


International Yield Spillovers

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Abstract

This article 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 proxy for foreign shocks' contribution to U.S. yields. A model that identifies U.S., Euro area, and U.K. shocks that move global yields suggests that foreign shocks account for at least 20% of the daily variation in long-term U.S. yields in recent years. We also document the predictability of long-term U.S. yields by the U.S.–foreign yield spread.

I. Introduction

Over the 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 several years were on average about 2 times higher, reaching levels around 0.7 in 2019 (see [Figure 1](#)).¹ Earlier studies suggest that the comovement between developed countries' 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 (see, e.g., [Goldberg and Leonard \(2003\)](#), [Gerko and Rey \(2017\)](#), [Rogers, Scotti, and Wright \(2018\)](#), and [Brusa, Savor, and Wilson \(2020\)](#)).

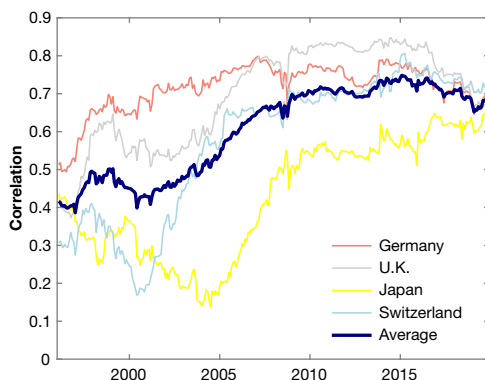
The behavior of long-term sovereign yields in advanced economies in recent years, however, has drawn increased attention to the possibility that U.S. yield

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¹In 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.

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

Figure 1 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 Jan. 1990 to Dec. 2019. The horizontal axis labels the end of the rolling window.



movements are significantly affected by foreign developments. For example, even as the FOMC tightened policy from Dec. 2015 to Dec. 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 time²: 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 long-term U.S. bonds more attractive relative to long-term foreign bonds.³

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

²See, for example, the discussion of the yield curve inversion in Aug. 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: a recession did occur in 2020, but the economic contraction is widely viewed as caused by a large, unanticipated negative shock, namely, the COVID-19 pandemic.

³This debate was already alive after the European debt crisis. Some investors reportedly argued that the slow postcrisis 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, June 2012, *The Economist* article “To strive, to seek, to find, and not to yield.”

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% of the variation in the long-term U.S. yield in the 1992–1996 period to representing 9% 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% in the 1992–1996 period to 30% 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. yields.

Second, we propose a measure of the magnitude of spillovers from foreign yields using a model that decomposes the U.S., Euro area, and U.K. long-term yield changes into three types 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 from Jan. 2010 to Aug. 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 (2018a). We further document that the share of variance of long-term U.S. yields explained by Euro area and U.K. shocks is nonnegligible in recent years. Our estimates suggest that between 20% and 25% of 10-year Treasury yield variations are accounted for by foreign (non-U.S.) shocks over the period of 2010 to 2017. 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, likely reflecting 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 predictive power 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. Our empirical evidence also shows that the spread between U.S. and foreign long-term yields contains predictive information beyond that of the short-maturity cycle variable of Cieslak and Povala (2015) and the “convergence gap” factor of Berardi, Markovich, Plazzi, and Tamoni (2021), which control for fluctuations in the expected U.S. short-rate at business cycle frequencies, suggesting that the foreign spread likely reflects moves in risk premia driven by foreign factors rather than moves in the expectations component of long-term rates. Finally, 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 largely through a portfolio balance channel.

This article 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.⁴ Gerko and Rey (2017) and Rogers et al. (2018), using high-frequency asset price 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

⁴There 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 Saprizza (2015), Neely (2015), Aizenman, Chinn, and Ito (2016), Bernanke (2017), Dedola, Rivolta, and Stracca (2017), and Curcuro, Kamin, Li, and Rodriguez (2018b).

long-term Treasury yields is mixed and mostly sides with the view that the U.S. sets the tone in international bond markets.⁵ 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 an insignificant effect on U.S. yields. By contrast, using an event-study approach, Curcuru et al. (2018a) do find evidence of spillovers from German yields to U.S. yields following policy communications from the ECB. In addition, 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.⁶

Existing research using an event-study methodology around monetary policy announcements, however, still leaves open the question how much of U.S. yield variation is accounted for by foreign shocks, because days with monetary policy announcements represent only a fraction of 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 model of global yields, presented in Section III, allows us to estimate the contribution of U.S. and foreign (non-U.S.) shocks to the total variance of yield changes. In addition, our article offers 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. Furthermore, while most of this literature has focused on the effect of central bank communications (i.e., monetary policy shocks) on global yields, our approaches (Sections II and III) strive to account for spillovers from other kinds of foreign news as well, in particular foreign data releases, as we need to account for *all* notable foreign events, not just foreign central bank communications, in order to quantify the foreign (non-U.S.) contribution to U.S. yield variation. A key aspect of our empirical strategy in Section III is the use of “daily” data; more specifically, we sample data at noon New York time, when the liquidity in both U.S. and European markets is very likely higher than overnight hours. Any higher frequency (such as hourly or 30-minutely) for our model could be less promising, as there would be times (hours) when one market is in overnight hours, and consequently in a less liquid state, while another market is in a daytime session. The use of “daily” data may also address the concern that some of the notable “events” might not be precisely resolved at a high frequency; for example, sometimes there could

⁵Similarly, Brusa et al. (2020) show that investors in equity markets in Germany, Japan, and the UK 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.

⁶The evidence of spillovers from foreign economies to the U.S. found in recent papers contrasts with the mild or nonexistent spillovers reported in earlier literature. We believe that one finding in our paper that helps reconcile the seeming inconsistency is that the overnight variance ratio has increased over time, which suggests that the foreign influence on U.S. yields has grown over time and that estimates using a more recent sample are more likely to find evidence of spillovers. For example, Curcuru et al. (2018a) find that the Euro area-to-U.S. coefficient is smaller in the early part of their sample (2005–2007) relative to their estimates based on postcrisis data (2010–2017).

be some additional price movements in the hours following a data release (without any other news), and some ECB and Fed communication events can occur over a span of more than an hour (statement release followed by press conference).

Lastly, this article is related to an older literature that studies interest rate linkages in a cointegration framework. Kirchgässner and Wolters (1993), for example, examine the cointegration of U.S., German, and other European short-term interest rates to test the “German Dominance” hypothesis, and Chinn and Frankel (1995) study 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 IV 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 article), 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 easily control for other known predictors of bond returns.

II. Decomposing Round-the-Clock Variations in Long-Term Treasury Yields

A 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: the changes in the Treasury yield around key macroeconomic data releases and monetary policy announcements, and changes in the Treasury yield outside of these windows, namely,

$$(1) \quad \Delta y_{t+1} = \Delta y_{a,t+1} + \Delta y_{na,t+1},$$

where $\Delta y_{a,t+1}$ is the yield change accrued around a narrow window bracketing major economic and policy announcements, and $\Delta y_{na,t+1}$ is the yield change outside of these windows.⁷ 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,

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

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

⁷This decomposition is analogous to the decomposition in Faust and Wright (2018) of bond returns earned around announcements and at other times.

announcements. We focus on the reaction of yields around the release of the FOMC statement and the following 14 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.⁸ We cumulate the change in yields around macroeconomic releases to a weekly frequency, so that $\Delta y_{a,t+1}$ represents the change in long-term Treasury yields during week $t + 1$. Our sample covers Jan. 1992 to De. 2019.⁹

Changes in the Treasury yield between time t and $t + 1$ can also be decomposed as

$$(3) \quad \Delta y_{t+1} = \Delta y_{o,t+1} + \Delta y_{d,t+1},$$

where $\Delta y_{o,t+1}$ and $\Delta y_{d,t+1}$ represent changes in the yield overnight and changes in the yield during the domestic daytime trading session, respectively. Since many of the most important foreign economic news are released outside of U.S. daytime trading hours, the contribution of yield changes overnight to the overall variation in the Treasury yield is a simple proxy for the degree of spillovers from news about foreign macroeconomic fundamentals and foreign economic policies to domestic long-term yields. The overnight variance ratio is, in turn, defined as

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

In our empirical analysis, we define overnight yield changes as the change in the 10-year Treasury yield between 8:00AM and 5:00PM 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 estimates of the economic news variance ratio and the overnight variance ratio for the full sample (1992–2019), the first 5 years of the sample (1992–1996), and the last 5 years of the sample (2015–2019). The last 2 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 parentheses, 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 column 1 of Table 1, from 1992 to 2019 the economic news variance ratio is around 20%. The subsample evidence, reported in columns 2 and 3,

⁸These announcements have been shown to be the most 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).

⁹The Supplementary Material contains a description and source of the data series used in this section as well as subsequent sections.

¹⁰The 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.

TABLE 1
Economic News and Overnight Variance Ratios

Table 1 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 Jan. 1992 to Dec. 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.

	Sample			Change	Wald Test
	1992–2019	1992–1996	2015–2019		
	1	2	3	4	5
Economic news	0.18	0.27	0.09	–0.18	37.82
Variance ratio	(0.02)	(0.03)	(0.01)		[0.00]
Overnight	0.20	0.13	0.31	0.18	20.46
Variance ratio	(0.01)	(0.01)	(0.04)		[0.00]

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% of the variation in yields between 1992 and 1996 to slightly below 10% 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.

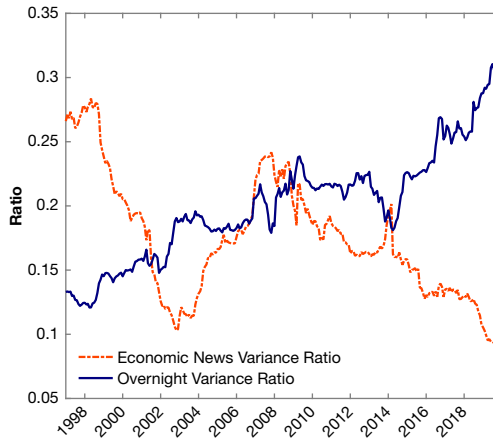
The decline in the economic variance ratio (and, equivalently, the rise in the share of yield variations from movements outside announcement windows) appears to reflect in large part the rise in the share of yield variations coming from overnight hours.¹¹ The second row of Table 1 shows that the share of yield variation due to overnight yield movements increased from percentages in the low teens in the 1992–1996 period to slightly above 30% in the 2015–2019 period. As indicated by the Wald statistic, the increase in the overnight variance ratio between 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 year $t - 5$ to year 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

¹¹Note that the share of variation in the Treasury yield outside announcement windows can be approximated as $\text{var}(\Delta y_{na})/\text{var}(\Delta y) \approx 1 - \text{var}(\Delta y_a)/\text{var}(\Delta y)$, because in practice $\text{cov}(\Delta y_a, \Delta y_{na}) \approx 0$. Also 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 yield changes during nonannouncement periods Δy_{na} , that is, $\Delta y_{na} = \Delta y_o + \Delta y_{na,d}$, where $\Delta y_{na,d}$ denotes the yield changes in nonannouncement periods that occur during the day time.

FIGURE 2
Economic News and Overnight Variance Ratios

Figure 2 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 Jan. 1992 to Dec. 2019, and a 5-year rolling window. The horizontal axis labels the end of the rolling window.



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 U.S. presidential elections. 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. As shown in Tables A.1 and A.2 of the Supplementary Material, this empirical regularity in long-term yields is robust to alternative overnight window definitions as well as to alternative event windows around key macroeconomic and monetary policy announcements.

¹²Swanson 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.

It is also 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 our exercise *are* more spotty during overnight hours, our measure of overnight yield changes uses only the yield before the start of the U.S. daytime session (namely, 8:00AM) and the yield at the previous close of the U.S. daytime session (namely, 5:00PM); therefore, as long as market-moving foreign news during the overnight hours is incorporated in Treasury prices before the start of U.S. daytime session, the overnight variance ratio will capture them. The other possibility (that overnight foreign news is not incorporated until after the start of U.S. daytime trading) would require a fairly strong belief in market inefficiency. 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 (7:30AM to 8:00AM). 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.

III. Decomposing Multi-Country Yield Changes

A. Identification by Heteroskedasticity

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

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

where Δy_t^i is the daily change in country i 's long-term yield ($i = 1, 2, 3 = \text{U.S., EA, U.K.}$). Our model assumes two types of individual country shocks. The shocks $\varepsilon_t^{\text{US}}, \varepsilon_t^{\text{EA}}, \varepsilon_t^{\text{UK}}$ 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^i , 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 $\Gamma_{ij} (i \neq j)$ are different from 0 or country yields are sensitive to the "other global" shock η_t .

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

$$(6) \quad \Delta y_t = \Gamma \varepsilon_t + \iota' \eta_t + \ell_t,$$

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

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 has a significant impact not only on yields in the country where news is emanating, but also on yields in other countries (Rogers, Scotti, and Wright (2014), Curcuru et al. (2018a)). But there are also many days without notable news. Our key identifying assumption 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,¹³ except for an overall scale factor, that is, the second moment of the ε_t^i shock satisfies

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

with $\sigma_i^* > \sigma_i$.¹⁴ In addition, we assume that the local shocks ε_t^i and the “other global” shock η_t are homoscedastic, namely,

$$(8) \quad \sigma_t^2(\varepsilon^i) = \sigma_{\varepsilon^i}^2, \quad \sigma_t^2(\eta) = \sigma_\eta^2.$$

Our approach is closely related to Wright (2012), who uses an identification-by-heteroskedasticity approach to estimate the effects of monetary policy shocks on other asset prices.¹⁵

To estimate the spillover effects using the model in equation (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 Jan. 2010 to Aug. 2017.¹⁶ Specifically, for each country, we define Δy_t^j as the change in the 10-year yield between 12:00PM of day t and 12:00PM of the previous day $t - 1$, all in New York time.

As detailed in Appendix A, 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 to assess the sensitivity of our key results to the specific criteria for determining notable news days.

While the model in equation (6) can be estimated for N countries, 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

¹³That is, the Γ matrix is the same on days with notable news and days without notable news.

¹⁴As 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 equation (5).

¹⁵Rigobon (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 shocks.

¹⁶The start of our sample is guided by the findings in Section II 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, for example, whether they are U.S. news or global news.

identified, but we can impose a few 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. Pasquariello and Vega (2007) and Evans and Lyons (2008), 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 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.

Aside from the “local” shock ℓ_t^i that does not affect other countries, ℓ_t^i is the only country- i shock in our model. At a deeper level, the ℓ_t^i shock can be further decomposed into a monetary policy shock, a growth shock, an inflation shock, a risk premium shock, and so forth.¹⁷ These components may very well have different propagation properties (impulse responses), which are beyond the purview of the present article. 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.

B. Preliminary Regressions

Intuitively, on days with notable country- i news (and no news for other countries), we can expect that $\sigma_i^* > \sigma_{j=1, \dots, N(\neq i)}, \sigma_{\ell, j=1, \dots, N}, \sigma_\eta$. Therefore, a rough estimate of Γ_{ji} can be obtained by running the following event-study type regression.¹⁸

¹⁷There are some empirical studies that estimate a VAR that model the 10-year Treasury yield using macroeconomic and financial variables. Cieslak and Pang (2021), 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.

¹⁸Rigobon and Sack (2004) note that event study regressions are a special case of their identification-by-heteroskedasticity approach.

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

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 1-day change in country j 's long-term yield on days when there is 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 3 countries in our sample (U.S., Euro area, and U.K.). Panel A presents the results for all days that contain notable macro news releases or central 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 is notable news about more than 1 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 was important news about those economies is

TABLE 2
Event-Study Estimates of Yield Spillovers

Table 2 presents the slope coefficient from the following event-study regression:

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

where t_i denotes the days in which there were notable news about country i and no important news about other countries, $\Delta y_{t_i}^j$ denotes the 1-day change in country j 's sovereign long-term yield on the days with news about country i ($j \neq i$). The regressions are estimated using daily time-synchronized data from Jan. 2010 to Oct. 2017. Days with notable news for more than 1 country are excluded from the sample. The estimates are obtained using ordinary least squares (OLS). Standard errors are reported in parentheses.

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

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^2 s. 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 equation (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 0.66 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 reasonable 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 of Table 2 is consistent with the findings in Curcuru et al. (2018a), namely, about half of the moves in long-term German Bund yields from Euro area monetary policy shocks are transmitted to long-term U.S. yields.

These results suggest that our selection 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 types 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.

C. Full Model Estimation

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

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

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

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,¹⁹ 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:

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

with the vector h_t given by

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

where $\Omega_i(\theta)$ are given in equation (11), d_{0t} is a dummy variable that is equal to 1 on days with no news, and d_{it} for $i > 0$ is a dummy variable that is equal to 1 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 country i as determined in Appendix A, excluding days with news for more than 1 country. Days classified as “other news” days are not included in T_1, T_2, T_3 (or in T_0) 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 1,348 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:

$$(14) \quad \text{Version 1: } \sigma_{\ell, 1} = \sigma_{\ell, 2} = \sigma_{\ell, 3},$$

$$(15) \quad \text{Version 2: } \sigma_\eta = 0.$$

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

¹⁹There are only a small number of days when news emerges for two or more countries.

TABLE 3
GMM Estimates of Yield Spillovers

Table 3 presents the GMM estimates of the following model:

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

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 12PM of day t and 12PM of the previous day $t - 1$, all in New York time. The sample covers the period from Jan. 2010 to Aug. 2017.

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 $\times 10^3$)	0
$\sigma_{\varepsilon,1}$	1.132 (0.407)	1.792 (0.315)
$\sigma_{\varepsilon,2}$	1.132	0.001 (488.7)
$\sigma_{\varepsilon,3}$	1.132	0.651 (1.137)

consistent with the event-study regressions (see Table 2).²⁰ 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.

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 $\sigma_{\varepsilon,1}, \sigma_{\varepsilon,2}, \sigma_{\varepsilon,3}$. The estimate of the size of U.S. local shock ($\sigma_{\varepsilon,1}$) in version 2 is a bit larger than the $\sigma_{\varepsilon,1}$ 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-U.S.) shocks, λ_{US}^f , which can be shown to be approximately equal to

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

²⁰GMM 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.

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). Note that λ_{US}^f depends not just on Γ_{ij} 's (which correspond to event study coefficients) but also on other quantities, including σ_i 's, σ_i^* 's, and T_i 's.

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, as elaborated in the Supplementary Material, 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 addition, 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. 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 nonnegligible 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.²¹

Finally, for a complete picture, we note that the corresponding numbers for the share of foreign shocks (noncountry-*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 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-U.K. 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.

IV. Evidence of Foreign Spillovers from Predictive Regressions

A. 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

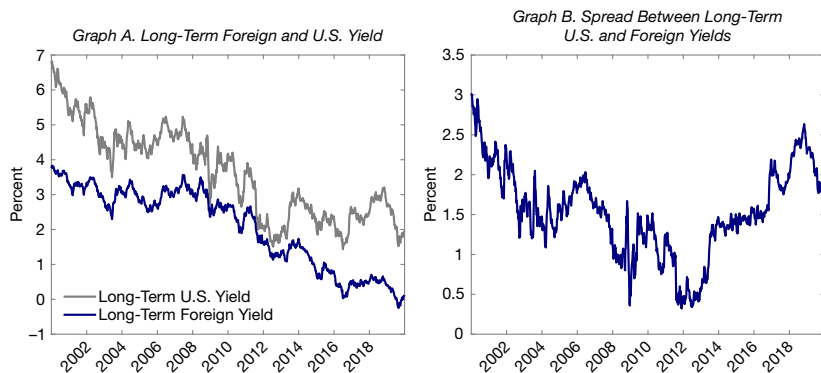
$$(17) \quad y_{f,t} = \sum_{c=1}^M w_{c,t} y_{c,t}^{(10)},$$

where the weight for country *c* is $w_{c,t} = \frac{GDP_{c,t}}{\sum_c^M GDP_{c,t}}$ and $M = 3$. Since GDP figures are quarterly and released with a delay of at least a quarter, the weight applied to the

²¹As shown in Sentana and Fiorentini (2001), assuming time-variation in volatility also allows the identification of all parameters. In a richer version of our model in which “the other” global shock η_t has a GARCH structure, we find the estimated η shocks are often small in magnitude, but can be sizable at certain times during our sample period, including the 2010–2011 period (Euro area debt crisis) and 2015 (PBOC’s yuan devaluation).

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

Graph A of Figure 3 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. Graph B plots the spread between long-term U.S. and foreign yields. The data is weekly and the sample period is Jan. 2000 to Dec. 2019.



weekly yields is constant within each quarter and corresponds to GDP figures from two quarters back to ensure that weights are known to the investor at time t . On average, over the period of 2000 to 2019, 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. Graph 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% to 0.1% over this period, a decline of more than 350 basis points. Similarly, the yield on long-term Treasury securities declined from 6.8% to 1.9% 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. Graph 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% with a standard deviation of 0.5% and a first-order autocorrelation coefficient of 0.989.

In what follows, we consider predictive regressions using the U.S.–foreign long-term yield spread as the main explanatory variable. An alternative approach in the literature for examining spillover of interest rates relies on the cointegration of long-term yields. If the cointegrating vector takes the form $z_t = y_{US,t} - \beta_1 y_{GE,t} - \beta_2 y_{UK,t} - \beta_3 y_{JP,t}$ with $\beta_i > 0$, our predictive regressions would be similar to the Vector Error-Correction model as in Kirchgässner and Wolters (1993) and Chinn and Frankel (1995).

B. Predictability of Intraday Moves in Long-Term Yields

Table 4 reports the results from predictive regressions of the form

TABLE 4
 Predictability of Intraday Changes in Long-Term Treasury Yields

Table 4 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 $\Delta y_{a,t+1}$), and the cumulative changes in the long-term yield outside of announcement windows (Panel C $\Delta y_{na,t+1}$). The key explanatory variable is the spread between the 10-year Treasury yield and the long-term foreign yield, $y_t - y_{f,t}$. The foreign yield is computed as the GDP-weighted average of German, Japanese, and U.K. 10-year zero-coupon yields. The control variables are the 10-year forward rate spread ($f_t - r_t$), the near-term forward spread ($E_t(r_{t+i}) - r_t$), the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{Aaa,t} - y_t$), and the effective duration of outstanding mortgage-backed securities (MBSDUR_t). All regressions include a constant and a dummy variable that is equal to 1 on 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 Jan. 2000 to Dec. 2019.

	Dependent Variable						
	Panel A. Δy_{t+1}			Panel B. $\Delta y_{a,t+1}$		Panel C. $\Delta y_{na,t+1}$	
	1	2	3	1	2	1	2
$y_t - y_{f,t}$	-1.59 [-2.46]	-3.49 [-3.12]		-0.04 [-0.14]	-0.24 [-0.56]	-1.55 [-2.75]	-3.25 [-3.64]
$f_t - r_t$		-1.24 [-2.34]	-0.17 [-0.49]		-0.17 [-1.16]		-1.07 [-2.15]
$E_t(r_{t+i}) - r_t$		1.60 [1.46]	-0.43 [-0.54]		-0.21 [-0.77]		1.81 [1.75]
$y_{Aaa,t} - y_t$		-1.56 [-2.22]	-1.51 [-1.92]		0.20 [-0.77]		-1.75 [-2.36]
MBSDUR _t		-0.83 [-2.06]	-0.86 [-2.20]		-0.08 [-0.75]		-0.75 [-1.90]
$R^2 \times 100$	0.496	1.424	0.533	-0.034	0.186	0.471	1.436
<i>T</i>	1,043	1,043	1,043	1,043	1,043	1,043	1,043

$$(18) \quad \Delta y_{t+1} = \alpha + \beta(y_t - y_{f,t}) + \gamma' \mathbf{x}_t + \varepsilon_{t+1}$$

and

$$(19) \quad \Delta y_{a,t+1} = \alpha_a + \beta_a(y_t - y_{f,t}) + \gamma'_a \mathbf{x}_t + \varepsilon_{a,t+1},$$

$$(20) \quad \Delta y_{na,t+1} = \alpha_{na} + \beta_{na}(y_t - y_{f,t}) + \gamma'_{na} \mathbf{x}_t + \varepsilon_{na,t+1}.$$

where Δy_{t+1} is the change in the long-term U.S. yield, $\Delta y_{a,t+1}$ is the yield change around macroeconomic and monetary policy announcements, and $\Delta y_{na,t+1}$ is the yield change outside of these windows.²² 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 \mathbf{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 II, we use intraday data on the 10-year on-the-run Treasury yield to compute the weekly cumulative change in the long-term yield around major macroeconomic and policy announcements $\Delta y_{a,t+1}$, and outside announcement times $\Delta y_{na,t+1}$. The weekly change in the 10-year yield is the sum of changes during announcement times and outside announcement times. The vector of controls \mathbf{x}_t

²²Recall from equation (1) that $\Delta y_{t+1} = \Delta y_{a,t+1} + \Delta y_{na,t+1}$. To our knowledge, Faust and Wright (2018) was the first to study the predictability of bond returns over announcement windows and nonannouncement windows separately.

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 ($y_{Aaa,t} - y_t$) 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 et al. (2016). We also include an indicator variable that equals 1 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 Jan. 2000 to Dec. 2019 and using OLS. We report t -statistics based on Newey and West (1987) standard errors with 52 lags to deal with the autocorrelation of the residuals.

Panel A of Table 4 presents the estimates for the overall weekly change in the long-term Treasury yield (i.e., equation (18)). 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 estimated coefficient 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 1.6 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, a 1-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.²³

The results in column 2 of Panel A of Table 4 show that the predictive power of the spread between U.S. and foreign long-term yields is robust to controlling for 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 (-3.49) 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 predictive ability of this single factor, as captured by the R^2 , is not only comparable to that of a regression that includes only 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 forward spread, near-term spread, the Aaa-Treasury spread, and MBS duration is 0.53%, as shown in column 3 of Panel A of Table 4. If we include the U.S.–foreign yields spread as a regressor, as shown in column 2, the R^2 increases threefold to 1.42%.

²³ Assuming that the U.S.–foreign long-term yield spread follows a first-order autoregressive process with an 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.

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, we find that the U.S.–foreign yield spread does not seem to predict yield fluctuations around macroeconomic announcements as the coefficient on this spread is not statistically or economically significant (Panel B), whereas the spread between U.S. and foreign yields is a strong predictor of future changes in the long-term yield outside of windows bracketing the release of key macroeconomic data (Panel C).

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 Litterman and Scheinkman (1991)), and have been shown to forecast bond returns around macroeconomic data releases (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–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 U.S.–foreign yield spread predict future changes in long-term Treasury yields outside macro announcements after controlling for the predictive power of the three PCs. The coefficient on the U.S.–foreign

TABLE 5
Predictive Regressions with Yield Curve Principal Components

Table 5 presents the estimated coefficients from predictive regressions. The dependent variables are weekly changes in the 10-year Treasury yield (Panel A $\Delta y_{t,t+1}$), cumulative weekly changes in 30-minute windows around macroeconomic and monetary policy announcements (Panel B $\Delta y_{\Delta,t,t+1}$), and the cumulative changes in yields outside of announcement windows (Panel C $\Delta y_{na,t,t+1}$). The key explanatory variable is the spread between the 10-year Treasury yield and the long-term foreign yield, $y_t - y_{f,t}$. The foreign yield is computed as the GDP-weighted average of German, Japanese, and U.K. 10-year zero-coupon yields. The control variables are 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 1 on 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 Jan. 2000 to Dec. 2019.

	Dependent Variable					
	Panel A. $\Delta y_{t,t+1}$		Panel B. $\Delta y_{\Delta,t,t+1}$		Panel C. $\Delta y_{na,t,t+1}$	
	1	2	1	2	1	2
$y_t - y_{f,t}$		-3.40 [-2.73]		-0.29 [-0.67]		-3.11 [-3.15]
L_t	-0.10 [-2.75]	0.06 [0.93]	-0.00 [-0.05]	0.01 [0.65]	-0.10 [-2.96]	0.05 [0.83]
S_t	0.28 [1.60]	0.35 [2.34]	0.14 [3.08]	0.15 [2.98]	0.13 [0.87]	0.20 [1.56]
C_t	-0.55 [-0.70]	-2.17 [-1.88]	0.24 [1.14]	0.10 [0.33]	-0.79 [-1.03]	-2.27 [-2.19]
$\bar{R}^2 \times 100$	0.382	1.117	0.165	0.107	0.385	1.110
T	1,043	1,043	1,043	1,043	1,043	1,043

yield spread remains highly significant and of roughly the same magnitude as in the specifications presented in Panel C of Table 4. Moreover, including the U.S.–foreign yield spread increases significantly the R^2 of the regression from 0.38% to 1.12%. 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.²⁴

Analogous to a cointegration equation, the U.S.–foreign yield spread can be interpreted as short-term departures of the long-term U.S. yield from an equilibrium relationship, or as a detrended component of the long-term yield.²⁵ We further investigate if the predictive ability of the U.S.–foreign yield spread withstands the inclusion of known factors that predict bond returns that are constructed using some type of detrending of nominal yields. In particular, we explore if the U.S.–foreign yield spread continues to predict the changes in the U.S. long-term yield once we include as control variables: i) the cycle measures of Cieslak and Povala (CP) (2015), which are defined as the portion of yields that is orthogonal to trend inflation; and ii) the “convergence gap” factor of Berardi et al. (2021), which is defined as the deviation of the current policy rate from the natural rate of interest rate. Since the CP cycle variables and the “convergence gap” are computed at a monthly frequency, we report results from these predictive regressions using monthly data starting in Jan. 2000 to Dec. 2019.²⁶

The results reported in column 1 of Table 6 show that a widening of the U.S.–foreign yield spread continues to predict a decline in the Treasury yield after controlling for the short-term cycle variable ($c_t^{(1)}$) and the 10-year interest rate cycle ($c_t^{(10)}$). In column 1 of Panels B and C, we continue to find that the U.S.–foreign yield spread is particularly informative about changes in Treasury yields outside announcement windows. Column 2 of Table 6 reports the regression results using the average cycle \bar{c}_t (computed using the portion of yields with maturity between 2 and 20 years that is orthogonal to trend inflation) instead of the maturity-specific cycle as an explanatory variable. We find that our results continue to be robust to the inclusion of this alternative regressor. Table 7 presents the results of predictive regressions where the vector of controls adds the “convergence gap” factor of Berardi et al. (2021) ($r_t^* - r_t$) to the set of economically motivated forecasting variables used in our baseline specification. As shown in Table 7, the spread between U.S. and foreign long-term yields continues to have a negative and statistically significant coefficient after controlling for the “convergence gap” factor of Berardi et al. (2021) and the other economically motivated control variables, with most of the predictive power explained by the predictability of changes in yields outside of windows bracketing macroeconomic and monetary

²⁴Using 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 null hypothesis that the U.S.–foreign yield spread does not have any incremental predictive power, namely, the R^2 increase of 1.12 is well outside the 95% bootstrap confidence interval [−0.096, 0.413].

²⁵We thank the referee for suggesting this interpretation of the U.S.–foreign yield spread.

²⁶See, Cieslak and Povala (2015) and Berardi et al. (2021) for the details about the construction of the CP yield cycle factors and the “convergence gap” factors, respectively.

TABLE 6
 Predictive Regressions Controlling for Short- and Long-Term Cycle Factors

Table 6 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 ($\Delta y_{a,t+1}$), and the cumulative monthly changes in yields outside of these windows ($\Delta y_{na,t+1}$). The key explanatory variable is the spread between the 10-year Treasury yield and the long-term foreign yield, $y_t - y_{f,t}$. The foreign yield is computed as the GDP-weighted average of German, Japanese, and U.K. 10-year zero-coupon yields. The control variables in column 1 are the Cieslak and Povala (2015) short-term cycle variable ($c_t^{(1)}$) and the maturity-specific interest rate cycle ($c_t^{(10)}$); and in column 2 the Cieslak and Povala (2015) short-term cycle factor ($\bar{c}_t^{(1)}$) and the average cycle factor (\bar{c}_t). We use Newey and West (1987) standard errors with 36 lags to deal with the autocorrelation of the residuals. *t*-statistics are reported in brackets. The sample is monthly and covers the period from Jan. 2000 to Dec. 2019.

	Dependent Variable					
	Panel A. Δy_{t+1}		Panel B. $\Delta y_{a,t+1}$		Panel C. $\Delta y_{na,t+1}$	
	1	2	1	2	1	2
$y_t - y_{f,t}$	-7.16 [-1.92]	-7.29 [-2.02]	-1.36 [-1.04]	-1.39 [-1.07]	-5.80 [-1.99]	-5.91 [-2.10]
$c_t^{(1)}$	2.09 [1.78]	2.38 [1.90]	1.26 [4.36]	1.35 [4.30]	0.83 [0.81]	1.03 [0.95]
$c_t^{(10)}$	-5.14 [-2.66]		-1.19 [-2.82]		-3.95 [-2.26]	
\bar{c}_t		-5.21 [-2.62]		-1.27 [-2.73]		-3.94 [-2.22]
$R^2 \times 100$	4.7	4.3	0.9	0.8	4.4	4.1
<i>T</i>	240	240	240	240	240	240

TABLE 7
 Predictive Regressions Controlling for Monetary Policy Cycles

Table 7 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 ($\Delta y_{a,t+1}$), and the cumulative monthly changes in yields outside of these windows ($\Delta y_{na,t+1}$). The key explanatory variable is the spread between the 10-year Treasury yield and the long-term foreign yield, $y_t - y_{f,t}$. The foreign yield is computed as the GDP-weighted average of German, Japanese, and U.K. 10-year zero-coupon yields. The control variables are the spread between the natural rate of interest rates and the real federal funds rate ($r_t^* - r_t$), the 10-year forward rate spread ($f_t - r_t$), the near-term forward spread ($\mathbb{E}_t(r_{t+i}) - r_t$), the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{Aaa,t} - y_t$), and the effective duration of outstanding mortgage-backed securities (MBS DUR_{*t*}). We use Newey and West (1987) standard errors with 36 lags to deal with the autocorrelation of the residuals. *t*-statistics are reported in brackets. The sample is monthly and covers the period from Jan. 2000 to Dec. 2019.

	Dependent Variable		
	Panel A. Δy_{t+1}	Panel B. $\Delta y_{a,t+1}$	Panel C. $\Delta y_{na,t+1}$
	$y_t - y_{f,t}$	-13.02 [-1.98]	-1.73 [-1.11]
$r_t^* - r_t$	2.15 [1.84]	-0.20 [-0.41]	2.35 [2.12]
$\mathbb{E}_t(r_{t+i}) - r_t$	5.01 [0.95]	-0.24 [-0.18]	5.25 [1.13]
$f_t - r_t$	-5.86 [-2.59]	-0.82 [-1.40]	-5.05 [-2.33]
$y_{Aaa,t} - y_t$	-3.69 [-1.91]	-1.40 [-1.56]	-2.29 [-1.01]
MBS DUR _{<i>t</i>}	-4.02 [-2.69]	-0.31 [-0.64]	-3.71 [-2.44]
$R^2 \times 100$	7.2	0.7	6.8
<i>T</i>	240	240	240

policy announcements. The robustness of our conclusions to controlling for the cycle variables of Cieslak and Povala (2015) and the “convergence gap” factor of Berardi et al. (2021) suggests that the foreign spread likely reflects moves in risk premia driven by foreign factors. For one thing, the short-maturity cycle variable

and the “convergence gap” factor control for fluctuations in the U.S. real short rate at business cycle frequencies and, as shown in Cieslak and Povala (2015) and Berardi et al. (2021), these factors capture the expectations component embedded in the U.S. long-term yield, which are likely not affected by foreign factors but by domestic conditions. Moreover, the CP factors will likely allow us to control for fluctuations in the risk premium driven by the U.S. business cycle.

The predictive power of the U.S.–foreign yield spread could depend on the underlying factors driving the moves in the spread. We explore if the predictive power of the spread between U.S. and foreign long-term yields increases when the overnight variance is higher than usual, which are likely periods that contain more information concerning the economic outlook abroad. As shown in Table A.9 of the Supplementary Material, when the overnight variance is above its long-run level, a widening of the U.S.–foreign long-term yield spread has a compressing effect that is about two-thirds larger relative to usual times. This result 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 economically larger subsequent moves in U.S. yields, supporting the idea that moves in foreign long-term yields spillover to U.S. long-term yields.

C. 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. Here, we further explore 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.²⁷ 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.

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

$$(21) \quad y_t^{(10)} = \frac{1}{10} \sum_{n=1}^{10} f_t^{(n)},$$

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

²⁷Various studies decomposing distant-horizon forward rates into short-rate expectations and term premia, including Kim and Wright (2005), find that moves in 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.

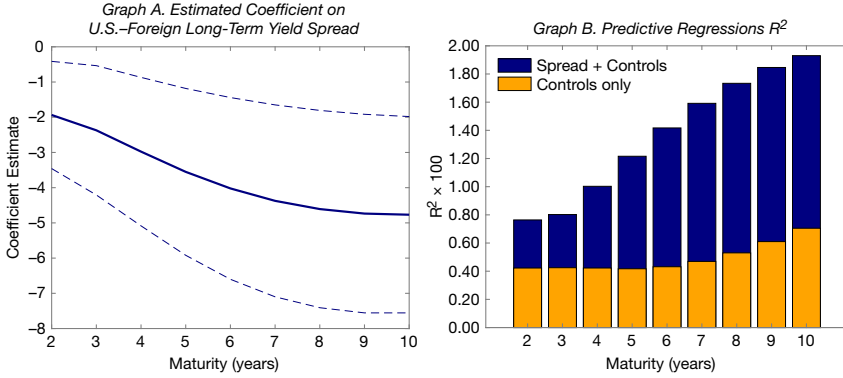
FIGURE 4

Predictability of U.S. Forward Rates by the U.S.–Foreign Long-Term Yield Spread

Graph A of Figure 4 plots the coefficient β_n from estimating the following predictive regression of 1-year forward rates:

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

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



$$(22) \quad \Delta f_{t+1}^{(n)} = \alpha_n + \beta_n (y_t - y_{f,t}) + \gamma_n' \mathbf{x}_t + \varepsilon_{t+1}^{(n)},$$

for $n = 2, \dots, 10$. Figure 4 plots the key coefficient of interest β_n along with a 90% confidence intervals and the associated R^2 for maturities $n = 2, \dots, 10$.²⁸

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 IV.B, the 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. Our sample is weekly from 2000 to 2019.

Graph A of Figure 4 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 toward far-forward rates. In particular, the estimated coefficients suggest that a 100 basis point widening 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.

Graph B of Figure 4 displays the R^2 s of the predictive regression in equation (22) 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^2 s. Graph B shows that including the U.S.–foreign yield spread as a predictor increases the R^2 and the additional forecasting power is

²⁸We 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).

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.²⁹

D. 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

$$(23) \quad rx_{t \rightarrow t+\tau} - rx_{f,t \rightarrow t+\tau} = \alpha + \beta(y_t - y_{f,t}) + \gamma' \mathbf{x}_t + \varepsilon_{t+\tau},$$

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

$$(24) \quad rx_{f,t \rightarrow t+\tau} = \sum_{c=1}^3 w_{c,t} rx_{c,t \rightarrow t+\tau}.$$

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

The coefficient estimates of [equation \(23\)](#) are obtained using weekly data from Jan. 2000 to Dec. 2019 and for holding period horizons of 1-week, 4-weeks, and 12-weeks ($\tau = 1, 4, 12$). 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 52 lags to deal with the autocorrelation of the residuals and overlapping observations when $\tau > 1$. As in [Section IV.B](#), we include in \mathbf{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 ($y_{Aaa,t} - y_t$), and the effective duration of MBS (MBSDUR_t).

The coefficient estimates, shown in [Table 8](#), suggest 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^2 s equal to 1%, 2%, and 5% for the 1-, 4-, and 12-week holding period horizons, respectively. The estimated coefficient is

²⁹Kearns et al. (2020) and Stedman (2020) also conclude that the yield curve spillovers largely occur through the term premium component of yields.

TABLE 8
 Predictability of Excess Returns on Long-Term Treasury Securities
 Over Long-Term Foreign Bonds

Table 8 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, $rx_{t-t+\tau} - rx_{t-t-\tau}$, with $rx_{t-t+\tau}$ defined as the GDP-weighted excess return on German, Japanese, and U.K. 10-year sovereign debt. Panels A–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 the long-term foreign yield, $y_t - y_{f,t}$. The foreign yield is computed as the GDP-weighted average of German, Japanese, and U.K. 10-year zero-coupon yields. The vector \mathbf{x}_t controls for the 10-year forward rate spread ($f_t - r_t$), the near-term forward spread ($E_t(r_{t+i}) - r_t$), the yield spread between Aaa-rated corporate bonds and Treasury securities ($y_{Aaa,t} - y_t$), and the effective duration of MBS (MBSDUR $_t$). All regressions include a constant and a dummy variable that is equal to 1 on 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 Jan. 2000 to Dec. 2019.

	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 [3.19]	20.83 [5.06]	6.51 [2.78]	15.96 [3.44]	4.46 [2.46]	10.56 [4.44]
$f_t - r_t$		7.05 [3.71]		5.66 [2.66]		3.62 [3.38]
$E_t(r_{t+i}) - r_t$		-10.34 [-2.67]		-6.84 [-1.59]		-3.52 [-1.42]
$y_{Aaa,t} - y_t$		2.18 [0.81]		3.16 [1.64]		1.68 [1.12]
MBSDUR $_t$		1.93 [1.22]		2.17 [1.44]		1.18 [1.28]
$\bar{R}^2 \times 100$	0.89	2.22	2.53	8.16	5.43	17.32
T	1,032	1,032	1,029	1,029	1,022	1,022

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 8, becomes more negative and remains highly statistically significant when we add control variables known to forecast U.S. bond returns.

E. Robustness Checks

This section summarizes additional exercises we performed to examine the robustness of our results. Full details are presented in the Supplementary Material. First, we show that the predictability we document is robust to reasonable variations to the way we compute the foreign yield. In particular, our results are robust to using the equally weighted average of German, Japanese, and U.K. long-term yields as well as to 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 (see Table A.4 of the Supplementary Material). We also perform our predictive regressions using monthly data and adding an additional 8 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 in U.S. long-term yields (see Table A.5 of the Supplementary Material).

We also show that the U.S.–foreign yield spread remains a strong predictor of the return on a strategy that is long on a long-term Treasury security and short on a long-term foreign bond after we control for the cycle variables of Cieslak and Povala (2015) and the “convergence gap” factor of Berardi et al. (2021). Finally, we 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. The return on this strategy is the same as that presented in equation (24) when the covered interest parity holds. Using this strategy, we also find 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.

V. Concluding Remarks

Yield spillover effects have been traditionally thought to run mainly from the U.S. to other countries. In this article, 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 the 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 IV. 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. Notable News Days Selection

To identify “notable news days” for the U.S., Euro area, and the U.K., 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. In particular, we identify days when financial news databases as well as internal

market intelligence reports attributed notable moves in Treasury, Bund, and/or Gilts markets to a specific development in the U.S., Euro area, or the U.K. For the U.S., we also identify days with sizable yield changes around major scheduled announcements using the reaction of the 10-year Treasury yield moves over a 30-minute window around the announcement. For the U.K. and Euro area, we use Citi Economic Surprise indices to help identify days with notable news, as these indices often show substantial changes on days with notable news.

The set of “notable days” can be classified as i) “data release” days, which are days with sizable market reactions to scheduled macro data releases; ii) “central bank communication” days, which are days with market-moving communications from central banks; and iii) “other news” days, which include days with notable moves in yields that are not linked to a data or monetary policy announcement (e.g., the 2016 presidential election) as well as days when it was not easy to point to a specific identifiable event but the financial press characterized as having been influenced by country i news.³⁰ We list the dates identified as days with notable news in Table A.3 of the Supplementary Material.

Appendix B. Identification of Model of Global Yields

To identify the model of global yields presented in equations (5), (7), and (8), note that we can identify $\Gamma_{(:,i)}$ and $\sigma_i^{*2} - \sigma_i^2$, with Γ_{ii} 's normalized to 1 from equation (11). Therefore, taking as $\Gamma_{(:,i)}$ as known, the following equation:

$$\Omega_0 = \sigma_1^2 \Gamma_{(:,1)} \Gamma'_{(:,1)} + \dots + \sigma_N^2 \Gamma_{(:,N)} \Gamma'_{(:,N)} + \sigma_\eta^2 u' + D \left(\left[\sigma_{\ell,1}^2, \dots, \sigma_{\ell,N}^2 \right] \right)$$

provides $N(N+1)/2$ equations to identify the following $2N+1$ unknown parameters $\sigma_1^2, \dots, \sigma_N^2, \sigma_\eta^2, \sigma_{\ell,1}^2, \dots, \sigma_{\ell,N}^2$. The full model is thus identified for values of N such that $N(N+1)/2 \geq 2N+1$, namely, $N \geq 4$. In our case, $N=3$, $3 \times 4/2 = 6 < 2 \times 3 + 1$. For this reason, we impose additional restrictions (i.e., equation (14) or equation (15)).

Supplementary Material

To view supplementary material for this article, please visit <http://doi.org/10.1017/S0022109022001429>.

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³⁰For 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’ speeches/comments.

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