vector autoregressions (VARs),[5] each involving the U.S. policy rate (the effective federal funds rate) and the policy rate of a foreign country. For each country, we estimate the following system:

$$\begin{bmatrix} i_t^{(p)} \\ i_t^{(p^*)} \end{bmatrix} = \left(\sum_{l=1}^{12} \Phi_l L^l \right) \begin{bmatrix} i_t^{(p)} \\ i_t^{(p^*)} \end{bmatrix} + \begin{bmatrix} \varepsilon_t^{(p)} \\ \varepsilon_t^{(p^*)} \end{bmatrix}. \tag{11}$$

where L denotes the lag operator, and the system includes twelve monthly lags of both the U.S. and foreign policy rates. We then perform Granger causality tests to determine whether the U.S. policy rate provides statistically significant predictive content for foreign policy rates, and vice versa. Japan is excluded from this analysis due to a prolonged period of policy rate at the zero lower bound, which undermines the power of the test.

Table 5 presents the p-values from Wald tests for Granger causality between the U.S. and foreign policy rates. In each case, the null hypothesis is that one country's policy rate does not Granger-cause the other. The results display a striking asymmetry: the U.S. policy rate Granger-causes foreign policy rates at the 1% significance level for all foreign policy rate, whereas there is no statistically significant evidence in the reverse direction. The smallest p-value for the hypothesis that a foreign policy rate Granger-causes the U.S. rate is 0.18, far from conventional significance thresholds. These findings suggest that monetary policy spillovers are primarily driven by changes in the U.S. policy rate, which are decided during the FOMC announcement windows.

[5] Since policy rates are set by central banks, there is no meaningful daily variation when the other central bank has not yet conducted its own meeting.

Table 5: p-values from Wald tests for Granger causality between the U.S. and foreign policy rates.

	AUD	CAD	CHF	EUR	GBP	NOK	NZD	SEK
US rate → foreign rate	0.00***	0.00***	0.00***	0.00***	0.00***	0.00***	0.00***	0.00***
Foreign rate → US rate	0.41	0.19	0.90	0.41	0.75	0.18	0.33	0.81

Notes. The table presents p-values from Wald tests for Granger causality. The null hypothesis is that one country's policy rate does not Granger-cause the other. The column head currency is the currency of the foreign policy rate.

5 Channels Behind the Secular Decline in US Yields

Why is the decline in global interest rates so concentrated during FOMC windows? We unpack this by investigating why U.S. yields decline during FOMC windows.

In this section, we present evidence supporting the "Fed response to news" channel, which captures the Federal Reserve's reaction to macroeconomic information arriving shortly before monetary policy decisions. This channel, recently emphasized by Bauer and Swanson (2023a,b), constitutes a key component of the monetary policy news observed on FOMC announcement days. Our results indicate that this mechanism also accounts for a substantial portion of the cumulative decline in yields over the past three decades.

5.1 FOMC Announcements Preceded by Macro News

We partition the FOMC windows into two disjoint subsets: those preceded by macroeconomic news announcements and those that are not. Given the Federal Reserve's dual mandate, we focus on Consumer Price Index (CPI) and Gross Domestic Product (GDP) announcements as the prime indicators of price and real activity developments, the two key variables in standard characterizations of monetary policy reaction functions such as the Taylor rule. An FOMC window is classified as being preceded by macroeconomic news if a CPI or GDP data release occurs within the four business days leading up to the FOMC meeting. The four-day cutoff is selected so the sample split between meetings preceded by macroeconomic news and those that are not is roughly equal (52% and 48%).

Similar to Equation (3), we define a pseudo R^2 to evaluate how much of the yield changes during FOMC windows can be explained by those that are preceded by macro news. The pseudo R^2 of macro news FOMC

windows is:[6]

$$pseudo\ R^2 \equiv 1 - \frac{\sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{macronewsFOMC}\right)^2}{\sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{macronewsFOMC}\right)^2 + \sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{non-macronewsFOMC}\right)^2}. \tag{12}$$

Table 6 reports both the pseudo R^2 and the cumulative change in yields within the macro-news FOMC windows. Impressively, the results indicate that the downward yield trend during FOMC windows is heavily concentrated in those preceded by macroeconomic news. The U.S. yield decline in the macro-news FOMC windows tracks the overall yield decline almost perfectly, resulting in a pseudo R^2 of 99%. For all other countries, the macro-news FOMC windows explain, on average, about 82% of the cumulative change variation. The macro-news FOMC windows also explain the data well in levels. The average decline across countries is -4.2%, with -3.1% occurring during the macro-news FOMC windows.

Note that macro-news and non-macro-news FOMC windows equally split the sample. As a benchmark, the pseudo R^2 and cumulative decline during the macro-news FOMC windows would be roughly 50% if the yield decline were evenly distributed. The fact that the pseudo R^2 values exceed this neutral benchmark suggests that the Fed's responses to macroeconomic news play a crucial role in shaping yield movements around FOMC meetings.

The impressive fit of the macro news FOMC windows can be visually seen in Figure 5, in which we also compare it with other channels in the next subsection.

Table 6: Contributions of macro-FOMC windows to overall FOMC windows.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD	Average
Pseudo R^2, %											
Macro news FOMC windows	93	66	97	97	77	87	67	69	62	99	82
Yield Change, %											
Macro news FOMC windows	-4.5	-3.6	-2.8	-3.1	-2.1	-2.0	-2.7	-2.6	-2.3	-5.4	-3.1
FOMC windows	-5.8	-6.1	-2.8	-3.8	-2.4	-2.5	-4.2	-4.8	-4.2	-5.6	-4.2

Notes. The first panel reports the percentage of yield variations within all FOMC windows explained by those preceded by U.S. macro announcements within four business days.

[6]Since $\nabla y_t^{FOMC} = \nabla y_t^{macronewsFOMC} + \nabla y_t^{non-macronewsFOMC}$, this definition is equivalent to Equation (3) when we partition days into two disjoint subsets.

5.2 Comparison with other channels

We next assess the fit of the macro news FOMC windows by comparing them with three alternative windows. This approach helps identify the specific mechanism driving the observed patterns around the macro news FOMC windows.

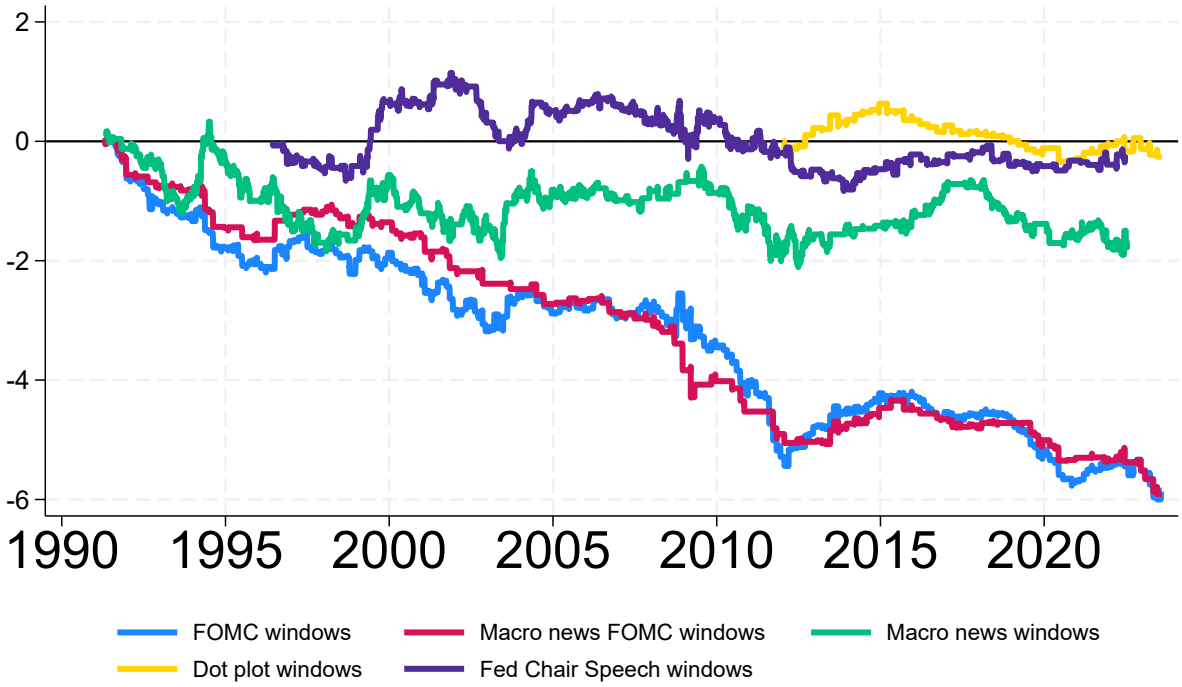
First, we compare with the macro data announcement window itself, defined as the three-day windows surrounding U.S. GDP or CPI releases. According to standard economic theories, long-term interest rate trends are determined by the economy's underlying fundamentals, and the market uses the latest macroeconomic data to update its assessment of these trends.

The second alternative is the FOMC windows associated with the release of the Summary of Economic Projections ("dot plot FOMC window"). The Summary of Economic Projections (SEP), published quarterly since 2012, outlines how Federal Reserve officials view the outlook for the U.S. economy and the likely future path of monetary policy. As part of the Fed's efforts to enhance policy transparency, the report presents each participant's forecasts of key variables such as GDP growth, unemployment, and inflation. Importantly, the SEP also includes the "dot plot," which shows each FOMC member's projection for the appropriate federal funds rate over the current year, the next two years, and the longer run.

Hillenbrand (2025) finds that declines in interest rates during FOMC windows are closely linked to changes in the dot plots over time—for instance, a drop in the longer-run median projected policy rate tends to coincide with a decline in the 10-year Treasury yield during these windows. He interprets this pattern as evidence of a "Fed information effect", whereby the market reacts not only to realized policy actions but also to new information conveyed by the SEP regarding the longer-run equilibrium interest rate or other measures of r^*. In particular, revisions in the dot plot provide signals that the market's perception of r^* may have changed.

Finally, we also consider a three-day window around Fed Chair speeches. As documented by Swanson and Jayawickrema (2023), asset price volatility—particularly for long-maturity yields—is often higher on Fed Chair speech dates than during FOMC meetings, as markets tend to react strongly to the Chair's statements. They argue that Fed Chair speeches may be more important than FOMC announcements for monetary policy transmission. For our purposes, Fed Chair speeches may convey more information about longer-term policy. We also view this as an alternative measure of the information effect. Notably, while the SEP is only available from 2012, records of Fed Chair speeches span a much longer period.

Figure 5: Cumulative 10Y US Yield Changes During Alternative Event Windows.



Notes. All announcement windows are days -1 to 1 relative to the announcement day. "FOMC" refers to the FOMC monetary policy announcements. "Macro FOMC" consists of the FOMC announcements preceded by U.S. CPI or GDP announcements by at most four business days. "Macro" consists of U.S. CPI or GDP announcement windows. "Dot plot" consists of the FOMC announcements that release dot plots. "Fed Chair speech" consists of the days that the Fed Chair delivers speeches.

We sum up daily yield changes over the Fed Chair speech window, defined as the three-day period $T-1$, T, and $T+1$, where T is the day of the speech. Fed Chair speech dates are sourced from the Federal Reserve Board of Governors' website, with records beginning on June 13, 1996.[7] There are 412 Fed Chair speech dates in our sample, whereas the number of FOMC announcements during the same period is 226.

Figure 5 plots the FOMC windows against the four alternatives for the case of 10-year US Treasury yield: 1) macro news FOMC windows, 2) macro news windows, 3) dot plot FOMC windows, and 4) Fed chair speech window. The statistics of pseudo R^2s and overall yield changes for all countries are summarized in Table 7.

[7]Depending on the country sample, some of the the 10-year bond yield series start from 1991 to 1994 (see Table 1), for this exercise, we cumulate the Fed Chair window only from 1996 due to data availability. The website is https://www.federalreserve.gov/newsevents/speeches.htm.

The figure clearly shows that macro news FOMC window dominates the others in explaining the FOMC window decline (and thus the overall decline). The macro news FOMC window closely track the FOMC windows over the whole sample, and as reported in Table 6 with a pseudo R^2 of 99%.

Interestingly, U.S. macroeconomic data announcement windows, considered in isolation, account for only a small share of the observed yield dynamics. These windows—including those associated with data revisions—comprise roughly 15% of all trading days, twice the frequency of FOMC announcement windows. Nevertheless, they explain less of the variation in yields than their FOMC counterparts. The cumulative yield decline during these windows is approximately 2%, with the overall decline over time generally remaining in the 0–2% range. As shown in Table 7, the average pseudo-R^2 measuring the share of total yield changes explained by macro news windows is 14%, closely mirroring the share of trading days on which such announcements occur. In several countries—including Switzerland, the Eurozone, Japan, New Zealand, and Sweden—the 10-year yield has increased, rather than declined, during macro news windows.

Similarly, dot plot FOMC windows do not account for the overall secular decline in yields. This is primarily for two reasons. First, the Summary of Economic Projections is available only from 2012, leaving a substantial portion of the sample unobserved. Second, yield changes since 2012 have been relatively small. From 2012 to the end of the sample, the decline in the 10-year Treasury yield during FOMC windows is less than -0.5%. Therefore, while dot plot FOMC windows capture post-2012 movements well, they do not drive the long-term decline. In Table 7, the pseudo R^2 for dot plot FOMC windows is close to 0%, as these windows do not capture the overall downward trend and omit a large portion of the sample.

The purple line in Figure 5 reports cumulative changes during Fed Chair speech windows. Overall, these windows do not explain the secular decline in interest rates. The cumulative sum is near zero at the end of the sample, with roughly half of the days slightly above zero and half slightly below. In Table 7, the average pseudo R^2 of how much of the overall yield change can be explained by Fed Chair speech windows across countries is only about 5%.

Notably, the fact that Fed Chair speech windows do not explain the secular trend of Treasury yields does not contradict Swanson and Jayawickrema (2023). In the Appendix, we show that the standard deviation of daily 10-year yields across countries during FOMC and Fed Chair speech windows is of similar magnitude and higher than on regular business days. Therefore, the findings of Swanson and Jayawickrema (2023) for the U.S. financial market can be generalized to the G10 countries. However, market reactions to Fed

Chair speeches exhibit a much less pronounced one-sided downward drift compared to those following FOMC announcements, and thus cannot account for the secular decline in government bond yields. As reported in Table 7, the cumulative decline during Fed Chair speech windows is only -0.3% for the U.S. and -0.2% for the average of G10 countries.

In summary, the FOMC meetings immediately preceded by macroeconomic data releases are primarily responsible for the global downward trend in interest rates. Macro data releases, dotplots, or Fed Chair Speeches alone do not lead to downward interest rate trends.

Table 7: Explanatory power of various alternative windows.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD	Average
Pseudo R^2, %											
FOMC window	80	72	67	84	58	19	86	85	74	94	73
Macro news window	17	16	7	9	17	7	8	17	4	35	14
Dot plot FOMC window	1	0	1	1	1	1	2	1	2	0	1
Fed Chair Speech window	0.8	4	3	5	4	15	6	7	1	2	5
Yield Change, %											
All Days	-8.1	-7.3	-4.0	-3.3	-8.4	-5.5	-2.8	-3.9	-6.1	-5.7	-5.5
FOMC window	-5.8	-6.1	-2.8	-3.8	-2.4	-2.5	-4.2	-4.8	-4.2	-5.6	-4.3
Macro news window	-1.5	-1.3	0.9	0.7	-2.7	2.4	0.1	1.9	1.0	-2.9	-0.2
Dot plot FOMC window	0.02	-0.1	0.3	0.4	0.1	-0.4	0.6	0.3	1.2	-0.2	-0.2
Fed Chair Speech window	-0.6	-1.7	-1.0	-0.8	-0.8	-1.4	-0.9	-1.7	-1.1	-0.8	-1.1

Notes. The first panel presents the percent variations of the actual yield explained by each event window. The second panel presents the total change in the yield within each event window.

5.3 Measuring Fed Response to News Using High Frequency Surprises

We next demonstrate the connection between macro news FOMC windows and high-frequency measures of monetary policy shocks. In particular, the macro news FOMC window is highly related to the "Fed response to news" component of Bauer and Swanson (2023a,b).

High-frequency identified monetary policy shocks (MPS) have become a standard tool in the empirical macro literature since Kuttner (2001) and Gürkaynak et al. (2005). We summarize the high-frequency changes in interest rate futures using the MPS measure from Bauer and Swanson (2023a,b), which is defined as the first principal component of 30-minute changes in the first four Eurodollar futures around FOMC announcements. As Bauer and Swanson (2023a,b) pointed out, part of the MPS is predictable by a set of macro variables and macro announcements before the FOMC meeting. Therefore, the MPS measure can be

"purified" or orthogonalized with the following regression:

$$MPS_t = \gamma_0 + \gamma_1 X_{t-} + u_t, \tag{13}$$

where X_{t-} is a set of macro variables known before the announcement date t. Bauer and Swanson (2023a,b) estimate this equation and categorize the error term u_t as the orthogonalized monetary policy shock, MPS_t^{orth}. For our purpose, we measure the "Fed responses to news" component as the predictable component of MPS.[8] In other words, we define

$$MPS_t^{Response} \equiv MPS_t - MPS_t^{orth}. \tag{14}$$

In Table 8, we first verify the relevance of the macro news FOMC window using a contemporaneous regression. We regress the 3-day change of the 10-year government bond yield on both $MPS_t^{Response}$ and MPS_t^{orth}. We estimate this regression separately on the macro news FOMC windows (if there is a CPI or GDP announcement within 4 days before the FOMC) and the non-macro news FOMC windows. The table summarizes the results for both the U.S. yield change and the rest of the G10 countries' yield change with a panel fixed effect regression:

$MPS_t^{Response}$ is a significant predictor of the 3-day yield change during the macro news FOMC windows but not the non-macro news FOMC windows. On the other hand, the MPS_t^{ORTH} is significant during the non-macro news FOMC windows but not during the macro news FOMC windows. The regression results suggest that the global yield decline during the macro FOMC windows is significantly driven by the Fed's response to the macroeconomic news.

We estimate the following equation to establish the relationship between the FOMC yield decline and the cumulative MPS:

$$\nabla y_t^{macronewsFOMC} = \gamma_0 + \gamma_1 CMPS_t^{ORTH} + \gamma_2 CMPS_t^{Response} + u_t, \tag{15}$$

where $\nabla y_t^{macronewsFOMC}$ is the sum of daily changes in a country's 10-year yield over macro news FOMC windows, and $CMPS_t^j = \sum_{s=0}^{t} MPS_s^j$ is the cumulative sum of the response component and orthogonalized

[8] With more VAR identification restrictions, Jarociński and Karadi (2020) document that more than 80% of the $MPS_t^{Response}$ variations are explained by the Fed response to macro news.

Table 8: Contemporaneous effects of monetary policy shocks on 10-year yield during FOMC windows

	G9		U.S.	
	Macro	No Macro	Macro	No Macro
MPS^{orth}	-0.2	0.3***	0.2	0.8***
	(0.2)	(0.1)	(0.2)	(0.2)
$MPS^{Response}$	1.1**	-0.2	1.5***	-0.1
	(0.5)	(0.5)	(0.4)	(0.5)
Const.	-0.0	-0.0	-0.0*	-0.0
	(0.0)	(0.0)	(0.0)	(0.0)
N	1134	1034	134	115
* $p < 0.1$	** $p < 0.05$	*** $p < 0.01$		

Notes. The table reports results for $\Delta_3 y_{i,t} = \alpha + \beta_1 MPS_t^{orth} + \beta_2 MPS_t^{Response} + u_{i,t}$. $\Delta_3 y_{i,t}$ denotes three-day changes in country i's 10-year yield over the FOMC window. For the "G9" columns, the equation is estimated for the panel of non-U.S. G9 countries with country fixed effects. For the "U.S." columns, the equation uses the U.S. yield only. "Macro" and "Non Macro" denote whether the regression is estimated using FOMC windows preceded by U.S. GDP or CPI announcements.

component of historical MPS up to t during the macro news FOMC windows.[9]

Table 9: Relative contributions of $CMPS^{ORTH}$ and $CMPS^{Response}$ in equation (15).

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD	Avg.
R^2 (%)	77	90	85	89	72	85	95	91	89	92	87
Semipartial correlation^2 (Response)	76	86	69	84	61	53	85	90	80	90	77
Semipartial correlation^2 (Orthogonal)	7	0	0	0	10	7	1	4	1	2	3

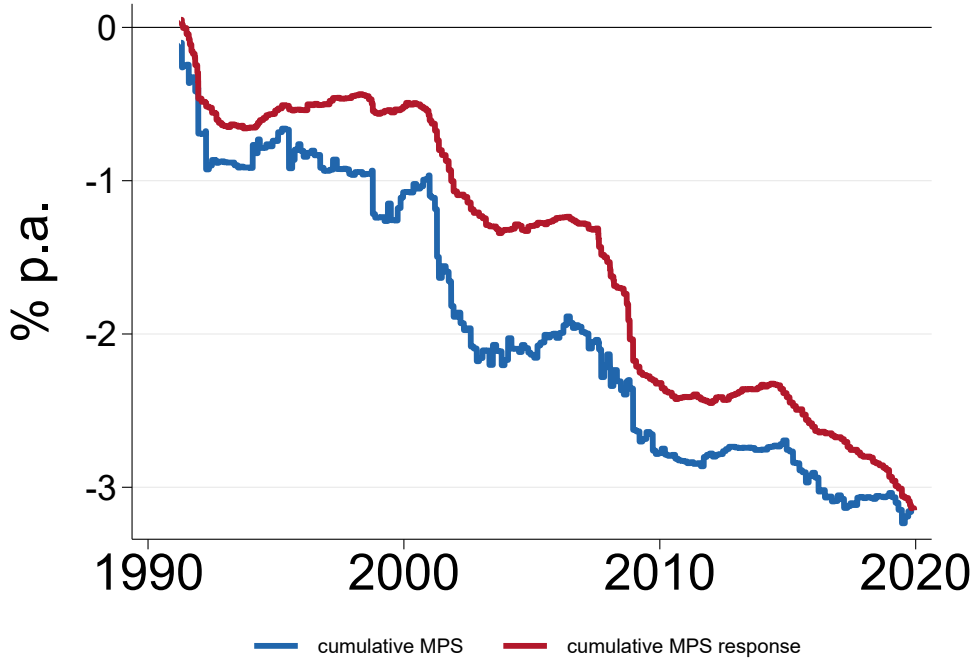
We fit the equation country by country. In Table 9, we report the overall R^2 of the regression and the semipartial squared correlation of each of the regressors.[10] It indicates that, in all cases, the cumulative yield change within macro news FOMC windows are well explained by the monetary policy shocks. The average R^2 across countries is 87%. For the semipartial correlation, it shows that adding $CMPS_t^{Response}$ to the regressions provides a much better fit to the cumulative 10-year yield decline. On the other hand, adding $CMPS_t^{ORTH}$ does a very little job of changing the fit.

[9] We focus on explaining the cumulative $\nabla y_t^{macronewsFOMC}$ for two reasons. First, while the macro news FOMC windows account for 50% of the FOMC windows, they account for about 80% of the yield decline across countries, and especially 99% of the yield decline in the U.S. Therefore, they capture significant variations. Second, by the OLS construction, MPS_t^{ORTH} sums up to zero, so there cannot be a downward trend over the full sample. Looking at the macro news FOMC windows avoids the mechanical sum-to-zero problem. In any case, our key focus is $CMPS_t^{Response}$.

[10] squared semipartial correlation measures the additional contribution of the regressor X_1 to explaining the Y variable, above and beyond another variable X_2

The analysis in this subsection indicates that the FOMC window declines are well connected with the "Fed response to news" component of MPS. Figure 6 presents the cumulative MPS change, along with the cumulative $MPS^{Response}$. The figure indicates that realized monetary policy shocks over the past three decades have been predominantly negative. This trend helps explain the secular decline in 10-year yields through the Fed's response to the news channel.

Figure 6: Cumulative US monetary policy shocks.



Notes. MPS refers to monetary policy shock measured by Bauer and Swanson (2023a). MPS response is the component of MPS defined in equation (14).

5.4 Interpretations of the "Fed response to news" effect on long-term yields

How does the significant association of the "Fed response to news" component with the macro news FOMC window inform us about the mechanism behind the systematic yield decline observed during FOMC windows?

To assess this point, we consider a standard Taylor rule:

$$i_t = r_t^* + \gamma_t^{\pi}(\pi_t - \pi_t^*) + \gamma_t^{Y}(outputgap_t^*) + u_t, \tag{16}$$

where i_t is the nominal policy rate, r_t^* is the natural rate of interest, π and π^* are the inflation rate and the inflation target and u_t is a monetary policy shock with zero mean. Similarly, the long-term interest rate is the sum of the expected future short-term rates and term premium:

$$y_t^{(n)} = \underbrace{\mathbf{E}_t\left[\frac{1}{n}\sum_{k=0}^{n-1} r_{t+k}^*\right] + \mathbf{E}_t\left[\frac{1}{n}\sum_{k=0}^{n-1}\gamma_{t+k}^{\pi}(\pi_t - \pi_t^*)\right] + \mathbf{E}_t\left[\frac{1}{n}\sum_{k=0}^{n-1}\gamma_{t+k}^{Y}(outputgap_{t+k}^*)\right]}_{\text{Risk Neutral Yield}_t \equiv \frac{1}{n}\mathbf{E}_t[i_t + i_{t+1} + \cdots + i_{t+n-1}]} + TP_t. \tag{17}$$

Thus, the long-term yield reflects the expected future path of the natural rate (r^*), the cumulative expected policy responses to inflation and output gaps, and the term premium. In Section 5, we perform a detailed term structure decomposition and find that the contribution of the term premium during FOMC windows is relatively modest. Hence, the dominant driver of yield movements during FOMC windows is the expected future path of short-term rates—that is, the risk-neutral yield.

However, several distinct channels may influence this risk-neutral yield: (1) expectations about the future path of the natural rate r^*; (2) expectations about future deviations of inflation from the target and the associated policy response coefficients; (3) expectations about the output gap and the corresponding reaction coefficients. As highlighted by Bauer and Swanson (2023a) and Bocola et al. (2024), the policy response coefficients themselves may vary over time, further amplifying the influence of macroeconomic expectations on yields.

To identify which of these channels primarily drives yield variation during FOMC windows, we examine the pseudo R^2 of yield changes conditional on the type of macroeconomic information released prior to each FOMC meeting. A macroeconomic announcement is classified as preceding an FOMC window if it occurs within four days before the meeting. The analysis considers CPI, industrial production (IP), unemployment (UN), GDP, and personal consumption expenditure (PCE) releases.

The results, presented in table 10, reveal a pronounced role for inflation-related announcements. FOMC windows preceded by CPI releases account for 28% of all windows but explain 81% of the variation in U.S. yields and 60% for the remaining G9 economies. When CPI and PCE announcements are considered jointly,

the explanatory power rises further—to 93% for the United States and 74% for the other G9 countries. This pattern indicates that FOMC windows preceded by inflation-related news disproportionately account for the observed movements in long-term yields.

By contrast, the explanatory power of other macroeconomic announcements, such as IP, UN, or GDP releases, remains modest and largely proportional to their frequency of occurrence in the sample. These results suggest that the information driving yield adjustments during FOMC windows is closely tied to inflation dynamics rather than to broader real-side developments.

Taken together, the evidence provides new insight into the mechanisms underlying the secular decline in global interest rates. Traditional explanations emphasize persistent real factors—such as demographics, productivity trends, or global savings imbalances—implying a limited role for monetary policy or inflation expectations in shaping long-run yield dynamics. Our results challenge this view. The strong sensitivity of yields to inflation-related news and evolving policy response coefficients suggests that expectations about monetary policy and inflation have been central to the secular adjustment of long-term interest rates over the past decades.

Table 10: Pseudo R^2 of various macro news FOMC windows (%).

	CPI	CPI+PCE	IP	UN	GDP
10 year US Treasury	81	93	13	1	18
average of G9 10 year Treasury	60	74	34	2	7
Share of $\frac{\text{this macro news FOMC windows}}{\text{total FOMC window}}$	28	44	26	15	25

Pseudo R^2 represents how much the total decline in yield during FOMC windows is explained by the particular macro news FOMC windows. Pseudo R^2 definition is definied in Equation (12).

6 Robustness: Placebo Tests

To investigate the statistical significance of the FOMC-based strategy, we compare the actual returns with those implied by randomly selected hypothetical FOMC windows.

Specifically, we conduct the test as follows. For each year y, randomly draw N_y Wednesdays from that year to serve as the hypothetical FOMC announcement days. This is because most FOMC announcements occur on Wednesday afternoon, U.S. Eastern time. We construct three-day windows around these Wednesdays

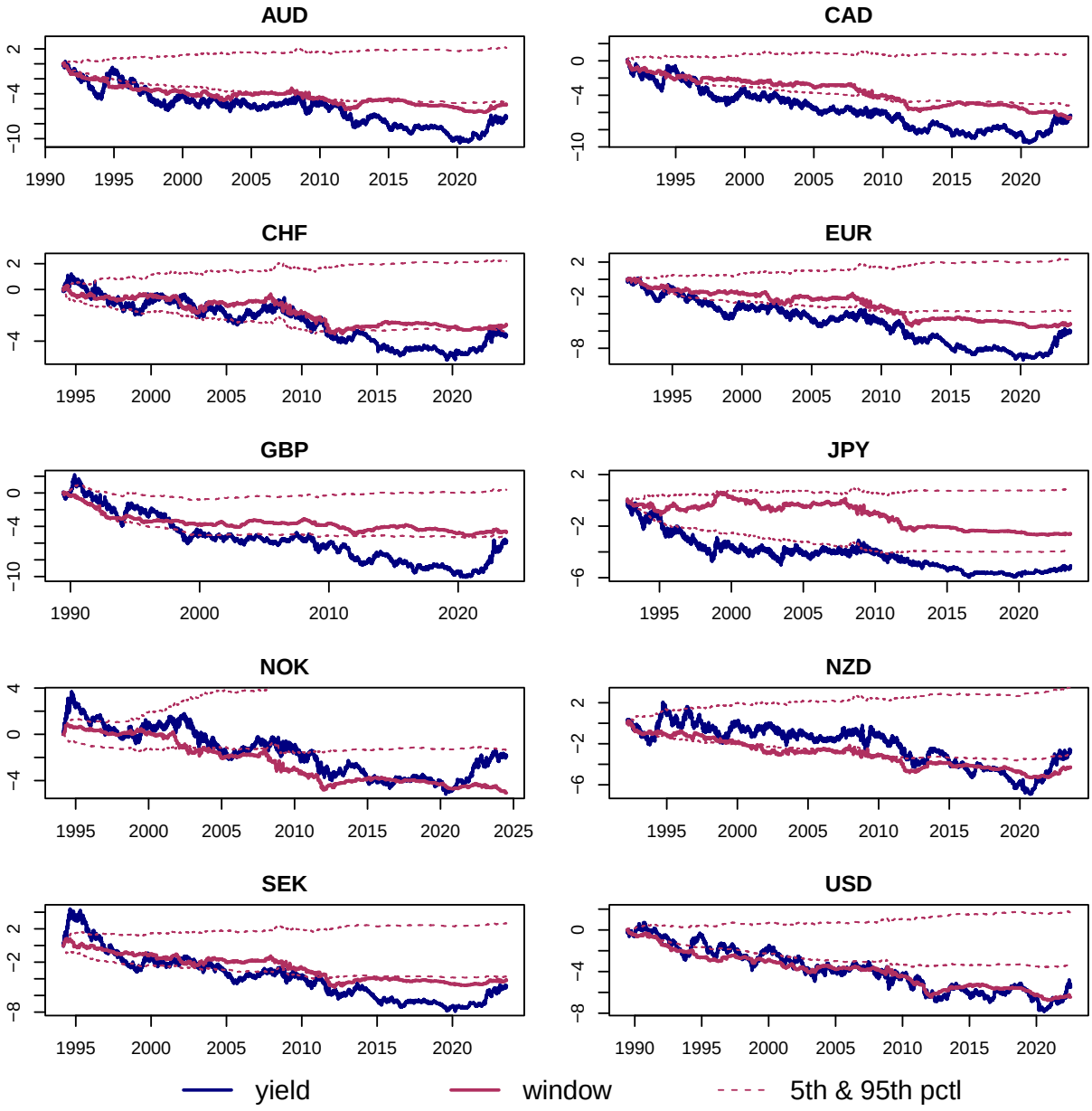
and compute cumulative daily yield changes within these windows. Then, we repeat the process.

Figure 7 presents the placebo test result based on 10,000 random samples. The 90% confidence bands denote the 5^{th} and 95^{th} percentiles of the cumulative yield changes during the placebo three-day windows. Except for Japan, the cumulative changes in the 10-year yields during the actual FOMC windows are at or lower than the lower bound of the confidence bands. In other words, the cumulative yield changes during FOMC windows are lower than 95% of the placebo sample paths, indicating that FOMC windows lead to significantly more negative yield changes than randomly selected event windows of the same length.

7 Conclusion

This study contributes to the literature by first documenting that and then analyzing how U.S. monetary policy announcements have driven the secular decline in long-term interest rates across advanced economies. Our empirical analysis demonstrates that FOMC announcement windows significantly influence global yields, with the effects being most pronounced when announcements are preceded by key macroeconomic news. Unlike monetary policy announcements of other central banks, U.S. monetary policy exhibits a unique capacity to shape international bond markets through signaling and expectation channels. By decomposing yield movements, we highlight the dominance of changes in risk-neutral rates over term premia in driving the global propagation of FOMC announcement effects. These findings challenge conventional views that attribute secular yield declines solely to structural economic shifts, emphasizing instead the pivotal role of learning about monetary policy reaction functions in a globally interconnected financial environment.

Figure 7: Cumulative yield changes over placebo windows.



Notes. Unit of the y-axis is percentage p.a.. Placebo tests for cumulative yield changes during the $t-1$-to-$t+1$ windows. We randomly pick N_y Wednesdays from each year y serving as the placebo announcement days. Then, construct the placebo announcement windows accordingly. The dashed lines denote the 5^{th} and 95^{th} percentiles from 10,000 placebo samples.

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Online Appendix (not for publication)

A FOMC Announcement Dates

Table A1 presents the scheduled FOMC dates in our sample. Table A2 presents the unscheduled FOMC announcement dates in our sample. Table A3 presents the numbers of G10 monetary policy announcement dates overlapping with FOMC windows.

B The Shifting Endpoint Model

B.1 Model setup

To decompose the long-term interest rates into term premia and expected future short-term interest rates, we adopt the affine term structure model developed by Bauer and Rudebusch (2020). The model is particularly appropriate for our analysis because it allows for the possibility of cointegration trends as state variables.[11]

We use upper case or bold letters to denote vectors. The state variable vector X_t is a $K_X \times 1$ vector X_t consisting of a trend vector $\boldsymbol{\tau}_t$ and a stationary vector $\tilde{X}_t$:

$$\begin{aligned} X_t =& \boldsymbol{\mu} + \Gamma \boldsymbol{\tau}_t + \tilde{X}_t, \\ \boldsymbol{\tau}_t =& \boldsymbol{\tau}_{t-1} + \boldsymbol{\eta}_t, \quad \boldsymbol{\eta}_t \sim \mathcal{N}(0, \Omega_\eta) \\ \tilde{X}_t =& \Phi \tilde{X}_{t-1} + \tilde{U}_t, \quad \tilde{U}_t \sim \mathcal{N}(0, \tilde{\Omega}), \end{aligned} \tag{A1}$$

where $\boldsymbol{\tau}_t$ is a $K_\tau \times 1$ random walk and $\tilde{X}_t$ is a $K_X \times 1$ stationary VAR(1). The shocks are i.i.d over time and $\boldsymbol{\eta}_t \perp \tilde{U}_t$. Define

$$Z_t \equiv \begin{bmatrix} \boldsymbol{\tau}_t^\top & X_t^\top \end{bmatrix}^\top, U_t \equiv \Gamma \boldsymbol{\eta}_t + \tilde{U}_t, \Omega \equiv \mathbf{E}[U_t U_t^\top] = \Gamma \Omega_\eta \Gamma^\top + \tilde{\Omega}.$$

[11] Bauer and Rudebusch (2020) show that once an otherwise standard affine term structure model incorporates the cointegration trend as a state variable, the long-term behaviors of expected interest rates and term premia are opposite to those implied by canonical models. The reason is that canonical term structure models assume stationary state processes, so expected interest rates at long horizons must converge to the unconditional mean.

Table A1: Scheduled FOMC meeting dates.

		Scheduled FOMC Meetings							
Year	N	1.	2.	3.	4.	5.	6.	7.	8.
1989	5				7-Jul	23-Aug	4-Oct	15-Nov	20-Dec
1990	8	8-Feb	28-Mar	16-May	5-Jul	22-Aug	3-Oct	14-Nov	18-Dec
1991	8	7-Feb	27-Mar	15-May	5-Jul	21-Aug	2-Oct	6-Nov	18-Dec
1992	8	6-Feb	1-Apr	20-May	2-Jul	19-Aug	7-Oct	18-Nov	23-Dec
1993	8	4-Feb	24-Mar	19-May	8-Jul	18-Aug	22-Sep	17-Nov	22-Dec
1994	8	4-Feb	22-Mar	17-May	6-Jul	16-Aug	27-Sep	15-Nov	20-Dec
1995	8	1-Feb	28-Mar	23-May	6-Jul	22-Aug	26-Sep	15-Nov	19-Dec
1996	8	31-Jan	26-Mar	21-May	3-Jul	20-Aug	24-Sep	13-Nov	17-Dec
1997	8	5-Feb	25-Mar	20-May	2-Jul	19-Aug	30-Sep	12-Nov	16-Dec
1998	8	4-Feb	31-Mar	19-May	1-Jul	18-Aug	29-Sep	17-Nov	22-Dec
1999	8	3-Feb	30-Mar	18-May	30-Jun	24-Aug	5-Oct	16-Nov	21-Dec
2000	8	2-Feb	21-Mar	16-May	28-Jun	22-Aug	3-Oct	15-Nov	19-Dec
2001	8	31-Jan	20-Mar	15-May	27-Jun	21-Aug	2-Oct	6-Nov	11-Dec
2002	8	30-Jan	19-Mar	7-May	26-Jun	13-Aug	24-Sep	6-Nov	10-Dec
2003	8	29-Jan	18-Mar	6-May	25-Jun	12-Aug	16-Sep	28-Oct	9-Dec
2004	8	28-Jan	16-Mar	4-May	30-Jun	10-Aug	21-Sep	10-Nov	14-Dec
2005	8	2-Feb	22-Mar	3-May	30-Jun	9-Aug	20-Sep	1-Nov	13-Dec
2006	8	31-Jan	28-Mar	10-May	29-Jun	8-Aug	20-Sep	25-Oct	12-Dec
2007	8	31-Jan	21-Mar	9-May	28-Jun	7-Aug	18-Sep	31-Oct	11-Dec
2008	8	30-Jan	18-Mar	30-Apr	25-Jun	5-Aug	16-Sep	29-Oct	16-Dec
2009	8	28-Jan	18-Mar	29-Apr	24-Jun	12-Aug	23-Sep	4-Nov	16-Dec
2010	8	27-Jan	16-Mar	28-Apr	23-Jun	10-Aug	21-Sep	3-Nov	14-Dec
2011	8	26-Jan	15-Mar	27-Apr	22-Jun	9-Aug	21-Sep	2-Nov	13-Dec
2012	8	25-Jan	13-Mar	25-Apr	20-Jun	1-Aug	13-Sep	24-Oct	12-Dec
2013	8	30-Jan	20-Mar	1-May	19-Jun	31-Jul	18-Sep	30-Oct	18-Dec
2014	8	29-Jan	19-Mar	30-Apr	18-Jun	30-Jul	17-Sep	29-Oct	17-Dec
2015	8	28-Jan	18-Mar	29-Apr	17-Jun	29-Jul	17-Sep	28-Oct	16-Dec
2016	8	27-Jan	16-Mar	27-Apr	15-Jun	27-Jul	21-Sep	2-Nov	14-Dec
2017	8	1-Feb	15-Mar	3-May	14-Jun	26-Jul	20-Sep	1-Nov	13-Dec
2018	8	31-Jan	21-Mar	2-May	13-Jun	1-Aug	26-Sep	8-Nov	19-Dec
2019	8	30-Jan	20-Mar	1-May	19-Jun	31-Jul	18-Sep	30-Oct	11-Dec
2020	7	29-Jan	29-Apr	10-Jun	29-Jul	16-Sep	5-Nov	16-Dec	
2021	8	27-Jan	17-Mar	28-Apr	16-Jun	28-Jul	22-Sep	3-Nov	15-Dec
2022	8	26-Jan	16-Mar	4-May	15-Jun	27-Jul	21-Sep	2-Nov	14-Dec

The log stochastic discount factor (SDF) m_{t+1} evolves as

$$m_{t+1} = -\delta_0 - \boldsymbol{\delta}_1^\top X_t - \frac{1}{2}\Lambda_t^\top \Lambda_t - \Lambda_t^\top \Omega^{-\frac{1}{2}} U_{t+1}. \tag{A2}$$

Table A2: Unscheduled FOMC Meeting Dates

		Unscheduled FOMC Meetings								
Year	N	1.	2.	3.	4.	5.	6.	7.	8.	9
1989	4				5-Jun	26-Jul	16-Oct	6-Nov		
1990	3	13-Jul	29-Oct	7-Dec						
1991	9	8-Jan	1-Feb	8-Mar	30-Apr	6-Aug	13-Sep	31-Oct	6-Dec	20-Dec
1992	2	9-Apr	4-Sep							
1993	0									
1994	1	18-Apr								
1995	0									
1996	0									
1997	0									
1998	1	15-Oct								
1999	0									
2000	0									
2001	3	3-Jan	18-Apr	17-Sep						
2002	0									
2003	0									
2004	0									
2005	0									
2006	0									
2007	0									
2008	2	22-Jan	8-Oct							
2009	0									
2010	0									
2011	0									
2012	0									
2013	0									
2014	0									
2015	0									
2016	0									
2017	0									
2018	0									
2019	0									
2020	2	3-Mar	15-Mar							
2021	0									

The price of risk is an affine function of Z_t:[12]

$$\Lambda_t = \Omega^{-\frac{1}{2}}(\Lambda_0 + \Lambda_1 Z_t). \tag{A3}$$

[12]We assume that Λ_1 satisfies $\Lambda_1 = \left[(I-\Phi)\Gamma, \Lambda_{12}\right]$, i.e., the first K_τ column of Λ_1 (the loading on $\boldsymbol{\tau}_t$) equals $(I-\Phi)\Gamma$ and the remaining K_X columns is an unrestricted $K_X \times K_X$ matrix Λ_{12}.

Table A3: Overlaps with FOMC announcement windows.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
MP start	1990	2000	2000	1999	1997	1999	1986	1999	1999
$N_{overlap}$	15	11	32	29	49	66	83	38	24
$\frac{N_{overlap}}{N_{FOMC}}$	0.05	0.04	0.10	0.09	0.16	0.21	0.27	0.12	0.08
$\frac{N_{overlap}}{N_{FOMC}}$ post-2000	0.05	0.05	0.20	0.14	0.22	0.31	0.27	0.18	0.11

Note that the SDF is driven by a $K_X \times 1$ dimensional shock with the same dimension as X_t, but is a combination of shocks to $\boldsymbol{\tau}_t$ and $\tilde{X}_t$. Although the trend $\boldsymbol{\tau_t}$ does not directly affect the observed yields, it affects risk premia by affecting the price of risk.

The model implies a standard expression in the affine term structure model literature that the log zero-coupon bond prices are affine in the state vector X_t:

$$y_t^{(n)} = A_n + B_n^\top X_t, \tag{A4}$$

where A_n and B_n are coefficients satisfying standard no-arbitrage recursions. The log risk-neutral bond prices solve

$$p_t^{(n),rn} = \ln \mathbf{E}_t \left[\exp \left\{ -y_t^{(1)} + p_{t+1}^{(n-1),rn} \right\} \right], \tag{A5}$$

which is an affine function of X_t and $\boldsymbol{\tau}_t$:

$$p_t^{(n),rn} = \mathcal{A}_n^{rn} + \mathcal{B}_n^{rn\top} X_t + \mathcal{C}_n^{rn\top} \boldsymbol{\tau}_t, \tag{A6}$$

and the coefficients solve the recursions described in Appendix B. The bond price and yield are related by $p_t^{(n)} = -n y_t^{(n)}$, and analogously for the risk-neutral and term premium components. Clearly, risk-neutral rates explicitly depend on $\boldsymbol{\tau}_t$.

B.2 Details of the No-Arbitrage Recursions

First, we show that the state vector Z_t evolves as

$$Z_t = \boldsymbol{\mu}_Z + \Phi_Z Z_{t-1} + V_t, \quad V_t \equiv \begin{bmatrix} \boldsymbol{\eta}_t \\ U_t \end{bmatrix}, \tag{A7}$$

with

$$\boldsymbol{\mu}_Z = \begin{bmatrix} 0 \\ (I-\Phi)\boldsymbol{\mu} \end{bmatrix}, \quad \Phi_Z = \begin{bmatrix} I_{K_\tau} & 0_{K_\tau \times K_X} \\ (I-\Phi)\Gamma & \Phi \end{bmatrix}, \quad \Omega_V \equiv \mathbf{E}[V_t V_t^\top] = \begin{bmatrix} \Omega_\eta & \Omega_\eta \Gamma^\top \\ \Gamma\Omega_\eta & \Omega \end{bmatrix}.$$

We rewrite Z_t as

$$Z_t = \begin{bmatrix} 0 \\ \boldsymbol{\mu} \end{bmatrix} + \begin{bmatrix} I_{K_\tau} & 0_{K_\tau \times K_X} \\ \Gamma & \Phi \end{bmatrix} \begin{bmatrix} \boldsymbol{\tau}_{t-1} \\ \tilde{X}_{t-1} \end{bmatrix} + \begin{bmatrix} \boldsymbol{\eta}_t \\ \Gamma \boldsymbol{\eta}_t + \tilde{U}_t. \end{bmatrix}$$

Note that

$$\begin{bmatrix} \boldsymbol{\tau}_{t-1} \\ \tilde{X}_{t-1} \end{bmatrix} = \begin{bmatrix} 1 & 0_{K_\tau \times K_X} \\ -\Gamma & I \end{bmatrix} \begin{bmatrix} \boldsymbol{\tau}_{t-1} \\ X_{t-1} \end{bmatrix} - \begin{bmatrix} 0 \\ \boldsymbol{\mu} \end{bmatrix}.$$

Substituting for $\boldsymbol{\tau}_{t-1}$ and $\tilde{X}_{t-1}$, we get $\boldsymbol{\mu}_Z, \phi_Z$ and V_t. Since $V_t \perp \tilde{U}_t$, the expression for Ω_V follows naturally.

Next, we show that the restriction

$$\Lambda_1 = \begin{bmatrix} (I-\Phi)\Gamma & \Lambda_{12} \end{bmatrix} \tag{A8}$$

implies the bond pricing equation $p_t^{(n)} = \mathcal{A}_n + \mathcal{B}_n^\top X_t$. We prove by guess-and-verify. The no-arbitrage recursion is

$$p_t^{(n)} = \mathbf{E}_t[m_{t+1}] + \mathbf{E}_t[p_{t+1}^{(n-1)}] + \frac{1}{2}\mathbf{Var}_t(m_{t+1}) + \frac{1}{2}\mathbf{Var}_t(p_{t+1}^{(n-1)}) + \mathbf{Cov}_t(m_{t+1}, p_{t+1}^{(n-1)}). \tag{A9}$$

Note that $\mathbf{E}_t[\cdot]$ refers $\mathbf{E}[\cdot|Z_t]$, and

$$\begin{aligned} \mathbf{E}_t[X_{t+1}] =& \boldsymbol{\mu} + \Gamma\boldsymbol{\tau}_t + \Phi\tilde{X}_t = \boldsymbol{\mu} + \Gamma\boldsymbol{\tau}_t + \Phi(X_t - \boldsymbol{\mu} - \Gamma\boldsymbol{\tau}_t) \\ =& (I-\Phi)(\boldsymbol{\mu} + \Gamma\boldsymbol{\tau}_t) + \Phi X_t. \end{aligned}$$

When $p_t^{(n)} = \mathcal{A}_n + \mathcal{B}_n^\top X_t$,

$$\mathbf{E}_t[m_{t+1}] + \frac{1}{2}\mathbf{Var}_t(m_{t+1}) = -\delta_0 - \boldsymbol{\delta}_1^\top X_t,$$

$$\mathbf{E}_t[p_{t+1}^{(n-1)}] = \mathcal{A}_{n-1} + \mathcal{B}_{n-1}^\top[(I-\Phi)(\boldsymbol{\mu} + \Gamma\boldsymbol{\tau}_t) + \Phi X_t],$$

$$\mathbf{Var}_t(p_{t+1}^{(n-1)}) = \mathcal{B}_{n-1}^\top \Omega \mathcal{B}_{n-1},$$

$$\mathbf{Cov}_t(m_{t+1}, p_{t+1}^{(n-1)}) = -\mathcal{B}_{n-1}^\top(\Lambda_0 + \Lambda_1 Z_t).$$

Note that $\Lambda_1 Z_t = \Lambda_{11}\boldsymbol{\tau}_t + \Lambda_{12} X_t$, and we hope to eliminate $\boldsymbol{\tau}_t$ from the right-hand side of the recursion. Collecting the terms involving $\boldsymbol{\tau}_t$, we should have

$$\mathcal{B}_{n-1}^\top[(I-\Phi)\Gamma - \Lambda_{11}] = 0, \quad \forall n.$$

So $\Lambda_{11} = (I-\Phi)\Gamma$ eliminates $\boldsymbol{\tau}_t$ from the right-hand side of Equation (A9).

Finally, we derive the bond pricing recursions. Equation (A9) together with Equation (A8) implies

$$\begin{aligned} p_t^{(n)} =& -\delta_0 - \boldsymbol{\delta}_1^\top X_t + \mathcal{A}_{n-1} + \mathcal{B}_{n-1}^\top[(I-\Phi)\boldsymbol{\mu} + \Phi X_t] && \text{(A10)} \\ &+ \frac{1}{2}\mathcal{B}_{n-1}^\top \Omega \mathcal{B}_{n-1} - \mathcal{B}_{n-1}^\top(\Lambda_0 + \Lambda_{12} X_t). && \text{(A11)} \end{aligned}$$

So,

$$\begin{aligned} \mathcal{A}_n =& \mathcal{A}_{n-1} - \delta_0 + \mathcal{B}_{n-1}^\top(I-\Phi)\boldsymbol{\mu} + \frac{1}{2}\mathcal{B}_{n-1}^\top \Omega \mathcal{B}_{n-1} - \mathcal{B}_{n-1}^\top \Lambda_0, && \text{(A12)} \\ \mathcal{B}_n^\top =& -\boldsymbol{\delta}_1^\top + \mathcal{B}_{n-1}^\top(\Phi - \Lambda_{12}). && \text{(A13)} \end{aligned}$$

The yields are

$$y_t^{(n)} = A_n + B_n^\top X_t, \tag{A14}$$

with $A_n = -\frac{1}{n}\mathcal{A}_n$ and $B_n = -\frac{1}{n}\mathcal{B}_n$.

The log risk-neutral bond prices solve

$$p_t^{(n),rn} = \ln \mathbf{E}_t\left[\exp\left\{-y_t^{(1)} + p_{t+1}^{(n-1),rn}\right\}\right]. \tag{A15}$$

We guess and verify that $p_t^{(n),rn}$ is an affine function of X_t and $\boldsymbol{\tau}_t$:

$$p_t^{(n),rn} = \mathcal{A}_n^{rn} + \mathcal{B}_n^{rn\top} X_t + \mathcal{C}_n^{rn\top} \boldsymbol{\tau}_t, \tag{A16}$$

and the coefficients solve the following recursions:

$$\begin{aligned}
\mathcal{A}_n^{rn} =& \mathcal{A}_{n-1}^{rn} - \delta_0 + \mathcal{B}_{n-1}^{rn\top}(I-\Phi)\boldsymbol{\mu} + \frac{1}{2}\mathcal{B}_{n-1}^{rn\top}\Omega\mathcal{B}_{n-1}^{rn} \\
&+ \frac{1}{2}\left(\mathcal{B}_{n-1}^{rn\top}\Gamma\Omega_\eta\mathcal{C}_{n-1}^{rn} + \mathcal{C}_{n-1}^{rn\top}\Omega_\eta\Gamma^\top\mathcal{B}_{n-1}\right) + \frac{1}{2}\mathcal{C}_{n-1}^{rn\top}\Omega_\eta\mathcal{C}_{n-1}^{rn}, \\
\mathcal{B}_n^{rn\top} =& -\boldsymbol{\delta}_1^\top + \mathcal{B}_{n-1}^{rn\top}\Phi, \\
\mathcal{C}_n^{rn\top} =& \mathcal{B}_{n-1}^{rn\top}(I-\Phi)\Gamma + \mathcal{C}_{n-1}^{rn\top}
\end{aligned} \tag{A17}$$

with $\mathcal{A}_1^{rn} = -\delta_0, \mathcal{B}_1^{rn} = -\boldsymbol{\delta}_1, \mathcal{C}_1^{rn} = \mathbf{0}$.

B.3 Excess Bond Return

The excess holding period return is

$$rx_{t+1}^{(n)} = p_{t+1}^{(n-1)} - p_t^{(n} - y_t^{(1)} = \mathcal{A}_{n-1} + \mathcal{B}_{n-1}^\top X_{t+1} - \mathcal{A}_n + \mathcal{B}_n^\top X_t - \delta_0 - \boldsymbol{\delta}_1^\top X_t.$$

Equation (A7) implies $X_{t+1} = (I-\Phi)(\boldsymbol{\mu}+\Gamma\boldsymbol{\tau}_t)+\Phi X_t+U_{t+1}$. Substituting for X_{t+1} and using the recursions for $\mathcal{A}$ and $\mathcal{B}$, we get

$$rx_{t+1}^{(n)} = \mathcal{B}_{n-1}^\top(\Lambda_0 + \Lambda_1 Z_t) - \mathcal{B}_{n-1}^\top\Omega\mathcal{B}_{n-1} + \mathcal{B}_{n-1}^\top U_{t+1}. \tag{A18}$$

Furthermore, we can substitute $X_{t+1} - (I-\Phi)(\boldsymbol{\mu}+\Gamma\boldsymbol{\tau}_t) - \Phi X_t$ for U_{t+1}:

$$\begin{aligned}
rx_{t+1}^{(n)} =& \mathcal{B}_{n-1}^\top(\Lambda_0 + \Lambda_1 Z_t) - \frac{1}{2}\mathcal{B}_{n-1}^\top\Omega\mathcal{B}_{n-1} + \mathcal{B}_{n-1}^\top\left(X_{t+1} - (I-\Phi)(\boldsymbol{\mu}+\Gamma\boldsymbol{\tau}_t) - \Phi X_t\right) \\
=& \mathcal{B}_{n-1}^\top(\Lambda_0 - (I-\Phi)\boldsymbol{\mu}) - \frac{1}{2}\mathcal{B}_{n-1}^\top\Omega\mathcal{B}_{n-1} \\
&+ \mathcal{B}_{n-1}^\top(\Lambda_{11}\boldsymbol{\tau}_t + \Lambda_{12}X_t - (I-\Phi)\Gamma\boldsymbol{\tau}_t - \Phi X_t) + \mathcal{B}_{n-1}^\top X_{t+1} \\
=& \mathcal{B}_{n-1}^\top(\Lambda_0 - (I-\Phi)\boldsymbol{\mu}) - \frac{1}{2}\mathcal{B}_{n-1}^\top\Omega\mathcal{B}_{n-1} + \mathcal{B}_{n-1}^\top(\Lambda_{12} - \Phi)X_t + \mathcal{B}_{n-1}^\top X_{t+1}.
\end{aligned} \tag{A19}$$

Finally, we add pricing errors to the excess bond returns as in Adrian et al. (2013). The coefficients Λ_0 and Λ_{12} can be transformed from regression coefficients of Equation (A18) or Equation (A19). To estimate Equation (A18), we regress $rx_{t+1}^{(n)}$ on Z_t and $\hat{U}_{t+1}$ subject to the constraint $\Lambda_1 = [(I - \Phi)\Gamma, \Lambda_{12}]$. To estimate Equation (A19), we regress $rx_{t+1}^{(n)}$ on X_t and X_{t+1} without any constraints. Either way, we obtain the same estimation results.

B.4 Estimating the Term Structure Model

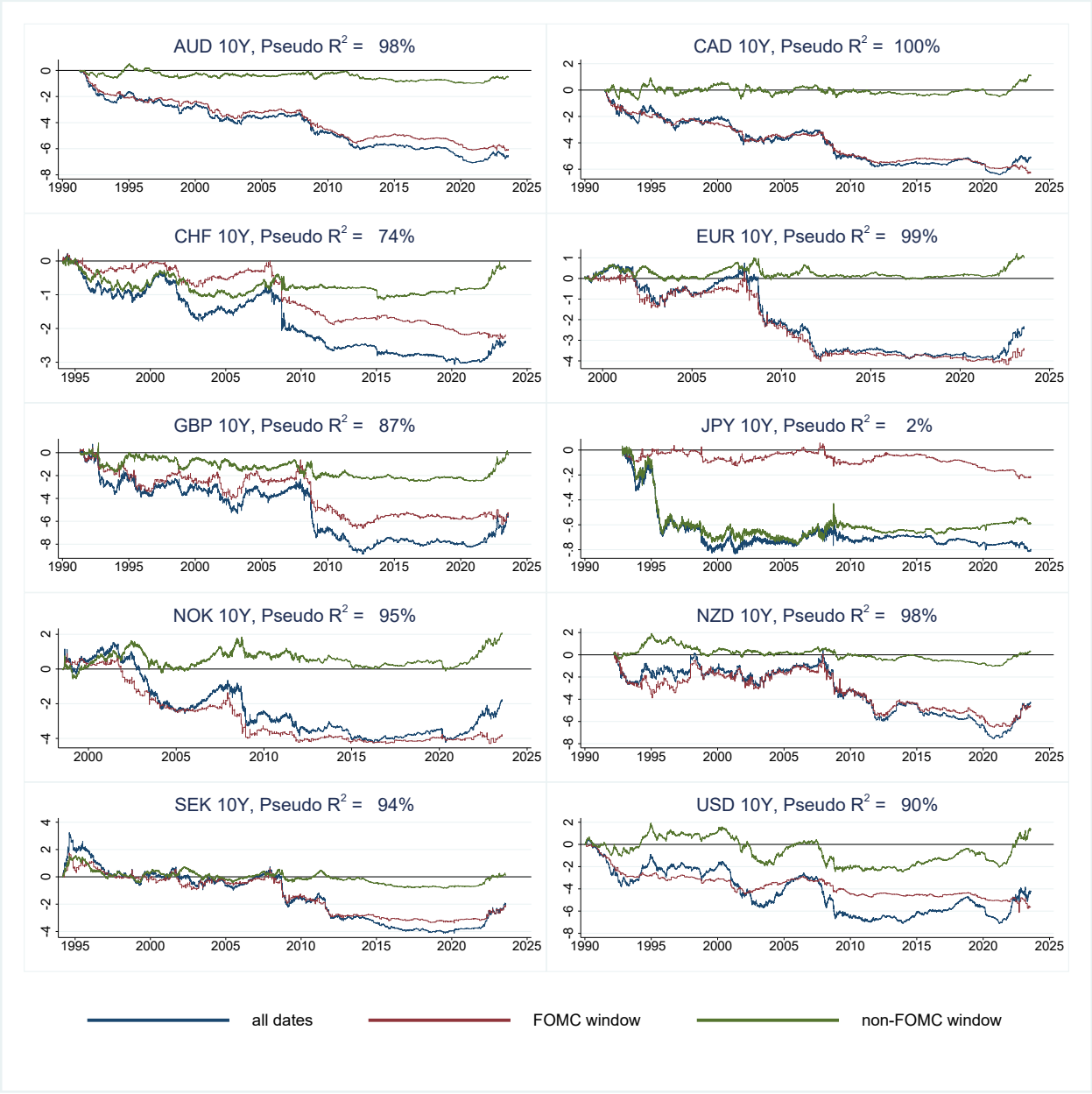
We estimate the term structure model separately for each country. Following Adrian et al. (2013), we take X_t as the first five principal components of the country's yield curve. To investigate the role of monetary policy, as described above, our empirical proxy for $\boldsymbol{\tau}_t$ is $[\nabla PC1_{loc,t}^{US}, \nabla PC1_{US,t}^{US}]^\top$, the first principal components of the cumulative daily changes in the cross-sections of domestic and U.S. zero-coupon yields during the three-day FOMC windows. For the U.S., $\boldsymbol{\tau}_t = \nabla PC1_{US,t}^{US}$. The results are quantitatively similar if we use changes during the FOMC window in individual yields, such as the 10-year yield. The previous sections have demonstrated that $\nabla PC1_{loc,t}^{US}$ and $\nabla PC1_{US,t}^{US}$ well account for the trends of sovereign yields, so we use them to proxy the persistent state variable of the yield curve over the past three decades. As a robustness check, in Appendix B, we demonstrate that the OSE model results are similar to those produced by an estimated shift endpoint model, which estimates $\boldsymbol{\tau}_t$ from the observed yield curve using statistical methods.

Given observed X_t and $\boldsymbol{\tau}_t$, we estimate the term structure model parameters using the regression algorithm described in Adrian et al. (2013). We estimate the parameters using end-of-month observations and plug the daily X_t and $\boldsymbol{\tau}_t$ into the model to get daily estimations of the risk-neutral yields and term premia. For the regressions, we select excess bond returns of the one-month holding period for maturities $n \in \{0.5, 1, 1.5, 2, 2.5, 3, 3.5, 4, 4.5, 5, 7, 10\}$ years. The linear regression approach requires no numerical optimization algorithms, making it much faster than the estimation approach in Bauer and Rudebusch (2020).

B.5 Term Structure Model Results

Figure A1 presents the sum of daily changes in the 10-year risk-neutral yields over each event window. Two patterns are noticeable. First, the risk-neutral rates for all countries have declined steadily over the respective sample periods. This is consistent with the result in Bauer and Rudebusch (2020) for U.S. Treasury

Figure A1: Cumulative changes in the 10-year risk-neutral rates during different event windows.



Notes. Unit of the y-axis is percentage p.a.. The figure presents the sum of daily changes in the 10-year risk-neutral yields over each event window. The pseudo R^2 is defined as $R^2 = 1 - \frac{\sum_{t=1}^{T}\left(y_t^{RNY} - \nabla y_t^{RNY,FOMC}\right)^2}{\sum_{t=1}^{T}\left(y_t^{RNY} - \nabla y_t^{RNY,nonFOMC}\right)^2 + \sum_{t=1}^{T}\left(y_t^{RNY} - \nabla y_t^{RNY,FOMC}\right)^2}$.

yields, arguing that once the term structure model accounts for the persistent interest rate dynamics, the implied risk-neutral yields inherit the trend behavior of observed yields. We extend their findings to an international setup and find consistent results. Second, FOMC windows are responsible for most of the variations in the 10-year risk-neutral yields. Switching back to the definition described by **??**, we use the pseudo R^2 to measure how well the FOMC-window series fits the actual risk-neutral yield. Overall, FOMC windows account for over 70% of the total variations in the 10-year risk-neutral yields. For Australia, Canada, and New Zealand, FOMC windows can explain almost all the variations in the 10-year risk-neutral yields. Again, Japan is an outlier.[13]

FOMC windows have much smaller impacts on the term premia than on the risk-neutral yields. Figure A2 presents the sum of daily changes in the 10-year term premium over each event window. Our model suggests that the term premium has remained stable across countries over the past three decades. Moreover, FOMC windows only explain small fractions of the total variations in the term premium. For example, the pseudo R^2 for Germany suggests that FOMC windows account for 5% of the total variation in the 10-year term premium, while Figure A1 shows that FOMC windows explain 90% of the total variation in the 10-year risk-neutral yield.

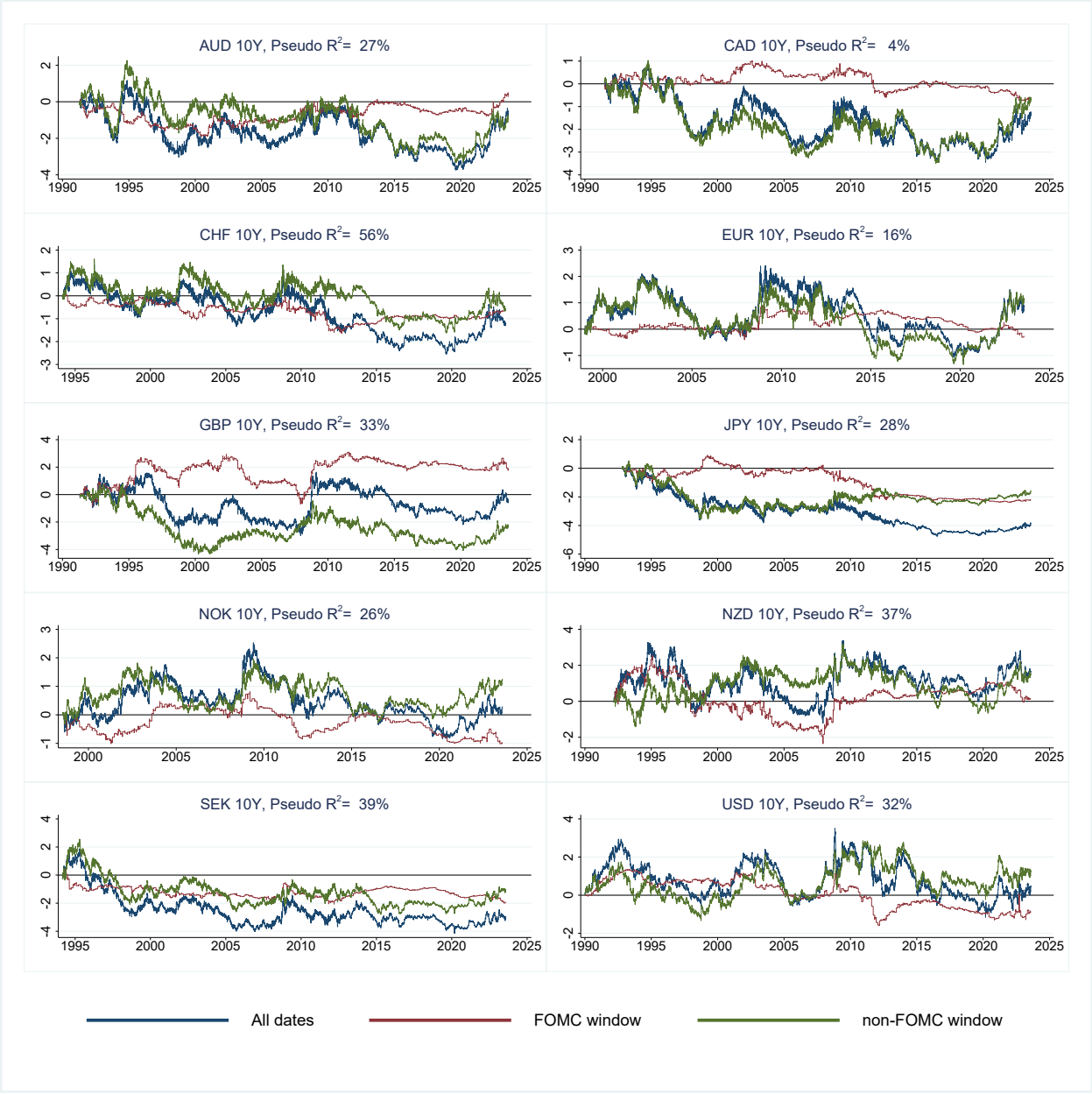
In summary, the responses of world long-maturity bond yields to U.S. monetary policy announcements are primarily attributable to changes in the expected paths of short-term interest rates. Although term premia also fluctuate significantly, their dynamics are predominantly driven by windows outside of U.S. monetary policy announcements windows.

B.6 Model Fit

In the main text, we estimated the daily model using the OSE method, and the empirical proxy for τ_t is the cumulative yield changes during FOMC announcement windows. As a robustness check, we compare the performances of the OSE and ESE models at the monthly frequency. The proxy for τ_t in the monthly model is the end-of-month observations of $[\nabla PC1^{US}_{loc,t}, \nabla PC1^{US}_{US,t}]^\top$. The ESE model uses τ_t estimated from the observed end-of-month yields. In Table A4, Table A5, and Table A6, we report the root mean squared errors of yield curve fitting. Specifically, we compute the squared difference between the yield data and model-

[13] In Appendix E, we compare the risk-neutral rates implied by the OSE model and the canonical FE model from Adrian et al. (2013), and investigate their dynamics during FOMC windows. We conclude that the FOMC windows account for substantial fractions of the total variations in world risk-neutral rates, regardless of the model for decomposing the yields.

Figure A2: Cumulative changes in the 10-year term premium during FOMC windows.



Notes. Unit of the y-axis is percentage p.a.. The figure presents the sum of daily changes in the 10-year term premium over each event window. The pseudo R^2 is defined as $R^2 = 1 - \frac{\sum_{t=1}^{T}\left(y_t^{TP} - \nabla y_t^{TP,FOMC}\right)^2}{\sum_{t=1}^{T}\left(y_t^{TP} - \nabla y_t^{TP,nonFOMC}\right)^2 + \sum_{t=1}^{T}\left(y_t^{TP} - \nabla y_t^{TP,FOMC}\right)^2}$.

implied yields. We then compute the square root of the average squared difference over the full sample. For each country, we report the RMSE associated with the 1-, 2-, ..., 10-year maturities and the mean RMSE across all maturities.

Table A4: RMSE of the monthly OSE model.

	AUD	CAD	CHF	DKK	EUR	GBP	JPY	NOK	NZD	SEK	USD
1	1.46	1.51	1.59	1.25	1.46	0.51	1.03	0.52	0.72	1.08	0.55
2	1.20	0.70	1.50	1.16	1.52	0.37	2.05	0.47	1.47	1.21	0.43
3	0.65	0.87	0.86	0.94	1.20	0.49	1.00	0.29	0.75	1.02	0.42
4	1.43	0.81	1.31	1.45	1.92	0.55	1.75	0.34	1.32	1.72	0.46
5	1.46	0.64	1.17	1.25	1.63	0.37	2.14	0.45	1.09	1.58	0.25
6	1.57	0.72	1.15	1.68	1.56	0.46	3.06	0.65	0.97	1.39	0.34
7	2.91	0.86	2.23	2.89	3.36	0.58	4.75	1.21	1.65	1.85	0.50
8	4.70	0.81	3.26	4.15	5.30	0.48	6.93	3.16	2.08	2.39	0.44
9	6.37	0.93	3.73	5.42	6.72	0.48	9.79	6.48	2.03	3.18	0.39
10	8.45	1.86	5.73	7.52	8.98	1.30	14.06	9.19	4.57	5.80	0.99
Mean	3.50	1.22	2.45	3.14	3.87	0.59	5.40	3.15	1.78	2.41	0.56

Notes: Root mean squared errors of yield curve fitting, in basis points. For each country, we report the RMSE for the 1-, 2-, ..., 10-year maturities and the mean RMSE across all maturities.

Table A5: RMSE of the monthly ESE model.

	AUD	CAD	CHF	DKK	EUR	GBP	JPY	NOK	NZD	SEK	USD
1	2.93	1.49	3.00	3.01	1.78	0.55	3.86	0.67	0.86	1.13	0.58
2	1.22	0.80	3.28	1.28	1.92	0.36	2.84	0.44	1.61	1.31	0.42
3	0.78	1.00	2.19	0.90	1.62	0.50	2.21	0.21	0.81	1.20	0.50
4	2.13	1.20	2.61	2.82	1.81	0.56	2.38	0.14	1.38	1.63	0.60
5	2.68	0.76	2.61	3.80	1.50	0.35	1.77	0.24	1.16	1.31	0.42
6	2.68	0.64	2.14	4.19	0.97	0.41	0.95	0.25	0.91	0.80	0.39
7	2.83	0.85	2.07	4.76	1.84	0.53	1.21	0.32	1.32	1.22	0.47
8	3.41	0.65	2.21	6.25	2.77	0.42	1.54	0.56	1.52	1.43	0.37
9	4.77	0.56	3.48	9.79	3.32	0.27	1.66	1.16	0.85	0.91	0.29
10	8.65	1.88	6.97	17.20	6.43	1.10	4.61	1.96	3.78	3.30	0.99
Mean	3.83	1.06	3.36	7.11	2.82	0.55	2.56	0.80	1.65	1.57	0.54

Notes: Root mean squared errors of yield curve fitting, in basis points. For each country, we report the RMSE for the 1-, 2-, ..., 10-year maturities and the mean RMSE across all maturities.

Table A6: RMSE of the daily OSE model.

	AUD	CAD	CHF	DKK	EUR	GBP	JPY	NOK	NZD	SEK	USD
1	2.97	4.48	2.75	2.22	1.90	1.29	3.19	1.05	0.79	1.19	0.85
2	1.26	3.35	2.86	1.49	2.31	1.01	2.63	0.68	1.52	1.35	0.66
3	0.73	3.44	1.82	0.93	2.33	1.12	2.42	0.83	0.80	1.11	0.99
4	2.06	3.20	2.33	2.35	2.38	1.23	2.33	1.10	1.36	1.91	1.04
5	2.51	2.75	1.99	2.56	1.82	1.05	1.75	0.87	1.05	1.61	0.97
6	2.47	2.40	1.71	2.31	1.06	0.86	1.23	0.39	0.69	0.94	1.05
7	2.78	2.19	1.86	2.33	1.69	0.74	1.46	0.61	1.28	1.29	1.20
8	3.63	2.03	2.19	2.74	2.50	0.51	1.79	0.41	1.64	1.52	1.27
9	5.33	2.19	3.44	4.25	3.10	0.29	1.95	1.64	1.23	0.84	1.31
10	9.41	3.17	6.23	8.51	6.10	1.22	4.62	2.87	3.78	3.66	1.59
Mean	3.55	3.45	2.81	3.15	2.63	1.03	2.34	1.16	1.59	1.91	1.16

Notes: Root mean squared errors of yield curve fitting, in basis points. For each country, we report the RMSE for the 1-, 2-, ..., 10-year maturities and the mean RMSE across all maturities.

Table A7: Standard deviation of daily yield changes.

	FOMC	FOMC without Dot Plot	Dot Plot	Chair	Other
AUD	11.11	11.59	8.19	10.68	10.06
CAD	9.82	10.47	5.46	9.28	8.72
CHF	9.70	10.44	5.32	9.13	8.60
EUR	10.66	11.23	7.06	9.88	9.25
GBP	5.83	5.80	5.95	5.31	6.64
JPY	9.19	9.98	3.86	8.28	7.94
NOK	9.94	10.65	6.29	10.32	9.84
NZD	10.05	10.58	7.02	9.29	9.09
SEK	10.76	11.45	6.72	9.59	9.60
USD	6.49	6.67	5.27	5.91	5.67

Notes. The table reports standard deviations of daily yield changes over three sets of dates: three-day event windows bracketing FOMC announcements, three-day event windows bracketing Fed Chair speeches, and other dates. The units are annualized basis points.

C Volatility of Yield in Fed Chair Speech dates

The weak relationship between Fed Chair speeches and the secular decline of government bond rates does not imply that these speeches have no impact on asset prices. As highlighted by Swanson and Jayawickrema (2023), Fed Chair speeches significantly influence various asset classes in the U.S., including futures, equities, and the entire term structure of government bond yields. In Table A7, we report the standard deviations of daily changes in world 5-year, 10-year, and 5-5 forward rates during different event windows. "FOMC" and "Chair" refer to the three-day event windows centered around the FOMC announcement dates and Federal Reserve Chair speech dates, respectively. "Other" pertains to days belonging to neither of these windows. All countries' government bond rate changes during the FOMC window are more volatile than those on normal days. Additionally, yield changes in eight out of ten countries during the Fed Chair speech window are more volatile than those during other days. Overall, the findings of Swanson and Jayawickrema (2023) for the U.S. financial market can be generalized to G10 countries, but the reactions to Fed Chair speeches are much less monotonic than those to FOMC announcements, thus failing to explain the secular decline of government bond rates.

D Covered Interest Parity Decomposition Results

Similar to the UIP decomposition, we can conduct a covered interest parity (CIP) decomposition

$$y_t^{(10)*} = y_t^{(10)} + \frac{1}{10}\left(f_{t,t+10} - s_t\right) + \eta_t,$$

where $FP_t \equiv \frac{1}{10}\left(f_{t,t+10} - s_t\right)$ is the forward premium, $f_{t,t+n}$ is the log n-year outright forward exchange rate, and s_t is the log spot exchange rate, all measured in foreign currency per dollar. η_t is the CIP deviation of government bond, in the recent literature, this deviation has been used extensively as a measure of the convenience yield of a country's government securities relative to other countries (see Jiang et al. (2021), Engel and Wu (2023), Du et al. (2018)). Corsetti et al. (2023) studies the 10-year convenience yield and connects it to the long-run risk of the U.S. government.

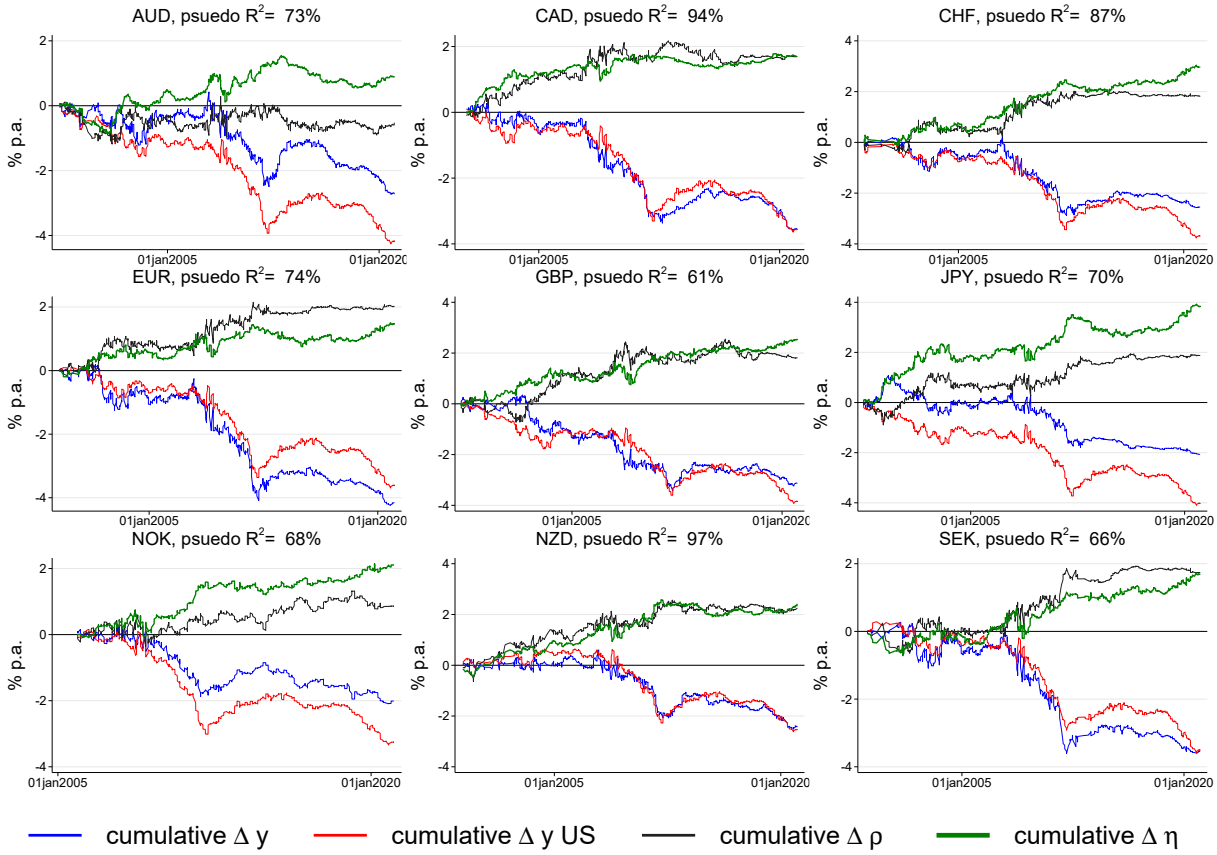
We rewrite the CIP condition in the form of daily changes

$$\Delta_1 y_t^{(10)*} = \Delta_1 y_t^{(10)} + \Delta_1 \frac{1}{10}\left(f_{t,t+10} - s_t\right) + \Delta_1 \eta_t. \tag{A20}$$

In Figure A3, we plot the cumulative sum of each variable in equation (A20). The start dates of this figure differ from those in the UIP figure due to the limited availability of long-maturity forward premium data in the early sample. Despite this slight discrepancy, Figure A3 consistently shows that cumulative changes in U.S. government interest rates spill over almost perfectly to foreign interest rates. In most cases, the U.S. convenience yield declines during the FOMC window, but this is offset by a decrease in the forward premium. This suggests that the dollar initially depreciates on impact, while the forward market anticipates future appreciation of the U.S. dollar following a decline in U.S. interest rates.

In summary, both UIP and CIP decompositions indicate a strong pass-through of U.S. interest rates to foreign interest rates during FOMC windows. In both cases, the typical shock absorbers—such as the exchange rate, forward premium, or convenience yield—play little to no role in adjusting to external conditions driven by the U.S..

Figure A3: CIP decomposition within the FOMC windows (Equation (A20))



Notes: The CIP condition is $FP_t - (y_t^{(10)} - y_t^{(10)*}) = \eta_t$. The figure plots the cumulative sums of daily changes in $f_{t,t+10y} - s_t$ (green), $y_t^{(10)} - y_t^{(10)*}$ (blue), and η_t (red) during FOMC windows. The interest rates are the 10-year government bond yields, and the exchange rates match the bond maturity.

E Additional Results

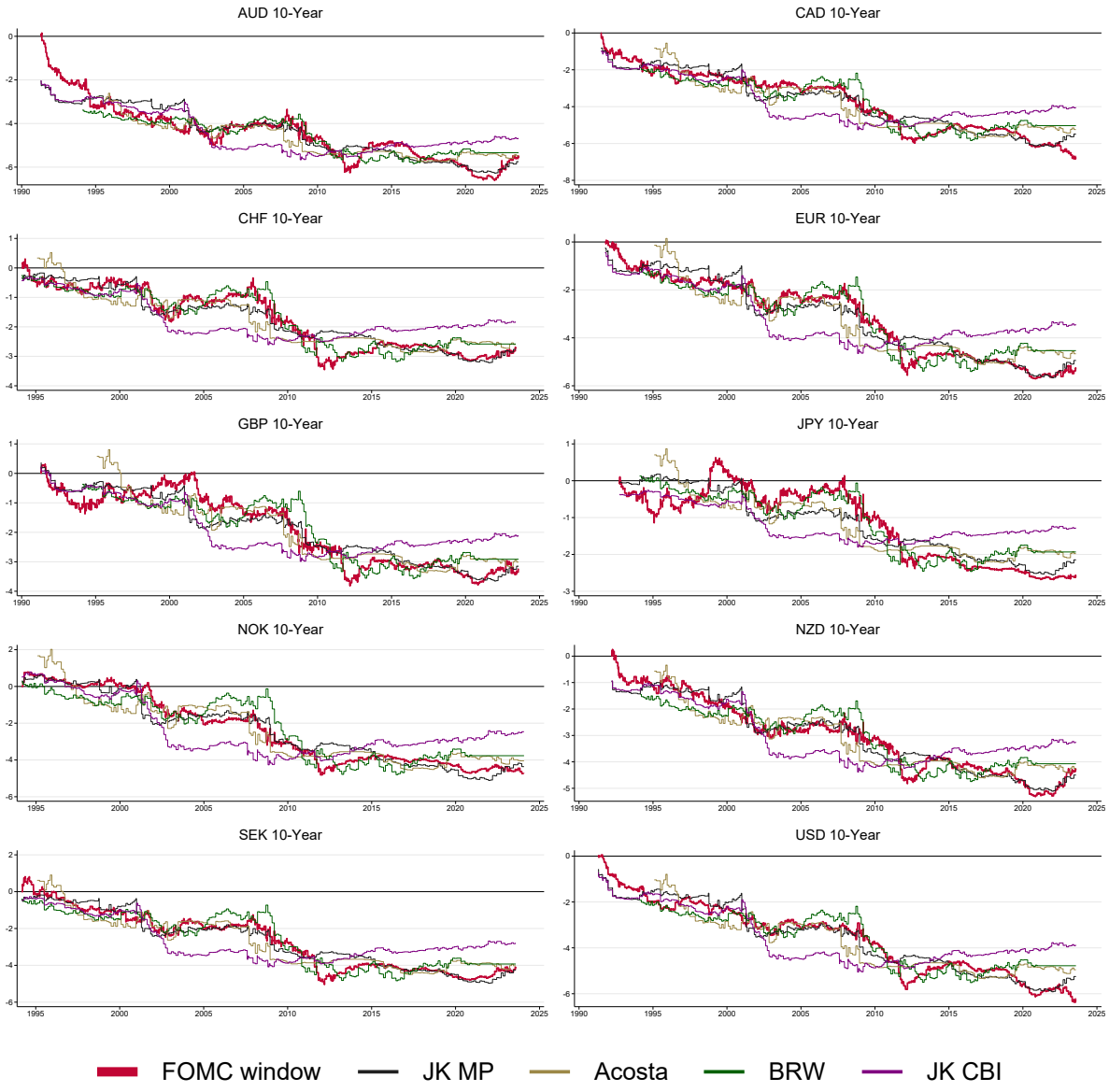
E.1 Purified Monetary Policy Shocks

In Figure A4, we present the fitted value from the cointegration regression $\nabla y_t^{FOMC} = \gamma_0 + \gamma_1 CMPS_t + u_t$ and the regression R^2. Here, MPS_t is a "purified" U.S. monetary policy shock, and $CMPS_t = \sum_{s=0}^{t} MPS_s$

is the sum of historical shocks up to t. In addition to the monetary policy shock and central bank information shock from Jarociński and Karadi (2020) investigated in the main text, we also consider the monetary policy shocks from Acosta (2023) and Bu et al. (2021). We omit the purified shocks that are OLS residuals, such as those in Bauer and Swanson (2023b) and Miranda-Agrippino and Ricco (2021). As a property of OLS, the residuals must sum to zero. Thus, the cumulative sum of this class of shocks must eventually end up at zero by construction, which may not necessarily reflect the true nature of cumulative monetary policy shocks. Interestingly, the cumulative sum of Bauer and Swanson (2023b) "orthogonalized MPS" is V-shaped, exhibiting a persistent downward trend in the early half of the sample before monotonically returning to zero to satisfy the OLS restriction.

Next, we plot the cumulative effects of each country's central bank monetary policy announcements on world 10-year yields. In each of Figure A5-A13, we fix the central bank and plot its cumulative announcement effects on G10 ten-year yields. All samples start in Feb 2000 such that the sample period covers all central banks' announcements in our record.

Figure A4: Cumulative yield changes and sums of U.S. HF monetary policy shocks.



Notes. The figure presents ∇y_t^{FOMC}, the fitted value from the cointegration regression $\nabla y_t^{FOMC} = \gamma_0 + \gamma_1 CMPS_t + u_t$, and the regression R^2. Here,MPS_t is a "purified" U.S. monetary policy shock, and $CMPS_t = \sum_{s=0}^{t} MPS_s$ is the sum of historical shocks up to t. "JK MP" and "JK CBI" are the monetary policy shock and central bank information shock in Jarociński and Karadi (2020); "Acosta" is the Federal funds rate shock from Acosta (2023); and "BRW" is the monetary policy shock from Bu et al. (2021).

Figure A5: Sum of daily yield changes during U.S. or Australian central bank announcement windows.

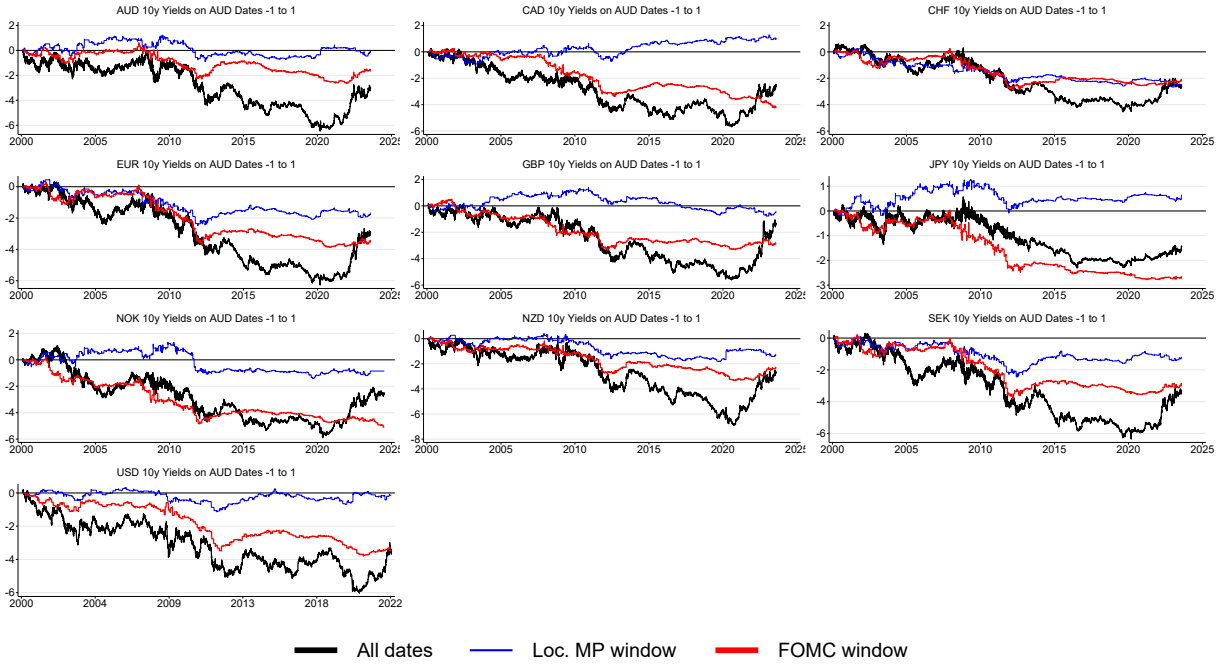


Figure A6: Sum of daily yield changes during U.S. or Canadian central bank announcement windows.

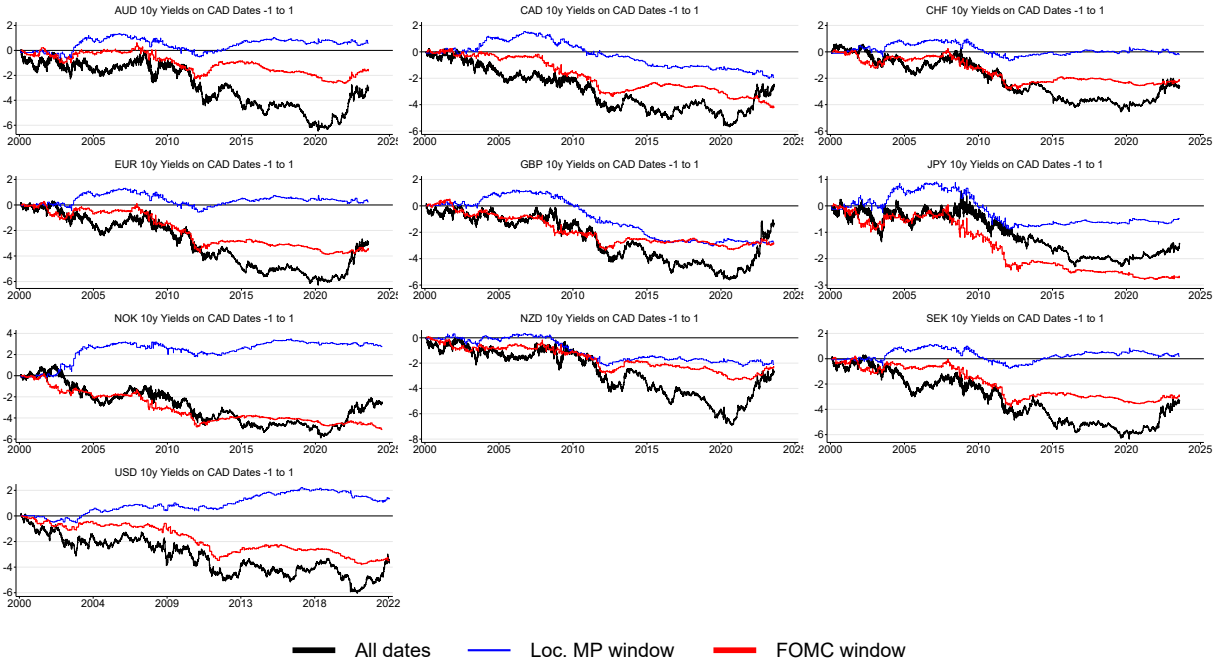


Figure A7: Sum of daily yield changes during U.S. or Swiss central bank announcement windows.

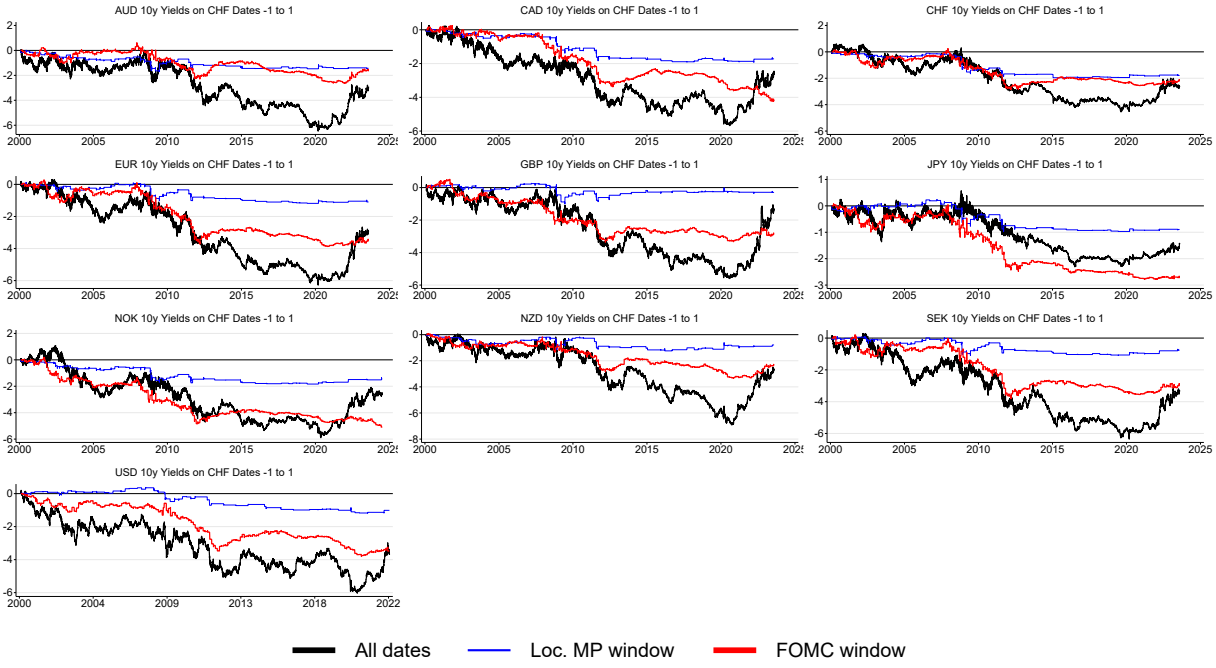


Figure A8: Sum of daily yield changes during U.S. or ECB announcement windows.

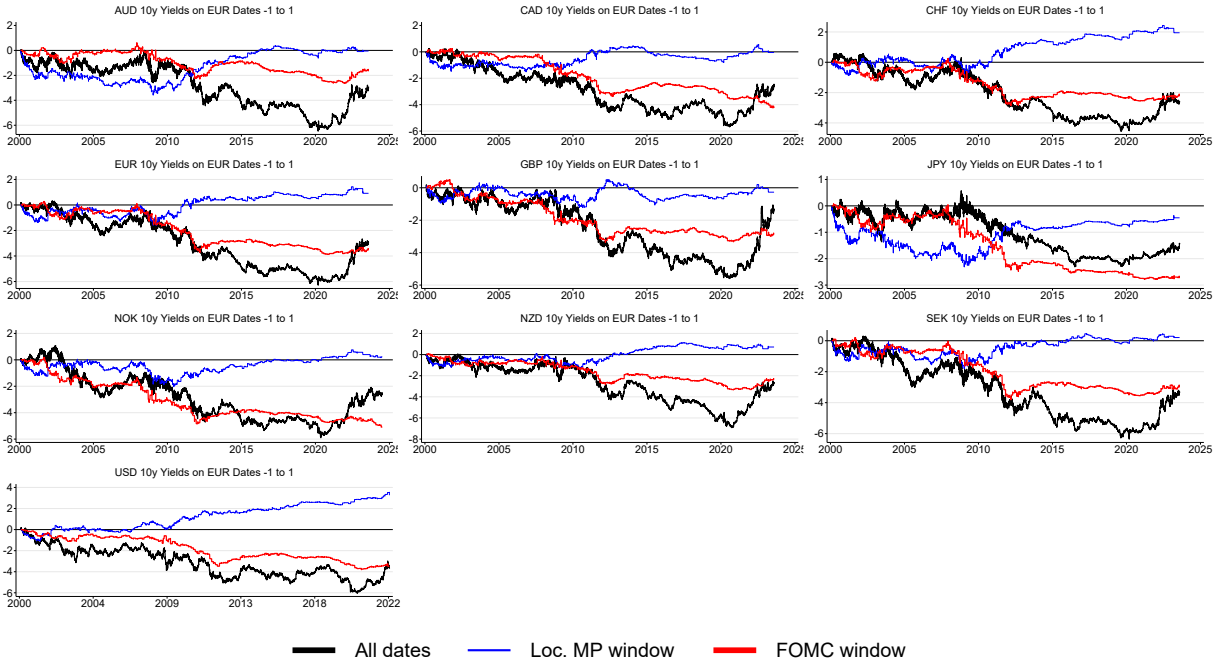


Figure A9: Sum of daily yield changes during U.S. or U.K. central bank announcement windows.

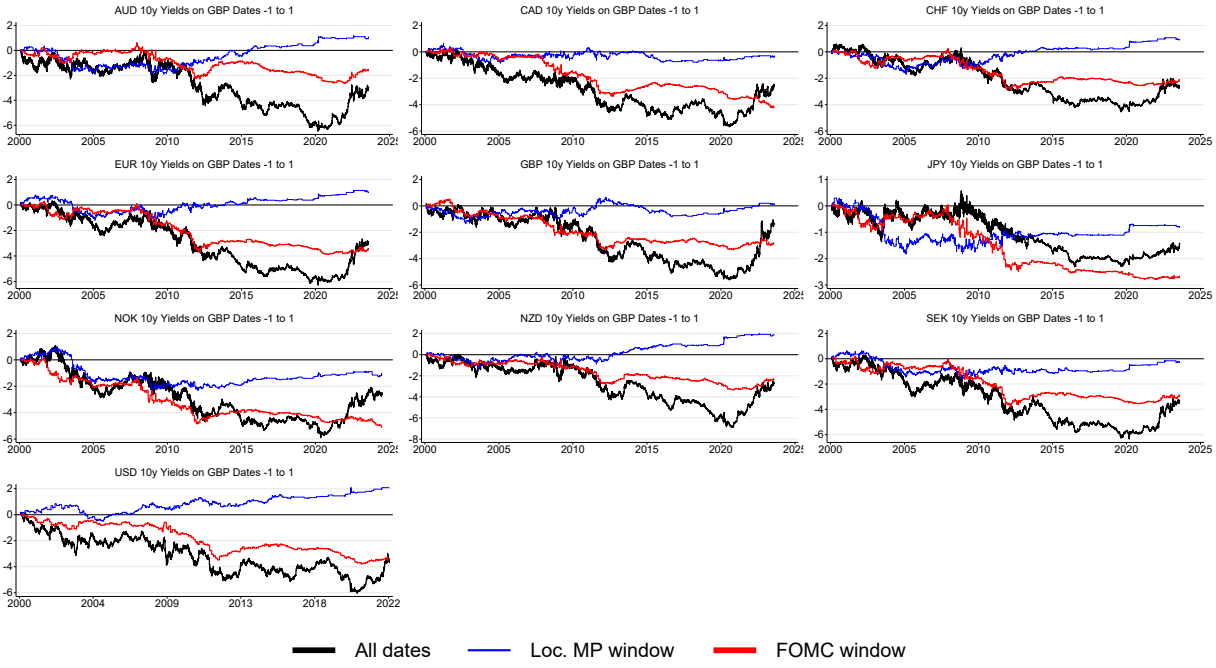


Figure A10: Sum of daily yield changes during U.S. or Japanese central bank announcement windows.

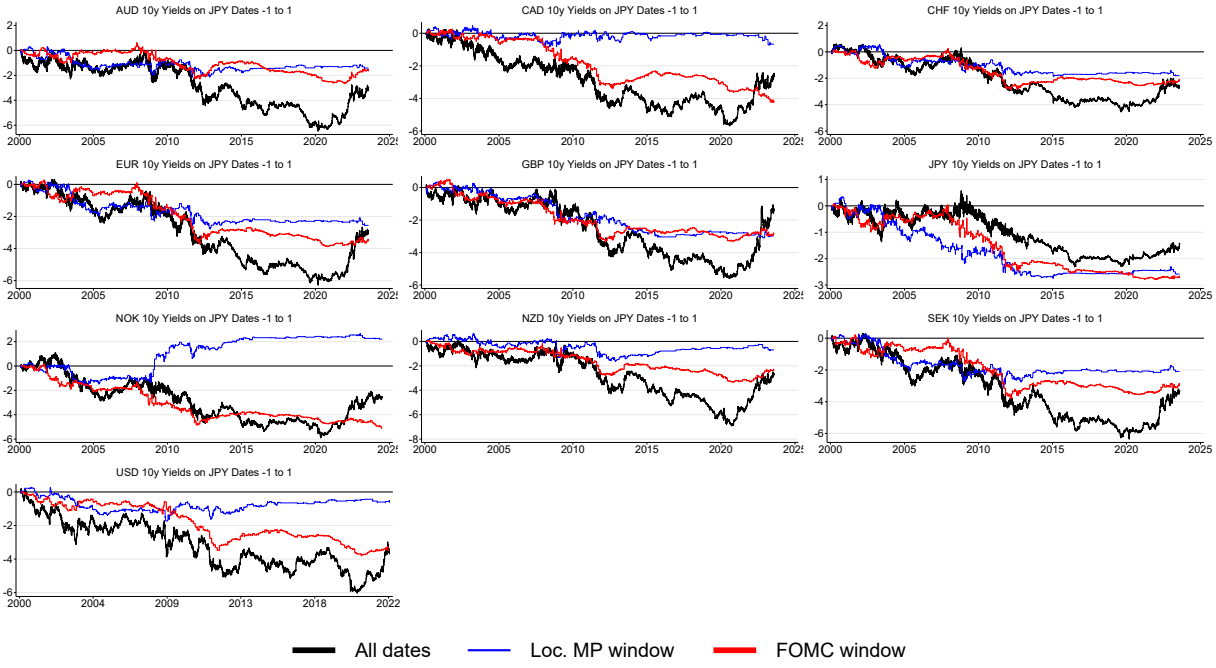


Figure A11: Sum of daily yield changes during U.S. or Norwegian central bank announcement windows.

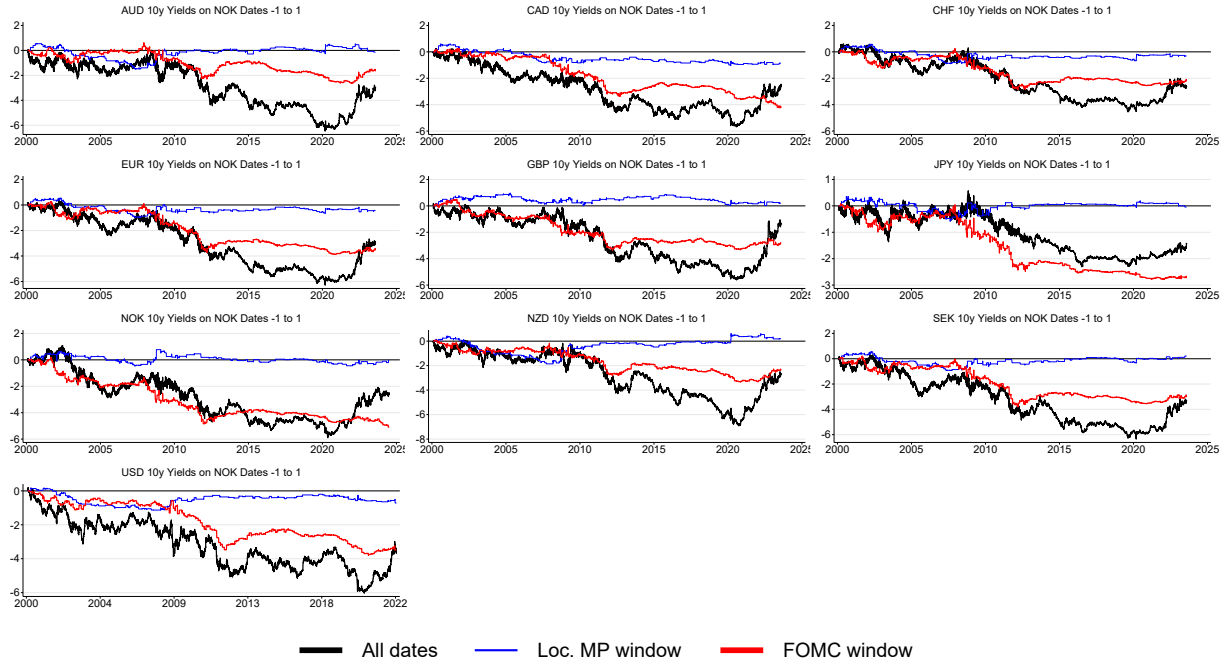


Figure A12: Sum of daily yield changes during U.S. or New Zealand central bank announcement windows.

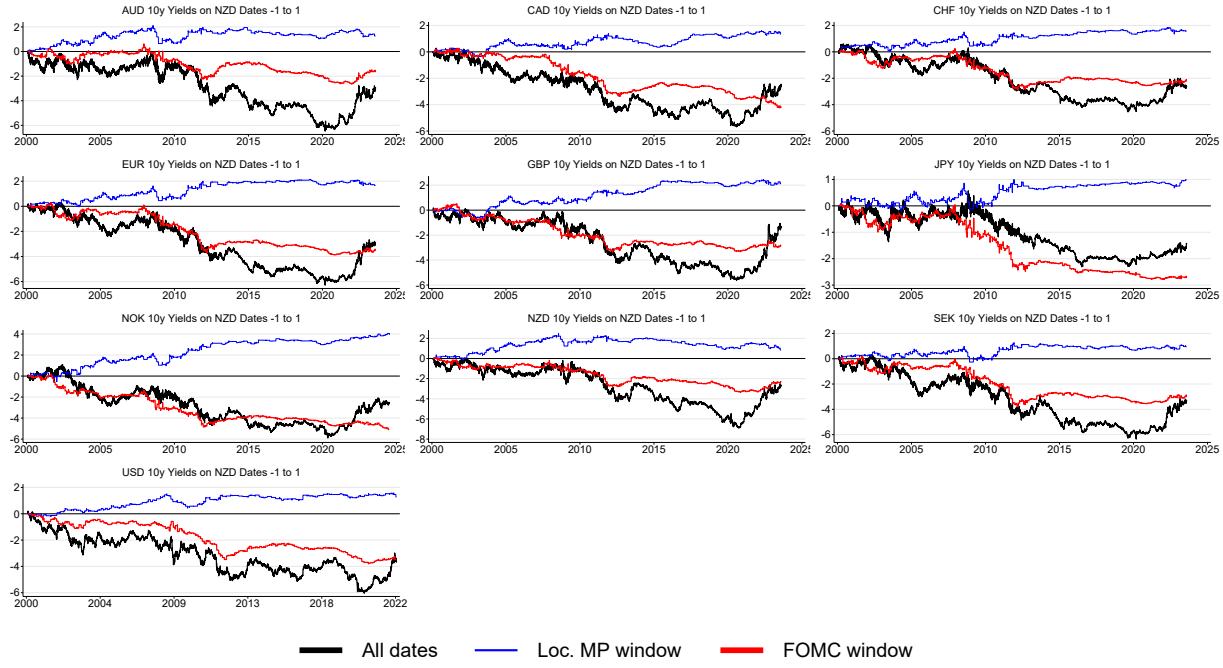


Figure A13: Sum of daily yield changes during U.S. or Swedish central bank announcement windows.

