Risk and Return in Government Bonds*

Matteo Leombroni Carolin Pflueger Adi Sunderam

October 23, 2026

Abstract

Does the compensation investors require for holding government bonds vary with their risk? We measure expected excess returns using market yields and professional forecasts of future yields, showing that they are tightly linked to a standard measure of risk: bonds' betas with respect to the stock market. In a month-by-maturity panel of U.S. Treasuries from 1988 to 2024, the correlation between expected excess returns and bond–stock betas is 67%. The correlation is driven by variation in market yields rather than forecasters' expectations. The relationship between realized excess returns and betas is weak in a univariate regression, but it becomes economically large and significant when we control for monetary-policy and macroeconomic forecast errors. Similar results obtain in an international panel of developed countries from 1989 to 2024. Our findings suggest that variation in risk premia driven by changing bond risk accounts for more than half of both the decline in U.S. long-term bond yields since the mid-1980s and their increase since 2020.

Keywords: long-term bond yields, bond–stock betas, safe assets, term premia, macroeconomic risks
JEL Classifications: E43, E52, E58

*Leombroni: leombron@bc.edu, Boston College, Carroll School of Management, Pflueger: cpflueger@uchicago.edu, University of Chicago, Harris School of Public Policy, NBER, and Federal Reserve Bank of Chicago, Sunderam: asunderam@hbs.edu, Harvard Business School and NBER. We thank Nick Bloom, John Campbell, Mike Golosov, Hanno Lustig, Alan Moreira, Kelly Shue (discussant), Andrea Vedolin (discussant), and Luis Viceira for helpful comments. We thank seminar participants at the Stanford SITE conference, Berkeley Haas, Wharton, San Francisco Fed, Boston Fed, MIT Sloan, Chicago Booth, DePaul University, Boston College, Boston University, NYU Stern, Columbia SIPA, the Princeton Macro-Finance Conference, the NBER Asset Pricing Summer Institute, and the Texas Macroeconomics conference for helpful comments. The views expressed here are those of the authors and do not necessarily reflect the views of others in the Federal Reserve System.

1. Introduction

Long-term Treasury yields experienced a multi-decade decline between the 1980s and the 2010s, followed by a dramatic reversal after the Covid pandemic. Whether these secular changes in bond yields were driven by changes in the natural rate of interest, the demand for safe assets, or the safety of Treasury bonds themselves is the subject of active debate. We show that the risk premium that investors require for holding bonds can be explained by a simple measure of risk, which is intuitively linked to risks in the macroeconomy.

Even when Treasury bonds are free of default, movements in long-term real rates and long-term inflation expectations can cause their prices to vary over time, potentially making them risky investments. We relate risk premia in long-term bond yields to an intuitive and economically meaningful measure of bond risks: the comovement between bonds and stocks. The "beta" of bonds can be viewed as a simple measure of how stock-like bonds are. It is a natural measure of risk, as stocks are often thought to be a good proxy for the marginal utility of investors.[1]The bond–stock beta is closely linked to the macroeconomy, as it is driven in part by the correlation between output and inflation (Campbell et al., 2026). We show empirically that the changing risks of Treasury bonds are priced by investors, and have contributed substantially to secular movements in long-term bond yields.

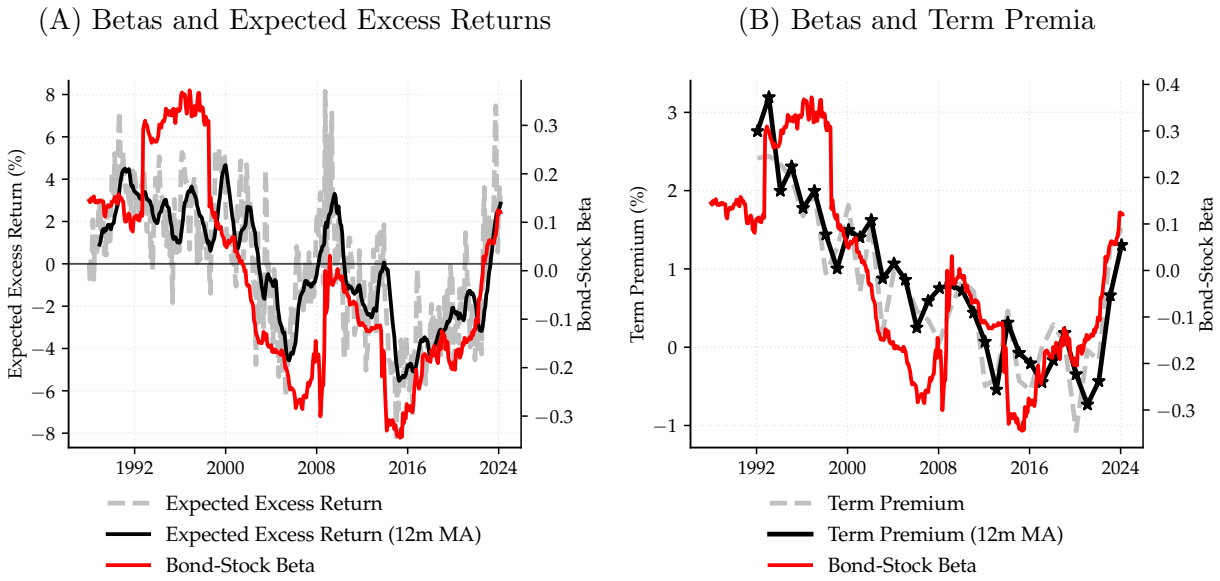
Figure 1 Panel A illustrates the close empirical relationship between expected excess returns – the compensation investors require for holding long-term bonds – and bond–stock betas for the 11-year Treasury bond.[2] The one-year expected excess return is constructed by comparing the current market yield on the 11-year Treasury to forecasts of the 10-year yield one year ahead. The red line in the figure captures the well-known fact that bonds changed from risky assets with positive betas to safe hedges with negative betas around 2000. The black line in the figure illustrates our main result: expected excess returns are highly correlated with this measure of bond risk. Before 2000, investors required a positive expected excess return, but after 2000 they were willing to hold Treasury bonds at negative expected excess returns. The figure also shows that the expected excess return required by investors has increased commensurately with bond risks after the Covid pandemic.

The relationship between risk and return is also reflected in a longer-horizon measure of expected bond excess returns. Figure 1 Panel B plots the 11-year Treasury yield minus survey forecasts of the expected average Treasury bill rate over the next 10 years—a measure

[1]The link is exact in the Capital Asset Pricing Model of Sharpe (1964) and Lintner (1965). More broadly, the logic of asset pricing described in Cochrane (2009) predicts that financial assets will be valued more highly if they pay out in high marginal utility states, such as recessions.

[2]Because forecasters report forecasts for the 10-year Treasury bond yield, the expected excess return on the 11-year bond is most naturally computed, and we use the 11-year Treasury bond throughout this figure. Betas are estimated over five-year rolling windows using monthly bond and stock excess returns.

Figure 1: Bond–Stock Betas, Bond Excess Returns, and Term Premia



Note: Panel A plots expected excess returns for the 11-year U.S. Treasury in dashed gray. One-year expected excess returns are computed according to Equation (7) using 4-quarter-ahead bond yield forecasts from the Blue Chip Financial Forecasts. The black line shows the 12-month moving average. In red, the figure plots rolling estimates of the stock market beta of the 11-year Treasury bond. The beta is computed by regressing monthly par bond excess returns on stock market excess returns over a backward-looking rolling window of 60 months, according to Equation (10). Panel B plots the same bond–stock beta in red and a measure of the long-horizon term premium in dashed gray. The term premium is the observed 11-year Treasury yield minus the expected average Treasury bill rate over the next 10 years, based on forecasts from the Survey of Professional Forecasters (SPF). The black line shows the 12-month moving average of the term premium. The sample runs from January 1988 to March 2024.

of the excess return expected over the life of the bond. Note that Panel B is based on different data than Panel A, as it does not require forecasts for long-term bond yields. Nonetheless, a similarly robust relationship emerges between long-horizon expected excess returns and bond–stock betas.

To guide the formal empirical analysis, we first develop a stylized term-structure model. In the model, there are two factors: (1) a standard term structure factor, such as the monetary policy rate or long-term inflation expectations, and (2) bond–stock betas. The model makes two predictions. First, expected excess returns rise with beta as long as the *market* demands compensation for beta risk. Rejecting the null of no relationship between expected excess returns and bond–stock betas is evidence that the market prices betas, not just that forecasters consider betas when making their forecasts. Second, if forecasters and the market have the same beliefs about the path of monetary policy, forecast errors are differenced out of expected excess returns but not realized excess returns. Betas will thus have a weaker relationship with realized returns than expected returns – as they do in the

data. The model predicts the relationship between realized excess returns and betas should reemerge when controlling for monetary policy and macroeconomic forecast errors.

Turning to the empirical analysis, we first formalize the empirical relationship between expected bond excess returns and rolling bond–stock betas shown in Figure 1. We use a month-by-maturity panel of U.S. Treasuries from January 1988 through March 2024. Expected excess returns are constructed using the average 4-quarter-ahead forecast of the 2-, 5-, and 10-year Treasury yields from Blue Chip Financial Forecasts (BCFF). Following Piazzesi et al. (2015) and Nagel and Xu (2023), we compute one-year expected bond excess returns analogously to one-year realized excess returns, replacing the realized bond yield one year ahead with its survey forecast. Bond–stock betas are computed from monthly bond and stock excess returns using five-year backward-looking rolling windows.

In our Treasury panel, we find a positive relationship between expected excess returns and betas that is economically and statistically significant. A unit increase in beta is associated with a 10 percentage point higher expected excess return. The magnitude and significance of the relationship are unchanged when we control for time trends, maturity and time fixed effects, predictors of realized bond returns like the term spread and the Cochrane-Piazzesi factor (Cochrane and Piazzesi, 2005), principal components of yields, or the Laubach-Williams estimate of the natural rate of interest (Laubach and Williams, 2003). The relationship is also similar for each bond maturity, and it does not rely on a specific choice of how the bond–stock beta is measured.

Next, we decompose expected excess returns into two components: one related to current market yields and the second related to forecasts of future long-term yields. Since expected excess returns depend positively on current market yields and negatively on forecasted future long-term yields, they load on bond–stock betas when current yields are more sensitive to betas than forecasted yields. This is what we find. Our results for expected excess returns are driven by the fact that market yields move in response to bond–stock betas. They are not simply a product of forecasters' methodologies.

We then show that the relationship between expected excess returns and bond–stock betas is not explained by inflation, simple forecasting rules based on mean reversion in yields, or slow learning. In other words, simple rules for forecasting long-term yields do not generate expected returns that are strongly correlated with betas, and our baseline finding is not simply an artifact of such rules.

Next, we study the relationship between realized excess returns and bond–stock betas. Under the full information rational expectations (FIRE) benchmark, the difference between realized and expected excess returns should merely reflect noise, and betas should therefore also predict realized bond excess returns. In practice, however, the raw relationship between

betas and realized bond excess returns is weak. We show that we can recover a strong, significant relationship by controlling for forecast errors about the path of monetary policy, GDP growth, and inflation. Put differently, yield surprises relative to forecasts were substantial in our sample, obscuring the relationship between betas and realized returns. Estimating our model with GMM, we can jointly match the relationship between one-year expected excess returns and betas, the relationship between longer-horizon term premia and betas, and the relationship between realized excess returns and betas controlling for monetary policy and macroeconomic surprises.

Finally, we show that these empirical facts carry over to an international panel of 12 developed economies from September 1989 to December 2024. Expected excess returns for non-U.S. countries are constructed using 12-month forecasts of 10-year government bond yields from Consensus Economics. The international panel adds important new time variation, with some countries, such as Japan, experiencing negative bond–stock betas in the mid-1990s, when U.S. betas were still positive. Matching this different pattern in betas, expected excess returns turned negative earlier in Japan than in the U.S. The estimated relationship between expected excess returns and betas is positive and indistinguishable from the U.S. coefficient for all countries in our sample, and is generally precisely estimated.

Our findings imply that a substantial portion of the variation in long-term yields can be attributed to the changing risks of bonds rather than the natural rate of interest, i.e., the rate at which monetary policy is neither expansionary nor contractionary (Wicksell, 1898; Bernanke, 2005; Summers, 2013). Our estimates provide new evidence on the fundamental drivers of long-term borrowing rates, suggesting that changing risk of Treasury bonds drove more than half of the four-decade decline in long-term nominal yields since the 1980s, and the majority of their increase since 2020.

1.1 Literature Review

We contribute to the large literature seeking to understand the decline in long-term interest rates across advanced economies since the 1980s, which has arguably had important implications for welfare (Gomez and Gouin-Bonenfant, 2024; Greenwald et al., 2026). A fruitful literature has sought to quantify the decline in the natural rate or "r-star" (e.g., Laubach and Williams, 2016; Bauer and Rudebusch, 2020; Bianchi et al., 2022a). As Reis (2026) highlights, these measures of r-star are generally based on the relationship between real GDP growth and interest rates and potentially understate the importance of risk and liquidity as drivers of secular changes in long-term bond yields.[3] Our contribution is to show

[3]A smaller literature in this area argues that increased demand for safe assets was important for the long-term decline in bond yields (Caballero et al., 2008; Del Negro et al., 2017). This literature has typically studied convenience yields in government bonds, as opposed to the term premia we focus on.

that changes in bonds' riskiness are closely linked to long-run changes in bond yields, with implications for their drivers going forward.

This paper speaks to the long-standing literature in asset pricing that studies the risks for which investors demand compensation. While the logic of asset pricing (Cochrane, 2009) implies that bonds should have negative expected excess returns if they act as hedges and pay out in high marginal utility states such as recessions, empirical evidence for this channel has so far been limited, making it difficult to link term premia to macroeconomic risk.[4] There is substantial evidence rejecting the expectations hypothesis – the idea that bonds have constant expected excess returns – by showing that realized bond excess returns vary predictably over time (e.g., Fama and Bliss, 1987; Campbell and Shiller, 1991; Cochrane and Piazzesi, 2005). Affine term structure models, building on Vasicek (1977) and Cox et al. (1985), allow for time-varying bond risk premia, but are largely silent on what macroeconomic risks those state variables represent.[5] Our approach is closest in spirit to evidence from currency markets, where Lustig and Verdelhan (2007) show that systematic exchange rate risk is priced and Kremens et al. (2025) show that expected currency returns line up with compensation for systematic risk. However, empirical evidence linking systematic risk to government bond pricing has so far been lacking.

Whether bonds are safe or risky plays a key role in macroeconomic models of the yield curve (e.g., Piazzesi and Schneider (2006); Rudebusch and Swanson (2012); Brunnermeier et al. (2024)). In Kekre et al. (2024), term premia reflect the risk that long-term bonds pose to specialized bond intermediaries. Gomez Cram et al. (2024) ask whether government debt inherits the risk of the fiscal surpluses that back it. Our paper provides some of the first evidence supporting and quantifying the premise that these models share: investors require additional compensation when the macroeconomic risks of bonds rise, and they accept negative expected excess returns when bonds become hedges.

Our results also relate to the literature analyzing the macroeconomic drivers of bond–stock comovements. Piazzesi and Schneider (2006) highlight that high inflation tends to predict poor future consumption growth, making nominal Treasury bonds unattractive assets. Baele et al. (2010) and Campbell et al. (2017) emphasize that bond risks, as measured by bond–stock betas, vary substantially over time. Campbell et al. (2020) show that bond–

[4]The classic Capital Asset Pricing Model (CAPM) of Sharpe (1964) and Lintner (1965) predicts that the market portfolio is the unique systematic factor and that expected excess returns are proportional to market beta. The empirical literature on the CAPM has largely focused on U.S. equities and found some but limited support for this idea (e.g., Black et al., 1972; Fama and MacBeth, 1973; Fama and French, 2004; Frazzini and Pedersen, 2014). Beyond equities, Fama and French (1993) show that realized bond and stock returns load on distinct term and default factors.

[5]Extensions incorporating macroeconomic risks, such as Ang and Piazzesi (2003) and Ang et al. (2006), take steps toward integrating bond markets with the broader economy, but do not use bond–stock betas.

stock betas are linked to inflation cyclicality, while Pflueger (2025) decomposes changes in bond–stock betas into changes in the nature of macroeconomic shocks and changes in monetary policy. Bianchi et al. (2022c) and Bianchi et al. (2022a) point to perceptions of monetary policy as a natural driver of bond risks, while Haddad et al. (2025) argue that bond risk is affected by market expectations of future quantitative easing. The evidence in this paper provides a new empirical link between bond–stock comovements and bond yields, making it even more relevant to understand the macroeconomic and policy drivers.

We build on tools from the recent literature investigating the role of survey expectations in asset prices. Following Piazzesi et al. (2015) and Nagel and Xu (2023), we define expected excess returns as those that would obtain if professional forecasts of long-term bond yields were realized. Different from Nagel and Xu (2023), who find that expected excess returns are hard to explain with many variables that predict realized returns, we show that they can be explained well by bond–stock betas.[6] In a large sample with rational expectations, the gap between expected and realized bond excess returns would merely be noise. In practice, however, forecast errors can be substantial and systematic (Cieslak, 2018). Such errors may arise from slow but rational learning (Farmer et al., 2024) or behavioral biases (Bianchi et al., 2022b; Shue et al., 2025; Molavi et al., 2025). Our measure of expected excess returns differences out forecast errors, allowing us to study how the market prices bond risks given its expectations. We find that bond–stock betas and forecast errors play complementary roles in explaining realized bond excess returns.

Beyond bonds, a large literature shows that survey expectations contain useful information about asset prices across markets, including equities (Bordalo et al., 2024; Delao and Myers, 2021). Different from the literature on subjective asset pricing in equities (e.g., Brav et al., 2005; Jensen, 2026), our findings are not a direct product of the methodologies used by forecasters, but provide evidence on how the market prices risk. Forecasters in the BCFF or Survey of Professional Forecasters (SPF) come from a variety of backgrounds and do not directly report betas in their analyses. We find that the relationship between market yields and bond–stock betas is crucial for explaining our results. Expectations data serve as a long-term or risk-neutral benchmark, which helps to uncover this empirical relationship.

[6] Couts et al. (2026) take a different approach, showing that institutional investors' long-run return expectations are consistent with their expectations for covariances, with a median forecast horizon of ten years. Different from them, we consider actual bond–stock betas and show that they are linked to actual market yields. Within fixed income, Pesch et al. (2024) link bond return volatility to expected excess returns, while Rogers (2026) uses derivatives to construct short-horizon expectations. We study longer-horizon risk premia for long-term bonds and find that betas are priced.

2. Simple Model of Bond Risks and Expectations

This section presents a stylized term structure model to guide the empirical analysis. Two key lessons emerge. First, the empirical link between bond–stock betas and expected excess returns is closely related to the compensation *markets* require for holding high-beta assets, not just the process macroeconomic forecasters use to predict future yields. This link holds even when markets and forecasters have beliefs about long-term inflation or monetary policy that differ from the econometrician's, and if forecasters either correctly weight or underweight betas when forming their long-term yield forecasts. Second, the gap between realized and expected bond excess returns reflects forecast errors about long-term bond yields. Thus, under certain conditions, controlling for monetary policy forecast errors can help uncover the relationship between realized returns and betas.

Letting t denote time in months, we assume for simplicity that the one-year nominal yield is driven by a single factor, f_t:

$$y_{1,t} \;=\; f_t, \qquad f_{t+12} = \rho_f f_t + \varepsilon_{t+12}. \tag{1}$$

The residual ε_{t+12} is assumed to be uncorrelated with time-t variables. The factor f_t can represent the short-term interest rate or longer-run variables like perceptions of the central bank's inflation target or the natural rate of interest, as in shifting end-point models (e.g., Kozicki and Tinsley, 2001; Bauer and Rudebusch, 2020). For simplicity and concreteness we refer to f_t as the "monetary policy factor" and the related forecast errors as "monetary policy forecast errors", keeping in mind that this factor may represent either variation over the monetary policy cycle or more persistent shifts in monetary credibility.

Letting n denote bond maturity in years, the n-year zero-coupon bond–stock beta is assumed to follow a simple mean-reverting process:

$$\beta_{n,t+12} \;=\; \rho_\beta \beta_{n,t} + \eta_{t+12}, \tag{2}$$

where η_{t+12} is uncorrelated with time-t variables but may be correlated with ε_{t+12}. In order to price bonds as they roll down the maturity structure, we make the additional auxiliary assumption that bond–stock betas are proportional to maturity: $\beta_{i,t} = \frac{i}{n}\beta_{n,t}$.[7] We allow markets and forecasters to have different beliefs. We denote rational expectations by $\mathbb{E}$, forecaster expectations by $\widetilde{\mathbb{E}}$, and market expectations by $\widehat{\mathbb{E}}$. Associated perceived parameters are defined analogously.

The key assumption is that the market prices bonds' comovement with the stock market:

[7] We make this assumption for simplicity, but our empirics do not depend on it.

Assumption 1. (Required Return) *Markets require an expected bond excess return proportional to their expected bond–stock beta with a price of risk $\lambda > 0$ and a positive perceived persistence of beta $\hat{\rho}_\beta > 0$: $\widehat{\mathbb{E}}_t xr_{n,t+12} = \lambda \widehat{\mathbb{E}}_t \beta_{n,t+12} = \lambda \hat{\rho}_\beta \beta_{n,t}$*

This assumption captures the central channel in a wide range of macroeconomic models of bond yields. The risk of bonds depends on how their values covary with the marginal utility of bond holders, which is typically closely correlated with the stock market in such models. We directly model the relationship between expected bond excess returns and bond–stock betas, capturing a broad range of microfoundations and agent setups.

Forecasters are assumed to predict future long-term yields using the market's pricing of beta but their own view of beta dynamics. Forecasters in the model anticipate that in one year the market will price bonds according to Assumption 1, while forecasting that beta mean-reverts with persistence $\tilde{\rho}_\beta \geq 0$. Allowing different λ parameters for markets and forecasters would act similarly.

Throughout, we make the simplifying assumption that forecasters and markets share the same beliefs about the monetary policy factor, f_t, i.e., $\tilde{\rho}_f = \hat{\rho}_f$, which may or may not coincide with the econometrician's, ρ_f. Assuming that the market's and forecasters' beliefs about monetary policy coincide is consistent with the evidence in Cieslak (2018), which shows that monetary policy forecast errors from survey forecasts are highly correlated with market-based monetary policy forecast errors using fed funds futures. We do not impose that forecaster beliefs about betas coincide with the market's to allow for the possibility that macroeconomic forecasters are less attuned to term premia than the market.

The expressions for realized, market-expected, and forecaster-expected log bond excess returns are then:

$$
\begin{aligned}
xr_{n,t+12} &= n y_{n,t} - (n-1) y_{n-1,t+12} - y_{1,t} && (3)\\
\widehat{\mathbb{E}}_t xr_{n,t+12} &= n y_{n,t} - (n-1) \widehat{\mathbb{E}}_t y_{n-1,t+12} - y_{1,t} && (4)\\
\widetilde{\mathbb{E}}_t xr_{n,t+12} &= n y_{n,t} - (n-1) \widetilde{\mathbb{E}}_t y_{n-1,t+12} - y_{1,t} && (5)
\end{aligned}
$$

Here, $y_{n,t}$ denotes the current market yield on an n-year zero coupon bond in month t and $y_{n-1,t+12}$ is the yield on an $(n-1)$-year zero coupon bond in month $t+12$. The current market yield enters all three expressions. They only differ in whether they incorporate realized or expected future yields.

Iterating the expression for the market-expected bond excess return gives:

$$y_{n,t} - \underbrace{\frac{1}{n}\sum_{i=0}^{n-1}\widehat{\mathbb{E}}_t y_{1,t+12i}}_{\text{expected short rate}} = \underbrace{\frac{1}{n}\sum_{i=1}^{n-1}\widehat{\mathbb{E}}_t xr_{n-i+1,t+12i}}_{\text{term premium}} = \lambda\left(\sum_{i=1}^{n-1}\widehat{\rho}_\beta^{\,i}\frac{n-i+1}{n^2}\right)\beta_{n,t} \tag{6}$$

Expression (6) shows that actual market yields depend on the market's expectations about (i) the path of short rates and (ii) the term premium, i.e., the path of excess returns over the life of the bond. Bond yields rise with bond–stock betas, provided that the market prices beta ($\widehat{\rho}_\beta > 0$, $\lambda > 0$).

The following proposition describes realized and expected bond excess returns.

Proposition 1. (Realized and expected excess returns) *There exists a function* $\Lambda_n(\hat{\rho}_\beta, \rho_\beta)$, $\frac{\partial \Lambda_n}{\partial \hat{\rho}_\beta} > 0$, $\frac{\partial \Lambda_n}{\partial \rho_\beta} < 0$, $\Lambda_n(0, \rho_\beta) = 0$, $\Lambda_n(\hat{\rho}_\beta, \hat{\rho}_\beta) = \hat{\rho}_\beta$ *such that:*[8]

a. Realized excess returns: $xr_{n,t+12} = \lambda\Lambda_n(\hat{\rho}_\beta, \rho_\beta)\beta_{n,t} - \frac{1-\hat{\rho}_f^{n-1}}{1-\hat{\rho}_f}\left((\rho_f - \hat{\rho}_f)f_t + \varepsilon_{t+12}\right) + u_{t+12}$, *where* u_{t+12} *is an uncorrelated residual.*

b. Expected excess returns: $\widetilde{\mathbb{E}}_t xr_{n,t+12} = \lambda\Lambda_n(\hat{\rho}_\beta, \tilde{\rho}_\beta)\beta_{n,t}$.

Proposition 1(a) says that realized bond excess returns increase with the bond–stock beta and decline with the market's ex-post forecast error. For instance, if we think of f_t as the policy rate, a positive surprise about the near-term path of monetary policy that was unpredictable to market participants lowers realized returns. Similarly, if we think of f_t as long-term inflation expectations, unexpected increases in the inflation outlook reduce realized returns. How strongly realized excess returns load on bond–stock betas depends on the market-perceived persistence of betas because this determines the required compensation for beta risk and hence actual market yields.[9]

Model expected excess returns, by contrast, depend only on bond–stock betas (Proposition 1(b)). Note that here the monetary policy factor f_t does not enter. Said differently, to the extent that forecasts provide a reasonable approximation of markets' perceptions about monetary policy, forecast errors are fully differenced out.[10] This makes it easier to detect

[8] The function is $\Lambda_n(\hat{\rho}_\beta, \rho_\beta) = \hat{\rho}_\beta + (\hat{\rho}_\beta - \rho_\beta)\frac{1}{n}\sum_{i=1}^{n-2}(n-i)\hat{\rho}_\beta^{\,i}$. See Appendix A.

[9] The actual persistence of betas may also enter, if market perceptions about the persistence of beta differ from the econometrician's.

[10] The assumption that $\tilde{\rho}_f = \hat{\rho}_f$ maximizes the benefits of studying expected excess returns as opposed to realized returns, as forecast errors about f_t are fully differenced out of expected excess returns. If forecaster and market beliefs about f_t are not identical but similar, the qualitative implications are similar, but the analytic expressions are more complicated.

a relationship between expected excess returns and bond–stock betas than between realized excess returns and betas.

We next make an additional assumption on the perceived persistence of betas.

Assumption 2. (Beta Perceptions) *Forecasters perceive bond–stock betas to be no more persistent than the market does:* $\widetilde{\rho}_\beta \leq \widehat{\rho}_\beta$.

Note that Assumption 2 nests two corner cases: (1) the case where forecasters share the market's expectations about beta ($\tilde{\rho}_\beta = \hat{\rho}_\beta$) and (2) the "pure macroeconomic forecasters" case where forecasters do not incorporate beta and hence term premia into their forecasts at all ($\tilde{\rho}_\beta = 0$). We then obtain the following corollary.

Corollary 1. *Under Assumption 2, the loading of expected excess returns onto bond–stock betas is positive provided that the market prices bond–stock betas, i.e.* $\lambda, \hat{\rho}_\beta > 0$.

Corollary 1 says that a positive relationship between expected excess returns and bond–stock betas indicates that the *market* prices bond–stock betas, but not necessarily whether forecasters perceive betas to be priced. Neither forecaster nor market rationality is required for Corollary 1, though full information rational expectations (FIRE) is nested as a special case. In this special case, the bond–stock beta loading equals $\hat{\rho}_\beta \times \lambda$.

Holding fixed market beliefs about the persistence of beta, as forecasters' perceived persistence of beta $\tilde{\rho}_\beta$ falls, the relationship between expected excess returns and beta strengthens. Intuitively, when betas are high, forecasters expect them to mean revert faster than the market. Thus, they expect yields to fall faster and hence expected returns to be higher. In other words, if the market prices beta, but forecasters do not, the loading of expected excess returns on betas will exceed $\hat{\rho}_\beta \times \lambda$. Conversely, if forecasters price beta but the market does not, i.e., Assumption 2 fails, the model predicts a negative beta loading, contrary to data.

We think Assumption 2 is a reasonable approximation of reality. In our data, forecasters are predominantly macroeconomic forecasters, with titles such as "Chief Economist." In major financial institutions, these forecasters typically work separately from portfolio allocation and trading functions and emphasize macroeconomic fundamentals like long-run inflation rather than explicit term-premium considerations. They forecast yields rather than expected excess returns and, unlike equity analysts, they do not report explicit beta calculations. Instead, they typically forecast macroeconomic fundamentals like GDP growth and inflation using VARs or structural models, and then relate bond yields to these fundamentals using so-called "fair value" models.[11] Thus, the model clarifies that a positive relationship

[11] See the Philadelphia Fed's survey of forecasting methods for an informative overview of the methods used by a similar set of macroeconomic forecasters. A specific fair-value model, the "Sudoku" model, used by Goldman Sachs since 2007, is described in detail here. We are grateful to Cary Leahey for confirming that this class of models has been used by major forecasters since the 1980s.

between expected excess returns and betas should be viewed as a rejection of the null that the market does not price betas (i.e., an acceptance of the alternative that $\lambda, \hat{\rho}_\beta > 0$).

While so far we have not made any assumptions that rule out full rationality, we next show that the gap between realized and expected excess returns becomes meaningful when FIRE is violated. We then obtain the following corollary:

Corollary 2. *If beliefs about the monetary policy factor deviate from FIRE ($\tilde{\rho}_f = \hat{\rho}_f \neq \rho_f$), then realized bond excess returns load onto bond–stock betas and serially correlated monetary policy forecast errors, whereas expected excess returns depend only on bond–stock betas.*

Corollary 2 says that, beyond finite-sample noise, there can be systematic reasons for expected excess returns to be more closely linked to bond–stock betas than realized excess returns. Realized returns compare today's market yield to the realized future yield, so systematic forecast errors generated by the market's beliefs about the persistence of f_t show up as ex-post return surprises. By contrast, in the construction of expected excess returns, beliefs shared by markets and forecasters affect both today's market yield and the forecast of next year's yield and hence cancel out. Thus, expectations inconsistent with FIRE are potentially significant for understanding realized excess returns, while we study the complementary component of bond yields determined by expected excess returns. In a sample like ours with persistent downward trends in interest rates and long-term inflation, it is plausible that repeated forecast errors led to repeated yield surprises, and hence a gap between expected and realized bond excess returns. The assumption that forecasters and markets make serially correlated forecast errors accords with a growing literature and could be driven by either behavioral biases or the difficulty of learning about persistent processes even in relatively large samples (Cieslak, 2018; Bauer et al., 2024; Farmer et al., 2024; Shue et al., 2025).

Overall, the model clarifies that a positive relationship between expected excess returns and bond–stock betas arises as long as the *market* prices betas. Forecaster rationality is not required, but also not ruled out. Deviations from FIRE can help rationalize a weaker empirical relationship between realized bond excess returns and bond–stock betas. In this case, expected bond excess returns are particularly useful to test whether the market prices bond–stock betas, and the model predicts that controlling for ex-post monetary forecast errors can help recover the relationship between realized excess returns and bond–stock betas.

3. U.S. Evidence

This section presents our main results for the U.S.

3.1 U.S. Data and Summary Statistics

The data for our U.S. sample come from two main sources. First, we obtain forecasts of future long-term bond yields from the Blue Chip Financial Forecasts (BCFF). The survey asks for forecasts of interest rates, including the federal funds rate and Treasury yields of several maturities. Participating institutions include large investment banks, insurance companies, and economic research consultancies, with between 30 and 50 institutions participating in each survey. Our sample is monthly, starting in January 1988, when a consistent set of Treasury yield forecasts becomes available. Our second main source of data is nominal Treasury par yields from Gürkaynak et al. (2007). We use par yields to match the object that BCFF participants forecast.

Expected Bond Excess Returns Following the data construction of Piazzesi et al. (2015) and Nagel and Xu (2023), we compute the expected bond excess return for an n-year coupon bond by combining the current market yield on the n-year bond with the forecasters' one-year-ahead expectation of the $(n-1)$-year yield. Specifically, we substitute forecasters' expectations of future bond yields from Blue Chip into the standard expression for a coupon bond return:

$$\tilde{E}_t xr_{n,t\to t+12} \quad = \quad dur_{n,t} y^{par}_{n,t} - (dur_{n,t} - 1)\tilde{E}_t y^{par}_{n-1,t+12} - y_{1,t}. \tag{7}$$

This is the analog of Equation (5) in the model for an n-year coupon bond with duration $dur_{n,t}$. $y^{par}_{n,t}$ denotes the current market bond yield, $\tilde{E}_t y^{par}_{n-1,t+12}$ denotes forecasters' prediction for the $(n-1)$-year yield in 12 months, and $y_{1,t}$ is the one-year yield, in annualized percent units.[12] Note that forecasters predict future yields, not future excess returns. The expected return rises with the current market yield and falls with the forecasted future yield.

Since Blue Chip contains forecasts for the 2-, 5-, and 10-year bond yields, we can construct expected excess returns for a panel of 3-, 6-, and 11-year government bonds. We proxy for the 12-month ahead forecast with 4-quarter-ahead forecasts, averaged across forecasters. We approximate the duration of the n-year bond selling at par as $dur_{n,t} = \left(1 - (1 + y^{par}_{n,t})^{(-n)}\right) / \left(1 - (1 + y^{par}_{n,t})^{(-1)}\right)$ (Campbell, 2017, pp. 236–237). Blue Chip forecasters are required to submit their responses at the end of the previous month. To make sure the information sets are consistent, the March survey is paired with the realized par yield at the end of February in Equation (7).

This procedure gives us a monthly series for U.S. expected bond excess returns over a

[12] For the one-year risk-free rate we use the market yield on U.S. Treasury Securities at one year constant maturity downloaded from FRED (ticker `DGS1`). We average the realized par yield over the last week of each month to reduce day-to-day noise because we do not know the exact date forecasters make their forecasts.

sample from January 1988 through March 2024. With monthly data and three different bond maturities ($n = 3, 6, 11$), we have roughly 1300 observations, though of course we must account for common variation across maturities and time in our regressions.

Realized Bond Excess Returns We construct realized 12-month bond excess returns for an n-year par bond:[13]

$$xr_{n,t\to t+12} = dur_{n,t}\, y^{par}_{n,t} - (dur_{n,t} - 1)\, y^{par}_{n-1,t+12} - y_{1,t}. \tag{8}$$

The summary statistics in Table I Panel A show that, in contrast to realized bond excess returns, expected excess returns on average were close to zero over our sample, but they exhibited substantial variation over time. The 12-month moving average of expected excess returns on the 3-year bond ranges from -1.21% to 1.55%, with a standard deviation of 0.71%.[14] The moving average of expected excess returns on the 11-year bond varies more, ranging from -5.53% to 4.67%, with a standard deviation of 2.70%.

One might wonder whether our measures of expected bond excess returns are affected by forecaster incentives or institutional constraints. A simple split by forecaster type suggests that this is not a first-order concern. We classify forecasters in the Blue Chip data into three groups: dealers (e.g., Goldman Sachs), buy-side firms (e.g., Swiss Re), and research firms (e.g., Oxford Economics). Despite their different incentives and customer bases, the average yield forecasts for the three subgroups are highly correlated. Across any two forecaster types, the pairwise correlation for the 11-year expected bond excess return is at least 0.89.[15]

Bond Betas The beta of Treasury bonds with respect to the stock market is estimated using monthly data. Monthly U.S. stock market excess returns are obtained from Ken French's website. The Treasury bond excess return from month t to $t+1$ is computed using nominal Treasury par yields from Gürkaynak et al. (2007) as

$$xr_{n,t+1} \;=\; dur_{n,t} y^{par}_{n,t} - \left(dur_{n,t} - \frac{1}{12}\right) y^{par}_{n,t+1} - y_{\text{rfr},t}, \tag{9}$$

[13] Consistent with our construction of expected excess returns, we average realized par yields over the last week of each month.

[14] To reduce month-to-month noise, we also construct a smoothed series by taking the average expected excess returns over the preceding twelve months. Higher-frequency fluctuations in expected bond excess returns could arise, for example, if forecasters update their forecasts incompletely month-to-month while actual bond yields fluctuate. Because we use expected excess returns primarily as the dependent variable in our regressions, this should simply act as white noise.

[15] Appendix Table B6 reports the full set of pairwise forecaster-type correlations across bond maturities. Forecaster-type yield forecasts are plotted in Appendix Figure B4.

where the monthly risk-free rate $y_{rfr,t}$ is from Ken French's website, and the duration of the n-year par bond in month t is computed assuming that bonds sell at par.

The beta of a bond with maturity n in month t is estimated by regressing Treasury bond excess returns on stock market excess returns using monthly data:

$$xr_{n,t} = a_{n,t} + \beta^{w}_{n,t} xr^{eq}_{t} + \varepsilon_{t}. \tag{10}$$

Our baseline estimate of $\beta^{w}_{n,t}$ uses a rolling 5-year window ($w = 60$ months). This window is consistent with evidence in Couts et al. (2026) on how institutional investors form their perceptions of bond–stock comovements, but we show below that our results are robust to varying this choice. Summary statistics are provided in Table I Panel B.

Figure 1 plots expected excess returns and bond–stock betas for the 11-year Treasury.[16] Expected excess returns were positive until the turn of the millennium, then turned negative after 2000. They fluctuated around zero during the global financial crisis and its aftermath. Negative expected bond excess returns prevailed during the 2010s and early 2020s, followed by a return to positive expected bond excess returns in 2023 and early 2024. The parallel to bond–stock betas in Figure 1 is visually apparent.

Betas move in a macroeconomically intuitive way, as documented by the previous literature. They were positive before the turn of the millennium, but negative thereafter. This is what one would expect if inflation tended to rise in recessions before 2000 ("stagflations") but tended to fall in recessions thereafter, since returns on nominal bonds are inversely related to expected inflation and stocks move with the business cycle.[17]

In our regressions below, our estimated betas are technically generated regressors. However, as Table I Panel B shows, they are precisely estimated.

Alternative Beta Estimates Our baseline bond–stock beta is estimated using monthly returns over a 5-year rolling window. In addition to this baseline measure, we estimate the beta using a shorter 1-year window ($w = 12$ months). In this case, we use daily bond and equity excess returns. Appendix Figure B2 plots the 5-year and 1-year rolling betas. The pairwise correlation between the two, averaged across bond maturities, is 0.78.

We also consider a "forward" version of the 1-year rolling beta, where we use the 1-year beta estimated at time $t + 12$, and hence over the period from t to $t + 12$. This shifts the 1-year beta in Figure B2 forward by one year. Expected excess returns are constructed at

[16] We also construct the same plot for the 3-year maturity and 6-year maturity in Appendix Figure B1.

[17] The correlation between annual CPI inflation and the output gap was negative at -0.24 between 1988q1 and 1999q4, but positive at 0.47 from 2000q1-2019q4. The 11-year bond–stock beta moved inversely, averaging 0.22 in the pre-2000 period and -0.15 in the post-2000 period. Correlations between bond–stock betas and rolling estimates of macroeconomic and monetary policy risk are reported in Appendix B.1.

Table I: Summary Statistics

Panel (A): Expected Excess Returns and Realized Excess Returns

	Expected Excess Returns			Expected Excess Returns (MA)			Realized XR		
	3Y	6Y	11Y	3Y	6Y	11Y	3Y	6Y	11Y
Average	0.09	-0.03	-0.05	0.08	-0.05	-0.10	1.03	2.06	2.83
Std	0.84	1.76	3.11	0.71	1.46	2.70	2.45	4.77	7.25
Min	-2.03	-4.23	-8.18	-1.21	-3.17	-5.53	-6.99	-13.20	-21.36
Max	2.17	4.42	8.16	1.55	3.06	4.67	6.52	12.60	17.60

Panel (B): Bond–Stock Beta

	5-y rolling β		
	3Y	6Y	11Y
Average	-0.01	-0.02	-0.02
Std	0.07	0.13	0.20
Min	-0.17	-0.26	-0.34
Max	0.14	0.25	0.37
Average S.E.	0.03	0.05	0.08

Note: Panel A reports summary statistics for expected excess returns, the 12-month moving average (MA) of expected excess returns, and realized excess returns for bonds with 3-, 6-, and 11-year maturities. Expected bond excess returns are computed according to Equation (7), using 4-quarter-ahead par bond yield forecasts from BCFF. Panel B reports summary statistics for rolling estimates of bond–stock betas for par bonds with 3-, 6-, and 11-year maturities. Bond–stock betas are computed as the regression coefficients of par bond excess returns on stock market excess returns over a backward-looking rolling window, as defined in Equation (10). We use monthly returns over a 60-month rolling window (5-year rolling β). The table reports the average estimate, standard deviation, minimum and maximum values, and average robust standard errors. The sample of expected excess returns runs from January 1988 to March 2024, so that the corresponding sample of realized excess returns runs through March 2025.

time t for the return horizon from t to $t+12$. The 1-year forward beta is therefore estimated using information contemporaneous with the horizon over which expected excess returns are defined. Bond prices, and hence expected excess returns, should in principle reflect investors' expectations of future betas. If investors have perfect foresight, the realized beta would be the relevant measure. In practice, betas are highly autocorrelated so these adjustments do not change our conclusions.

We also estimate betas using a 5-year window with daily returns, which is a smoother version of the baseline beta. For betas estimated using daily data we consider an alternative where we apply the Dimson (1979) adjustment to account for non-synchronous trading.[18] Finally, we construct bond–stock betas using exponentially decaying weights with a half-life of three years. These beta estimates are shown in Appendix Figure B3. Our main results

[18]The Dimson beta is obtained by including two lags and two leads of the market return in the regression and summing the corresponding coefficients. We estimate $xr_{n,d} = \alpha_n + \sum_{k=-2}^{2} \beta_{n,k}\, xr_{d+k}^{mkt} + \varepsilon_{n,d}$, and define the Dimson beta as $\beta_n^{Dimson} = \sum_{k=-2}^{2} \beta_{n,k}$.

are robust across all of these alternatives.

Other Variables Our analysis controls for additional predictors of bond excess returns. We construct a measure of bond return variance using the rolling variance of bond excess returns, computed over a 5-year window of monthly returns. We further include controls for the shape of the yield curve. Specifically, we compute the level, slope, and curvature factors as the first three principal components of zero-coupon nominal bond yields with maturities of 1, 2, 5, 7, 10, 15, and 20 years. We define the term spread as the difference between the n-year par yield and the one-year yield.[19] The term spread is a traditional predictor of bond excess returns (Campbell and Shiller, 1991), though its influence has been found to be weaker for expected excess returns than for realized ex-post bond excess returns (Nagel and Xu, 2023). We also include the tent-shaped factor from Cochrane and Piazzesi (2005), where we apply the originally estimated coefficients to monthly Fama–Bliss discount bond price data. We also control for the natural rate of interest, r-star. We use the Laubach-Williams estimates of r-star, available from the Federal Reserve Bank of New York website and constructed following Laubach and Williams (2003).

Long-Term Forecasts We use long-term forecasts of CPI inflation and the Treasury bill rate from the Survey of Professional Forecasters (SPF). The SPF is a quarterly survey conducted in the middle month of each quarter. So, for example, we match Q1 SPF long-term forecasts to the March BCFF, collected at the end of February. The survey provides forecasts of the average CPI inflation rate over the next ten years at a quarterly frequency. We also use forecasts of the average U.S. Treasury bill rate over the next ten years, which are available at an annual frequency.

3.2 U.S. Expected Excess Returns vs. Betas

Our stylized model predicts that expected excess returns increase with bond–stock betas, provided that the market prices bond–stock betas and that forecasters put weakly less weight on betas than the market (Corollary 1). We now provide evidence for this prediction using our monthly U.S. panel, estimating the following panel regression:

$$\tilde{E}_t xr_{n,t\to t+12} = \alpha + \lambda\beta_{n,t} + X'_{n,t}\delta + \eta_{n,t}, \tag{11}$$

where $\beta_{n,t}$ denotes the estimated rolling beta available at time t, and $X_{n,t}$ is a vector of control variables. We are careful that betas and other yield curve-based controls are known at the time when forecasts are made. For concreteness, consider the March observation

[19] In our specification, we use the maturity-specific term spread: for the n-year expected excess return, we use the n-year term spread.

of $\tilde{E}_t xr_{n,t\to t+12}$. This measure is constructed using the March Blue Chip survey, which forecasters submit at the end of February. Hence, we use bond betas and other predictors at the end of February.

We estimate the regression in a panel dataset, where the unit of observation is bond maturity $n \in \{3, 6, 11\}$ by month t. Betas are estimated over 5-year rolling backward-looking windows. Before turning to the regression results, Figure 2 visualizes the data by pooling observations across maturities and plotting expected excess returns against the corresponding bond–stock betas. Each maturity is distinguished by a different marker. The figure shows a clear positive relationship, both over time and across maturities. The relationship also holds across different sample periods, which are distinguished by different colors, and is not driven by specific bond maturities or outliers.

Table II shows the results of estimating Equation (11). Throughout the paper, we use Driscoll–Kraay standard errors for panel regressions and Newey–West standard errors for time-series regressions. We use 27 lags, following the truncation rule suggested by Lazarus et al. (2018), and evaluate statistical significance using fixed-b critical values (Kiefer and Vogelsang, 2002, 2005).

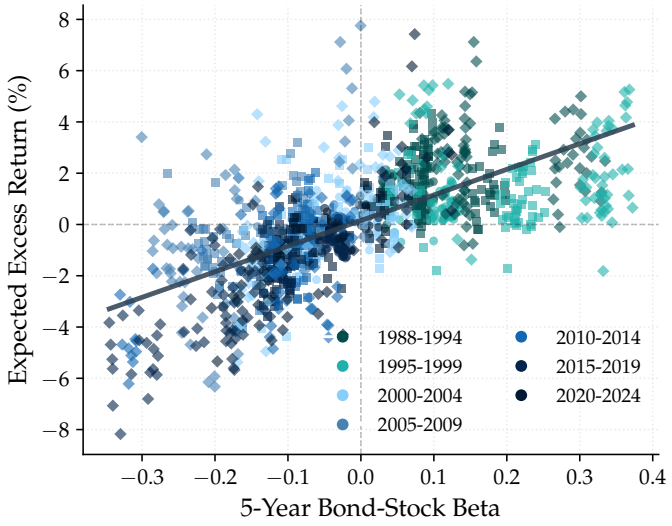
Column (1) in Table II shows that a unit increase in beta is associated with an increase in expected excess returns of 10 percentage points per year. The R^2 of the regression is high at 44%, corresponding to a correlation coefficient of 67%. As a simple back-of-the-envelope calculation, the peak-to-trough decline in the beta of the 11-year bond from roughly 0.37 to -0.34 can explain a decline in the expected bond excess return from $0.37 \times 10\% = 3.72\%$ to $-0.34 \times 10\% = -3.45\%$. For comparison, the moving average of expected excess returns in Figure 1 Panel A declined from 4.67% to -5.53%.

We next show that the relationship between betas and expected excess returns is economically and statistically unchanged when we include a host of controls. Column (2) includes maturity fixed effects to allow for the possibility that some bond maturities may persistently have higher or lower expected excess returns, for example, because of differences in liquidity. Column (3) adds maturity-specific time trends, constructed by interacting a linear time trend with maturity indicators, in addition to maturity fixed effects. Thus, our result is not driven simply by downward drifts over our sample in expected excess returns and betas.

Column (4) adds the variance of bond returns, estimated as the annualized variance of monthly bond excess returns over a rolling 5-year window.[20] The estimated coefficient on the beta remains similar, and the R^2 increases only slightly. The coefficient on bond return variance is positive but not statistically significant.

[20] We multiply the variance by 100 to make its magnitude more easily comparable to the bond–stock beta. The variance ranges from 0.02 to 1.28.

Figure 2: Bond Expected Excess Returns and Bond–Stock Betas



Note: The figure plots expected excess returns against their corresponding bond–stock betas. 1-year expected excess returns are computed according to Equation (7) using 4-quarter-ahead bond yield forecasts from the Blue Chip Financial Forecasts. Betas are computed by regressing monthly par bond excess returns on stock market excess returns over a backward-looking rolling window of 60 months, according to Equation (10). The figure pools observations across the different maturities. Maturities are distinguished by marker shape: three-year bonds (circles), six-year bonds (squares), and eleven-year bonds (diamonds). Sample periods are distinguished by color. The sample spans January 1988 to March 2024.

Columns (5) and (6) add controls for known predictors of realized bond returns. Column (5) adds the term spread, defined as the difference between the bond yield at the relevant maturity (3, 6, or 11 years) and the 1-year bond yield. Column (6) controls for the Cochrane–Piazzesi factor (CP). Column (7) controls for the first three principal components of the yield curve: level, slope, and curvature. Column (8) controls for the Laubach-Williams estimate of the natural rate of interest (r-star). Throughout the table, the R^2 only increases modestly from Column (1) to Column (8), emphasizing the substantial explanatory power of betas for expected excess returns.

Column (9) flexibly controls for both time-series patterns and maturity-specific characteristics by including time and maturity fixed effects. The estimated coefficient remains statistically significant and similar in magnitude to those reported in the other columns. Thus, the relationship between bond–stock betas and expected excess returns holds both over time and across maturities.

It is well known that there is significant common variation in bond yields across maturities. This implies that there is limited independent cross-sectional variation and most of the identification in Table II comes from the time series. To show that our results do not depend on the panel structure of the data in Table II, we therefore estimate the regression

Table II: Expected Excess Returns and Stock Market Betas

Dependent Variable: Expected Excess Returns									
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Bond-Stock Beta 5YR	9.99***	9.98***	8.00***	9.86***	10.13***	9.06***	7.75***	9.20***	12.11***
	(1.62)	(1.61)	(2.06)	(1.43)	(1.63)	(1.81)	(1.56)	(1.50)	(1.53)
Bond Variance				1.79					
				(1.16)					
Term Spread					0.27				
					(0.17)				
CP Factor						0.12			
						(0.09)			
Level							0.54***		
							(0.17)		
Slope							-0.23		
							(0.13)		
Curvature							0.20		
							(0.14)		
r-star								0.31	
								(0.17)	
Constant	0.16								
	(0.20)								
Maturity FE		✓	✓	✓	✓	✓	✓	✓	✓
Time Trend × Maturity FE			✓						
Time FE									✓
R-squared	0.44	0.44	0.47	0.46	0.46	0.45	0.50	0.46	0.87
Observations	1305	1305	1305	1305	1305	1305	1305	1305	1305

Note: The table reports estimates from regressions of expected excess returns on bond–stock betas and controls, as in Equation (11):

$$\tilde{E}_t xr_{n,t\to t+12} = \alpha + \lambda\beta_{n,t} + X'_{n,t}\delta + \eta_{n,t}.$$

The dependent variable is the 1-year expected excess bond return. The bond–stock beta $\beta_{n,t}$ is estimated from monthly bond and stock excess returns over a 5-year rolling window. Column (1) is the baseline; Columns (2)–(3) add maturity fixed effects and maturity-specific time trends; Columns (4)-(7) add the bond return variance (multiplied by 100), the term spread, the Cochrane–Piazzesi factor and the yield-curve level, slope, and curvature; Column (8) controls for Laubach-Williams estimates of the natural rate of interest (r-star), and Column (9) includes maturity and time fixed effects. Driscoll–Kraay standard errors with 27 lags are reported in parentheses. Stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample consists of monthly data from January 1988 to March 2024.

separately for each maturity in Table III. The relationship between the bond–stock beta and expected excess returns is both statistically and economically significant for every maturity. The estimated loadings on the bond–stock beta are similar across maturities: 7.96 for the 3-year Treasury (Column (1)), 8.62 for the 6-year (Column (4)), and 10.79 for the 11-year (Column (7)). The R^2 is 41%, 39%, and 47%, respectively, so the bond–stock beta explains a somewhat larger fraction of the variation in expected excess returns for the 11-year bond.

Robustness We present a battery of robustness checks in the Appendix. Table B2 replicates all the specifications in Table II using the noisier but faster-updating 1-year rolling beta. We find that the 1-year rolling beta has similarly high explanatory power for expected

Table III: Expected Excess Returns and Stock Market Betas: Individual Maturities

	3-year			6-year			11-year		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Beta 5YR	7.96***	6.19***	5.17***	8.62***	6.82***	5.59***	10.79***	8.78***	6.85***
	(1.11)	(1.34)	(1.11)	(1.56)	(1.98)	(1.78)	(1.70)	(2.30)	(1.86)
Level			0.33***			0.61**			1.11***
			(0.11)			(0.24)			(0.34)
Slope			-0.06			-0.25*			-0.36
			(0.06)			(0.14)			(0.23)
Curvature			0.11			0.22			0.18
			(0.08)			(0.15)			(0.25)
Constant	0.18*	0.52**	0.15*	0.12	0.70	0.07	0.14	1.12	0.07
	(0.10)	(0.21)	(0.08)	(0.20)	(0.48)	(0.16)	(0.33)	(0.77)	(0.29)
Time Trend		✓			✓			✓	
R-squared	0.41	0.45	0.54	0.39	0.41	0.49	0.47	0.49	0.55
Observations	435	435	435	435	435	435	435	435	435

Note: The table reports the regression estimates for equation (11) estimated maturity by maturity:

$$\tilde{E}_t xr_{n,t\to t+12} = \alpha + \lambda \hat{\beta}^w_{n,t} + X'_{n,t}\delta + \eta_{n,t}.$$

The bond–stock beta $\hat{\beta}^w_{n,t}$ is estimated from monthly bond and stock excess returns over a 5-year rolling window. Columns (2), (5), and (8) include a time trend; columns (3), (6), and (9) control for the first three principal components of the yield curve (level, slope, and curvature), each standardized to have mean zero and unit standard deviation. Newey–West standard errors with 27 lags are reported in parentheses. Stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample consists of monthly data from January 1988 to March 2024.

bond excess returns, though the coefficients and R^2s are somewhat attenuated. This is what one would expect if markets use a version of bond–stock betas with roughly 5-year rolling windows, as suggested by the analysis of Couts et al. (2026). The table also adds a column where we use the "forward" one-year beta estimated at time $t+12$ rather than at time t, i.e., using information that is contemporaneous with the horizon over which expected excess returns are constructed, but not available to investors at time t. In this case, we find a slightly smaller loading and a lower R^2 than for the backward-looking one-year beta used in the other columns, but our results remain robust. Appendix Table B3 shows robustness to the other alternative measures of bond–stock betas described in Section 3.1. The results are virtually the same for any specification. Table B4 reports the maturity-by-maturity regressions using one-year rolling betas, showing similar results to Table III.

3.2.1 Expected Return Decomposition

We have documented a strong empirical relationship between expected excess returns and bond–stock betas in U.S. data. To understand the source of this relationship, we next decompose the expected excess return in Equation (7) into its main components. We test whether betas are priced in market yields, in forecasted yields, or both. This decomposition speaks to Corollary 1 in Section 2, where we saw that expected excess returns should be related to bond–stock betas as long as betas are priced by market yields (Assumption 1)

and forecasts load equally or less onto bond–stock betas (Assumption 2). In this section, we show that these two conditions hold.

Define as $\widetilde{i}_t^n$ forecasters' expected average Treasury bill rate over the next n years:

$$\widetilde{i}_t^n = \widetilde{E}_t \left[\frac{1}{n} \sum_{j=0}^{n-1} i_{t+12j} \right],$$

where we maintain our convention that t is measured in months. We can then rewrite the expected excess return in Equation (7) as

$$\widetilde{E}_t xr_{n,t\to t+12} = dur_{n,t} \left(y_{n,t} - \widetilde{i}_t^n \right) - (dur_{n,t} - 1) \left(\widetilde{E}_t y_{n-1,t+12} - \widetilde{i}_t^n \right) + \left(\widetilde{i}_t^n - y_{1,t} \right). \quad (12)$$

Expressing yields relative to the expected Treasury bill rate offers two main advantages. First, it removes the common low-frequency component in interest rates without using forecasters' expectations of long-term yields and hence term premia. Second, as Equation (6) illustrates, the difference between the long-term yield and the expected average Treasury bill rate provides a proxy for the term premium – the average expected excess return over the life of the bond. This exercise therefore complements the results in Table II, which focus on the one-year expected excess return.[21]

We measure $\widetilde{i}_t^{10}$ using data from the SPF, which reports forecasts of the average Treasury bill rate over the next 10 years. It is natural to match these forecasts to the 10-year yield forecast, and hence to the expected return on the 11-year bond. The SPF 10-year forecasts are only available annually, which means the sample is smaller than in our baseline results, with 33 observations for the 11-year bond and 99 observations for the U.S. panel.[22]

We regress the first two terms on the right-hand side of Equation (12) onto bond–stock betas. Table IV Panel (A) reports results only using the 11-year bond because this maturity maps most cleanly into the long-horizon Treasury bill rate forecast reported by the SPF. In column (1), the dependent variable is the spread between the current 11-year bond yield and the expected average Treasury bill rate over the next ten years. This variable, a proxy for the expected excess return over the life of the bond, is also plotted in Figure 1 Panel B. The regression coefficient is equal to 3.60 and is statistically significant. Hence, actual market yields load positively onto bond–stock betas, as predicted by the model if betas move

[21] Molavi et al. (2025) argue that the spread between long-term bond yields and the expected short rate over the lifetime of the bond reflects errors in linking forecasts for yields of different maturities. We show that this spread is more strongly related to betas than the spread between the forecasted future long-term yield and the forecasted Treasury bill rate.

[22] This exercise combines forecasts from Blue Chip and the SPF. This is reasonable because forecasts that appear in both surveys are highly correlated.

term premia. Column (2) shows that a similar relationship holds when controlling for a time trend. This is consistent with the idea that markets price variation in bond–stock betas into yields.

Columns (3) and (4) find a positive but weaker relationship for forecasted yields, as predicted by the model. The forecasted yield minus the expected average Treasury bill rate loads positively onto bond–stock betas, but the loading is smaller in magnitude than the loading for actual market yields and becomes insignificant when we control for a time trend.

Columns (5) and (6) show that market yields have a significantly more positive relationship with bond–stock betas than yield forecasts. This holds with or without controlling for a time trend. Note also that the difference in Columns (5) and (6) does not use the SPF data on long-horizon Treasury bill rate forecasts, so the difference in loadings is not an artifact of that data. This difference in the sensitivity of market yields and yield forecasts to bond–stock betas directly translates into the relationship we find for expected excess returns.[23]

Panel (B) shows that the decomposition results for the 11-year maturity carry through to the full panel. The SPF only reports the average expected Treasury bill rate over the next 10 years, and we therefore use this long-horizon forecast for all bond maturities in Panel (B). The results show that the same patterns hold in the panel regression.

Overall, the decomposition of expected excess returns shows that the relationship between bond–stock betas and expected excess returns is driven by actual market yields, consistent with Corollary 1. Thus, our results primarily capture how markets price beta risk, not simply the methodologies forecasters use to predict future yields.

In Appendix B.7, we conduct a different decomposition to study whether convenience yields in U.S. Treasuries contribute to our results. We construct a measure of expected excess returns on Aaa corporate bonds, which should have limited convenience benefits.[24] Expected excess returns on Aaa corporate bonds are also highly correlated with the bond–stock beta of U.S. Treasuries, suggesting that convenience yields do not play a key role in explaining our results.

3.2.2 Relation to Inflation Uncertainty and Inflation Expectations

In this section, we analyze the impact of inflation on the relation between bond–stock betas and expected excess returns. We use inflation-related information in two ways. First, we

[23]Because expected excess returns are a duration-weighted difference between the market yield and the yield forecast, the difference in their loadings on bond–stock betas is further amplified for expected excess returns. See Appendix Table B5 for a decomposition of the duration-weighted yields and forecasts.

[24]Since Aaa bonds are virtually default-free, but offer holders fewer benefits in terms of liquidity, regulation, and collateralization than Treasuries, the yield spread between Aaa corporate bonds and Treasuries is often used as a measure of convenience yields (Krishnamurthy and Vissing-Jorgensen, 2012).

Table IV: Expected Excess Return Decomposition and Bond–Stock Betas

Panel (A): 11-Year Yield

Dependent Variable:	$y - \tilde{E}i^{10y}$		$\tilde{E}y - \tilde{E}i^{10y}$		$y - \tilde{E}y$	
	(1)	(2)	(3)	(4)	(5)	(6)
Beta 5YR	3.60***	2.56**	2.19**	0.89	1.37***	1.36***
	(0.55)	(0.86)	(0.60)	(0.65)	(0.24)	(0.38)
Time trend		✓		✓		✓
R-squared	0.63	0.71	0.34	0.53	0.45	0.45
Observations	33	33	33	33	435	435

Panel (B): Panel of All Maturities

Dependent Variable:	$y - \tilde{E}i^{10y}$		$\tilde{E}y - \tilde{E}i^{10y}$		$y - \tilde{E}y$	
	(1)	(2)	(3)	(4)	(5)	(6)
Beta 5YR	4.91***	4.01**	3.24***	1.94*	1.70***	1.77***
	(0.61)	(1.15)	(0.69)	(0.88)	(0.30)	(0.48)
Time trend × Maturity FE		✓		✓		✓
Maturity FE	✓	✓	✓	✓	✓	✓
R-squared	0.50	0.55	0.40	0.49	0.32	0.34
Observations	99	99	99	99	1305	1305

Note: The table reports the results of the regression

$$Y_{n,t} = \alpha_n + \lambda\, \beta_{n,t} + X'_{n,t}\delta + \varepsilon_{n,t},$$

for $Y_{n,t} \in \{\, y_{n,t} - \tilde{i}_t^{10},\ \tilde{E}_t y_{n-1,t+12} - \tilde{i}_t^{10},\ y_{n,t} - \tilde{E}_t y_{n-1,t+12} \,\}$, where $y_{n,t}$ is the observed yield on the n-year bond, $\tilde{E}_t y_{n-1,t+12}$ is the survey forecast of the $(n-1)$-year yield one year ahead, $\tilde{i}_t^{10}$ is the SPF forecast of the average Treasury bill rate over the next ten years, $\beta_{n,t}$ is the five-year rolling bond–stock beta, $X_{n,t}$ is a vector of control variables, and α_n are maturity fixed effects (a single intercept in Panel (A), which uses one maturity). The dependent variable $Y_{n,t}$ is the observed-yield spread $y_{n,t} - \tilde{i}_t^{10}$ in columns 1–2, the forecasted-yield spread $\tilde{E}_t y_{n-1,t+12} - \tilde{i}_t^{10}$ in columns 3–4, and the gap between observed and forecasted yields $y_{n,t} - \tilde{E}_t y_{n-1,t+12}$ in Columns 5–6. Panel (A) uses the 11-year bond alone ($n = 11$) and includes a single intercept, with Newey–West standard errors. Panel (B) pools the 3-, 6-, and 11-year bonds and includes maturity fixed effects, with Driscoll–Kraay standard errors. Even-numbered columns include maturity-specific time trends. Because $\tilde{i}_t^{10}$ is observed annually, columns 1–4 are estimated at the annual frequency and use standard errors with 7 lags, while columns 5–6 use the full monthly sample and standard errors with 27 lags. Standard errors are reported in parentheses; stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample runs from January 1988 to March 2024; Columns 1–4, which use the SPF forecast $\tilde{i}_t^{10}$, start in 1992, when the SPF ten-year Treasury bill rate forecast becomes available.

show that our results are robust to controlling for the level and volatility of inflation, alleviating concerns that they might simply reflect views on the direction of inflation. Second, inflation forecasts provide another way to show that the relationship between bond–stock betas and expected excess returns depends on whether the market prices betas, not on whether forecasters do.

Columns (2) and (3) in Table V show that inflation volatility, measured as the standard deviation of CPI inflation over a 5-year rolling window, is not significantly related to expected bond excess returns. Columns (4) and (5) show that the estimated loading on beta is

unchanged when controlling for inflation volatility.

We next relate expected excess returns to long-term inflation expectations. We use expected inflation over the next 10 years from the SPF.[25] In column (6), we find a positive relationship between expected inflation and expected excess returns: a one percentage point increase in expected inflation is associated with a 2.07 percentage point increase in expected excess returns. The R^2 is 18%, which is substantially lower than the baseline explanatory power of bond–stock betas of 44%, reported in column (1). Columns (8)-(9) show that the loading on expected inflation is driven out by betas whether or not we interact expected inflation with maturity fixed effects. The loading on beta is similar to the baseline and remains highly statistically significant. This is what we would expect if bond–stock betas exhibit some joint movement with inflation expectations over our sample, but markets price bond–stock betas.

We next use inflation forecasts to construct an alternative proxy for expected excess returns. This proxy is immune to concerns that we are merely capturing the term premium implications of a particular forecasting methodology because inflation forecasts should not reflect term premium considerations. However, since it takes market yields as an input, this proxy should still correlate with bond–stock betas if markets price beta.

We construct inflation-based pseudo-expected excess returns in two steps. First, for each maturity, we regress the Blue Chip forecast of the future bond yield on the SPF's forecast of 10-year expected inflation. We take the fitted value as a yield forecast driven purely by long-run inflation expectations, $\widetilde{E}_t^{\pi} y_{n-1,t+12}$.[26] Second, we substitute this forecast for the survey yield forecast in Equation (7) to obtain the pseudo-expected excess return $\widetilde{E}_t^{\pi} xr_{n,t\to t+12}$. See Appendix B.2.1 for details.

We then regress the pseudo-expected excess returns on bond–stock betas in column (10) of Table V. The coefficient in column (10), 18.95, is positive and substantially larger than the baseline estimate of 9.98. This is what we would expect based on Corollary 1. A positive relationship between bond–stock betas and expected excess returns is predicted as long as markets price betas, and this relationship is amplified if forecasts put low weight on term premia.

[25] These expectations can also be thought of as a measure of perceived shifting end-points (Kozicki and Tinsley, 2001).

[26] This procedure captures the spirit of so-called fair value models used by some forecasters. For example, Goldman Sachs reported using its "Sudoku" model: "Sudoku estimates the fair value of 10-year yields based on macroeconomic factors (short rates, activity growth, inflation) ..." (Goldman Sachs Global Viewpoint, "Sudoku Gets Bigger and Forward-Looking!", June 2007). Similarly, J.P. Morgan reported: "Our fair value model [...] values 10y real yield as a function of the real Fed funds rate, inflation volatility ..." (J.P. Morgan, "Flows & Liquidity: An Update of Our Fair Value Models", June 2022).

Table V: Expected Excess Returns, Bond–Stock Betas, and Inflation

Dependent Variable:	Expected Excess Returns									XR CPI 10Y
Control Variable:		CPI Std				CPI 10Y				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Beta 5YR	9.98***			10.08***	10.08***			8.48***	8.19***	18.95*
	(1.61)			(1.55)	(1.55)			(1.82)	(2.02)	(8.32)
Control		0.04		0.28		2.07***		0.29		
		(0.45)		(0.23)		(0.34)		(0.53)		
Control × Mat. 3			0.01		0.24**		0.98***		0.20	
			(0.20)		(0.11)		(0.11)		(0.29)	
Control × Mat. 6			0.11		0.38		1.80***		0.16	
			(0.39)		(0.24)		(0.29)		(0.64)	
Control × Mat. 11			-0.01		0.22		3.42***		0.70	
			(0.80)		(0.42)		(0.69)		(0.88)	
Maturity FE	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
R-squared	0.44	0.00	0.00	0.45	0.45	0.18	0.22	0.39	0.40	0.16
Observations	1305	1305	1305	1305	1305	390	390	390	390	390

Note: The table reports estimates from regressions of expected excess returns on bond–stock betas and inflation-related controls, as in Equation (11):

$$\tilde{E}_t xr_{n,t\to t+12} = \alpha_n + \lambda\beta_{n,t} + X'_{n,t}\delta + \eta_{n,t},$$

where $\beta_{n,t}$ denotes the 5-year rolling beta, $X_{n,t}$ is a vector of inflation-related controls, and α_n are maturity fixed effects. Column (1) reports the baseline with the beta alone. Columns (2)–(5) use the 5-year rolling standard deviation of CPI inflation (CPI Std), and columns (6)–(9) long-run expected inflation from the SPF (CPI 10Y), the expected average CPI inflation over the next ten years. Within each block the control enters first without the beta and then together with it; in the even columns it enters with a single common loading ("Control" row), in the odd columns with a separate slope per maturity ("Control × Mat." rows). Column (10) instead estimates

$$\widetilde{E}^{\pi}_t xr_{n,t\to t+12} = \alpha_n + \lambda\beta_{n,t} + \eta_{n,t},$$

where the dependent variable is the inflation-based pseudo-expected excess return (XR CPI 10Y), which replaces the survey yield forecast in Equation (7) with one implied by long-run inflation expectations (see Appendix B.2.1). Columns (1)–(5) are monthly, January 1988 to March 2024; Columns (6)–(10) are quarterly, since the SPF forecasts are observed once per quarter, and run from 1991:Q4, when the SPF 10-year inflation forecast starts, to 2024:Q1. Driscoll–Kraay standard errors (27 lags for the monthly sample, 14 lags for the quarterly sample) are in parentheses; stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values.

3.2.3 Controlling for Standard Forecaster Biases

We next show that standard biases in yield forecasts do not explain our baseline result. The literature has documented substantial biases in how markets and forecasters form expectations about short- and long-term interest rates (Kozicki and Tinsley, 2001; Cieslak, 2018; Farmer et al., 2024; Shue et al., 2025). As Proposition 1(b) shows, if forecasters and markets have similar biases, expected excess returns difference out these biases, leaving behind the relationship between betas and term premia. However, if forecasters and markets have different biases and that difference is correlated with betas, it could affect our results. We show

that the relationship between bond–stock betas and expected excess returns is unaffected if we control for the most common forecaster biases, suggesting this is not the case.

We start by controlling for the forecast bias documented by Shue et al. (2025): forecasted changes in long-term bond yields overreact to forecasted changes in short-term interest rates. In Column (2) of Table VI, we regress expected excess returns on bond–stock betas and the expected change in the federal funds rate. We construct the expected change in the federal funds rate over the next 4 quarters using Blue Chip forecasts. This matches our construction of expected excess returns, which uses 4-quarter-ahead forecasts of bond yields. The coefficient on the expected change in the federal funds rate is negative and statistically different from zero, but adds little additional explanatory power over the baseline in column (1). However, controlling for the expected change in the federal funds rate has little effect on the coefficient on bond–stock betas, which remains close to the baseline estimate reported in column (1). Column (3) repeats this specification, but interacts the expected change in the federal funds rate with maturity fixed effects, to account for the possibility that forecast biases matter more for longer-duration bonds. The coefficient on the bond–stock beta remains unchanged.

Finally, we consider the possibility that forecasters might expect bond yields to mean-revert to their recent levels, as argued by Cieslak (2018) and Farmer et al. (2024) (FNS).[27] To capture this possibility, we form pseudo-expected excess returns where we replace BCFF yield forecasts with forecasts mechanically based on past yields. If our results primarily reflect a bias towards mean-reversion in yields, the pseudo-expected excess return should drive out bond–stock betas in our main regression. Columns (4)–(9) show that this is not the case, mitigating these concerns. Columns (4) and (5) use a simple rolling AR(1) model for yields, while Columns (6) and (7) use a moving average of yields over the past 60 months. We control for the expected excess return implied by the slow-learning model of Farmer et al. (2024) (FNS) in columns (8) and (9).[28] We find that expected excess returns do not load onto FNS expected excess returns and continue to load onto betas, suggesting our results are not driven by slow learning or serially correlated forecast errors. We will next see that this is different for realized excess returns.

[27] Appendix B.2.2 describes the construction of these forecasts in detail.

[28] The FNS expected excess return is what an econometrician with more knowledge of factor dynamics than the market would predict. In FNS, the expectations hypothesis holds under the market's expectations, i.e., perceived bond risk premia are constant. However, they vary from the perspective of an outside econometric forecaster with knowledge of factor dynamics that the market is trying to learn. For instance, this forecaster would understand that a decline in the permanent component of yields would lead the market to make serially correlated errors, leading to predictably positive excess returns. Thus, in our model, the FNS expected excess return most closely resembles the loading of actual excess returns onto the term structure factor f_t due to biased expectations in Proposition 1(a). As we saw in Corollary 2, the loading onto the term structure factor should be differenced out from expected excess returns.

Table VI: Expected Excess Returns and Simple Yield Forecasts

Dependent Variable: Expected Excess Returns									
Control Variable:		$\tilde{E}_t[\Delta FF]$		AR Roll		MA 60		FNS	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Beta 5YR	9.98***	9.32***	9.14***	10.34***	10.32***	10.14***	10.20***	10.74***	10.77***
	(1.61)	(1.43)	(1.40)	(1.65)	(1.64)	(1.64)	(1.63)	(1.89)	(1.90)
Control		-0.60***		0.14		0.07***		0.02	
		(0.15)		(0.12)		(0.02)		(0.19)	
Control × Mat. 3			-0.19*		0.02		-0.01		-0.18
			(0.09)		(0.06)		(0.03)		(0.12)
Control × Mat. 6			-0.57***		0.09		0.03		0.03
			(0.16)		(0.12)		(0.02)		(0.21)
Control × Mat. 11			-1.07***		0.21		0.10***		0.19
			(0.26)		(0.15)		(0.03)		(0.30)
Maturity FE	✓	✓	✓	✓	✓	✓	✓	✓	✓
R-squared	0.44	0.48	0.49	0.45	0.45	0.46	0.47	0.49	0.49
Observations	1305	1305	1305	1305	1305	1305	1305	435	435

Note: The table reports estimates from regressions of expected excess returns on bond–stock betas and proxies for behavioral biases, as in Equation (11):

$$\tilde{E}_t xr_{n,t\to t+12} = \alpha_n + \lambda\beta_{n,t} + X'_{n,t}\delta + \eta_{n,t},$$

where $\beta_{n,t}$ denotes the 5-year rolling beta available at time t, $X_{n,t}$ is the control variable of the column, and α_n are maturity fixed effects, included in every column. In each pair of columns, the control enters first with a single (common) loading (the "Control" row) and then with maturity-specific loadings and no common term (the "Control × Mat. 3/6/11" rows). Column (1) is the baseline specification with the bond–stock beta alone. Columns (2)–(3) control for the Blue Chip forecast of the change in the federal funds rate over the next four quarters, $\tilde{E}_t[\Delta FF]$. Columns (4)–(9) control for pseudo-expected excess returns, constructed by replacing the survey yield forecast in our baseline measure with a purely statistical forecast of future bond yields: a rolling AR(1) forecast estimated over a 15-year window (AR, columns (4)–(5)), a 5-year (60-month) trailing moving-average forecast (MA, columns (6)–(7)), and the FNS forecast (FNS, columns (8)–(9)). The construction of each pseudo-expected excess return is detailed in Appendix B.2.2. Columns (1)–(7) use monthly data, while columns (8)–(9) use quarterly data, since the FNS measure is observed once per quarter. Driscoll–Kraay standard errors are reported in parentheses, using 27 lags for the monthly specifications and 15 lags for the quarterly specifications. Stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample period is January 1988 to March 2024.

3.3 U.S. Realized Excess Returns and Bond–Stock Betas

In the previous section, we showed that bond–stock betas are strongly related to expected excess returns. Under the full information rational expectations (FIRE) benchmark, the difference between realized and expected excess returns should be white noise, and betas should similarly predict realized bond excess returns in a sufficiently large sample. In practice, however, interest rate surprises relative to forecasters' expectations may be substantial in a sample with a persistent decline in long-term bond yields, such as ours. Such surprises could be due to either behavioral biases (Cieslak, 2018; Bordalo et al., 2020; Shue et al., 2025; Molavi et al., 2025) or the difficulty of learning about the structure of the economy

(Farmer et al., 2024; Molavi et al., 2024). We first show that the raw relationship between realized bond excess returns and bond–stock betas is weak. We then show that the relationship strengthens when we control for forecaster surprises about the path of monetary policy and the macroeconomy.

We estimate predictive regressions of the form:

$$xr^{par}_{n,t\to t+12} = \alpha + \gamma\,\beta_{n,t} + X'_{n,t}\delta + \varepsilon_{n,t+12}. \tag{13}$$

Here, $xr^{par}_{n,t\to t+12}$ denotes the realized excess return on a bond with maturity n between t and $t+12$, while $\beta_{n,t}$ denotes the bond–stock beta estimated using information available through time t. Under FIRE, we would expect the estimated coefficient γ to be similar in magnitude to the λ estimated in Table II in a sufficiently large sample.

Column (1) of Table VII shows a weak and statistically insignificant relationship between betas and future realized excess returns. The raw relationship between bond–stock betas and future realized excess returns is not significantly different from zero, in sharp contrast to what we find when using expected excess returns as the dependent variable in column (2) of Table II.[29]

The remaining columns of the table drill down into the difference between expected and realized bond excess returns based on the identity below, which follows directly from the definitions of realized and expected bond excess returns:

$$xr^{par}_{n,t\to t+12} = \tilde{E}_t xr_{n,t\to t+12} - (dur_{n,t} - 1)\underbrace{\left(y^{par}_{n-1,t+12} - \tilde{E}_t y^{par}_{n-1,t+12}\right)}_{\text{yield surprise}}. \tag{14}$$

The realized bond excess return equals the expected bond excess return minus the yield surprise, scaled by the bond's duration at the end of the holding period (i.e., $dur_{n,t} - 1$). An immediate implication is that controlling for the contemporaneous yield surprise (scaled by duration) should recover the strong, positive relationship between excess returns and bond–stock betas.

This is exactly what we see in column (2) of Table VII. Controlling for the scaled yield surprise, the coefficient on bond–stock betas becomes positive, strongly statistically significant, and almost identical to the comparable result in Table II column (2). The scaled yield surprise enters with a coefficient of almost exactly minus one, as implied by Equation (14). The R^2 is 0.91. Given the results in Table II and the identity in Equation (14), this is not

[29] Given this contrast, it is natural to ask whether expected excess returns forecast realized excess returns. In Appendix Table B7, we show that they do, which implies that the component of expected excess returns that is orthogonal to bond–stock betas forecasts realized returns in our sample.

surprising. Controlling for the surprises (i.e., forecast errors) about long-term bond yields recovers the baseline relationship between bond–stock betas and expected excess returns. We next turn to understanding the drivers of yield surprises.

We find that surprises about the path of short-term rates and the macroeconomy can recover the relationship between realized excess returns and bond–stock betas.[30] This is visible in columns (3) through (6), where we first control for surprises in the Treasury bill rate, GDP, and CPI individually and then jointly.[31] For consistency, we scale these by the same $(dur_{n,t} - 1)$ factor. Controlling for the T-bill surprise alone recovers a strongly positive and statistically significant bond–stock beta coefficient of 4.62. Moreover, the R^2 rises to 50%, substantially above the 2% obtained in column (1). Accounting only for surprises in the monetary policy path can explain about half the variation in realized bond excess returns, and recovers a strong relationship between bond–stock betas and future realized bond excess returns. However, the fact that the R^2 is lower in column (3) than column (2) indicates that long-term yield surprises are not fully explained by short-rate surprises.

Next, columns (4) through (6) show that CPI and GDP growth surprises capture further variation in long-term yield surprises. Columns (4) and (5) show that GDP surprises on their own also help recover the realized return–beta relationship, but CPI surprises do not, suggesting that not all macroeconomic news equally influences long-term yield surprises. Column (6) shows that, once we include all macroeconomic and short-term interest rate surprises jointly, we recover a positive and statistically significant relationship between bond–stock betas and future realized excess returns. Since macroeconomic forecasters predict short-term interest rates, inflation, and GDP growth without regard to term premia, these surprises are the natural empirical analog of ex-post forecast errors about term structure factors in the model. Consistent with the prediction of Proposition 1(a), we recover the empirical relationship between realized excess returns and bond–stock betas once we control for the analog of forecast errors about f_t.[32]

Column (7) adds innovations in the bond–stock beta to the specification in column (6).[33]

[30] Most of the results in the paper are invariant to the construction of the bond–stock beta, but the results in columns (3)–(6) are somewhat sensitive. If we use other constructions, the coefficients are generally similar, but the statistical significance fluctuates. The finding in column (2) that controlling for long-term yield surprises recovers the beta-realized return relationship is strongly robust. But short rate and macroeconomic surprises may not always explain enough of the long-term yield surprise to recover the relationship.

[31] The Treasury bill surprise is defined as the difference between the realized Treasury bill rate and the rate forecasted 12 months prior, which is closely related to monetary policy forecast errors and contains only negligible term premia. The GDP and CPI surprises are computed as the differences between the actual GDP growth and CPI inflation and their respective BCFF forecasts.

[32] Bianchi et al. (2022b) show that a machine learning approach can uncover real-time ex-ante estimates for ex-post professional forecast errors.

[33] Innovations in beta are defined as $\beta_{n,t+12} - \hat{\rho}_n \, \beta_{n,t}$, where $\hat{\rho}_n$ is the in-sample 12-month autocorrelation of the bond–stock beta for maturity n, estimated by regressing $\beta_{n,t+12}$ on $\beta_{n,t}$ separately for each maturity.

Table VII: Realized Excess Returns and Bond–Stock Betas

Dependent Variable: Realized Excess Return									
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Beta	1.25	9.89***	4.62**	6.93**	-0.76	4.43**	4.54***	5.13**	1.20
	(3.25)	(1.68)	(1.86)	(3.00)	(2.71)	(1.58)	(1.54)	(1.82)	(2.17)
Yield Surprise		-0.99***							
		(0.02)							
T-bill Surprise			-0.52***			-0.43***	-0.41***	-0.42***	-0.42***
			(0.07)			(0.06)	(0.06)	(0.05)	(0.06)
GDP Surprise				-0.19***		-0.04	-0.05*	-0.04	-0.04
				(0.04)		(0.03)	(0.03)	(0.03)	(0.03)
CPI Surprise					-0.27***	-0.13***	-0.13***	-0.11***	-0.12***
					(0.05)	(0.03)	(0.02)	(0.04)	(0.03)
Innovations in Beta							-7.91		
							(6.56)		
Term Spread								1.27***	
								(0.39)	
CP Factor									0.39*
									(0.19)
Maturity FE	✓	✓	✓	✓	✓	✓	✓	✓	✓
R-squared	0.02	0.91	0.50	0.14	0.22	0.54	0.55	0.59	0.56
Observations	1305	1305	1305	1305	1305	1305	1305	1305	1305

Note: The table reports the estimates of Equation (13):

$$xr^{par}_{n,t\to t+12} = \alpha + \gamma\,\beta_{n,t} + X'_{n,t}\delta + \varepsilon_{n,t+12}.$$

Here, $xr^{par}_{n,t\to t+12}$ denotes the realized excess return on a bond with maturity n between t and $t+12$, while $\beta_{n,t}$ denotes the bond–stock beta estimated using information available through time t. Column (2) controls for the yield surprise; columns (3)–(5) control for the T-bill, GDP, and CPI surprises one at a time; and column (6) for all three jointly. Surprises are realized values at $t+12$ minus BCFF forecasts made at t, scaled by $(dur_{n,t}-1)$. Columns (7)–(9) each add one control to column (6): innovations in beta, the term spread, and the Cochrane–Piazzesi factor. The term spread and the Cochrane–Piazzesi factor are measured at t. Driscoll–Kraay standard errors with 27 lags are reported in parentheses. Stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample consists of monthly data with forecast dates t from January 1988 to March 2024, so that realized returns run through March 2025.

According to Proposition 1(a), innovations in beta may also affect the empirical relationship between realized excess returns and bond–stock betas. However, this is not the case in the data, as the coefficient on the bond–stock beta remains broadly unchanged in column (7).

The remainder of Table VII provides evidence on standard predictors of realized bond excess returns. Proposition 1(a) suggests these standard predictors can forecast bond excess returns either because they are correlated with ex-post surprises or because they are correlated with the compensation for beta risk that investors require. Column (8) adds the well-known term spread variable of Campbell and Shiller (1991) and Fama and Bliss (1987), finding that it enters positively and significantly. However, it leaves the coefficients on the bond–stock beta and short-rate and macro surprises largely unchanged. This suggests that the term spread's forecasting power arises from a correlation with surprises about long-term bond yields that are orthogonal to betas and surprises to short rates and the macroeconomy.

By contrast, column (9) shows that the tent-shaped factor of Cochrane and Piazzesi (2005) enters positively and significantly and drives out the bond–stock beta, suggesting that the factor is correlated with compensation that investors require for holding positive-beta assets, as in Campbell et al. (2017). However, we have already seen that the bond–stock beta is not driven out by the Cochrane and Piazzesi (2005) factor in our baseline results for expected excess returns in Table II, suggesting that in addition to expected risk compensation the tent-shaped factor may also be correlated with ex-post long-term yield surprises. Hence, by distinguishing between expected and realized bond excess returns, we also provide new insights on the economic drivers of standard predictors of realized excess returns on bonds.

3.4 GMM Model Estimation

Finally, we provide a simple estimation of the model. The goal is to understand whether reasonable values for the price of beta risk and the actual and perceived persistence of beta can reconcile our main pieces of empirical evidence.

For estimation, we impose two simplifying restrictions. First, we estimate the forecaster-perceived beta persistence, $\tilde{\rho}_\beta$, freely but assume that markets are rational about the persistence of beta ($\hat{\rho}_\beta = \rho_\beta$). Thus, we allow the data to speak to whether macroeconomic forecasters over- or under-weight the role of bond–stock betas when predicting long-term bond yields.[34] Second, we permit markets and forecasters to have incorrect beliefs when they forecast the term structure factor, $\tilde{\rho}_f, \hat{\rho}_f \neq \rho_f$. However, as in Section 2, we impose the simplification that forecasters have the same beliefs as the market, i.e., $\tilde{\rho}_f = \hat{\rho}_f$. This second restriction is consistent with our evidence that proxies for biases have little impact on the relationship between expected excess returns and bond–stock betas (Tables V and VI), but ex-post interest rate and macro surprises are helpful in recovering the relationship between bond–stock betas and realized bond excess returns (Table VII). Intuitively, this restriction reflects that macroeconomic forecasters accurately capture the market's view of the economy and monetary policy (hence $\tilde{\rho}_f = \hat{\rho}_f$). However, macroeconomic forecasters' primary responsibility is not to form a view about term premia (hence $\tilde{\rho}_\beta \neq \hat{\rho}_\beta$).

There are three free parameters in the estimation: the price of risk (λ), the actual persistence of beta (ρ_β), which is assumed to equal the market's perception, $\hat{\rho}_\beta$, and forecasters' perceived persistence of beta ($\tilde{\rho}_\beta$). We estimate these parameters with four target moments: (i) the empirical persistence of beta; (ii) the relationship between bond–stock betas and

[34] Anchoring the market-perceived beta persistence to the actual persistence is necessary because otherwise only the product $\lambda\hat{\rho}_\beta$ is strongly identified.

Table VIII: Target Moments for GMM Estimation

Moment	Data	Model
Expected excess returns on $\beta_{n,t}$	9.99	9.99
	(1.62)	
Long-term yield minus expected T-bill rate on $\beta_{n,t}$	3.60	3.16
	(0.55)	
Realized excess returns on $\beta_{n,t}$, net of surprises	4.43	6.27
	(1.58)	
12-month persistence of bond–stock betas	0.93	0.98
	(0.06)	

Note: The table lists the model and data moments for the GMM estimation. Model moments are for a zero-coupon bond with maturity $n = 10$ years. Data standard errors in parentheses are those reported in the source tables, Table II column (1), Table IV Panel A column (1), and Table VII column (6). The persistence of bond–stock betas is the regression coefficient of bond–stock betas on their own 12-month lag in a panel of all bond maturities with $n \leq 10$ from 1975 to 2025 with Driscoll–Kraay standard errors with 18 lags.

expected excess returns (Table II column (1));[35] (iii) the relationship between bond–stock betas and the spread between the long-term yield and the expected Treasury bill rate over the next ten years (Table IV Panel A column (1)); and (iv) the relationship between realized bond excess returns and betas after controlling for short-rate and macroeconomic surprises (Table VII column (6)). This last moment provides a more precise estimate of the bond–stock beta loading for realized excess returns if innovations to bond–stock betas are uncorrelated with monetary policy and macroeconomic forecast errors.[36] For the GMM estimation, the monetary policy persistence $\tilde{\rho}_f = \hat{\rho}_f$ is then left unrestricted and does not need to be estimated.

The empirical moments and their model counterparts are presented in Table VIII. The model matches the data relatively well; all model moments are within roughly one standard error of the data. The estimated persistence of beta is slightly higher than in the data. The model exactly matches the coefficient from regressing expected excess returns on bond–stock betas through forecasters' perceived beta persistence $\tilde{\rho}_\beta$. The model also fits the coefficient from a regression of the difference between long-term yields and the expected T-bill rate onto bond–stock betas. In the model, the coefficient from regressing realized excess returns onto betas is slightly larger than in the data, but the difference is not statistically significant.

Table IX shows the estimated model parameter values. The estimated price of beta risk

[35] We target our benchmark estimate from Table II, which uses different bond maturities. However, the coefficient from a time series regression using just the 11-year bond in Table III is almost identical. A further gap between the model and the empirical evidence arises because the data use 11-year par bonds, while the model uses 10-year zero coupon bonds. For a simple model such as ours, this second discrepancy has only a quantitatively small effect on the target moments.

[36] In practice, innovations to bond–stock betas are roughly uncorrelated with forecast errors and do not affect the relationship matched by the GMM, see Table VII column (7).

is 6.43% annualized, squarely within the range for the empirical equity premium. The null of zero price of beta ($\lambda = 0$) is rejected at conventional levels of statistical significance. The actual and market-perceived persistence of beta is high, in line with the empirical persistence of beta. Forecasters' perceived persistence is lower, and the difference is marginally statistically significant. Note that this difference, which we assumed in the model (Assumption 2), is not imposed on the estimation but is instead an estimation result. Economically, the low forecaster-perceived persistence of beta is the way the model reconciles the fact that betas are more strongly related to expected excess returns than to the spread between the long-term yield and the expected T-bill rate. Intuitively, if forecasters under-weight term premia arising from bond–stock betas, the expected future yield moves less with beta than the current market yield, consistent with our results in Table IV. This quantitatively strengthens the relationship between expected excess returns and betas. Even though the gap between forecaster and market perceptions of bond–stock betas is only marginally statistically significant, its magnitude is nonetheless informative. Forecasters do not have to substantially underestimate the persistence of beta to fit the data.

3.5 Implications for Long-term Movements in Yields

We conclude this section with a simple back-of-the-envelope calculation for understanding the long decline in long-term bond yields since the 1980s, and their recent reversal. We estimate term premia for 10-year nominal and real (inflation-indexed) zero-coupon bonds following the term premium in Equation (6) and the model parameter estimates in Table IX, proxying for β_n with the 5-year rolling monthly beta.[37] To calculate model-implied real term premia, we use analogously calculated inflation-indexed bond–stock betas and the same parameter values.

The estimated model implies that term premia are quantitatively meaningful and account for more than half of both the long decline and the recent increase in long-term yields since 2020.[38] Declining inflation expectations and other trends arguably also played a substantial role in determining long-term yields over the past four decades, but they have likely been less important in the post-2020 period despite the substantial volatility of realized inflation.

[37] Including as a moment the relationship between realized returns and betas reduces the estimated impact of betas on term premia. Thus, these calculations are more conservative than a calculation based purely on our evidence in Table II.

[38] The model-implied nominal term premium exhibits a substantial peak-to-trough decline of 3.5 percentage points from 2.3% in March 1986 to −1.2% in May 2015. Over the same period, the nominal 10-year zero coupon yield fell by 5.3 percentage points, from 7.5% to 2.2%, so that the decline in the model-implied term premium accounts for more than half of the decline in the long-term yield. Between January 2020 and September 2025 the ten-year nominal yield rose by 2.7 percentage points, and the model-implied term premium rose by 1.6 percentage points. These simple comparisons are of course only illustrative and depend on start and end points, since actual yields exhibit short-term volatility around these longer-term patterns.

Table IX: Parameter Estimates

		Estimate
Price of bond–stock beta risk	λ	6.43
		(1.35)
Market and actual beta persistence	$\widehat{\rho}_\beta = \rho_\beta$	0.98
		(0.05)
Forecasters' perceived persistence	$\widetilde{\rho}_\beta$	0.83
		(0.09)
Forecaster misperception	$\widehat{\rho}_\beta - \widetilde{\rho}_\beta$	0.14
		(0.09)
J-statistic		2.62
p-value		0.11

Note: GMM parameter estimates obtained by matching the four moments in Table VIII. The specification imposes $\widehat{\rho}_\beta = \rho_\beta$, so that markets perceive beta persistence correctly, and uses a diagonal weighting matrix consisting of the inverse squared empirical standard errors. Parameter standard errors in parentheses are delta-method sandwich standard errors. The J-statistic is compared to a χ^2_1 distribution.

Overall, an estimated version of our simple model reconciles the empirical evidence from expected excess returns, yields versus long-horizon T-bill forecasts, and realized excess returns. The model implies that term premia in long-term bonds have varied substantially, contributing both to the long decline in long-term bond yields since the 1980s and the recent increase in long-term yields.

4. International Panel Evidence

We now show that our main empirical results also hold in an international panel of twelve developed countries: Canada, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, the U.K., and the U.S.

4.1 International Data

We construct expected excess returns following Equation (7). To obtain bond yield forecasts for the international panel, we turn to Consensus Economics. For each month, we use the average forecast of the 10-year yield 12 months ahead, which is the only bond maturity consistently reported across countries in Consensus Economics. Thus, we have a country-by-month panel. Consensus Economics reports the exact survey date, which typically falls on the second or third Monday of each month.[39] The Consensus Economics sample begins

[39]The definition of the Consensus Economics forecasts is largely comparable to that of the Blue Chip forecasts described above. For the U.S., the two series have a correlation of 0.99. Minor differences may arise because: (i) Consensus Economics elicits forecasts of the yield 12 months ahead, whereas Blue Chip reports forecasts of the average yield in the quarter 4 quarters ahead; (ii) survey dates may differ slightly within the month; and (iii) the surveyed forecasters do not perfectly overlap.

in September 1989 and ends in December 2024. For some countries, Consensus Economics forecasts become available at later dates. Coverage starts in 1994 for the Netherlands, Spain, and Sweden, and in 1998 for Norway and Switzerland.

We obtain actual month-end 10-year government bond yields, 1-year government bond yields, Treasury bill rates, and total stock return indexes from Global Financial Data (GFD).[40] Outside the U.S., we generally do not have full yield curves. We therefore use the 10-year bond yield as a substitute for the 11-year bond yield in Equation (7). For the U.S., we use the 11-year government par bond yield from Gürkaynak et al. (2007), as in our baseline U.S. analysis.[41] To proxy for the risk-free rate in Equation (7), we use the 1-year government bond yield.

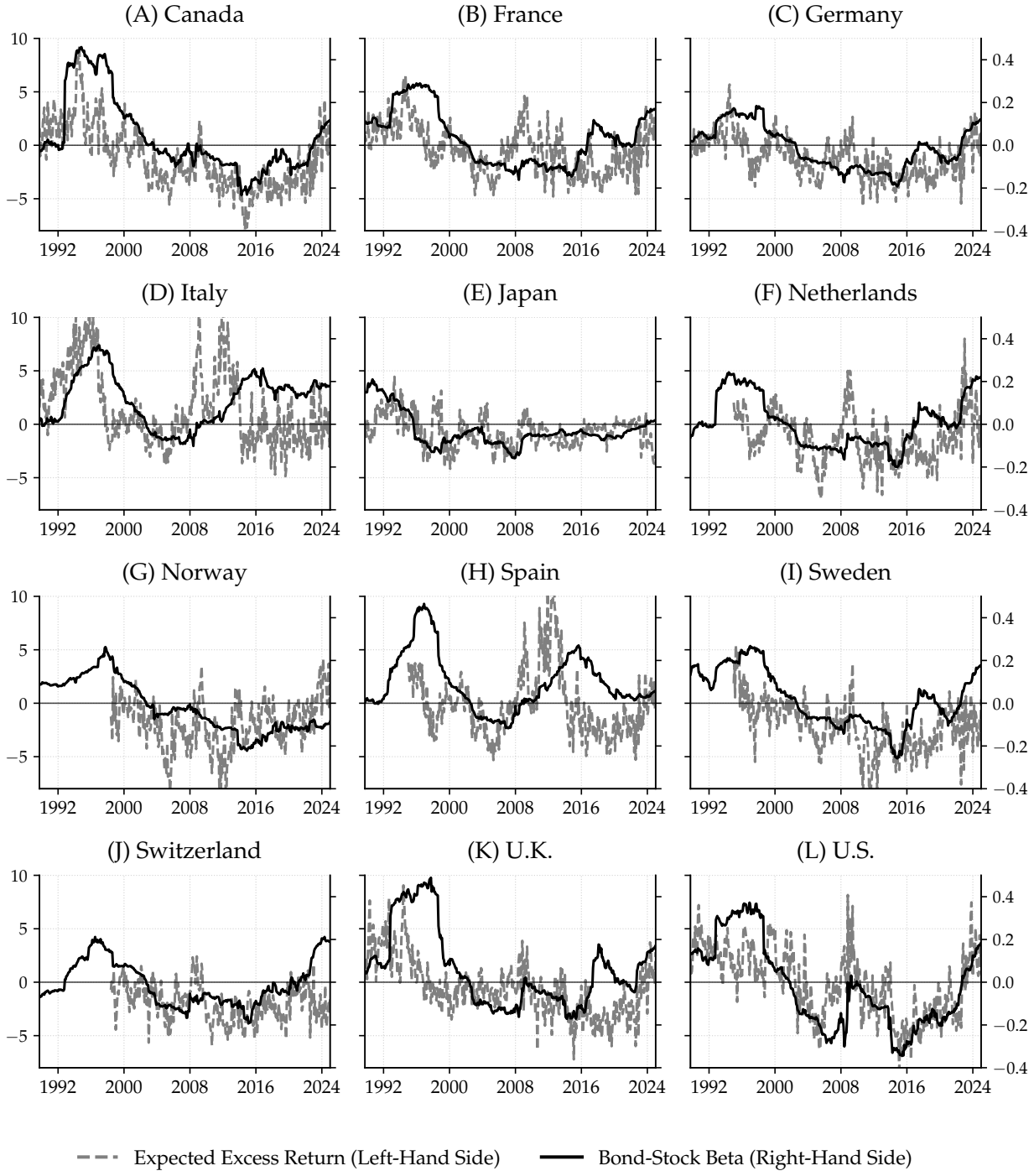
We construct bond–stock betas in the international panel using monthly realized excess returns on 10-year government bonds and monthly excess returns on the domestic stock market, with both stock returns and the risk-free rate in local currency. As an alternative specification, we also compute bond–stock betas using returns on the U.S. stock market in excess of the U.S. T-bill, which corresponds to currency-hedged excess returns on the U.S. stock market for local currency investors up to second-order hedging approximation error.

Expected excess returns and rolling betas for the international data are plotted in Figure 3, and summary statistics are reported in Appendix Table B9. Each panel of Figure 3 plots data for one country, with expected excess returns plotted as a dashed gray line (left-hand y-axis) and beta plotted as a solid black line (right-hand y-axis). The figure illustrates the usefulness of bringing in data from other countries. First, a close correlation between expected bond excess returns and betas is visible consistently across countries. For example, the univariate correlation between smoothed expected excess returns and bond–stock betas equals 84% for Canada and 68% for Germany. Second, cross-country differences in betas are also mirrored in expected bond excess returns. One case in point is Japan, which experienced negative bond–stock betas substantially earlier than most other countries, and correspondingly saw an earlier switch to negative expected bond excess returns.

[40] For Italy, the series of the total return index from GFD stops in 2019. We supplement this series with the total return index for the Italian market from Datastream. For a small subset of countries, the 1-year government bond yield is not available over the full sample. In those cases, we instead use the three-month Treasury bill rate. The 1-year rate is not available for Canada until 2007, for the Netherlands until 1999, and for Norway until 2003.

[41] For the U.S., we average daily bond yields over the last week of each month to reduce the volatility of outliers and use the average BCFF forecast of 10-year Treasury yields. These choices ensure consistency with the U.S. analysis but are not crucial for the results.

Figure 3: Expected Excess Returns and Betas by Country



Note: The figure reports estimated expected excess returns and betas for each country. The sample runs from September 1989 through December 2024. We compute expected bond excess returns using survey forecasts from Consensus Economics and financial data from Global Financial Database. The only exception is the U.S., where we use our baseline expected excess returns based on the BCFF. Bond–stock betas are estimated using monthly bond and stock excess returns from the Global Financial Database. We use a 5-year (60-month) rolling window. Consensus Economics data are not available from the start of the sample for a number of countries: data for the Netherlands, Spain, and Sweden start in December 1994, while data for Norway and Switzerland start in May 1998. We use the same x-axis and y-axis for all panels.

4.2 International Expected Excess Returns vs. Betas

We next estimate the relationship between expected excess returns and betas across countries and over time via the following international panel regression:

$$\tilde{E}_t xr_{c,t\to t+12} = \alpha + \lambda\beta_{c,t} + X'_{c,t}\delta + \eta_{c,t}, \tag{15}$$

where c indexes countries, $\tilde{E}_t xr_{c,t\to t+12}$ denotes the expected excess return for the 11-year bond in country c over a 12-month forecast horizon, $\beta_{c,t}$ is the 5-year rolling beta estimated using information available at time t, and $X_{c,t}$ is a vector of control variables. The specification mirrors Equation (11), with the cross-sectional dimension given by countries rather than maturities.

Table X shows that the estimated relationship between bond–stock betas and expected bond excess returns is significantly positive and economically meaningful for the international panel. Column (1) starts with the univariate regression without controls or fixed effects, finding an estimated coefficient of 8.73. Its magnitude is comparable to the U.S. results reported in Table II. A one-standard-deviation increase in the beta is associated with a 0.15 $\times$ 8.73 = 1.31 percentage point increase in the expected excess return. For comparison, the standard deviation of expected excess returns is 2.73.[42]

The relationship between expected bond excess returns and betas in the international panel is robust to a wide range of controls and fixed effects. In column (2), we include country fixed effects, which absorb time-invariant country-specific characteristics, such as persistent differences in liquidity or the special status of U.S. Treasury bonds. The estimated coefficient is still large and equal to 8.07. In column (3), we add country-specific linear time trends. The coefficient in this specification is equal to 6.37.

In column (4), we add the rolling variance of bond excess returns as a control. In the international panel, we again find a positive but insignificant coefficient on the bond return variance, similar to the U.S. regression in Table II. The coefficient on bond–stock betas and the R^2 are again little changed. In column (5), we control for the term spread. The coefficient on beta remains unchanged, while the coefficient on the term spread is close to zero. Column (6) controls for the Laubach-Williams estimate of the natural rate of interest (r-star). We use the U.S. r-star estimate for all countries in our sample because country-

[42]The estimated constant in column (1) is statistically significantly negative. This may be due to the fact that data constraints force us to approximate the 11-year yield in the expected excess return Equation (7) with a 10-year yield, leading us to understate expected excess returns on average. If the actual 11-year yield $y^{par}_{n,t}$ is under-estimated by 10 bps and bond duration is roughly 9 years, this could explain the estimated -90 bps constant merely through measurement bias due to the unavailability of complete yield curves for the international sample. On average, the U.S. 11-year par yield exceeds the U.S. 10-year par yield by 7 bps in Gürkaynak et al. (2007) data.

level estimates are not available for most countries in our sample.[43] The loading on beta is virtually unchanged after controlling for r-star. This is not surprising, since expected excess returns already subtract out the expected future long-term bond yield, which is likely to contain similar variation in r-star. Overall, this column suggests time-varying bond risks and risk-neutral r-star are complementary drivers of low-frequency movements in bond yields.

Columns (7) and (8) add increasingly restrictive fixed effects. Column (7) controls for time fixed effects, thereby exploiting only cross-country variation. The coefficient on bond–stock betas declines somewhat in magnitude, but remains statistically significant and on the order of the equity premium. The fact that the coefficient remains large and significant indicates that, while countries comove over time, countries that tend to have more positive bond–stock betas also tend to have higher expected excess returns. Only when we include both time and country fixed effects in column (8) does the estimated coefficient on beta become smaller in magnitude, but it remains highly statistically significant. In this specification, a large share of the variation is absorbed by fixed effects, so it is not surprising that results change somewhat, even though the slope remains statistically significant and large.

So far, the results use betas constructed using domestic stock market returns. However, if investors are primarily exposed to the U.S. equity market, the relevant beta may instead be the one computed with respect to U.S. stock returns. For this reason, we replicate the same set of regressions using betas of bond returns with respect to the U.S. stock market. The results are reported in Appendix Table B10 and are very similar to the baseline results. A natural explanation is that stock market returns are highly correlated across countries, so the choice between domestic and U.S. stock returns has little effect on our results.

Individual Countries We next show that the relationship between expected bond excess returns and betas is consistently present for all countries, and not driven by a small number of countries. For this, we estimate Equation (15) separately for each country using betas computed with respect to the domestic stock market, thereby only exploiting time-series variation. The consistent results across countries are visible in Figure 4, which reports the estimated coefficient $\hat{\lambda}_c$ together with its 95% confidence interval for each country. The figure shows that the estimated coefficient is positive for all countries. The magnitude of the coefficient varies across countries: it is the largest in Canada, Germany, and the United States, where it is close to 10; it lies between 5 and 10 for France, Japan, the Netherlands, Sweden, and the United Kingdom; and it is smaller in Italy, Norway, Spain, and Switzerland,

[43] Measures of r-star are also likely to be highly correlated across countries. The Federal Reserve Bank of New York provides Holston-Laubach-Williams estimates for Canada and the euro area, whose correlations with the U.S. r-star are 0.92 and 0.86, respectively. For the international panel, we do not have enough yield curve data to construct the Cochrane-Piazzesi factor or level, slope and curvature factors.

Table X: Expected Excess Returns and Stock Market Betas: International Panel

Dependent Variable: Expected Excess Returns								
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Bond-Stock Beta	8.73***	8.07***	6.37***	7.01***	8.09***	7.58***	6.17***	4.16***
	(1.70)	(1.59)	(2.08)	(1.62)	(1.58)	(1.62)	(1.99)	(1.13)
Bond Variance				1.17				
				(0.94)				
Term Spread					0.19			
					(0.16)			
r-star						0.25		
						(0.22)		
Constant	-0.91***							
	(0.26)							
Country FE		✓	✓	✓	✓	✓		✓
Time Trend × Country FE			✓					
Time FE							✓	✓
R-squared	0.22	0.29	0.37	0.30	0.30	0.30	0.56	0.63
Observations	4690	4690	4690	4690	4690	4690	4690	4690

Note: The table reports the results of Equation (15):

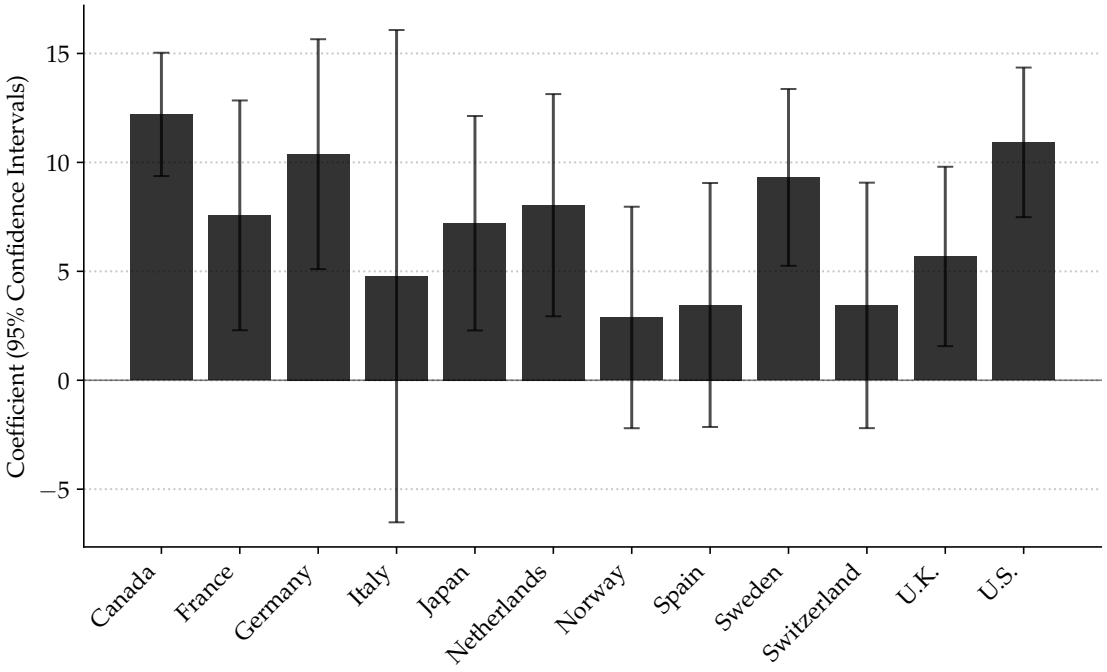
$$\tilde{E}_t xr_{c,t\to t+12} = \alpha + \lambda\beta_{c,t} + X'_{c,t}\delta + \eta_{c,t},$$

estimated using our international panel. The dependent variable is the expected one-year bond excess return, constructed from Consensus Economics twelve-month-ahead forecasts of the ten-year government bond yield. The only exception is the U.S., where we use our baseline expected excess returns based on the BCFF. The bond–stock beta $\beta_{c,t}$ is estimated from monthly bond and stock excess returns over a five-year (60-month) rolling window, using domestic stock market excess returns. Column (1) is the pooled baseline; Columns (2)–(3) add country fixed effects and country-specific linear time trends; Columns (4)–(6) add the bond return variance (multiplied by 100), the term spread, the Laubach-Williams estimates of the natural rate of interest, r-star (with country fixed effects); and Columns (7)–(8) include time fixed effects and country-plus-time fixed effects. The bond return variance is the annualized realized variance of monthly bond excess returns over the five-year window, multiplied by 100; the term spread is the difference between the ten-year and one-year government bond yields. Driscoll–Kraay standard errors with 27 lags are reported in parentheses. Stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The data are monthly and span September 1989 to December 2024.

where it remains positive but is not always statistically significant. This shows that the results in Table X are not driven by sovereign default risk.

Differences in magnitude partly reflect differences in sample length and spikes in expected bond excess returns during the European debt crisis. Italy and Spain, most exposed to the European debt crisis, exhibit substantial volatility in both beta and expected returns, as can be seen in Figure 3. For both countries, betas and expected excess returns display a very strong comovement from the start of their samples (1990 for Italy and 1994 for Spain) to 2014. After 2014, this relationship weakens: the beta remains positive, while expected excess returns turn negative. This break coincides with the announcement of the ECB's quantitative easing (QE) program in January 2015. One interpretation is that QE substantially lowered sovereign spreads and term premia in these countries, leading to persistently low expected

Figure 4: Country-by-Country Regressions of Expected Excess Returns onto Beta



Note: The figure plots the estimated coefficient from Equation (15), estimated separately for each country. The figure also reports 95% confidence intervals constructed using Newey–West standard errors with 27 lags. The data are monthly, and for most countries, the sample spans September 1989 to December 2024. A few countries use shorter samples, as detailed in Table B9.

returns.

In contrast, Norway and Switzerland display a weaker relationship between expected bond excess returns and bond–stock betas likely because their data start in 1998, shortly before bond–stock betas became negative. There is less time series variation in beta over the shorter samples for these countries, so it is unsurprising that the relationship is estimated to be weaker.

Figure B6 presents the corresponding country-by-country regressions when bond–stock betas are computed with respect to U.S. stock excess returns. Overall, the results are robust and very similar. However, the estimated coefficients are somewhat smaller in magnitude relative to the baseline results, with the difference being particularly pronounced for Italy and Spain. This may suggest that the domestic stock market is the relevant market portfolio for investors in the local bond market, particularly for those two countries.

4.3 Realized Excess Returns and Betas, International Panel

The international panel results confirm the robust relationship between bond–stock betas and expected excess returns that we found for the U.S. panel in Section 3.2. We now test

the relationship between future realized excess returns and bond–stock beta, mirroring the discussion in Section 3.3. Similarly to Equation (13) for the U.S. panel, we estimate predictive regressions of the form:

$$xr^{par}_{c,t\to t+12} = \alpha + \gamma\,\beta_{c,t} + X'_{c,t}\delta + \varepsilon_{c,t+12}. \tag{16}$$

Here, $xr^{par}_{c,t\to t+12}$ denotes the realized excess return on a 10-year bond of country c between t and $t+12$, while $\beta_{c,t}$ denotes the bond–stock beta estimated using information available through time t.

The results are presented in Table XI. In the first column, where we do not include any controls, we find a coefficient γ on bond–stock beta of 6.07. This estimate is smaller in magnitude than the corresponding coefficient for expected excess returns (8.73) and is not statistically significant. The R^2 of the regression is also low, at 2%. Thus, as in the U.S. results, bond–stock betas have a noisy relationship with future realized excess returns in the absence of controls.

We then control for surprises in long-term bond yields, short-term rates, and macroeconomic variables. In column (2), we control for the scaled yield surprise, defined as the difference between the realized 10-year yield and its forecasted value. We find a loading on bond–stock beta of 8.56, virtually identical to the coefficient estimated for expected excess returns. The loading on the yield surprise is negative and equal to -0.94. As discussed in Section 3.3, Equation (14) implies that once we control for the scaled yield surprise the coefficient γ should be close to the λ estimated in Table X, while the coefficient on the scaled yield surprise should be close to -1. This is what we find.

In columns (3)–(7), we control for surprises in short-term rates and macroeconomic news.[44] In column (3), once we control for short-term rate surprises, we find a positive and statistically significant coefficient on beta, equal to 7.41. The R^2 also increases to 36%. These results suggest that movements in short-term rates relative to forecasts were correlated with the level of bond–stock beta and added substantial noise to the relationship between realized excess returns and beta.

Column (4) adds GDP surprises, while column (5) adds CPI surprises. In interpreting these results, we caution that CPI and GDP growth forecast horizons do not generally match the horizons for expected excess returns or Treasury bill and yield forecasts, so CPI and GDP growth surprises are likely to be noisier in the international panel. Similar to our findings for the U.S., controlling for GDP surprises increases the loading on bond–stock

[44]For macroeconomic surprises, the realized real GDP growth and CPI inflation are from the Global Macro Database (GMD) of Müller et al. (2025).

Table XI: Realized Returns and Betas: International Evidence

Dependent Variable: Realized Excess Return							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Beta	6.07	8.56***	7.41**	7.07	3.29	5.56	4.21
	(4.77)	(1.57)	(3.45)	(4.96)	(4.21)	(3.24)	(3.43)
Yield Surprise		-0.94***					
		(0.03)					
Short-Rate Surprise			-0.50***			-0.39***	-0.40***
			(0.08)			(0.06)	(0.06)
GDP Surprise				-0.11**		0.01	0.01
				(0.04)		(0.03)	(0.03)
CPI Surprise					-0.48***	-0.30***	-0.31***
					(0.05)	(0.05)	(0.05)
Country FE							✓
R-squared	0.02	0.88	0.36	0.05	0.27	0.44	0.45
Observations	4690	4690	4689	4573	4519	4401	4401

Note: The table reports the results of Equation (16) estimated using our international panel:

$$xr^{par}_{c,t\to t+12} = \alpha + \gamma\,\beta_{c,t} + X'_{c,t}\delta + \varepsilon_{c,t+12},$$

where $xr^{par}_{c,t\to t+12}$ denotes the realized excess return on a 10-year bond of country c between t and $t+12$, while $\beta_{c,t}$ denotes the five-year rolling bond–stock beta estimated using information available through time t. Driscoll–Kraay standard errors with 27 lags are reported in parentheses. Stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample consists of monthly data from September 1989 to December 2024. The number of observations differs across columns because some surprises are not available for the full sample: GDP forecasts start in December 1991 for the U.S., December 1992 for Germany, and December 1993 for Japan; CPI forecasts for the U.K. start in December 2003; and the short-rate surprise is missing for one observation for the Netherlands.

beta—although the coefficient is not statistically significant—and only modestly increases the R^2. Controlling for inflation surprises alone reduces the loading on beta. In column (6), we jointly control for short-term rate and macroeconomic surprises. We find a positive though not statistically significant coefficient on bond–stock beta of 5.56.

In Column (7), we add country fixed effects to the set of surprises included in column (6). The loading on bond–stock beta falls to 4.21 and is not statistically distinguishable from zero. This result suggests that the positive relationship between bond–stock beta and realized excess returns is also driven by cross-country variation: countries with higher bond–stock betas also tend to experience higher realized excess returns.

Overall, the results in this section confirm the findings from the U.S.: the relationship between bond–stock betas and realized returns is weaker than the relationship with expected excess returns. However, controlling for forecasters' surprises about long-term yields restores the relationship. In the international panel, evidence is more mixed about whether controlling for surprises about short-term rates and macroeconomic outcomes is also sufficient to restore the relationship.

5. Conclusion

We show that bond risks, measured by bond–stock betas, are a central determinant of bond risk premia priced in markets when subjective interest rate expectations are accounted for. Using survey forecasts to isolate ex-ante expectations, we document a strong and stable relationship between bond–stock betas and expected bond excess returns in U.S. data and in an international panel of developed economies. In both settings, the estimated market price of risk is economically large. We show that this robust relationship reflects the fact that markets price bond–stock beta risk into market yields. Finally, we show that the relationship between realized bond excess returns and betas is weak. This reflects forecast errors, which may be generated by learning or biases in expectation formation and are absorbed by our construction of expected excess returns. Once we control for surprises in short-term rates and macroeconomic variables, the relationship between bond–stock betas and realized returns re-emerges, particularly in the U.S. sample.

Our findings provide a new perspective on the secular decline, and more recent increase, in long-term yields. Our estimated link between expected excess returns and bond–stock betas attributes more than one half of the secular decline in long-term interest rates since the 1980s to a macroeconomically relevant measure of risk. Moreover, rising bond risk can account for the majority of the increase in 10-year Treasury rates since 2020. In contrast to standard estimates, our results suggest that term premia had already turned negative in the early 2000s, prior to the global financial crisis.

This paper opens up avenues for future research. Government bond markets are a cornerstone of financial markets. Our finding that subjective bond risk premia increase with bond risks is likely to have implications for financial investors, optimal portfolio choice, and fiscal and monetary policy.

References

Acharya, V. V. and T. Laarits (2025): "When do Treasuries earn the convenience yield? A hedging perspective," .

Ang, A. and M. Piazzesi (2003): "A No-Arbitrage Vector Autoregression of Term Structure Dynamics with Macroeconomic and Latent Variables," *Journal of Monetary Economics*, 50, 745–787.

Ang, A., M. Piazzesi, and M. Wei (2006): "What Does the Yield Curve Tell Us about GDP Growth?" *Journal of Econometrics*, 131, 359–403.

Baele, L., G. Bekaert, and K. Inghelbrecht (2010): "The Determinants of Stock and Bond Return Comovements," *Review of Financial Studies*, 23, 2374–2428.

Bauer, M. D., C. E. Pflueger, and A. Sunderam (2024): "Perceptions About Monetary Policy," *Quarterly Journal of Economics*, 139, 2227–2278.

Bauer, M. D. and G. D. Rudebusch (2020): "Interest Rates under Falling Stars," *American Economic Review*, 110, 1316–1354.

Bernanke, B. S. (2005): "The Global Saving Glut and the U.S. Current Account Deficit," Remarks at the Sandridge Lecture, Virginia Association of Economists, Richmond, Virginia.

Bianchi, F., M. Lettau, and S. C. Ludvigson (2022a): "Monetary policy and asset valuation," *Journal of Finance*, 77, 967–1017.

Bianchi, F., S. C. Ludvigson, and S. Ma (2022b): "Belief Distortions and Macroeconomic Fluctuations," *American Economic Review*, 112, 2269–2315.

——— (2022c): "A Structural Approach to High-Frequency Event Studies: The Fed and Markets as Case History," Working Paper 30072, National Bureau of Economic Research, revised August 2026.

Black, F., M. C. Jensen, and M. Scholes (1972): "The Capital Asset Pricing Model: Some Empirical Tests," in *Studies in the Theory of Capital Markets*, ed. by M. C. Jensen, New York: Praeger, 79–121.

Bordalo, P., N. Gennaioli, R. La Porta, and A. Shleifer (2024): "Belief overreaction and stock market puzzles," *Journal of Political Economy*, 132, 1450–1484.

Bordalo, P., N. Gennaioli, Y. Ma, and A. Shleifer (2020): "Overreaction in macroeconomic expectations," *American Economic Review*, 110, 2748–2782.

Brav, A., R. Lehavy, and R. Michaely (2005): "Using expectations to test asset pricing models," *Financial Management*, 34, 31–64.

Brunnermeier, M. K., S. Merkel, and Y. Sannikov (2024): "Safe assets," *Journal of Political Economy*, 132, 3603–3657.

Caballero, R. J., E. Farhi, and P.-O. Gourinchas (2008): "An equilibrium model of "global imbalances" and low interest rates," *American Economic Review*, 98, 358–393.

Campbell, J. Y. (2017): *Financial decisions and markets: A course in asset pricing*, Princeton University Press.

Campbell, J. Y., C. Pflueger, and L. M. Viceira (2020): "Macroeconomic drivers of bond and equity risks," *Journal of Political Economy*, 128, 3148–3185.

——— (2026): "Bond-stock comovements," *Annual Review of Financial Economics, forthcoming*.

Campbell, J. Y. and R. J. Shiller (1991): "Yield spreads and interest rate movements: A bird's eye view," *Review of Economic Studies*, 58, 495–514.

Campbell, J. Y., A. Sunderam, and L. M. Viceira (2017): "Inflation Bets or Deflation Hedges? The Changing Risks of Nominal Bonds," *Critical Finance Review*, 6, 263–301.

Cieslak, A. (2018): "Short-rate expectations and unexpected returns in Treasury bonds," *Review of Financial Studies*, 31, 3265–3306.

Cochrane, J. (2009): *Asset pricing: Revised edition*, Princeton University Press.

Cochrane, J. H. and M. Piazzesi (2005): "Bond Risk Premia," *American Economic Review*, 95, 138–160.

Couts, S. J., A. S. Goncalves, and J. A. Loudis (2026): "The Subjective Risk and Return Expectations of Institutional Investors," *Working Paper, University of Southern California, Ohio State University, and Notre Dame University*.

Cox, J. C., J. E. Ingersoll, and S. A. Ross (1985): "A Theory of the Term Structure of Interest Rates," *Econometrica*, 53, 385–407.

Del Negro, M., D. Giannone, M. P. Giannoni, and A. Tambalotti (2017): "Safety, liquidity, and the natural rate of interest," *Brookings Papers on Economic Activity*, 2017, 235–316.

Delao, R. and S. Myers (2021): "Subjective cash flow and discount rate expectations," *Journal of Finance*, 76, 1339–1387.

Dimson, E. (1979): "Risk measurement when shares are subject to infrequent trading," *Journal of Financial Economics*, 7, 197–226.

Fama, E. F. and R. R. Bliss (1987): "The Information in Long-Maturity Forward Rates," *American Economic Review*, 77, 680–692.

Fama, E. F. and K. R. French (1993): "Common Risk Factors in the Returns on Stocks and Bonds," *Journal of Financial Economics*, 33, 3–56.

——— (2004): "The Capital Asset Pricing Model: Theory and Evidence," *Journal of Economic Perspectives*, 18, 25–46.

Fama, E. F. and J. D. MacBeth (1973): "Risk, Return, and Equilibrium: Empirical Tests," *Journal of Political Economy*, 81, 607–636.

Farmer, L. E., E. Nakamura, and J. Steinsson (2024): "Learning about the Long Run," *Journal of Political Economy*, 132, 3334–3377.

Frazzini, A. and L. H. Pedersen (2014): "Betting against beta," *Journal of Financial Economics*, 111, 1–25.

Gomez, M. and É. Gouin-Bonenfant (2024): "Wealth inequality in a low rate environment," *Econometrica*, 92, 201–246.

Gomez Cram, R., H. Kung, and H. Lustig (2024): "Government Debt in Mature Economies: Safe or Risky?" in *Kansas City Federal Reserve Symposium*, 113.

Greenwald, D. L., M. Leombroni, H. Lustig, and S. Van Nieuwerburgh (2026): "Winners and Losers When Interest Rates Change," Working Paper, NYU, Stanford University, and Boston College.

Gürkaynak, R. S., B. Sack, and J. H. Wright (2007): "The US Treasury yield curve: 1961 to the present," *Journal of Monetary Economics*, 54, 2291–2304.

Haddad, V., A. Moreira, and T. Muir (2025): "Whatever it takes? The impact of conditional policy promises," *American Economic Review*, 115, 295–329.

Jensen, T. I. (2026): "Subjective risk and return," *Working Paper, Yale School of Management*.

Kekre, R., M. Lenel, and F. Mainardi (2024): "Monetary policy, segmentation, and the term structure," *NBER Working Paper wp32324*.

Kiefer, N. M. and T. J. Vogelsang (2002): "Heteroskedasticity-autocorrelation robust standard errors using the Bartlett kernel without truncation," *Econometrica*, 70, 2093–2095.

——— (2005): "A new asymptotic theory for heteroskedasticity-autocorrelation robust tests," *Econometric Theory*, 21, 1130–1164.

Kozicki, S. and P. A. Tinsley (2001): "Shifting endpoints in the term structure of interest rates," *Journal of Monetary Economics*, 47, 613–652.

Kremens, L., I. Martin, and L. Varela (2025): "Long-Horizon Exchange Rate Expectations," *Journal of Finance*, 80, 3695–3724.

Krishnamurthy, A. and A. Vissing-Jorgensen (2012): "The Aggregate Demand for Treasury Debt," *Journal of Political Economy*, 120, 233–267.

Laubach, T. and J. C. Williams (2003): "Measuring the natural rate of interest," *Review of Economics and Statistics*, 85, 1063–1070.

——— (2016): "Measuring the natural rate of interest redux," *Business Economics*, 51, 57–67.

Lazarus, E., D. J. Lewis, J. H. Stock, and M. W. Watson (2018): "HAR Inference: Recommendations for Practice," *Journal of Business & Economic Statistics*, 36, 541–574.

Lintner, J. (1965): "The Valuation of Risk Assets and the Selection of Risky Investments in Stock Portfolios and Capital Budgets," *Review of Economics and Statistics*, 47, 13–37.

Lustig, H. and A. Verdelhan (2007): "The cross section of foreign currency risk premia and consumption growth risk," *American Economic Review*, 97, 89–117.

Molavi, P., A. Tahbaz-Salehi, and A. Vedolin (2024): "Model complexity, expectations, and asset prices," *Review of Economic Studies*, 91, 2462–2507.

——— (2025): "Expectations and the Term Structure of Interest Rates," *NBER Working Paper wp35766*.

Müller, K., C. Xu, M. Lehbib, and Z. Chen (2025): "The Global Macro Database: A New International Macroeconomic Dataset," Working Paper 33714, National Bureau of Economic Research.

Nagel, S. and Z. Xu (2023): "Dynamics of subjective risk premia," *Journal of Financial Economics*, 150, 103713.

Pesch, D., P. Whelan, and I. Piatti (2024): "Subjective Risk Premia in Bond and FX Markets," *Working Paper, Queen Mary University of London and CUHK Business School*.

Pflueger, C. (2025): "Back to the 1980s or not? The drivers of inflation and real risks in Treasury bonds," *Journal of Financial Economics*, 167, 104027.

Piazzesi, M., J. Salomao, and M. Schneider (2015): "Trend and Cycle in Bond Premia," *Working Paper, Stanford University and University of Minnesota*.

Piazzesi, M. and M. Schneider (2006): "Equilibrium yield curves," *NBER Macroeconomics Annual*, 21, 389–472.

Reis, R. (2026): "The four r-stars: From interest rates to inflation and back," in *Medium and Long-Run Trends in Interest Rates: Causes and Implications for Monetary Policy*, Banco Central de Chile.

Rogers, R. (2026): "Risk-based Interest Rate Expectations," *Working Paper, London School of Economics*.

Rudebusch, G. D. and E. T. Swanson (2012): "The bond premium in a DSGE model with long-run real and nominal risks," *American Economic Journal: Macroeconomics*, 4, 105–143.

Sharpe, W. F. (1964): "Capital Asset Prices: A Theory of Market Equilibrium under Conditions of Risk," *Journal of Finance*, 19, 425–442.

Shue, K., R. Townsend, and C. Wang (2025): "Long-Short Interest Rate Confusion," *Working paper, Yale, UC San Diego, and University of Notre Dame.*

Summers, L. H. (2013): "Remarks at the IMF Fourteenth Annual Research Conference in Honor of Stanley Fischer," Speech at the Fourteenth Annual Research Conference, International Monetary Fund.

Vasicek, O. (1977): "An Equilibrium Characterization of the Term Structure," *Journal of Financial Economics*, 5, 177–188.

Wicksell, K. (1898): *Geldzins und Güterpreise*, Jena: Gustav Fischer, English translation: *Interest and Prices*, Macmillan, London, 1936.

A. Model Proofs

Proof. **Proof of Proposition 1:** The zero-coupon bond yield satisfies

$$\begin{aligned} y_{n,t} &= \underbrace{\frac{1}{n}\widehat{\mathbb{E}}_t \sum_{i=0}^{n-1} f_{t+12i}}_{\text{risk-neutral component}} + \underbrace{\frac{1}{n}\sum_{i=1}^{n-1} \widehat{\mathbb{E}}_t xr_{n-i+1,t+12i}}_{\text{term-premium component}} \\ &= \frac{1}{n}\widehat{\mathbb{E}}_t \sum_{i=0}^{n-1} f_{t+12i} + \frac{\lambda}{n}\sum_{i=1}^{n-1} \widehat{\mathbb{E}}_t \beta_{n-i+1,t+12i} \\ &= \frac{1}{n}\widehat{\mathbb{E}}_t \sum_{i=0}^{n-1} f_{t+12i} + \frac{\lambda}{n}\sum_{i=1}^{n-1} \frac{n-i+1}{n}\widehat{\mathbb{E}}_t \beta_{n,t+12i} \\ &= \frac{1}{n}\sum_{i=0}^{n-1} \hat{\rho}_f^i f_t + \frac{\lambda}{n}\sum_{i=1}^{n-1} \frac{n-i+1}{n}\hat{\rho}_\beta^i \beta_{n,t} \\ &= \frac{1-\hat{\rho}_f^n}{1-\hat{\rho}_f}\frac{1}{n} f_t + \lambda S_n(\hat{\rho}_\beta)\beta_{n,t}, \end{aligned} \tag{A.2}$$

where

$$S_n(\rho_\beta) \equiv \frac{1}{n^2}\sum_{i=1}^{n-1}(n-i+1)\rho_\beta^i. \tag{A.3}$$

S_n satisfies the telescoping property:

$$n^2 S_n(\rho_\beta) - \rho_\beta (n-1)^2 S_{n-1}(\rho_\beta) = n\rho_\beta, \qquad S_1(\rho_\beta) = 0. \tag{A.4}$$

For part a), realized bond excess returns satisfy

$$xr_{n,t+12} = n y_{n,t} - (n-1) y_{n-1,t+12} - y_{1,t}. \tag{A.5}$$

Substituting the expression for the actual yield into the excess-return identity and using $y_{1,t} = f_t$ gives

$$\begin{aligned} xr_{n,t+12} &= \lambda n S_n(\hat{\rho}_\beta)\beta_{n,t} - \lambda(n-1)S_{n-1}(\hat{\rho}_\beta)\beta_{n-1,t+12} \\ &\quad + \left[\frac{1-\hat{\rho}_f^n}{1-\hat{\rho}_f} f_t - \frac{1-\hat{\rho}_f^{n-1}}{1-\hat{\rho}_f} f_{t+12} - f_t\right]. \end{aligned} \tag{A.6}$$

Actual state dynamics imply

$$f_{t+12} = \rho_f f_t + \varepsilon_{t+12}, \qquad \beta_{n-1,t+12} = \frac{n-1}{n}\beta_{n,t+12} = \frac{n-1}{n}\left(\rho_\beta \beta_{n,t} + \eta_{t+12}\right).$$

Therefore,

$$\begin{aligned} xr_{n,t+12} &= \lambda\left[nS_n(\hat{\rho}_\beta) - \frac{(n-1)^2}{n}\rho_\beta S_{n-1}(\hat{\rho}_\beta)\right]\beta_{n,t} \\ &+ \left[\frac{1-\hat{\rho}_f^n}{1-\hat{\rho}_f} - \rho_f\frac{1-\hat{\rho}_f^{n-1}}{1-\hat{\rho}_f} - 1\right]f_t - \frac{1-\hat{\rho}_f^{n-1}}{1-\hat{\rho}_f}\varepsilon_{t+12} + u_{t+12}, \end{aligned} \tag{A.7}$$

where

$$u_{t+12} \equiv -\lambda\frac{(n-1)^2}{n}S_{n-1}(\hat{\rho}_\beta)\eta_{t+12}. \tag{A.8}$$

Since

$$\frac{1-\hat{\rho}_f^n}{1-\hat{\rho}_f} - 1 = \hat{\rho}_f\frac{1-\hat{\rho}_f^{n-1}}{1-\hat{\rho}_f},$$

the loading on f_t is

$$(\hat{\rho}_f - \rho_f)\frac{1-\hat{\rho}_f^{n-1}}{1-\hat{\rho}_f}.$$

Define

$$\Lambda_n(\hat{\rho}_\beta, \rho_\beta) = \hat{\rho}_\beta + \frac{(n-1)^2}{n}(\hat{\rho}_\beta - \rho_\beta)S_{n-1}(\hat{\rho}_\beta). \tag{A.9}$$

By the telescoping property of S_n then

$$xr_{n,t+12} = \lambda\Lambda_n(\hat{\rho}_\beta, \rho_\beta)\beta_{n,t} + (\hat{\rho}_f - \rho_f)\frac{1-\hat{\rho}_f^{n-1}}{1-\hat{\rho}_f}f_t - \frac{1-\hat{\rho}_f^{n-1}}{1-\hat{\rho}_f}\varepsilon_{t+12} + u_{t+12}, \tag{A.10}$$

as claimed. The properties for Λ_n follow from the definition (A.9), and $S_n(0) = 0$, and $S_n' > 0$.

For part b), note that under forecasters' beliefs, the expected $n-1$-period yield in 12 months is given by

$$\widetilde{\mathbb{E}}_t y_{n-1,t+12} = \frac{1-\hat{\rho}_f^{n-1}}{1-\hat{\rho}_f}\frac{\tilde{\rho}_f}{n-1}f_t + \lambda S_{n-1}(\hat{\rho}_\beta)\tilde{\rho}_\beta\frac{n-1}{n}\beta_{n,t}, \tag{A.11}$$

using the maturity-proportionality assumption $\beta_{n-1,t} = \frac{n-1}{n}\beta_{n,t}$. The rest of the proof is analogous to part i). This proves the Proposition. □

B. Empirical Appendix

B.1 Correlations with Macroeconomic Drivers

We now describe the construction of the inflation-output gap correlation and Fed Funds rate volatility, and how these macroeconomic and monetary policy risk proxies correlate with our benchmark bond–stock betas. We measure the quarterly output gap as $x_q =$

$100\,(\log Y_q - \log Y_q^*)$, where Y_q is real GDP (FRED: GDPC1) and Y_q^* is the Congressional Budget Office's estimate of potential real GDP (FRED: GDPPOT) (accessed Oct 2, 2026). We use two inflation measures. Because we use CPI inflation throughout the paper, we focus on the 12-month percent change in the CPI for all urban consumers (FRED: CPIAUCSL), constructed in the same way as in the rest of the paper. Following Campbell et al. (2020), we also use GDP deflator inflation, which is the annualized quarterly log change in the GDP price deflator (FRED: GDPDEF), $\pi_q = 400\,(\log P_q - \log P_{q-1})$. There is an ambiguity with respect to the timing of output and asset prices. We follow the beginning-of-quarter timing convention of Campbell et al. (2020) throughout the paper and align output reported by FRED for quarter q with inflation measured at the end of quarter $q-1$.

To illustrate the inverse relationship between inflation-output correlations and nominal bond–stock betas, we first report inflation-output gap correlations over the 1988q1-1999q4 and 2000q1-2019q4 subsamples. Panel A of Table B1 shows that the inflation-output gap correlation was negative from 1988 to 1999 (-0.24 for CPI inflation and -0.11 for GDP deflator inflation), and turned positive from 2000 to 2019 (0.47 and 0.49). The sign switch around 2000 coincides with the switch in the average bond-stock beta from positive (0.22) to negative (-0.15). After 2020, the beta rises toward zero (-0.05 on average) while the inflation-output gap correlation remains positive. The positive inflation–output gap correlation post-2020 might at first seem to contradict the generally inverse movements between bond–stock betas and inflation-output gap correlations. However, the post-2020 change in bond–stock betas has been concentrated in real or inflation-indexed bonds, while the inflation component of bond–stock betas has changed less (Campbell et al., 2026), consistent with the inflation-output gap correlation also remaining stable.

Table B1: Bond-Stock Beta and Macroeconomic Risks

Panel A: Inflation-Output Gap Correlation by Subsample

	Corr(Inflation, Output Gap)			
Sample	GDP Deflator	CPI	Obs.	Mean Beta
1988Q1–1999Q4	-0.11	-0.24	48	0.22
2000Q1–2019Q4	0.49	0.47	80	-0.15
2020Q1–2024Q1	0.39	0.47	17	-0.05

Panel B: Correlation of Bond–Stock Beta with 5-Year Rolling Macroeconomic Risks

Five-Year Rolling Macro Variable	Corr. with Beta	Obs.
Std. Δ Fed Funds rate	0.41	435
Corr(CPI inflation, output gap)	-0.23	145
Corr(GDP deflator inflation, output gap)	-0.24	145

Note: Panel A reports the correlation between inflation and the next quarter's output gap, $\text{Corr}\left(\pi_q, x_{q+1}\right)$, where $x_q = 100\left(\log Y_q - \log Y_q^*\right)$ is the quarterly log output gap based on the CBO's estimate of potential real GDP. GDP deflator inflation is the annualized quarterly log change in the GDP deflator, measured quarterly. CPI inflation is the 12-month percent change in the CPI at the quarter end and paired with the output gap of the following quarter. Mean Beta is the average five-year rolling bond-stock beta of the 11-year Treasury bond at quarter end. Panel B reports the correlation of the five-year rolling bond-stock beta of the 11-year Treasury bond with five-year rolling macroeconomic risks. The sample is January 1988 to March 2024.

Panel B of Table B1 shows correlations between our benchmark bond–stock beta for the 11-year Treasury bond with rolling measures of macroeconomic risks. We use the five-year rolling inflation–output gap correlations and the five-year rolling standard deviation of 12-month changes in the Federal Funds rate. Bond–stock betas are positively correlated with Federal Funds rate volatility (0.41), and negatively correlated with the five-year rolling inflation-output gap correlation (-0.23 for CPI inflation, -0.24 for GDP deflator inflation). Viewing the Federal Funds rate volatility as a simple proxy for monetary policy uncertainty, these correlations support the interpretation that bond–stock betas are linked to macroeconomic risks, in particular, the risk of stagflations and monetary policy uncertainty.

B.2 Additional Data Construction Details

This subsection provides additional details for the construction of pseudo expected excess returns used in Table V and the simple expectations models used in Table VI.

B.2.1 Inflation-Based Pseudo Expected Excess Returns

In Table V in the main paper we show that bond-stock betas are even more strongly related to pseudo expected excess returns, when the long-term yield forecast is replaced by a projection of yields onto long-term inflation expectations. We now detail the construction of these pseudo expected excess returns.

We proceed in two steps. First, let $\widetilde{\pi}_t^{10}$ denote the SPF forecast of the average CPI

inflation rate over the next ten years,

$$\widetilde{\pi}_t^{10} = \widetilde{E}_t \left[\frac{1}{10} \sum_{j=0}^{9} \pi_{t+12j} \right].$$

For each maturity n, we regress the Blue Chip forecast of the future $(n-1)$-year yield on $\widetilde{\pi}_t^{10}$ and take the fitted value as an inflation-implied yield forecast: $\widetilde{E}_t^{\pi} y_{n-1,t+12}^{par}$.

We use the same ten-year expected inflation forecast for all maturities, including the three- and six-year bonds, because this is the long-horizon inflation forecast with the longest data availability. The maturity-specific coefficients map it onto the level and variation of the corresponding bond yield forecast, which also makes the resulting measure comparable in magnitude to our baseline.

Second, we substitute this inflation-implied forecast for the survey yield forecast in Equation (7) to obtain the inflation-based pseudo expected excess return,

$$\widetilde{E}_t^{\pi} xr_{n,t\to t+12} = dur_{n,t} y_{n,t}^{par} - (dur_{n,t} - 1)\, \widetilde{E}_t^{\pi} y_{n-1,t+12}^{par} - y_{1,t}.$$

Through the fitted forecast, this measure varies over time only with the current bond yield, the short rate, and long-run expected inflation, but does not require a forecaster assessment of future term premia.

B.2.2 Simple Models of Future Long-Term Yields

Autoregressive (AR) Forecast. At each date, we assume forecasters predict future bond yields with a first-order autoregressive (AR(1)) model estimated in real time. For each of the two-, five-, and ten-year par yields—the maturities relevant for the three-, six-, and eleven-year bonds—we estimate

$$y_{m,s}^{par} = c_t + \rho_t\, y_{m,s-1}^{par} + \varepsilon_{m,s}$$

by OLS on monthly data over the trailing fifteen-year (180-month) window ending at date t. The implied twelve-month-ahead forecast of the $(n-1)$-year yield is

$$\widehat{E}_t^{AR} y_{n-1,t+12}^{par} = \mu_t + \rho_t^{12} \left(y_{n-1,t}^{par} - \mu_t \right), \qquad \mu_t = \frac{c_t}{1-\rho_t},$$

where μ_t is the long-run mean implied by the rolling AR(1). Because the regression uses only data available through date t, the forecast is formed in real time. The AR-based pseudo expected excess return replaces $\widetilde{E}_t y_{n-1,t+12}^{par}$ in Equation (7) with this forecast:

$$\widehat{E}_t^{AR} xr_{n,t\to t+12} = dur_{n,t}\, y_{n,t}^{par} - (dur_{n,t} - 1)\, \widehat{E}_t^{AR} y_{n-1,t+12}^{par} - y_{1,t}.$$

Moving-Average (MA) Forecast. As a second purely statistical benchmark, we forecast the future yield with a trailing moving average. The MA forecast of the $(n-1)$-year

yield is its average over the preceding w months,

$$\widehat{E}_t^{MA} y_{n-1,t+12}^{par} = \frac{1}{w} \sum_{j=0}^{w-1} y_{n-1,t-j}^{par}.$$

In the paper we use a sixty-month window ($w = 60$, i.e. five years). The MA-based pseudo expected excess return is constructed exactly as the AR one, replacing $\widetilde{E}_t y_{n-1,t+12}^{par}$ by $\widehat{E}_t^{MA} y_{n-1,t+12}^{par}$ in Equation (7).

Farmer–Nakamura–Steinsson (FNS) Forecast. We also construct pseudo-expected excess returns using forecasts based on Farmer et al. (2024) (FNS). We use their model with a structural break, which is used to construct Figure 9 in their paper and which the authors identify as their preferred specification for fitting the yield curve. We use the full term structure of forecasts for the 3-month T-bill to construct the average expected short-term rate over the life of the bond. We compute the model-implied yield for an n-year bond as

$$y_{n,t}^{FNS} = \frac{1}{4n} \sum_{h=0}^{4n-1} \mathbb{E}_t^{FNS} i_{t+h},$$

where n denotes maturity in years, i_{t+h} is the quarterly 3-month T-bill rate h quarters ahead, and $\mathbb{E}_t^{FNS}$ denotes expectations under the FNS model. Since the forecasts are quarterly, the expected short-term rate over the life of an n-year bond is computed as the average over the subsequent $4n$ quarters.

Applying the model one year ahead and using the law of iterated expectations, we compute the one-year-ahead expected $(n-1)$-year yield as:

$$E_t^{FNS} y_{n-1,t+12}^{FNS} = \frac{n\, y_{n,t}^{FNS} - y_{1,t}^{FNS}}{n-1}. \tag{A.12}$$

The measure entering Table VI in the main paper is computed entirely within the FNS model: we replace *both* the current and the forecasted yield in Equation (7) with their FNS counterparts. For the three-, six-, and eleven-year bonds, we use the FNS model-implied current yield $y_{n,t}^{FNS}$ and the FNS one-year-ahead forecast of the $(n-1)$-year yield, $E_t^{FNS} y_{n-1,t+12}^{FNS}$ (Equation (A.12)):

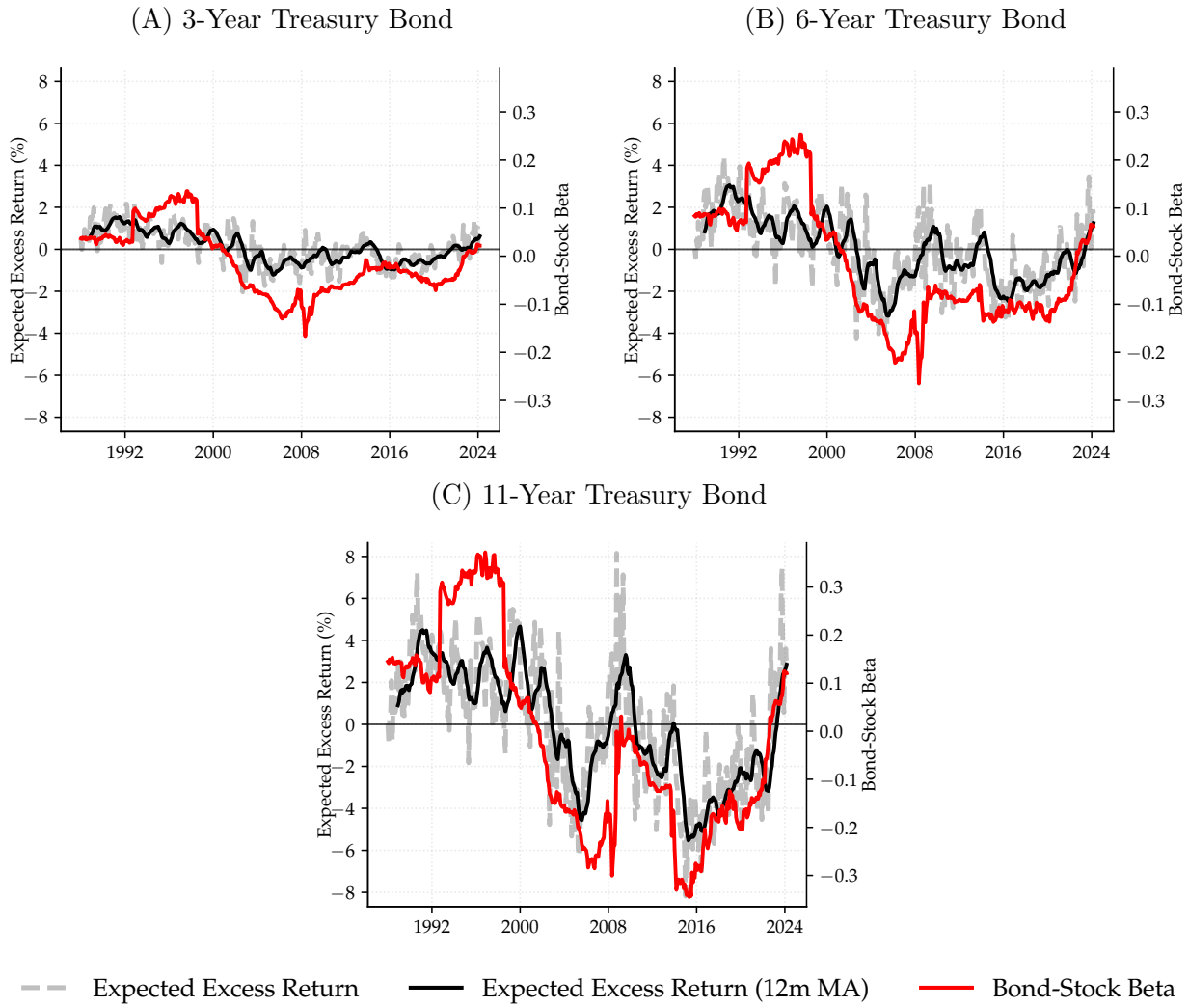
$$E_t^{FNS} xr_{n,t\to t+12} = dur_{n,t}\, y_{n,t}^{FNS} - (dur_{n,t} - 1)\, E_t^{FNS} y_{n-1,t+12}^{FNS} - y_{1,t}.$$

Because both the level and the forecast come from the FNS model, this isolates the expected excess return implied by the model, purged of the observed market yield. Finally, note that the term premium in their model is constant. We do not explicitly include it in our construction, since it would be absorbed by the intercept in our regressions and would not affect any of our coefficient estimates.

B.3 Robustness: Expected Excess Returns onto Betas

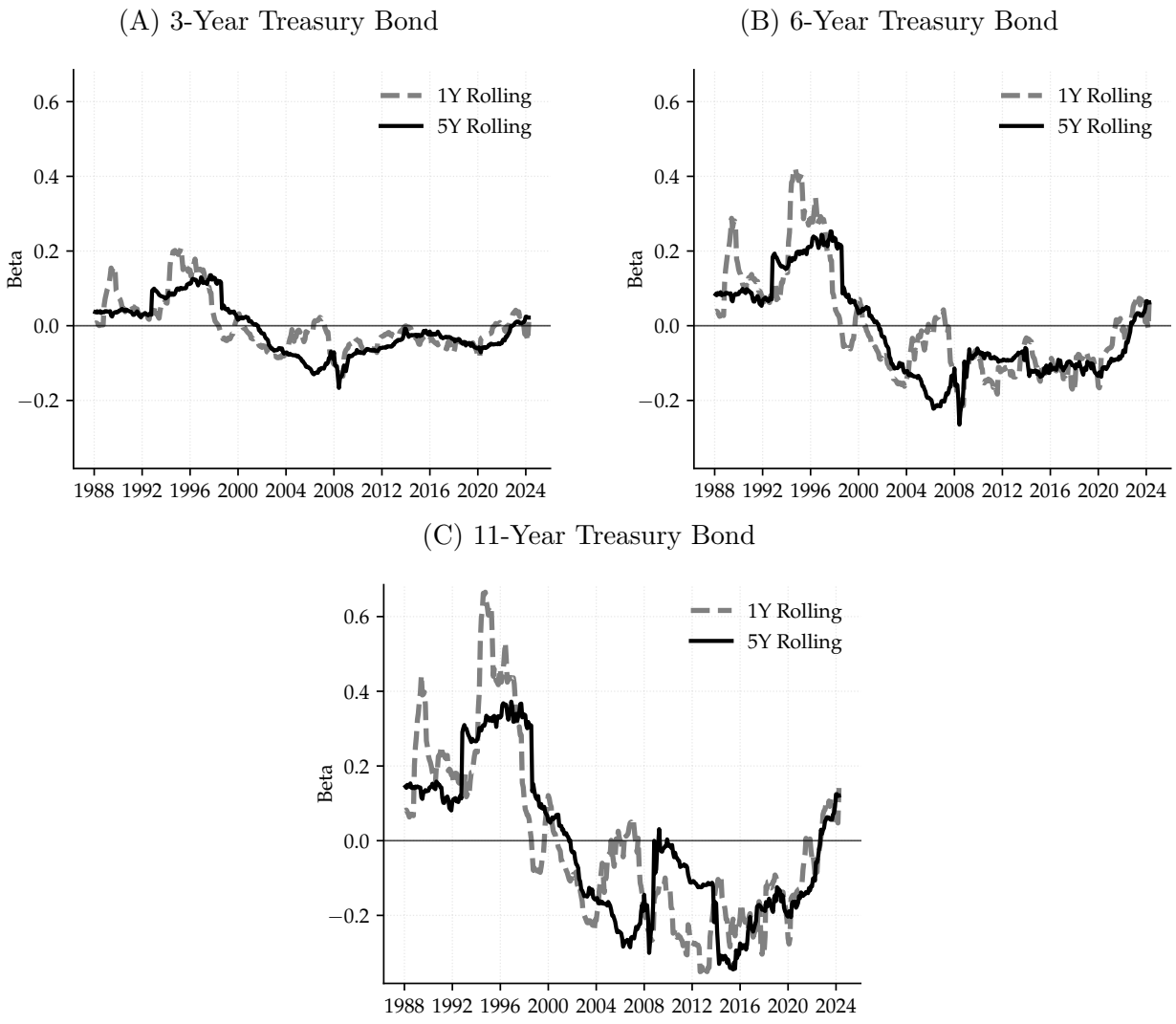
In this section, we provide additional robustness for our benchmark result in Table II. Figure B1 plots the time series of expected excess returns vs. five-year rolling bond-stock betas for all bond maturities, with Panel C corresponding to Panel A in Figure 1 in the main paper. We next show that our main result is robust to different ways of measuring bond–stock betas. Figure B2 compares our baseline bond–stock beta, computed from monthly returns over a five-year backward-looking rolling window, with betas computed from daily returns over one-year rolling windows. Figure B3 shows betas computed from daily returns, with and without the Dimson (1979) adjustment for non-synchronous trading, and betas that use exponentially decaying weights rather than a fixed rolling window. Tables B2 and B3 run the regressions of Equation 11 using alternative betas and show that the relationship between expected excess returns and bond–stock betas is robust. In particular, Table B2 shows that we can also use a forward-looking measure of bond–stock betas realized over the expected return horizon. Table B4 extends the time series regressions for each bond maturity individually in Table III to the one-year rolling betas, and shows that the coefficient on beta in the univariate regressions is similar for each maturity.

Figure B1: Expected Excess Returns: U.S. Treasury Bonds



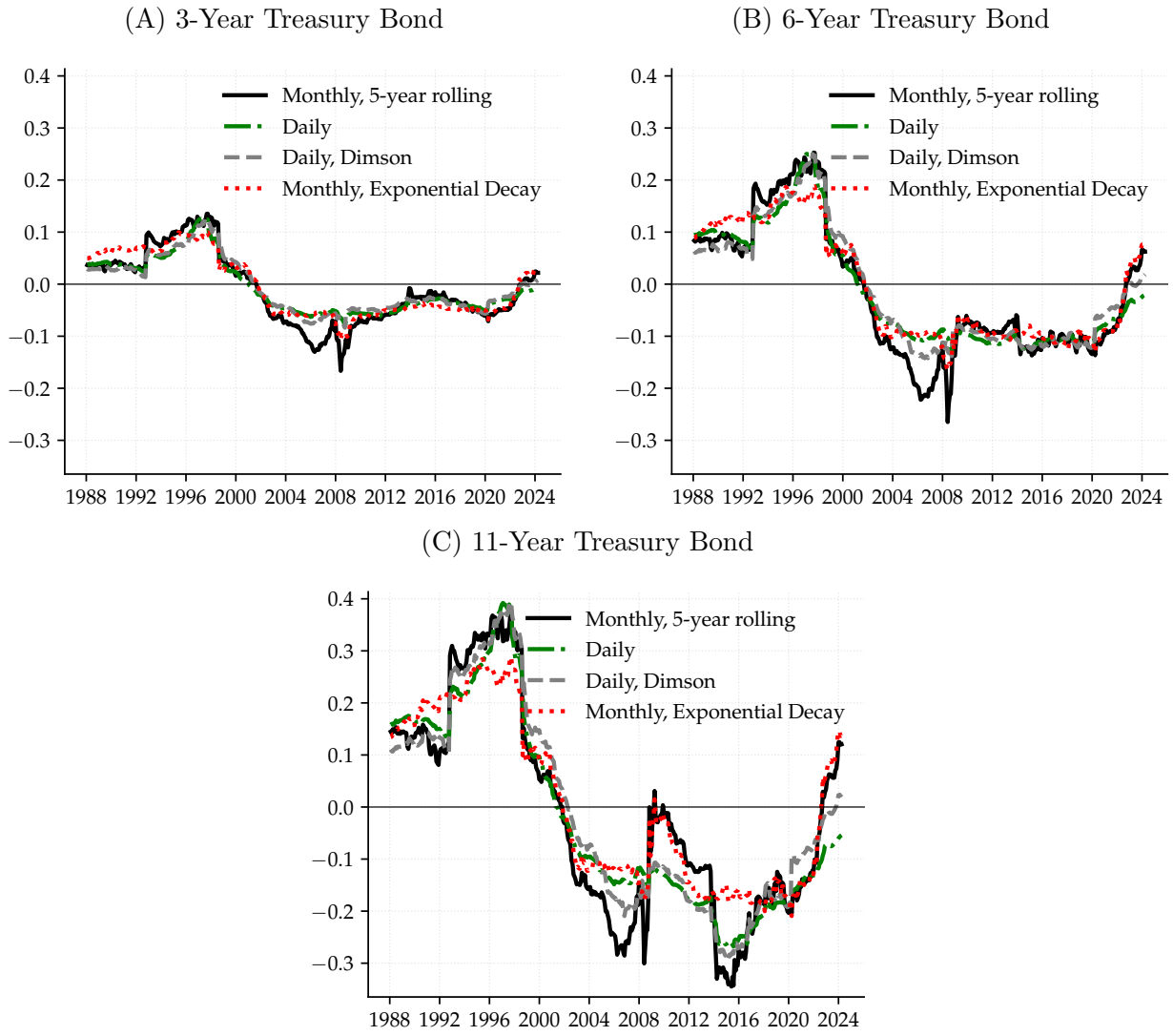
Note: This figure plots one-year expected excess returns for U.S. Treasury par bonds of three- (Panel A), 6- (Panel B), and 11-year (Panel C) maturities. Expected excess returns are computed according to equation (7). We use 4-quarter par bond yield forecasts from the Blue Chip Financial Forecasts. The dashed gray line shows monthly expected excess returns, and the black line shows the 12-month moving average. In red, the figure plots rolling estimates of stock market betas for the corresponding maturities. Betas are computed by regressing monthly par bond excess returns onto stock market excess returns over a backward-looking rolling window of 60 months according to Equation (10). We resample the data to a monthly frequency by taking the estimate from the last trading day of each month. The sample runs from January 1988 to March 2024.

Figure B2: Five-Year and One-Year rolling Betas, U.S.



Note: This figure plots rolling estimates of stock market betas for par Treasury bonds of three- (Panel A), 6- (Panel B), and 11-year (Panel C) maturities. Betas are computed by regressing par bond excess returns onto stock market excess returns over a backward-looking rolling window according to Equation (10). The dashed gray line uses 12-month rolling windows of daily returns, and the black line uses 60-month rolling windows of monthly returns. We resample the data to a monthly frequency by taking the estimate from the last trading day of each month. The sample runs from January 1988 to March 2024.

Figure B3: Alternative Betas, U.S.



Note: This figure plots alternative estimates of stock market betas for par Treasury bonds of three-year (Panel A), six-year (Panel B), and eleven-year (Panel C) maturities. Our baseline estimate (black line) uses five-year rolling regressions of monthly par bond excess returns on stock market excess returns. The dashed green line reports betas estimated using a five-year rolling regressions of daily par bond excess returns on stock market excess returns. The dash gray lines also uses daily data with the Dimson (1979) adjustment, which includes two lags and two leads of the market return. The red dash–dot line reports betas estimated from five-year rolling regressions using exponentially decaying weights with a half-life of three years. Par bond excess returns are computed from Gürkaynak et al. (2007) nominal Treasury bond par yields. Stock market excess returns are from Ken French's website. We resample the estimates to a monthly frequency by taking the estimate from the last trading day of each month. The sample runs from January 1988 to March 2024.

Table B2: Expected Excess Returns and One-Year Rolling Bond-Stock Betas

Dependent Variable: Expected Excess Returns										
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Bond-Stock Beta 1YR	7.14***	7.13***	4.20**	7.42***	7.79***	5.74***	4.20***	6.12***	7.78***	
	(1.55)	(1.53)	(1.86)	(1.56)	(1.67)	(1.65)	(1.40)	(1.30)	(1.71)	
Bond-Stock Beta 1YR (forward)										6.06***
										(1.36)
Bond Variance				2.09**						
				(0.77)						
Term Spread					0.47*					
					(0.22)					
CP Factor						0.21*				
						(0.10)				
Level							0.81***			
							(0.19)			
Slope							-0.33*			
							(0.17)			
Curvature							0.07			
							(0.16)			
r-star								0.44*		
								(0.24)		
Constant	0.02									
	(0.25)									
Maturity FE		✓	✓	✓	✓	✓	✓	✓	✓	✓
Time Trend × Maturity FE			✓							
Time FE									✓	
R-squared	0.29	0.29	0.37	0.34	0.33	0.33	0.41	0.32	0.82	0.21
Observations	1305	1305	1305	1305	1305	1305	1305	1305	1305	1305

Note: The table reports the estimates for the regression of expected excess returns on bond-stock beta as in equation (11):

$$\tilde{E}_t xr_{n,t\to t+12} = \alpha + \lambda \hat{\beta}^w_{n,t} + X'_{n,t}\delta + \eta_{n,t}.$$

We use the 1-year rolling bond-stock beta estimated using a regression of daily par bond excess returns on daily equity excess returns. The first 9 columns mirror Table II. Column (10) uses the beta estimated on daily data from time t to time $t+12$. The estimate is therefore forward-looking rather than backward-looking. Driscoll–Kraay standard errors with 27 lags are reported in parentheses. Stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample consists of monthly data from January 1988 to March 2024.

Table B3: Expected Excess Returns and Stock Market Betas, Alternative Betas

Dependent Variable: Expected Excess Returns						
	(1)	(2)	(3)	(4)	(5)	(6)
Beta 5YR (Monthly)	9.98***					
	(1.61)					
Beta 1YR		7.13***				
		(1.53)				
Beta 5YR (Dimson)			10.58***			
			(1.80)			
Beta 1YR (Dimson)				7.13***		
				(1.23)		
Beta 5YR (Daily)					10.87***	
					(1.84)	
Beta HL 3Y (Monthly)						12.48***
						(1.67)
Constant	0.20*	0.09	0.15	0.09	0.17*	0.16**
	(0.10)	(0.11)	(0.09)	(0.10)	(0.09)	(0.07)
Maturity FE	✓	✓	✓	✓	✓	✓
R-squared	0.44	0.29	0.42	0.34	0.43	0.48
Observations	1305	1305	1305	1305	1305	1305

Note: The table reports regression estimates of Equation (11) using alternative measures of beta:

$$\tilde{E}_t xr_{n,t\to t+12} = \alpha_n + \lambda\beta_{n,t} + \eta_{n,t},$$

where the dependent variable is the expected excess return, $\beta_{n,t}$ is the beta measure of the column, and α_n are maturity fixed effects, included in every column. Column (1) is the baseline five-year rolling beta estimated from monthly returns. Column (2) uses a beta estimated from daily returns over a one-year rolling window. Columns (3) and (4) use Dimson betas estimated from daily returns over five-year and one-year rolling windows, respectively, which account for non-synchronous trading by including two leads and two lags of the market return and summing the associated coefficients. Column (5) uses a beta estimated from daily returns over a five-year rolling window. Column (6) uses a beta constructed with exponentially decaying weights with a half-life of three years, estimated using monthly data. Driscoll–Kraay standard errors with 27 lags are reported in parentheses; stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample consists of monthly observations from January 1988 to March 2024.

Table B4: Expected Excess Returns and Stock Market Betas: U.S., Individual Maturities, One-Year Rolling Beta

	3-year			6-year			11-year		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Beta 1YR	7.07***	5.05***	4.12***	6.51***	4.08**	2.85*	7.37***	4.15*	1.76
	(1.30)	(1.42)	(1.44)	(1.49)	(1.63)	(1.59)	(1.60)	(2.09)	(1.72)
Level			0.37***			0.82***			1.80***
			(0.10)			(0.24)			(0.51)
Slope			-0.08			-0.30			-0.45
			(0.08)			(0.17)			(0.31)
Curvature			0.11			0.14			-0.05
			(0.07)			(0.15)			(0.30)
Constant	0.09	0.55**	0.09	-0.01	0.97*	-0.02	-0.01	1.96*	-0.04
	(0.11)	(0.23)	(0.09)	(0.24)	(0.49)	(0.20)	(0.44)	(0.94)	(0.35)
Time Trend		✓			✓			✓	
R-squared	0.34	0.41	0.50	0.27	0.34	0.44	0.30	0.37	0.47
Observations	435	435	435	435	435	435	435	435	435

Note: The table reports the regression estimates for equation (11) estimated maturity by maturity:

$$\tilde{E}_t xr_{n,t\to t+12} = \alpha + \lambda \hat{\beta}^w_{n,t} + X'_{n,t}\delta + \eta_{n,t}.$$

The bond–stock beta $\hat{\beta}^w_{n,t}$ is estimated from daily bond and stock excess returns over a one-year rolling window. Columns (2), (5), and (8) include a time trend; Columns (3), (6), and (9) control for the first three principal components of the yield curve (level, slope, and curvature), each standardized to have mean zero and unit standard deviation. Newey–West standard errors with 27 lags are reported in parentheses. Stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample consists of monthly data from January 1988 to March 2024.

B.4 Robustness: Expected Excess Return Decomposition

Table B5 provides robustness for Table IV in the main paper. One might wonder whether the decomposition is sensitive to using the 10-year forecast of the T-bill rate from the SPF. Here, we show that even just using raw yields, the actual market yield is significantly more strongly related to bond-stock betas than the yield forecast (both duration-weighted). The magnitudes of the coefficients in columns (3) through (6) are of course different, because we now multiply by bond duration. We again see that duration-weighted market yields load more strongly onto bond-stock betas than duration-weighted forecasted yields. Hence, market yields move with bond-stock betas, which is why expected excess returns are positively related to bond-stock betas.

Table B5: Decomposition of Risk Premium

Dependent Variable:	$\tilde{E}_t xr_{n,t\to t+12}$		$dur_{n,t}\, y_{n,t}$		$(dur_{n,t}-1)\,\tilde{E}_t y_{n-1,t+12}$		$y_{1,t}$	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Beta 5YR	9.98***	8.00***	58.60***	18.29	37.78***	4.64	10.83***	5.65
	(1.61)	(2.06)	(11.49)	(11.36)	(8.46)	(6.91)	(2.45)	(3.38)
Maturity FE	✓	✓	✓	✓	✓	✓	✓	✓
Time trend × Maturity FE		✓		✓		✓		✓
R-squared	0.44	0.47	0.76	0.91	0.81	0.94	0.36	0.55
Observations	1305	1305	1305	1305	1305	1305	1305	1305

Note: The table decomposes the loading of expected excess returns on the bond–stock beta into the loadings of its three components. By construction (Equation 7), the expected excess return is

$$\tilde{E}_t xr_{n,t\to t+12} = dur_{n,t}\, y_{n,t} - (dur_{n,t}-1)\,\tilde{E}_t y_{n-1,t+12} - y_{1,t},$$

the current yield on the n-year bond scaled by its duration, minus the survey forecast of the $(n-1)$-year yield one year ahead scaled by duration less one, minus the one-year yield. Each pair of columns regresses one of these objects on the five-year rolling bond–stock beta,

$$Y_{n,t} = \alpha_n + \lambda\beta_{n,t} + X'_{n,t}\delta + \varepsilon_{n,t},$$

for $Y_{n,t}$ equal to the expected excess return $\tilde{E}_t xr_{n,t\to t+12}$ (Columns 1–2), the duration-scaled current yield $dur_{n,t}\, y_{n,t}$ (Columns 3–4), the duration-scaled yield forecast $(dur_{n,t}-1)\,\tilde{E}_t y_{n-1,t+12}$ (Columns 5–6), and the one-year yield $y_{1,t}$ (Columns 7–8), where α_n are maturity fixed effects, included in every column. The odd columns have no controls ($X_{n,t}$ is empty); in the even columns $X_{n,t}$ is a linear time trend, one for each maturity. Because the identity is linear and every column uses the same sample and controls, the beta coefficient in Columns (1)–(2) equals the coefficient in Columns (3)–(4) minus those in Columns (5)–(6) and (7)–(8). Unlike Table IV in the main text, yields are not expressed relative to the SPF forecast of the average T-bill rate, so the current yield and the yield forecast enter in levels. The panel pools the three-, six-, and eleven-year bonds. Driscoll–Kraay standard errors with 27 lags are reported in parentheses; stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample consists of monthly observations from January 1988 to March 2024.

B.5 Forecaster Groups

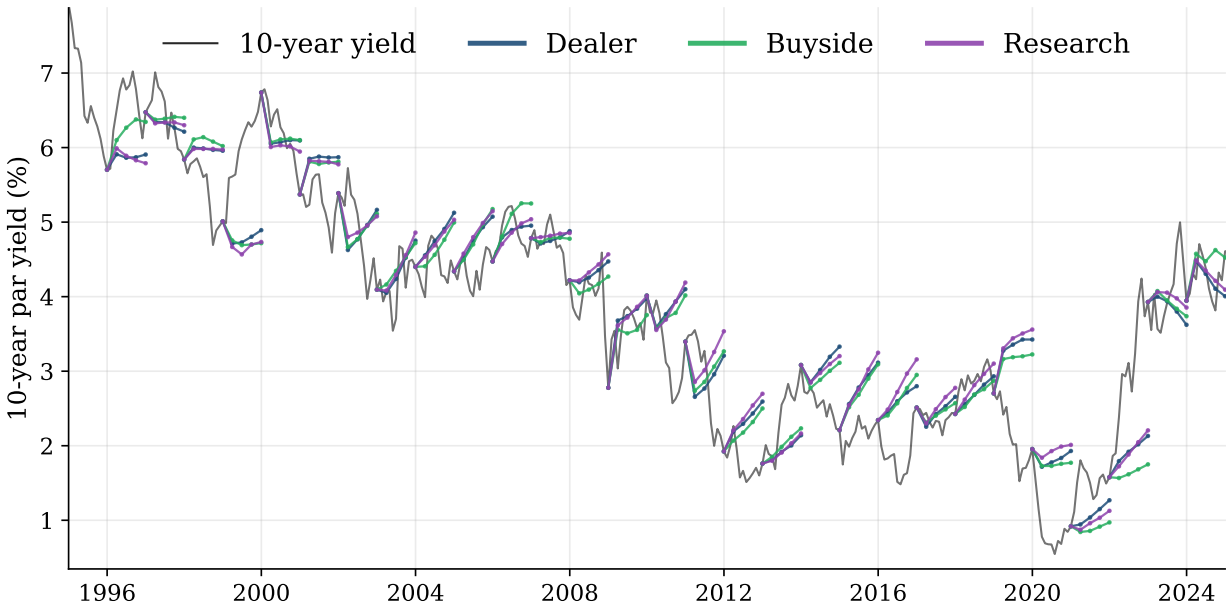
We now document the correlation between expected excess returns and long-term yield expectations across different types of forecasters. We split forecasters into dealers (e.g. JP Morgan), buy side (e.g. Swiss Re) and research (e.g. Oxford Economics). The forecaster institution name is available starting in 1992, so we can do to this split from 1995 onwards. Table B6 provides the correlations for expected excess returns across forecaster groups, showing that these correlations are high. Figure B4 displays the source of this high correlation across forecasters, showing the forecasted yield path as a hair plot against actual 10-year yields for different forecaster groups. It is notable that all three forecaster groups predicted a declining yield path before 2000 (i.e. positive expected excess returns) but increasing yield paths after 2000 (i.e. all three groups implied negative excess returns). Overall, our main result is not specific to buy or sell side forecasters, but robust for different forecaster groups.

Table B6: Correlations of Expected Excess Returns Across Forecaster Groups

	3Y			6Y			11Y		
	Dealer	Buy-side	Research	Dealer	Buy-side	Research	Dealer	Buy-side	Research
Dealer	1.00	0.89	0.92	1.00	0.87	0.92	1.00	0.89	0.92
Buy-side	0.89	1.00	0.88	0.87	1.00	0.87	0.89	1.00	0.90
Research	0.92	0.88	1.00	0.92	0.87	1.00	0.92	0.90	1.00

Note: The table reports correlations in expected bond excess returns across different groups of forecasters. Using information from Blue Chip on the firm associated with each forecaster, we classify forecasters into three groups: dealers, buy-side firms, and research institutes. Based on their forecasts, we compute group-level expected excess returns as in Equation (7).

Figure B4: 10-Year Bond Yield Hair Plot by Forecaster Type



Note: The figure plots forecasts of the 10-year yield, from one-quarter-ahead to five-quarters-ahead horizons, by forecaster group. Using information from Blue Chip on the firm associated with each forecaster, we classify forecasters into three groups: dealers, buy-side firms, and research institutes. Based on their individual forecasts, we compute group-level forecasts of the 10-year yield. The figure reports forecasts made in December of each year in the sample for which individual forecasts are available. The sample runs from 1995 to 2024.

B.6 Realized vs Expected Excess Returns

In Table B7, we study the relationship between realized excess returns and expected excess returns. We find that expected excess returns predict future realized excess returns with a coefficient of 0.55. We also use the smoother version of expected excess returns and find a coefficient of 0.80. These results show that expected excess returns contain information about subsequent realized excess returns.

Table B7: Realized and Expected Excess Returns

Dependent Variable: Future Realized Excess Return				
	(1)	(2)	(3)	(4)
Expected Excess Returns	0.54**	0.55**		
	(0.23)	(0.23)		
Expected Excess Returns (MA)			0.78**	0.80**
			(0.30)	(0.28)
Constant	1.97**		2.00**	
	(0.73)		(0.73)	
Maturity FE		✓		✓
R-squared	0.05	0.07	0.07	0.09
Observations	1305	1305	1272	1272

Note: The table reports the estimates of equation:

$$xr_{n,t\to t+12} = b_0 + b_1\, \tilde{E}xr_{n,t\to t+12} + X'_{n,t}\delta + \varepsilon_{n,t+12},$$

where $xr_{n,t\to t+12}$ is the realized excess bond return from time t to time $t+12$ and $\tilde{E}xr_{n,t\to t+12}$ are the expected excess returns for the corresponding time period (t to $t+12$), measured at time t. Columns (1) and (2) use raw expected excess returns while Columns (3) and (4) use the smoother twelve month moving average. Driscoll–Kraay standard errors with 27 lags are reported in parentheses. Stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample consists of monthly data with forecast dates t from January 1988 to March 2024, so that realized returns run through March 2025; Columns (3) and (4) start in December 1988, when the 12-month moving average is first available.

B.7 Convenience Yield

In this appendix, we examine the potential role of the convenience yield in explaining part of our results. The convenience yield is the premium that Treasury securities command over otherwise similar securities because of their liquidity and safety attributes. Krishnamurthy and Vissing-Jorgensen (2012), for example, show that Treasury yields are lower than yields on Aaa corporate bonds and attribute this spread to the convenience yield of U.S. Treasuries. More recently, Acharya and Laarits (2025) document that the Treasury convenience yield tends to be lower when the covariance of Treasury returns with aggregate stock market returns is higher. Building on this insight, it is natural to quantify the extent to which variation in the convenience yield contributes to the relationship we document between expected excess returns and the bond–stock beta.

To assess the potential role of convenience yield in explaining our results, we compute the expected excess returns on Aaa corporate bonds. The BCFF provides forecasts of Aaa corporate bond yields throughout our sample, allowing us to construct a measure of expected excess returns for Aaa corporate bonds. For the observed Aaa corporate bond yield, y_t^{AAA}, we use Moody's Seasoned Aaa Corporate Bond Yield from FRED (ticker `DAAA`), averaged over the last week of each month, as we do for Treasury par yields. For the expected Aaa corporate bond yield, $\tilde{E}_t y_{t+12}^{AAA}$, we use the four-quarter-ahead BCFF forecast, averaged across forecasters and timed consistently with the Treasury yield forecasts in Equation (7).

We can then compute the expected excess returns on Aaa corporate bonds by replacing the Treasury yields in Equation (7) with the Aaa corporate bond yields:

$$\tilde{E}_t xr_{n,t\to t+12}^{AAA} = dur_{n,t}\, y_t^{AAA} - (dur_{n,t} - 1)\, \tilde{E}_t y_{t+12}^{AAA} - y_{1,t}. \tag{A.13}$$

We compare the Aaa expected excess returns with the expected excess returns $\tilde{E}_t xr_{n,t\to t+12}$ on the 11-year Treasury ($n = 11$), and use the duration of the 11-year Treasury, $dur_{11,t}$, for both bonds for comparability. Using a different duration for the Aaa corporate bond yield would effectively scale up the expected Aaa corporate bond excess return by a constant and thereby leave correlations with our baseline variables unchanged. The two expected excess returns are plotted in Figure B5. The figure shows that the expected excess returns on the two assets are highly correlated, with a sample correlation of 0.89.

Using the current Aaa corporate bond yield and its expected value, we can rewrite the expected excess return on the U.S. Treasury in Equation (7) as

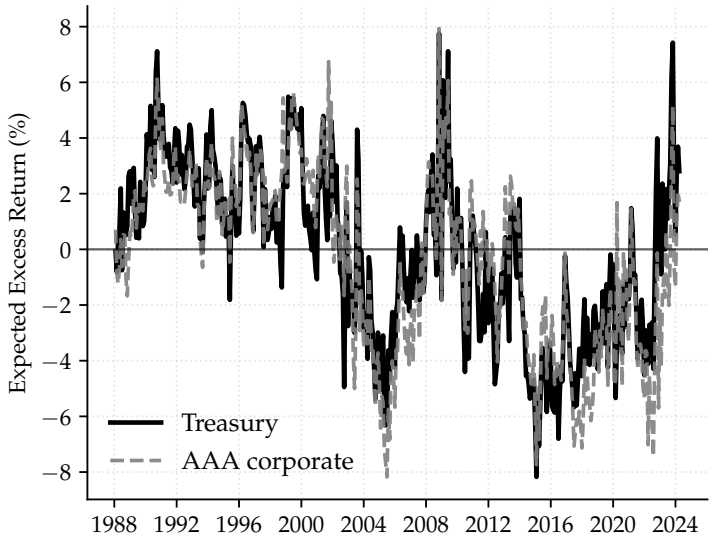
$$\begin{aligned}\tilde{E}_t xr_{n,t\to t+12} = &\underbrace{dur_{n,t}\, y_t^{AAA} - (dur_{n,t}-1)\,\tilde{E}_t y_{t+12}^{AAA} - y_{1,t}}_{\tilde{E}_t xr^{AAA}_{n,t\to t+12}} \\ &- \underbrace{\left[dur_{n,t}\left(y_t^{AAA} - y_{n,t}\right) - (dur_{n,t}-1)\left(\tilde{E}_t y_{t+12}^{AAA} - \tilde{E}_t y_{n-1,t+12}\right)\right]}_{\tilde{E}_t xr^{CY}_{n,t\to t+12}}.\end{aligned} \tag{A.14}$$

As the equation shows, the expected excess return on a Treasury bond can be decomposed into the expected excess return on an Aaa corporate bond, $\tilde{E}_t xr^{AAA}_{n,t\to t+12}$, minus the expected return on a strategy that is long the Aaa corporate bond and short the U.S. Treasury bond, $\tilde{E}_t xr^{CY}_{n,t\to t+12}$. Intuitively, U.S. Treasuries offer a lower yield than otherwise comparable Aaa corporate bonds because of their convenience value. The expected return on the long-Aaa, short-Treasury strategy therefore captures the return that investors forgo in exchange for holding the U.S. Treasury due to its convenience yield.

Table B8 regresses each of these components on our bond–stock beta for the 11-year U.S. Treasury bond. Column (1) reports the loading of the expected excess return on the 11-year Treasury bond on the bond–stock beta, 10.79. In Column (3), we regress the expected excess return on Aaa corporate bonds on the bond–stock beta and obtain a coefficient of 10.73, which is extremely close to the coefficient for Treasury expected excess returns in Column (1). Column (5) reports the results from regressing the expected excess return on the long-Aaa, short-Treasury strategy on the bond–stock beta. The estimated coefficient is -0.06 and is not statistically significant. Because Equation (A.14) is an identity and all columns use the same sample, the difference between the coefficients in Columns (3) and (1) is exactly the coefficient reported in Column (5). A similar relation holds when we include a time trend in Columns (2), (4), and (6).

Overall, the results show the loadings of expected excess returns on Aaa corporate bonds and U.S. Treasuries are virtually identical, indicating that the bulk of the variation we document is not driven by variation in the convenience yield.

Figure B5: Expected Excess Returns: U.S. Treasuries vs Aaa Corporate Bonds



Note: The figure plots the expected excess returns $\tilde{E}_t xr_{n,t\to t+12}$ on the 11-year U.S. Treasury (solid line), constructed as in Equation (7), and $\tilde{E}_t xr^{AAA}_{n,t\to t+12}$ on Aaa corporate bonds (dashed line), constructed as in Equation (A.13), with $n = 11$. The expected yields on the U.S. Treasury and on Aaa corporate bonds are the BCFF forecasts. The observed Aaa corporate bond yield is the Moody's seasoned Aaa corporate bond yield. Both expected excess returns use the duration of the 11-year Treasury for comparability. The sample consists of monthly data from January 1988 to March 2024.

Table B8: Convenience Yield and Bond–Stock Beta

Dependent Variable:	$\tilde{E}_t xr_{n,t\to t+12}$		$\tilde{E}_t xr^{AAA}_{n,t\to t+12}$		$\tilde{E}_t xr^{CY}_{n,t\to t+12}$	
	(1)	(2)	(3)	(4)	(5)	(6)
Beta 5YR	10.79***	8.78***	10.73***	7.66***	-0.06	-1.12
	(1.70)	(2.30)	(1.50)	(2.00)	(0.90)	(1.82)
Time trend		✓		✓		✓
R-squared	0.47	0.49	0.44	0.48	0.00	0.02
Observations	435	435	435	435	435	435

Note: The table decomposes the loading of expected excess returns on the bond–stock beta into the loadings of the expected excess return on Aaa corporate bonds and of the expected return on a strategy that is long the Aaa corporate bond and short the U.S. Treasury. By construction (Equation (A.14)),

$$\tilde{E}_t xr_{n,t\to t+12} = \tilde{E}_t xr^{AAA}_{n,t\to t+12} - \tilde{E}_t xr^{CY}_{n,t\to t+12},$$

where $\tilde{E}_t xr^{AAA}_{n,t\to t+12} = dur_{n,t}\, y_t^{AAA} - (dur_{n,t} - 1)\, \tilde{E}_t y^{AAA}_{t+12} - y_{1,t}$ and $\tilde{E}_t xr^{CY}_{n,t\to t+12} = dur_{n,t}(y_t^{AAA} - y_{n,t}) - (dur_{n,t} - 1)(\tilde{E}_t y^{AAA}_{t+12} - \tilde{E}_t y_{n-1,t+12})$. Each pair of columns regresses one of these objects on the five-year rolling bond–stock beta,

$$Y_{n,t} = \alpha + \lambda \beta_{n,t} + X'_{n,t}\delta + \varepsilon_{n,t}.$$

All columns use the 11-year Treasury bond ($n = 11$). The odd columns have no controls ($X_{n,t}$ is empty). In the even columns $X_{n,t}$ is a linear time trend. Newey–West standard errors with 27 lags are reported in parentheses. Asterisks denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The sample consists of monthly observations from January 1988 to March 2024.

B.8 Additional Results: International Panel

Here, we provide summary statistics for the international panel (Table B9). The relationship between expected excess returns and bond-stock betas is robust if we measure betas with respect to the U.S. stock market rather than local stock market excess returns. Table B10 shows the analogous regressions to Table X using bond betas onto U.S. stock returns. Figure B6 is analogous to Figure 4 in the main paper, but uses a measure of betas of bond excess returns with respect to U.S. stock excess returns. We find that results are similar for either version of bond-stock betas.

Table B9: Summary Statistics: International Panel

Panel (A): Expected Excess Returns

	Pooled	Canada	France	Germany	Italy	Japan	Netherlands	Norway	Spain	Sweden	Switzerland	U.K.	U.S.
Average	-0.76	-1.26	-0.34	-1.13	1.53	-0.67	-1.28	-1.87	-0.10	-2.17	-1.92	-0.57	-0.06
Std	2.73	2.94	2.22	1.87	3.44	1.44	2.36	2.31	3.10	2.37	1.70	2.62	3.16
Min	-9.73	-8.50	-4.80	-5.71	-4.87	-4.17	-6.96	-9.24	-5.32	-9.73	-5.90	-7.22	-8.18
Max	13.03	9.37	6.63	5.68	11.80	4.44	8.00	4.15	13.03	5.44	2.42	9.06	8.16
N	4690	424	424	424	424	424	360	320	361	361	320	424	424
First Year	1989	1989	1989	1989	1989	1989	1994	1998	1994	1994	1998	1989	1989

Panel (B): Expected Excess Returns (Moving Average)

Average	-0.75	-1.27	-0.35	-1.14	1.56	-0.64	-1.26	-1.92	-0.07	-2.10	-1.88	-0.56	-0.07
Std	2.37	2.63	1.82	1.51	3.04	1.09	1.92	1.71	2.73	1.97	1.23	2.27	2.73
Min	-6.49	-6.49	-3.60	-3.87	-2.81	-2.75	-5.30	-6.20	-4.19	-6.21	-4.44	-4.33	-5.53
Max	8.83	6.53	4.60	3.02	8.83	2.10	2.97	2.10	8.81	3.53	0.97	5.19	4.67

Panel (C): Realized Excess Returns

Average	2.11	2.10	2.02	1.77	2.64	1.76	2.19	1.10	2.83	2.58	1.67	1.67	2.86
Std	7.06	6.37	6.86	6.59	9.80	4.84	6.70	5.92	8.56	7.40	5.15	7.21	7.30
Min	-29.99	-15.75	-25.17	-24.22	-29.99	-21.48	-25.66	-16.51	-25.61	-26.94	-15.61	-25.52	-21.36
Max	24.44	15.62	16.77	15.24	24.44	14.92	14.83	13.82	21.90	18.16	12.78	16.42	17.60

Panel (D): Bond-Stock Betas

Average	0.02	0.04	0.04	-0.01	0.11	-0.03	-0.00	-0.07	0.10	-0.00	-0.03	0.06	-0.02
Std	0.15	0.18	0.13	0.10	0.12	0.08	0.12	0.09	0.14	0.13	0.09	0.18	0.20
Min	-0.34	-0.23	-0.16	-0.19	-0.10	-0.16	-0.20	-0.22	-0.12	-0.26	-0.19	-0.17	-0.34
Max	0.49	0.46	0.29	0.18	0.37	0.21	0.24	0.22	0.46	0.27	0.21	0.49	0.37
Average S.E.	0.05	0.07	0.05	0.04	0.05	0.04	0.05	0.05	0.05	0.05	0.05	0.07	0.08

Panel (E): β w.r.t. US Stock Returns

Average	0.01	0.04	0.02	-0.01	0.09	0.02	-0.02	-0.03	0.09	-0.01	-0.04	0.02	-0.02
Std	0.15	0.17	0.13	0.13	0.12	0.09	0.13	0.13	0.16	0.18	0.09	0.16	0.20
Min	-0.34	-0.22	-0.22	-0.26	-0.17	-0.12	-0.28	-0.27	-0.18	-0.33	-0.21	-0.23	-0.34
Max	0.55	0.49	0.28	0.25	0.39	0.31	0.24	0.29	0.55	0.45	0.19	0.36	0.37
Average S.E.	0.06	0.07	0.07	0.06	0.08	0.05	0.06	0.06	0.07	0.07	0.05	0.08	0.08

Note: The table reports summary statistics for expected excess returns (Panel A), the one-year moving average of expected excess returns (Panel B), one-year holding period realized excess returns (Panel C), bond-stock betas using local equity (Panel D) and bond-stock betas using US equity (Panel E) for different countries. Expected bond excess returns are computed according to equation 7. We use 12-month par bond yield forecasts from Consensus Economics. The only exception is the U.S., where we use our baseline expected excess returns based on the BCFF. Bond yields, T-bill rates, and stock returns are obtained from GFD. Bond–stock betas are computed as the regression coefficient of monthly par bond excess returns on stock excess returns over a backward-looking rolling window, as defined in Equation 10. We use 60-month rolling windows (five-year rolling β). Par bond excess returns are constructed using ten-year par yields, and stock market excess returns are computed using domestic stock market index returns. The table reports the average estimate, standard deviation, minimum and maximum values, and, for the betas in Panels D and E, the average heteroskedasticity-robust standard errors. Panel A also reports the number of observations and the first year for which forecasts are available. The sample is monthly and spans September 1989 to December 2024; all panels use the months for which an expected excess return is available.

Table B10: Excess Returns and Bond-Stock (U.S.) Market Betas: International Panel

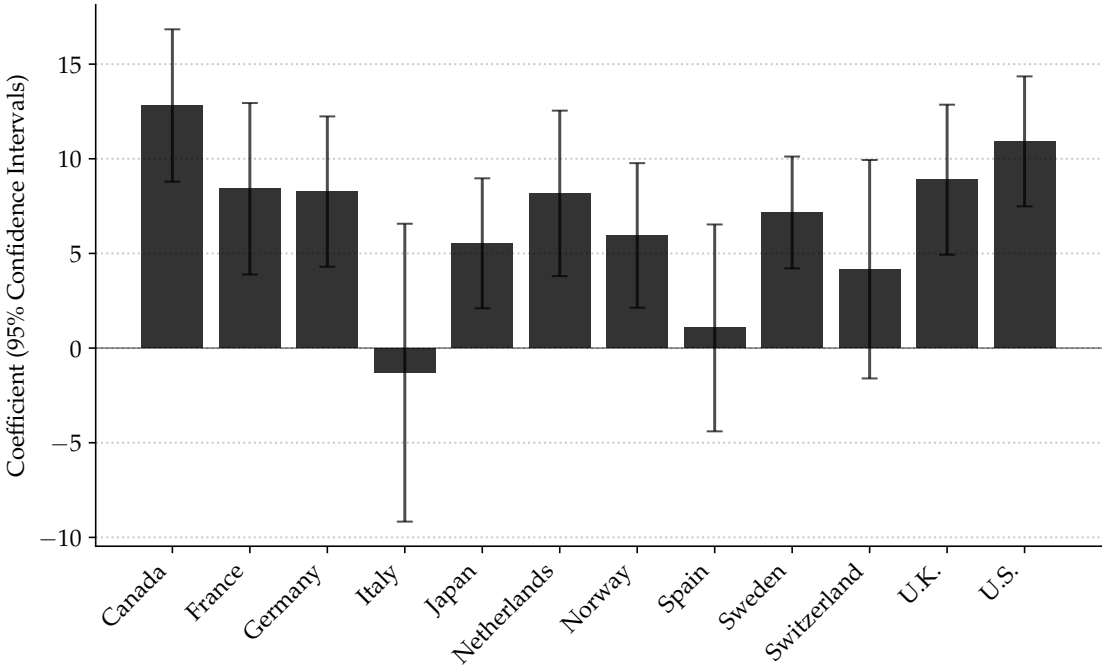
Dependent Variable: Expected Excess Returns								
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Bond-Stock Beta (U.S.)	8.36***	7.64***	6.08***	6.56***	7.68***	7.28***	6.25**	3.82**
	(1.68)	(1.53)	(1.98)	(1.44)	(1.54)	(1.53)	(2.38)	(1.52)
Bond Variance				1.38				
				(0.93)				
Term Spread					0.21			
					(0.16)			
r-star						0.16		
						(0.20)		
Constant	-0.88***							
	(0.25)							
Country FE		✓	✓	✓	✓	✓		✓
Time Trend × Country FE			✓					
Time FE							✓	✓
R-squared	0.21	0.29	0.37	0.30	0.30	0.30	0.55	0.62
Observations	4690	4690	4690	4690	4690	4690	4690	4690

Note: The table reports the results of Equation (15),

$$\tilde{E}_t xr_{c,t\to t+12} = \alpha + \lambda \beta_{c,t}^{US} + X_{c,t}'\delta + \eta_{c,t},$$

estimated using our international panel. It is analogous to Table X in the main text, but the bond–stock beta $\beta_{c,t}^{US}$ is estimated with respect to the U.S. stock market rather than the domestic one: it is the slope of country c's monthly bond excess return on the U.S. stock market excess return over a five-year (60-month) rolling window. The dependent variable is the expected one-year bond excess return, constructed from Consensus Economics twelve-month-ahead forecasts of the ten-year government bond yield. The only exception is the U.S., where we use our baseline expected excess returns based on the BCFF. Column (1) is the pooled baseline; Columns (2)–(3) add country fixed effects and country-specific linear time trends; Columns (4)–(6) add the bond return variance (multiplied by 100), the term spread, and the Laubach-Williams estimate of the natural rate of interest, r-star (with country fixed effects); and Columns (7)–(8) include time fixed effects and country-plus-time fixed effects. The bond return variance is the annualized realized variance of monthly bond excess returns over the five-year window, multiplied by 100; the term spread is the difference between the ten-year and one-year government bond yields; and r-star is the U.S. estimate, used for all countries. Driscoll–Kraay standard errors with 27 lags are reported in parentheses; stars denote significance at the 10, 5, and 1 percent levels based on Kiefer–Vogelsang fixed-b critical values. The data are monthly and span September 1989 to December 2024; a few countries use shorter samples, as detailed in Table B9.

Figure B6: International Panel: Expected Returns on Bond-U.S. Stock Return Beta



Note: The figure plots the estimated coefficient $\hat{\lambda}_c$ from Equation (15), estimated separately for each country by regressing the expected one-year bond excess return on the bond–stock beta, using only time-series variation. It is analogous to Figure 4 in the main text, but the beta is estimated with respect to the U.S. stock market rather than the domestic one: the slope of the country's monthly bond excess return on the U.S. stock market excess return over a five-year (60-month) rolling window. The bars report 95% confidence intervals constructed using Newey–West standard errors with 27 lags. The data are monthly and, for most countries, span September 1989 to September 2025; a few countries use shorter samples, as detailed in Table B9.