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

Boris Hofmann† Zehao Li‡ Steve Pak Yeung Wu§

March, 2025

Abstract

We demonstrate that the three days surrounding U.S. monetary policy announcements (FOMC windows) account for approximately 70% of the persistent declines in cross-country long-end interest rates—exceeding the influence of domestic central banks. The decline in interest rates during FOMC windows is closely associated with identified monetary policy shocks and FOMC windows that are preceded by Macro News announcements in the US. The information channel, proxied by dot plot announcements and Fed Chair speech days, does not produce a comparable drift in global interest rates. These findings underscore the pivotal role of inference U.S. monetary policy rule in understanding long-end interest rate movements.

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

JEL Codes: E43, E52, F42

*We are grateful to Dun Jia, Peter Hördahl, John Rogers, Phurichai Rungcharoenkitkul, Andreas Schrimpf, Eric Swanson, and Yicheng Wang for helpful discussions. All errors are ours. The views expressed in this paper are those of the authors and not necessarily the views of the BIS. Li is funded by the National Natural Science Foundation of China No. 72173111.

†Bank for International Settlements. Email: boris.hofmann@bis.org.

‡Chinese University of Hong Kong, Shenzhen. Email: lizehao@cuhk.edu.cn.

§University of California, San Diego. Email: stevepywu@gmail.com.

1 Introduction

Over the past few decades, long-term interest rates have displayed a marked secular decline across all major advanced economies. This downward trend is often 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)). At the same time, there is a growing literature examining long-term interest rates through the lens of business cycle fluctuations, highlighting in particular the role of monetary policy. For example, Rey (2015) explores the influence of the global financial cycle on interest rates. 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 secular 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 secular decline in long-term interest rates over the past couple of decades across advanced economies is largely driven by market dynamics in the three-day windows around U.S. Federal Open Market Committee (FOMC) meetings. Specifically, changes in bond yields in FOMC announcement windows account on average for over 70% of the total decline in the 10-year yields of Australia, Canada, the euro area, Norway, New Zealand, Sweden, and the U.S. While our analysis primarily focuses on the 10-year yields of G10 countries, our results are consistent across a broad range of maturities and extend to countries outside the G10 currency group.

Figure 1 illustrates that the global secular decline in interest rates is mainly driven by U.S. monetary policy announcements, which surprisingly explain more of the long-term variation in sovereign yields than monetary policy announcements of domestic central banks. The charts show, as examples, the cumulative daily yield changes of the 10-year Euro (German) and New Zealand government bond yields, representing the largest and smallest non-U.S. economies in our sample, respectively. The figures further display the cumulative yield changes over 3-day U.S. FOMC meeting windows and over the corresponding 3-day domestic central bank meeting 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

[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.

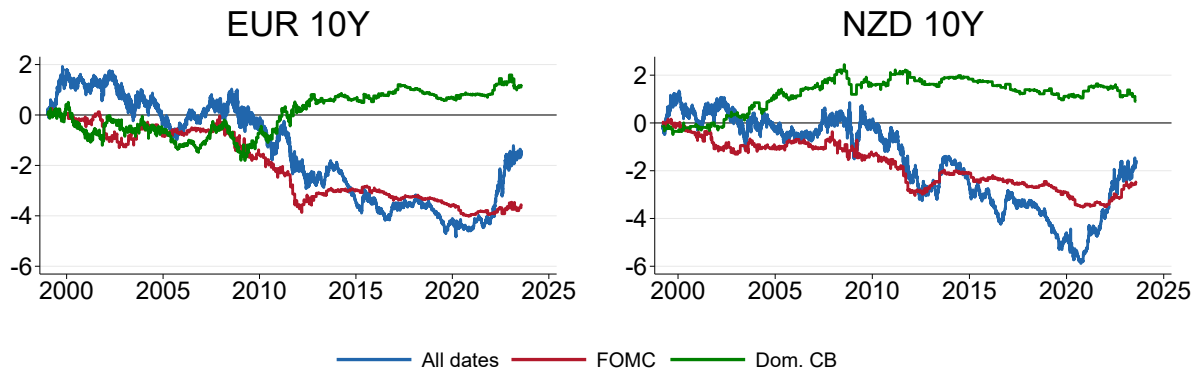


Figure 1: Cumulative changes in nominal 10-year Euro (German) and New Zealand yields during FOMC or domestic monetary policy announcement windows.

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.

overall decline in interest rates; and third, the cumulative changes in yields during domestic central bank monetary policy windows has not contributed to the overall interest rate trend. In summary, the secular decline in long-term interest rates is primarily linked to U.S. monetary policy announcements rather than domestic policy announcements.

In this paper, we conduct a systematic analysis of the explanatory power of monetary policy announcements from the Federal Reserve and other central banks in shaping sovereign yield trends and reveal a significant contrast. Notably, FOMC announcement windows account for the bulk of the observed declines in sovereign interest rates across countries. On average, yield changes in G10 countries are explained by the FOMC windows, domestic central bank windows, and non-monetary policy windows by 64%, 5%, and 31%, respectively. These results are consistent with well-documented international monetary policy spillovers, but they underscore an important distinction: while monetary policy announcements from all central banks generate immediate cross-country effects, only U.S. FOMC announcements are consistently 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 pressure associated with U.S. announcements.

To examine the underlying channels linking global interest rates to FOMC announcements, we conduct a series of decompositions. We begin by assessing whether yield movements during FOMC announcement windows are primarily driven by news about the setting of US policy rates or the market's inferences about the macroeconomic fundamentals and future Fed actions from the monetary policy announcements. Utilizing state-of-the-art high-frequency identified 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 component.

In light of Swanson and Jayawickrema (2023), which highlights the impact of Federal Reserve Chair speeches, we also partition our sample into three-day windows around Fed Chair speech dates and non-speech dates, treating the speech events as another proxy for the information channel. The Fed Chair's speeches provide an interesting opportunity to distinguish the two channels because the 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). Again, the difference emerges when looking at the cumulative sum of daily changes. Although global bond yields fluctuate significantly during Fed Chair speech days, the sums of daily changes 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 introduce short-term volatility, 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 that incorporates 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, which

has driven 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. In addition, we find that a 1% decline in the 10-year real rate in the U.S. during FOMC announcement windows is associated with an average decline of 77 basis points in foreign 10-year real interest rates, with no significant interest rate reversal in the subsequent 10 years. This suggests that the impact of FOMC announcements on global interest rates is long-lasting rather than merely transitory.

Literature review. Since the seminal work by Kuttner (2001), Gürkaynak et al. (2005), Bernanke and Kuttner (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. These shock measures are commonly referred to as "FOMC shocks". Recent influential studies, including Nakamura and Steinsson (2018) and Hanson and Stein (2015), have provided compelling evidence for the non-neutrality of monetary policy. 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 examine the monetary policy announcement effect across multiple central banks. Second, we show that FOMC announcement windows explain a significant portion of the global secular trend in interest rates, while announcements from other central banks do not produce comparable effects. Third, we offer various decompositions 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 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), which focus on the Federal Reserve. Additionally, Gerko and Rey (2017), Jarociński (2022), Miranda-Agrippino and Nenova (2022), and Kearns et al. (2023) highlight the central roles of the Federal Reserve, the European Central Bank, and the Bank of England. 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 safe asset prices—a phenomenon we term the "global safety trend".

The literature on the secular decline of natural rates highlights 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 present alternative investment opportunities, raising natural rates in advanced economies (Obstfeld (2023)), while the limited supply of safe assets—especially U.S. government bonds—drives up their price and lowers returns (Bernanke (2005), Bárány et al. (2023) Caballero et al. (2008); Del Negro et al. (2017); Krishnamurthy and Vissing-Jorgensen (2012)). Market power also plays a role, as increased concentration can depress future investment demand but redirect dividends from labor to capital owners, with 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 bring in a new monetary explanation to the debate about the drivers of the trend in equilibrium long-term interest rate.

The remainder of the paper is organized as follows. Section 2 describes the sovereign yield data and central bank event dates. Section 3 describes our baseline empirical methods and results. Section 3.2 compares the spillover effects of various central banks on international sovereign yields and contrasts the contemporaneous effects with the cumulative effects. Section 4 shows that the realized U.S. monetary policy actions are more important than the "information channel" in explaining world interest rate trends.

Section 5 decomposes our baseline results into the responses of risk-neutral rates and term premia. Section 6 investigates the cumulative effects of FOMC announcements on real interest rates. Section 7 examines the statistical significance of cumulative effects of FOMC announcements. Section 8 concludes.

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.

2.2 Central Bank Announcements

FOMC Announcement Dates The Federal Open Market Committee (FOMC) makes U.S. monetary policy decisions. Since 1981, the FOMC has typically held eight scheduled meetings per year. Most monetary policy decisions since 1994 have been made during these *scheduled* meetings, while a few were made during *unscheduled* meetings. In contrast, unscheduled meetings accounted for a large fraction of changes in the federal funds rate before 1994. Some of the unscheduled meetings were not followed by immediate policy actions or a statement. The public learned about these meetings with a significant time lag. These meetings

Table 1: Sources of daily sovereign yields.

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

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

are excluded from the list of FOMC announcement dates.

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

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

Other Central Bank Monetary Policy Announcements Whenever possible, we collect monetary policy announcement dates from the respective central bank's website. We supplement the calendar with records from `Investing.com` and verify that the latter coincides with the official central bank calendar when

the two samples overlap. Most of the monetary policy announcement samples start in 2000, and the latest starting date is September 16, 2004, for the Swiss National Bank. According to Albagli et al. (2019), the non-U.S. monetary policy announcement dates seldom coincide with FOMC monetary policy announcements.

3 Interest Rate Dynamics and Monetary Policy Announcements

We examine the dynamics of the 10-year government bond yield 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 the methodologies of Hanson and Stein (2015) and Hillenbrand (2025), we define 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 log 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 countries.[2] 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, with the U.K. yield experiencing the largest decrease at 8.24%. These trends align with a substantial body of literature documenting the secular decline of long-term interest rates and the natural interest rate, r^*.

The striking findings emerge from the decomposition between FOMC windows and non-FOMC windows. 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 decline and that within FOMC windows converge at around negative five percentage points. More interestingly, this 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 various countries is closely associated with the declines occurring solely 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% by the end of the sample. The strong association between the actual yield dynamics and those within FOMC windows is particularly impressive given that FOMC meetings occur only eight times a year, meaning the cumulative yield adjustments during FOMC windows only happen 24 days annually—less than 7% of the trading days in a year.

Quantitatively, we evaluate the model fit using a pseudo R^2 measure, which is reported in each subtitle. The measure is computed as one minus the ratio of the squared fitting error of the FOMC-window series to

[2]These countries represent advanced economies with the most heavily traded government bond and exchange rate markets, including Australia (AUD), Canada (CAD), Switzerland (CHF), Germany (EUR), the United Kingdom (GBP), Japan (JPY), Norway (NOK), New Zealand (NZD), Sweden (SEK), and the U.S. (USD).

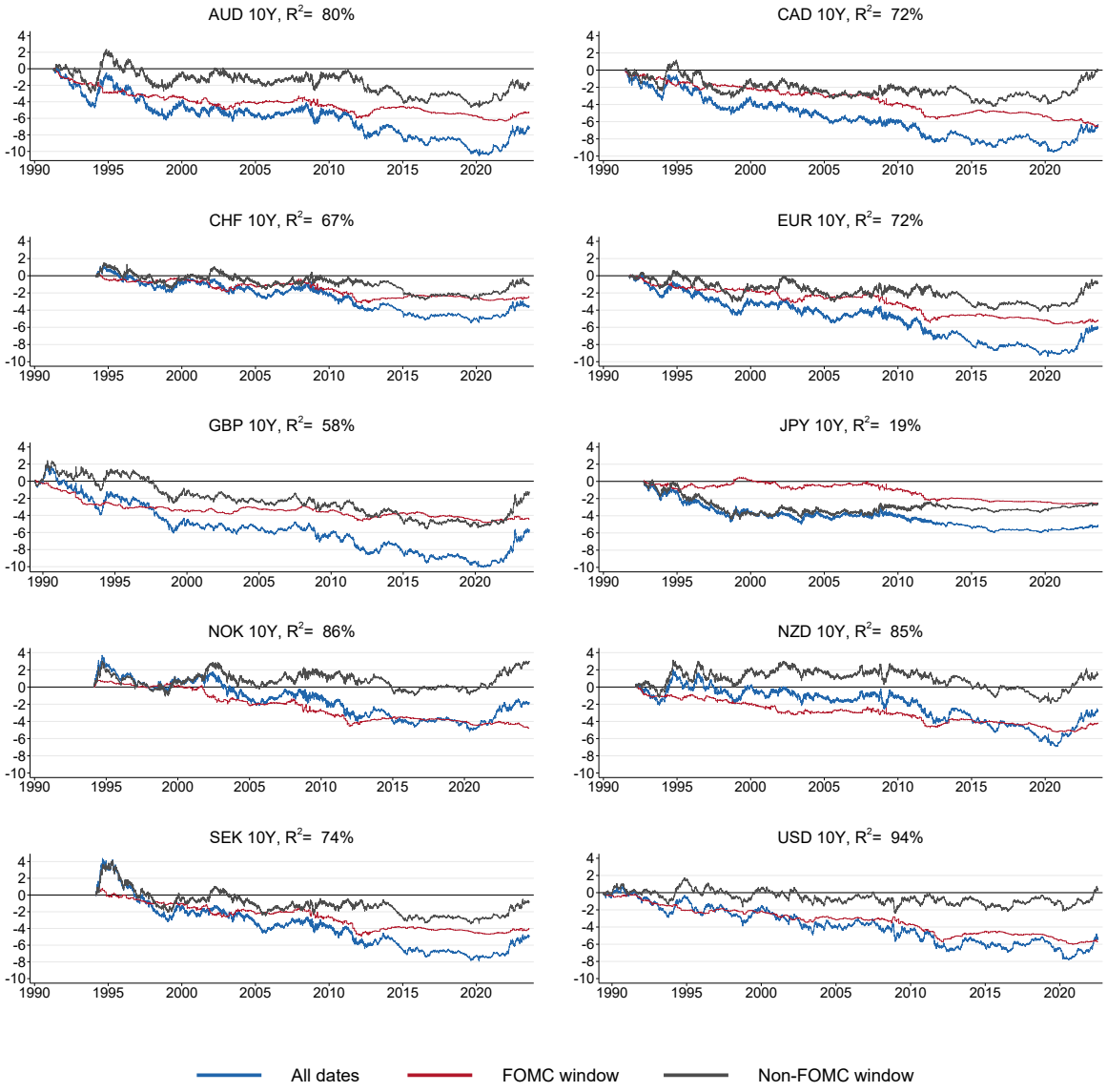


Figure 2: Cumulative changes in nominal 10-year yields during FOMC announcement windows.

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

the sum of the squared fitting 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 as analogous to R^2 and is bounded between zero and one. A larger pseudo R^2 indicates that greater error occurs 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 more effectively.

Overall, the FOMC series fit the actual yields quite well. The most impressive one is the U.S. 10-year yield. The pseudo R^2 indicates that the FOMC windows explain 94% of the total variation in the U.S. 10-year yield. Additionally, the FOMC windows demonstrate impressive explanatory power for yields in Australia, Canada, Switzerland, Germany, Norway, New Zealand, and Sweden, accounting for over two-thirds of the observed total variations.[3]

3.2 Other Central Banks' Monetary Policy Announcements

In the previous section, we have shown that yield changes over the US FOMC windows can explain a significant share of the long-term decline of global interest rates. Is our baseline result a unique feature of U.S. monetary policy, or can other central banks' announcement dates also explain the long-term trends of domestic or foreign interest rates? In this section, we first show that the cross-country yield *long-run decline* is a unique feature of the U.S. FOMC windows. Then, we demonstrate that our findings are not contradictory to the established results in the literature that other central banks' announcements also have significant spillover effects on world interest rates, but these effects are *contemporaneous*.

Do world interest rates also persistently decline during non-U.S. central bank announcement windows? We repeat the exercise described by Equation (2) but replace the FOMC announcement dates with other central banks' announcement dates, and then we investigate whether the announcement windows can explain interest rate trends internationally or at least domestically.

In Table 2, we use the pseudo R^2 and root mean square error (RMSE) to summarize how well the

[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.

cumulative sum of daily yield changes within each central bank's monetary policy announcement windows can fit the observed 10-year yield data.[4] In comparison, we also report the measures for the sums of daily yield changes during the FOMC windows and the remaining days. The RMSE for the cumulative sum of daily yield changes of country i within an event window W is defined as:

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

In the upper panel, 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 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 the majority of the total variation in the observed yield. The contributions of the domestic monetary policy announcements to the cumulative yield changes are universally much smaller than the FOMC windows.

Alternatively, we report the RMSEs associated with the three event windows. A small RMSE indicates that the series is close to the actual yield. For all countries, the FOMC windows generate a smaller RMSE than the domestic monetary policy windows, indicating that the FOMC announcements have performed better in explaining the yield dynamics. In fact, in Germany, New Zealand, and Sweden, the RMSE of the domestic monetary announcement windows is larger than that of the non-monetary policy announcement windows.

In the lower panel of Table 2, we present results for the sample beginning when domestic central bank monetary policy announcement data becomes available. This approach prevents underestimating the domestic central bank's contribution due to a smaller number of event windows. For instance, ECB announcement data are only available after 1999, following the establishment of the ECB. By and large, 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, providing a visualization of the summary table in Table 2. We begin the cumulative sum when domestic monetary policy data becomes available. Similar to what is observed in Figure 2, the

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

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
					After domestic MP announcement data is available				
MP 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.

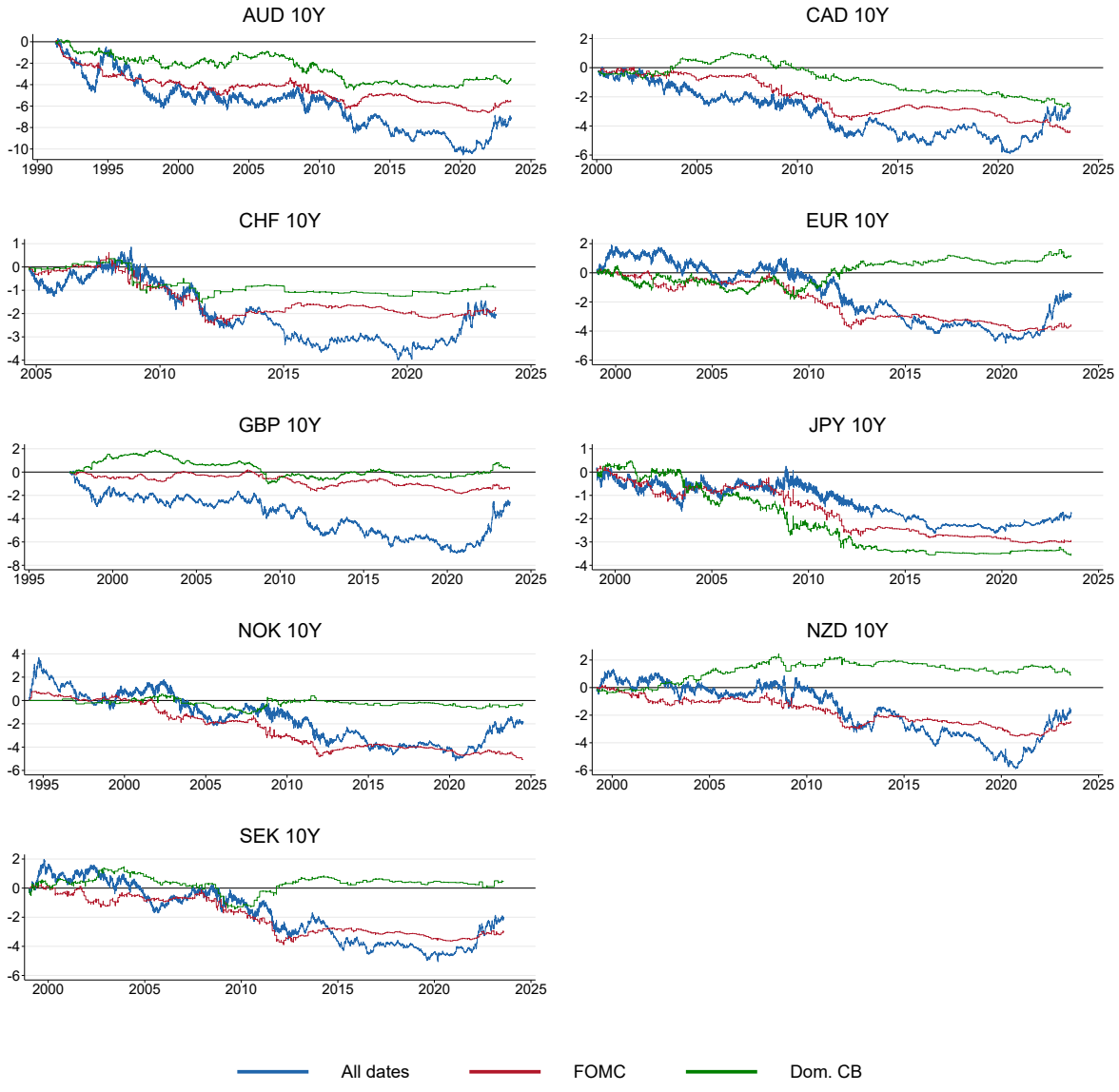


Figure 3: Cumulative changes in nominal 10-year yields during FOMC or domestic CB announcement windows.

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.

overall secular decline in the 10-year yield is accompanied by a consistent downward drift during FOMC windows. This contributes to the relatively higher pseudo R^2 and lower RMSE associated with these windows. Interestingly, a systematic downward drift is not commonly observed on other central bank monetary announcement dates. While there are some downward movements in the cumulative yield changes on the Reserve Bank of Australia, Bank of Canada, and Swiss National Bank's announcement dates, the cumulative yield change remains nearly flat during the Bank of England, Norges Bank, and Sveriges Riksbank announcement windows. More surprisingly, upward trends are observed during the European Central Bank and Reserve Bank of New Zealand's announcement windows.

We summarize the idea of Figure 3 with a panel regression, which is reported in Table 3. We quantitatively examine the systematic differences between FOMC windows, other central bank monetary announcement windows, and other regular days by dummy variables. The estimated regression is:

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

where $\Delta_3 y_{i,t}$ is the 3-day change in the 10-year yield in country i, and $\mathbf{1}_{c,t}$ is an indicator for 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 yield of all countries and all days after 2000. We control for country-fixed effects, and the standard errors are clustered at the event level.

Among all the central banks, only the Fed has a significantly negative effect on cross-country yields during announcement windows, with an additional 0.87 basis point drop relative to non-FOMC window days. Furthermore, we test whether all central banks except the Fed have zero effects on cross-country yields. We find very weak evidence against the null hypothesis (F=1.05 with a p-value of 0.39).

Overall, the main message of this section is that 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 a similarly persistent effect on domestic or other countries' long-term interest rates.

What explains the dominant role of the FOMC window in global long-term interest rates trends? Mechanically, it results from large and significant interest rate spillovers from the U.S. to other countries (see e.g.

Table 3: Average three-day 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 $\Delta_3 y_{i,t} = \alpha + \sum_c \beta_c \mathbf{1}_{c,t} + u_{i,t}$, where $\Delta_3 y_{i,t}$ is the 3-day change in the 10-year yield in country , and $\mathbf{1}_{c,t}$ is an indicator for whether t belongs to the announcement window of country c. Each column reports $\hat{\beta}_c$ and se($\hat{\beta}_c$). 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.

Obstfeld (2015), Hofmann and Takats (2015), Kearns et al. (2023), Albagli et al. (2019)). 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 effect of the decline of U.S. long-term interest rates during the FOMC windows on other countries long-term interest rates. The findings are summarized in Table 4. We first conduct a panel fixed effect regression of the change of daily 10-year yield change of a foreign country during the FOMC windows on the change in US yields:

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

This regression captures the contemporaneous spillovers of changes in U.S. bond yields during the FOMC window. We observe a significant positive monetary spillover. We find that when the U.S. bond yields decrease by 1 percentage point during FOMC meetings, foreign bond yields decline by an average of 0.88 percentage points.

We then examine the longer-term spillovers. Specifically, for each FOMC meeting at date t, we compute the 10-year yield change from $t+1$ to t+10-year.[5] We regress this 10-year horizon yield change on the daily change of US yield during the FOMC meetings:

$$\Delta_{10y} y_{i,t} = \alpha_i + \beta \Delta y_{i,t} + u_{i,t+10y}, \tag{7}$$

[5]When computing the 10-year horizon yield change, we exclude the yield change that happens during the future FOMC windows.

Table 4: Regression table of foreign country yield change contemporaneously and 10-year from event dates.

	Contemporaneous regression	10 year horizon local projection
Δy_{US}	0.886***	2.50*
	(0.055)	(1.15)
R^2	0.335	0.449
adj. R^2	0.336	0.447
within R^2	0.336	0.01
N	6,447	2,385

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

The premise is that if the monetary policy spillover is temporary, we would not see a persistent effect on foreign interest rates. We find the estimated coefficient to be around one, even if only marginally significant. The long-lasting effects of monetary policy documented here are consistent with findings from Hanson and Stein (2015), who identified a similar pattern using forward data, as well as recent studies by Jordà et al. (2020) and Fukui et al. (2023), which identify monetary policy shocks through a pegged currency restriction.

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.

4 The Underlying Channel: Fed Response to News

Why is the global yield decline so concentrated during FOMC windows? In this section, we present evidence supporting a "Fed response to news" channel. The inference of monetary policy stance has long been a key focus for market practitioners. Asset prices are inherently forward-looking; when a monetary policy decision is made, investors must not only grasp the immediate implications for the macroeconomy but also comprehend the underlying rationale behind the policy and any additional information released alongside the decision. Such insights can significantly influence expectations regarding the future trajectory of interest rates. These inferences are commonly referred to as the "Fed response to news" channel and the "Fed information" channel.

We first present evidence that the cross-country yield decline surrounding FOMC meetings is significantly

concentrated around those meetings preceded by macro news announcements, aligning with the "Fed response to news" perspective articulated by Bauer and Swanson (2023a). Next, we compare and contrast this channel with the Fed communication channel proposed by Hillenbrand (2025), which utilizes dot-plot information and Fed Chair speech insights. Finally, we demonstrate that the yield decline is associated with a measure of the "Fed response to news" component within the monetary policy surprise framework.

4.1 FOMC Dates Preceded by Macro News Announcements

To examine the underlying transmission mechanism, we partition the FOMC windows into two distinct sub-windows: those preceded by macroeconomic news announcements and those that are not. Given the Federal Reserve's dual mandate, we focus on Consumer Price Index (CPI) and Gross Domestic Product (GDP) announcements, as these variables are primary determinants of policy interest rates under the Taylor rule and are widely referenced in the New Keynesian framework. An FOMC window is classified as being preceded by macroeconomic news if a CPI or GDP announcement occurs within the four business days leading up to the FOMC meeting. The four-day cutoff is selected to ensure an equal split of the sample between meetings preceded by macroeconomic news and those that are not.

Figure 4 reports the cumulative sum of daily yield changes across the two sub-windows. The blue line in each subfigure refers to the cumulative change in FOMC windows. The red and green lines are those preceded by macro news and those that are not, respectively. We denote these two windows as macro news FOMC window and non-macro news FOMC window. The pseudo R^2 measuring the percentage of total variations during FOMC windows explained by the macro-FOMC windows is reported in the subfigure title.

The figure indicates that the downward yield trend during FOMC windows is heavily concentrated on those preceded by macro news. In the USD case, the yield decline in the macro news FOMC window tracks the yield decline almost perfectly from the start date to the end date with a pseudo R^2 of 99%. For all other countries, The macro news FOMC window explains, on average, about 80% of the variation. Note that the macro news and non-macro news FOMC windows equally split the sample. So, as a benchmark, the pseudo R^2 should be 50% if the yield decline is evenly distributed. The fact that all pseudo R^2 values exceed this neutral benchmark suggests that how the Fed responds to macroeconomic news plays a crucial role in shaping yield movements around FOMC meetings.

Interestingly, the macro news windows themselves do not explain much of the observed yield dynamics. In

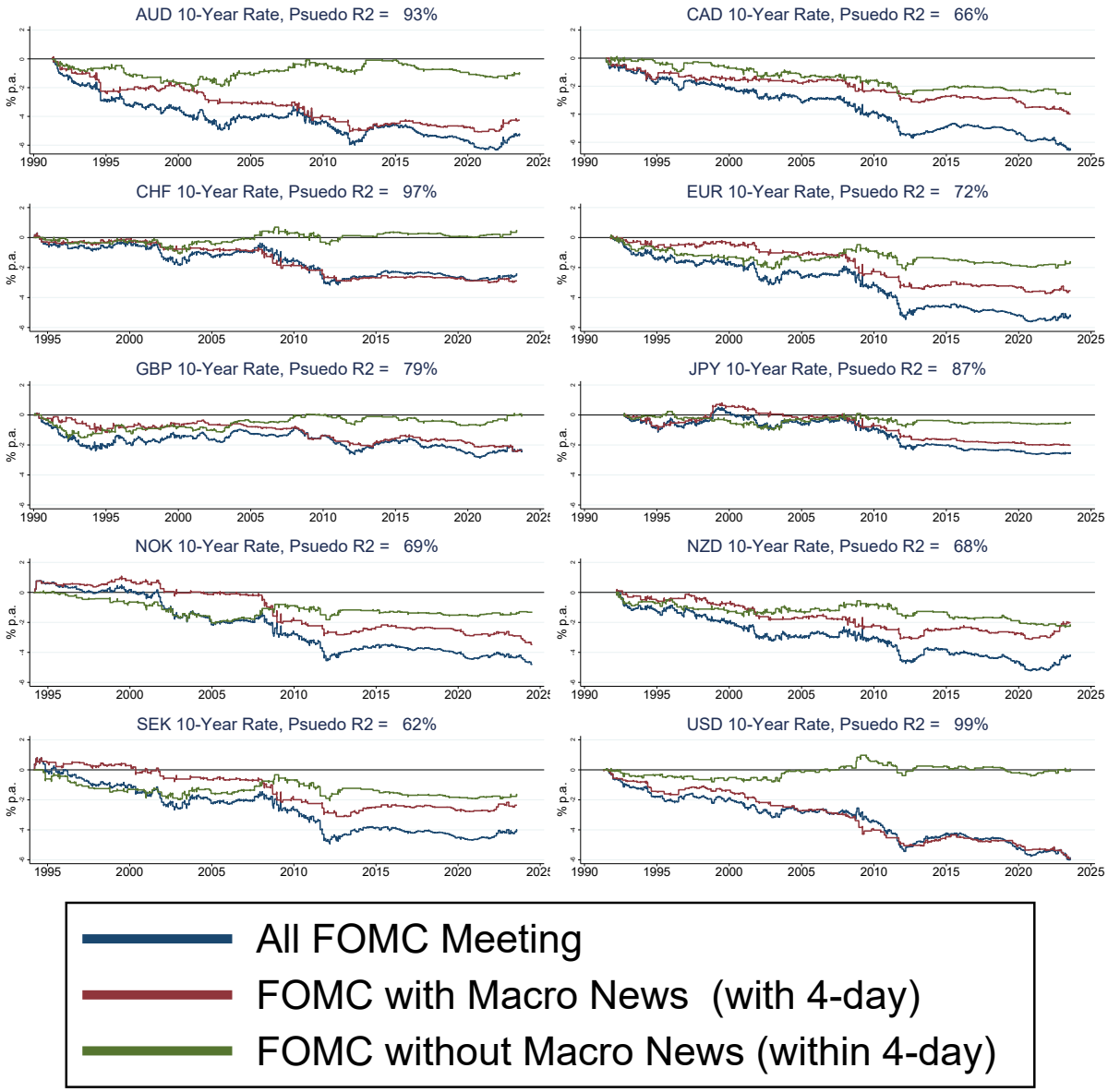


Figure 4: Cumulative changes in nominal 10-year yields during FOMC macro news windows.

Notes. Unit of the y-axis is percentage p.a.. The pseudo R^2 measures the percent variations in the "All FOMC Meeting" explained by the "FOMC with Macro News" series.

Figure 5, we compare the yield dynamics during FOMC announcement windows and U.S. news announcement windows, where the latter is defined as from the day before to the day after U.S. GDP or CPI announcement days. The macro news announcement windows constitute 20% of all trading days, twice as frequent as the FOMC announcement windows. Nevertheless, the macro news windows explain less yield dynamics than the FOMC windows. In some countries, such as Switzerland, the Eurozone, Japan, New Zealand, and Sweden, the 10-year yield increased during many of the macro news windows.

In summary, the FOMC meetings immediately preceded by macroeconomic data releases are mainly responsible for the global downward interest rate trends. Macro data releases alone do not lead to downward interest rate trends. This does not necessarily imply that the financial market fails to internalize macroeconomic news. Instead, there seems to be a "wait and see" effect, whereby the market internalizes the macro shocks only after observing the Fed's response to the news.

4.2 Macro News and the Dot Plot

In this subsection, we compare the macro news FOMC window with the Fed information effect (or Fed communication effect) as suggested by Hillenbrand (2025). We show that the Fed information effect is unlikely to explain a significant share of the secular decline observed.

We measure the Fed information effect using two events. First, we follow Hillenbrand (2025) to make use of the "dot plot" from the Summary of Economic Projections. Second, we look at an alternative set of windows when the Fed chair delivers speeches to the public - the Fed chair speech window.

As part of the efforts by the Fed to improve policy transparency, the Summary of Economic Projections and the dot plot were introduced in 2012. Since then, every quarter-end FOMC meeting announcement comes with economic projections and, more importantly, the interest rate projections of the FOMC members, which is known as the dot plot. Because the FOMC meets eight times a year, 50% of the FOMC meetings post-2012 come with the dot plot release.

We evaluate the Fed information effect using the dot plot in Figure 6. The figure reports the cumulative yield change during FOMC windows, the cumulative yield change during dot plot FOMC windows, and the cumulative yield change during macro news FOMC windows.

Figure 6 shows three messages. First, the blue and the red lines show that the secular yield decline has been evident since the sample start date in the early 1990s, and the macro news FOMC window tracks it

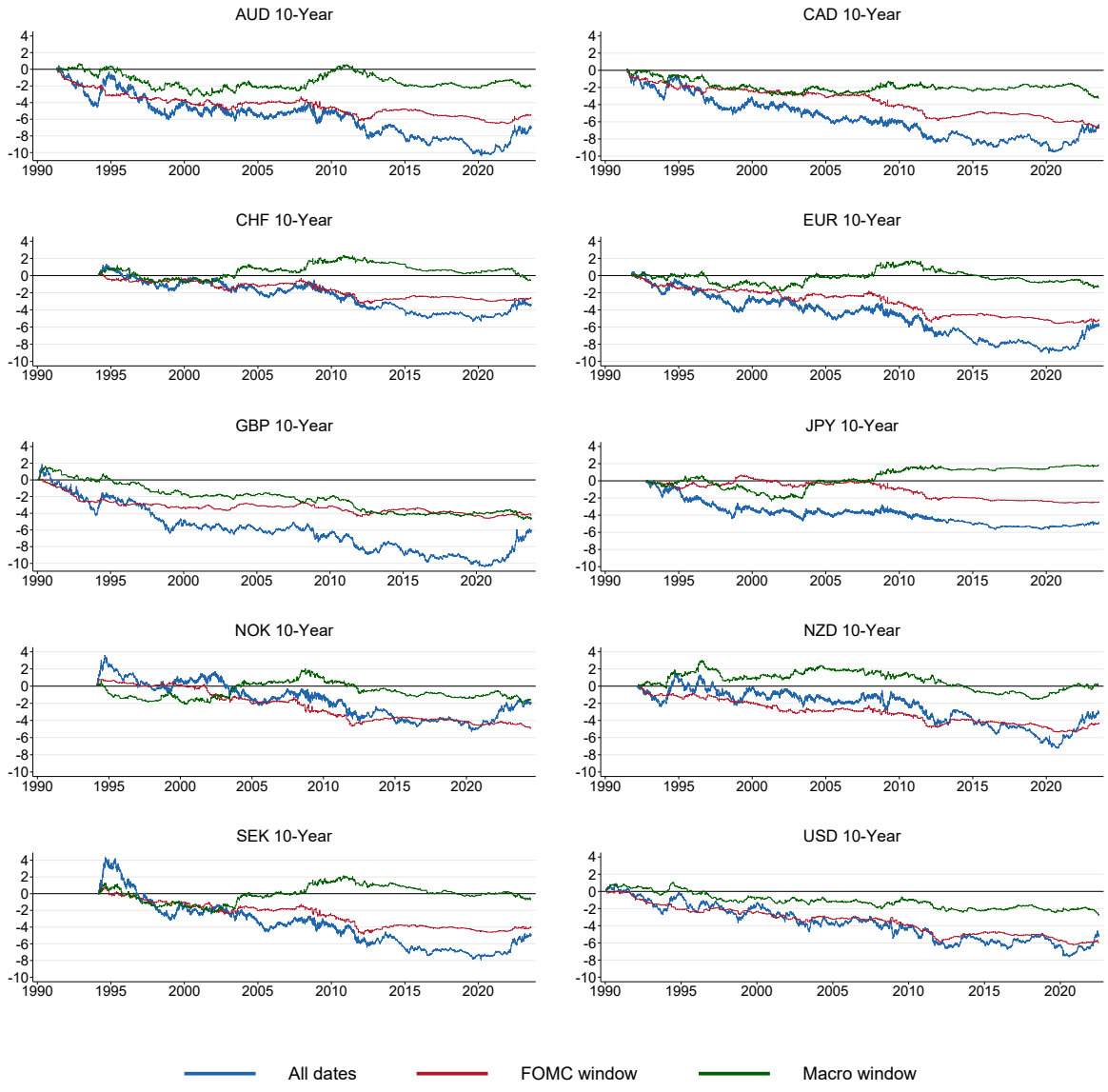


Figure 5: news announcement windows.

Notes. The macro news announcement windows consist of $t-1, t, t+1$ of U.S. GDP or CPI announcements. The figure plots the sums of daily yield change over the full sample, three-day FOMC announcement windows, and three-day macro announcement windows. The macro announcements are the releases of U.S. GDP and CPI data.

well. For most countries, there were about 4 percentage points decline in yield observed in the twenty-year sample before 2012. Second, the yield decline during the FOMC windows from 2012 onwards is relatively mild, which is about -0.5 percentage points on average, and the cumulative decline during the dot plot FOMC window (green lines) features a slight upward trend for six out of the ten countries in our sample. Finally, the pseudo R^2 that measures how well the yield changes are explained by the dot plot FOMC windows shows that the fit is very poor, ranging from zero to 2%, when the number of dot plot FOMC windows account for about 16% of the total FOMC windows in the sample since the 1990s.

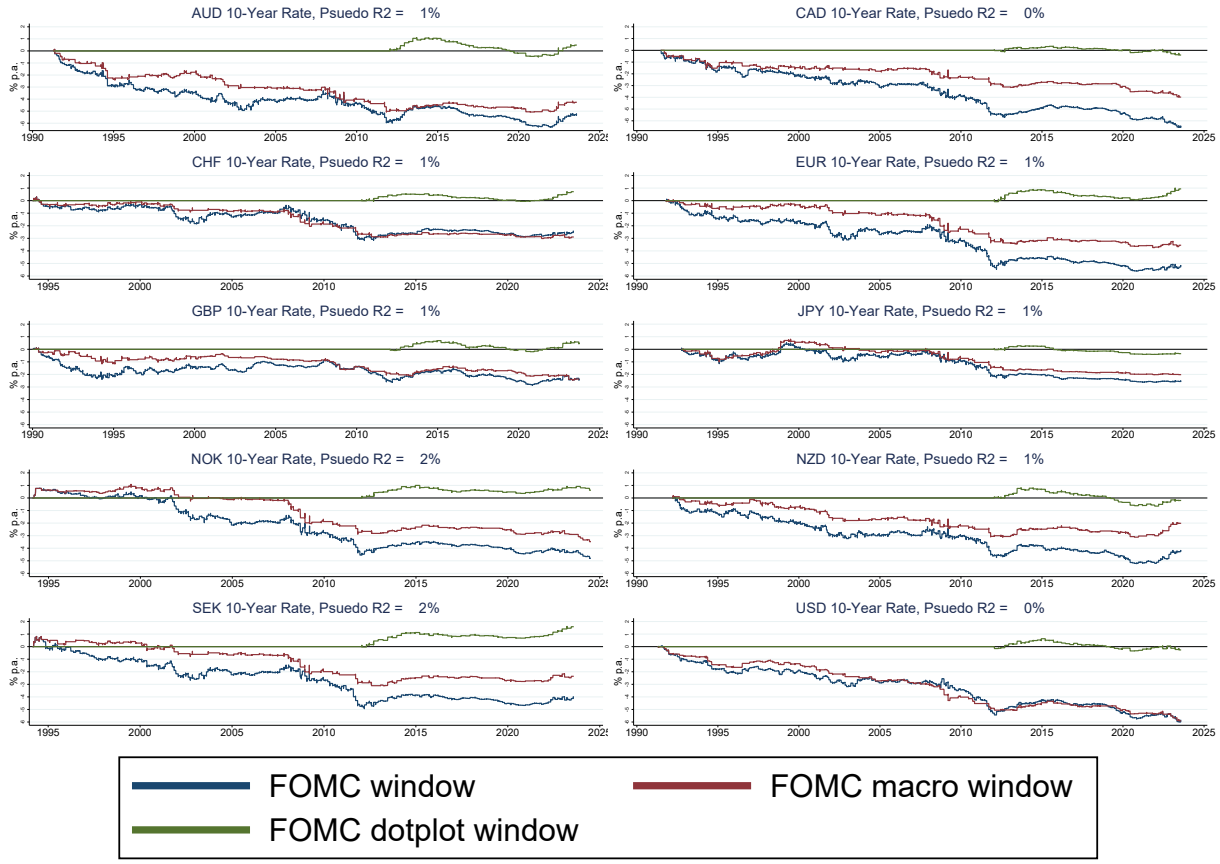


Figure 6: Cumulative changes in nominal 10-year yields during FOMC macro news windows.

Notes. Unit of the y-axis is percentage p.a.. The pseudo R^2 measures the percent variations in the "FOMC window" explained by the "FOMC dot plot window" series.

4.3 Do Actions Speak Louder Than Words? Fed Chair speech windows

As documented by Swanson and Jayawickrema (2023), the volatility of asset prices, 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. In this exercise, we interpret Fed Chair speeches as a relatively pure information channel since there is no obvious policy action taken during the speeches.

We cumulatively sum the daily yield changes during Fed Chair speech windows and non-Fed Chair speech windows, comparing them with the actual yield dynamics. The Fed Chair speech dates are sourced from the Federal Reserve Board of Governors' website, with documentation beginning on June 13, 1996.[6]

Figure 7 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 greater number of Chair speeches, they are disproportionately less associated with overall yield dynamics than FOMC announcements. For all the countries' yields, the cumulative declines over the Fed Chair window do not exceed 2%. 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 rates does not imply that these speeches have no impact on asset prices. As highlighted by Swanson and Jayawickrema (2023), Fed Chair speeches significantly influence various asset classes in the U.S., including futures, equities, and the entire term structure of government bond yields. In Table 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. All countries' government bond rate changes during the FOMC window are more volatile than those on

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

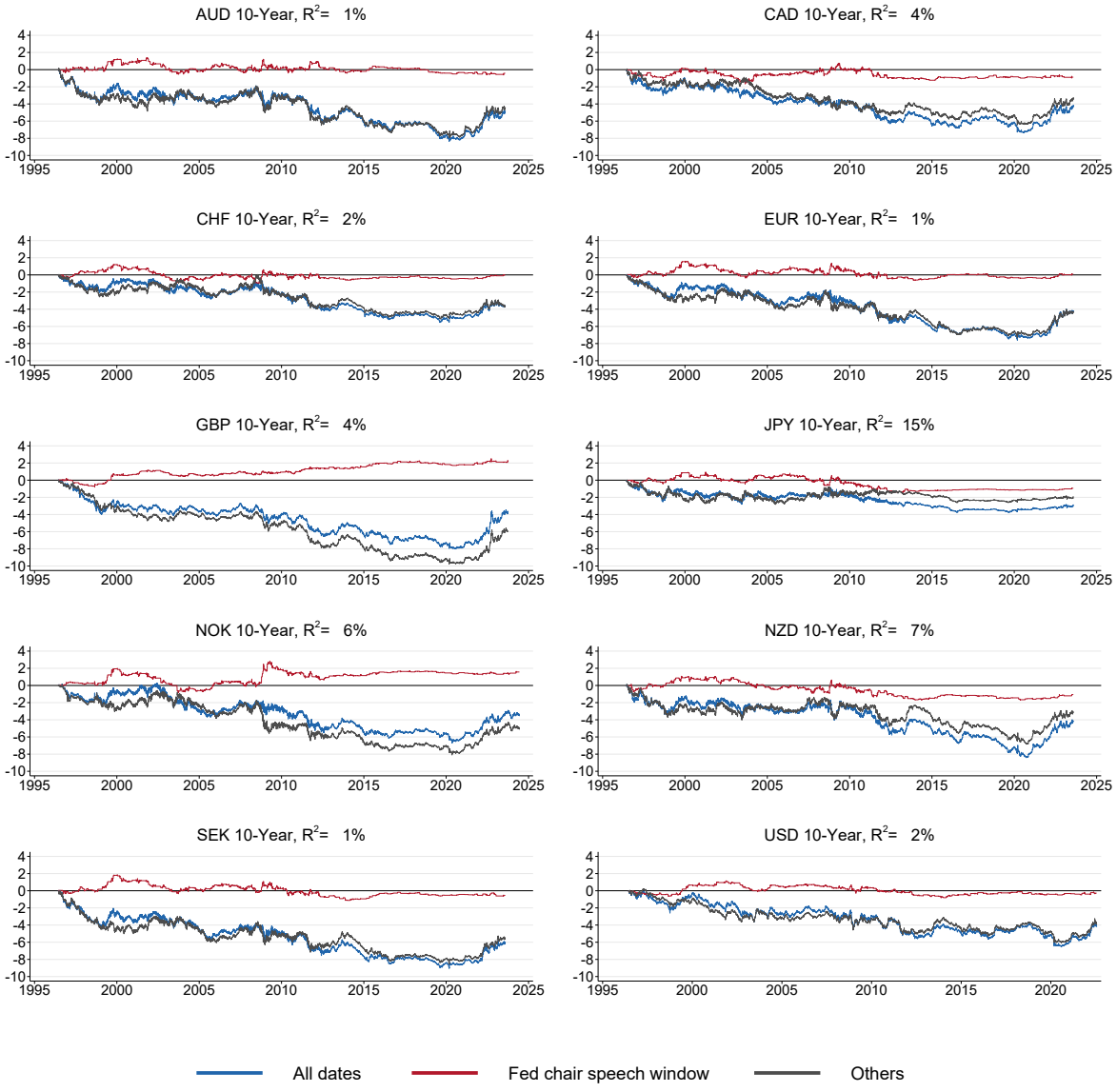


Figure 7: Cumulative 10-year yield changes during Fed Chair speech windows.

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.

normal days. Additionally, yield changes in eight out of ten countries during the Fed Chair speech window are more volatile than those during other days. Overall, the findings of Swanson and Jayawickrema (2023) for the U.S. financial market can be generalized to G10 countries, but the reactions to Fed Chair speeches are much less monotonic than those to FOMC announcements, thus failing to explain the secular decline of government bond rates.

Interestingly, Swanson and Jayawickrema (2023) demonstrate that FOMC announcements have larger impacts on short-maturity interest rates than Chair speeches, while the latter is more important for long-maturity interest rates, measured by contemporaneous changes in the 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 7 seems to contradict Swanson and Jayawickrema (2023). However, the reason for the contradiction is that we study the 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 out each other in the summation.

4.4 Measuring Fed Response to News Using High Frequency Surprises

Having established that much of the cumulative yield change variation is concentrated in the FOMC windows that are preceded by macro news announcements, in this subsection, we make use of a measure from the high-frequency identification literature to directly connect the macro news windows with the "Fed response to news" channel.

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

$$MPS_t = \gamma_0 + \gamma_1 X_{t-} + u_t, \tag{8}$$

where X_{t-} is a set of macro variables known before the announcement date t. Bauer and Swanson (2023a) estimate this equation and categorize the error term u_t as the orthogonalized monetary policy shock, MPS_t^{orth}. For our purpose, we measure the "Fed responses to news" component as the predictable component of MPS.[7] In other words, we define

$$MPS_t^{Response} \equiv MPS_t - MPS_t^{orth}. \tag{9}$$

In Table 6, we first verify the relevance of the macro news FOMC window using a contemporaneous regression. We regress the 3-day yield change of the 10-year government bond rate on both $MPS_t^{Response}$ and MPS_t^{orth}. We estimate this regression separately on the macro news FOMC windows (if there is a CPI or GDP announcement within 4 days before the FOMC) and the non-macro news FOMC windows. The table summarizes the results for both the U.S. yield change and the rest of the G10 countries' yield change with a panel fixed effect regression. $MPS_t^{Response}$ is a significant predictor of the 3-day yield change during the macro news FOMC windows but not the non-macro news FOMC windows. On the other hand,

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

Table 6: Contemporaneous effects of monetary policy shocks.

	G9		U.S.	
	Macro	No Macro	Macro	No Macro
MPS^{orth}	-0.2	0.3***	0.2	0.8***
	(0.2)	(0.1)	(0.2)	(0.2)
$MPS^{Response}$	1.1**	-0.2	1.5***	-0.1
	(0.5)	(0.5)	(0.4)	(0.5)
Const.	-0.0	-0.0	-0.0*	-0.0
	(0.0)	(0.0)	(0.0)	(0.0)
R^2	0.0	0.0	0.1	0.1
R^2_within	0.0	0.0		
N	1134	1034	134	115
* $p < 0.1$	** $p < 0.05$	*** $p < 0.01$		

Notes. The table reports results for $\Delta_3 y_{i,t} = \alpha + \beta_1 MPS_t^{orth} + \beta_2 MPS_t^{Response} + u_{i,t}$. $\Delta_3 y_{i,t}$ denotes three-day changes in country i's 10-year yield over the FOMC window. For the "G9" columns, the equation is estimated for the panel of non-U.S. G9 countries with country fixed effects. For the "U.S." columns, the equation uses the U.S. yield only. "Macro" and "No Macro" denote whether the regression is estimated using FOMC windows preceded by U.S. GDP or CPI announcements.

the MPS_t^{ORTH} is significant during the non-macro news FOMC windows but not during the macro news FOMC windows. The regression results suggest that the global yield decline during the macro FOMC windows is precisely driven by the Fed's response to the macroeconomic news. Note that $MPS_t^{Response}$ is part of the 30-minute innovations in the interest rate futures, so it is consistent with the interpretation that the financial market does not internalize the macro news until the Fed responds to it.

In Figure 8, we estimate the following equation to establish the relationship between the FOMC yield decline and the cumulative MPS:

$$\nabla y_t^{macronewsFOMC} = \gamma_0 + \gamma_1 CMPS_t^{ORTH} + \gamma_2 CMPS_t^{Response} + u_t, \tag{10}$$

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

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

We fit the equation country by country. In Figure 8, the cumulative yield change during the macro new FOMC windows is indicated by the red lines. The fitted value of $\gamma_1 CMPS_t^{ORTH}$ and $\gamma_2 CMPS_t^{Response}$ are plotted as the black lines and purple lines, respectively. The figure clearly shows that the fitted value from $CMPS_t^{Response}$ tracks the 10-year yield decline over the macro news window very well. The subfigure title reports the semipartial correlation of the two regressors. It indicates that, in all cases, adding $CMPS_t^{Response}$ to the regressions provides a much better fit to the cumulative 10-year yield decline. On the other hand, adding $CMPS_t^{ORTH}$ does a very little job of changing the fit.

4.5 Fed Response to News and the Associated Decline

While the cumulative 10-year yield change is explained by yield change at the narrow set of FOMC windows preceded by macro news, it doesn't give rise to a secular "decline" automatically. Take a standard Taylor rule as an example:

$$i_t = r_t^* + \gamma^\pi(\pi_t - \pi_t^*) + \gamma^Y(outputgap_t^*) + u_t, \tag{11}$$

where i_t is the nominal policy rate, r_t^* is the natural rate of interest, π and π^* are the inflation rate and the inflation target and u_t is the monetary policy shock. Assume we start from the state where inflation is at the target and zero output gap, then $i_t = r_t^*$. Now if the latest announced inflation is 1% higher than the inflation target and if γ^π is perceived as 1.5, the market would expect the Fed to raise the interest rate by 1.5% (without interest rate smoothing). However, if the Fed turns out to be raising the interest rate only by 1.2%, there will be a monetary policy surprise measured as -0.3%. In the equation above, there are two sources of the monetary policy surprise. The first one is u_t, so the -0.3% surprise is a pure exogenous monetary policy shock. The second one is that the -0.3% surprise reveals the true γ^π is 1.2, and so the

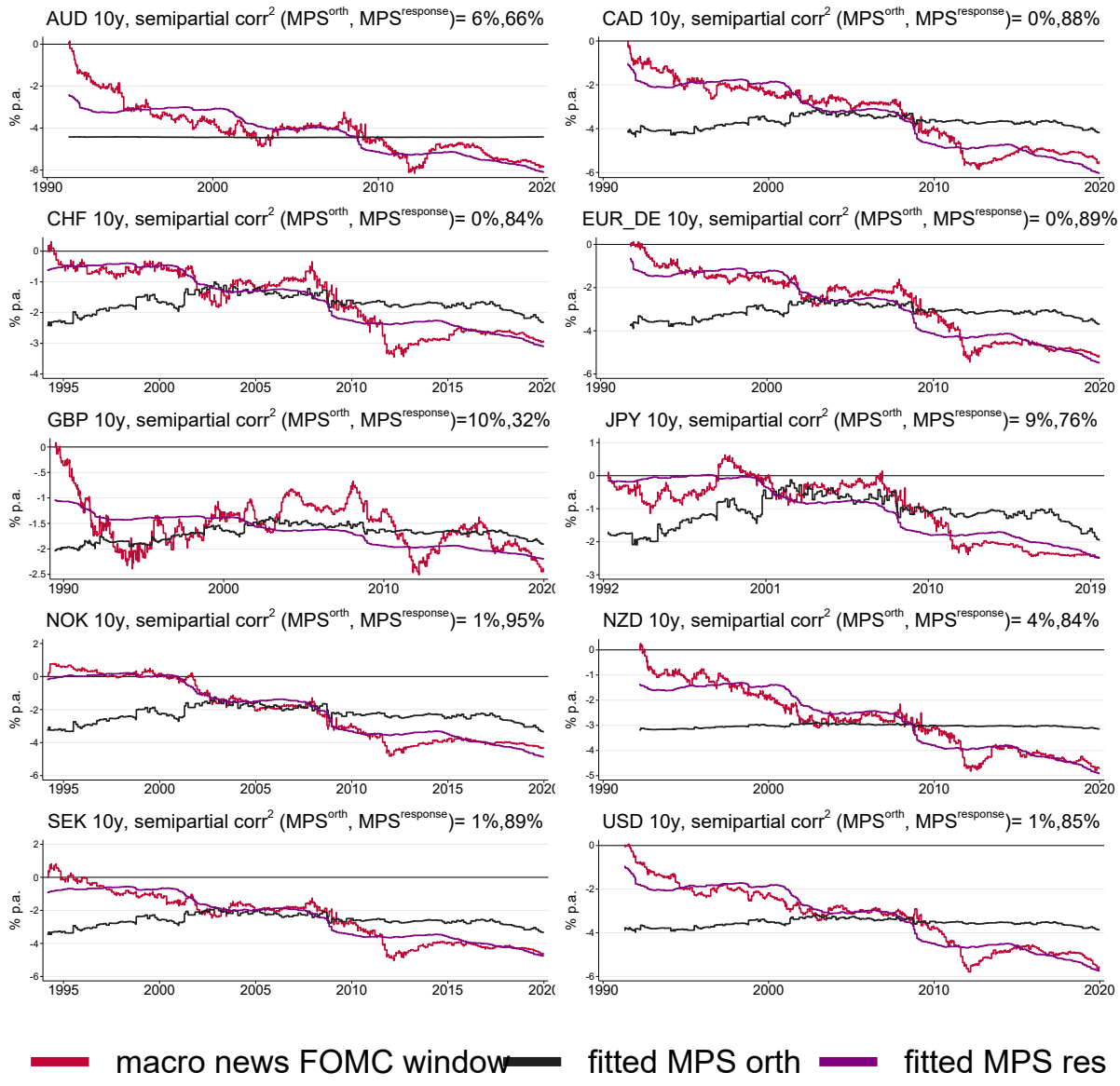


Figure 8: Cumulative changes in nominal 10-year yields during FOMC macro news windows.

Notes. Unit of the y-axis is percentage p.a.. TBD.

market update coefficient accordingly. According to the MPS decomposition in the previous section, the MPS in the macro news FOMC windows is mostly associated with the second one.

Figure 9 presents the cumulative MPS change, along with cumulative $MPS^{Response}$ and the cumulative MPS during macro news FOMC windows. The figure indicates that realized monetary policy shocks over the past three decades have been predominantly negative, regardless of the measure used. This trend helps explain the secular decline in 10-year yields within the Fed's Response to News channel. While the market has generally perceived the Fed as more hawkish, over time, it has gradually adjusted its expectations, recognizing that the Fed has been more dovish than initially anticipated. This interpretation is supported by narratives such as Taylor's 2007 Jackson Hole speech, which highlights a significant deviation from the Taylor Rule between 2002 and 2005, criticizing the Fed for being excessively accommodative. Similarly, the "too low for too long" narrative has been prevalent in discussions of post-2008 monetary policy.

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.[8]

[8] 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.

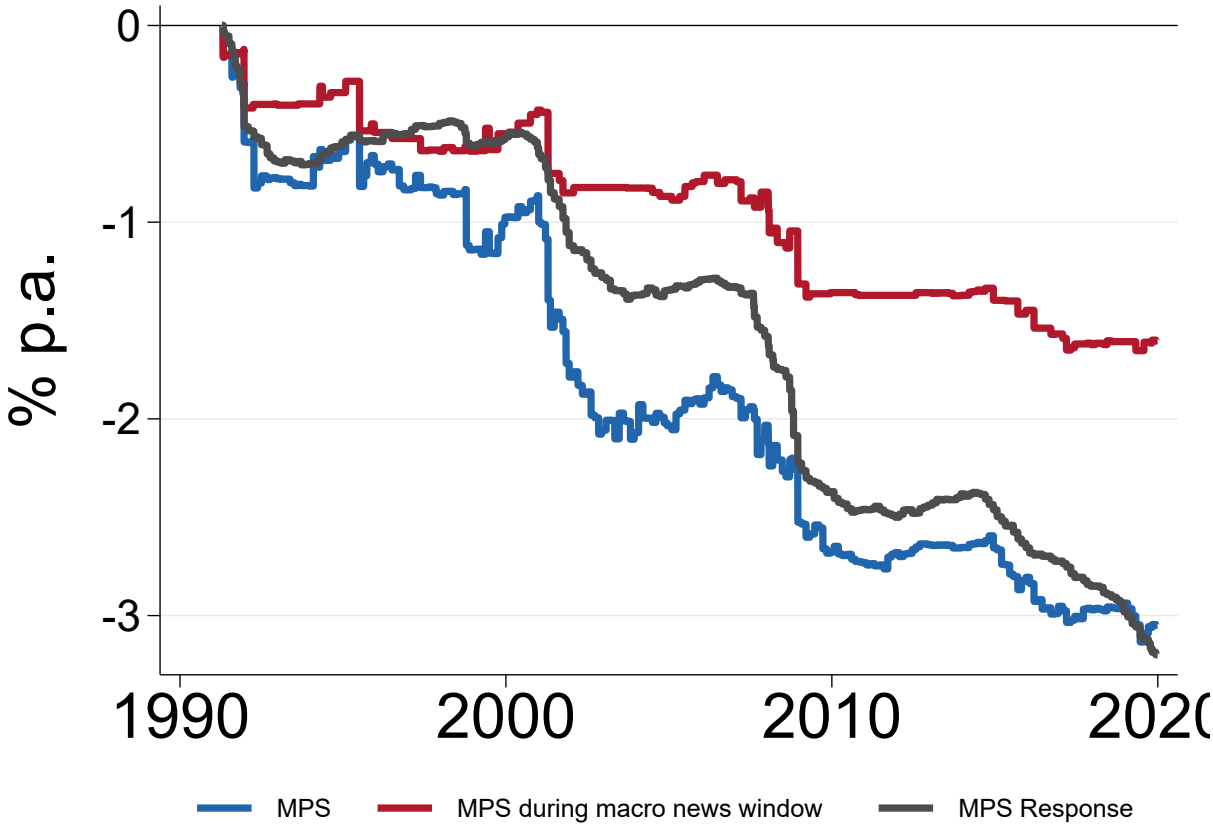


Figure 9: Cumulative changes in nominal 10-year yields during FOMC macro news windows.

Notes. Unit of the y-axis is percentage p.a.. TBD.

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

$$\begin{aligned} X_t =& \boldsymbol{\mu} + \Gamma \boldsymbol{\tau}_t + \tilde{X}_t, \\ \boldsymbol{\tau}_t =& \boldsymbol{\tau}_{t-1} + \boldsymbol{\eta}_t, \quad \boldsymbol{\eta}_t \sim \mathcal{N}(0, \Omega_\eta) \\ \tilde{X}_t =& \Phi \tilde{X}_{t-1} + \tilde{U}_t, \quad \tilde{U}_t \sim \mathcal{N}(0, \tilde{\Omega}), \end{aligned} \tag{12}$$

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

$$m_{t+1} = -\delta_0 - \boldsymbol{\delta}_1^\top X_t - \frac{1}{2} \Lambda_t^\top \Lambda_t - \Lambda_t^\top \Omega^{-\frac{1}{2}} U_{t+1}. \tag{13}$$

The price of risk is an affine function of Z_t:[9]

$$\Lambda_t = \Omega^{-\frac{1}{2}} (\Lambda_0 + \Lambda_1 Z_t). \tag{14}$$

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:

$$y_t^{(n)} = A_n + B_n^\top X_t, \tag{15}$$

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

[9] We assume that Λ_1 satisfies $\Lambda_1 = \begin{bmatrix} (I - \Phi)\Gamma, \Lambda_{12} \end{bmatrix}$, 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

$$p_t^{(n),rn} = \ln \mathbf{E}_t \left[\exp \left\{ -y_t^{(1)} + p_{t+1}^{(n-1),rn} \right\} \right], \tag{16}$$

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

$$p_t^{(n),rn} = \mathcal{A}_n^{rn} + \mathcal{B}_n^{rn\top} X_t + \mathcal{C}_n^{rn\top} \boldsymbol{\tau}_t, \tag{17}$$

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

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.[10] 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

$$y_t = \gamma_0 + \gamma_1 \tau_t + u_t \tag{18}$$

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

[10] 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 7. 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.[11] 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

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

Table 7: Cointegration tests: 10-year yields 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	-3.20	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 10 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

$$R^2 \equiv 1 - \frac{\sum_{t=1}^{T} \left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC} \right)^2}{\sum_{t=1}^{T} \left(\nabla y_t^{FOMC} - \nabla y_t^{TP,FOMC} \right)^2 + \sum_{t=1}^{T} \left(\nabla y_t^{FOMC} - \nabla y_t^{RNY,FOMC} \right)^2}. \tag{19}$$

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.

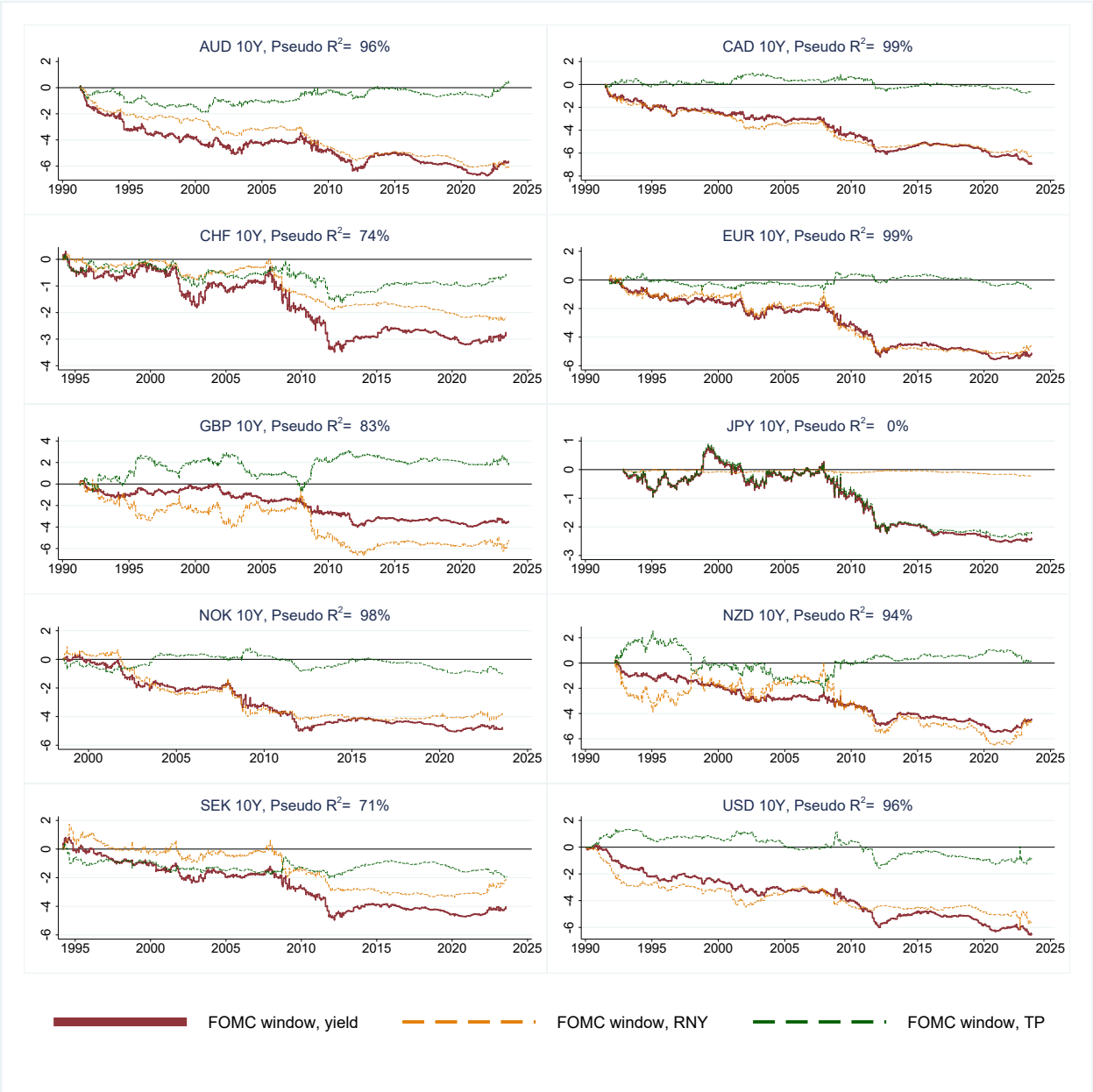


Figure 10: FOMC window: risk neutral yield (RNY) vs. term premia (TP).

Notes. Unit of the y-axis is percentage p.a.. The figure decomposes the FOMC-window cumulative change in the 10-year yield into cumulative changes in the risk-neutral yield (RNY) and the term premium (TP) components. The pseudo R^2 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 Yield and Term Premia Change During/Between FOMC Windows

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 11 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 12 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

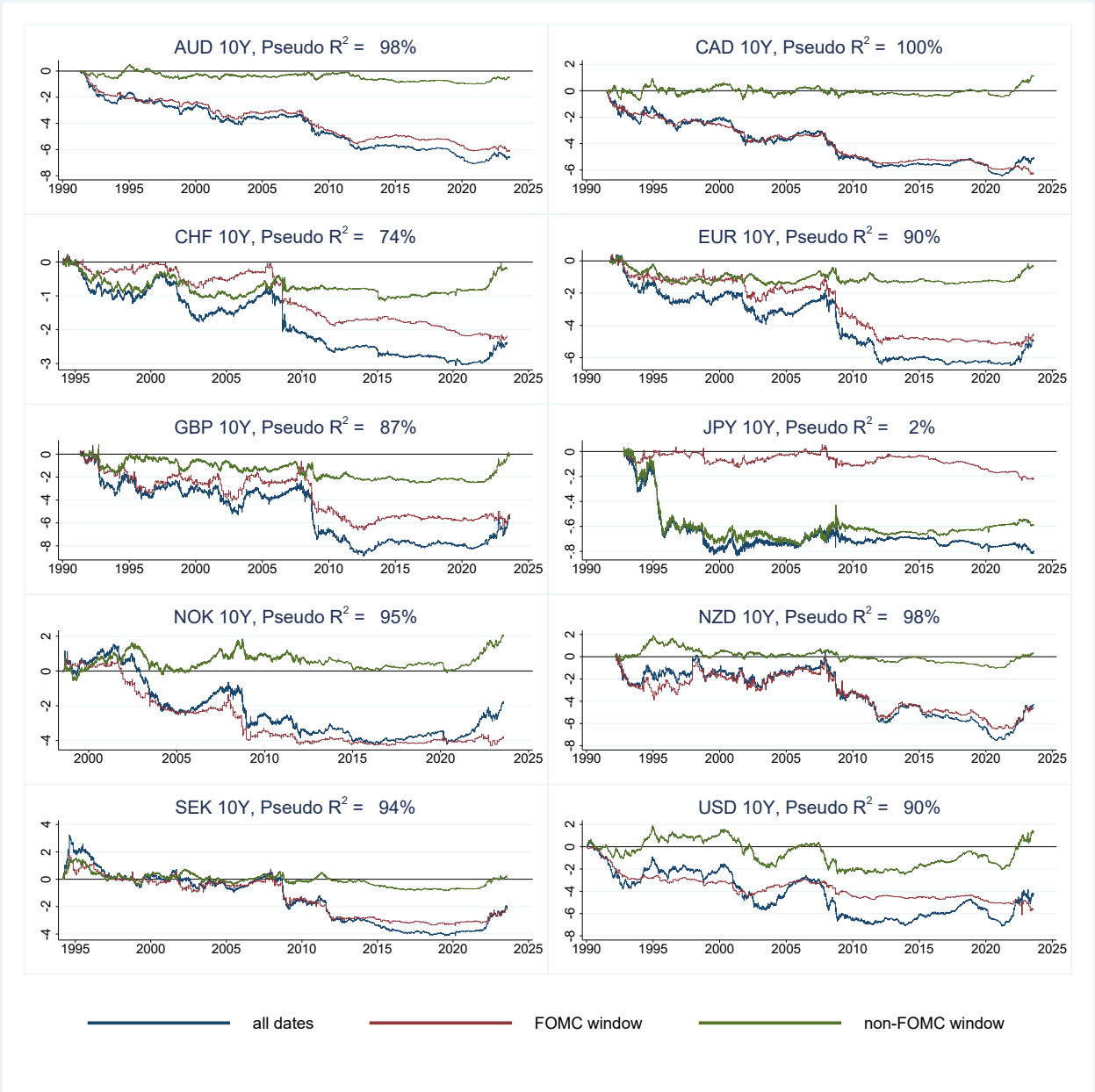


Figure 11: Cumulative changes in the 10-year risk-neutral rates during different event windows.

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

R^2 for Germany suggests that FOMC windows account for 5% of the total variation in the 10-year term premium, while Figure 11 shows that FOMC windows explain 89% of the total variation in the 10-year risk-neutral yield.

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

In this section, we assess whether our findings for nominal long-term interest rates carry over to long-term real interest rates. First, we re-examine our baseline exercise using real interest rates constructed from inflation-protected securities. Our key findings indicate that the declines in long-term interest rates across countries, which are concentrated around FOMC meetings, also extend to real rates. Second, we demonstrate that the interest rate declines during FOMC meetings have a persistent effect; specifically, the real interest rate declines observed during these meetings are not offset by changes in real interest rates over the subsequent 10 years.

To establish a measure of ex-ante inflation expectations and real interest rates, we utilize inflation-protected securities, such as Treasury Inflation-Protected Securities (TIPS) in the U.S. The spread between the nominal government bond rate and the inflation-protected interest rate is often regarded as a market-based inflation expectation, 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. Consequently, we are able to obtain real interest rates only for Australia, Canada, Germany, Sweden, the U.K., and the U.S., excluding Japan, Switzerland, Norway, and New Zealand from our analysis. The TIPS yield data are sourced from Bloomberg.

In Figure 13, 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. For Australia (AUD) and Canada (CAD), some flat regions are observed due to a lack of TIPS data. Overall,

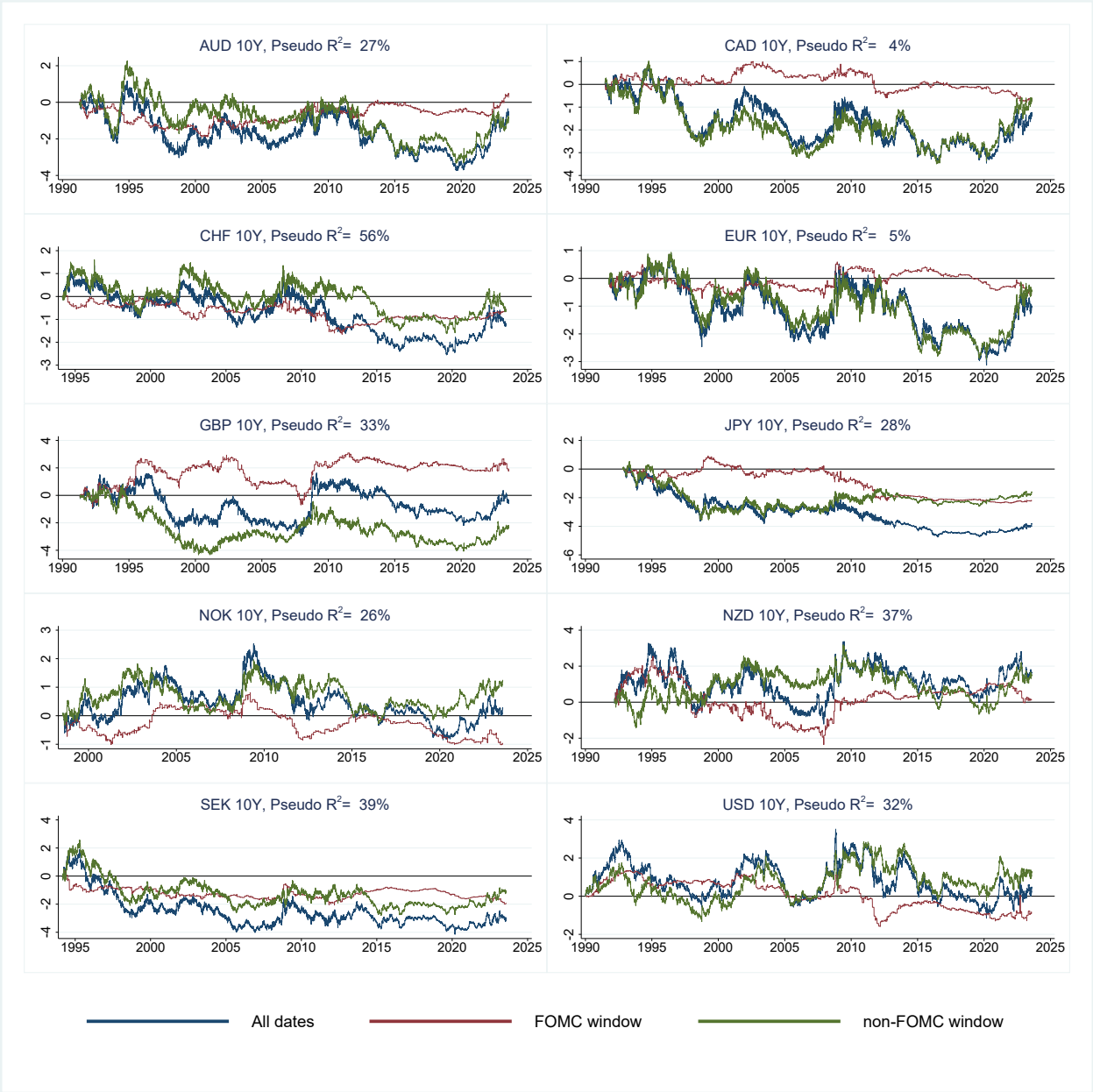


Figure 12: Cumulative changes in the 10-year term premium during FOMC windows.

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

the figure demonstrates that the decline in real interest rates across countries is again heavily concentrated during FOMC windows. This reinforces the evidence presented in Figure 2, indicating that the findings carry over to real rates, albeit over a shorter sample period.

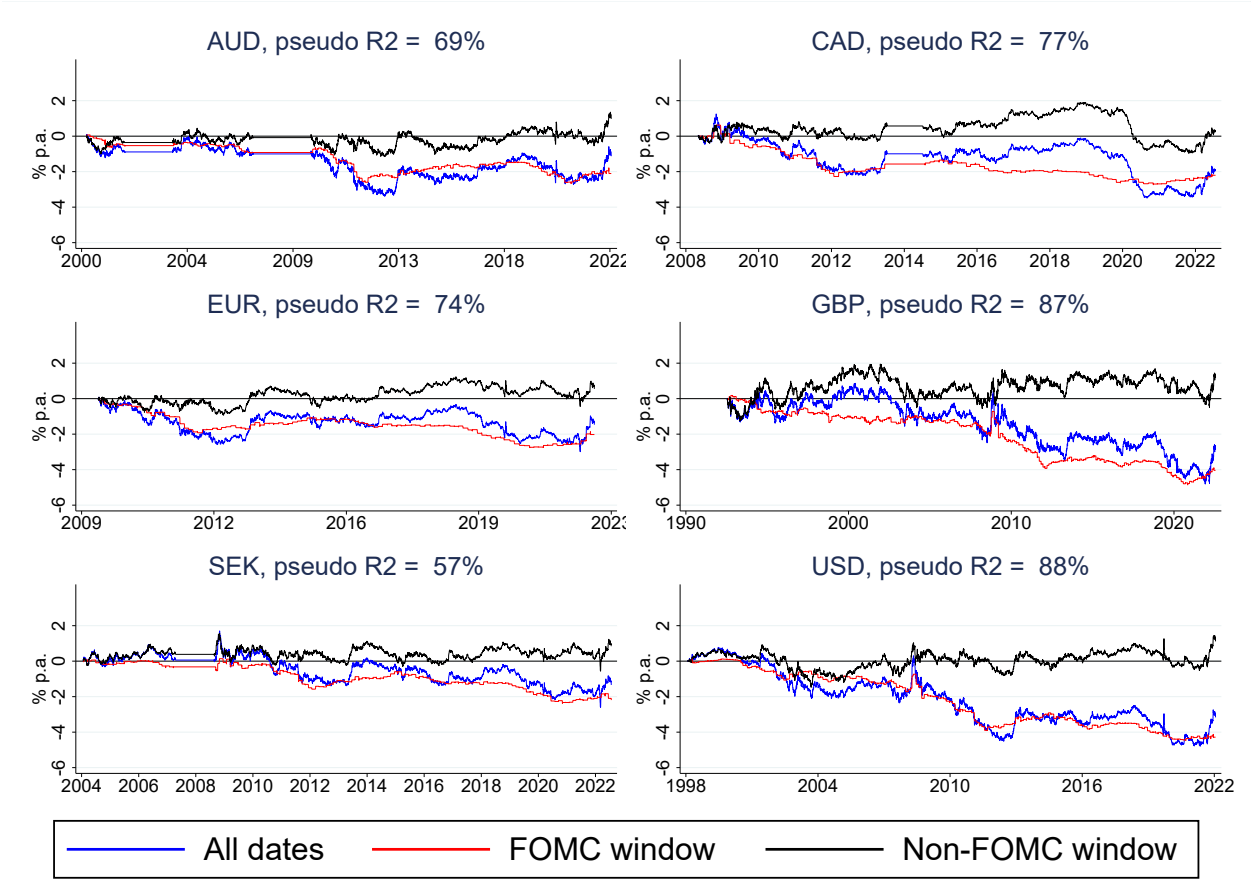


Figure 13: Cumulative changes in real 10-year yields during FOMC announcement windows.

Notes. Unit of the y-axis is percentage p.a.. The real rates is 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 $R^2 = 1 - \frac{\sum_{t=1}^{T}(y_t - \nabla y_t^{FOMC})^2}{\sum_{t=1}^{T}(y_t - \nabla y_t^{nonFOMC})^2 + \sum_{t=1}^{T}(y_t - \nabla y_t^{FOMC})^2}$.

7 Placebo Tests

To investigate the statistical significance of the FOMC-based strategy, we compare the actual returns with those implied by randomly selected hypothetical FOMC windows. Specifically, we 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^{th} and 97.5^{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 15 presents the placebo test result based on 10,000 random samples. The 90% confidence bands denote the 5^{th} and 95^{th} percentiles of the cumulative yield changes during the placebo three-day windows. Except for Japan, the cumulative changes in the 10-year yields during the actual FOMC windows are at or lower than the lower bound of the confidence bands. In other words, the cumulative yield changes during FOMC windows are lower than 95% of the placebo sample paths, indicating that FOMC windows lead to significantly more negative yield changes than randomly selected event windows of the same length.

8 Conclusion

This paper provides a comprehensive analysis of the secular relationship between long-term interest rates and monetary policy announcements, with a particular focus on the effects of U.S. Federal Open Market Committee (FOMC) meetings on global yield dynamics. Our findings underscore the significant influence of U.S. monetary policy on the long-term interest rates of advanced economies, revealing that over 70% of the variation in 10-year government bond yields across G10 currencies can be attributed to FOMC announcement windows.

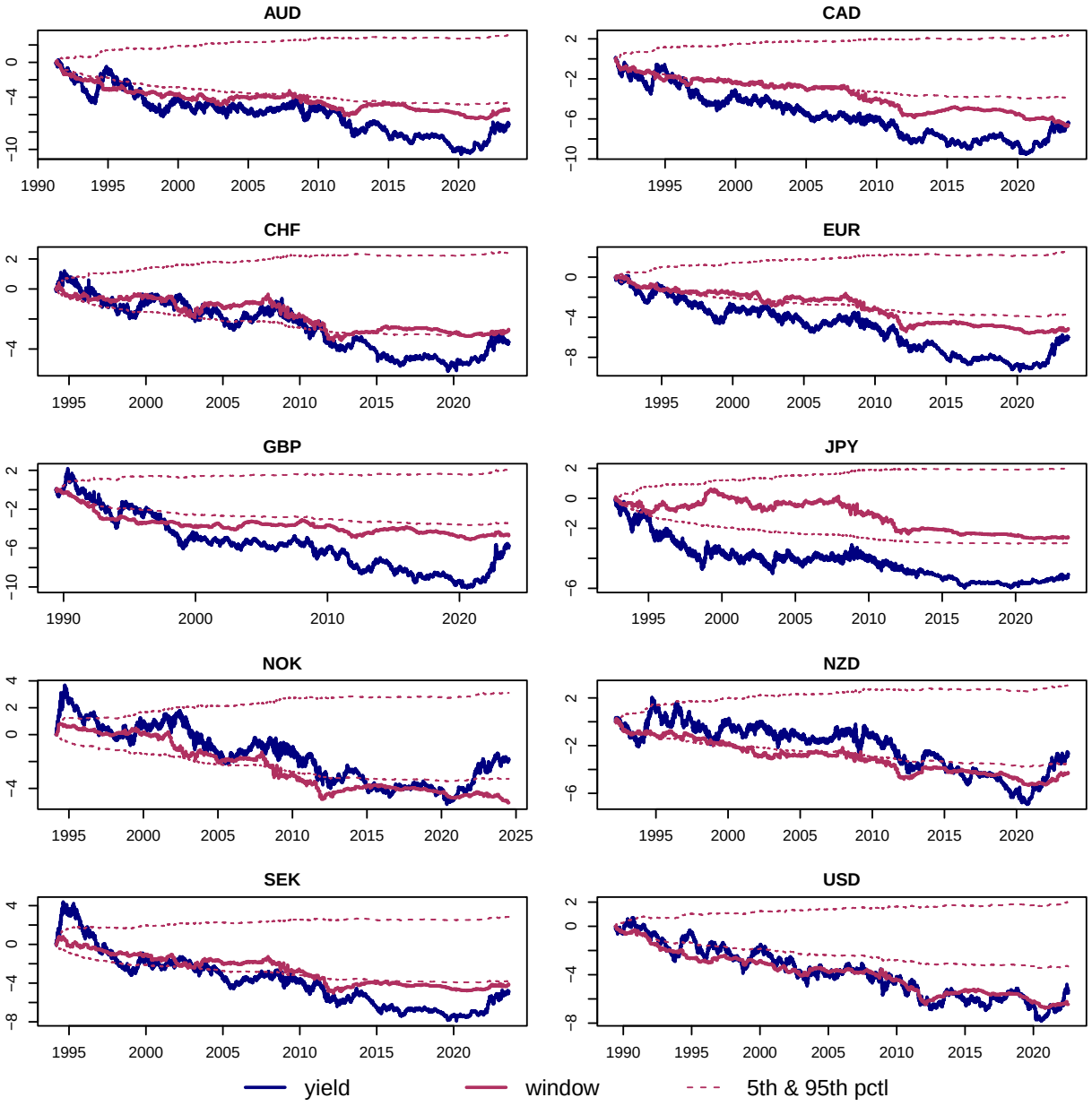


Figure 14: Placebo test 1.

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.

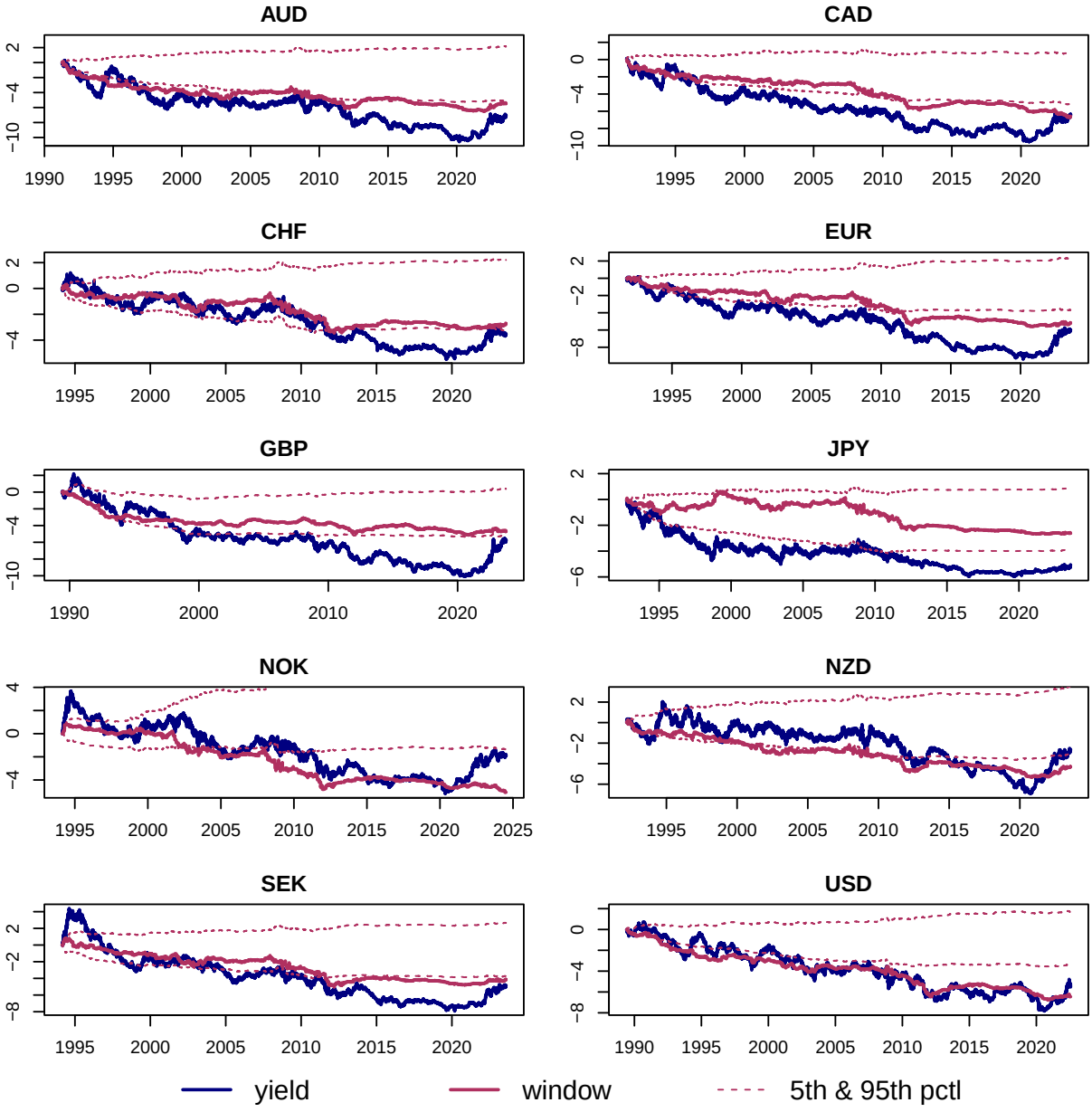


Figure 15: Placebo test 2.

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.

We demonstrate that the persistent decline in interest rates observed globally correlates closely with these U.S. announcements, while domestic central bank announcements account for a minor share of the overall trend. This highlights a unique characteristic of U.S. monetary policy: its ability to generate systematic long-term declines in global interest rates, in contrast to the cyclical effects produced by non-U.S. policy announcements.

Our exploration of the underlying mechanisms reveals that the decline in risk-neutral rates during FOMC windows plays a crucial role in this process. By employing high-frequency data and decomposing the shocks, we find that the impact of monetary policy is primarily transmitted through changes in risk-neutral rates rather than term premia, which remain stable during these periods. Furthermore, our investigation into the contributions of real interest rates indicates that the effects of FOMC announcements extend well beyond immediate market reactions, emphasizing their long-lasting nature.

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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

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

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

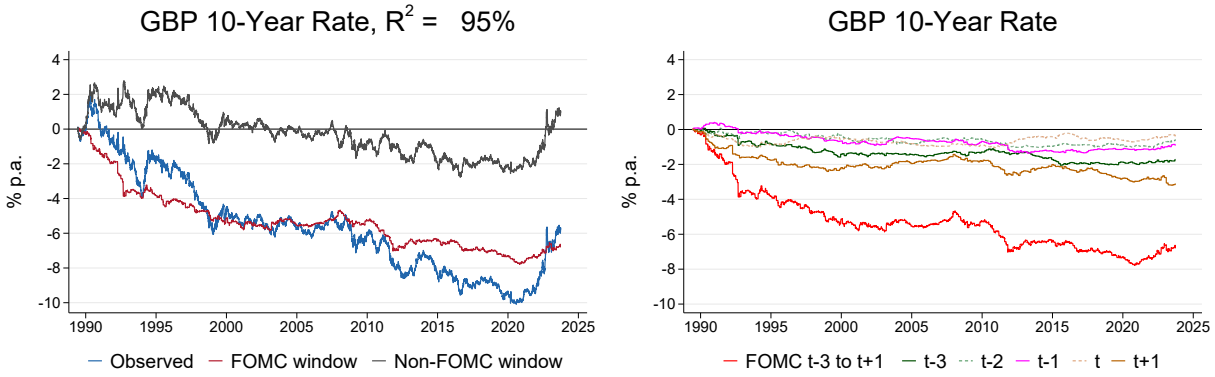


Figure A1: U.K. yields during extended FOMC windows.

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.

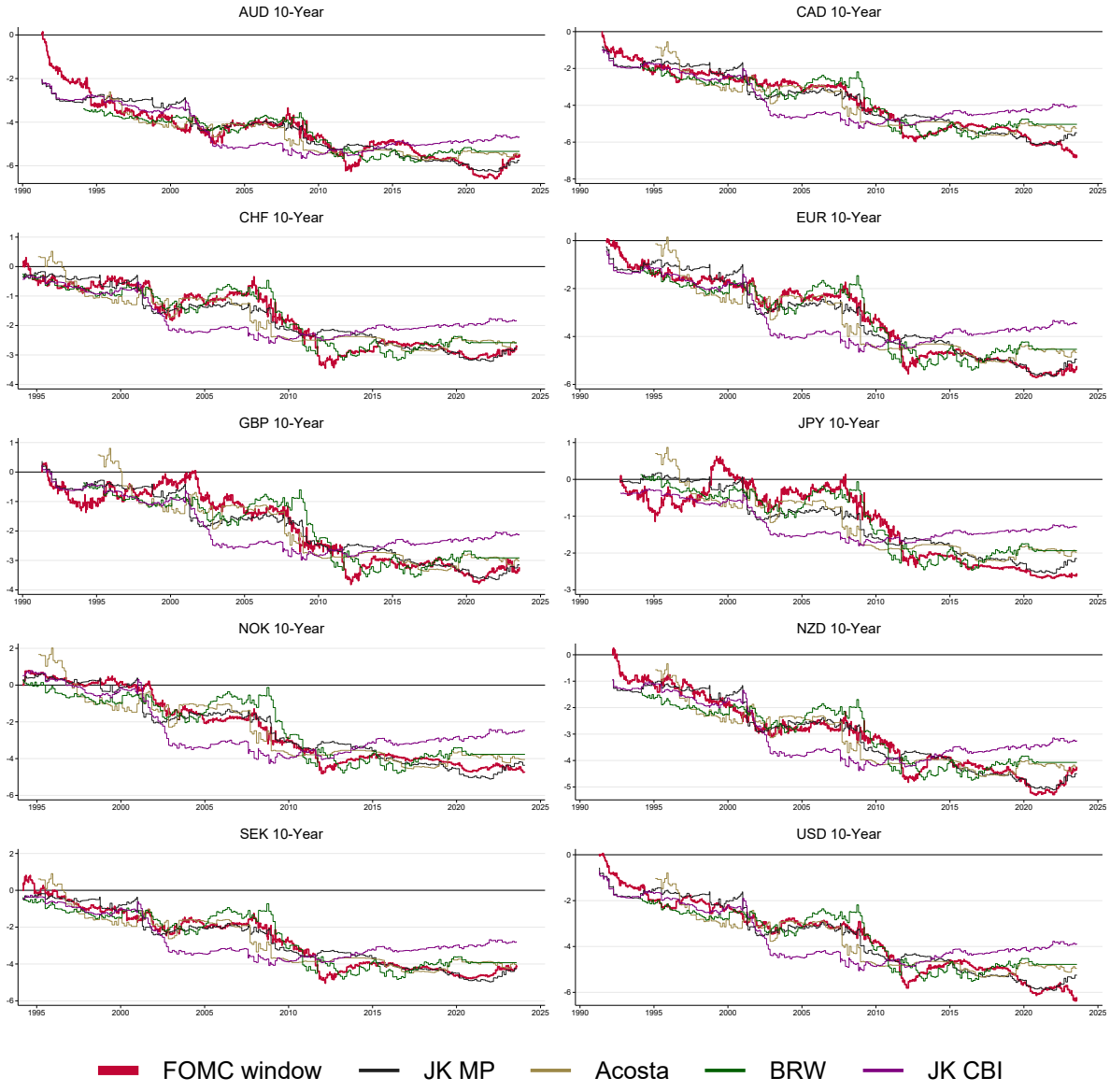


Figure A2: Cumulative yield changes and sums of U.S. HF monetary policy shocks.

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

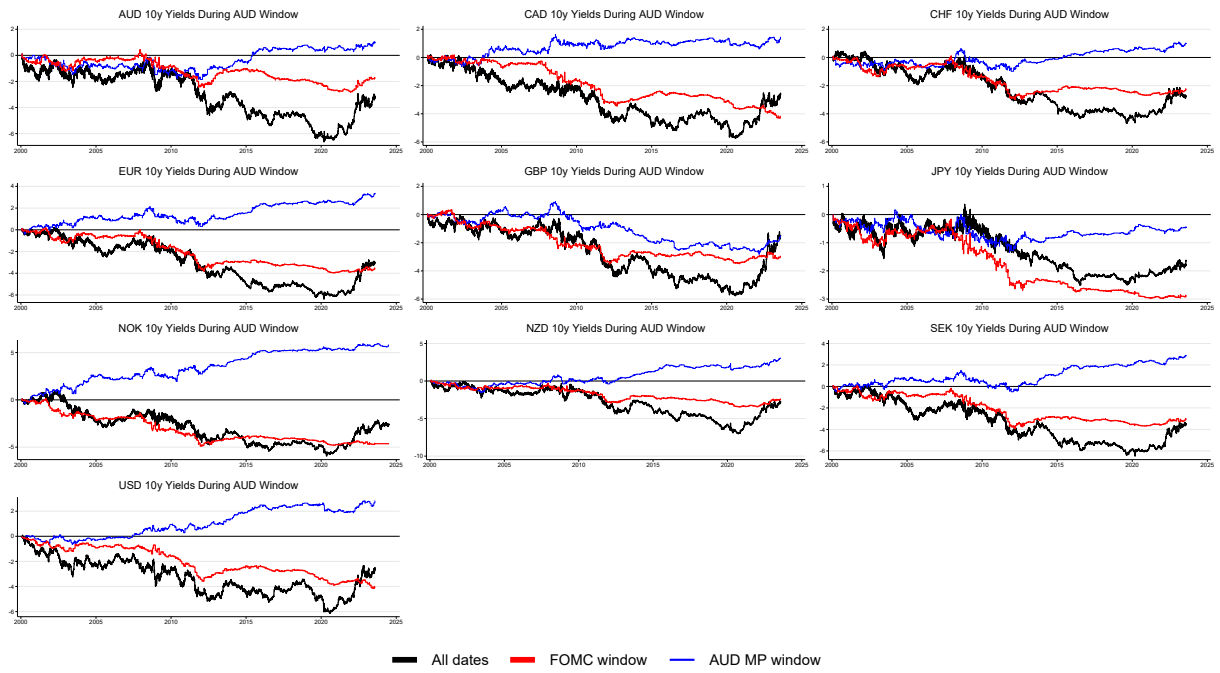


Figure A3: Sum of daily yield changes during U.S. or Australian central bank announcement windows.

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.

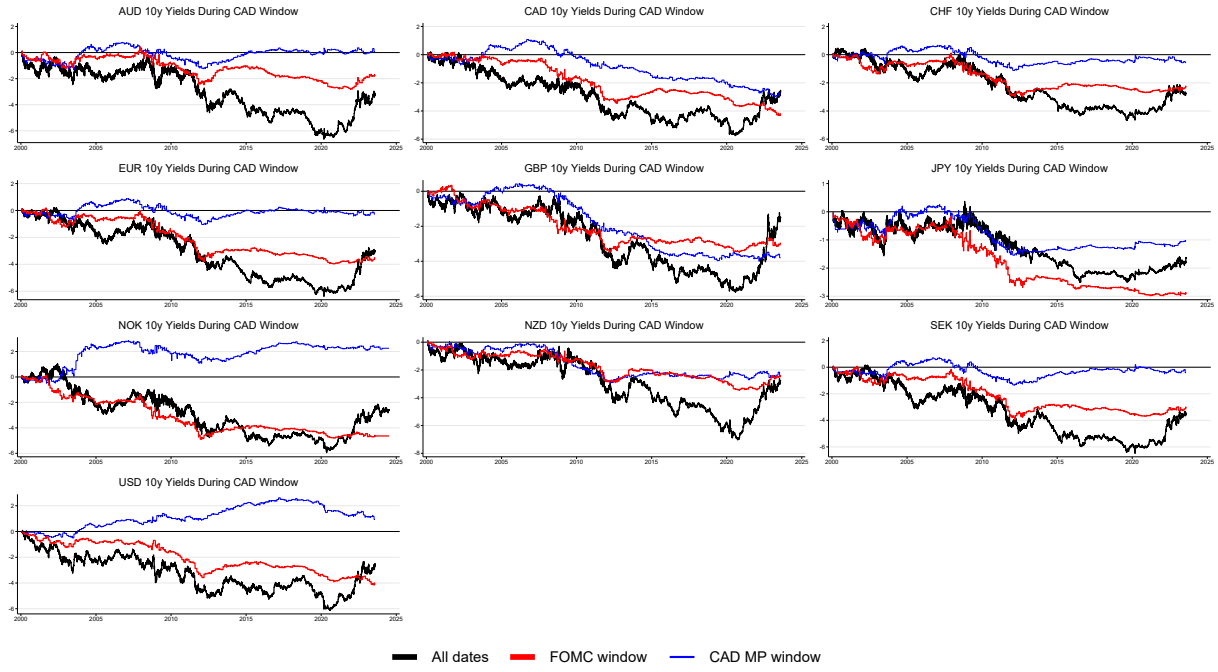


Figure A4: Sum of daily yield changes during U.S. or Canadian central bank announcement windows.

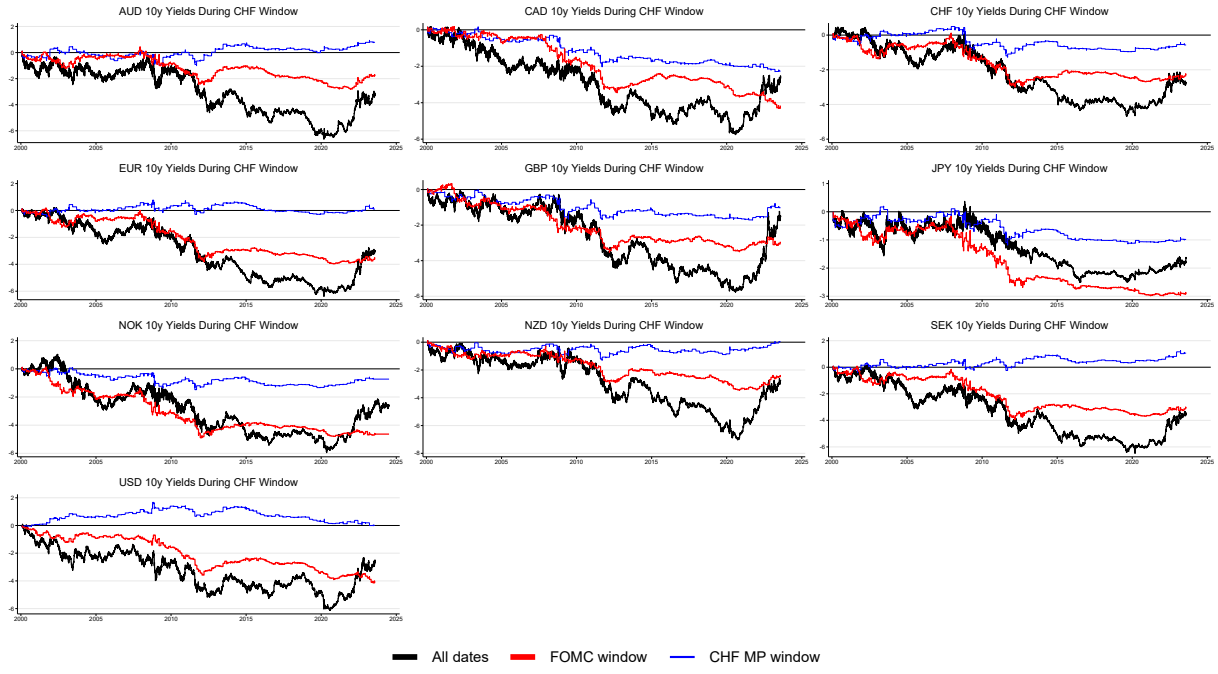


Figure A5: Sum of daily yield changes during U.S. or Swiss central bank announcement windows.

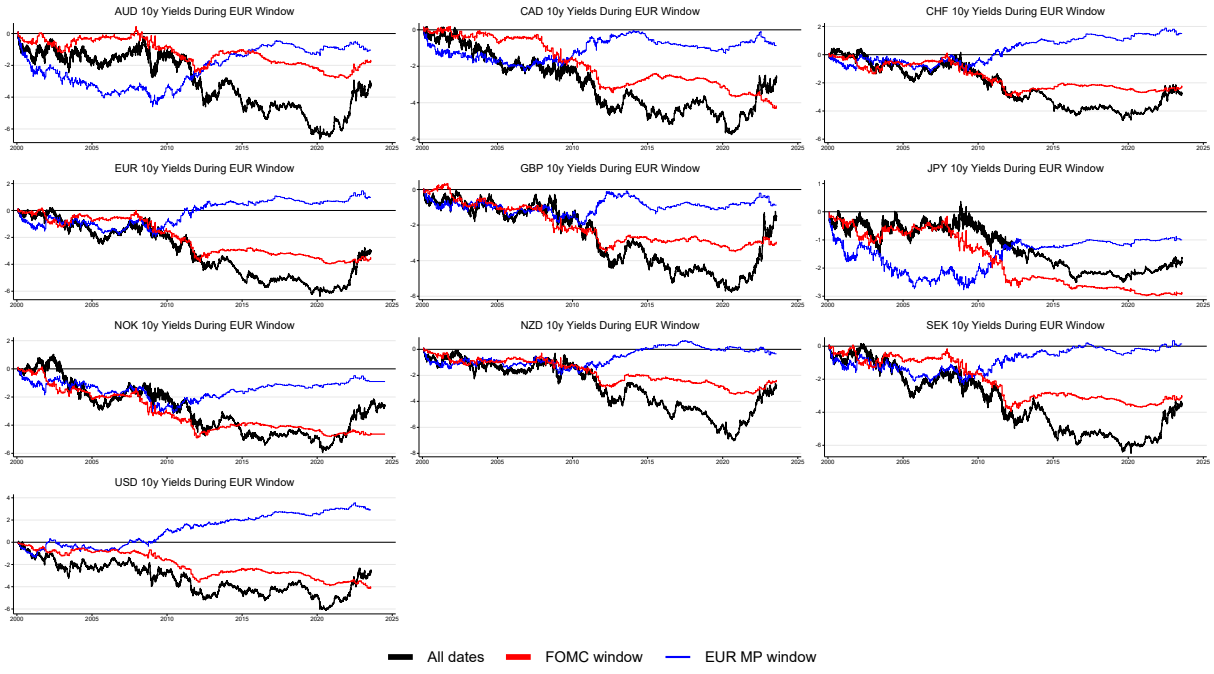


Figure A6: Sum of daily yield changes during U.S. or ECB announcement windows.

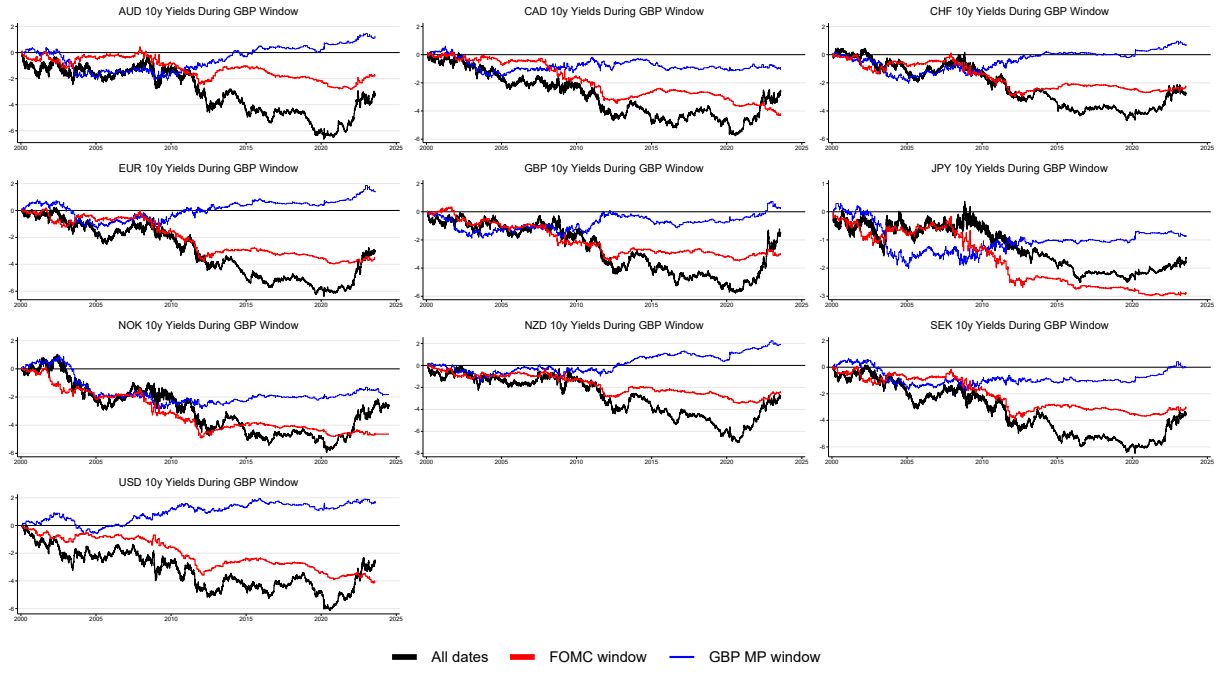


Figure A7: Sum of daily yield changes during U.S. or U.K. central bank announcement windows.

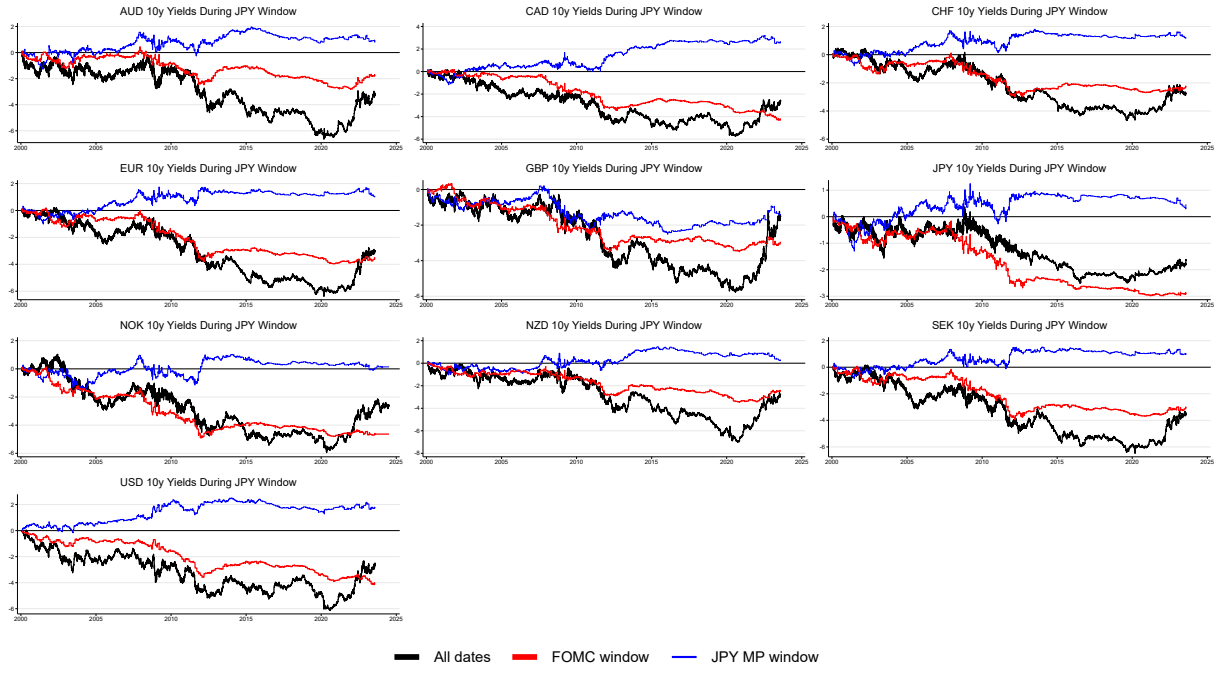


Figure A8: Sum of daily yield changes during U.S. or Japanese central bank announcement windows.

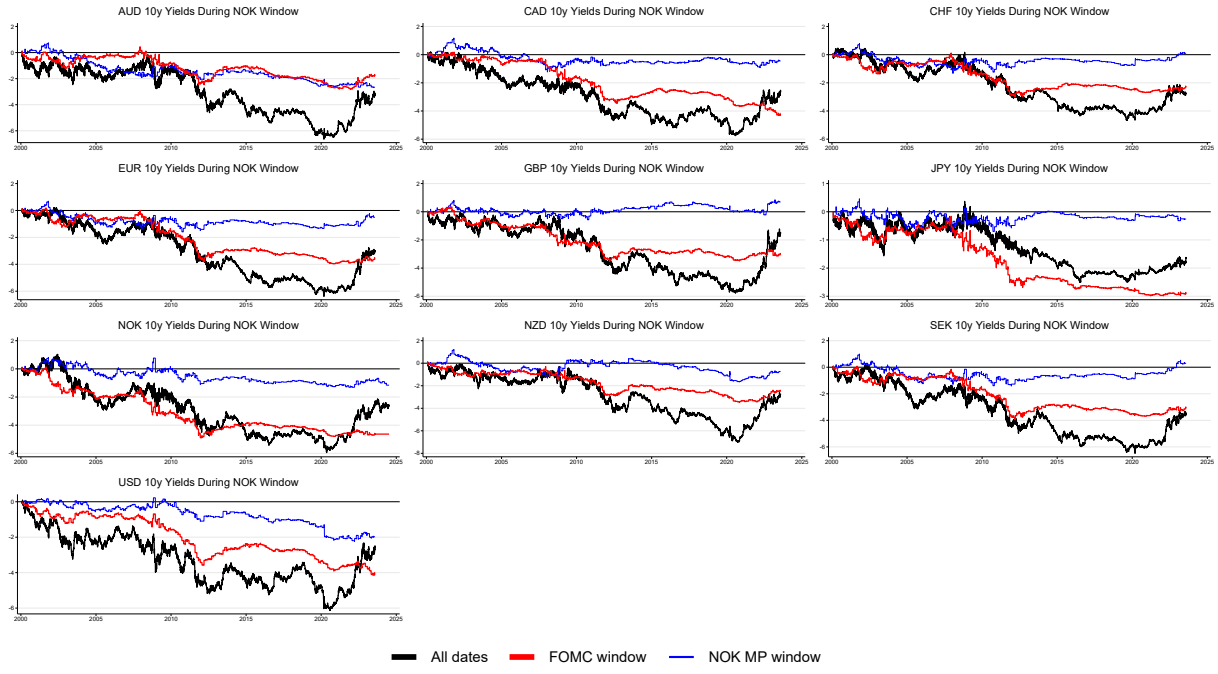


Figure A9: Sum of daily yield changes during U.S. or Norwegian central bank announcement windows.

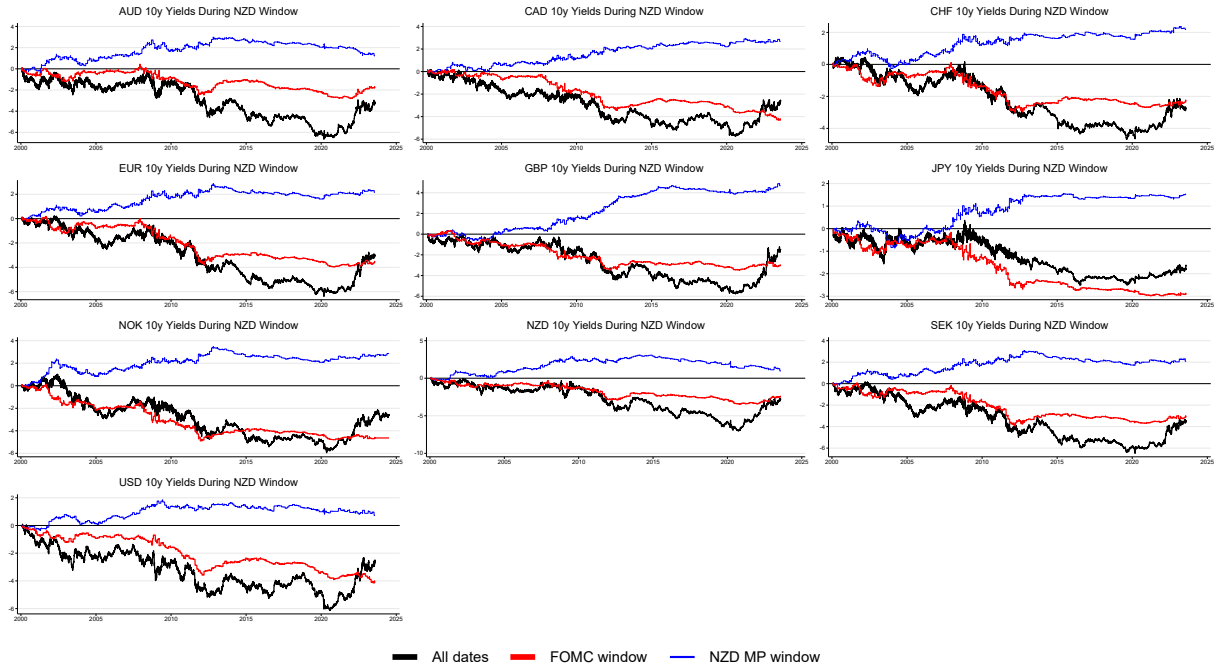


Figure A10: Sum of daily yield changes during U.S. or New Zealand central bank announcement windows.

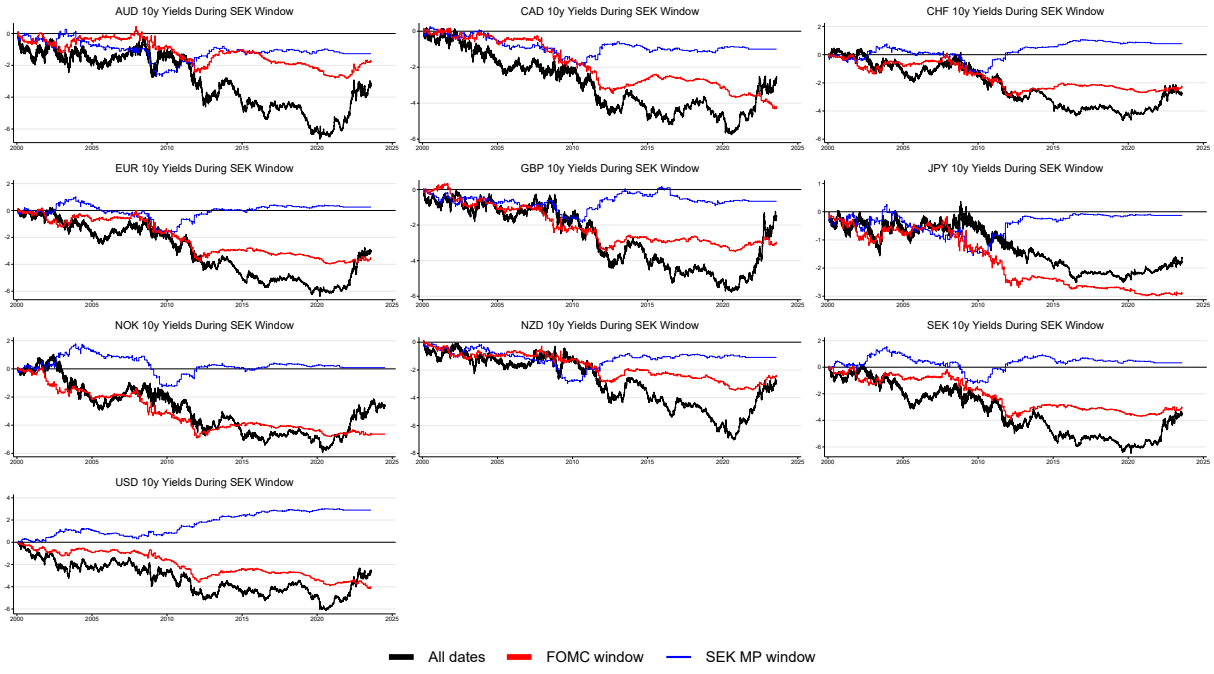


Figure A11: Sum of daily yield changes during U.S. or Swedish central bank announcement windows.

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.

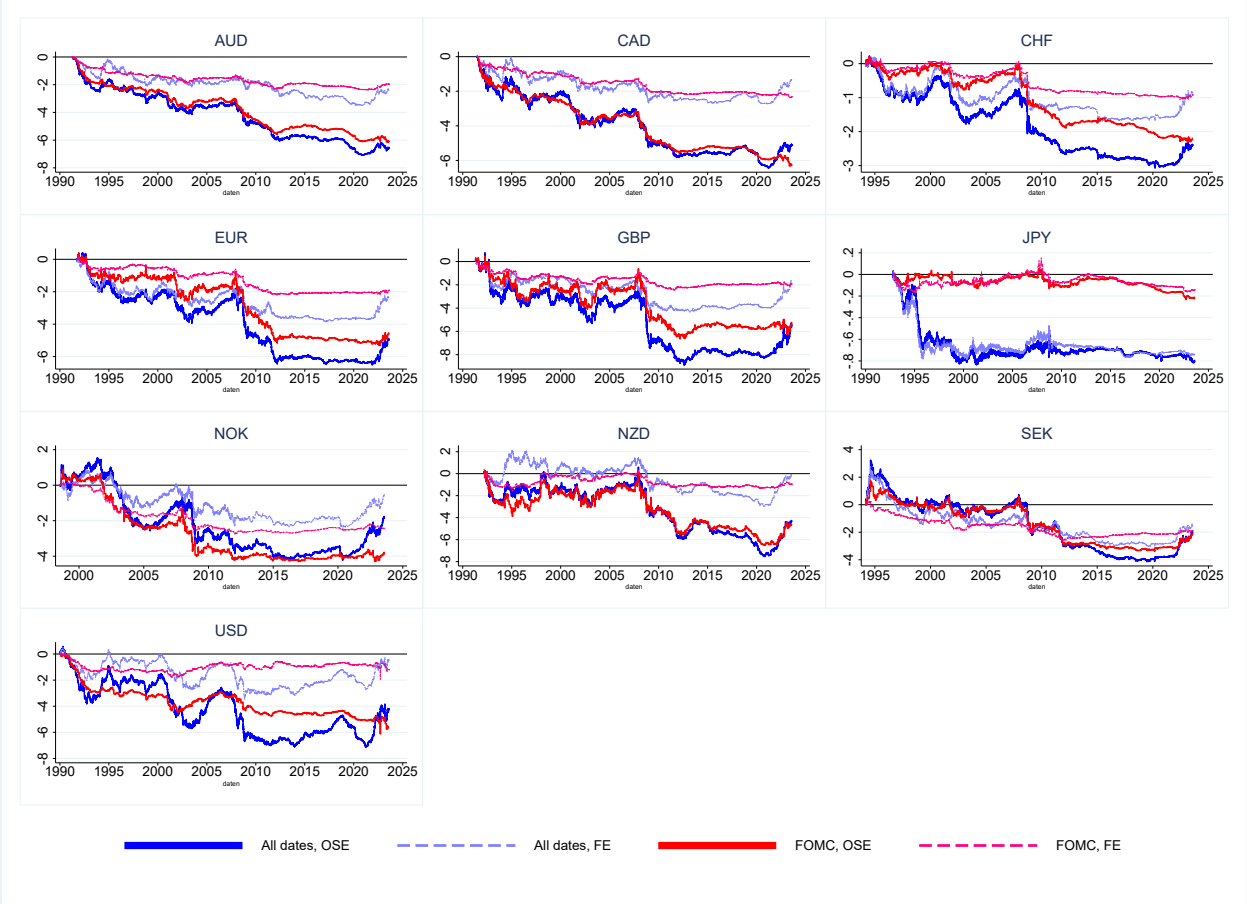


Figure A12: Dynamics of the 10-yr risk-neutral yield: OSE vs. FE.

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.

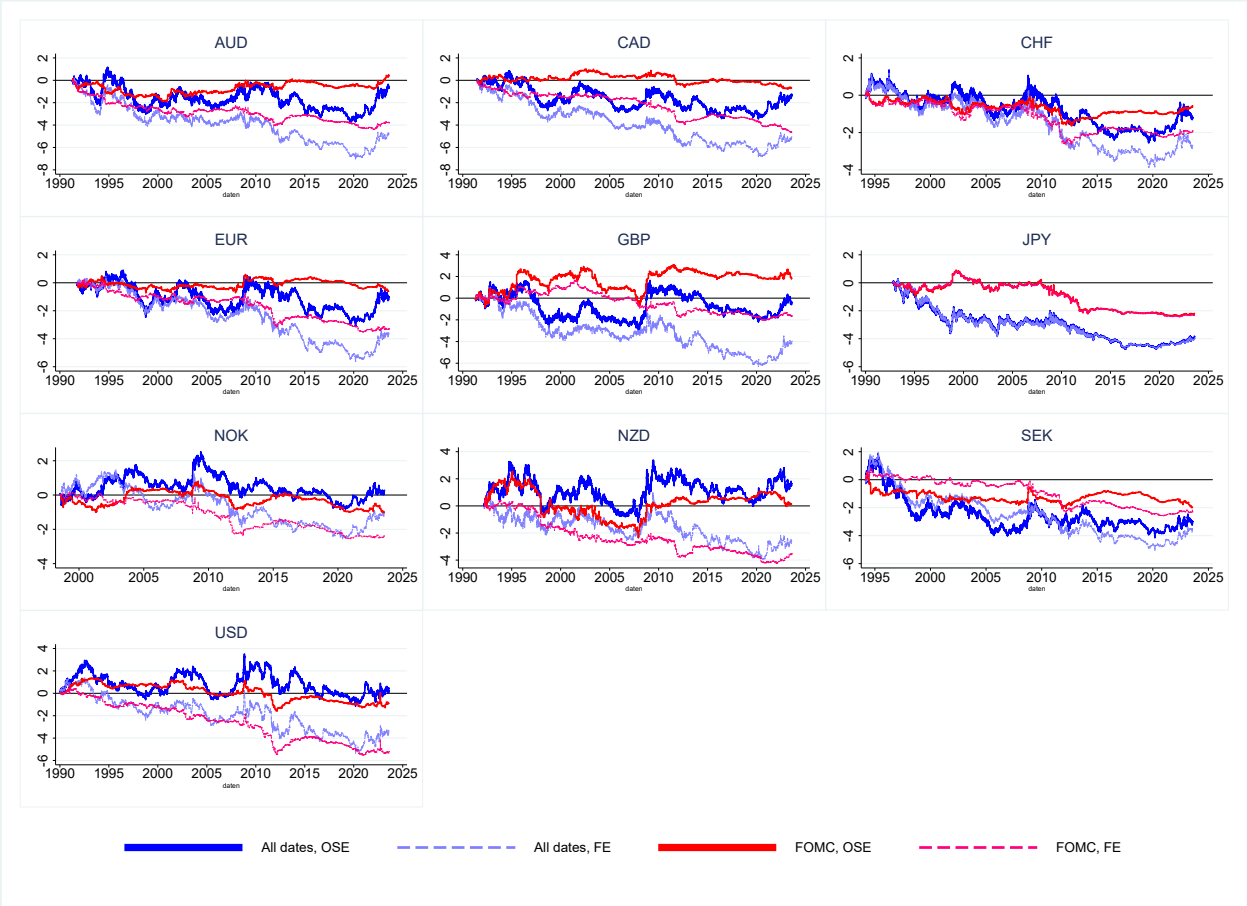


Figure A13: Dynamics of the 10-yr term premium: OSE vs. FE.

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.