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International Yield Spillovers*

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Abstract

This paper investigates spillovers from foreign economies to the U.S. through changes in long-term Treasury yields. We document a decline in the contribution of U.S. domestic news to the variance of long-term Treasury yields and an increased importance of overnight yield changes—a rough proxy for the contribution of foreign shocks to U.S. yields—over the past decades. Using a model that identifies U.S., Euro area, and U.K. shocks that move global yields, we estimate that foreign (non-U.S.) shocks account for at least 20% of the daily variation in long-term U.S. yields in recent years. We argue that spillovers occur in large part through bond term premia by showing that a low *level* of foreign yields relative to U.S. yields predicts a decline in distant forward U.S. yields and higher returns on a strategy that is long on a long-term Treasury security and short on a long-term foreign bond.

Keywords: Bond risk premia, foreign spillovers, event study, identification by heteroskedasticity, predictability.
JEL Classifications: E52, F37, G12, G15.

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Introduction

Over past three decades, long-term interest rates across advanced economies not only experienced a secular decline, but also appeared to exhibit more frequent synchronized high-frequency fluctuations. While correlations between monthly changes in long-term U.S. yields and monthly changes in long-term yields in Germany, Japan, the U.K., and Switzerland were on average about 0.4 in the early 1990s, these correlations in the past few years were on average about two times higher, reaching levels north of 0.7 in 2019 (see Figure 1).[1] Earlier studies suggest that the comovement between developed sovereign bond yields is mainly driven by powerful financial spillovers from U.S. monetary policy to the rest of the world and the influential effect of news about U.S. macroeconomic fundamentals on foreign financial markets.[2]

The behavior of long-term sovereign yields in advanced economies in recent years, however, has drawn increased attention to the possibility that U.S. yield movements are significantly affected by foreign developments. For example, even as the FOMC tightened policy between December of 2015 and the end of 2018, the 10-year Treasury yield remained low by historical standards. In fact, the 10-year Treasury yield became sufficiently low that the spread between the 10-year yield and the 3-month yield turned negative in May of 2019. While a negative spread is often interpreted as a signal of a future recession, some commentators have suggested that the spillover effects from foreign yields may have played a significant role in the inversion of the yield curve at that point in time:[3] the idea is that low levels of 10-year yields in advanced foreign economies, such as Germany and Japan, put downward pressure on the 10-year Treasury yield by making longer-term U.S. bonds more attractive relative to longer-term foreign bonds.[4]

This paper provides new empirical evidence that links the movements of long-term U.S. yields

[1]In contrast, correlations between 1-year international yields, on average, remained below 0.2 over the past decade, and do not exhibit the upward trend seen in long-term yields correlations.

[2]See, for example, Goldberg and Leonard (2003), Gerko and Rey (2017), Rogers, Scotti, and Wright (2018), and Brusa, Savor, and Wilson (2020).

[3]See, for example, the discussion of the yield curve inversion in the August 21, 2019, JPMorgan commentary article titled "Reassessing the Inverted Yield Curve." The debate about whether the decline in the spread between the 10-year yield and the 3-month yield was predicting a recession in 2018 and 2019 is now moot: the recession did occur in 2020, but the economic contraction is widely viewed as caused by a large, unanticipated negative shock (i.e., the COVID-19 pandemic).

[4]This debate was already alive after the European debt crisis. Some investors reportedly argued that the slow post-crisis growth and aggressive monetary stimulus in Europe had pushed European long-term yields to ultra-low levels, leading investors to buy long-term U.S. government bonds for the higher income they offer compared to European sovereign debt. See, for example, the June, 2012 The Economist article "To strive, to seek, to find, and not to yield."

to spillovers from yields in advanced foreign economies based on three alternative methodologies. We begin by constructing two simple variance ratios: the economic news variance ratio, as defined here, is the variance of 10-year Treasury yield changes accrued around a narrow window bracketing the release of major U.S. economic and policy announcements relative to the overall variance of the changes in the 10-year Treasury yield; and the overnight variance ratio, as defined here, is the variance of 10-year Treasury yield changes outside of U.S. daytime trading hours—when investors likely receive information mostly about foreign economies—relative to the overall variance of the changes in the 10-year Treasury yield. We find that the economic news variance ratio declined from explaining 30 percent of the variation in the long-term U.S. yield between 1992 and 1996 to representing 8 percent of the variation in the 2015-2019 period. Perhaps more remarkably, the overnight variance ratio—which we take as a rough proxy for the contribution of foreign shocks to U.S. yields—increased from 13 percent in the 1992-1996 period to 30 percent in the 2015-2019 period. These findings provide suggestive evidence that the role of news about domestic fundamentals in explaining moves in long-term Treasury yields has been declining over the past three decades, and that spillovers from foreign economies to long-term U.S. yields have a significant and increasing role in explaining fluctuations in long-term U.S. rates.

Second, we propose a measure of the magnitude of spillovers from foreign yields using a model that decomposes U.S., Euro area, and U.K. long-term yield changes into three kinds of shocks: a country shock that moves bond yields globally, an idiosyncratic country shock (i.e., shock that only affects its own country yield), and "other global" shock. Country shocks that move yields globally are visible on days with influential monetary policy announcements and macro data releases but, in light of the high degree of correlation between yields in our sample of countries, it stands to reason that these shocks are also present on days without notable economic releases in these countries. We posit that while the pattern of the response of global yields to these shocks is the same for days with notable news and days without notable news, their overall magnitudes are larger on notable news days; this assumption allows us to estimate the model using the identification-by-heteroskedasticity technique of Rigobon (2003), Rigobon and Sack (2004), and Wright (2012). Using time-synchronized data on daily changes in U.S., German, and U.K. long-term yields for January 2010 through August of 2017 and a set of days with notable news, we estimate that a shock that lowers Euro area (U.K.) long-term yields by 100 basis points will lead to a decline in U.S. long-term yields of about 50 (40) basis points, roughly consistent with the event-study estimates in Curcuru, De Pooter, and Eckerd (2018). We further document that the share of variance of long-term U.S. yields explained by Euro area and U.K. shocks is non-negligible. Our estimates suggest that between 20 to 25 percent of 10-year Treasury yield variations are accounted for by foreign (non-U.S.) shocks over the 2010-2017 period. This figure is likely a lower bound on the true degree of spillovers from foreign yields to U.S. yields, as the effects from other economies, such as Japan and China, are either estimated to be very small or unaccounted for in our measure, reflecting the limitations of our model.

Third, we provide evidence that the downward pressure on U.S. yields from the low level of yields in advanced foreign economies (relative to U.S. yields) also manifests itself in terms of *predictable* variations in U.S. yields. We explore this effect by running predictive regressions of weekly changes in long-term U.S. yields on the spread between the 10-year Treasury yield and the 10-year foreign yield. Our measure of long-term foreign yield is a GDP-weighted average of yields on government debt for three advanced foreign economies, Germany, Japan, and the U.K., which have safety and liquidity features that are somewhat comparable to U.S. Treasury securities. The predictive regressions show that after a widening of the U.S.–foreign long-term yield spread, investors expect Treasury yields to decline over the following week, even after controlling for factors capturing the U.S. business cycle—the near-term spread (Engstrom and Sharpe, 2019), the forward spread (Fama and Bliss, 1987), the Aaa-Treasury spread (Krishnamurthy and Vissing-Jorgensen, 2012), and the effective duration of mortgage-backed securities (MBS) (Hanson, 2014; Malkhozov, Mueller, Vedolin, and Venter, 2016). The predictive power of the U.S.–foreign yield spread is economically and statistically significant for future changes in long-term Treasury yields outside of windows bracketing the release of key U.S. economic releases, whereas it does not seem to predict yield fluctuations around U.S. macroeconomic and policy announcements. Interestingly, the predictability of the U.S.–foreign long-term yield spread increases when the overnight variance of U.S. yields is higher than usual, which are times when shifts in the spread between long-term U.S. and long-term foreign yields are likely driven by information concerning the economic outlook abroad. The predictive ability of the U.S.–foreign yield spread raises the question of whether it reflects predictable movements in short-term rate expectations or predictable movements in term premia. Starting from the premise that distant nominal forward rates are mostly driven by time-varying term premia, we document the predictability of forward rates for different horizons. Our results show that the U.S.–foreign yield spread is a stronger predictor of distant forward rates than short-forward rates. Similarly, we find that the U.S.–foreign long-term yield spread is a strong predictor of the excess return on a strategy that takes a long position in a long-term U.S. bond and a short position in a long-term foreign bond. Taken together, these empirical results suggest that the U.S.–foreign long-term yield spread is more informative about term premia than about future short rates, supporting the idea that spillovers to long-term U.S. yields likely occur through a portfolio balance channel.

Related Literature. This paper is related to several strands of the literature. One is those that study the presence of a global factor driving yields across advanced economies.[5] Diebold, Li, and Yue (2008) find that global yield factors linked to macroeconomic fundamentals appear to explain a significant fraction of country yield curve dynamics. Furthermore, Dahlquist and Hasseltoft (2013) find that a global Cochrane-Piazzesi (CP) factor has a strong forecasting power for bond returns in both the U.S. and industrial countries, and Jotikasthira, Le, and Lundblad (2015) find that a global inflation factor and the level of U.S. yields drive the comovement between international yields. The global factors identified in these papers, however, are closely related to bond risk premia and monetary policy in the U.S. Dahlquist and Hasseltoft (2013) show that their global CP factor is highly correlated with the U.S. CP factor and Jotikasthira, Le, and Lundblad (2015) find that the level factor in the U.S. yield curve is the most important contributor to the correlation between U.S. and German yields. This paper differs from these contributions in that, building on the observation that a large fraction of what drives long-term yields in the U.S. and other advanced economies is "global," much of our focus is on taking apart this global component, separately identifying the contribution of shocks emanating from the U.S. and from advanced foreign economies.

This paper also builds on and extends the literature that studies the international transmission of foreign and U.S. macroeconomic and monetary policy announcements in global capital markets.[6] Gerko and Rey (2017) and Rogers, Scotti, and Wright (2018), using high-frequency asset price

[5] There is also a sizable literature that analyzes multi-country yield curves and exchange rates in no-arbitrage term structure models, which have "global" factors and "local" factors. See, for example, Backus, Foresi, and Telmer (2001), Ahn (2004), Sarno, Schneider, and Wagner (2012), and Kaminska, Meldrum, and Smith (2013).

[6] There is a large literature focusing on the international transmission of U.S. monetary policy shocks to advanced and emerging economies. See, for example, Kim (2001), Bowman, Londono, and Sapriza (2015), Neely (2015), Aizenman, Chinn, and Ito (2016), Dedola, Rivolta, and Stracca (2017), Bernanke (2017), and Curcuru, Kamin, Li, and Rodriguez (2018).

movements around monetary policy events as an external instrument to identify monetary policy shocks in a structural VAR, find strong evidence of important spillovers from U.S. monetary policy to bond risk premia in Germany, Japan and the U.K. On the other hand, their evidence on spillovers from monetary policy actions in advanced foreign economies to long-term Treasury yields is mixed and mostly sides with the view that the U.S. sets the tone in international bond markets.[7] Furthermore, Goldberg and Leonard (2003) find that, while many U.S. economic news had significant effects on German yields, German and Euro area economic news generally had insignificant effect on U.S. yields. By contrast, using an event-study approach, Curcuru, De Pooter, and Eckerd (2018) do find evidence of spillovers from German yields to U.S. yields following policy communications from the ECB. Consistent with their results, we find that the response of U.S. yields to foreign shocks is economically and statistically significant. In addition, Dilts Stedman (2020) finds evidence of spillovers from the Euro area and Bank of England unconventional monetary policy measures to U.S. yields, particularly after 2015, and Kearns, Schrimpf, and Xia (2020) find significant evidence of spillovers from ECB announcements, while the spillovers from the actions of other advanced economy central banks, including the Bank of England and the Bank of Japan, are estimated to be mild. However, existing research using an event-study methodology around monetary policy announcements still leave open the question how much of U.S. yield variation is accounted for by foreign shocks because days with ECB announcements represent only a fraction the total number of business days. By imposing more structure to the model and including the behavior of yields on days without notable news, our empirical approach allows us to estimate the contribution of U.S. and foreign (non-U.S.) shocks to the total variance of yield changes. Our paper also distinguishes from the current literature by offering complementary evidence based on the overnight variance ratio, which exploits the round-the-clock trading in the Treasury market and highlights the fluctuations in long-term yields outside U.S. trading hours. In addition, while most of this literature has focused on the effect of central bank communications (i.e., monetary policy shocks) on global yields, our two approaches consider the spillovers from both macroeconomic and monetary policy announcements.

Lastly, this paper is related to an older literature that studies interest rate linkages in a coin-

[7] Relatedly, Brusa, Savor, and Wilson (2020) show that investors in equity markets in Germany, Japan, and the U.K. demand a high risk premium around FOMC announcements, but U.S. equity markets seem unmoved by decisions of the European Central Bank (ECB), the Bank of Japan and the Bank of England.

tegration framework. Kirchgässner and Wolters (1993), for example, examines the cointegration of U.S., German, and other European short-term interest rates to test the "German Dominance" hypothesis, and Chinn and Frankel (1995) studies the relative influence of U.S. and Japanese real interest rates on the determination of rates in Pacific Rim countries using an error correction model. The predictive power of the U.S.–foreign long-term yield spread documented in Section 3 can be viewed analogous to the empirical evidence of the presence of cointegrating vector documented in this literature. However, these studies have focused on shorter-maturity interest rates (as opposed to longer-maturity interest rates that are the focus of our paper), and we are not aware of studies in this framework that focus on examining the influence of other countries on U.S. interest rates. Furthermore, compared to cointegration approaches, our predictive regressions allow us to more manageably control for other known predictors of bond returns.

The paper is organized as follows. Section 1 describes how to measure the two variance ratios, the economic news variance ratio and overnight variance ratio, and documents their trajectory over the past three decades. Section 2 presents a measure of the degree of spillovers to U.S. yields based on heteroskedasticity of long-term yields around notable news events. Section 3 documents the predictive ability of the U.S.–foreign long-term yield spread for changes in long-term U.S. yields. The last Section concludes. An Appendix contains details of the data sources and variable definitions, details of the identification assumptions and criteria for selecting notable news days utilized in Section 2, and robustness checks.

1 Decomposing Round-the-Clock Variations in Long-Term Yields

One simple way to gauge the contribution of domestic macroeconomic and monetary policy announcements to the overall variation in long-term yields is to decompose the change in yields between time t and $t+1$ into two components as

$$\Delta y_{t+1} = \Delta y_{\mathrm{a},t+1} + \Delta y_{\mathrm{na},t+1}, \tag{1}$$

where $\Delta y_{\text{a},t+1}$ is the yield change accrued around a narrow window bracketing the release of major economic and policy announcements, and $\Delta y_{\text{na},t+1}$ is the yield change outside of theses windows.[8] Using this decomposition, we construct the economic news variance ratio as the variance of long-term yields around economic announcements relative to the overall variance of long-term yields,

$$\frac{\text{Var}(\Delta y_{\text{a},t+1})}{\text{Var}(\Delta y_{t+1})}. \tag{2}$$

This ratio measures the importance of domestic macroeconomic and policy announcements in explaining the variation in long-term yields.

More specifically, we define Δy_{t+1} as the weekly change in the yield on the most recently issued 10-year Treasury security. We use intraday yields on the 10-year on-the-run Treasury security to construct the change in yields between 5 minutes before to 25 minutes after major U.S. macroeconomic and policy announcements. We focus on the reaction of yields around the release of the FOMC statement and the following fourteen major releases: nonfarm payrolls, CPI, PPI, retail sales, PCE, durable goods orders, initial unemployment claims, industrial production, ISM manufacturing, capacity utilization, real GDP, Michigan consumer confidence, leading economic indicators, and new home sales.[9] We cumulate the change in yields around macroeconomic releases to a weekly frequency, so that $\Delta y_{\text{a},t+1}$ represents the change in long-term Treasury yields during week $t+1$. Our sample covers January of 1992 to December of 2019.

Alternatively, yield changes can be decomposed as

$$\Delta y_{t+1} = \Delta y_{\text{o},t+1} + \Delta y_{\text{d},t+1}, \tag{3}$$

where $\Delta y_{\text{o},t+1}$ and $\Delta y_{\text{d},t+1}$ represent changes in the yield overnight and changes in the yield during the domestic daytime trading session, respectively. The overnight variance ratio is defined as

$$\frac{\text{Var}(\Delta y_{\text{o},t+1})}{\text{Var}(\Delta y_{t+1})}. \tag{4}$$

[8] This decomposition is analogous to the decomposition in Faust and Wright (2018) of bond returns earned around announcements and at other times.

[9] These announcements have been shown to be influential for bond returns in previous studies such as Fleming and Remolona (1999), Balduzzi, Elton, and Green (2001), Andersen, Bollerslev, Diebold, and Vega (2007), Faust, Rogers, Wang, and Wright (2007), Swanson and Williams (2014), and Faust and Wright (2018).

This ratio provides a rough measure of the degree of spillovers from news about foreign macroeconomic fundamentals and economic policies to domestic long-term yields, since many of the most important foreign economic news are released outside of U.S. daytime trading hours. We define overnight yield changes as the change in the 10-year Treasury yield between 8 a.m. and 5 p.m. of the previous business day.[10] To match the weekly frequency of the data, we cumulate the overnight changes over each week in the sample.

Table 1 reports the estimates of the economic news variance ratio and the overnight variance ratio for the full sample (1992–2019), the first five years of the sample (1992–1996), and the last five years of the sample (2015–2019). The last two columns of Table 1 show the difference in the economic variance ratio and the overnight variance ratio between the early and the late sample as well as the Wald statistic testing the null hypothesis that the contribution of these news to the variance of long-term U.S. yields has remained constant. Newey and West (1987) standard errors are provided in parenthesis, and the p-values associated with the Wald test are provided in brackets. These values are heteroskedasticity-robust and allow for serial correlation up to 52 lags.

As shown in the first column of Table 1, from 1992 to 2019 the economic news variance ratio is around 20 percent. The sub-sample evidence, reported in columns (2) and (3), shows that the economic variance ratio—the fraction of the variance in yields explained by yield fluctuations around economic news releases—has declined from representing close to 30 percent of the variation in yields between 1992 and 1996 to about 10 percent in the 2015-2019 period. The Wald test shows that the decline in the economic news variance ratio between the early and the late parts of the sample is highly statistically significant. All in all, the evidence suggesting a decreasing role of fluctuations in yields around U.S. macro announcements is striking.

The decline in the economic variance ratio—and, equivalently, the rise in the share of yield variations from movements outside announcement windows—appears to reflect in large part the rise in the share of yield variations coming from overnight hours.[11] Table 1 shows that the share of yield variation due to overnight yield movements increased from percentages in the low teens in

[10] The Treasury market is an over-the-counter market that is open (almost) around the clock. Therefore, there are not official opening and closing times for daytime trading sessions.

[11] We have $\mathrm{Var}(\Delta y_{\mathrm{na}})/\mathrm{Var}(\Delta y) \approx 1 - \mathrm{Var}(\Delta y_{\mathrm{a}})/\mathrm{Var}(\Delta y)$, because $\mathrm{cov}(\Delta y_{\mathrm{a}}, \Delta y_{\mathrm{na}}) \approx 0$. And note that, since there are practically no U.S. macro data releases or policy announcements during overnight hours, the overnight yield changes can be viewed as a component of the the yield changes during non-announcement periods Δy_{na}, i.e., $\Delta y_{\mathrm{na}} = \Delta y_{\mathrm{o}} + \Delta y_{\mathrm{na,d}}$, where $\Delta y_{\mathrm{na,d}}$ denotes the yield changes in non-announcement periods that occur during the day time.

the 1992-1996 period to slightly above 30 percent in the 2015-2019 period. As indicated by the Wald statistic, the increase in the overnight variance ratio between the 1992–1997 and 2015–2019 is statistically significant.

Figure 2 provides a more detailed look at the evolution of the variance ratios by plotting the economic news variance ratio (dotted line) and the overnight variance ratio (solid line) from 1992 to 2019 using a 5-year rolling window; the variance ratios plotted at time t are computed using weekly data from t-5 years to t. As can be seen, the overnight variance ratio has trended up more or less steadily over time, though the increase appears a bit faster in the more recent period, which followed developments such as the ECB and the Bank of Japan (BoJ) setting negative policy interest rates and launching asset purchase programs targeting a wide range of long-duration assets. The economic news variance ratio, on the other hand, shows an overall decline over the 1992–2019 period, with a notable dip and bounce-back in the 2000s. Interestingly, this variance ratio was higher during the effective lower bound (ELB) period (2008–2015) than in the period after the Federal Reserve started increasing the target for the federal funds rate.[12]

Admittedly, overnight yield changes are only a rough measure of spillovers from foreign economies to U.S. yield changes. Some important foreign news, such as the ECB press conference following its policy announcement, arrive during the daytime U.S. trading session, and some U.S. economic news occur during overnight trading hours as is, for example, the case of the outcome of the U.S. presidential election. Even so, the evidence that the contribution of movements in yields during overnight trading hours not only increased significantly over the past three decades but has surpassed the contribution of moves around major domestic economic announcements is striking, and suggests that spillovers from foreign economies to long-term U.S. yields have a significant and increasing role in explaining fluctuations in long-term U.S. rates.

It is important to note that the increased contribution of overnight changes in yields over our sample period is not explained by the possibly lower liquidity of the Treasury market during the overnight hours in the earlier part of the sample. While intraday data on yields that are used in

[12] Swanson and Williams (2014) document that shorter maturity yields were less sensitive to economic data releases during the ELB period, especially following the introduction of date-based forward guidance. On the other hand, these authors find that longer-term yields such as the 10-year yield, which is our focus, were less affected by the ELB (See, Swanson, 2018, for evidence including the last years of the ELB period). At the same time, the volatility compression effect due to the ELB, if there is any, can be expected to appear in both the numerator and the denominator of the variance ratios, therefore the variance ratios would not be particularly influenced by the ELB.

our exercise *are* more spotty during overnight hours, our measure of overnight yield changes utilizes only the yield before the start of U.S. daytime session (namely, 8:00 a.m.) and the yield at the previous close of U.S. daytime session (namely, 5:00 p.m.); therefore, as long as market-moving foreign news during the overnight hours are incorporated in Treasury prices before the start of U.S. daytime session, our measure would capture them. The other possibility – that overnight foreign news are not incorporated until after the start of U.S. daytime trading – would require a fairly strong belief in market inefficiency.[13]

2 Decomposing Multi-Country Yield Changes

2.1 Identification by Heteroskedasticity

In the empirical exercise that follows, we assume that dynamics of U.S., Euro area (EA) and U.K. long-term yields can be written as

$$\begin{bmatrix} \Delta y_t^{US} \\ \Delta y_t^{EA} \\ \Delta y_t^{UK} \end{bmatrix} = \begin{bmatrix} 1 & \Gamma_{12} & \Gamma_{13} \\ \Gamma_{21} & 1 & \Gamma_{23} \\ \Gamma_{31} & \Gamma_{32} & 1 \end{bmatrix} \begin{bmatrix} \varepsilon_t^{US} \\ \varepsilon_t^{EA} \\ \varepsilon_t^{UK} \end{bmatrix} + \begin{bmatrix} 1 \\ 1 \\ 1 \end{bmatrix} \eta_t + \begin{bmatrix} \ell_t^{US} \\ \ell_t^{EA} \\ \ell_t^{UK} \end{bmatrix}, \tag{5}$$

where Δy_t^i is the daily change in country i's long-term yield ($i = 1, 2, 3 =$US, EA, UK). Our model assumes two types of individual country shocks. The shocks $\varepsilon_t^{US}, \varepsilon_t^{EA}, \varepsilon_t^{UK}$ are shocks that arise from country i and affect not only their own country yields but also yields in other countries; the spillover from country i shock to country j is given by Γ_{ji}. In this sense, ε-shocks are global shocks. The ℓ_t shocks—the other individual country shocks—are purely local shocks (i.e., idiosyncratic shocks) that affect only their own country yields. Yields are also assumed to be driven by a third unobserved shock, η_t, that affects yields in all countries. This shock gauges global shocks not captured by ε_t^j such as those emanating from other economies or regions, e.g., Asia or Middle East. Although we assume that all shocks are uncorrelated, long-term yields will be correlated as long as Γ_{ij} are different from zero or country yields are sensitive to the "other global" shock η_t.

[13] In his analysis of the Treasury market, Fleming (1997) shows that U.S. trading jumps higher in the first-half hour of New York trading (7:30 a.m. to 8 a.m.). The time window we use to compute the overnight change in yields, in turn, likely captures the response of domestic investors to foreign news, even if domestic traders reacted with some lag to overnight foreign news.

The model (5) can be also written in a matrix form for general N countries:

$$\Delta y_t = \Gamma \varepsilon_t + \iota' \eta_t + \ell_t, \tag{6}$$

where $\Delta y_t = [y_t^1, ..., y_t^N]'$, $\varepsilon_t = [\varepsilon_t^1, ..., \varepsilon_t^N]$, $\ell_t = [\ell_t^1, ..., \ell_t^N]$, and $\iota = [1, ..., 1]'$.

The ε_t^i shocks are more or less visible on days with notable news in country i such as notable central bank communications and macro data releases. Studies using event-study approaches have documented that such news have significant impact not only on yields in the country where news are emanating, but also on yields in other countries (Rogers, Scotti, and Wright, 2014; Curcuru, De Pooter, and Eckerd, 2018). But there are also many days without notable news. The key identifying assumption of this paper is that, on days without notable news, the ε_t^i shock affects the country-i yield and other country yields in the same way as on days with notable news,[14] except for an overall scale factor, i.e., the second moment of the ε_t^i shock satisfies

$$\begin{aligned} \sigma_t^2(\varepsilon^i) &= \sigma_i^{*2} \quad \text{(days with country } i \text{ news)} \\ &= \sigma_i^2 \quad \text{(days without country } i \text{ news)}, \end{aligned} \tag{7}$$

with $\sigma_i^* > \sigma_i$.[15] In addition, we assume that the local shocks ℓ_t^i and the "other global" shock η_t are homosckedastic, i.e.,

$$\sigma_t^2(\ell^i) = \sigma_{\ell,i}^2, \qquad \sigma_t^2(\eta) = \sigma_\eta^2. \tag{8}$$

Our approach is closely related to Wright (2012), who used an identification-by-heteroskedasticity technique to estimate the effects of monetary policy shocks on other asset prices.[16]

To estimate the spillover effects using the model (5), we use data on 10-year Treasury, Bund and Gilt futures to compute the change in yields over a day in a time-synchronized manner from January 2010 to August 2017.[17] Specifically, for each country we define Δy_t^i as the change in the

[14] That is, the Γ matrix is the same on days with notable news and days without notable news.

[15] As will be clear below, we do not need to make further assumptions about the distribution of ε_t^i or the other disturbances to obtain the estimates of the parameters in (5).

[16] Rigobon (2003) first introduced the idea of identifying shocks using heteroskedasticity, and Rigobon and Sack (2004) applied this technique to identify the response of asset prices to U.S. monetary policy.

[17] The start of our sample is guided by the findings in Section 1 suggesting that spillovers are more pronounced in recent years. We stop our sample before the onset of U.S.-China trade tensions, as it may be unclear how to classify news related to these events, e.g., whether they are U.S. news or global news.

10-year yield between 12 p.m. of day t and 12 p.m. of the previous day $t-1$, all in New York time.[18]

As detailed in Appendix A.2, the days with notable news for each country in our sample (i.e., U.S., Euro area, U.K.) are determined using several sources, including yield changes over a narrow window encompassing macro data releases and central bank communications, market intelligence reports, and Bloomberg news. Regardless of the specific criteria for determining these dates, there will always be some days on the margin that could be debated whether the news and the associated yield changes are "notable" enough; therefore, we also perform some robustness checks that examine how sensitive the quantities of interest, such as the fraction of the variance of U.S. yields accounted for by non-U.S. shocks, are to the specific criteria for determining notable news days.

While the model (6) can be estimated for any number of countries N, as shown in Appendix B, important parts of the model are not identified if we have $N = 2$, for example, just the U.S. and the Euro area. The case with $N = 3$ is still not fully identified, but we can impose a bit more restrictions to draw useful inferences. Our choice of $N = 3$, with the U.S., Euro area, and the U.K., for this analysis reflects practical considerations including our desire to keep the model small and manageable, and the importance of these markets for global fixed income markets. We do not include Japan (JGB market), partly because the active trading hours in Japan are far apart from the U.S. trading hours, thus there could be more concern whether the arrival of notable news in the U.S. could be reflected in the same-day synchronized yield changes in Japan, and vice versa. The "other global" shock η_t in our model attempts to capture contributions from Japan, among other possible contributions.

One advantage of using daily time-synchronized data relative to an event study analysis that uses intraday data, such as 30-minute windows encompassing announcement events, is that it may better capture the impounding of information across different bond markets. Evans and Lyons (2008) and Pasquariello and Vega (2007), for example, show that following important scheduled announcements there are further trades that reflect the process of information being discussed by market participants and incorporated into prices. In addition, many central bank communications occur over a window longer than 30 minutes. For example, both the Federal Reserve and the ECB statements have been followed by press conferences which could be also market-moving. Going

[18] Appendix A presents a more detailed description of the variables's definitions and sources.

beyond narrow windows (like 30 minutes) also allows for country-i news getting incorporated into other country yields with some possible delay. For example, FOMC statements are usually released during the afternoon hours in the U.S., which would be evening hours in Europe, during which the European market might not be as liquid as during its own daytime hours.

Excluding the "local" shock ℓ_t^i that does not affect other countries, ε_t^i is the only country-i shock in our model. At a deeper level, the ε_t^i shock can be further decomposed into a monetary policy shock, a growth shock, an inflation shock, a risk premium shock, etc.[19] These components may very well have different propagation properties (impulse responses), which are beyond the purview of the present paper. Our goal here is to analyze the *contemporaneous response* of various countries' longer-term yields to country i in a parsimonious manner; for that purpose, we are assuming that more detailed components of country-i shock have the same response patterns (the Γ matrix) insofar as the contemporaneous effect on other country yields are concerned. In the regression below, we get to examine this assumption.

2.2 Preliminary Regressions

Intuitively, on days with notable country-i news (and no news for other countries), we can expect that $\sigma_i^* > \sigma_{j=1,\ldots,N(\neq i)}, \sigma_{\ell,j=1,\ldots,N}, \sigma_\eta$. Therefore, a rough estimate of Γ_{ji} can be obtained by running the following event-study type regression,[20]

$$\Delta y_{t_i}^j = \alpha + \beta \Delta y_{t_i}^i + e_{t_i}^j, \quad \text{for } j \neq i \tag{9}$$

where t_i denotes the days in which there are notable news about country i and no important news about other countries, $\Delta y_{t_i}^j$ denotes the one-day change in country j's long-term yield on days when there are news about country i. The slope coefficient β provides a rough estimate of the response of country j's long-term yields to country i's shock, Γ_{ji}.

Table 2 presents the estimates of β for our sample of three countries, U.S., Euro area and the U.K. Panel A presents the results for all days that contain notable macro news releases or central

[19]There are many empirical studies with 10-year Treasury yields that include macroeconomic and financial variables in a VAR setup. Cieslak and Pang (2020), for example, propose a VAR model of U.S. yields and equity prices driven by monetary, growth, and risk-premium news. D'Amico, King, and Wei (2016) include U.S. and German equity and bond prices, and identify local and foreign growth, inflation, and risk aversion shocks using sign restrictions.

[20]Rigobon and Sack (2004) had noted that event study regressions are a special case of their identification-by-heteroskedasticity approach.

bank communication events, while Panels B and C present estimates with days with notable central bank communications only and days with notable macroeconomic data releases only, respectively. We exclude days when there are notable news about more than one country. The regressions are estimated using daily time-synchronized data and ordinary least squares (OLS).

The results presented in Panel A of Table 2 show that a rise in yields in the Euro area or the U.K. on days when there were important news about those economies are accompanied by an increase in U.S. yields. This positive comovement is highly statistically significant and explains a large fraction of the variation in U.S. yields over those days as suggested by the high R^2s. Similarly, as shown in the second (third) column, the point estimates suggest that Euro area (U.K.) yields also move together with U.K. (Euro area) and U.S. yields on days with important news about those economies.

Panels B and C of Table 2 present the estimates of the slope coefficient in (9) distinguishing between monetary policy communications and macroeconomic releases. The magnitude of the slope coefficients for central bank communication days are roughly similar to those obtained for days with important macroeconomic news. For example, the slope coefficient β for the spillover effect of U.S. news to Euro area and U.K. yields are 0.50 and 0.68 for central bank communications and 0.47 and 0.61 for macro news.

The magnitude of spillovers from the Euro area to the U.S. is smaller than that to the U.K. for both monetary policy communications and macroeconomic announcements. This seems sensible as the U.K. and Euro area economies are relatively more tightly connected than the U.S. and Euro area economies. Lastly, the magnitude of the spillovers from Euro area yield moves to the long-term U.S. yield in Panel B is consistent with the findings in Curcuru, De Pooter, and Eckerd (2018), namely, about half of the moves in long-term German bund yields from Euro area monetary policy shocks is transmitted to long-term U.S. yields.

These results suggest that our choice of dates with notable country i shocks that have a global effect is supported by the comovements in long-term yields. Moreover, the similar patterns of spillovers around central bank news (monetary policy shock) and around macroeconomic news (growth shock, inflation shock) provide support for grouping different type of news together to consider a single shock for country i as we do in our setting. The similar patterns may be suggesting that these conceptually distinct sources of country i shocks are impacting longer-term yields in country i and other countries in large part through the term premium channel; Hanson and Stein (2015), for example, have proposed such a mechanism based on investors "reach for yield" behavior.

2.3 Full Model Estimation

We now estimate the model shown in (5) using the generalized method of moments (GMM). To this end, note that the variance-covariance matrix of Δy_t is given by,

$$\Omega_0 = \sigma_1^2 \Gamma_{(:,1)}\Gamma'_{(:,1)} + \cdots + \sigma_N^2 \Gamma_{(:,N)}\Gamma'_{(:,N)} + \sigma_\eta^2 \iota\iota' + D([\sigma_{\ell,1}^2, ..., \sigma_{\ell,N}^2]) \quad (10)$$

$$\Omega_i = \Omega_0 + (\sigma_i^{*2} - \sigma_i^2)\Gamma_{(:,i)}\Gamma'_{(:,i)}, \quad (11)$$

where Ω_0 is the variance-covariance matrix of Δy_t on days with no notable news for any of the countries in our sample (i.e., U.S., Euro area, U.K.), Ω_i is the variance-covariance matrix on days in which there is notable news about country i but not about other countries,[21] and $D(v)$ denotes a diagonal matrix whose diagonal elements are given by the vector v.

Collectively denoting the model parameters by the vector θ, we have the following GMM moment conditions,

$$\mathbb{E}(h_t(\theta)) = 0, \quad (12)$$

with the vector h_t given by,

$$h_t = \begin{bmatrix} d_{0t} vech(\Delta y_t \Delta y_t') - (T_0/T) vech(\Omega_0(\theta)) \\ d_{1t} vech(\Delta y_t \Delta y_t') - (T_1/T) vech(\Omega_1(\theta)) \\ d_{2t} vech(\Delta y_t \Delta y_t') - (T_2/T) vech(\Omega_2(\theta)) \\ d_{3t} vech(\Delta y_t \Delta y_t') - (T_3/T) vech(\Omega_3(\theta)) \end{bmatrix} \quad (13)$$

where $\Omega_i(\theta)$ are given in (11), d_{0t} is a dummy variable that is equal to one on days with no news, and d_{it} for $i > 0$ is a dummy variable that is equal to one on days with country-i news (and no other news). The number of days with news for each country, T_i, is defined as $T_i = \sum_t d_{it}$ for $i = 0, 1, 2, 3$, and the full sample is given by $T = T_0 + T_1 + T_2 + T_3$. For our baseline estimation, T_i ($i = 1, 2, 3$) includes "macro data release" days and "central bank communication" days for

[21] There are only a small number of days when news emerge for two or more countries.

country i as determined in Appendix A.2, excluding days with news for more than one country. Days classified as "other news" days are not included in T_1, T_2, T_3 in our baseline results because the determination of these dates as "notable days" may be more open to debate, since some of them may not have a cleanly identifiable event to point to. In the end, we have 1348 days for T_0, 118 days for T_1 (U.S.), 92 days for T_2 (Euro area), and 86 days for T_3 (U.K.). In one of our robustness checks, we include "other news" days as part of T_1, T_2, T_3.

As discussed in Appendix B, for $N = 3$ we can identify Γ and $\sigma_i^{*2} - \sigma_i^2$, but we can only identify 6 out of the 7 parameters characterizing Ω_0 (i.e., $\sigma_i, \sigma_{\ell,i}$, for $i = 1, 2, 3$ and σ_η) under the current assumptions. To identify all parameters of the model, we estimate two versions of the model with the following additional restrictions:

$$\text{Version 1}: \qquad \sigma_{\ell,1} = \sigma_{\ell,2} = \sigma_{\ell,3}, \tag{14}$$

$$\text{Version 2}: \qquad \sigma_\eta = 0. \tag{15}$$

Version 1 is based on the consideration that data indicate that "global" shocks are more important than idiosyncratic (local) shocks, at least in accounting for the variance of yield changes in these countries; therefore, we consider imposing fairly simple structure on local shocks. Version 2, by setting $\sigma_\eta = 0$, allows to free up the parameters $\sigma_{\ell,1}, \sigma_{\ell,2}, \sigma_{\ell,3}$; this was motivated by our finding, discussed below, that "other global" shocks appeared to be only weakly identified in practice.

As shown by the estimates of Γ presented in Table 3, the spillovers from foreign countries to Treasury yields are statistically significant for both alternative specifications (Versions 1 and 2) and the magnitudes of the spillovers are roughly consistent with the event-study regressions (see Table 2).[22] For example, $(\Gamma_{12}, \Gamma_{13})$ estimates from Version 1 are equal to (0.53, 0.41), while for Version 2 are equal to (0.50, 0.43), both roughly similar to the slope coefficients presented in Table 2, namely, (0.60,0.56). We also find that the estimated size of ε_t^i shock on country-i news days is the largest for the U.S., and smallest for the U.K. ($\sigma_1^* > \sigma_2^*, \sigma_3^*$); this accords with the general perception that news coming from the U.S. are often more prominent than those coming from the other two economies in our sample.

[22] GMM standard errors are obtained using a Newey and West (1987) weighting matrix with 60 lags (business days). The results are not sensitive to the choice of lag length.

The estimate of σ_η in Version 1 is very small, with a large standard error, indicating that the η_t shock is not very well identified in our setup. This motivates setting $\sigma_\eta = 0$ in Version 2, which frees up $\sigma_{\ell,1}, \sigma_{\ell,2}, \sigma_{\ell,3}$. The estimate of the size of U.S. local shock ($\sigma_{\ell,1}$) in Version 2 is a bit larger than the $\sigma_{\ell,1}$ estimate in Version 1, while it is slightly smaller for the Euro area and the U.K.

A key quantity of our interest is the share of total U.S. yield variance accounted for by foreign (i.e. non-US) shocks, which can be shown to be approximately equal to

$$\lambda_{US}^{f} \approx 1 - \frac{T_1(\sigma_1^{*2} + \sigma_{\ell,1}^2) + (T_0 + T_2 + T_3)(\sigma_1^2 + \sigma_{\ell,1}^2)}{T_0[\Omega_0]_{11} + T_1[\Omega_1]_{11} + T_2[\Omega_2]_{11} + T_3[\Omega_3]_{11}}, \tag{16}$$

where $[\Omega_i]_{11}$ denotes the (1,1) element of matrix Ω_i, and $\Omega_0, \Omega_1, \Omega_2, \Omega_3$ are given in equations (10) and (11).

The estimated parameters in Table 3 imply a λ_{US}^{f} value equal to 0.20 and 0.22 for Version 1 and Version 2, respectively. These values appear fairly robust to the definition of "notable news days." For example, when we estimate the model redefining the news days such that we have a smaller number of Euro area news, we obtain λ_{US}^{f} value of 0.22 and 0.25 for Version 1 and 2, respectively.[23] In addition, when we implement our model including "other news" days in T_1, T_2, T_3, we obtain λ_{US}^{f} estimates of 0.25 for both Version 1 and Version 2. In sum, about 20 to 25% of U.S. 10-year yield variations are accounted for by foreign (non-U.S.) shocks. This is a non-negligible magnitude, and indicates a significant amount of foreign influence on U.S. yields. In fact, these are likely *underestimates* of the true number: we should expect other countries, including Japan and China, to also have some effect on U.S. yields, but the "other global" factor is not pinned down well in our setting likely due to the limitations of the model.[24]

Finally, for a complete picture, we note that the corresponding numbers for the share of foreign shocks (non country-i shocks) in country i variance for the Euro area and the U.K.—λ_{EA}^{f} and λ_{UK}^{f}—based on the estimates in Table 3, are 0.24 and 0.23 for λ_{EA}^{f} with Version 1 and Version 2, respectively, and 0.50 for λ_{UK} with both Version 1 and Version 2. So the share of Euro area yield

[23] In this exercise we used a more stringent criteria for defining "notable" ECB news, which reduced T_2 from 92 to 69.

[24] Utilizing the fact that introducing time-variation in volatility helps with identification (Sentana and Fiorentini, 2001), we have also explored a richer version of the model in which "the other" global shock η_t has a GARCH structure, with a QML estimation. We find in that case that the estimated η shocks are often small in magnitude, but can be sizeable at certain times during our sample period, including the 2010–2011 period (Euro area debt crisis) and 2015 (PBOC's yuan devaluation).

variance accounted for by foreign shocks is comparable to that of the U.S., while the corresponding estimate for the U.K. is notably higher, with half of U.K. yield variance being attributed to non-UK shocks. The higher share for U.K. variation explained by foreign shocks seems plausible in light of the smaller size of the U.K.'s economy relative to the U.S. and the Euro area.

3 Evidence of Foreign Spillovers from Predictive Regressions

This section explores whether predictable variations in long-term Treasury yields are associated with low levels of yields in advanced foreign economies relative to U.S. yields. In particular, we test if the knowledge of the spread between U.S. and foreign long-term yields is useful in forecasting changes in long-term Treasury yields.

3.1 Constructing the U.S.–Foreign Long-Term Yield Spread

We define the long-term yield on foreign sovereign debt as the GDP-weighted average of German, Japanese, and U.K. 10-year zero-coupon yields,

$$y_{f,t} = \sum_{c=1}^{M} w_{c,t}\, y_{c,t}^{(10)} \tag{17}$$

where the weight for country c is $w_{c,t} = \frac{\text{GDP}_{c,t}}{\sum_c^M \text{GDP}_{c,t}}$ and $M = 3$. Since GDP figures are quarterly and released with a delay of at least a quarter, the weight applied to the weekly yields are constant within each quarter and correspond to GDP figures from two quarters back to ensure that weights are known to the investor at time t. On average, over the 2000-2019 period, the weight for Germany is 0.32, the weight for Japan is 0.45, and the weight for the U.K. is 0.23. These weights are relatively constant throughout our sample period.

The U.S.–foreign long-term yield spread is computed as the spread between the U.S. 10-year zero-coupon Treasury yield and the GDP-weighted foreign yield. Panel (a) of Figure 3 displays the level of the long-term foreign yield along with the long-term U.S. yield from 2000 to 2019. As shown in this figure, the long-term foreign yield fell from 3.8 percent to 0.1 percent over this period, a decline of more than 350 basis points. Similarly, the yield on long-term Treasury securities declined from 6.7 percent to 2.0 percent over the same period, reaching multi-decade lows. The correlation coefficient between weekly changes in U.S. and foreign yields is 0.8, suggesting the presence of common factors driving short-run fluctuations in U.S. and foreign long-term yields. Panel (b) of Figure 3 displays the spread between U.S. and foreign long-term yields. As can be seen in this figure, the spread between these yields is positive throughout our sample and averages about 1.5 percent with a standard deviation of 0.5 percent and a first-order autocorrelation coefficient of 0.989.

We consider the predictive regressions discussed below using the U.S–foreign long-term yield spread (and control variables) as a simple and tractable way to explore the influence of foreign yields on the predictable variation of U.S. yields. An alternative approach in the literature for examining the influence of other countries' interest rates on a country's interest rate is the cointegration approach, as in Kirchgässner and Wolters (1993) and Chinn and Frankel (1995). In our context, that approach would take the form

$$\Delta y_t = \alpha - B z_t + \zeta_1 \Delta y_{t-1} + \zeta_2 \Delta y_{t-2} + \cdots + \epsilon_t, \tag{18}$$

where y_t is the vector of yields of various countries, e.g., $y_t = [y_t^{US}, y_t^{EA}, y_t^{UK}, ...]'$, and z_t is cointegrating vector, $z_t = A' y_t$. If the cointegrating vector takes the form

$$z_t = y_t^{US} - a_1 y_t^{EA} - a_2 y_t^{UK} - \cdots, \tag{19}$$

that would probe a similar effect as our predictive regressions.

3.2 Predictability of Intraday Moves in Long-Term Yields

Table 4 reports the results from predictive regressions of the form,

$$\Delta y_{t+1} = \alpha + \beta \left(y_t - y_{f,t} \right) + \gamma' \boldsymbol{x}_t + \epsilon_{t+1}, \tag{20}$$

and,

$$\Delta y_{\mathrm{a},t+1} = \alpha_{\mathrm{a}} + \beta_{\mathrm{a}} \left(y_t - y_{f,t} \right) + \gamma'_{\mathrm{a}} \boldsymbol{x}_t + \epsilon_{\mathrm{a},t+1} \tag{21}$$

$$\Delta y_{\mathrm{na},t+1} = \alpha_{\mathrm{na}} + \beta_{\mathrm{na}} \left(y_t - y_{f,t} \right) + \gamma'_{\mathrm{na}} \boldsymbol{x}_t + \epsilon_{\mathrm{na},t+1}. \tag{22}$$

where Δy_{t+1} is the change in the long-term U.S. yield, $\Delta y_{\mathrm{a},t+1}$ is the yield changes around macroeconomic and monetary policy announcements, and $\Delta y_{\mathrm{na},t+1}$ is the yield changes outside of these windows.[25] The main predictive variable is the spread between the 10-year Treasury yield and the 10-year foreign yield $(y_t - y_{f,t})$. The vector $\boldsymbol{x}_t$ contains bond return forecasting variables identified in the literature that at the same time capture the U.S. business cycle.

As in Section 1, we use intraday data on the yield on the 10-year on-the-run Treasury security to compute the weekly cumulative change in the long-term yield around major macroeconomic and policy announcements $\Delta y_{\mathrm{a},t+1}$, and outside announcement times $\Delta y_{\mathrm{na},t+1}$. The weekly change in the 10-year yield is the the sum of changes during announcement times and outside announcement times. The vector of controls $\boldsymbol{x}_t$ contains the 10-year forward rate spread $(f_t - r_t)$ as in Fama and Bliss (1987), the near-term forward spread $(\mathbb{E}_t(r_{t+j}) - r_t)$ of Engstrom and Sharpe (2019) as a measure of expectations for the near-term path of the U.S. monetary policy rate, the yield spread between Aaa-rated corporate bonds and Treasury securities $(y_{\mathrm{Aaa},t} - y_t)$ to capture shifts in domestic demand for the liquidity and safety of long-term Treasury securities documented in Krishnamurthy and Vissing-Jorgensen (2012), and the effective duration of outstanding MBS (MBSDUR_t) to control for shifts in the demand for long-term Treasury securities of U.S. MBS investors in response to changes in expectations for future household refinancing documented in Hanson (2014) and Malkhozov, Mueller, Vedolin, and Venter (2016). We also include an indicator variable that equals to one if there was a Treasury auction over the forecasting period to capture the change in yields due to the higher liquidity of the newest issued security (Krishnamurthy, 2002). The regressions are estimated using weekly data from January of 2000 to December of 2019 and using ordinary least squares (OLS). We report t-statistics based on Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals.

[25] Recall from equation (1) that $\Delta y_{t+1} = \Delta y_{\mathrm{a},t+1} + \Delta y_{\mathrm{na},t+1}$. To our knowledge, Faust and Wright (2018)) was the first to study the predictability of bond returns over announcement windows and non-announcement windows separately.

Panel A of Table 4 presents the estimates for the overall weekly change in the long-term Treasury yield, namely, equation (20). The results in column (1) show that the spread between U.S. and foreign long-term yields has a negative and statistically significant coefficient. The coefficient of -1.67 in column (1) of Panel A suggests that after a 100 basis point increase in the spread between U.S. and foreign long-term yields, investors expect Treasury yields to decline by 1.7 basis points over the following week. While this effect is small, the persistence of the U.S.–foreign yield spread means that a widening of the spread can lead to economically significant declines in Treasury yields over the following months. In particular, the coefficient estimates suggest that a one standard deviation increase in the U.S.–foreign yield spread, around a 50 basis point move, is expected to be followed by a 36 basis point decline in Treasury yields over the next year.[26]

The results in column (2) of Panel A show that the predictive power of the spread between U.S. and foreign long-term yields is robust to controlling for the predictability of the forward spread, the near-term spread, the Aaa-Treasury spread, and MBS duration. Interestingly, the coefficient on the U.S.–foreign yield spread becomes about twice more negative (-3.62) once we control for variables that are not only predictors of bond returns but are also linked to the U.S. business cycle, relative to the specification without any controls. Including these variables likely reduces the noise in yield fluctuations unrelated to foreign fluctuations and improves the predictive power of the U.S.–foreign yield spread.

The predictability of this single factor, as captured by the R^2, is not only comparable with that of a regression that only includes the set of control variables, but it also adds predictive power over and above the other bond return predictors included in the regression. In particular, the R^2 from a regression using the the forward spread, near-term spread, the Aaa-Treasury spread, and MBS duration is 0.55%, as shown in Column (3) of Panel A. If we include the U.S.–foreign yields spread as a regressor, as shown in column (2), the R^2 increases threefold to 1.53%.

Column (1) of Panels B and C of Table 4 shows that all the forecasting power of the spread between U.S. and foreign long-term yields is explained by its ability to forecast changes in the long-term Treasury yield in windows outside of domestic macroeconomic and policy announcements. In particular, in panel B we find that the U.S.–foreign yield spread does not seem to predict

[26] Assuming that the U.S.–foreign long-term yield spread follows a first-order autoregressive process with autoregressive coefficient ρ, the cumulative effect of a move of size σ in this spread translates into an expected move in U.S. yields of about $\beta \frac{1-\rho^n}{1-\rho} \sigma$ over the next n weeks.

yield fluctuations around macroeconomic announcements as the coefficient on this spread is not statistically or economically significant, whereas the spread between U.S. and foreign yields, as shown in column (1) of Panel C, is a strong predictor of future changes in the long-term yield outside of windows bracketing the release of key macroeconomic data. Results reported in column (2) of Panel B show that adding control variables does not change the lack of predictability of the U.S.–foreign yield spread of long-term yield changes around important U.S. economic announcements. In contrast, as shown in column (2) of Panel C, adding controls increases the statistical and economic significance of the U.S.–foreign yield spread as predictor of long-term yield changes outside of windows with domestic economic releases.

Another way to further assess the predictive power of the U.S.–foreign yield spread is to use the first three principal components (PCs) of the U.S. yield curve as control variables. While the three PCs are less theoretically motivated than the controls we use in Table 4, these three components, often labeled level, slope, and curvature, explain almost all of the variation in yields (see, for example, Litterman and Scheinkman (1991)), and have been shown to forecast bond returns around macroeconomic data releases (see Faust and Wright (2018)). These PCs can be also viewed as encompassing well known yield curve variables, such as the short-term yield, the Cochrane and Piazzesi (2005) factor, and some control variables used above like the forward spread and the near-term spread. The results in column (1) in Panels A, B and C of Table 5 show that indeed the three principal components (L_t, S_t, and C_t) are informative predictors of weekly Treasury yield changes at times of news announcements and at times outside announcement windows. More importantly, column (2) of Panel C shows that the ability of the U.S.–foreign yield spread to predict future changes in long-term Treasury yields outside macro announcements is robust to controlling for the predictive power of the three principal components. For one thing, the coefficient on the U.S.–foreign yield spread remains highly significant and of roughly the same magnitude as in the specifications presented in panel C of Table 4. For another, including the U.S.-foreign yield spread increases significantly the R^2 of the regression from 0.47% to 1.30%. As in Table 4, we continue to find a lack of predictive power of the U.S.–foreign yield spread to changes in the long-term Treasury yield around macro announcements. [27]

[27] Using the changes in the zero-coupon 10-year Treasury yield as dependent variable along with Bauer and Hamilton (2018) bootstrap estimates for the critical values of the t-statistics, we also find that the U.S.–foreign yield spread is strongly statistically significant (p-value = 0.011). Similarly, the large rise in the R^2 is quite implausible under the

3.3 Does the Predictive Power of the U.S.–foreign Long-Term Yield Spread Vary Over Time?

From a relative pricing perspective, a widening in the U.S.–foreign longer-term yield spread should predict declines in U.S. yields, regardless of whether the widening is due to U.S. developments or foreign developments. At a more detailed level, while both negative foreign news that depress foreign yields and positive news that raise U.S. yields would lead to a widening of the U.S.–foreign yield spread, the predictability of this spread could be different depending on the underlying factors driving the moves in the spread. We take a simple approach to empirically examine whether the source of movements in the U.S.–foreign spread matters, and estimate a conditional version of (20), (21), and (22) allowing the regression coefficient β to be a linear function of the overnight variance of yields, namely,

$$\Delta y_{t+1} = \alpha + \Big(\beta_0 + \beta_1 \operatorname{Var}_t(y_{\mathrm{o}})\Big) \left(y_t - y_{f,t}\right) + \gamma' \boldsymbol{x}_t + \epsilon_{t+1}, \tag{23}$$

where $\operatorname{Var}_t(y_{\mathrm{o}})$ is the variance of changes in U.S. yields overnight standardized to have a zero mean and a standard deviation equal to one. Intuitively, in periods when the overnight variance is higher than usual, one may expect the spread between U.S. and foreign long-term yields to contain more information concerning the economic outlook abroad than in times when the overnight variance is lower than usual. As a consequence, the U.S.–foreign long-term yield spread would be expected to weigh more heavily on the U.S. yield when overnight volatility is higher than usual.

The variance of yields overnight is computed using an exponentially weighted moving average of squared changes in long-term yields overnight.[28] The vector of control variables, $\boldsymbol{x}_t$, includes the forward spread, the near-term spread, the Aaa-Treasury spread, MBS duration, overnight volatility, and an indicator variable that equals to one if there was a Treasury auction. Our sample is weekly from January of 2000 to December of 2019 and our inference is performed using Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals.

null hypothesis that the U.S.–foreign yield spread does not have any incremental predictive power, namely, the R^2 increase of 1.11 is well outside the 95% bootstrap confidence interval $[-0.096, 0.413]$.

[28] We define overnight yield changes as the change in the 10-year Treasury yield between 8 a.m. and 5 p.m. of the previous business day and cumulate these overnight changes over a week. The variance of yields overnight is computed using weekly overnight changes and setting the smoothing parameter such that the overnight variance has a half-life of around 25 weeks.

The results presented in column (1) of Table 6 show that the long-term Treasury yield is expected to experience a more pronounced decline following a widening of the U.S.–foreign long-term yield spread, when the overnight variance is above its average level. In particular, we find that the coefficient on the interaction term between overnight variance and the U.S.–foreign yield spread is negative and statistically significant. As shown in column (2) of Table 6 the coefficient on the interaction term remains negative and statistically once we control for known predictors of U.S. Treasury returns. The point estimates suggest that when overnight variance is one standard deviation above its long-run level, the widening of the U.S.–foreign long-term yield spread has a compressing effect that is about two-thirds larger relative to usual times. All in all, the evidence of time-varying predictability suggests that movements in the U.S.–foreign term spread on days with a larger flow of macroeconomic and policy news from abroad leads to larger subsequent moves in U.S. yields.

3.4 The Predictability of U.S. Forward Rates

Our predictability regressions show that the U.S.–foreign long-term yield spread predicts future movements in long-term U.S. yields, in particular those that are not linked to the release of U.S. macroeconomic news and in periods when the overnight variance is high. Here we ask whether these results reflect predictable movements in term premia or predictable movements in expected future short rates. We start from the premise that changes in distant nominal forward rates are mostly driven by time-varying term premia and estimate the predictability of forward rates for different horizons.[29] If the U.S.–foreign yield spread were informative about future short rates, the predictability on long-term rates would arise mainly from short-forward rate components of long-term yields. In contrast, if we find that the evidence for predictability gets stronger as we increase the forward rate horizon, that can be suggestive evidence that the U.S.–foreign yield spread is more informative about term premia than about future short rates.

To perform this forecasting exercise we use data on nominal Treasury zero-coupon yields from Gürkaynak, Sack, and Wright (2007) to construct U.S. forward rates. As in Section 3.2, the

[29] Various studies decomposing distant-horizon forward rates into short-rate expectations and term premia, including Kim and Wright (2005), find that distant-horizon nominal forward rates are in large part driven by movements in term premia. In addition, Hanson and Stein (2015) show that around FOMC announcements, when investors receive information about the path of policy rates, far-forward rates are mainly driven by news about future term premia.

U.S.–foreign long-term yield spread is the key explanatory variable in the regressions and we control for the forward spread, the near-term spread, the Aaa-Treasury spread, and MBS duration. We estimate the predictive regressions using weekly data from 2000 to 2019. We obtain estimates of the coefficients using OLS and perform inference using Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals.

We start by documenting that the predictability reported in Section 3.2 using on-the-run yields is also evident when we use the 10-year zero-coupon Treasury yield,

$$\Delta y_{t+1}^{(10)} = \alpha + \underbrace{\beta}_{-4.31 \text{ t-stat}=-3.60} \left(y_t - y_{f,t}\right) + \gamma' \boldsymbol{x}_t + \epsilon_{t+1}, \quad R^2 \times 100 = 1.7 \tag{24}$$

The coefficient on the spread between U.S. and foreign long-term yields is negative and the t-statistic shows that we can safely reject the null hypothesis that the U.S.–foreign yield spread does not predict future movements in long-term Treasury yields. The predictive R^2 is about the same magnitude as the one reported in Table 5 for the on-the-run yield.

The 10-year zero coupon yield can be decomposed into 1-year forward rates as follows

$$y_t^{(10)} = \frac{1}{10} \sum_{n=1}^{10} f_t^{(n)} \tag{25}$$

where $f_t^{(n)}$ is 1-year forward rate for the n-th year, with $f_t^{(1)}$ denoting the 1-year yield. We now turn to the predictability of these forward rates and estimate,

$$\Delta f_{t+1}^{(n)} = \alpha_n + \beta_n \left(y_t - y_{f,t}\right) + \gamma_n' \boldsymbol{x}_t + \epsilon_{t+1}^{(n)}, \tag{26}$$

for $n = 2, \ldots, 10$. Figure 4 plots the key coefficient of interest β_n along with a 90% confidence intervals and the associated R^2 for maturities $n = 2, \ldots, 10$.[30]

Figure 4(a) shows that a widening of the U.S.–foreign long-term yield spread predicts a subsequent decline in forward rates, and the predicted decline is more pronounced as we move towards far-forward rates. In particular, the estimated coefficients suggest that a 100 basis point widening

[30] We do not report the results for $n = 1$ because, as shown in Swanson and Williams (2014), short-term rates were constrained by the ELB and this constraint might bias our estimates of β_n. The ELB effect is less of a concern for longer maturities and, in unreported results, we show that the results reported in this section are robust to using a sample that ends before the ELB was said to be binding, i.e., 2011.

of the U.S.–foreign yield spread is followed by a 5 basis point decline the distant forward rates, while it predicts only a 2 basis point decline in the forward rate 1-to-2-years ahead.

Figure 4(b) displays the R^2s of the predictive regression (26) and those of predictive regressions that only include the control variables. The additional predictive power added by including the U.S.–foreign yield spread can be gauged by the difference in R^2s. Figure 4(b) shows that including the U.S.–foreign yield spread as a predictor increases the R^2 and the additional forecasting power is higher for more distant forward rates. These empirical results suggest that the U.S.–foreign long-term yield spread is more informative about term premia than about future short rates, supporting the hypothesis that spillovers to U.S. long-term rates are likely occurring through bond risk premia.[31]

3.5 The Predictability of Returns on Long-Term Treasury Over Long-Term Foreign Bonds

One potential explanation for the ability of the U.S.–foreign yield spread to predict the long-term U.S. yield might be related to shifts in demand for long-term Treasury securities. A decline in long-term foreign yields could boost demand for higher-yielding U.S. securities as investors would be attracted to the higher expected returns from investing in U.S. relative to foreign sovereign long-term bonds.

To test this hypothesis, we explore the ability of the U.S.–foreign long-term yield spread to predict the excess return on a 10-year Treasury security over a 10-year foreign bond by running the following predictive regression,

$$rx_{t\to t+\tau} - rx_{\mathrm{f},t\to t+\tau} = \alpha + \beta\left(y_t - y_{f,t}\right) + \gamma' \boldsymbol{x}_t + \epsilon_{t+\tau}, \tag{27}$$

where $rx_{t\to t+\tau}$ is the return on a 10-year Treasury in excess of the U.S. short-term rate, and $rx_{\mathrm{f},t\to t+\tau}$ is the excess return on a 10-year foreign bond that we define as the GDP-weighted excess

[31] Kearns, Schrimpf, and Xia (2020) and Dilts Stedman (2020) also conclude that the yield curve spillovers largely occur through the term premium component of yields.

return on German, Japanese, and U.K. 10-year sovereign debt,

$$rx_{\mathrm{f},t\to t+\tau} = \sum_{c=1}^{3} w_{c,t} rx_{c,t\to t+\tau}. \tag{28}$$

The key predictive variable is the spread between U.S. and foreign long-term yields ($y_t - y_{f,t}$), and $\boldsymbol{x}_t$ is a vector of control variables.

The coefficient estimates of (27) are obtained using weekly data from January of 2000 to December of 2019 and for holding period horizons of 1-week, 4-weeks, and 12-weeks (τ = 1, 4, 12). Returns on U.S. and foreign 10-year bonds are computed using zero-coupon U.S., German, Japanese and U.K. yields. To perform inference we rely on Newey and West (1987) standard errors with at least 26 lags to deal with the autocorrelation of the residuals and overlapping observations when $\tau > 1$. As in Section 3.2, we include in $\boldsymbol{x}_t$ the 10-year forward rate spread ($f_t - r_t$), the near-term forward spread ($\mathbb{E}_t(r_{t+j}) - r_t$), the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{\mathrm{Aaa},t} - y_t$), and the effective duration of MBS (MBSDUR_t).

Column (1) of Table 7 includes the U.S.–foreign yield spread as the only explanatory variable for different horizons. The coefficient estimates show that the U.S.–foreign long-term yield spread is a strong predictor of the return on long-term U.S. bonds relative to long-term foreign bonds with R^2s equal to 1 percent, 4 percent, and 10 percent for the 1-, 4-, and 12-week holding period horizons, respectively. The estimated coefficient is positive and statistically significant for all horizons suggesting that a widening of the spread between U.S. and foreign long-term yields leads to higher expected returns on long-term Treasury securities relative to long-term foreign bonds. The coefficient on the U.S.–foreign yield spread, as shown in column (2) of Table 7, becomes more negative and remains highly statistically significant when we add control variables known to forecast U.S. bond returns.

We also consider the predictability of the return on a 10-year Treasury yield in excess of the currency-hedged return on a 10-year foreign bond. In particular, we compute the currency-hedged τ-period return on country c as,

$$ret^{h}_{\mathrm{c},t\to t+\tau} = ret_{\mathrm{c},t\to t+\tau} + (e_t - f_t) \tag{29}$$

where $ret_{\mathrm{c},t\to t+\tau}$ is the country c's bond return, e_t is the spot exchange rate expressed as foreign currency per USD, and f_t is the τ-period forward exchange rate at time t. As in (28), we define the currency-hedged return on a long-term foreign bond as the GDP-weighted of currency-hedged returns on 10-year German, Japanese, and U.K. bonds.[32]

Table 8 presents predictive regressions where the dependent variable is the return on a 10-year Treasury security in excess of the currency-hedged return on a 10-year foreign bond, namely, $ret_{t\to t+\tau} - ret^{h}_{\mathrm{f},t\to t+\tau}$. Consistent with the results presented in Table 7, we find that the coefficient on the U.S.–long-term yield spread is positive and highly statistically significant, and remains a strong predictor even after controlling for the forward rate spread, the near-term forward spread, the yield spread between Aaa-rated corporate bonds and Treasury securities, and the effective duration of MBS. Overall, we find suggestive evidence that returns on long-term U.S. bonds are expected to rise relative to currency-hedged returns on foreign bonds when the U.S.–foreign long-term yield spread widens, likely boosting the demand for U.S. Treasury securities and pushing down U.S. bond yields.

3.6 Robustness Checks

This section summarizes additional exercises we performed to examine the robustness of our results. Full details are presented in the Appendix. We show that the predictability we document is robust to reasonable variations to the way we compute the foreign yield. In particular, we show that our results are robust to using the equally-weighted average of German, Japanese, and U.K. long-term yields as well as including yields on Swiss and French sovereign debt. Using these alternative proxy measures for the foreign yield produces results qualitatively and quantitatively similar to our baseline estimates. We also perform our predictive regressions using monthly data and adding an additional eight years of monthly observations. Consistent with the results using weekly data, we continue to find that a widening of the U.S.–foreign yield spread predicts future declines zin U.S.

[32]If the covered interest parity holds, we can replace $e_t - f_t$ with the spread in short-term rates $r_t - r_{c,t}$ and show that the return on a 10-year Treasury in excess of the currency-hedged return on a foreign bond is equivalent to the excess excess return on a long-term Treasury security over a long-term foreign bond,

$$ret_{t\to t+\tau} - ret^{h}_{\mathrm{c},t\to t+\tau} = \underbrace{(ret_{t\to t+\tau} - r_t)}_{rx_{t\to t+\tau}} - \underbrace{(ret_{\mathrm{c},t\to t+\tau} - r_{\mathrm{c},t})}_{rx_{\mathrm{c},t\to t+\tau}},$$

which is the dependent variable in the predictive regression (27).

long-term yields.

4 Concluding Remarks

Yield spillover effects have been traditionally thought to run mainly from the U.S. to other countries. In this paper, we present various pieces of evidence suggesting that there are also significant spillovers from foreign economies to the U.S. through changes in long-term yields, and that their importance has grown over time. We show that the share of U.S. yield variation accounted for by overnight yield changes—a rough proxy for foreign contribution to U.S. yield movements—has increased since 1990s. Using synchronized daily data on 10-year yield changes in the U.S., Euro area, and U.K., and a selection of dates with notable yield moves in these countries, we estimate an identification-by-heteroskedasticity model, which indicates that at least 20 to 25% of daily variations in 10-year U.S. yields in recent years are due to foreign shocks. The spillover effects occur not only through contemporaneous yield changes but also through predictable yield changes. We find that following a widening of the U.S.–foreign long-term yield spread Treasury yields tend to decline.

Conceptually, the yield spillover effects we document here appear to operate mainly through the term premium channel as opposed to expectations channel. For example, negative news in Europe would depress European yields, which in turn would make U.S. Treasuries relatively more attractive depressing U.S. term premiums, as opposed to negative European news darkening the U.S. economic outlook and lowering the expected path of the federal funds rate. This observation is consistent with the greater degree of comovement between longer-term international yields than shorter-maturity international yields as well as the evidence presented in Section 3. Still, in light of the limited amount of existing work in this area, more remains to be learned about the mechanisms underlying the yield spillover effects and their ramifications for understanding U.S. and international yield curve movements.

Appendix

A Data

A.1 Financial Markets Data

We collect intraday data on yields on the most recently issued 10-year Treasury security, namely, yields on the on-the-run Treasury security, from Bloomberg. Zero-coupon yields on the 10-year Treasury are obtained from the smoothed yield curve of Gürkaynak, Sack, and Wright (2007) updated by the Federal Reserve Board.

Our empirical analysis also relies on long-term yields for advanced foreign economies. In particular, we collect zero-coupon yields on 10-year sovereign debt for Germany, Japan, the U.K., Switzerland and France. Our dataset contains weekly data from January of 2000 to December 2019 and monthly data that goes back to January of 1992. The data comes from the the Bundesbank for Germany, the Japanese Ministry of Finance for Japan, the Bank of England for the U.K., and Refinitiv through Haver Analytics.

To construct synchronized daily changes in long-term yields on U.S., Euro area, and U.K. sovereign debt, we collect intraday prices of the 10-year Treasury note, German Bund, and U.K. Gilt futures contracts. The daily change in the country's long-term yield is obtained as the percentage change in the futures price between 12 p.m. Eastern Time (EST) of the current day and 12 p.m. EST of the previous day multiplied by minus one, and by the inverse of the modified duration of the cheapest to deliver. The data are available from January 2010 to August 2017.

Lastly, we also collect data on three-month Treasury yields from CRSP;[33] yields on Aaa-rated 10-year corporate bond yields from Moody's; quarterly GDP data for each country expressed in U.S. dollars at purchasing power parity from the OECD (2020); Euro, Japanese yen, British pound sterling, and USD LIBOR as well as data on spot and forward exchange rates are obtained from Bloomberg Finance, LP. The date and the time of macroeconomic announcements and central bank policy decisions in the U.S. are also obtained from Bloomberg Finance, LP.

[33] Center for Research in Security Prices, CRSP U.S. Treasury Database, Wharton Research Data Services, http://www.whartonwrds.com/datasets/crsp/.

A.2 Notable News Days

In this Appendix, we elaborate on our determination of "notable news days" for the U.S., Euro area, and the U.K. for the sample period January 2010 to August 2017. To identify notable news dates for each country, we use several sources including intraday data on the 10-year yield, Citi Economic Surprise indices, Bloomberg and other financial news, and internal daily market intelligence reports prepared by FRBNY.

More specifically, we search financial news databases as well as internal market intelligence reports to find days in which notable moves in Treasury, Bund, and/or Gilts markets are attributed to a specific development in the U.S., Euro area, or the U.K. In the case of the U.S., we also look at 10-year Treasury yield changes over 30-minute windows surrounding major scheduled announcement events to pick out those with sizeable yield changes. And we examine Citi Economic Surprise indices for U.K. and Euro area to pick out days in which these indices displayed notable changes, and check with news sources to determine if those dates could indeed be viewed as notable news days.

We classify "notable news" days as "data release" days, "central bank communication" days, or "other news" days. The "data release" days are dates in which there were sizable market reactions to scheduled macro data releases. In the U.S., many of notable "data release" days are the days of the employment report, but there were also days in which other releases, such as retail sales, ISM, CPI, had notable market reactions. In the Euro area, there are relatively fewer notable data release days, but data such as PMIs (Euro area's and member countries') and CPIs, have at times generated significant market reactions. In the U.K., data releases including labor market data, GDP, and CPI, have had notable market reactions.

Notable "central bank communication" days are days with market-moving communications from country-i central banks, i.e., Federal Reserve, ECB, and Bank of England. Often times, these were days with announcements following scheduled committee meetings, but some of these days pertain to other type of central bank communications, such as the releases of the minutes and policymaker's speeches testimonies.

The remaining notable news days are grouped as "other news" days. These include days with notable identifiable news other than data releases and central bank communications, for example, the 2016 presidential election day and some fiscal policy news days. Also included are some days in which it was not easy to point to a specific identifiable event but market commentaries characterized as having been influenced by country i news.[34] In addition, some of these "other news" days had data releases whose 30-minute event window yield changes were not large, but news reports and market commentaries had interpreted as important driver of yields those days. We list the dates identified as days with notable news in Table A.1.

B Identification

In this appendix, we discuss the identification of the model consisting of eq. (5), (7), and (8). Note that from eq. (11) for $i = 1, ..., N$, we can identify $\Gamma_{(:,i)}$ and $\sigma_i^{*2} - \sigma_i^2$, with Γ_{ii}'s normalized to 1. Therefore, in the equation

$$\Omega_0 = \sigma_1^2 \Gamma_{(:,1)} \Gamma'_{(:,1)} + \cdots + \sigma_N^2 \Gamma_{(:,N)} \Gamma'_{(:,N)} + \sigma_\eta^2 \iota\iota' + D([\sigma_{\ell,1}^2, ..., \sigma_{\ell,N}^2])$$

we can think of $\Gamma_{(:,i)}$ as known, and treat $\sigma_1^2, ..., \sigma_N^2, \sigma_\eta^2, \sigma_{\ell,1}^2, ..., \sigma_{\ell,N}^2$ as unknowns to solve for. This is $2N + 1$ unknowns, whereas there are $N(N+1)/2$ equations (the number of unique elements of the Ω_0 matrix). The full model is thus identified for values of N such that $N(N+1)/2 \geq 2N + 1$, i.e., for $N \geq 4$. Note that for $N = 2$, $2 \cdot 3/2 = 2 < 2 \cdot 2 + 1 = 5$, and for $N = 3$, $3 \cdot 4/2 = 6 < 2 \cdot 3 + 1$. In order to have an econometrically identified model for the case of our interest ($N = 3$), we impose additional restrictions, as in eq. (14) or eq. (15).

Note that additional moment restrictions can be obtained by considering days in which more than one country has a notable news. For example, denoting by $\Omega_{i,j}$ the variance covariance matrix for the days on which an i, j pair of countries had notable news, we have

$$\Omega_{i,j} - \Omega_0 = (\sigma_i^{*2} - \sigma_i^2) \Gamma_{(:,i)} \Gamma'_{(:,i)} + (\sigma_j^{*2} - \sigma_j^2) \Gamma_{(:,j)} \Gamma'_{(:,j)}.$$

However, such conditions do not help further identify the model, as they can be viewed as linear combinations of existing moment conditions (eq. (10) and eq. (11)).

[34] For example, during the height of European debt crisis, some days with large change in yields were attributed to developments pertaining to that crisis (therefore a Euro area news), which had "drips" of related news coverage, rather than a single notable event such as a government announcement or officials's speeches/comments.

C Robustness of Predictive Regressions

This Appendix reports in more detail the results of several robustness checks to the predictability of the U.S.–foreign long-term yield spread documented in Section 3.

C.1 Alternative Measures of the Long-term Foreign Yield

The key explanatory variable in the predictive regressions is the spread between U.S. 10-year yield and our measure of foreign long-term yields. Our empirical estimates are based on a GDP-weighted average of German, Japanese, and U.K. 10-year zero-coupon yields. In this section we explore the robustness of the predictive power of the U.S.–foreign long-term yield spread to sensible variations to computing the foreign long-term yield.

Table A.2 presents the estimates of predictive regressions that use the foreign yield computed as an equally-weighted average of German, Japanese, and U.K. long-term yields. The results in column (1) of Panel A show that the coefficient on the U.S.–foreign long-term yield spread has also a negative and statistically significant coefficient and, as shown in column (1) of Panels B and C, its predictive power is borne out of its ability to forecast changes in long-term Treasury yields in windows outside of economic announcements. The results in column (2) confirm our finding that the widening of the U.S.–foreign long-term yield spread has a more pronounced effect on U.S. Treasury long-term yields when overnight volatility is above its average level than when it is at its long-run level.

To check if the results are robust to using a larger set of countries, we include data on Swiss and French sovereign debt to compute the the foreign yield. As shown in Table A.3, our finding that *low* levels of foreign yields put downward pressure on U.S. long-term yields remains robust to using a larger set of countries to construct the foreign long-term yield. We continue to find that the U.S.–foreign yield spread does not seem to predict yield fluctuations around economic announcements, whereas it is a strong predictor of U.S. yields outside of windows bracketing the release of key macroeconomic data.

C.2 Data Frequency

The predictive regressions use weekly data to document the predictive ability of the U.S.–foreign spread. We have repeated our regressions using lower-frequency data, namely, monthly. We present in Table A.4 the estimated coefficients that are a counterpart of those shown in Table 4 with weekly data. The results are largely consistent with the estimates using weekly data; the predictability of the U.S.–foreign long-term yield spread is statistically and economically significant and robust to controlling for usual predictors of bond returns (Panel A). We also find that the predictability is significant for moves outside major economic announcements but not for changes in yields around macroeconomic and monetary policy announcements.

C.3 Small Sample

Given the persistence of the U.S.–foreign yield spread, one particular concern might be that the results might be suffering from small-sample bias. We extend our sample in slightly more than eight years to cover May of 1991 to December of 2019. Going to a monthly frequency and extending the data also reduces the persistence of the U.S.–foreign yield spread to levels below 0.95, which Ferson, Sarkissian, and Simin (2003) show is a threshold under which the t-statistics are well behaved and we can undertake inference on the coefficients of persistence regressors. In this exercise we use as control variables the three PCs of the U.S. yield curve since we have data on this variables for the sample under consideration here.

As shown in Table A.5, using a longer sample we continue to find that the U.S.–foreign yield spread predicts a decline in long-term Treasury yields. This spread is particularly informative about changes in Treasury yields outside announcement windows. Consistent with our evidence of an increasing role of spillovers to understanding moves in U.S. long-term yields, the coefficients on the U.S.–foreign yield spread are smaller than those using data starting in the year 2000 (see Table A.4).

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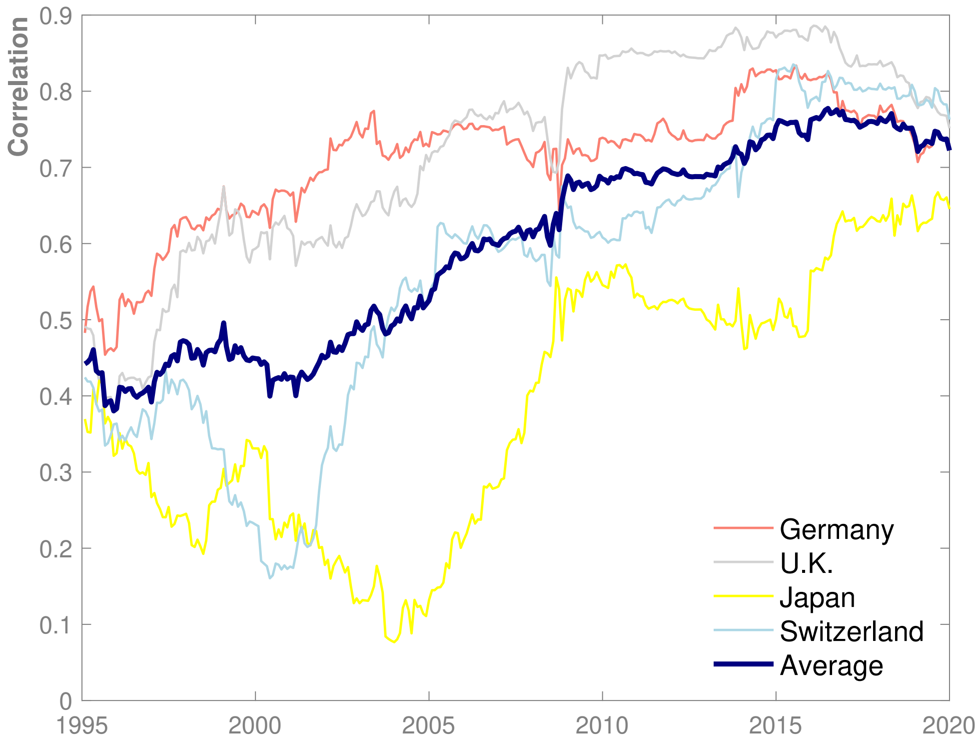


Figure 1. Rolling Correlation Between Monthly Changes in Long-Term Yields on U.S. and Foreign Sovereign Bonds

This figure plots the correlation between monthly changes in the 10-year Treasury yield and 10-year yields on government securities of Germany, Japan, the U.K., and Switzerland along with the average of these correlation coefficients. These are computed using a 5-year rolling window and monthly changes in 10-year yields from January 1990 to December of 2019. The horizontal axis labels the end of the rolling window.

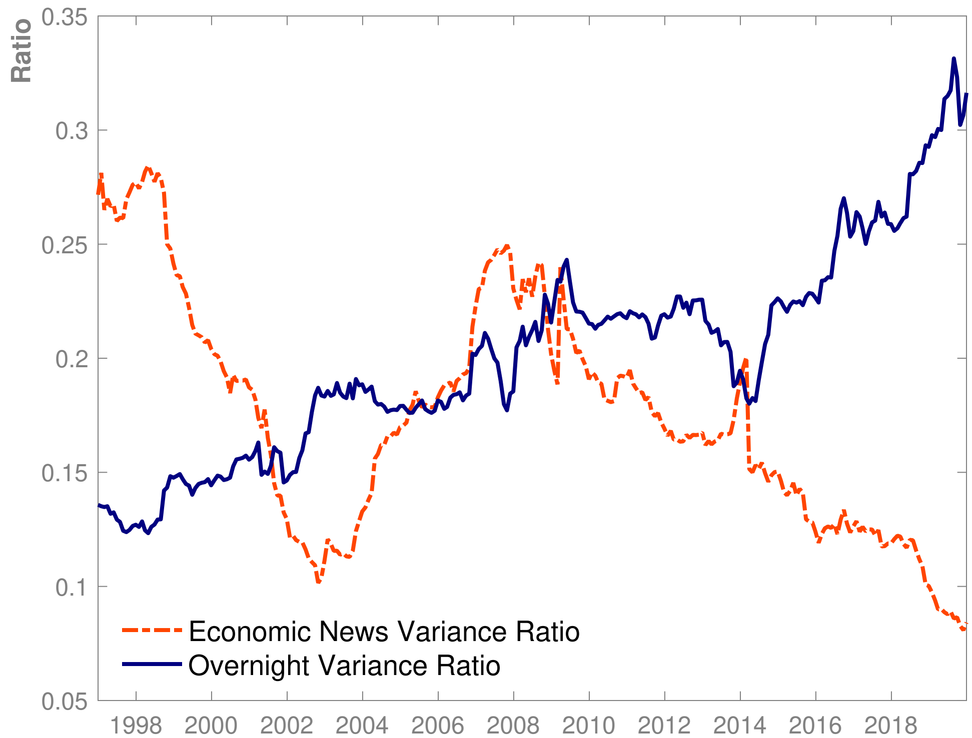
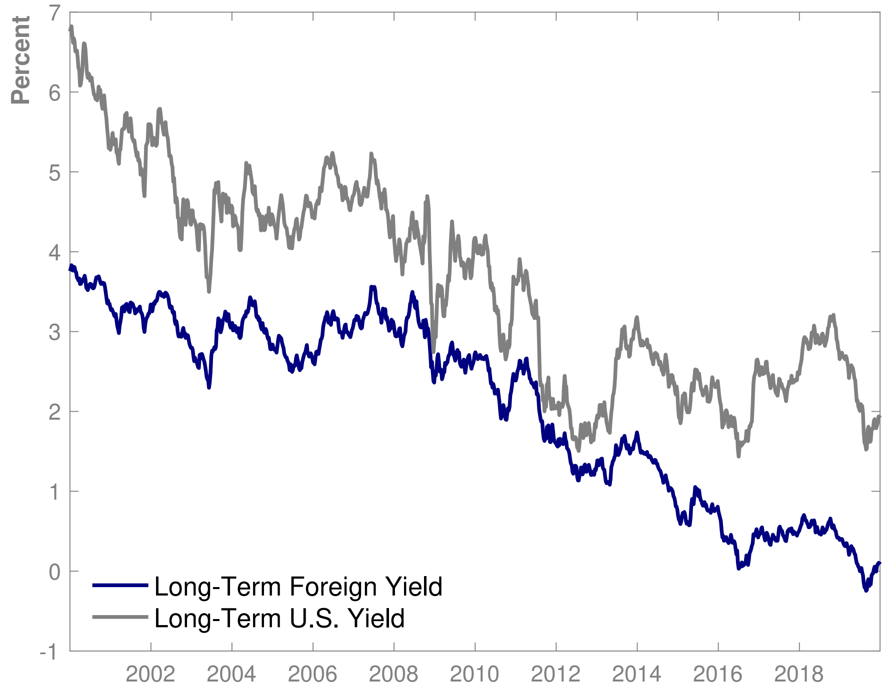
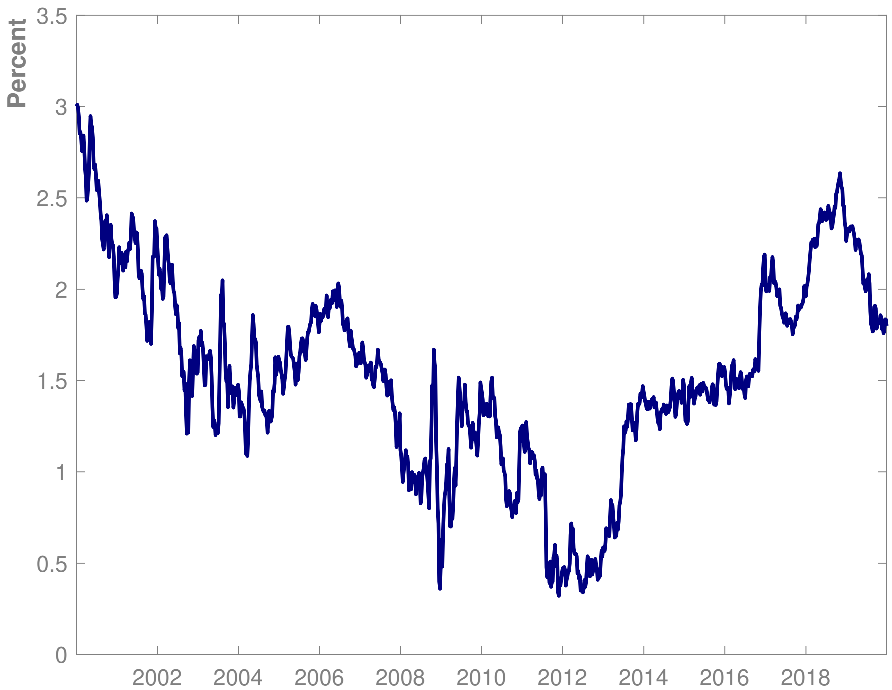


Figure 2. Economic News and Overnight News Variance Ratios
This figure plots the economic news variance ratio (dotted line) defined as the fraction of the variance in the 10-year Treasury yield explained by fluctuations in yields accrued around the release of domestic macroeconomic and monetary policy announcements; and the overnight variance ratio (solid line) defined as the variance of long-term Treasury yield changes outside of U.S. daytime trading hours relative to the overall variance of the changes in long-term yields. The variance ratios are computed using weekly data from January of 1992 to December of 2019, and a 5-year rolling window. The horizontal axis labels the end of the rolling window.



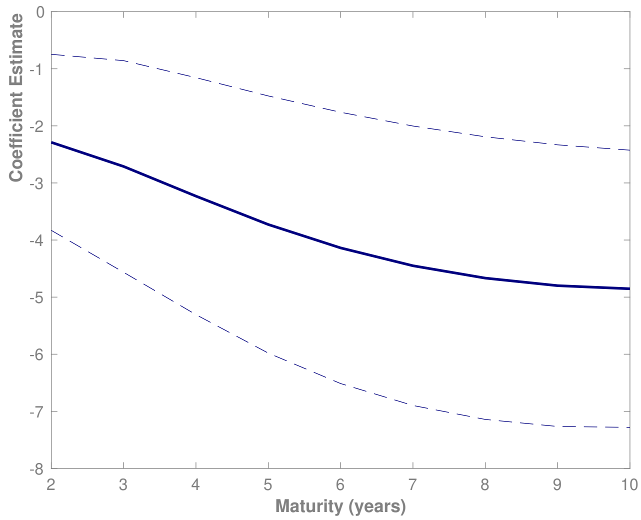
(a) Long-Term Foreign and U.S. Yield



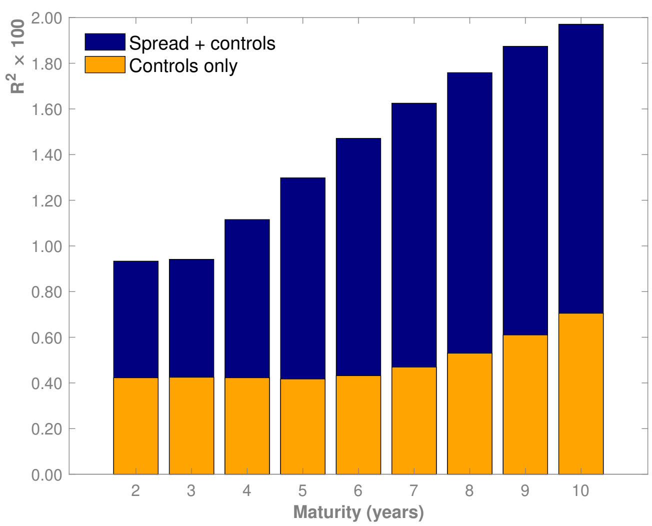
(b) Spread between Long-Term U.S. and Foreign Yields

Figure 3. Long-Term U.S. and Foreign Yields

Panel (a) of this figure shows the long-term yield on foreign sovereign debt computed as the GDP-weighted average of German, Japanese, and U.K. 10-year zero-coupon yields along with the 10-year Treasury yield. Panel (b) plots the spread between long-term U.S. and foreign yields. The data is weekly and the sample period is January 2000 to December 2019.



(a) Estimated coefficient on U.S.–foreign long-term yield spread



(b) Predictive regression R^2

Figure 4. Predictability of U.S. Forward Rates by the U.S.–foreign Long-Term Yield Spread

The upper-panel of this figure plots the coefficient β_n from estimating the following predictive regression of one-year forward rates,

$$\Delta f_{t+1}^{(n)} = \alpha_n + \beta_n \left(y_t - y_{f,t} \right) + \gamma_n' \boldsymbol{x}_t + \epsilon_{t+1}^{(n)},$$

for $n = 2, \ldots, 10$.. The dashed lines, based on Newey and West (1987) standard errors, show 90% confidence intervals. The lower-panel presents the associated R^2 for maturities $n = 2, \ldots, 10$.

Table 1. Economic News and Overnight Variance Ratios

	Sample				
	1992-2019	1992-1996	2015-2019	Change	Wald Test
	(1)	(2)	(3)	(3)−(2)	
Economic News Variance Ratio	0.18	0.28	0.08	-0.19	43.18
	(0.02)	(0.03)	(0.01)		[0.00]
Overnight Variance Ratio	0.20	0.13	0.32	0.18	22.73
	(0.01)	(0.01)	(0.04)		[0.00]

This table reports the economic news variance ratio defined as the fraction of the variance in the 10-year Treasury yield explained by fluctuations in yields accrued around the release of domestic macroeconomic and monetary policy announcements; and the overnight variance ratio defined as the variance of 10-year Treasury yield changes outside of U.S. daytime trading hours relative to the overall variance of the changes in the long-term yield. Our sample is weekly and covers January of 1992 to December of 2019. The table reports GMM Newey and West (1987) standard errors in parentheses. These values are heteroskedasticity-robust and allow for serial correlation up to 52 lags. The last column reports the Wald statistic testing the null hypothesis that the variance ratios remained constant and the associated p-values in brackets.

Table 2. Event-Study Estimates of Yield Spillovers

	Response of country j's long-term yield								
	U.S.			Euro area			U.K.		
Country i	β	R^2	T	β	R^2	T	β	R^2	T
	Panel A: Central Bank Communications and Data Releases								
U.S.				0.49	0.68	119	0.63	0.78	120
				(0.03)			(0.03)		
Euro Area	0.60	0.65	96				0.75	0.73	93
	(0.05)						(0.05)		
U.K.	0.56	0.57	86	0.45	0.52	88			
	(0.05)			(0.05)					
	Panel B: Central Bank Communications								
U.S.				0.50	0.65	39	0.66	0.78	39
				(0.06)			(0.06)		
Euro Area	0.58	0.63	67				0.70	0.69	64
	(0.06)						(0.06)		
U.K.	0.50	0.52	47	0.43	0.47	43			
	(0.08)			(0.07)					
	Panel C: Data Releases								
U.S.				0.47	0.70	83	0.61	0.80	84
				(0.03)			(0.03)		
Euro Area	0.64	0.69	29				0.87	0.85	29
	(0.08)						(0.07)		
U.K.	0.63	0.61	46	0.49	0.57	46			
	(0.07)			(0.07)					

This table presents the slope coefficient from a the following event-study regression,

$$\Delta y^j_{t_i} = \alpha + \beta \Delta y^i_{t_i} + e^j_{t_i}, \quad \text{for } j \neq i$$

where t_i denotes the days in which there were notable news about country i and no important news about other countries, $\Delta y^j_{t_i}$ denotes the one-day change in country j's sovereign long-term yield on the days with news about country $i (\neq j)$. The regressions are estimated using daily time-synchronized data from January of 2010 to October of 2017. Days with notable news for more than one country are excluded from the sample. The estimates are obtained using ordinary least squares (OLS). Standard errors are reported in parenthesis.

Table 3. GMM Estimates of Yield Spillovers

θ	Version 1	Version 2
Γ_{31}	0.411 (0.035)	0.453 (0.037)
Γ_{21}	0.534 (0.041)	0.590 (0.044)
Γ_{12}	0.525 (0.066)	0.502 (0.059)
Γ_{32}	0.730 (0.066)	0.658 (0.068)
Γ_{13}	0.411 (0.054)	0.427 (0.065)
Γ_{23}	0.325 (0.058)	0.276 (0.077)
σ_1^*	8.758 (0.457)	8.720 (0.392)
σ_2^*	6.627 (0.600)	7.000 (0.428)
σ_3^*	5.648 (0.412)	5.497 (0.384)
σ_1	3.590 (0.488)	3.255 (0.311)
σ_2	3.030 (0.646)	3.323 (0.286)
σ_3	2.879 (0.400)	3.110 (0.356)
σ_η	0.001 (1.369 10^3)	0
$\sigma_{\ell,1}$	1.132 (0.407)	1.792 (0.315)
$\sigma_{\ell,2}$	1.132	0.001 (488.7)
$\sigma_{\ell,3}$	1.132	0.651 (1.137)

This table presents the GMM estimates of the following model,

$$\begin{bmatrix} \Delta y_t^{US} \\ \Delta y_t^{EA} \\ \Delta y_t^{UK} \end{bmatrix} = \begin{bmatrix} 1 & \Gamma_{12} & \Gamma_{13} \\ \Gamma_{21} & 1 & \Gamma_{23} \\ \Gamma_{31} & \Gamma_{32} & 1 \end{bmatrix} \begin{bmatrix} \varepsilon_t^{US} \\ \varepsilon_t^{EA} \\ \varepsilon_t^{UK} \end{bmatrix} + \begin{bmatrix} \ell_t^{US} \\ \ell_t^{EA} \\ \ell_t^{UK} \end{bmatrix} + \begin{bmatrix} 1 \\ 1 \\ 1 \end{bmatrix} \eta_t$$

where Δy_t^i is the daily change in country i's long-term yield with the U.S. $i = 1$, the Euro Area $i = 2$ and the U.K. $i = 3$. We assume that the second moments of country i's ε_t^i shock satisfy,

$$\sigma_t(\varepsilon^i) = \begin{cases} \sigma_i^{*2} & \text{(days with country}-i\text{ news)} \\ \sigma_i^2 & \text{(days without country}-i\text{ news)} \end{cases}$$

For each country we define Δy_t^i as the change in the long-term country i yield between 12 p.m. of day t and 12 p.m. of the previous day $t-1$, all in New York time. The sample covers the period from January 2010 to August of 2017.

Table 4. Predictability of Intraday Changes in Long-Term Treasury Yields

	Dependent Variable						
	Panel A: Δy_{t+1}			**Panel B: $\Delta y_{\mathbf{a},t+1}$**		**Panel C: $\Delta y_{\mathbf{na},t+1}$**	
	(1)	(2)	(3)	(1)	(2)	(1)	(2)
$y_t - y_{f,t}$	-1.67	-3.70		-0.04	-0.19	-1.63	-3.50
	[-2.54]	[-3.56]		[-0.16]	[-0.44]	[-2.71]	[-4.24]
$f_t - r_t$		-1.30	-0.17		-0.15		-1.15
		[-2.53]	[-0.49]		[-0.99]		[-2.40]
$\mathbb{E}_t(r_{t+j}) - r_t$		1.76	-0.43		-0.32		2.08
		[1.62]	[-0.54]		[-1.06]		[1.98]
$y_{\text{Aaa},t} - y_t$		-1.53	-1.51		0.08		-1.61
		[-2.16]	[-1.92]		[0.31]		[-2.12]
MBSDUR_t		-0.82	-0.86		-0.10		-0.72
		[-2.01]	[-2.20]		[-1.03]		[-1.84]
$R^2 \times 100$	0.552	1.533	0.528	-0.039	0.296	0.527	1.486
T	1043	1043	1043	1043	1043	1043	1043

This table presents the estimated coefficients from predictive regressions. The dependent variables are weekly changes in the 10-year Treasury yield (Panel A Δy_{t+1}), cumulative weekly changes in 30-minute windows around macroeconomic and monetary policy announcements (Panel B $\Delta y_{\text{a},t+1}$), and the cumulative changes in the long-term yield outside of announcement windows (Panel C $\Delta y_{\text{na},t+1}$). The key explanatory variable is the spread between the 10-year Treasury yield and the long-term foreign yield computed in (17), $y_t - y_{f,t}$. The control variables are the the 10-year forward rate spread ($f_t - r_t$), the near-term forward spread ($\mathbb{E}_t(r_{t+j}) - r_t$), the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{\text{Aaa},t} - y_t$), and the effective duration of outstanding mortgage-backed securities (MBSDUR_t). All regressions include a constant and a dummy variable that is equal to one the weeks when there is an auction. We use Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals. t-statistics are reported in brackets. The sample is weekly and covers the period from January 2000 to December of 2019.

Table 5. Yield Curve Principal Components Regressions

	Dependent Variable					
	Panel A: Δy_{t+1}		**Panel B:** $\Delta y_{\mathbf{a},t+1}$		**Panel C:** $\Delta y_{\mathbf{na},t+1}$	
	(1)	(2)	(1)	(2)	(1)	(2)
$y_t - y_{f,t}$		-3.67		-0.36		-3.32
		[-3.15]		[-0.82]		[-3.55]
L_t	-0.10	0.06	0.01	0.02	-0.11	0.04
	[-2.90]	[1.04]	[0.35]	[1.16]	[-3.22]	[0.76]
S_t	0.28	0.36	0.14	0.15	0.14	0.21
	[1.62]	[2.41]	[2.71]	[2.66]	[0.90]	[1.67]
C_t	-0.56	-2.33	0.32	0.15	-0.87	-2.47
	[-0.70]	[-2.10]	[1.28]	[0.44]	[-1.14]	[-2.51]
$R^2 \times 100$	0.378	1.250	0.246	0.207	0.470	1.297
T	1043	1043	1043	1043	1043	1043

This table presents the estimated coefficients from predictive regressions. The dependent variables are weekly changes in the 10-year Treasury yield (Panel A Δy_{t+1}), cumulative weekly changes in 30-minute windows around macroeconomic and monetary policy announcements (Panel B $\Delta y_{\mathrm{a},t+1}$), and the cumulative changes in yields outside of announcement windows (Panel C $\Delta y_{\mathrm{na},t+1}$). The key explanatory variable is the spread between the 10-year Treasury yield and the long-term foreign yield computed in (17), $y_t - y_{f,t}$. The control variables are the the three principal components of the term structure of U.S. interest rates. All regressions include a constant and a dummy variable that is equal to one the weeks when there is an auction. We use Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals. t-statistics are reported in brackets. The sample is weekly and covers the period from January 2000 to December of 2019.

Table 6. Time-Varying Predictability of the U.S.–Foreign Yield Spread

	Dependent Variable	
	Δy_{t+1}	
	(1)	(2)
$y_t - y_{f,t}$	-2.30	-3.35
	[-3.07]	[-3.56]
$(y_t - y_{f,t}) \times \mathrm{Var}_t(y_o)$	-1.73	-2.10
	[-1.96]	[-2.02]
$f_t - r_t$		-0.95
		[-1.90]
$\mathbb{E}_t(r_{t+j}) - r_t$		1.32
		[1.26]
$y_{\mathrm{Aaa},t} - y_t$		-2.41
		[-2.16]
MBSDUR_t		-0.85
		[-2.49]
$V_t(y_o)$	1.50	2.64
	[1.34]	[1.72]
$R^2 \times 100$	1.084	1.854
T	1043	1043

This table presents the estimated coefficients from predictive regressions. The dependent variable is the weekly change in the 10-year Treasury yield (Δy_{t+1}). The key explanatory variable is the spread between the 10-year Treasury yield and the long-term foreign yield, $y_t - y_{f,t}$. We allow the predictability of the U.S.–foreign yield spread to vary with the variance of changes in the long-term U.S. yield overnight ($\mathrm{Var}(y_o)$). The overnight variance of the yield is computed using an exponentially weighted moving average and it is standardized to have a zero mean, and a standard deviation equal to one. Column (2) includes as control variables the forward spread ($f_t - r_t$), the near-term forward spread ($\mathbb{E}_t(r_{t+j}) - r_t$), the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{\mathrm{Aaa},t} - y_t$), and the effective duration of outstanding mortgage-backed securities (MBSDUR_t). All regressions include a constant and a dummy variable that is equal to one the weeks when there is an auction. We use Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals. t-statistics are reported in brackets. The sample is weekly and covers the period from January 2000 to December of 2019.

Table 7. Predictability of Excess Returns on Long-Term Treasury Securities over Long-Term Foreign Bonds

	Holding Period Return					
	Panel A: 1-week		**Panel B: 4-week**		**Panel C: 12-week**	
	(1)	(2)	(1)	(2)	(1)	(2)
$y_t - y_{f,t}$	8.49	20.83	6.51	15.96	4.46	10.56
	[2.91]	[4.68]	[2.52]	[3.32]	[2.35]	[4.44]
$f_t - r_t$		7.05		5.66		3.62
		[3.46]		[2.61]		[3.22]
$\mathbb{E}_t(r_{t+j}) - r_t$		-10.34		-6.84		-3.52
		[-2.42]		[-1.46]		[-1.21]
$y_{\text{Aaa},t} - y_t$		2.18		3.16		1.68
		[1.01]		[2.07]		[1.36]
MBSDUR_t		1.93		2.17		1.18
		[1.10]		[1.30]		[1.15]
$R^2 \times 100$	0.89	2.22	2.53	8.16	5.43	17.32
T	1032	1032	1029	1029	1022	1022

This table presents the estimated coefficients from predictive regressions. The dependent variable is the τ-period excess return on a 10-year Treasury security over the excess return on a 10-year foreign bond, $rx_{t\to t+\tau} - rx_{\text{f},t\to\tau}$, with $rx_{\text{f},t\to t+\tau}$ defined as the GDP-weighted excess return on German, Japanese, and U.K. 10-year sovereign debt. Panels A, B and C present results for 1-, 4-, and 12-week holding periods, respectively. The key explanatory variable is the spread between the 10-year Treasury yield and our measure of long-term foreign yield, $y_t - y_{f,t}$. The vector $\boldsymbol{x}_t$ controls for the 10-year forward rate spread ($f_t - r_t$), the near-term forward spread ($\mathbb{E}_t(r_{t+j}) - r_t$), the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{\text{Aaa},t} - y_t$), and the effective duration of MBS (MBSDUR_t). All regressions include a constant and a dummy variable that is equal to one the weeks when there is an auction. t-statistics reported in brackets are obtained using Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals and with overlapping observations for holding periods above 1-week. The sample is weekly and covers the period from January 2000 to December of 2019.

Table 8. Predictability of Returns on Long-Term Treasury Securities over Currency-Hedged Returns on Long-Term Foreign Bonds

	Holding Period Return					
	Panel A: 1-week		Panel B: 4-week		Panel C: 12-week	
	(1)	(2)	(1)	(2)	(1)	(2)
$y_t - y_{f,t}$	11.24	21.39	9.72	17.95	7.66	12.47
	[4.20]	[5.12]	[3.62]	[3.67]	[3.92]	[4.92]
$f_t - r_t$		6.11		4.94		2.84
		[3.10]		[2.22]		[2.35]
$E_t(r_{t+j}) - r_t$		-9.76		-7.63		-4.24
		[-2.46]		[-1.60]		[-1.45]
$y_{Aaa,t} - y_t$		2.11		3.42		2.12
		[0.81]		[2.17]		[1.61]
MBSDUR_t		2.36		2.16		1.15
		[1.41]		[1.23]		[1.05]
$R^2 \times 100$	1.59	2.44	5.72	9.59	15.22	20.86
T	1043	1043	1040	1040	1032	1032

This table presents the estimated coefficients from predictive regressions. The dependent variable is the τ-period return on a 10-year Treasury security in excess of the currency-hedged return on a 10-year foreign bond, $ret_{t\to t+\tau} - ret^h_{\mathrm{f},t\to t+\tau}$, with $ret^h_{\mathrm{f},t\to t+\tau}$ defined as the GDP-weighted currency-hedged return on German, Japanese, and U.K. 10-year sovereign debt. Panels A, B and C present results for 1-, 4-, and 12-week holding periods, respectively. The key explanatory variable is the spread between the 10-year Treasury yield and our measure of long-term foreign yield, $y_t - y_{f,t}$. The vector $\boldsymbol{x}_t$ controls for the 10-year forward rate spread ($f_t - r_t$), the near-term forward spread ($\mathbb{E}_t(r_{t+j}) - r_t$), the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{\mathrm{Aaa},t} - y_t$), and the effective duration of MBS (MBSDUR_t). All regressions include a constant and a dummy variable that is equal to one the weeks when there is an auction. t-statistics reported in brackets are obtained using Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals and with overlapping observations for holding periods above 1-week. The sample is weekly and covers the period from January 2000 to December of 2019.

Table A.1. List of Notable News Days for the January 2010-August 2017 period in the U.S., Euro area and U.K.

US – data releases
20100223 20100305 20100402 20100604 20100715 20100730 20100806 20100819 20100824 20100901 20100903 20100909
20101006 20101008 20101105 20101210 20101214 20110105 20110107 20110120 20110201 20110204 20110303 20110310
20110412 20110504 20110506 20110526 20110601 20110614 20110630 20110708 20110729 20110801 20110818 20110829
20110902 20111007 20120119 20120203 20120517 20120601 20120613 20120802 20120803 20120814 20120907 20121207
20130117 20130306 20130308 20130403 20130405 20130412 20130503 20130516 20130531 20130603 20130605 20130705
20130715 20130717 20130718 20130731 20130801 20130802 20130823 20130903 20130906 20131101 20131105 20131108
20131204 20140110 20140116 20140203 20140218 20140307 20140417 20140430 20140513 20140602 20140625 20140730
20140813 20140905 20141015 20141205 20150114 20150206 20150306 20150401 20150402 20150508 20150519 20150605
20150702 20150805 20151002 20151014 20151106 20151201 20160212 20160301 20160415 20160603 20160729 20160906
20161216 20170106 20170119 20170614 20170703
US – central bank communications
20100715 20100722 20100803 20100811 20100827 20100922 20101104 20101206 20110608 20110810 20110922 20120126
20120314 20120404 20120823 20120831 20120914 20121005 20121213 20130523 20130620 20130621 20130624 20130711
20130822 20130905 20130916 20130919 20130924 20131027 20131031 20140320 20140331 20140410 20140710 20140918
20141218 20150224 20150319 20150805 20150820 20150918 20151029 20160317 20160922 20161215 20170105 20170119
20170214 20170301 20170316 20170504 20170615 20170712
US – other news
20100223 20100604 20100811 20100824 20100922 20101115 20101207 20101208 20110201 20110303 20110412 20110729
20110808 20110810 20110818 20110902 20110922 20120409 20120803 20130503 20130620 20130705 20130903 20130905
20130919 20130924 20131017 20131204 20140110 20141015 20141218 20161102 20161109 20161114 20170123 20170301
20170614
Euro area – data releases
20100223 20100723 20100802 20100914 20100923 20110112 20110201 20110901 20110907 20110930 20111024 20111220
20120201 20120222 20120322 20120423 20120502 20120815 20120924 20130214 20130221 20130321 20130603 20130724
20130813 20130822 20130903 20130924 20131031 20131202 20131204 20140602 20141030 20141120 20150106 20150227
20150423 20150602 20150731 20160229 20160428 20170103 20170428 20170504 20170629 20170728
Euro area – central bank communications
20100506 20100507 20100610 20100902 20101201 20101202 20110118 20110303 20110505 20110804 20110808 20110909
20111006 20111011 20111012 20111103 20111201 20111208 20120607 20120705 20120706 20120726 20120727 20120802
20120803 20120821 20120904 20120906 20121206 20130110 20130207 20130307 20130506 20130704 20131107 20131205
20140206 20140306 20140403 20140514 20140605 20140825 20141021 20141106 20141117 20141121 20141124 20150122
20150123 20150309 20150310 20150429 20150505 20150507 20150511 20150519 20150603 20150903 20151021 20151022
20151102 20151112 20151113 20151120 20151203 20160121 20160310 20160630 20160908 20160909 20160912 20161004
20161005 20161123 20161208 20170309 20170427 20170608 20170627 20170628

Table A.1. List of Notable News Days in the U.S., Euro area and U.K. (cont.)

Euro area – other news
20100211 20100223 20100427 20100430 20100504 20100510 20100514 20100604 20100907 20101123 20101130 20110128 20110201 20110303 20110418 20110628 20110629 20110711 20110721 20110727 20110809 20110810 20110818 20110902 20110914 20110919 20110922 20110927 20111004 20111005 20111006 20111017 20111021 20111027 20111031 20111101 20111109 20111114 20111209 20111212 20111216 20120210 20120222 20120410 20120523 20120530 20120607 20120619 20120625 20120705 20120802 20120803 20120815 20120926 20121017 20130226 20130705 20130903 20130924 20131204 20140514 20150106 20150622 20150629 20150707 20150710 20161114 20170301
U.K. – data releases
20100119 20100122 20100126 20100225 20100420 20100518 20100723 20100819 20101021 20101026 20110118 20110125 20110201 20110202 20110218 20110225 20110322 20110405 20110412 20110504 20110517 20110628 20110729 20111206 20120601 20120625 20120807 20120815 20121017 20121025 20130301 20130321 20130425 20130503 20130603 20130620 20130709 20130717 20130815 20130903 20130919 20130926 20131105 20131112 20131113 20131218 20140110 20140117 20140122 20140219 20140521 20140708 20140715 20141014 20141015 20141218 20150106 20150218 20150318 20150519 20150528 20150617 20150818 20150916 20151007 20160712 20160901 20161027 20161207 20170103 20170321 20170613 20170614 20170718
U.K. – central bank communications
20100210 20100223 20100512 20100728 20100811 20100824 20100922 20101020 20101110 20101116 20110216 20110503 20110511 20110810 20110817 20111006 20120209 20120222 20120418 20120620 20120705 20120718 20120802 20130213 20130220 20130307 20130315 20130617 20130625 20130705 20130905 20130918 20130927 20131113 20140212 20140514 20140613 20140618 20140813 20140818 20141017 20141112 20150121 20150422 20150714 20150806 20160119 20160524 20160630 20160701 20160714 20160804 20160808 20160810 20161007 20170202 20170615 20170620 20170803 20170804
U.K. – other news
20100811 20100824 20100922 20110201 20110412 20110628 20110729 20110810 20110818 20110902 20120222 20120523 20120607 20120705 20120815 20121017 20130503 20130620 20130705 20130903 20130905 20130919 20130924 20131017 20131105 20140110 20140514 20141015 20141218 20150106 20160620 20160624 20160627 20170614

This table presents a list of dates identified as days with notable news. Some of these dates are one business day after the calendar day of the event to align with the change in yields from 12 p.m. to 12 p.m., New York time. For example, an FOMC event that happened at 2 p.m. on day t would be recorded as day $t+1$.

Table A.2. Equally-Weighted Foreign Long-Term Yields

	Dependent Variable					
	Panel A: Δy_{t+1}		**Panel B:** $\Delta y_{\mathbf{a},t+1}$		**Panel C:** $\Delta y_{\mathbf{na},t+1}$	
	(1)	(2)	(1)	(2)	(1)	(2)
$y_t - y_{f,t}$	-3.39	-3.25	-0.45	-0.38	-2.93	-2.87
	[-3.32]	[-3.50]	[-0.97]	[-0.83]	[-3.33]	[-3.67]
$(y_t - y_{f,t}) \times \text{Var}_t(y_o)$		-2.45		-0.07		-2.38
		[-2.40]		[-0.29]		[-2.47]
$f_t - r_t$	-1.10	-0.86	-0.22	-0.15	-0.88	-0.70
	[-2.27]	[-1.71]	[-1.47]	[-0.97]	[-2.00]	[-1.48]
$\mathbb{E}_t(r_{t+j}) - r_t$	1.28	0.93	-0.21	-0.27	-1.76	1.19
	[1.25]	[0.93]	[-0.70]	[-0.86]	[-1.76]	[1.24]
$y_{\text{Aaa},t} - y_t$	-1.71	-3.04	0.05	0.32	-1.76	-3.36
	[-2.25]	[-2.72]	[0.22]	[0.87]	[-1.76]	[-3.31]
MBSDUR_t	-0.61	-0.66	-0.07	-0.11	-0.54	-0.56
	[-1.56]	[-2.00]	[-0.66]	[-0.98]	[-1.41]	[-1.76]
$R^2 \times 100$	1.170	1.666	0.360	0.261	0.979	1.563
T	1043	1043	1043	1043	1043	1043

This table presents the estimated coefficients from predictive regressions. The dependent variables are weekly changes in the 10-year Treasury yield (Panel A Δy_{t+1}), cumulative weekly changes in 30-minute windows around macroeconomic and monetary policy announcements (Panel B $\Delta y_{\text{a},t+1}$), and the cumulative changes in yields outside of these windows (Panel C $\Delta y_{\text{na},t+1}$). The key explanatory variable is the spread between the 10-year Treasury yield and the long-term foreign yield. The control variables are the the 10-year forward rate spread ($f_t - r_t$), the near-term forward spread ($\mathbb{E}_t(r_{t+j}) - r_t$), the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{\text{Aaa},t} - y_t$), and the effective duration of outstanding mortgage-backed securities ($MBSDUR_t$). All regressions include a constant and a dummy variable that is equal to one the weeks when there is an auction. We use Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals. t-statistics are reported in brackets. The sample is weekly and covers the period from January 2000 to December of 2019.

Table A.3. Robustness to Using a Wider Set of Foreign Countries

	Value-weighted			Equally-weighted		
	Δy_{t+1}	$\Delta y_{\mathrm{a},t+1}$	$\Delta y_{\mathrm{na},t+1}$	Δy_{t+1}	$\Delta y_{\mathrm{a},t+1}$	$\Delta y_{\mathrm{na},t+1}$
$y_t - y_{f,t}$	-3.30	-0.16	-3.13	-2.56	-0.36	-2.20
	[-2.92]	[-0.31]	[-3.67]	[-2.48]	[-0.68]	[-2.55]
$f_t - r_t$	-1.15	-0.14	-1.01	-0.84	-0.18	-0.65
	[-2.37]	[-0.83]	[-2.31]	[-1.86]	[-1.18]	[-1.62]
$\mathbb{E}_t(r_{t+j}) - r_t$	1.45	-0.34	1.79	0.80	-0.26	1.06
	[1.41]	[-1.05]	[1.83]	[0.83]	[-0.87]	[1.20]
$y_{\mathrm{Aaa},t} - y_t$	-1.71	0.07	-1.78	-1.85	0.03	-1.88
	[-2.24]	[0.28]	[-2.20]	[-2.26]	[0.13]	[-2.19]
MBSDUR$_t$	-0.67	-0.09	-0.58	-0.56	-0.06	-0.50
	[-1.73]	[-0.95]	[-1.53]	[-1.45]	[-0.54]	[-1.29]
$R^2 \times 100$	1.199	0.289	1.143	0.831	0.326	0.677
T	1043	1043	1043	1043	1043	1043

This table presents the estimated coefficients from predictive regressions. The dependent variables are weekly changes in the 10-year Treasury yields (Δy_{t+1}), cumulative weekly changes in 30-minute windows around macroeconomic and monetary policy announcements ($\Delta y_{\mathrm{a},t+1}$), and the cumulative weekly changes in yields outside of these windows ($\Delta y_{\mathrm{n},t+1}$). The key explanatory variable is the spread between the 10-year Treasury yield and the long-term foreign yield computed as the equally and GDP-weighted average of German, Japanese, French, U.K. and Swiss 10-year yields. The control variables are the the 10-year forward rate spread ($f_t - r_t$), the near-term forward spread ($\mathbb{E}_t(r_{t+j}) - r_t$), the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{\mathrm{Aaa},t} - y_t$), and the effective duration of outstanding mortgage-backed securities ($MBSDUR_t$). All regressions include a constant and a dummy variable that is equal to one the weeks when there is an auction. We use Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals. t-statistics are reported in brackets. The sample is weekly and covers the period from January 2000 to December of 2019.

Table A.4. Predictability Using Monthly Data 2000–2019

	Dependent Variable					
	Panel A: Δy_{t+1}		**Panel B:** $\Delta y_{\mathbf{a},t+1}$		**Panel C:** $\Delta y_{\mathbf{na},t+1}$	
	(1)	(2)	(1)	(2)	(1)	(2)
$y_t - y_{f,t}$	-8.01	-15.74	0.23	-0.61	-8.24	-15.13
	[-2.51]	[-2.69]	[0.22]	[-0.34]	[-2.85]	[-3.01]
$f_t - r_t$		-5.03		-0.58		-4.45
		[-2.04]		[-0.86]		[-1.85]
$\mathbb{E}_t(r_{t+j}) - r_t$		6.22		-1.16		7.38
		[1.22]		[-0.83]		[1.42]
$y_{\text{Aaa},t} - y_t$		-7.49		-1.51		-5.98
		[-2.97]		[-0.91]		[-1.83]
MBSDUR_t		-3.25		-0.39		-2.86
		[-1.81]		[-1.08]		[-1.67]
$R^2 \times 100$	3.901	8.013	-0.774	1.537	4.053	6.590
T	240	240	240	240	240	240

This table presents the estimated coefficients from predictive regressions. The dependent variables are monthly changes in the 10-year Treasury yield (Panel A Δy_{t+1}), cumulative monthly changes in 30-minute windows around macroeconomic and monetary policy announcements (Panel B $\Delta y_{\text{a},t+1}$), and the cumulative monthly changes in yields outside of these windows (Panel C $\Delta y_{\text{na},t+1}$). The key explanatory variable in the regressions reported in the table is the spread between the 10-year Treasury yield and the long-term foreign yield computed in (17), $y_t - y_{f,t}$. The control variables are the the 10-year forward rate spread ($f_t - r_t$), the near-term forward spread ($\mathbb{E}_t(r_{t+j}) - r_t$, the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{\text{Aaa},t} - y_t$), and the effective duration of outstanding mortgage-backed securities ($MBSDUR_t$). We use Newey and West (1987) standard errors with 12 lags to deal with the autocorrelation of the residuals. t-statistics are reported in brackets. The sample is monthly and covers the period from January 2000 to December of 2019.

Table A.5. Predictability Using Monthly Data 1991–2019

	Dependent Variable		
	Δy_{t+1}	$\Delta y_{\mathrm{a},t+1}$	$\Delta y_{\mathrm{na},t+1}$
$y_t - y_{f,t}$	-5.91	0.36	-6.27
	[-1.86]	[0.41]	[-2.33]
L_t	-0.19	0.00	-0.19
	[-1.22]	[0.01]	[-1.48]
S_t	2.08	0.84	1.24
	[2.01]	[2.38]	[1.34]
C_t	-14.88	0.91	-15.79
	[-2.02]	[0.43]	[-2.30]
$R^2 \times 100$	3.468	0.480	3.486
T	344	344	344

This table presents the estimated coefficients from predictive regressions. The dependent variables are monthly changes in the 10-year Treasury yield (Δy_{t+1}), cumulative monthly changes in 30-minute windows around macroeconomic and monetary policy announcements ($\Delta y_{\mathrm{a},t+1}$), and the cumulative monthly changes in yields outside of these windows ($\Delta y_{\mathrm{na},t+1}$). The key explanatory variable is the spread between the 10-year Treasury yield and the foreign long-term yields computed in (17), $y_t - y_{f,t}$. The control variables are the the three principal components of the term structure of U.S. interest rates. We use Newey and West (1987) standard errors with 12 lags to deal with the autocorrelation of the residuals. t-statistics are reported in brackets. The sample is monthly and covers the period from May 1991 to December of 2019.