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Monetary Policy and the Secular Decline in Long-Term Interest Rates: A Global Perspective*

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Abstract

We demonstrate that almost 70% of the secular decline in long-term interest rates across advanced economies between the early 1990s and 2023 occurred in the three days surrounding U.S. monetary policy announcements (FOMC windows). By contrast, other central banks' announcements had only limited effects, if any, on the long-run direction of long-term interest rates, both domestically and across countries. The persistent global effect of the FOMC window reflects the combination of the concentration of declines in U.S. bond yields in this window and large interest rate spillovers from the U.S. to other countries. We further find that the decline in interest rates during FOMC windows is closely associated with pure monetary policy shocks and not with information effects. Moreover, the rate decline on FOMC announcement days is primarily driven by changes in real and expected short rates rather than inflation expectations and term premia. These findings highlight the pivotal role of U.S. monetary policy news in shaping global long-term interest rate dynamics.

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

JEL Codes: E43, E52, F42

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1 Introduction

Over the past three decades, long-term interest rates in major advanced economies have fallen significantly. This secular decline is commonly attributed to structural factors, including falling productivity growth, excess global saving, demographic shifts, and a decline in capital investment opportunities (see Bernanke (2005), Carvalho et al. (2016), and Summers (2014) among others). At the same time, there is a growing literature examining the link between long-term interest rates and monetary policy (see Cochrane and Piazzesi (2002), Gürkaynak et al. (2005), Hanson and Stein (2015), Nakamura and Steinsson (2018), Brooks et al. (2018), Adrian et al. (2024). In a recent paper, Hillenbrand (2025) provides evidence suggesting that, in the U.S., narrow windows surrounding monetary policy meetings account for much of the secular decline in long-term rates over recent decades.

In this paper, we empirically examine the link between long-term interest rates and monetary policy announcements worldwide. Focusing on the 10-year government bond yields of the G10 currencies,[1] we show that the global secular decline in long-term interest rates over the past three decades is largely driven by market dynamics in the three-day windows around U.S. Federal Open Market Committee (FOMC) monetary policy announcements. Specifically, changes in long-term bond yields in FOMC announcement windows account on average for almost 70% of the total decline in the 10-year yields of Australia, Canada, the euro area, Norway, New Zealand, Sweden, and the United States. By contrast, other central banks' announcements are generally not associated with any persistent effects on long-term interest rates, both domestically and internationally.

Figure 1 illustrates these findings for the case of the euro area and New Zealand, the largest and smallest economies covered by our analysis, respectively. The charts show the cumulative daily yield changes of the 10-year Euro (German) and New Zealand government bond yields together with the cumulative yield changes over 3-day U.S. FOMC announcement windows and over the corresponding 3-day domestic central bank announcement windows. Three key messages emerge from the data: first, there is an overall consistent decline of 10-year bond yields; second, the cumulative decline during FOMC windows closely aligns with the overall decline in interest rates; and third, the cumulative changes in yields during domestic central bank monetary policy announcement windows has not contributed to the overall interest rate trend. In summary,

[1] The G10 currencies include the Australian Dollar, the Canadian Dollar, the Swiss Franc, the Euro, the British Pound, the Japanese Yen, the Norwegian Krone, the Swedish Krona, and the U.S. Dollar. We use German bond yields for the Euro. Our study focuses on the 10-year bond yield because it is the longest maturity commonly available across currencies.

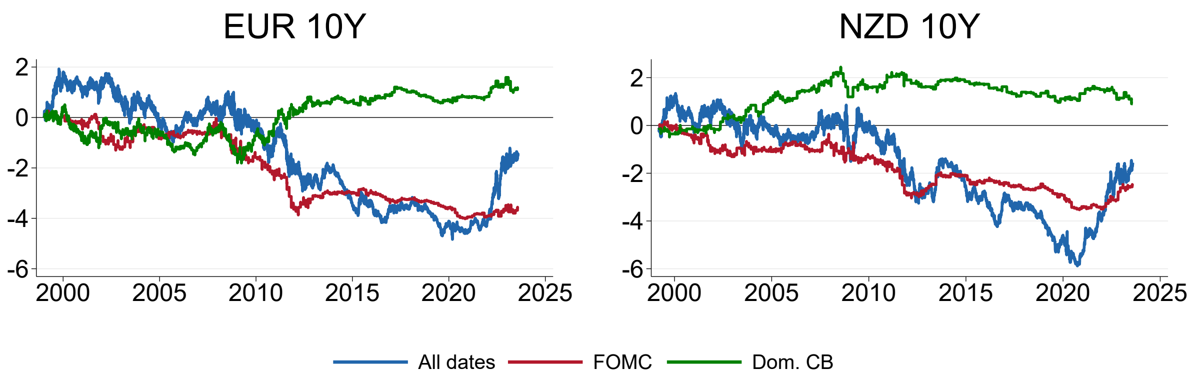

Figure 1: Cumulative changes in nominal 10-year Euro (German) and New Zealand government bond yields during FOMC and domestic monetary policy announcement windows.

Notes. Unit of the y-axis is percentage p.a. The sample starts from 2000, which is later than the baseline sample period in our main analysis due to the availability of non-US central bank announcement data. The figure plots the cumulative sum of the daily changes in the 10-year yield over the full sample or monetary policy announcement windows. "All dates" refers to the actual yield, "FOMC" represents three-day windows around FOMC announcements, and "Dom. CB" represents three-day windows around domestic monetary policy announcements.

the secular decline in long-term interest rates is primarily linked to U.S. monetary policy announcements but not to domestic policy announcements.

These results align with well-documented international monetary policy spillovers but highlight an important distinction: while monetary policy announcements from major non-U.S. central banks also generate cross-country effects (Kearns et al. (2023)), only FOMC announcements are associated with a sustained long-term decline in global interest rates. In contrast, non-U.S. central bank announcements tend to produce cyclical or even rising cumulative effects, different from the persistent downward trend in observed yields. This finding results from the combination of two effects: the large and significant interest rate spillovers from the U.S. to other countries (see e.g. Obstfeld (2015), Hofmann and Takats (2015), Kearns et al. (2023), Albagli et al. (2019)) and the large cumulative negative effect of FOMC announcements on U.S. long-term bond yields previously documented by Hillenbrand (2025) and also demonstrated here.

To examine the underlying channels linking global interest rates to FOMC announcements, we assess long-term interest rate dynamics from different perspectives. We begin by evaluating whether yield movements during FOMC announcement windows are driven by news about the setting of policy rates (monetary policy shock) or news about the macroeconomic outlook inferred from the policy decision (information shock).

Utilizing state-of-the-art high-frequency shocks from Jarociński and Karadi (2020), Bu et al. (2021), and Acosta (2023), we find that the decline in interest rates during FOMC windows closely aligns with the "purified" monetary shock series that is free of the "Fed information effect". Regressions of the cumulative yield change on the decomposition of the "purified" monetary shock and Fed information shock from Jarociński and Karadi (2020) show that the "purified" monetary shock explains substantially more variation in the yield change than the information shock.

In light of Swanson and Jayawickrema (2023), who highlight the market impact of Federal Reserve Chair speeches, we assess the impact of these speeches on global bond yields. To this end, we partition our sample into three-day windows around Fed Chair speech dates and non-speech dates. The speech events serve as another proxy for the information channel as policy rate does not change when the Chair delivers the speeches. In terms of mean absolute values and standard deviation of daily yield changes, our findings for G10 currency yields are consistent with those of Swanson and Jayawickrema (2023). However, while global bond yields fluctuate significantly during Fed Chair speech days, the cumulative daily changes in the three-day speech window hardly account for any of the secular decline in world interest rates over the past three decades. This suggests that while the information channel may drive short-term fluctuations, it does not explain the long-term trend in global interest rates.

We further investigate the contributions of changes in short-rate expectations and in term premia to the cumulative responses of long-term interest rates to U.S. monetary policy announcements. Using a dynamic term structure model incorporating trends as outlined by Bauer and Rudebusch (2020), we estimate daily risk-neutral rates and term premia for each country. Our analysis yields two key results. First, the persistent decline in global interest rates during FOMC announcement windows is primarily driven by reductions in risk-neutral rates, while term premia remain stationary during these periods. Second, FOMC windows account for over 85% of the total variation in world 10-year risk-neutral rates across all dates. Consistent with Albagli et al. (2019), this finding highlights a significant risk-neutral rate channel in monetary transmission driving the secular decline in global long-term interest rates.

Finally, while our main results focus on nominal yields, we demonstrate that similar findings hold for real rates using data from inflation-protected securities. The cumulative impact of FOMC news on global interest rates has hence been driven by declines in long-term real interest rates, rather than inflation expectations.

Literature review

Since the seminal work by Kuttner (2001), Bernanke and Kuttner (2005), Gürkaynak et al. (2005), Gürkaynak et al. (2005), and Gertler and Karadi (2015), a rich body of literature has leveraged high-frequency changes in interest rates or interest rate futures to identify monetary policy shocks and assess their effects on various economic variables. Important refinements to the shock measures have been presented by Bauer and Swanson (2023b), Jarociński and Karadi (2020), Altavilla et al. (2019), Bu et al. (2021), Acosta (2023) and Boehm and Kroner (2024). These contributions enhance our understanding of how monetary policy impacts financial markets and the broader economy. Our paper is closely related to the work of Hillenbrand (2025), who demonstrated that the three-day window surrounding FOMC monetary policy announcement dates accounts for much of the secular decline in long-term interest rates in the U.S. We make three key contributions to this literature. First, we systematically study the effects of monetary policy on long-term interest rates across the G10 countries. Second, we emphasize the role of FOMC announcements in explaining the global secular trend in interest rates, while announcements from other central banks do not have comparable effects, even on their domestic yields. Third, we assess various hypotheses on the nature of the relationship between monetary policy announcements and the secular decline in interest rates. Specifically, we find that this relationship cannot be explained by information channels, term premiums, or inflation expectations. In essence, our analysis parallels the approaches of Nakamura and Steinsson (2018) and Hanson and Stein (2015) but instead investigates the cumulative effects across countries.

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 Obstfeld (2015), Hofmann and Takats (2015), Miranda-Agrippino and Ricco (2021), Albagli et al. (2019), Dedola et al. (2017), Gilchrist et al. (2019), Miranda-Agrippino and Rey (2020), and Brusa et al. (2020), who focus on the role of the Federal Reserve in generating global monetary policy spillovers. Additionally, Gerko and Rey (2017), Jarociński (2022), Miranda-Agrippino and Nenova (2022), and Kearns et al. (2023) highlight that also other major central banks besides the Federal Reserve, in particular the European Central Bank, can induce significant cross-border monetary policy spillovers. Rather than 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 global bond yields.

Our analysis is further connected to the literature on the secular decline of the natural rate of interest, defined as the level of the real interest rate that will prevail in steady state.[2] This literature identifies several key macroeconomic and financial drivers. Demographic shifts, particularly changes in fertility and mortality rates, significantly influence the natural rate by affecting economic growth, dependency ratios, and aggregate saving for retirement (Auclert et al. (2021); Carvalho et al. (2016); Gagnon et al. (2021)). On the financial side, international capital flows and the scarcity of safe assets are critical factors. Emerging markets offer alternative investment opportunities, raising natural rates in advanced economies (Obstfeld (2023)), while the limited supply of safe assets—especially U.S. government bonds—pushes up their prices and lowers returns (Bernanke (2005); Bárány et al. (2023); Caballero et al. (2008); Del Negro et al. (2017); Krishnamurthy and Vissing-Jorgensen (2012)). Additionally, market power plays a role, as increased concentration can suppress future investment demand while redirecting dividends from labor to capital owners, leading to mixed effects on the natural rate (Ball and Mankiw (2023); Eggertsson et al. (2019); Platzer and Peruffo (2022)). Finally, productivity growth, by increasing the marginal product of capital, raises interest rates to incentivize lending (Cesa-Bianchi et al. (2022); Mankiw (2022)). While we do not directly estimate natural rates in this paper, we introduce a new monetary perspective to the debate regarding the drivers of the trend in equilibrium long-term interest rates.

The remainder of the paper is organized as follows. Section 2 describes the data. Section 3 describes our baseline empirical methods and findings, assessing the cumulative effects of G10 currency central banks' monetary policy announcements on international sovereign yields Section 4 assess the role of pure monetary policy vs information effects in driving the global effects of FOMC announcements. Section 5 investigates the role of risk-neutral rates and term premia in driving the overall response of long-term bond yields to FOMC news. Section 6 investigates the cumulative effects of FOMC announcements on real interest rates. Finally, Section 7 examines the statistical significance of cumulative FOMC announcements effects based on placebo tests. Section 8 concludes.

[2]For a recent assessment of the concept, see Benigno et al. (2024).

2 Data

2.1 Daily sovereign yields

Our main source of daily sovereign yields is Bloomberg, following Du et al. (2018). 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 yearly maturities from 1 to 10 years, except for 6, 8, and 9 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.

Table 1: Sources of daily sovereign bond 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: All charts 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.

2.2 Central bank announcements

FOMC announcement dates. The Federal Open Market Committee (FOMC) is in charge of conducting U.S. monetary policy. Since 1981, it has typically held eight scheduled meetings per year. Most monetary policy decisions since 1994 were 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 therefore excluded from our list of FOMC announcement dates.

In line with Hillenbrand (2025), our FOMC announcement dates correspond to the time when the public received 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 is in June 1989. After 1994, monetary policy decisions were predominantly made during scheduled FOMC meetings, with the Fed releasing a statement. Consequently, we utilize the publication dates of these statements as FOMC announcement date. In Appendix A, we list the all FOMC announcement dates covered by our analysis.

Fed chair speeches. We obtain the dates of the Federal Reserve Chair speeches from the Federal Reserve Board's official website: https://www.federalreserve.gov/newsevents/speeches.htm. The website documents the dates of speeches by the Federal Reserve Board Chair, Vice Chair, and other officials from 1996 onwards. Swanson and Jayawickrema (2023) demonstrate that the speeches by the Chair are much more important than those by the Vice Chair or other officials for interest rates and stock prices, so we focus on world interest rate dynamics around the Chair speech dates.

Other central banks' monetary policy announcements. Whenever possible, we collect monetary policy announcement dates from the respective central bank's website. We supplement these sources with records from `Investing.com` and verify that the latter coincide with the official central bank calendar when the two samples overlap. Most of the monetary policy announcement samples start in 2000, with the latest starting date being September 16, 2004, for the Swiss National Bank. As already highighted by Albagli et al.

(2019), the non-U.S. monetary policy announcement dates seldom coincide with FOMC monetary policy announcements.

3 Monetary policy announcements and long-term interest rates

We examine the dynamics of 10-year government bond yields during different event windows. The sample period is divided into two parts: central bank monetary policy announcement windows (MP windows) and non-monetary policy announcement windows (non-MP windows). When the Federal Reserve is the central bank of concern, we refer to these as FOMC windows and non-FOMC windows. Following Hanson and Stein (2015) and Hillenbrand (2025), we define monetary policy (MP) 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. We cumulatively sum the daily yield changes of the 10-year government bond yield over each event window. Due to data availability, most of our analysis focuses on nominal yields to maintain a longer sample, while Section 6 is dedicated to real yields using inflation-protected bonds.

Formally, the cumulative change series is defined 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 denote daily dates, t_0 is the first date of the sample, y_s is the n-year Treasury zero coupon yield on date s, $\mathbf{1}_{s \in W}$ is an indicator function for the set W, and $W \in \{MP, nonMP\}$ is either the set of monetary policy announcement window dates or remaining dates outside of central bank monetary policy announcement windows. Since the two event windows are disjoint and span the full sample, for each time t, the total change in the yield relative to the initial value equals the cumulative sum of yield changes over the MP windows plus the cumulative sum of yield changes over the non-MP windows:

$$y_t - y_0 = \nabla y_t^{AllDates} = \nabla y_t^{MP} + \nabla y_t^{nonMP} \tag{2}$$

In the rest of the paper, we subtract the initial values from the observed yields, so that $\nabla y_t^{AllDates}$ and y_t can be used interchangeably.

3.1 FOMC Announcements

In Figure 2, we plot the cumulative sum of daily changes in the 10-year government yields during FOMC windows for the G10 currency countries. For each country, the blue line represents the observed cumulative yield change up to time t, the red line indicates the cumulative yield change within FOMC windows, and the black line reflects the cumulative yield change during non-FOMC windows. At any time t, the cumulative changes within FOMC and non-FOMC windows together equal the observed cumulative yield change (red line + black line = blue line).

The blue lines depict the cumulative changes in observed yields relative to the initial value, showing steady declines in 10-year government bond yields over the last three decades. The average decline across countries is approximately 4.7 percentage points.These trends align with a substantial body of literature documenting the secular decline of long-term interest rates and the natural interest rate, r^*.

A striking pattern emerges from the decomposition of the total yield change into FOMC window and non-FOMC window changes. As noted by Hillenbrand (2025), the decline in the 10-year U.S. Treasury yield is primarily explained by the changes within the FOMC windows. Both the overall yield decline and that within FOMC windows are around negative five percentage points. More interestingly, the charts reveal a new result, namely that the pattern extends to the 10-year government bond yields of other countries. A significant portion of the decline in 10-year government bond rates across the G10 currency countries is driven by the declines occurring within the FOMC windows. For instance, the decline in observed yields and the decline within FOMC windows for Canadian government bonds are both exactly 6.3 percentage points by the end of the sample. The strong association between the actual yield dynamics and those within FOMC windows is particularly striking given that FOMC meetings occur only eight times a year, meaning the cumulative yield adjustments during FOMC windows only happen on 24 days per year — less than 7% of the yearly trading days.

We evaluate how well the yield dynamics during the FOMC window fit the overall yield dynamics using a pseudo R^2 measure, which is reported in the subtitle of each chart panel. The measure is computed as one minus the ratio of the squared fitting error of the FOMC-window series to the sum of the squared fitting

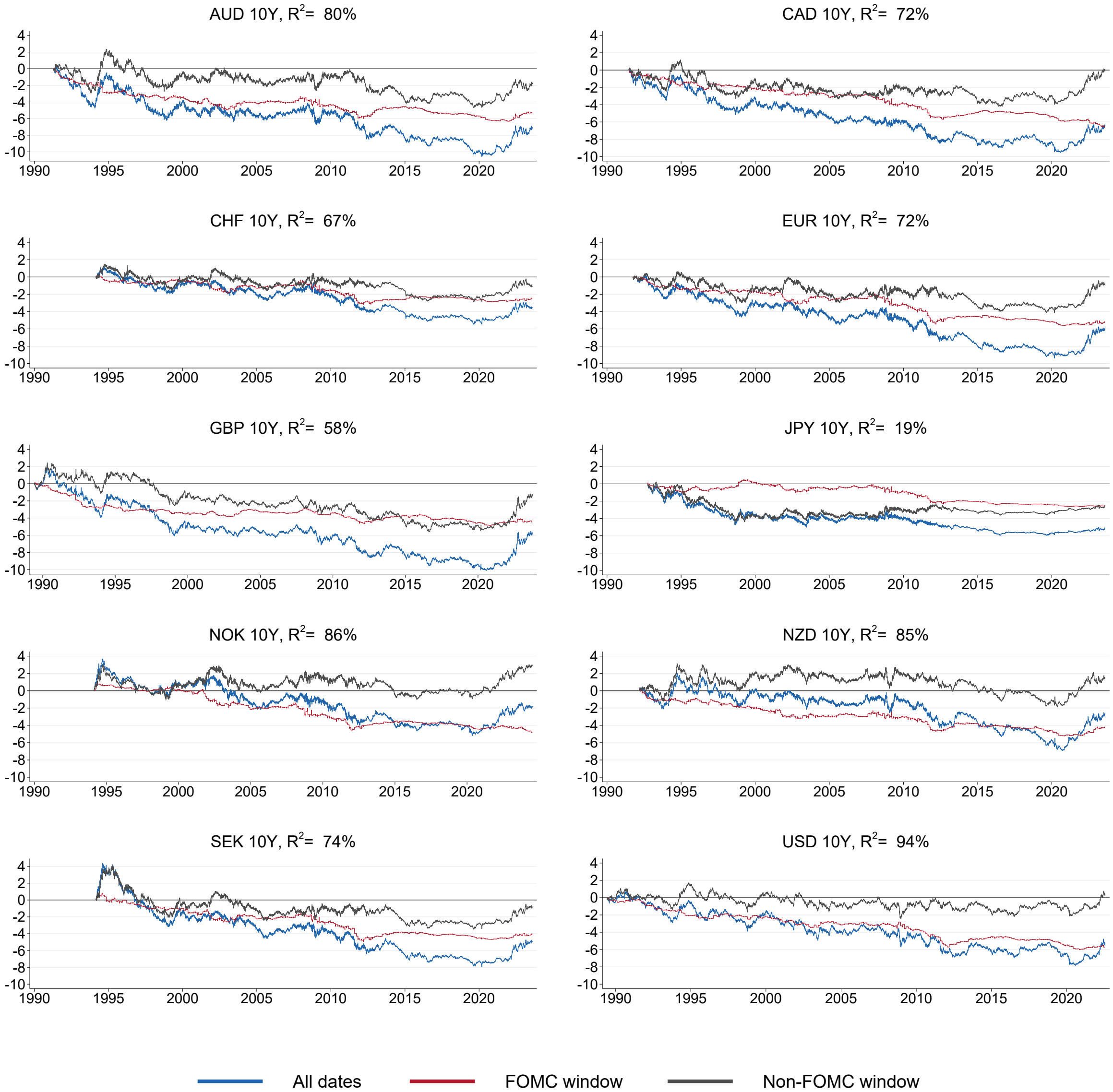

Figure 2: Cumulative changes in nominal 10-year yields during FOMC announcement windows.

Notes. Unit of the y-axis is percentage p.a.. The cumulative change in the 10-year yield y_t during event window W is $\nabla y_t^W = \sum_{s=t_0+1}^{t}(y_s - y_{s-1})\mathbf{1}_{s\in W}$. The FOMC window consists of $\{t-1, t, t+1\}$ if date t is an FOMC announcement date. The non-FOMC window complements the FOMC window. The pseudo R^2 is defined as $1 - \frac{\sum_{t=1}^{T}(y_t - \nabla y_t^{FOMC})^2}{\sum_{t=1}^{T}(y_t - \nabla y_t^{nonFOMC})^2 + \sum_{t=1}^{T}(y_t - \nabla y_t^{FOMC})^2}$.

errors of both the FOMC and non-FOMC window series:

$$pseudoR^2 \equiv 1 - \frac{\sum_{t=1}^{T}\left(y_t - \nabla y_t^{FOMC}\right)^2}{\sum_{t=1}^{T}\left(y_t - \nabla y_t^{nonFOMC}\right)^2 + \sum_{t=1}^{T}\left(y_t - \nabla y_t^{FOMC}\right)^2} \tag{3}$$

Intuitively, the pseudo R^2 can be interpreted analogous to R^2 and is bounded between zero and one. A larger pseudo R^2 indicates that greater fitting errors occur during non-FOMC windows, suggesting that the FOMC windows provide a better fit to the observed yields. Since the FOMC and non-FOMC windows cover the entire sample period, a higher pseudo R^2 also implies that the FOMC series fits the observed yield data better.

Overall, the FOMC series fits the actual yields quite well, with the best fit observed for the U.S. 10-year yield. The pseudo R^2 indicates that the FOMC windows explain 94% of the total variation in the U.S. 10-year yield over the sample period. Additionally, the FOMC windows exhibit very high explanatory power for yields in Australia, Canada, Switzerland, Germany, Norway, New Zealand, and Sweden, accounting for over two-thirds of the observed total variations. For Japan and the U.K, the fit is worse, at 19% and 58%, respectively.[3]

3.2 Other central banks' monetary policy announcements

Do world interest rates also persistently decline during non-U.S. central bank announcement windows? We repeat the exercise of the previous subsection but replace in Equation (2) the FOMC announcement dates with those of other central banks and then investigate their role in domestic and international interest rate trends.

To assess 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, we report in Table 2 the respective pseudo R^2 and root mean square error (RMSE).[4] For comparison, we also report results for the cumulative change during the FOMC windows and during the remaining days. The RMSE for the cumulative sum of daily

[3]In Appendix B, we demonstrate that the limited variation in the U.K. yield during FOMC windows can be substantially improved when the FOMC window is extended backward to three days before the announcement.

[4]In Appendix D, we plot the dynamics of world 10-year yields during each central bank's announcement windows analogous to Figure 2.

yield changes of country i within an event window W is defined as:

$$RMSE_i = \sqrt{\frac{1}{T}\sum_t \left(y_{i,t} - \nabla y_{i,t}^W\right)^2}. \tag{4}$$

A small RMSE indicates that the series is close to the actual yield.

In the upper panel of Table 2, we report the results for the full sample. The three rows of pseudo R^2 sum up to 100% for each column. The first row of the pseudo R^2 panel is analogous to the ones reported in Figure 2. Note that here we divide the non-FOMC window into the domestic monetary policy window and the days that are neither FOMC nor domestic monetary policy windows, so the denominator is not the same as that in Figure 2. Despite this difference, the results are similar, and the FOMC windows explain most of the variation in the observed yields. The contributions of the domestic monetary policy announcements to the cumulative yield changes are generally much smaller than the FOMC windows. On average, the fractions of G10 currency yield variations explained by the FOMC windows, domestic central bank windows, and non-monetary policy announcement windows are 64%, 5%, and 31%, respectively. Similarly, the FOMC windows generally generate a smaller RMSE than the domestic monetary policy windows, indicating that the FOMC announcements have performed better in explaining the yield dynamics. The exception is Japan, where the domestic monetary policy announcement windows play a more important role in explaining the yield dynamics than FOMC announcements.

In the lower panel of Table 2, we present results for the sample that begins when data on domestic central bank monetary policy announcements become available. This approach prevents the underestimation of the domestic central bank's contribution due to a smaller number of event windows. For example, ECB announcement data are only available from 1999 when the ECB was established. Overall, the pseudo R^2 and RMSE values remain very similar to those reported in the upper panel.

Figure 3 illustrates the cumulative daily changes in the 10-year yield during FOMC windows and domestic central bank windows, upon which the summary results in Table 2 are based. The series of cumulative changes begins when data on domestic monetary policy become available. The blue line represents the total cumulative change over the sample period, the red line indicates the change during the FOMC window, and the green line reflects changes during the domestic monetary policy window. Consistent with the results shown in Figure 2, the secular decline in the 10-year yield across countries is generally associated with a

Table 2: Fitting errors of the cumulative announcement effects of central banks.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
					Full Sample				
	Pseudo R^2 (%)								
FOMC	69	78	81	69	62	20	82	76	68
Domestic	25	5	8	3	2	44	1	11	2
Other dates	6	17	12	29	36	36	17	13	30
	RMSE (% p.a.)								
FOMC	2.00	2.19	1.13	2.16	2.76	3.12	1.25	1.35	1.79
Domestic	3.72	3.31	1.60	4.49	3.59	2.84	2.32	2.76	4.01
Other dates	7.33	4.82	2.62	3.53	3.64	3.80	3.35	2.26	2.90
					Sample starting with the first domestic MP announcement				
Sample start	90/1/21	00/2/3	00/1/20	99/1/7	97/6/6	99/1/18	86/1/12	99/3/6	99/1/4
	Pseudo R^2 (%)								
FOMC	69	79	82	69	60	22	82	77	70
Domestic	25	5	8	3	2	48	1	11	2
Other dates	6	16	10	28	38	30	17	11	27
	RMSE (% p.a.)								
FOMC	2.00	2.43	1.34	2.39	3.10	3.31	1.25	1.38	1.78
Domestic	3.72	3.78	1.95	5.07	4.06	2.98	2.32	3.05	4.30
Other dates	7.33	5.51	3.20	3.94	3.92	4.25	3.35	2.48	3.16

Notes. The table reports the pseudo R^2 and RMSE for each event window to explain the total variation in the observed yield. $RMSE = \sqrt{\frac{1}{T}\sum_t \left(y_t - \nabla y_t^W\right)^2}$, where y_t is the actual level of the yield, and ∇y_t^W is the cumulative changes in y_t during three-day windows bracketing the announcements of FOMC or the domestic central bank, or the other days.

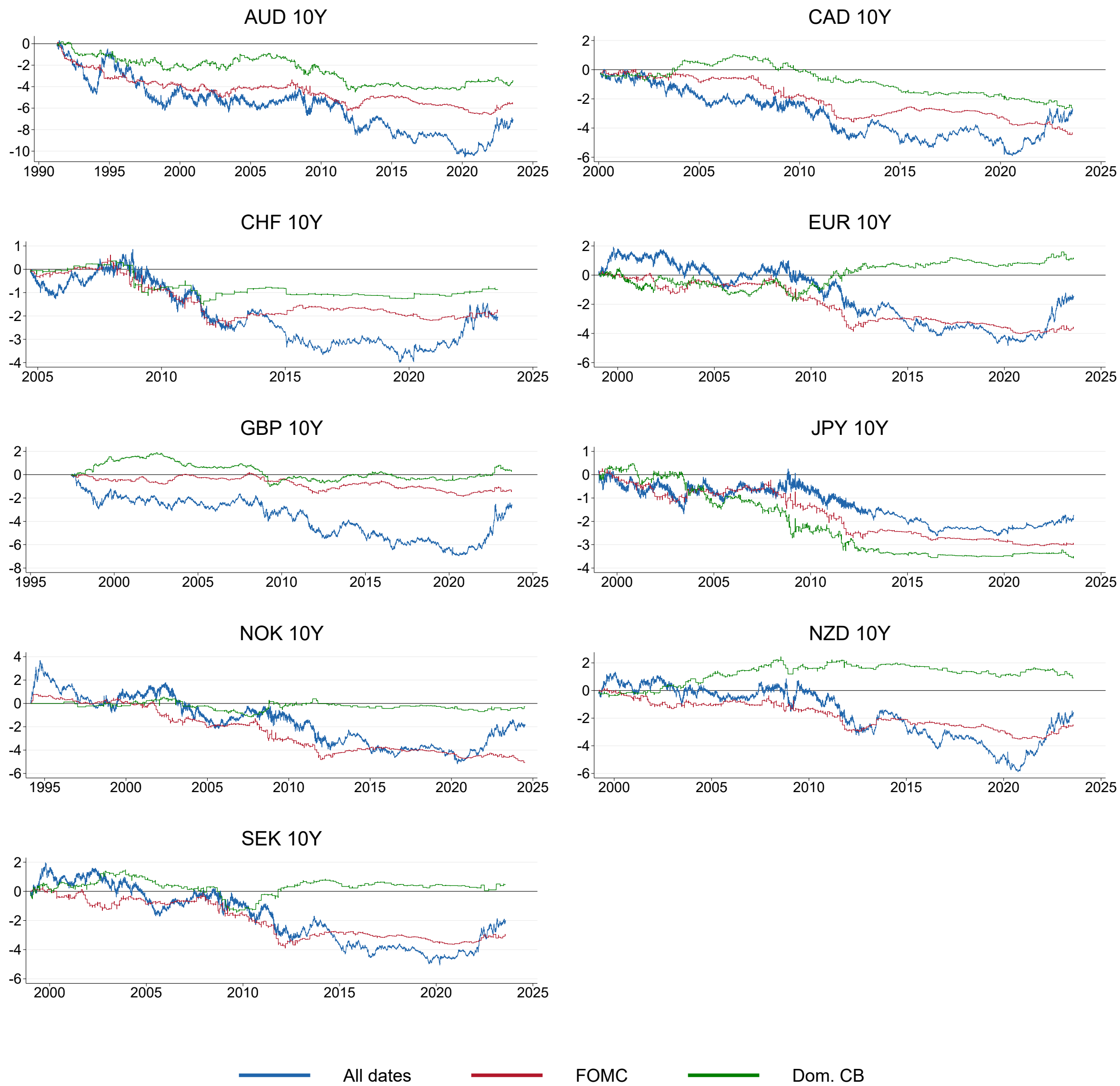

Figure 3: Cumulative changes in nominal 10-year yields during FOMC or domestic CB announcement windows.

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.

significant downward drift during FOMC windows, contributing to the relatively higher pseudo R^2 and lower RMSE associated with these windows. Interestingly, we do not observe a systematic downward drift around other central bank monetary announcement dates. The exception is Japan, where we find a substantial negative cumulative change during Bank of Japan monetary policy announcements, which exceeds that observed during FOMC windows. There are also downward movements in cumulative yield changes around the Reserve Bank of Australia, Bank of Canada, and Swiss National Bank announcement dates, but these are smaller than those associated with FOMC dates. The cumulative yield change remains nearly flat during the announcement windows of the Bank of England, Norges Bank, and Sveriges Riksbank. Positive cumulative yield changes are registered during the European Central Bank and Reserve Bank of New Zealand announcement windows.

As a more formal test of the link between overall yield changes and yield changes during monetary policy announcement windows, we estimate the following panel regression:

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

where $\Delta y_{i,t}$ is the change in the 10-year yield in country i over the FOMC window, and $\mathbf{1}_{c,t}$ is a dummy indicating whether t belongs to the 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. The equation is estimated using the 10-year yields over the period after 2000. We control for country-fixed effects (α_i) and the standard errors are clustered at the event level.

The results reported in Table 3 confirm that only FOMC announcements have a significant negative effect on cumulative yield changes across countries, with an additional 0.87 basis point drop relative to non-FOMC window days. For all other central bank announcements, the effect is essentially insignificant. When testing whether all central banks except the Fed have zero effects on cross-country yields, the null hypothesis is not rejected (F=1.05 with a p-value of 0.39).

In sum, a large share of the trend decline in global long-term interest rates occurred in the FOMC window, i.e. in the days around U.S. monetary policy announcements. By contrast, monetary policy announcements of other central banks did not exert any similar effects on domestic or other countries' long-term interest rates.

Table 3: Average yield changes during CB windows and non-CB windows.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
β	-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 of all non-US central bank coefficients are jointly zero: $F = 1.05$, $p = 0.39$.									

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The table reports the results from the panel regression $\Delta y_{i,t} = \alpha + \sum_c \beta_c \mathbf{1}_{c,t} + u_{i,t}$, where $\Delta y_{i,t}$ is the 2-day change in the 10-year yield in country i, 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 $\text{se}(\hat{\beta}_c)$. The last row reports the F-statistic and p-value for all non-U.S. central banks jointly have zero coefficients. The standard errors are clustered at the event level. The units are basis points per annum.

What explains the dominant role of the FOMC window in global long-term interest rate trends? Mechanically, it results from large and significant interest rate spillovers from the U.S. to other countries (see, e.g. Obstfeld (2015), Hofmann and Takats (2015), Albagli et al. (2019), Kearns et al. (2023)). Through these spillover effects, the cumulative negative effect of FOMC announcements on U.S. long-term bond yields transmits globally.

To substantiate this point, we assess the link between U.S. and other countries' long-term interest rates, both only during the FOMC windows and over the full sample. We conduct a panel fixed effect regression of the change of daily 10-year yield change of a foreign country on the change in US bond yields:

$$\Delta y_{i,t} = \alpha_i + \beta \Delta y_{US,t} + u_{i,t}, \tag{6}$$

The results reported in Table 4 indicate sizable and statistically significant spillover effects. We find that when the U.S. bond yields decrease by one percentage point during FOMC meetings, foreign bond yields decline, on average, by around 0.9 percentage points.

While these considerations help explain the persistent effects of FOMC announcement effects on global yields, they leave open the economic forces driving the results. In the following sections, we zoom in on the FOMC announcement effects on global yields, trying to uncover the underlying mechanisms by looking at long-term interest rate dynamics from different perspectives.

Table 4: U.S. spillovers in long-term bond yields.

	FOMC window regression	Full sample regression
Δy_{US}	0.886***	0.902***
	(0.055)	(0.02)
R^2	0.335	0.322
adj. R^2	0.336	0.322
within R^2	0.336	0.322
N	6,447	64,269

*p<0.1; **p<0.05; ***p<0.01.
Standard errors in parentheses are clustered by date.

4 Monetary policy vs information effects

Hillenbrand (2025) provides evidence suggesting that U.S. long-term yields decline when the Federal Reserve cuts interest rates and lowers its long-run forecasts for the Federal funds rate, as released through the "dot plots" in the Survey of Economic Projections (SEP). This finding suggests that the secular decline in long-term interest rates was driven by an information channel rather than by monetary policy itself. However, it is important to note that the dot plots have only been published since 2012, by which time most of the reductions in long-term rates linked to FOMC announcements had already occurred.

In this context, we seek to clarify the relative roles of information effects versus pure monetary policy effects in the cumulative impact of FOMC announcements on global bond yields. We conduct two exercises. First, we examine whether "purified" high-frequency U.S. monetary shocks can explain ∇y_t^{FOMC}. We utilize high-frequency monetary policy shocks from the recent literature that are designed to capture unexpected monetary policy changes. Second, we analyze a Fed Chair speech window, interpreting it as a monetary information shock. Fed Chair speeches do not coincide with FOMC announcements but provide insights into future policy actions, allowing us to identify the effects of the information channel while holding FOMC actions constant. If interest rate trends arise from global financial markets learning about macroeconomic fundamentals and the Fed's policy inclinations through these speeches, we would expect to observe similar patterns in world interest rates during both Fed Chair speech windows and FOMC announcement windows. However, in both exercises, we do not find a significant role for the information channel in shaping the long-term behavior of world interest rates.

4.1 Purified monetary policy shocks vs information shocks

High-frequency identified monetary shocks (MPS) have become a standard tool in the empirical macro literature since Kuttner (2001) and Gürkaynak et al. (2005). Papers such as Nakamura and Steinsson (2018) and Bauer and Swanson (2023a) suggest there are non-monetary shock components embedded in the shock, such as the “Fed information effect” and the “Fed response to news effect”. We first examine how well the raw high-frequency changes in interest rate futures can explain the interest rate dynamics during FOMC windows. Then, we examine the relationship between the FOMC-window interest rate trend and decomposed monetary shocks developed by Jarociński and Karadi (2020).

We summarize the high-frequency changes in interest rate futures using the MPS measure from Bauer and Swanson (2023a), which is defined as the first principal component of 30-minute changes in the first four Eurodollar futures around FOMC announcements. We estimate the following equation:

$$\nabla y_t^{FOMC} = \gamma_0 + \gamma_1 CMPS_t + u_t, \tag{7}$$

where ∇y_t^{FOMC} are the cumulative daily changes in a country's 10-year yield over FOMC windows, and $CMPS_t = \sum_{s=0}^{t} MPS_s$ is the cumulative sum of historical MPS up to t. In Figure 4, we report both the fitted value and ∇y_t^{FOMC}. The fitted values are almost indistinguishable from ∇y_t^{FOMC}, suggesting that the cumulative U.S. monetary policy shocks can indeed explain the cumulative changes in world sovereign yields during FOMC windows.

The literature has documented that the high-frequency changes in the interest rate futures around FOMC announcements do not purely reflect the unexpected changes in the monetary policy stance. They also incorporate information effects reflecting inference about the central banks assessment of the economic outlook embedded in the monetary policy decision. Several recent papers propose methods to “purify” the high-frequency shocks and estimate the “unexpected change in the monetary policy stance”. We utilize the purified monetary policy shocks from Jarociński and Karadi (2020), which decomposes the first principal component of ED1-ED4 shocks into a pure monetary policy shock (JK MP) and a central bank information shock (JK CBI). The former captures the unexpected changes in the U.S. monetary policy stance, while the latter reflects the market's inference about the Fed's information regarding macroeconomic conditions and outlook.

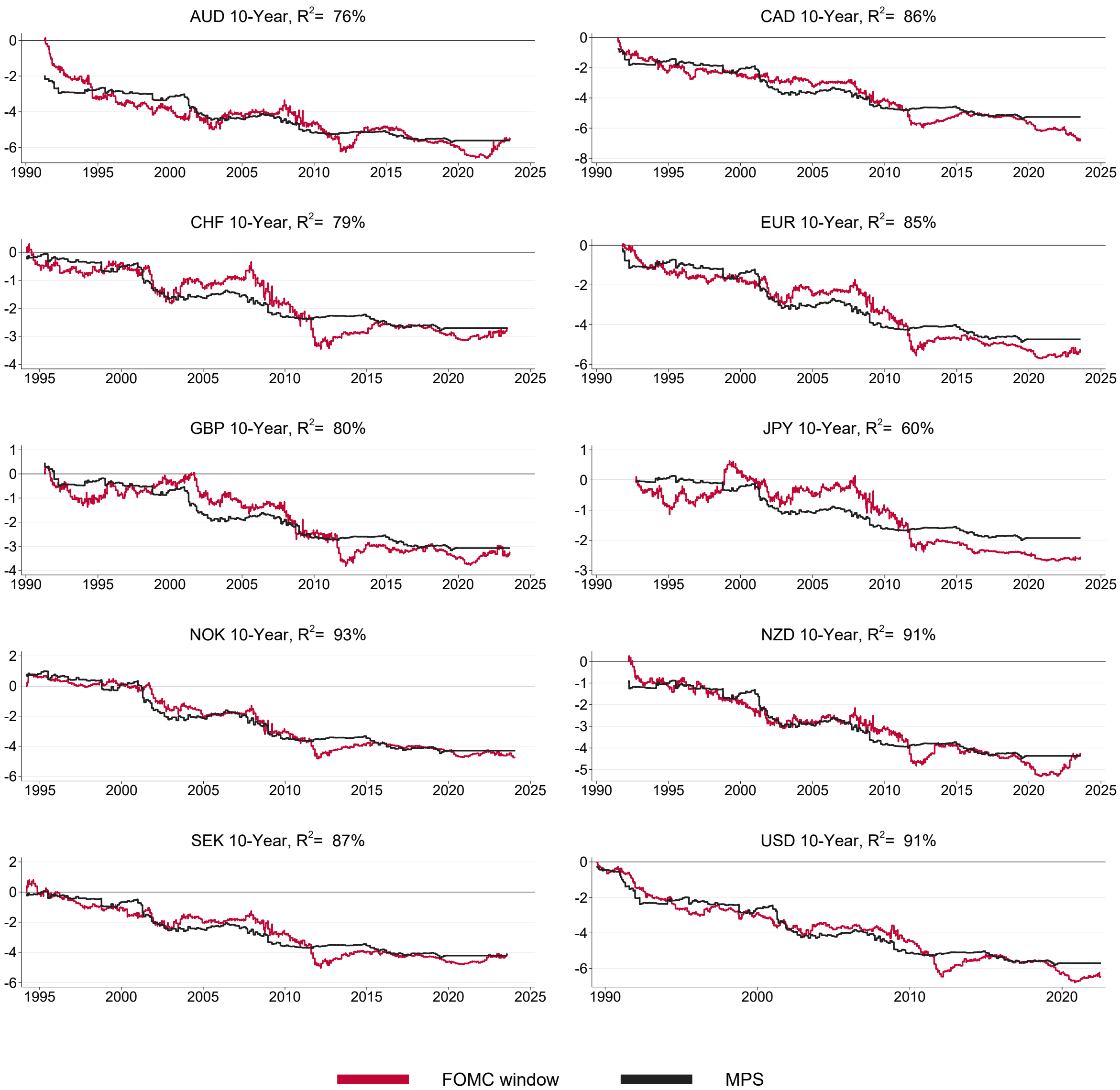

Figure 4: Cumulative yield changes and cumulative U.S. HF monetary policy shocks.

Notes. Unit of the y-axis is percentage p.a.. The figure presents ∇y_t^{FOMC} and the fitted value from the cointegration regression $\nabla y_t^{FOMC} = \gamma_0 + \gamma_1 CMPS_t + u_t$, where MPS_t is the first principal component of 30-min changes in ED1-ED4 around FOMC announcements, and $CMPS_t = \sum_{s=0}^{t} MPS_s$ is the sum of historical shocks up to t.

By substituting the cumulative sum of purified monetary policy shocks and of central bank information shocks respectively for Equation (7), we compute the fitted values and report them in Figure 5. Both shocks feature cumulative downward trends and track the FOMC window trend reasonably well. For each panel, we report in the panel title the semi-partial correlation squared, which corresponds to the marginal R^2 gain when either the pure monetary policy shock or the central bank information shock is added to the model while controlling for the other. The pure monetary policy shock contributes a substantial R^2 gain way beyond that of the central bank information shock. Conversely, the latter offers little additional explanatory power when the pure monetary shock is already included. This suggests that the pure monetary policy shock accounts for a larger portion of the variation observed in FOMC windows compared to the central bank information component. In Appendix D, we demonstrate that the cumulative monetary policy shocks from Acosta (2023) and Bu et al. (2021) produce similar results.

In sum, the analysis suggests that the dynamics of world interest rates during FOMC announcement windows are mostly driven by unexpected changes in the U.S. monetary policy stance, i.e. pure monetary policy shocks. The market's inference about the Federal Reserve's information does not seem to explain why world interest rates persistently decline during FOMC windows.

4.2 Fed Chair speech windows

Swanson and Jayawickrema (2023) provide evidence suggesting that the volatility of financial variables, especially long-maturity yields, is often higher during Fed Chair speech dates than during the FOMC meeting announcements. They argue that Fed Chair speeches are more important than FOMC announcements for monetary policy transmission. We interpret Fed Chair speeches as a relatively pure information channel since there is no policy action taken during the speeches. Their cumulative impact on bond yields hence provides another test of the relevance of the information effect in global yield trends.

We assess this point by cumulating the daily yield changes during Fed Chair speech windows and non-Fed Chair speech windows, comparing them with the actual yield dynamics. [5] Figure 6 plots the cumulative sum of daily yield changes over the Fed Chair speech windows. There are 412 Fed Chair speech dates in our sample, whereas the number of FOMC announcements during the same period is 226. Despite the larger

[5] Depending on the country sample, the 10-year bond yield series start from 1991 to 1994 (see Table 1), but we cumulate the Fed Chair window only from 1996 due to data availability.

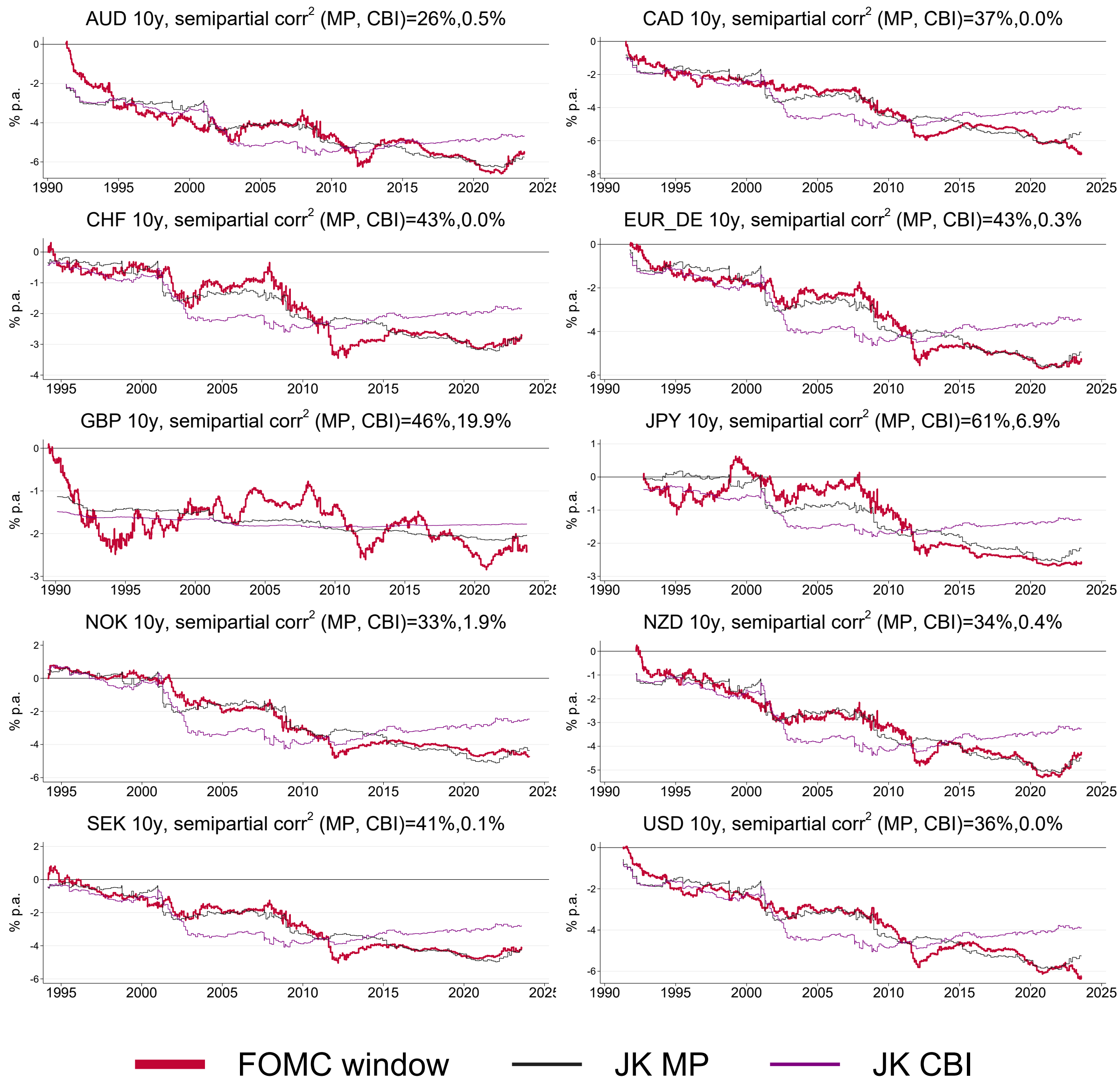

Figure 5: Cumulative yield changes and cumulative purified U.S. HF monetary policy shocks.

Notes. Unit of the y-axis is percentage p.a.. The figure presents ∇y_t^{FOMC}, the fitted value from the cointegration regression $\nabla y_t^{FOMC} = \gamma_0 + \gamma_1 CMPS_t + u_t$. 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). The subtitle of each figure reports the semipartial correlation squared of the two variables.

number of Chair speeches, they display a much weaker association with overall yield dynamics than FOMC announcements. For all countries, the cumulative yield declines over the Fed Chair window is on average close to zero. Moreover, the cumulative yield changes during these windows account for minimal fractions of the total variations in actual 10-year yields, as measured by the pseudo R^2. For example, the Fed Chair speech windows contribute less than 2% to the total variations of the 10-year yields in Australia, Switzerland, Germany, the U.K., Sweden, and the U.S. Therefore, the majority of the yield decline over the sample period is explained by non-Chair speech windows, which include FOMC windows.

The weak relationship between Fed Chair speeches and the secular decline of government bond yields does not imply that these speeches have no impact. 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 5, 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. In all countries, government bond yields are more volatile during the FOMC window than on normal days. Moreover, in eight out of ten countries yield changes are more volatile during the Fed Chair speech window than during other days. Overall, the findings of Swanson and Jayawickrema (2023) for the U.S. can be generalized to G10 currency 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.

Interestingly, Swanson and Jayawickrema (2023) demonstrate that FOMC announcements have larger impacts on short-maturity interest rates than Chair speeches while the latter are more important for long-maturity interest rates. Illustrating that the 10-year yields in various countries have declined much more during FOMC announcement than the Fed Chair speech windows, Figure 6 seems to contradict this finding. However, the reason for the contradiction is that we study the cumulative sum of actual daily yield changes during the event windows instead of the average absolute changes or standard deviation of daily changes. Although interest rates change drastically during the Fed Chair speeches, as shown by Table 5 and Swanson and Jayawickrema (2023), the directions of those changes are more evenly distributed across positive and negative values, so they cancel each other out in the summation.

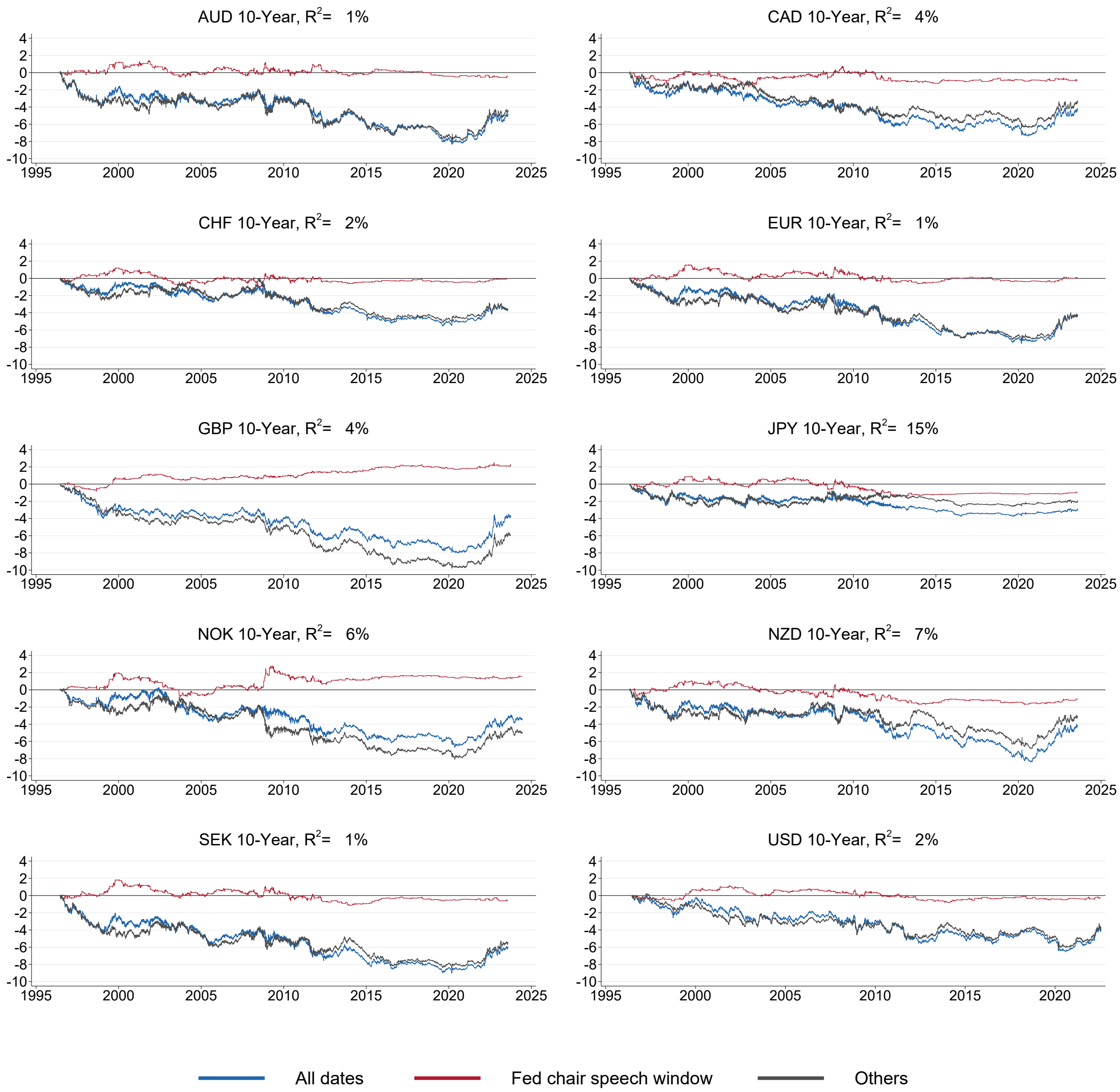

Figure 6: Cumulative 10-year yield changes during Fed Chair speech windows.

Notes. Unit of the y-axis is percentage p.a.. A Fed Chair speech window is from the day before the speech to the day after the speech. "Others" refers to all other dates (including FOMC announcement windows). The pseudo R^2 is defined as $R^2 = 1 - \frac{\sum_{t=1}^{T}\left(y_t - \nabla y_t^{chair}\right)^2}{\sum_{t=1}^{T}\left(y_t - \nabla y_t^{non-chair}\right)^2 + \sum_{t=1}^{T}\left(y_t - \nabla y_t^{chair}\right)^2}$.

Table 5: Standard deviation of daily yield changes.

	3-month			5-year			10-year			5-5 forward		
	FOMC	Chair	Other	FOMC	Chair	Other	FOMC	Chair	Other	FOMC	Chair	Other
AUD	7.32	6.61	7.40	11.43	10.74	10.09	11.11	10.71	10.05	12.72	11.92	11.61
CAD	7.36	7.06	6.78	10.77	9.46	9.20	9.82	9.31	8.71	10.71	10.18	9.60
CHF	7.14	6.21	5.61	9.94	9.41	8.55	9.70	9.26	8.57	11.03	10.39	9.94
EUR	5.58	4.45	5.39	11.24	10.55	9.49	10.66	10.04	9.22	11.74	10.70	10.24
GBP	7.05	5.83	7.67	8.08	6.24	7.60	7.52	5.84	7.25	10.95	7.69	10.78
JPY	4.44	3.79	4.11	9.75	8.59	8.10	9.19	8.36	7.93	10.48	9.30	9.24
NOK	6.43	5.95	5.78	10.33	10.26	9.55	10.03	10.37	9.91	11.61	13.54	13.52
NZD	7.27	6.34	7.37	11.25	9.68	9.53	10.05	9.34	9.08	13.70	12.06	13.34
SEK	6.39	5.74	5.82	10.92	10.04	9.63	10.76	9.73	9.58	12.41	10.61	10.95
USD	6.59	4.12	4.26	6.84	6.18	5.85	6.56	6.06	5.72	7.41	6.70	6.36

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.

5 Interest rate expectations and term premia

Hanson and Stein (2015) found that monetary policy has a surprisingly large effect on far-ahead real forward rates through term premia, reflecting a reaching-for-yield channel. In this section we assess whether our findings can be explained by this channel based on decompositions of long-term rates into term premia and expected future short-term interest rates. We investigate which of these two are the main drivers of the decline during the FOMC windows and show that the majority of the decline is associated with changes in expected future short-term rates.

5.1 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.[6]

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

[6] 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.

consisting of a trend vector $\boldsymbol{\tau}_t$ and a stationary vector $\tilde{X}_t$:

$$
\begin{aligned}
X_t =& \boldsymbol{\mu} + \Gamma \boldsymbol{\tau}_t + \tilde{X}_t, \\
\boldsymbol{\tau}_t =& \boldsymbol{\tau}_{t-1} + \boldsymbol{\eta}_t, \quad \boldsymbol{\eta}_t \sim \mathcal{N}(0, \Omega_\eta) \\
\tilde{X}_t =& \Phi \tilde{X}_{t-1} + \tilde{U}_t, \quad \tilde{U}_t \sim \mathcal{N}(0, \tilde{\Omega}),
\end{aligned} \tag{8}
$$

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

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

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

$$
m_{t+1} = -\delta_0 - \boldsymbol{\delta}_1^\top X_t - \frac{1}{2} \Lambda_t^\top \Lambda_t - \Lambda_t^\top \Omega^{-\frac{1}{2}} U_{t+1}. \tag{9}
$$

The price of risk is an affine function of Z_t:[7]

$$
\Lambda_t = \Omega^{-\frac{1}{2}} (\Lambda_0 + \Lambda_1 Z_t). \tag{10}
$$

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:

$$
y_t^{(n)} = A_n + B_n^\top X_t, \tag{11}
$$

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

[7] 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}.

prices solve

$$p_t^{(n),rn} = \ln \mathbf{E}_t \left[\exp \left\{ -y_t^{(1)} + p_{t+1}^{(n-1),rn} \right\} \right], \tag{12}$$

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

$$p_t^{(n),rn} = \mathcal{A}_n^{rn} + \mathcal{B}_n^{rn\top} X_t + \mathcal{C}_n^{rn\top} \boldsymbol{\tau}_t, \tag{13}$$

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

5.2 Model estimation

Trend Proxy. We assume that $\boldsymbol{\tau}_t$ is observable, which is referred to as the "observed shifting endpoint" (OSE) model in Bauer and Rudebusch (2020). Our empirical proxy for $\boldsymbol{\tau}_t$ is the first principal component of cumulative changes in the yield curve during FOMC windows.[8] We justify the choice with the following cointegration analysis.

Table A10 has established that the cumulative changes in yields during the Fed's monetary policy announcement windows are the best fit for the long-run trends of interest rates worldwide. This is a restricted version of the cointegration regression

$$y_t = \gamma_0 + \gamma_1 \tau_t + u_t \tag{14}$$

with constraints $\gamma_0 = 0, \gamma_1 = 1$. How well can the cumulative FOMC effects account for world interest rate trends in an unconstrained regression? According to the Beveridge and Nelson (1981) decomposition, an interest rate time series can be written as a random walk plus a stationary series, and the random walk is interpreted as the "trend". In this regard, we test whether u_t is stationary when τ_t is the cumulative yield change during FOMC windows.

We estimate γ_0 and γ_1 using the dynamic OLS estimator of Stock and Watson (1993). The data are end-of-month observations of each series. Then, we estimate $\hat{u}_t = y_t - \hat{\gamma}_0 - \hat{\gamma}_1 \tau_t$. We apply three stationarity

[8] The time series of cumulative changes during FOMC windows are almost parallel across maturities, so the first principal component is similar to a simple cross-sectional average and explains close to 100% of the total variations. Therefore, we use the first principal component of the FOMC-window series to represent the common trend of the yield curve.

tests to $\hat{u}_t$: the Augmented Dickey-Fuller (ADF), Phillips-Perron (PP), and Müller-Watson low-frequency (LFST) tests. We report the test statistics for ADF and PP and the p-value for LFST in Table 6. For each column, the dependent variable is the country's 10-year yield. In specification I, $\tau_t = \nabla PC1^{US}_{loc,t}$ is the first principal component of cumulative changes in the local yield curve during U.S. monetary policy announcement windows. For each country, either the ADF and PP test statistics significantly reject the unit root hypothesis, or the LFST p-value fails to reject the stationarity hypothesis. Therefore, the cumulative changes in local yields during FOMC windows appear to explain the trend of each country's 10-year yield well.[9] In specification II, $\tau_t = \nabla PC1^{US}_{US,t}$ is the first principal component of the cumulative changes in the U.S. yield curve during U.S. monetary policy announcement windows. Since it is common to all countries, we interpret $\nabla PC1^{US}_{US,t}$ as a global interest rate trend. For all countries, the ADF, PP, and LFST tests uniformly support the stationarity of u_t. Interestingly, the global trend $\nabla PC1^{US}_{US,t}$ outperforms the local trend $\nabla PC1^{US}_{loc,t}$ in explaining the interest rate trends of some countries, such as Switzerland, Japan, Norway, and Sweden. The dynamics of U.S. interest rates during FOMC announcement windows are highly informative about world interest rate trends. In specification III, we use $\boldsymbol{\tau}_t = [\nabla PC1^{US}_{loc,t}, \nabla PC1^{US}_{US,t}]^\top$ to explain world interest rate trends. The performance is better than the previous two specifications: the cointegration residual seems to be more stationary. Motivated by this fact, we use the bivariate specification of $\boldsymbol{\tau}_t$ for our term structure model.

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 $[\nabla PC1^{US}_{loc,t}, \nabla PC1^{US}_{US,t}]^\top$, 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., $\boldsymbol{\tau}_t = \nabla PC1^{US}_{US,t}$. 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 $\nabla PC1^{US}_{loc,t}$ and $\nabla PC1^{US}_{US,t}$ 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 C, 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

[9] In Appendix D, we report results using the first principal component of the local yield curve as the dependent variable.

Table 6: Cointegration tests: 10-year yields and cumulative yield changes during FOMC windows.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
I: $y_t = \gamma_0 + \gamma_1 \nabla PC1^{US}_{loc,t} + u_t$										
constant	-3.20	5.94	-1.64	-4.56	-7.74	-1.56	2.12	-4.63	-3.88	-2.01
	(0.86)	(0.23)	(0.29)	(0.36)	(0.71)	(0.26)	(0.41)	(0.76)	(0.44)	(0.22)
$\nabla PC1^{US}_{loc,t}$	1.42	0.09	1.37	1.36	1.63	0.96	1.63	2.01	1.41	1.36
	(0.14)	(0.01)	(0.09)	(0.06)	(0.10)	(0.10)	(0.34)	(0.15)	(0.09)	(0.05)
R^2	0.78	0.76	0.87	0.92	0.84	0.60	0.76	0.77	0.84	0.93
SD	1.07	1.48	0.74	0.79	1.04	0.86	1.18	0.94	1.19	0.50
$\hat{\rho}$	0.96	0.91	0.96	0.96	0.96	0.97	0.95	0.94	0.97	0.88
Half-life	15.9	7.0	18.5	15.3	17.2	20.4	14.8	11.1	24.9	5.4
ADF	-3.32**	-4.27***	-2.40	-2.66*	-2.72*	-2.54	-1.76	-3.26**	-2.29	-5.17***
PP	-17.57**	-32.27***	-10.89	-15.43**	-13.20*	-9.87	-6.47	-21.30***	-9.44	-50.19***
LFST	0.49	0.76	0.38	0.61	0.29	0.11	0.13	0.79	0.34	0.91
II: $y_t = \gamma_0 + \gamma_1 \nabla PC1^{US}_{US,t} + u_t$										
constant	-1.83	-3.63	-3.88	-4.78	-3.99	-2.35	-2.38	-0.58	-5.23	-2.01
	(0.52)	(0.37)	(0.43)	(0.46)	(0.39)	(0.37)	(0.80)	(0.56)	(0.64)	(0.22)
$\nabla PC1^{US}_{US,t}$	1.55	1.73	1.33	1.80	1.83	0.85	1.99	1.34	2.02	1.36
	(0.12)	(0.09)	(0.09)	(0.09)	(0.09)	(0.09)	(0.24)	(0.12)	(0.16)	(0.05)
R^2	0.80	0.89	0.85	0.89	0.89	0.75	0.82	0.75	0.82	0.93
SD	0.99	0.76	0.67	0.82	0.79	0.65	0.89	0.94	1.25	0.50
$\hat{\rho}$	0.96	0.93	0.96	0.96	0.94	0.95	0.93	0.95	0.97	0.88
Half-life	15.6	10.0	16.0	16.9	11.5	12.7	10.0	14.7	26.6	5.4
ADF	-3.22**	-3.46***	-2.59*	-2.78*	-3.26**	-2.73*	-2.62*	-2.93**	-2.89**	-5.17***
PP	-18.04**	-22.94***	-13.28*	-15.03**	-21.95***	-14.95**	-11.59*	-15.80**	-10.60	-50.19***
LFST	0.45	0.93	0.79	0.74	0.59	0.37	0.09	0.20	0.70	0.91
III: $y_t = \gamma_0 + \gamma_1 \nabla PC1^{US}_{loc,t} + \gamma_2 \nabla PC1^{US}_{US,t} + u_t$										
constant	-2.10	-3.82	-3.72	-4.69	-0.91	-2.75	-7.00	-3.02	-4.35	-2.01
	(0.75)	(1.20)	(0.51)	(0.35)	(0.93)	(0.26)	(1.06)	(1.05)	(0.72)	(0.22)
$\nabla PC1^{US}_{loc,t}$	0.34	-0.00	0.01	0.88	-1.16	-0.79	-1.79	1.12	0.85	1.36
	(0.37)	(0.01)	(0.31)	(0.34)	(0.34)	(0.19)	(0.40)	(0.48)	(0.43)	(0.05)
$\nabla PC1^{US}_{US,t}$	1.16	1.76	1.30	0.65	3.05	1.48	4.08	0.63	0.78	
	(0.39)	(0.23)	(0.29)	(0.47)	(0.38)	(0.15)	(0.47)	(0.35)	(0.65)	
R^2	0.81	0.90	0.92	0.93	0.92	0.87	0.93	0.79	0.87	0.93
SD	0.98	0.76	0.67	0.78	0.78	0.63	0.79	0.90	1.19	0.50
$\hat{\rho}$	0.96	0.93	0.96	0.96	0.93	0.93	0.89	0.94	0.98	0.88
Half-life	15.9	9.9	16.2	16.3	9.1	9.9	6.0	12.1	27.5	5.4
ADF	-3.18**	-3.46***	-2.59*	-2.75*	-3.98***	-2.75*	-3.65***	-3.08**	-3.01**	-5.17***
PP	-17.82**	-22.90***	-13.14*	-15.27**	-28.81***	-18.94**	-20.09**	-19.37**	-9.45	-50.19***
LFST	0.47	0.92	0.76	0.72	0.70	0.70	0.18	0.61	0.46	0.91

Notes. *p<0.1; **p<0.05; ***p<0.01. The table reports cointegration coefficients, Newey-West standard errors, and stationarity test results for the residual $\hat{u}_t$ for the cointegration regression $y_t = \gamma_0 + \boldsymbol{\gamma}_1^\top \boldsymbol{\tau}_t + u_t$. y_t denotes the 10-year yield of a given country. $\nabla PC1^{US}_{loc,t}, \nabla PC1^{US}_{US,t}$ denote the first principal component of cumulative changes in the local or U.S. yields during the U.S. central bank announcement windows. *SD*: the standard deviation of $\hat{u}_t$; $\hat{\rho}$: the AR(1) coefficient of $\hat{u}_t$; Half-life: the half-life of $\hat{u}_t$; ADF: the t-statistic for the augmented Dickey-Fuller test, and the null is that $\hat{u}_t$ contains a unit root; PP: the Phillips-Perron statistic, and the null is that $\hat{u}_t$ contains a unit root; LFST: the p-value for the Müller-Watson low-frequency stationarity test, and the null is that $\hat{u}_t$ is stationary.

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).

5.3 Decomposing ∇y_t^{FOMC}

We use our OSE model to decompose daily yields into the risk-neutral rates and term premia, and then compute the sum of their daily changes during the FOMC windows. This exercise estimates the contributions of the risk-neutral and term premium components to ∇y_t^{FOMC}, and Figure 7 presents the results for the 10-year yields. The series "all dates, yield" and "FOMC window, yield" replicate the observed yield and 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 and term premium during FOMC windows. The risk-neutral component explains most of the variations in the FOMC series for all countries except for Japan. The risk-neutral component for Japan has remained almost constant because the Japanese short rate has been stuck at zero during most of our sample period. On the other hand, the term premium component for all countries except for Japan exhibits limited variation during the FOMC windows and thus explains little variations in the cumulative responses of long-term sovereign yields 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

$$R^2 \equiv 1 - \frac{\sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC}\right)^2}{\sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{TP,FOMC}\right)^2 + \sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC}\right)^2}. \tag{15}$$

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 95% for Australia, Canada, Germany, and the U.S.

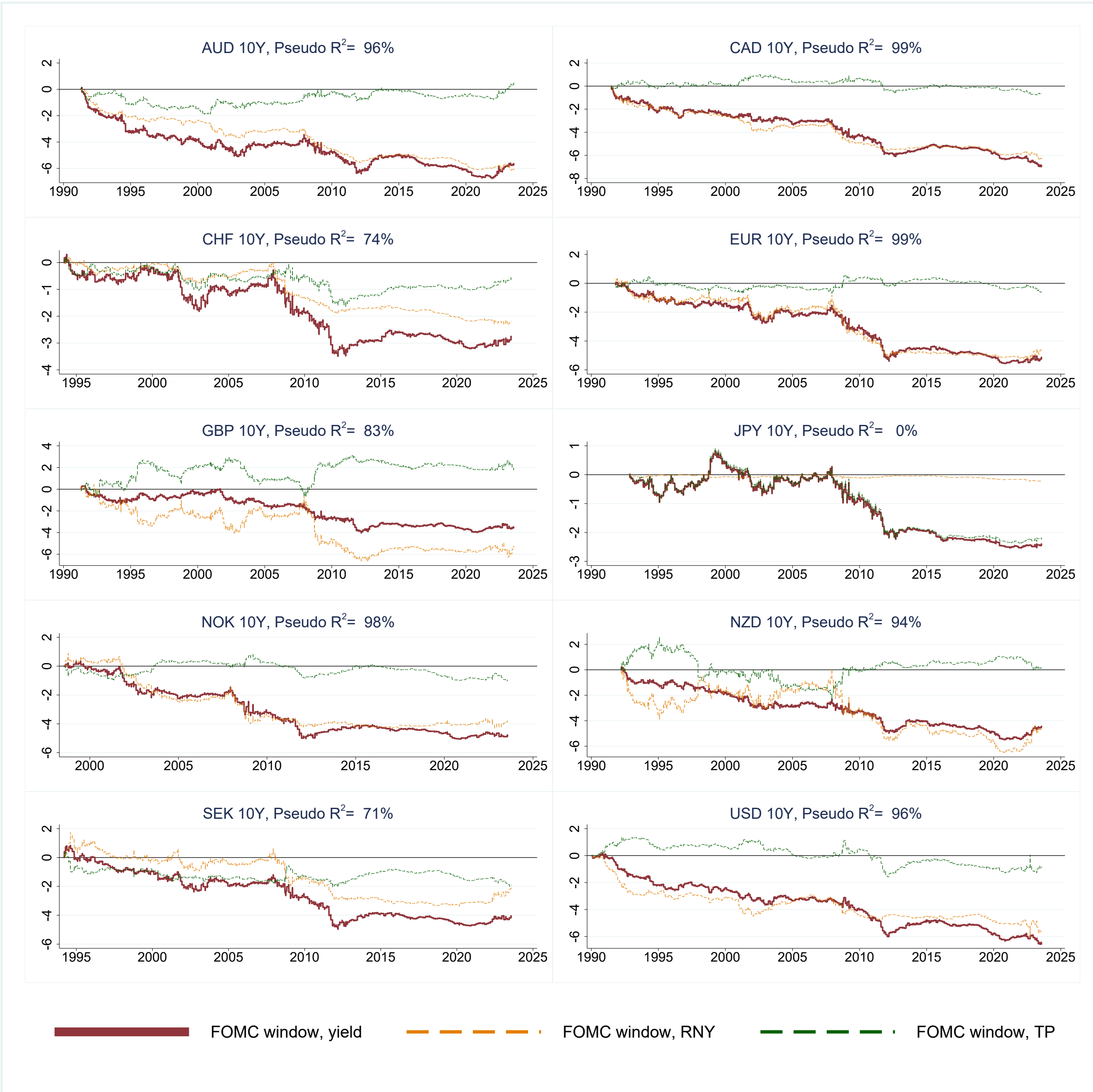

Figure 7: FOMC window: risk neutral yields (RNY) vs. term premia (TP).

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 is defined as $R^2 = 1 - \frac{\sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC}\right)^2}{\sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{TP,FOMC}\right)^2 + \sum_{t=1}^{T}\left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC}\right)^2}$.

5.4 Risk neutral yields and term premia in the FOMC window

Sovereign yield dynamics *during* FOMC windows can be mostly explained by changes in the risk-neutral yields and are nearly unrelated to term premium dynamics. On the other hand, how much do the risk-neutral yields and term premia change *between* FOMC meetings?

To understand the importance of FOMC windows for determining the dynamics of risk-neutral yields and term premia, we plot the sums of daily changes in the two components during FOMC windows and the complement dates.

Figure 8 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 yields, arguing that once the term structure model accounts for the persistent interest rate dynamics, the implied risk-neutral yields should 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 Equation (3), 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. The result might be mechanical because our trend proxy, which is a key component of the risk-neutral rate, only changes during the FOMC windows. In Appendix D, 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.

FOMC windows have much smaller impacts on the term premia than on the risk-neutral yields. Figure 9 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 8 shows that FOMC windows explain 89% of the total variation in the 10-year

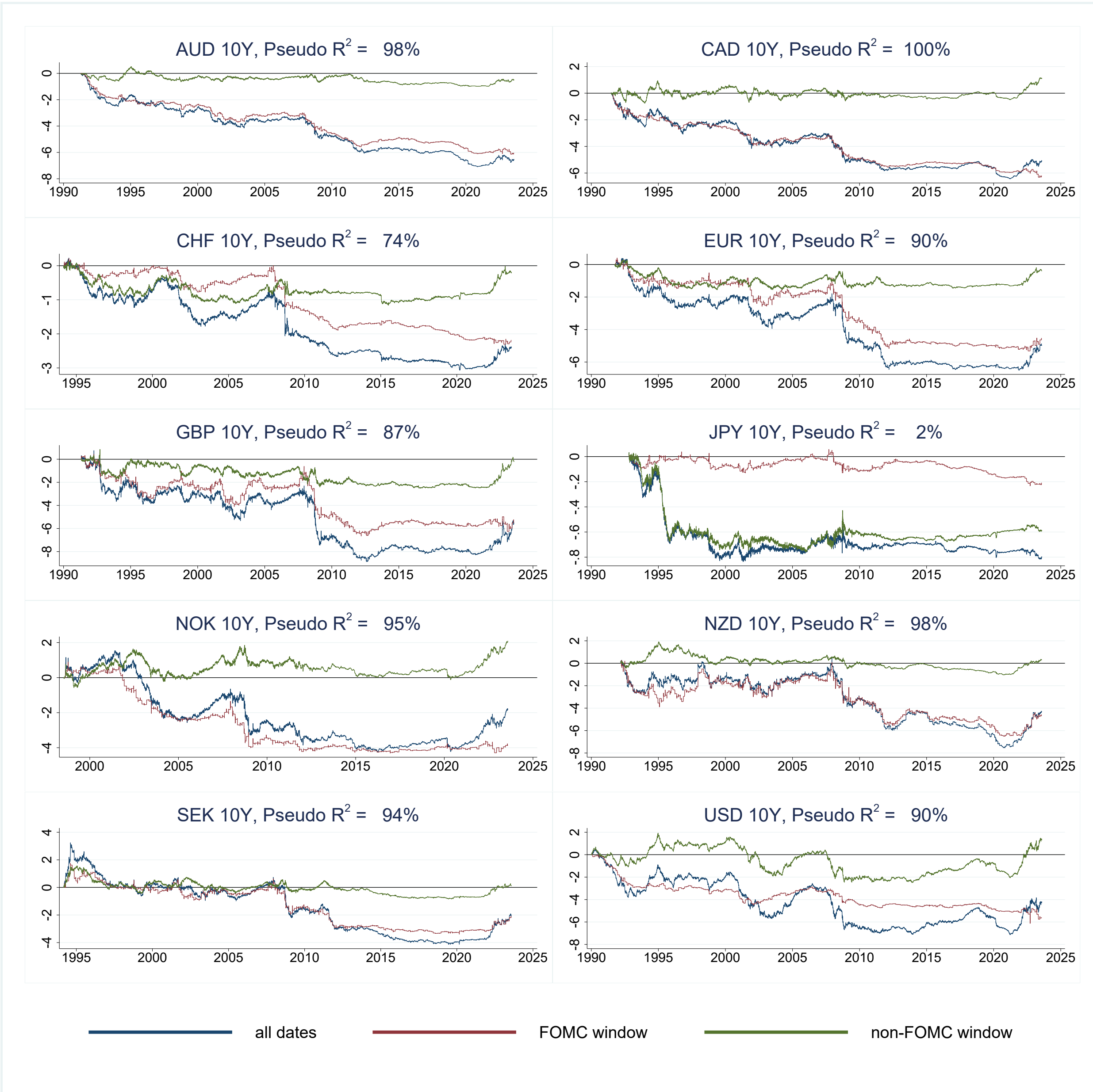

Figure 8: Cumulative changes in the 10-year risk-neutral rates during different event windows.

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 $R^2 = 1 - \frac{\sum_{t=1}^{T}\left(y_t^{RNY} - \nabla y_t^{RNY,FOMC}\right)^2}{\sum_{t=1}^{T}\left(y_t^{RNY} - \nabla y_t^{RNY,nonFOMC}\right)^2 + \sum_{t=1}^{T}\left(y_t^{RNY} - \nabla y_t^{RNY,FOMC}\right)^2}$.

risk-neutral yield.

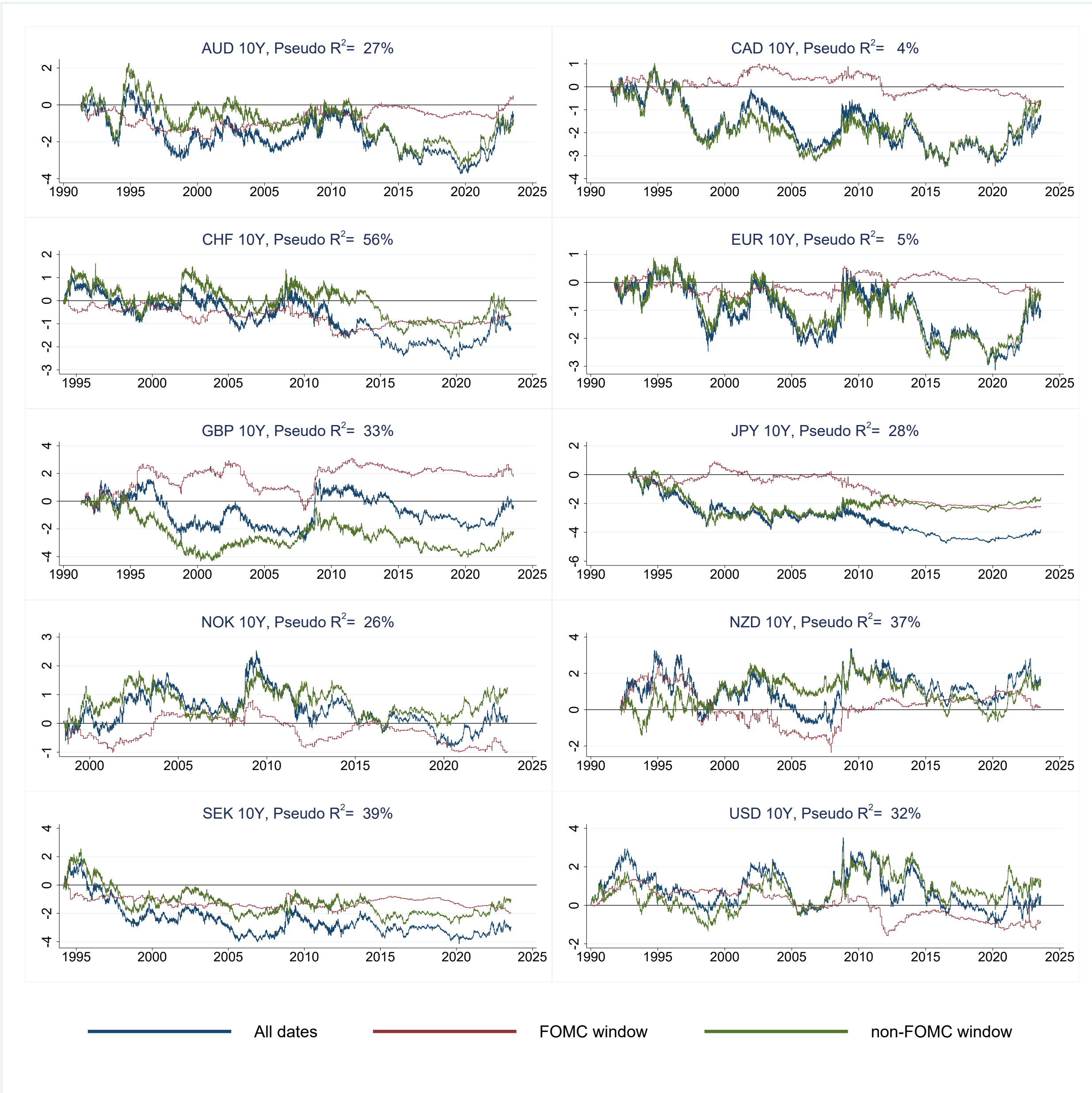

Figure 9: Cumulative changes in the 10-year term premium during FOMC windows.

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 $R^2 = 1 - \frac{\sum_{t=1}^{T}\left(y_t^{TP} - \nabla y_t^{TP,FOMC}\right)^2}{\sum_{t=1}^{T}\left(y_t^{TP} - \nabla y_t^{TP,nonFOMC}\right)^2 + \sum_{t=1}^{T}\left(y_t^{TP} - \nabla y_t^{TP,FOMC}\right)^2}$.

In summary, U.S. monetary policy significantly affects world interest rate trends, a feature not shared by the monetary policies of other countries. Furthermore, the responses of world long-maturity rates 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 factors other than U.S. monetary policy announcements.

6 Global long-term real interest rates and US monetary policy

Large parts of the decline in long-term bond yields over the past three decades can be attributed to a fall in long-term real interest rates (see e.g. Benigno et al. (2024). Against this background, we assess whether our findings for nominal long-term interest rates carry over to long-term real interest rates.

As a measure of ex-ante long-term real interest rates, we use 10-year inflation-protected securities, such as Treasury Inflation-Protected Securities (TIPS) in the U.S. The spread between the nominal government bond yield and the inflation-protected yield is a market-based measure of long-term inflation expectations, commonly referred to as the “breakeven inflation rate”. However, the inflation-protected bond market has a relatively short history, and not all countries in our sample issue inflation-protected securities. We are able to obtain market-based long-term real interest rates only for Australia, Canada, Germany, Sweden, the U.K. and the U.S., but not for Japan, Switzerland, Norway, and New Zealand from the analysis. The TIPS yield data are sourced from Bloomberg.

In Figure 10, we plot the observed cumulative changes in 10-year real interest rates, along with changes during FOMC and non-FOMC windows. Note that the start date of this figure differs from that of Figure 2 due to data availability. Also, for Australia (AUD) and Canada (CAD), some flat data are observed due to a lack of TIPS data. Overall, the figures demonstrate that the decline in long-term real interest rates across countries is again heavily concentrated during FOMC windows. Also for real rates, the rate changes during the FOMC window fits the actual yields quite well as reflected in the high pseudo R^2s reported in the panel titles. The best fit obtains again for the U.S. 10-year real yield, with a pseudo R^2 of 88% of the total variation in the U.S. 10-year yield. Also for the other countries, the FOMC windows have very high explanatory power accounting for more than half, and except for Sweden even for more than two thirds of the observed total yield variations. This reinforces the evidence presented in Figure 2, indicating that the findings carry over to real rates, albeit over a shorter sample period.

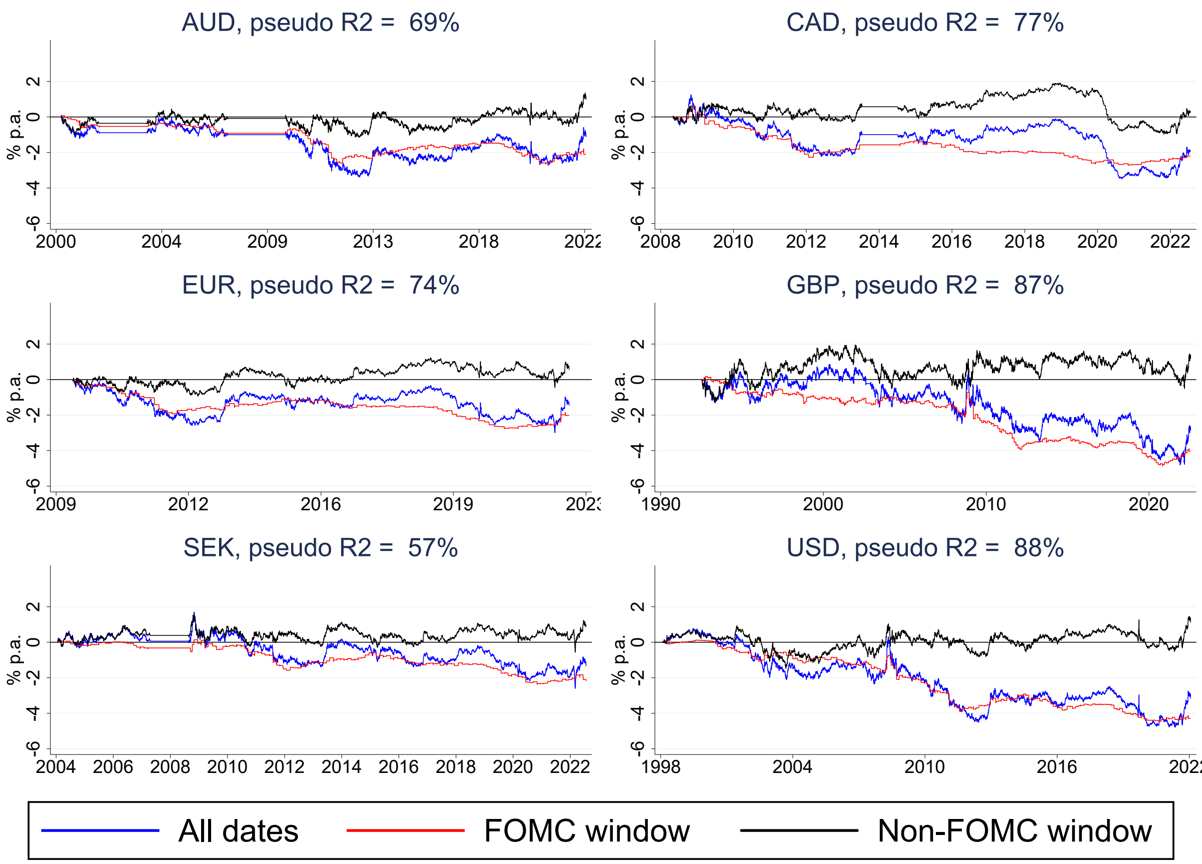

Figure 10: Cumulative changes in real 10-year yields during FOMC announcement windows.

Notes. Unit of the y-axis is percentage p.a.. The real rates are constructed using inflation-protected securities. The cumulative change in the 10-year yield y_t during event window W is $\nabla y_t^W = \sum_{s=t_0+1}^{t}(y_s - y_{s-1})\mathbf{1}_{s\in W}$. The FOMC window consists of $\{t-1, t, t+1\}$ if date t is an FOMC announcement date. The non-FOMC window complements the FOMC window. The pseudo R^2 is defined as $R^2 = 1 - \frac{\sum_{t=1}^{T}(y_t - \nabla y_t^{FOMC})^2}{\sum_{t=1}^{T}(y_t - \nabla y_t^{nonFOMC})^2 + \sum_{t=1}^{T}(y_t - \nabla y_t^{FOMC})^2}$.

The mechanism is the same as for nominal long-term bond yields. Significant international spillovers of U.S. real rate changes transmit the cumulative negative FOMC announcement effect on U.S. real yields globally. To demonstrate this point, we estimate Equation (6) for long-term real rates instead of nominal rates. The results reported in Table 7 indicate sizable and statistically significant spillovers from the U.S. to other countries. According to our estimates, a one percentage point fall of U.S. long-term real rates during FOMC announcement windows is associated with a 0.77 percentage point reduction on long-term real yields across other countries. The effect is even somewhat stronger when we consider all observations available.

Table 7: U.S. spillovers in long-term real bond yields.

	FOMC window regression	Full sample regression
Δy_{US}	0.769***	0.825***
	(0.025)	(0.013)
R^2	0.591	0.533
adj. R^2	0.590	0.533
within R^2	0.591	0.533
N	2,019	20,702

*p<0.1; **p<0.05; ***p<0.01.
Standard errors in parentheses are clustered by date.

7 Placebo tests

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

Placebo 1 For each year y, let N_y denote the number of FOMC announcements in that year. Randomly draw N_y trading days from that year to serve as the hypothetical FOMC announcement days. Construct hypothetical FOMC windows from $t-5$ to $t+1$. Compute the cumulative changes in the stock index during or outside the hypothetical FOMC windows. Repeat this exercise 10,000 times and compute the 2.5$^{\text{th}}$ and 97.5$^{\text{th}}$ percentiles.

Placebo 2 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. The remaining procedures are identical to those in Placebo 1.

Figure 12 presents the placebo test result based on 10,000 random samples. The 90% confidence bands denote the 5$^{\text{th}}$ and 95$^{\text{th}}$ 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.

8 Conclusion

The findings of this paper highlight a significant influence of U.S. monetary policy announcements on long-run global trends in long-term interest rates. Almost 70% of the total decline in 10-year government bond yields across G10 currencies over the past three decades is attributable to FOMC announcement effects. By contrast, other central banks' announcements have played only a minor role in the global secular decline in long-term interest rates, with measurable domestic effects only in a few countries.

We offer various assessments of the nature of the relationship between FOMC announcements and the secular decline in interest rates. Specifically, we find that this relationship is driven by monetary policy shocks, changes in expected interest rates and changes in real interest rates rather than information effects, term premia or inflation expectations. All this implies that changes in the stance of monetary policy are at the heart of the relationship.

How can monetary policy have such persistent negative effects on long-term nominal and real interest rates? A fully fledged assessment of this question is beyond the scope of this paper. One possible mechanism is through information feedback loops between central banks and financial markets under imperfect information which could ultimately influence real outcomes (Rungcharoenkitkul and Winkler (2021)). The empirical findings presented in this paper would seem to make the case for further research on this question.

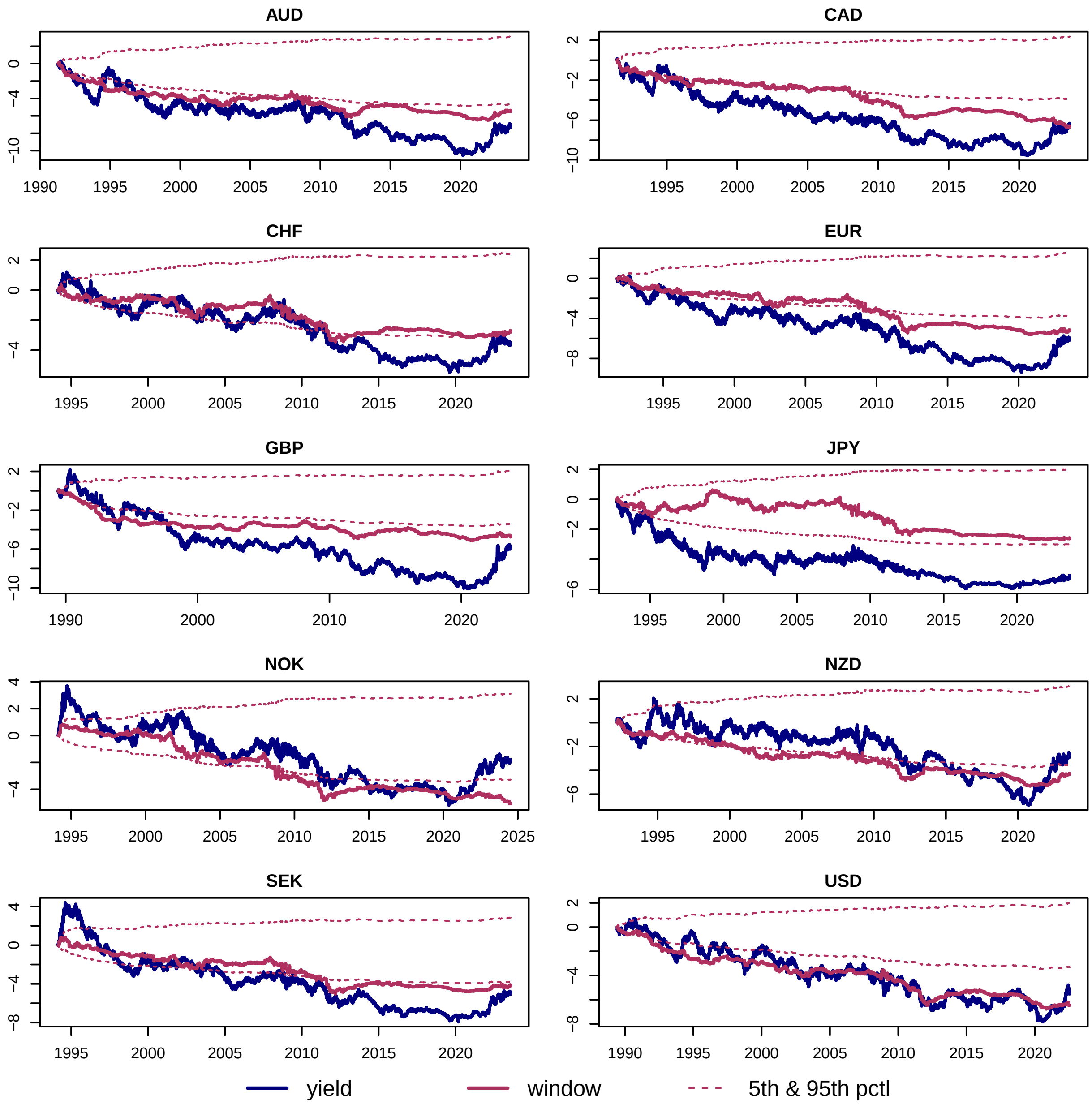

Figure 11: Placebo test 1.

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 business days from each year y serving as the placebo announcement days. Then, we construct the placebo announcement windows accordingly. The dashed lines denote the 5^{th} and 95^{th} percentiles from 10,000 placebo samples.

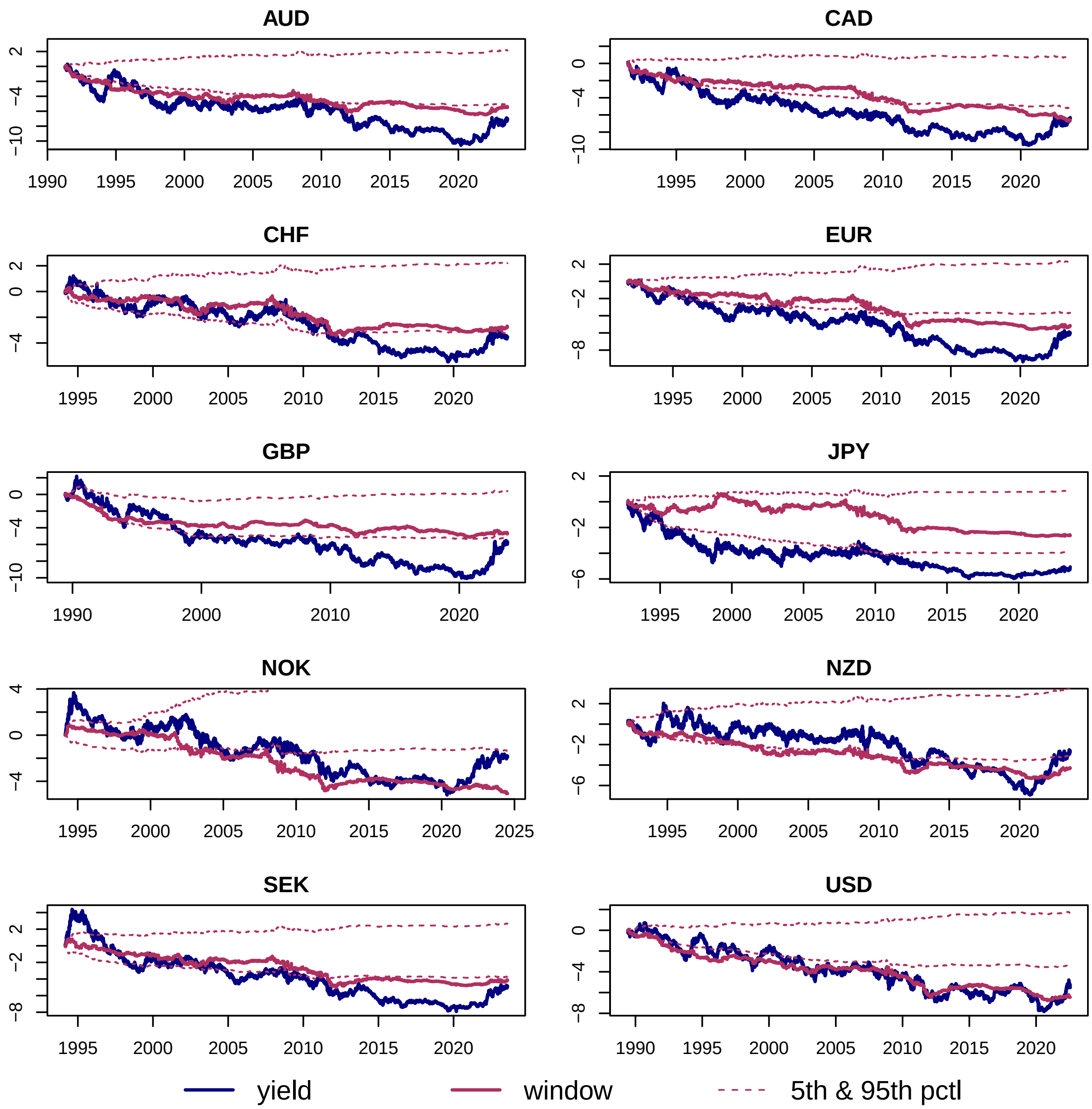

Figure 12: Placebo test 2.

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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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.

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

Table A2: Unscheduled FOMC Meeting Dates

Year	N	Unscheduled FOMC Meetings 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									

B U.K. yields during FOMC windows

Figure 2 seems to suggest that FOMC announcement windows explain little variations in the U.K. 10-year yield. Here, we demonstrate significant pre-announcement drift in the U.K. yield before FOMC announcements and substantial improvements in the explanatory power when the announcement window is

Table A3: Overlaps with FOMC announcement windows.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
MP start	1990	2000	2004	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

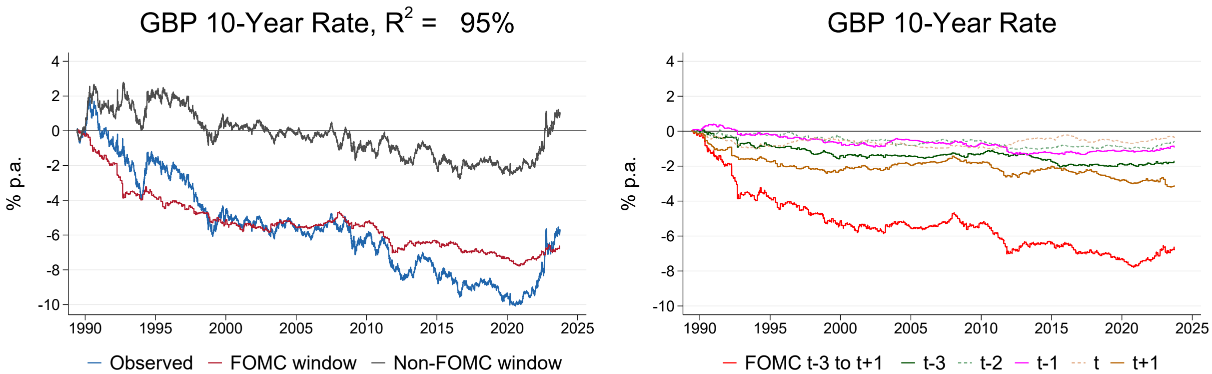

Figure A1: U.K. yields during extended FOMC windows.

Notes. The FOMC window extends from 3 days before the announcement to the day after. The left panel shows the cumulative yield changes during or outside the extended FOMC windows. The right panel decomposes the FOMC-window series into the sums of daily changes on each day in the FOMC window.

extended backward a little bit to capture the drift.

We extend the FOMC announcement window to start 3 days before the announcement and end on the day after the announcement. Figure A1 shows that the cumulative change in the U.K. 10-year yield during the extended FOMC window explains 95% of the total variation in the data, increasing from 58% during the baseline FOMC window. Therefore, a substantial fraction of the U.K. yield's responses to U.S. monetary policy announcements occur during $t-2$ and $t-3$.

To analyze which days in the window are more important for the total response, we compute the sums of daily yield changes for each day in the extended FOMC window. According to Figure A1, the total response of the U.K. 10-year yield is primarily attributable to changes on days $t+1$ and $t-3$. The yield also declines systematically on days $t-2$ to t, but less so than on the former two days. This explains the limited variation in the yield during the baseline FOMC window observed in Figure 2.

C The shifting endpoint model

C.1 Details of the no-arbitrage recursions

First, we show that the state vector Z_t evolves as

$$Z_t = \boldsymbol{\mu}_Z + \Phi_Z Z_{t-1} + V_t, \quad V_t \equiv \begin{bmatrix} \boldsymbol{\eta}_t \\ U_t \end{bmatrix}, \tag{A1}$$

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

$$\Lambda_1 = \begin{bmatrix} (I-\Phi)\Gamma & \Lambda_{12} \end{bmatrix} \tag{A2}$$

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

$$p_t^{(n)} = \mathbf{E}_t[m_{t+1}] + \mathbf{E}_t[p_{t+1}^{(n-1)}] + \frac{1}{2}\mathbf{Var}_t(m_{t+1}) + \frac{1}{2}\mathbf{Var}_t(p_{t+1}^{(n-1)}) + \mathbf{Cov}_t(m_{t+1}, p_{t+1}^{(n-1)}). \tag{A3}$$

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 (A3).

Finally, we derive the bond pricing recursions. Equation (A3) together with Equation (A2) implies

$$\begin{aligned}p_t^{(n)} =& -\delta_0 - \boldsymbol{\delta}_1^\top X_t + \mathcal{A}_{n-1} + \mathcal{B}_{n-1}^\top[(I-\Phi)\boldsymbol{\mu} + \Phi X_t] && \text{(A4)}\\ &+ \frac{1}{2}\mathcal{B}_{n-1}^\top\Omega\mathcal{B}_{n-1} - \mathcal{B}_{n-1}^\top(\Lambda_0 + \Lambda_{12}X_t). && \text{(A5)}\end{aligned}$$

So,

$$\begin{aligned}\mathcal{A}_n =&\mathcal{A}_{n-1} - \delta_0 + \mathcal{B}_{n-1}^\top(I-\Phi)\boldsymbol{\mu} + \frac{1}{2}\mathcal{B}_{n-1}^\top\Omega\mathcal{B}_{n-1} - \mathcal{B}_{n-1}^\top\Lambda_0, && \text{(A6)}\\ \mathcal{B}_n^\top =& -\boldsymbol{\delta}_1^\top + \mathcal{B}_{n-1}^\top(\Phi - \Lambda_{12}). && \text{(A7)}\end{aligned}$$

The yields are

$$y_t^{(n)} = A_n + B_n^\top X_t, \tag{A8}$$

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

$$p_t^{(n),rn} = \ln \mathbf{E}_t \left[\exp \left\{ -y_t^{(1)} + p_{t+1}^{(n-1),rn} \right\} \right]. \tag{A9}$$

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

$$p_t^{(n),rn} = \mathcal{A}_n^{rn} + \mathcal{B}_n^{rn\top} X_t + \mathcal{C}_n^{rn\top} \boldsymbol{\tau}_t, \tag{A10}$$

and the coefficients solve the following recursions:

$$\begin{aligned}
\mathcal{A}_n^{rn} =& \mathcal{A}_{n-1}^{rn} - \delta_0 + \mathcal{B}_{n-1}^{rn\top}(I-\Phi)\boldsymbol{\mu} + \frac{1}{2}\mathcal{B}_{n-1}^{rn\top}\Omega\mathcal{B}_{n-1}^{rn} \\
& + \frac{1}{2}\left(\mathcal{B}_{n-1}^{rn\top}\Gamma\Omega_\eta\mathcal{C}_{n-1}^{rn} + \mathcal{C}_{n-1}^{rn\top}\Omega_\eta\Gamma^\top\mathcal{B}_{n-1}\right) + \frac{1}{2}\mathcal{C}_{n-1}^{rn\top}\Omega_\eta\mathcal{C}_{n-1}^{rn}, \\
\mathcal{B}_n^{rn\top} =& -\boldsymbol{\delta}_1^\top + \mathcal{B}_{n-1}^{rn\top}\Phi, \\
\mathcal{C}_n^{rn\top} =& \mathcal{B}_{n-1}^{rn\top}(I-\Phi)\Gamma + \mathcal{C}_{n-1}^{rn\top}
\end{aligned} \tag{A11}$$

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

C.2 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 (A1) 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

$$rx_{t+1}^{(n)} = \mathcal{B}_{n-1}^\top(\Lambda_0 + \Lambda_1 Z_t) - \mathcal{B}_{n-1}^\top\Omega\mathcal{B}_{n-1} + \mathcal{B}_{n-1}^\top U_{t+1}. \tag{A12}$$

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

$$
\begin{aligned}
rx_{t+1}^{(n)} =& \mathcal{B}_{n-1}^{\top}(\Lambda_0 + \Lambda_1 Z_t) - \frac{1}{2}\mathcal{B}_{n-1}^{\top}\Omega\mathcal{B}_{n-1} + \mathcal{B}_{n-1}^{\top}\left(X_{t+1} - (I-\Phi)(\boldsymbol{\mu} + \Gamma\boldsymbol{\tau}_t) - \Phi X_t\right) \\
=& \mathcal{B}_{n-1}^{\top}(\Lambda_0 - (I-\Phi)\boldsymbol{\mu}) - \frac{1}{2}\mathcal{B}_{n-1}^{\top}\Omega\mathcal{B}_{n-1} \\
&+ \mathcal{B}_{n-1}^{\top}(\Lambda_{11}\boldsymbol{\tau}_t + \Lambda_{12}X_t - (I-\Phi)\Gamma\boldsymbol{\tau}_t - \Phi X_t) + \mathcal{B}_{n-1}^{\top}X_{t+1} \\
=& \mathcal{B}_{n-1}^{\top}(\Lambda_0 - (I-\Phi)\boldsymbol{\mu}) - \frac{1}{2}\mathcal{B}_{n-1}^{\top}\Omega\mathcal{B}_{n-1} + \mathcal{B}_{n-1}^{\top}(\Lambda_{12} - \Phi)X_t + \mathcal{B}_{n-1}^{\top}X_{t+1}.
\end{aligned}
\tag{A13}
$$

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 (A12) or Equation (A13). To estimate Equation (A12), 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 (A13), we regress $rx_{t+1}^{(n)}$ on X_t and X_{t+1} without any constraints. Either way, we obtain the same estimation results.

C.3 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 $[\nabla PC1_{loc,t}^{US}, \nabla PC1_{US,t}^{US}]^{\top}$. 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-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.

D Additional results

D.1 Purified monetary policy shocks

In Figure A2, 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.

D.2 Yield dynamics during central bank announcement windows

Contemporaneous effects In Table A7, we report pairwise estimation results using the 10-year yield as the dependent variable. Each row fixes the country i, whose 10-year yield is the dependent variable; each column fixes the country j, whose monetary policy shock is the independent variable. To ensure that we have records of all central banks' announcement dates, our sample starts from February 2000. The diagonal cells report the marginal effects of each central bank's monetary policy shocks on the domestic 10-year yield.

Does the U.S. monetary policy shock have stronger spillover effects than other countries' monetary policy shocks? Each row of Table A7 reports the responses of the row country's 10-year yield to each column country's monetary policy shock. For every country, there exists a foreign monetary policy shock that has a stronger spillover effect than the U.S. monetary policy shock. For example, the German 10-year yield responds to the Swiss monetary policy shock with a coefficient of 0.81 and responds to the U.S. monetary policy shock with a coefficient of 0.69. The R^2 associated with the Swiss monetary policy shock is more than twice as large as that associated with the U.S. monetary policy shock.

Table A7: Regressions for daily yield changes: 10-year yields.

Yields \ CB	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
AUD	0.82	0.61	0.92	0.67	0.63	1.05	0.61	0.57	0.95	0.71
se	0.05	0.09	0.16	0.06	0.07	0.14	0.09	0.09	0.10	0.16
R^2	0.59	0.21	0.36	0.38	0.28	0.33	0.24	0.28	0.44	0.17
CAD	0.39	0.78	0.60	0.40	0.28	0.72	0.40	0.27	0.53	0.58
se	0.05	0.05	0.11	0.04	0.06	0.10	0.07	0.07	0.08	0.11
R^2	0.23	0.62	0.29	0.25	0.13	0.27	0.19	0.11	0.26	0.18
CHF	0.48	0.56	0.91	0.62	0.55	0.92	0.46	0.38	0.79	0.51
se	0.04	0.07	0.11	0.04	0.08	0.10	0.08	0.07	0.08	0.13
R^2	0.34	0.27	0.61	0.51	0.36	0.39	0.25	0.22	0.46	0.14
EUR	0.52	0.60	0.81	0.75	0.68	0.95	0.56	0.45	0.86	0.69
se	0.04	0.07	0.13	0.05	0.08	0.12	0.09	0.08	0.09	0.13
R^2	0.33	0.28	0.45	0.58	0.41	0.36	0.27	0.25	0.41	0.20
GBP	0.38	0.49	0.72	0.36	0.51	0.70	0.27	0.26	0.47	0.51
se	0.05	0.06	0.16	0.05	0.06	0.11	0.08	0.08	0.08	0.12
R^2	0.18	0.23	0.28	0.23	0.28	0.23	0.08	0.08	0.20	0.13
JPY	0.35	0.42	0.59	0.44	0.37	0.79	0.40	0.37	0.69	0.43
se	0.04	0.06	0.13	0.04	0.07	0.09	0.08	0.06	0.07	0.13
R^2	0.30	0.24	0.32	0.40	0.23	0.43	0.24	0.28	0.41	0.13
NOK	0.52	0.62	0.86	0.67	0.54	0.82	0.69	0.32	1.04	0.46
se	0.05	0.07	0.14	0.08	0.08	0.12	0.12	0.09	0.10	0.11
R^2	0.33	0.21	0.51	0.46	0.28	0.25	0.43	0.13	0.52	0.10
NZD	0.47	0.47	0.77	0.49	0.38	0.81	0.48	0.58	0.74	0.51
se	0.06	0.06	0.18	0.05	0.07	0.10	0.07	0.08	0.07	0.15
R^2	0.23	0.21	0.31	0.30	0.16	0.27	0.16	0.34	0.38	0.12
SEK	0.56	0.61	0.87	0.68	0.62	0.88	0.57	0.43	0.93	0.58
se	0.05	0.07	0.15	0.04	0.07	0.11	0.10	0.08	0.09	0.14
R^2	0.37	0.30	0.48	0.50	0.38	0.31	0.28	0.23	0.46	0.13
USD	0.25	0.38	0.35	0.53	0.54	0.40	0.33	0.26	0.40	0.81
se	0.05	0.06	0.07	0.04	0.06	0.11	0.07	0.05	0.07	0.09
R^2	0.11	0.18	0.14	0.43	0.37	0.13	0.18	0.15	0.19	0.42

Notes. The table reports $\hat{\beta}_1$, $se(\hat{\beta}_1)$, and R^2 from the regression $\Delta_3 y_{r,t}^{(n)} = \beta_0 + \beta_1 \Delta_3 y_{c,t}^{(2)} + u_t$, where $\Delta_3 y_{r,t}^{(n)}$ is the three-day change in the yield of the row country bracketing the column country's monetary policy announcements, and $\Delta_3 y_{c,t}^{(2)}$ is the same measure for the column country's 2-year yield. The regression is estimated separately for each $\{r, c\}$ pair.

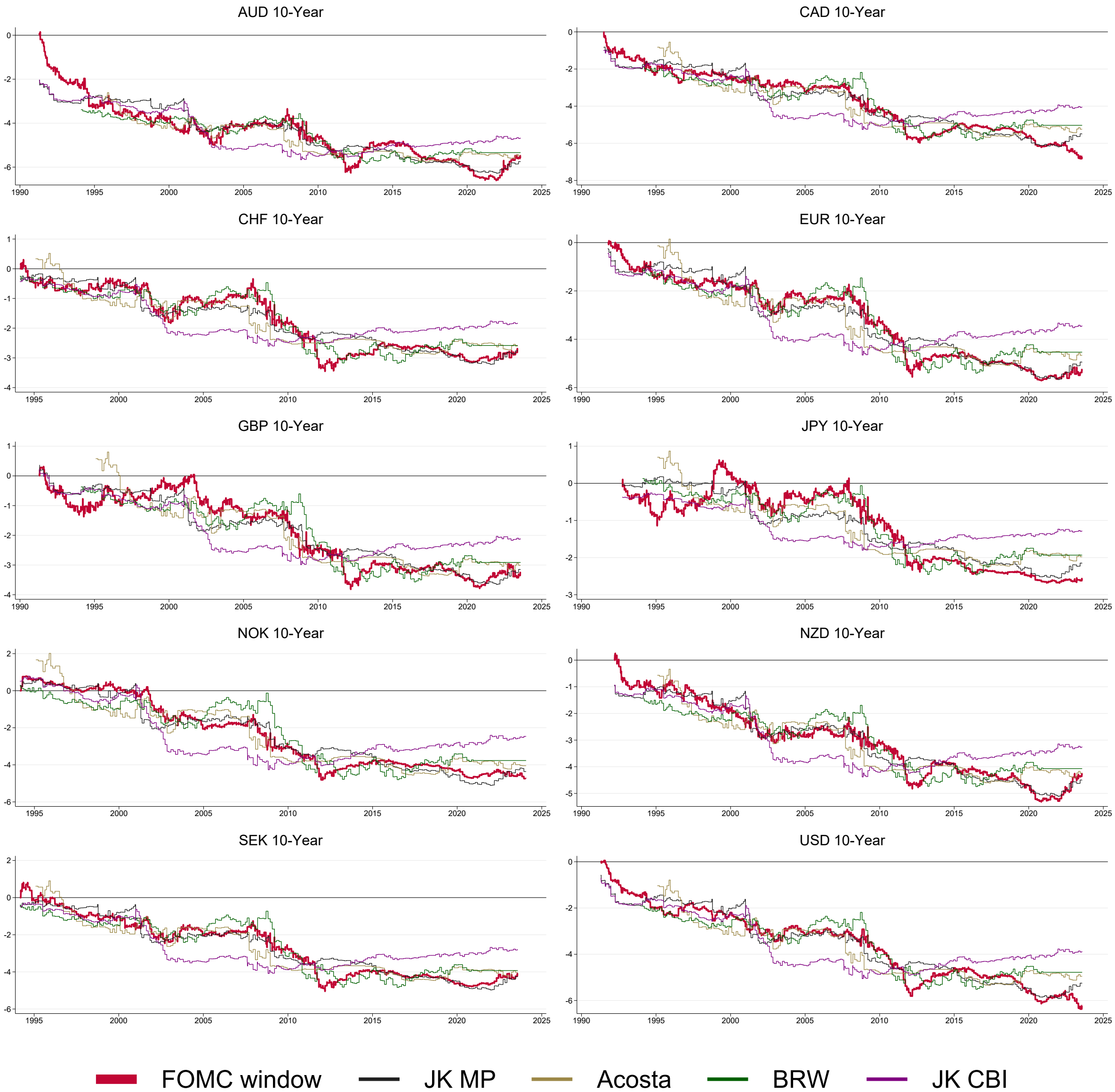

Figure A2: Cumulative yield changes and sums of U.S. HF monetary policy shocks.

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).

Table A8 reports the estimation results for

$$\Delta_3 y_{r,t}^{(n)} = \beta_0 + \beta_1 \Delta_3 y_{c,t}^{(n)} + u_t, \tag{A14}$$

where $\Delta_3 y_{r,t}^{(n)}$ is the three-day change in the n-year yield of the row country bracketing the column country's monetary policy announcements, and $\Delta_3 y_{c,t}^{(n)}$ is the same measure for the column country's n-year yield. The regression is estimated separately for each $\{r, c\}$ pair. In the main text, we estimated a similar equation using the column country's 2-year rate shock $\Delta_3 y_{c,t}^{(2)}$ as the independent variable. Here, we match the maturities of the interest rates on both sides of the equation like in Table 2 of Kim and Ochoa (2023). Although their sample runs from January 2010 to October 2017, their results are consistent with ours.

The results are quantitatively similar to those in the main text. Monetary policy shocks of all countries have significant spillover effects on other countries' 10-year yields, indicating strong contemporaneous comovements among world long-term interest rates. Moreover, the coefficient on the U.S. monetary policy shock is not the most significant, and the R^2 associated with the U.S. shock is not the largest within each row.

Cumulative effects Table A9 reports the pseudo R^2 between the cumulative announcement effects of column central banks and row 10-year yields.

Table A10 reports the RMSE for the column country's monetary policy windows to fit the row country's 10-year yield.

Next, we plot the cumulative effects of each country's central bank monetary policy announcements on world 10-year yields. In each of Figure A3-A11, 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.

D.3 Cointegration tests

Table A11 reports estimation results for the cointegration equation

$$y_t = \gamma_0 + \boldsymbol{\gamma}_1^\top \boldsymbol{\tau}_t + u_t, \tag{A15}$$

Table A8: Regressions for daily yield changes: 10-year yields.

Yields \ CB	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
AUD	1.00	0.85	1.12	0.91	0.72	1.23	0.86	0.86	0.87	0.88
se	0.00	0.09	0.10	0.04	0.07	0.07	0.06	0.05	0.05	0.11
R^2	1.00	0.41	0.72	0.68	0.36	0.64	0.52	0.66	0.71	0.41
CAD	0.48	1.00	0.75	0.53	0.42	0.85	0.62	0.47	0.54	0.71
se	0.04	0.00	0.06	0.05	0.06	0.06	0.07	0.07	0.06	0.10
R^2	0.41	1.00	0.60	0.45	0.27	0.53	0.50	0.34	0.52	0.43
CHF	0.59	0.78	1.00	0.82	0.61	1.09	0.73	0.59	0.78	0.76
se	0.04	0.07	0.00	0.02	0.07	0.04	0.06	0.07	0.03	0.09
R^2	0.58	0.52	1.00	0.89	0.42	0.79	0.67	0.54	0.86	0.49
EUR	0.64	0.80	0.97	1.00	0.69	1.12	0.87	0.66	0.92	0.94
se	0.04	0.07	0.05	0.00	0.07	0.04	0.06	0.08	0.04	0.08
R^2	0.57	0.50	0.85	1.00	0.39	0.73	0.71	0.54	0.89	0.59
GBP	0.51	0.71	0.88	0.50	1.00	0.85	0.56	0.48	0.50	0.64
se	0.05	0.06	0.07	0.04	0.00	0.06	0.06	0.07	0.06	0.09
R^2	0.37	0.49	0.57	0.42	1.00	0.50	0.37	0.31	0.43	0.32
JPY	0.43	0.61	0.76	0.59	0.48	1.00	0.61	0.51	0.64	0.65
se	0.04	0.06	0.05	0.03	0.06	0.00	0.06	0.06	0.05	0.08
R^2	0.51	0.49	0.72	0.69	0.37	1.00	0.60	0.54	0.68	0.46
NOK	0.63	0.73	0.93	0.86	0.56	1.04	1.00	0.55	0.91	0.79
se	0.04	0.08	0.06	0.04	0.08	0.07	0.00	0.11	0.06	0.11
R^2	0.56	0.29	0.79	0.74	0.28	0.59	1.00	0.40	0.74	0.46
NZD	0.68	0.69	0.96	0.67	0.51	1.02	0.74	1.00	0.68	0.68
se	0.07	0.07	0.09	0.04	0.06	0.07	0.08	0.00	0.06	0.12
R^2	0.55	0.43	0.64	0.55	0.27	0.62	0.41	1.00	0.61	0.32
SEK	0.67	0.79	1.00	0.92	0.61	1.08	0.88	0.67	1.00	0.92
se	0.04	0.07	0.06	0.03	0.07	0.04	0.05	0.08	0.00	0.09
R^2	0.60	0.49	0.83	0.90	0.33	0.68	0.72	0.56	1.00	0.52
USD	0.37	0.43	0.48	0.65	0.41	0.50	0.48	0.36	0.46	1.00
se	0.05	0.07	0.10	0.04	0.06	0.07	0.05	0.04	0.06	0.00
R^2	0.28	0.24	0.35	0.62	0.20	0.29	0.41	0.29	0.48	1.00

Notes. The table reports $\hat{\beta}_1$, $se(\hat{\beta}_1)$, and R^2 from the regression $\Delta_3 y_{r,t}^{(n)} = \beta_0 + \beta_1 \Delta_3 y_{c,t}^{(n)} + u_t$, where $\Delta_3 y_{r,t}^{(n)}$ is the three-day change in the n-year yield of the row country bracketing the column country's monetary policy announcements, and $\Delta_3 y_{c,t}^{(n)}$ is the same measure for the column country's n-year yield. The regression is estimated separately for each $\{r, c\}$ pair.

where y_t denotes the level factor of a country's yield curve. We consider three specifications of $\boldsymbol{\tau}_t$. I: $\tau_t = \nabla PC1_{loc,t}^{US}$ is the first principal component of cumulative changes in the local yield curve during FOMC announcement windows. II: $\tau_t = \nabla PC1_{US,t}^{US}$ is the first principal component of cumulative changes in the U.S. yield curve during FOMC announcement windows. III: $\boldsymbol{\tau}_t = [\nabla PC1_{loc,t}^{US}, \nabla PC1_{US,t}^{US}]^\top$. Along with the

Table A9: Pseudo R^2 between the cumulative announcement effects of central banks and 10-year yields.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
AUD	27	2	3	38	8	90	1	8	12	28
CAD	3	23	40	15	10	37	70	2	8	75
CHF	94	3	13	8	9	94	47	8	5	75
EUR	73	1	1	5	3	89	34	6	2	77
GBP	12	81	6	10	11	90	18	11	4	79
JPY	14	63	38	57	59	73	73	6	10	89
NOK	35	13	9	30	55	15	40	17	4	93
NZD	57	51	1	5	5	40	5	9	14	62
SEK	60	2	2	7	10	91	21	1	3	70
USD	7	9	5	12	6	28	52	6	12	72

Notes. The table reports the pseudo $R^2 = 1 - \frac{\sum_{t=1}^{T}\left(y_t - \nabla y_t^{MP}\right)^2}{\sum_{t=1}^{T}\left(y_t - \nabla y_t^{nonMP}\right)^2 + \sum_{t=1}^{T}\left(y_t - \nabla y_t^{MP}\right)^2}$, where y_t is the actual level of the yield of the row country, and ∇y_t^{MP} and ∇y_t^{nonMP} are the cumulative changes in y_t during or outside the three-day windows bracketing the announcement of the column central bank. The sample is post-2000. The unit is percentage points.

Table A10: Fitting the 10-year 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	0.80	1.22	0.95	1.99	1.31	2.47	1.32	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	4.24	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.

estimation results, we also report stationarity test statistics for u_t. For the augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests, we report the test statistics, and the null hypothesis is that u_t contains a unit root; for the Müller-Watson low-frequency stationarity test (LFST), we report the p-value, and the null hypothesis is that u_t is stationary. The results suggest that the level factors of G10 countries are cointegrated

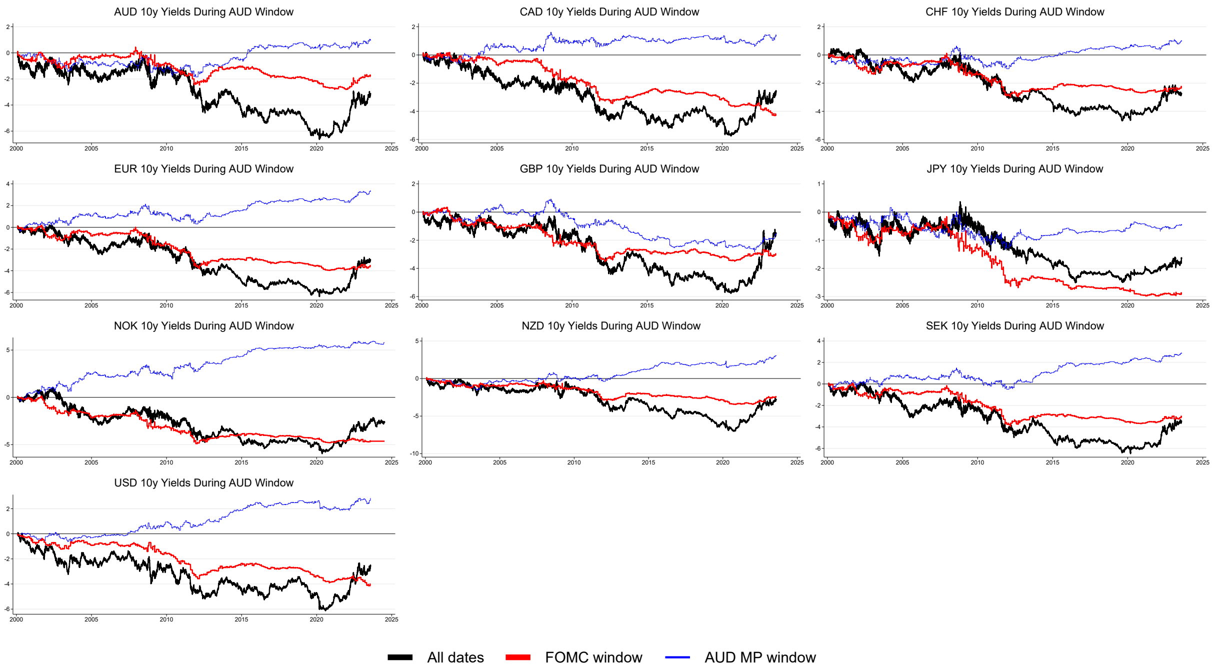

Figure A3: Sum of daily yield changes during U.S. or Australian central bank announcement windows.

with the cumulative effects of FOMC announcements, and the residuals are stationary.

D.4 OSE vs. FE

In the main text, we demonstrated that the FOMC windows explain most of the total variations in the risk-neutral rate. Such a result might be mechanical because our trend proxy, which is a key component of the risk-neutral rate, only changes during the FOMC windows. Here, 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. In Figure A12, the thick solid lines indicate the results of the OSE model, and the thin dashed lines indicate the results of the FE model. Both models imply persistent declines in the risk-neutral rates in all countries, but the ones implied by the OSE model decline more. Moreover, the FOMC windows can also explain the total variations in the FE risk-neutral rates well. Therefore, 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.

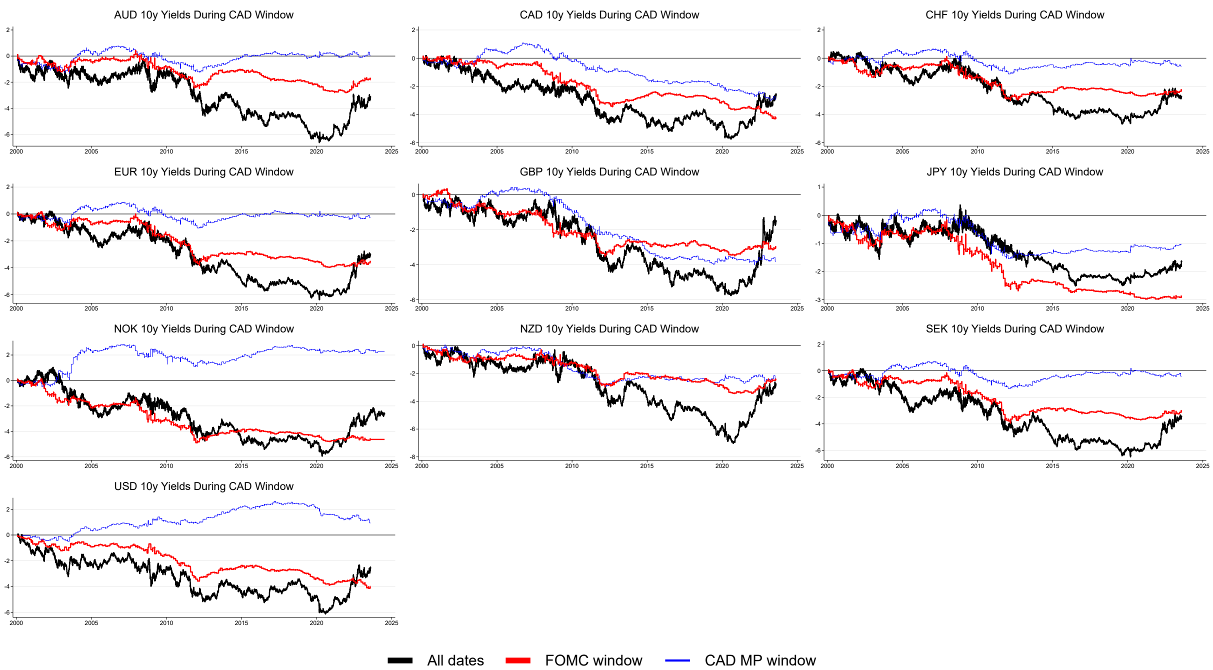

Figure A4: Sum of daily yield changes during U.S. or Canadian central bank announcement windows.

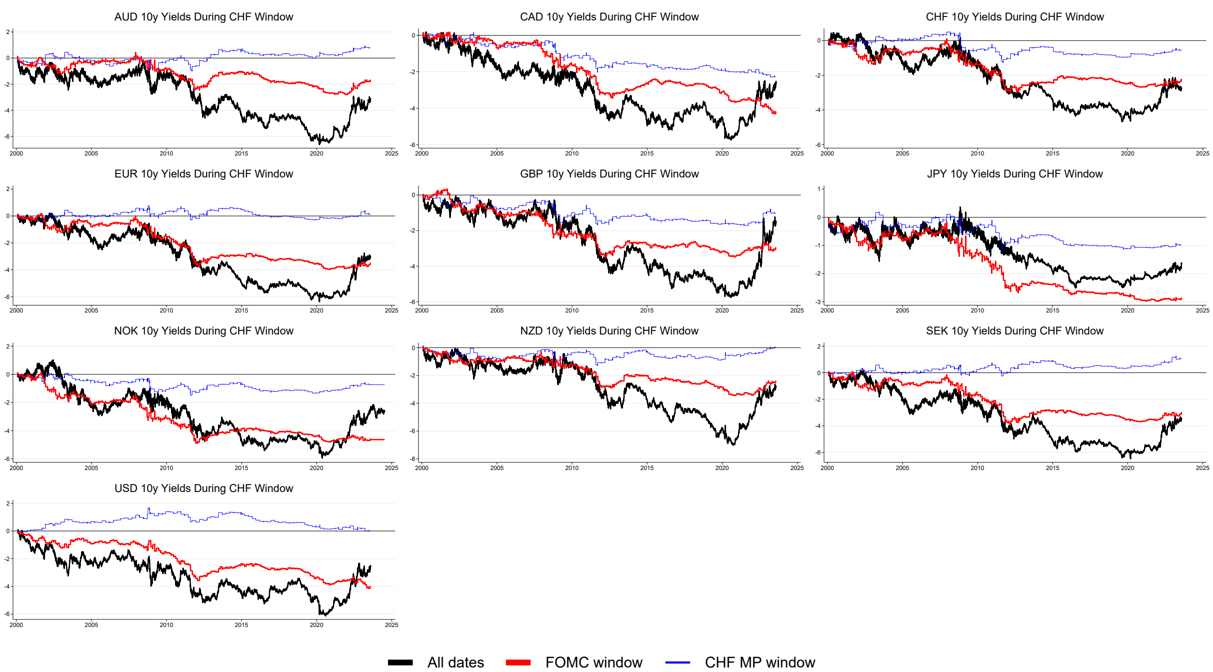

Figure A5: Sum of daily yield changes during U.S. or Swiss central bank announcement windows.

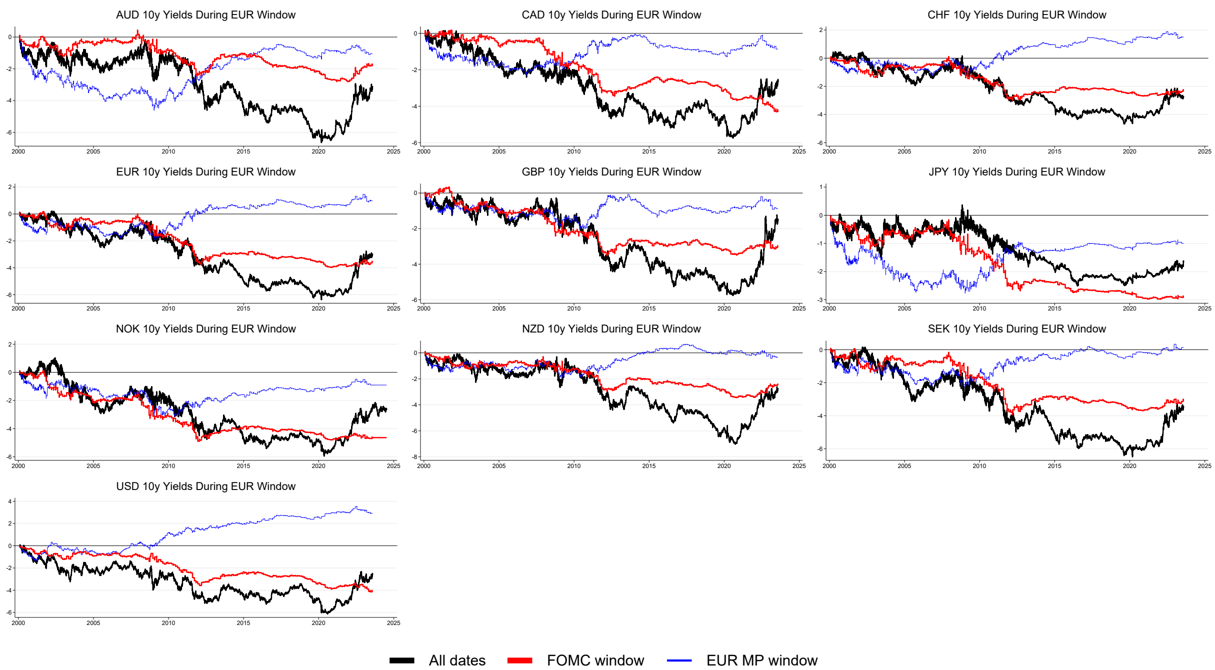

Figure A6: Sum of daily yield changes during U.S. or ECB announcement windows.

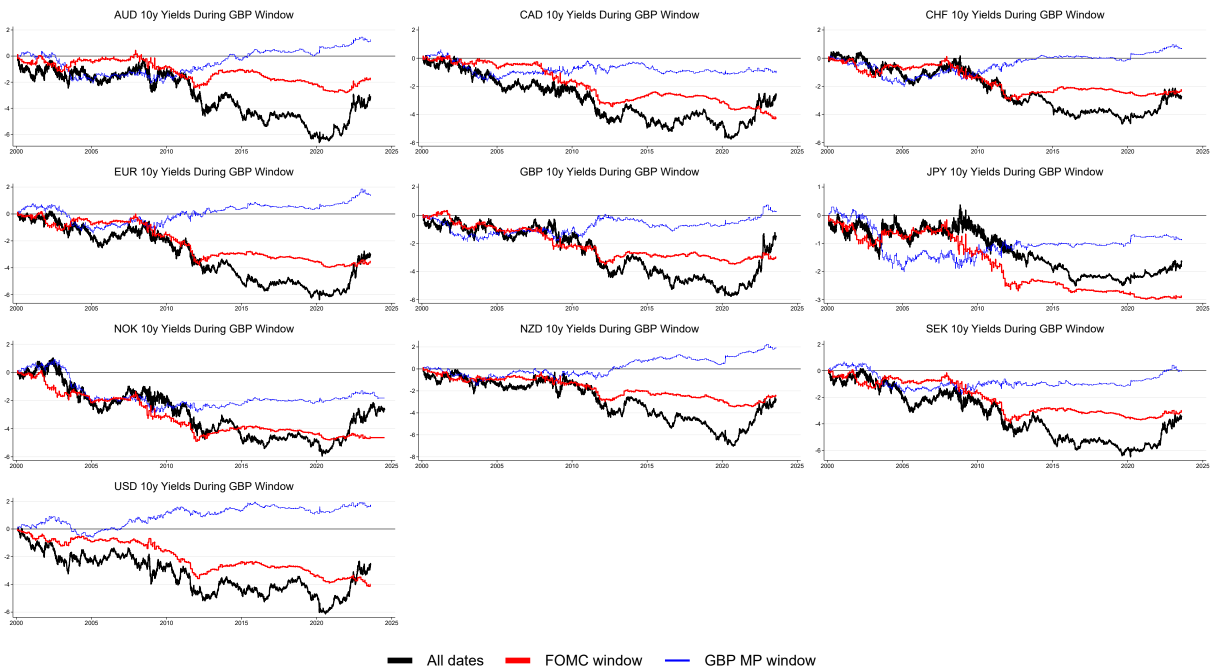

Figure A7: Sum of daily yield changes during U.S. or U.K. central bank announcement windows.

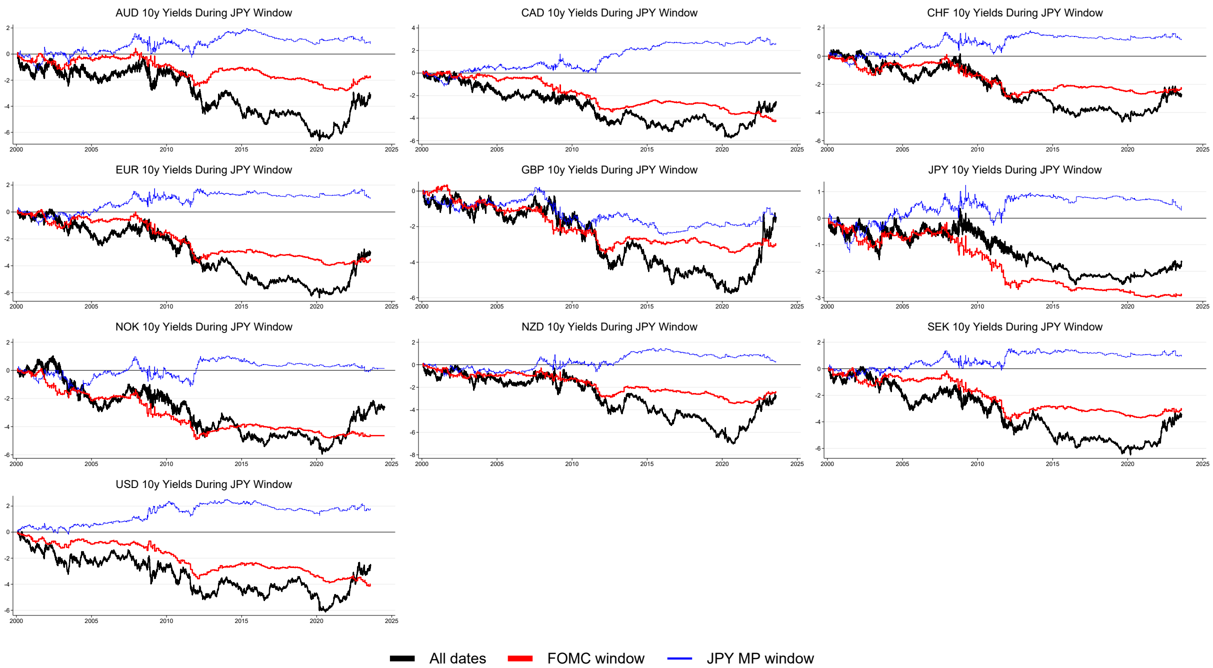

Figure A8: Sum of daily yield changes during U.S. or Japanese central bank announcement windows.

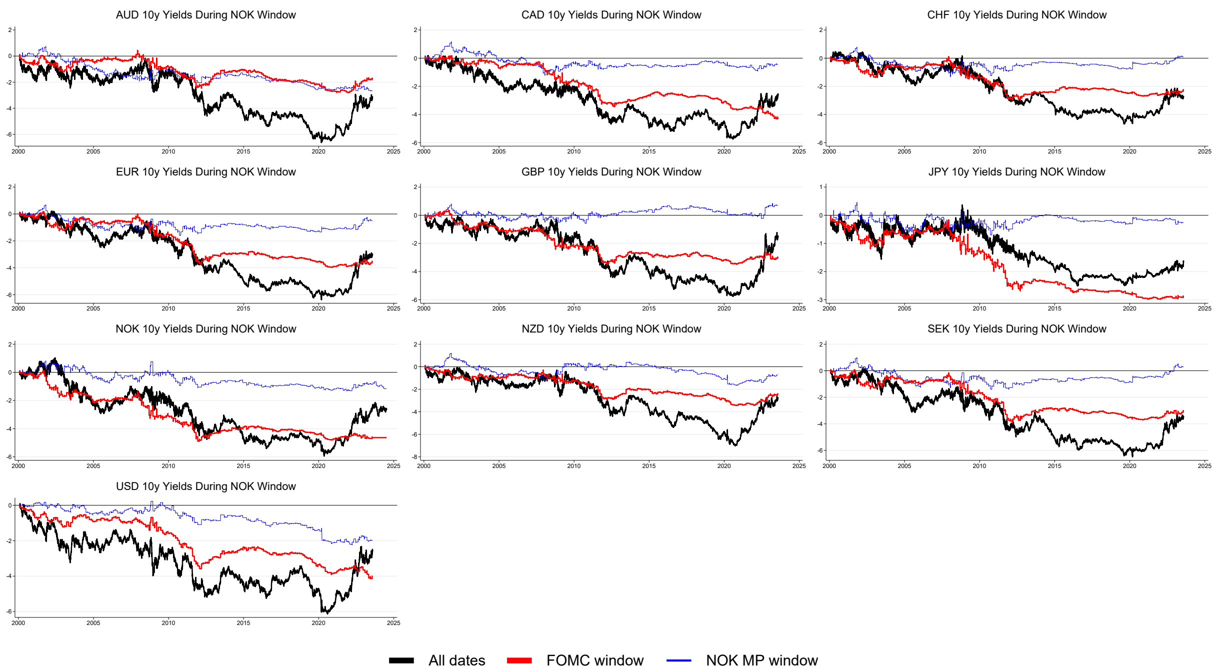

Figure A9: Sum of daily yield changes during U.S. or Norwegian central bank announcement windows.

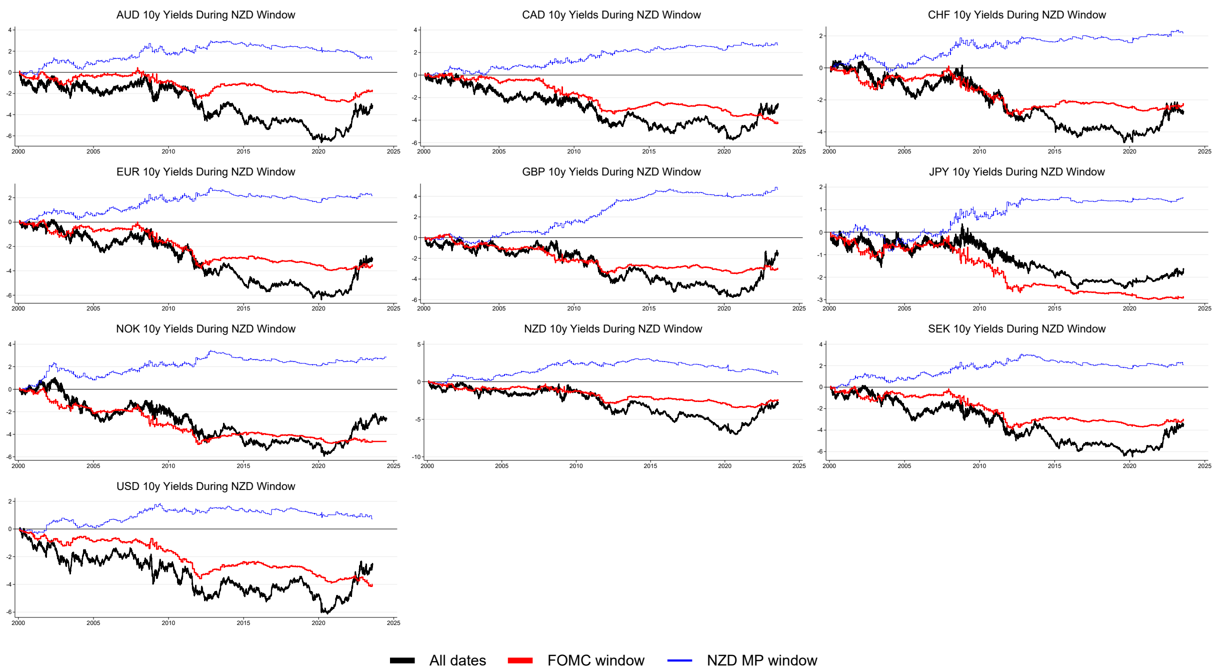

Figure A10: Sum of daily yield changes during U.S. or New Zealand central bank announcement windows.

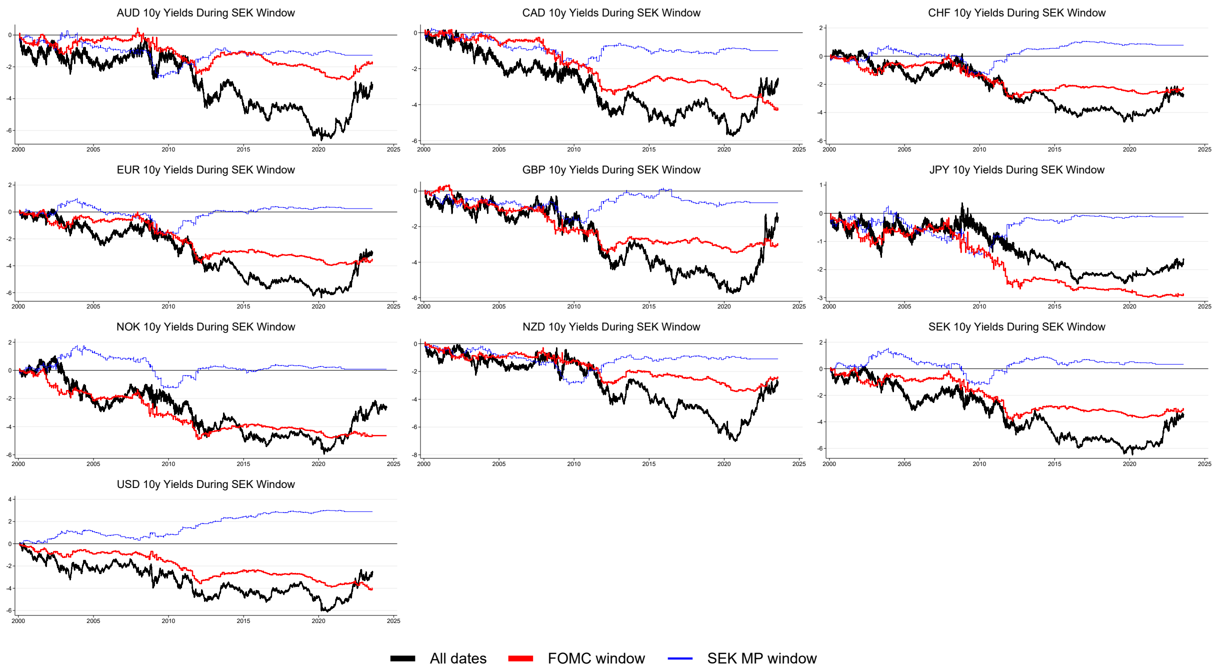

Figure A11: Sum of daily yield changes during U.S. or Swedish central bank announcement windows.

Table A11: Cointegration tests: PC1 and cumulative yield changes during FOMC announcement windows.

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
I: $y_t = \gamma_0 + \gamma_1 \nabla PC1^{US}_{loc,t} + u_t$										
constant	-3.23	4.88	-1.95	-5.23	-9.76	-1.20	0.44	-6.18	-4.46	-1.91
	(0.90)	(0.61)	(0.26)	(0.28)	(0.46)	(0.23)	(0.15)	(0.72)	(0.44)	(0.23)
$\nabla PC1^{US}_{loc,t}$	1.35	0.06	1.22	1.34	1.83	0.64	0.97	2.24	1.37	1.33
	(0.15)	(0.02)	(0.08)	(0.05)	(0.07)	(0.09)	(0.05)	(0.15)	(0.10)	(0.05)
R^2	0.73	0.20	0.88	0.94	0.93	0.45	0.93	0.80	0.84	0.93
SD	1.08	8.09	0.67	0.75	0.74	0.80	0.78	1.06	1.17	0.48
$\hat{\rho}$	0.96	0.14	0.97	0.96	0.94	0.97	0.96	0.95	0.98	0.91
Half-life	17.8	0.4	21.1	15.5	10.5	21.7	18.9	12.9	31.0	7.1
ADF	-2.98**	-16.64***	-2.58*	-3.40**	-3.38**	-2.87**	-2.58*	-4.10***	-3.54***	-4.67***
PP	-15.82**	-323.14***	-11.73*	-18.20**	-25.50***	-10.44	-11.49*	-21.59***	-9.29	-39.17***
LFST	0.29	0.26	0.55	0.39	0.15	0.09	0.64	0.95	0.38	0.63
II: $y_t = \gamma_0 + \gamma_1 \nabla PC1^{US}_{US,t} + u_t$										
constant	-1.86	-0.74	-3.93	-5.50	-5.37	-1.91	-4.54	-1.64	-5.78	-1.91
	(0.52)	(1.57)	(0.42)	(0.39)	(0.29)	(0.36)	(0.60)	(0.60)	(0.67)	(0.23)
$\nabla PC1^{US}_{US,t}$	1.44	0.91	1.17	1.77	1.99	0.60	1.81	1.49	1.97	1.33
	(0.12)	(0.39)	(0.09)	(0.08)	(0.07)	(0.09)	(0.15)	(0.14)	(0.16)	(0.05)
R^2	0.77	0.10	0.82	0.91	0.94	0.62	0.80	0.76	0.81	0.93
SD	0.99	8.15	0.69	0.82	0.66	0.64	1.01	1.11	1.29	0.48
$\hat{\rho}$	0.96	0.14	0.97	0.96	0.92	0.95	0.97	0.97	0.98	0.91
Half-life	15.1	0.3	19.8	17.1	7.8	14.7	19.9	19.6	30.8	7.1
ADF	-3.44**	-7.94***	-2.59*	-3.31**	-4.36***	-2.98**	-2.61*	-3.62***	-3.05**	-4.67***
PP	-19.30**	-325.56***	-12.68*	-16.42**	-36.29***	-13.96*	-11.44*	-15.49**	-10.63	-39.17***
LFST	0.31	0.18	0.95	0.69	0.88	0.27	0.94	0.24	0.75	0.63
III: $y_t = \gamma_0 + \gamma_1 \nabla PC1^{US}_{loc,t} + \gamma_2 \nabla PC1^{US}_{US,t} + u_t$										
constant	-2.44	2.82	-2.78	-5.29	-6.02	-2.29	1.88	-4.80	-4.58	-1.91
	(0.82)	(6.03)	(0.50)	(0.32)	(0.78)	(0.25)	(0.97)	(1.05)	(0.71)	(0.23)
$\nabla PC1^{US}_{loc,t}$	0.32	0.05	0.62	1.09	0.28	-0.95	1.25	1.48	1.08	1.33
	(0.41)	(0.06)	(0.31)	(0.29)	(0.31)	(0.21)	(0.17)	(0.56)	(0.38)	(0.05)
$\nabla PC1^{US}_{US,t}$	1.15	0.37	0.56	0.33	1.68	1.35	-0.50	0.55	0.38	NA
	(0.41)	(1.13)	(0.28)	(0.41)	(0.36)	(0.18)	(0.33)	(0.44)	(0.58)	NA
R^2	0.77	0.27	0.90	0.94	0.95	0.81	0.94	0.81	0.87	0.93
SD	0.97	8.12	0.66	0.75	0.64	0.58	0.79	1.04	1.18	0.48
$\hat{\rho}$	0.96	0.14	0.97	0.96	0.92	0.93	0.96	0.95	0.98	0.91
Half-life	16.0	0.4	21.3	16.1	7.9	9.4	18.4	14.6	33.0	7.1
ADF	-3.37**	-16.62***	-2.65*	-3.42**	-4.36***	-3.78***	-2.58*	-4.05***	-3.24**	-4.67***
PP	-18.70**	-323.43***	-12.37*	-17.89**	-36.27***	-20.06**	-10.01	-19.68**	-9.28	-39.17***
LFST	0.38	0.18	0.77	0.46	0.76	0.73	0.36	0.80	0.43	0.63

Notes. *p<0.1; **p<0.05; ***p<0.01. The table reports cointegration coefficients, Newey-West standard errors, and stationarity test results for the residual $\hat{u}_t$ for the cointegration regression $y_t = \gamma_0 + \boldsymbol{\gamma}_1^\top \boldsymbol{\tau}_t + u_t$. y_t denotes the level factor of a country's yield curve. $\nabla PC1^{US}_{loc,t}, \nabla PC1^{US}_{US,t}$ denote the first principal component of cumulative changes in the local or U.S. yields during the U.S. central bank announcement windows. *SD*: the standard deviation of $\hat{u}_t$; $\hat{\rho}$: the AR(1) coefficient of $\hat{u}_t$; Half-life: the half-life of $\hat{u}_t$; ADF: the t-statistic for the augmented Dickey-Fuller test, and the null is that $\hat{u}_t$ contains a unit root; PP: the Phillips-Perron statistic, and the null is that $\hat{u}_t$ contains a unit root; LFST: the p-value for the Müller-Watson low-frequency stationarity test, and the null is that $\hat{u}_t$ is stationary.

Figure A12: Dynamics of the 10-yr risk-neutral yield: OSE vs. FE.

Notes. The FE model is the original Adrian et al. (2013) model, and the OSE model uses $[\nabla PC1^{US}_{loc,t}, \nabla PC1^{US}_{US,t}]^{\top}$ to proxy the persistent state variable in the shifting endpoint model.

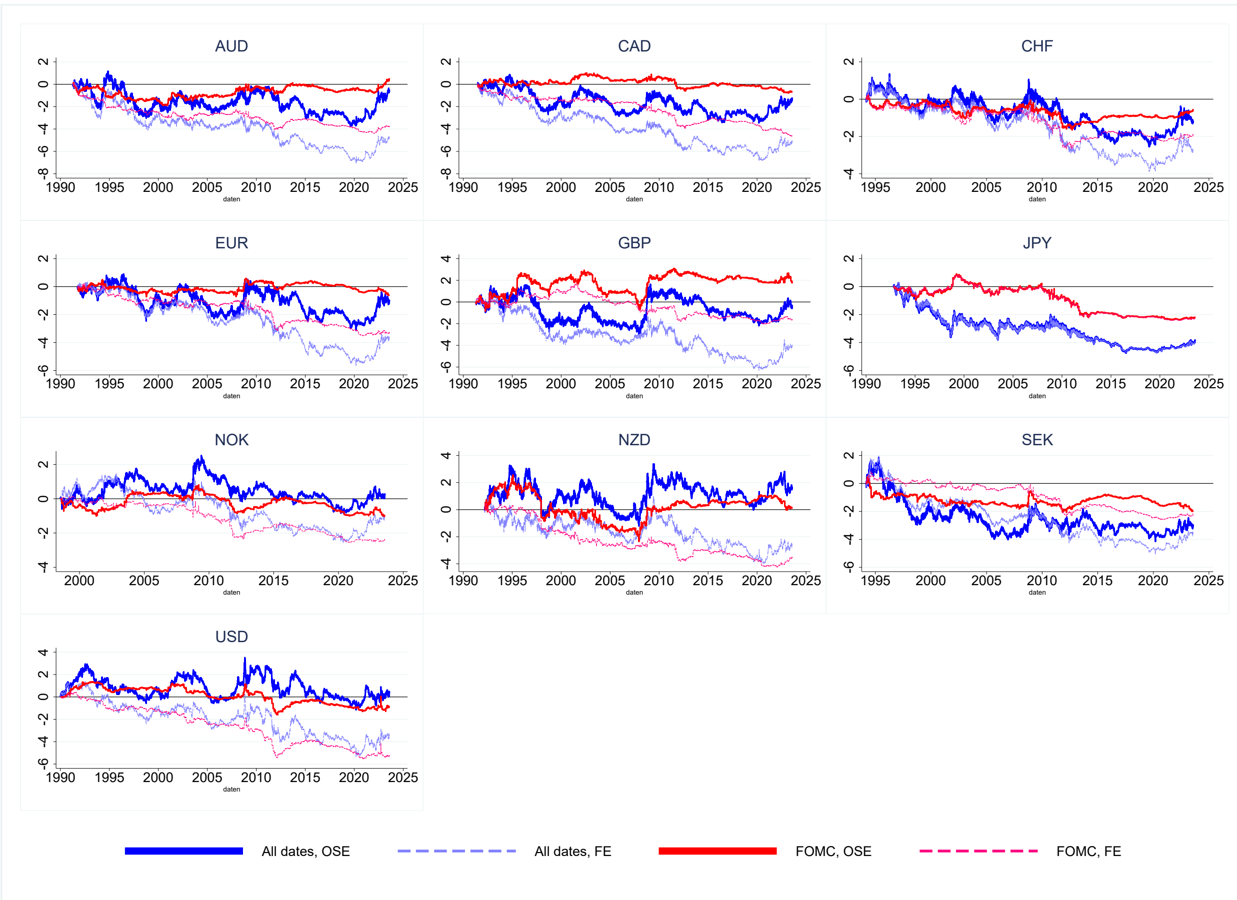

Figure A13: Dynamics of the 10-yr term premium: OSE vs. FE.

Notes. The FE model is the original Adrian et al. (2013) model, and the OSE model uses $[\nabla PC1^{US}_{loc,t}, \nabla PC1^{US}_{US,t}]^{\top}$ to proxy the persistent state variable in the shifting endpoint model.

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