Risk and Return in Government Bonds*

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Abstract

As the risks of government bonds change over time, does this change the compensation that investors require for holding bonds? While realized excess returns show little relation to bond risk, we find that subjective expected excess returns constructed from professional forecasts of future long-term yields are tightly linked to bonds' stock market betas, consistent with a CAPM-style relation. In a U.S. month-by-maturity panel from 1988–2024, five-year rolling bond–stock betas are highly correlated with subjective excess returns, with a correlation of 66%. The estimated market price of risk is economically large, comparable to the equity premium, and stable when controlling for time and maturity fixed effects. Realized excess returns are predicted by subjective excess returns, but this predictability is driven by higher-frequency variation and not by betas. Similar results hold in an international panel of developed countries from 1989–2024. During quantitative easing episodes, the price of bond risk declines, consistent with an increased investor willingness to bear risk. The implied ten-year U.S. term premium fell substantially from the 1980s to the 2010, turning negative in 2001, earlier than conventional estimates.

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1. Introduction

Long-term government bonds are often viewed as safe assets, yet even in the absence of default risk they can expose investors to substantial interest-rate and inflation risk. In the United States in the 1970s and 1980s, government bonds were risky in the sense that they exhibited strong comovement with stocks. By contrast, in the 2000s, they were largely hedges, negatively correlated with stocks. Bond returns have again recently become more positively correlated with stock returns in the U.S. and other advanced economies. These shifts raise a natural question: How do changes in bond risk translate into the returns investors require for holding long-term government bonds?

We provide new evidence on the link between expected excess returns on government bonds and their underlying risk exposures, shedding light on the determinants of bond risk premia. In the classic capital asset pricing model (CAPM) of Sharpe (1964) and Lintner (1965), expected excess returns are proportional to an asset's beta with respect to the market portfolio. We show new evidence that expected excess returns on government bonds are well described by a CAPM-style relation once subjective interest rate expectations are incorporated. Using forecasts of future government bond yields by professional forecasters, we construct one-year expected bond excess returns purged of interest-rate surprises. We find that these subjective expected excess returns are tightly related to bond–stock betas, with an implied price of risk close to the equity premium. The relationship holds over time and across maturities in U.S. data, and in an international panel of developed countries spanning nearly four decades.

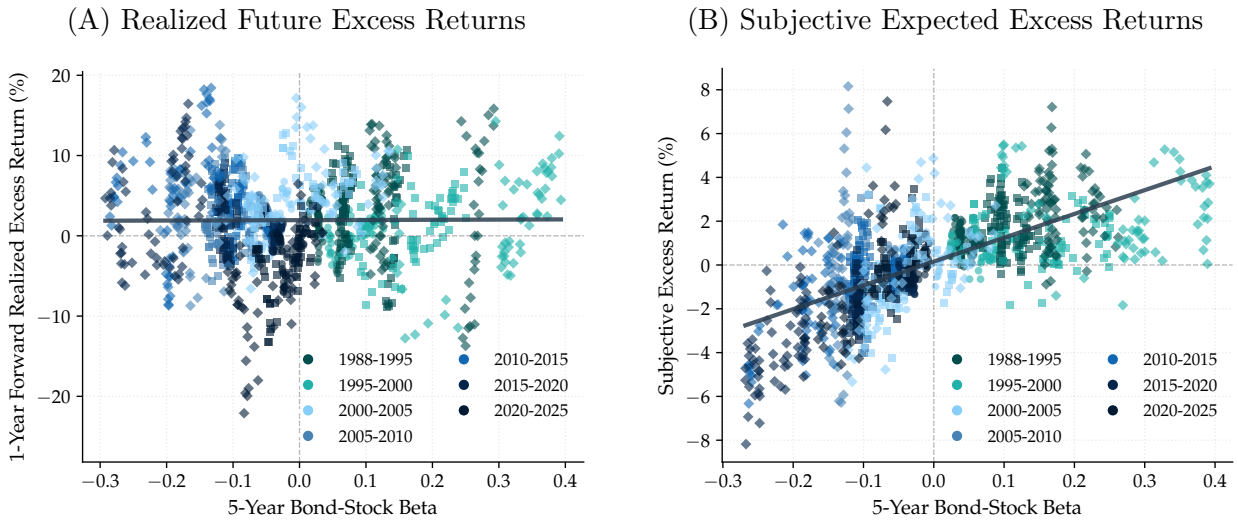
Figure 1 illustrates the importance of using subjective expected bond excess returns to uncover this result for the U.S. Panel A plots realized one-year bond excess returns on the y-axis against bond-stock betas on the x-axis. Each dot corresponds to one month and bond maturity combination.[1] The risks of Treasury bonds have varied substantially over time. In broad strokes, Treasury bonds turned from being risky with positive betas before 2000, to being hedges with negative bonds-stock betas after 2000, though there is also substantial higher-frequency variation.[2] According to the CAPM, one would expect a positive relationship between bond excess returns and the risk of bonds, but no such relationship is visible in Panel A.

By contrast, we see a strongly positive relationship in Panel B, where bond excess returns are cleansed from the effects of yield surprises using professional survey forecasts of

[1] Betas are estimated over five-year rolling windows from daily bond and stock excess returns. We use three, six, and eleven-year par bond excess returns and the matching betas to match the bond maturities available in the Blue Chip of Financial Forecasters (BCFF).

[2] See Campbell et al. (2026) for a review of variation in bond-stock betas and their economic drivers.

Figure 1: Bond Excess Returns and Bond-Stock Betas



Note: Panel A plots realized excess returns on U.S. Treasury bonds—computed over the subsequent one-year period—against their five-year stock-market betas for different maturities. Panel B plots subjective expected excess returns against the corresponding five-year stock-market betas. Maturities are distinguished by marker shape: three-year bonds (circles), six-year bonds (squares), and eleven-year bonds (diamonds). The sample spans January 1988 to March 2024.

future bond yields. When bond-stock betas were negative in the 2000s, forecasters expected negative excess returns on government bonds. Conversely, when betas were positive and bonds were risky, forecasters expected positive excess returns. The slope of the relationship in Panel B is 11% annualized with a raw correlation of 66%. This finding is consistent with the way major financial institutions active in bond markets and represented in the BCFF sample discuss bond risk premia. For example, Goldman Sachs Economics Research from June 2025 noted that the "record-high correlation between stock returns and Treasuries" creates a "need for greater risk premium on bonds to incentivize demand."[3]

We start by documenting the relationship between subjective expected bond excess returns and rolling bond-stock betas for the U.S. in a bond maturity-by-month panel from January 1988 through March 2024. We use the average four-quarter forecast of two-, 5-, and 10-year Treasury bond yields from Blue Chip Financial Forecasters (BCFF). Following Piazzesi et al. (2015) and Nagel and Xu (2023), expected one-year bond excess returns are computed analogously to realized one-year excess returns, simply replacing the realized bond yield in one year by its forecast. Bond-stock betas are computed from daily bond and stock excess returns over either five-year or one-year backward-looking rolling windows. We find substantial variation in both subjective expected bond excess returns and betas.

[3]Zu, Bill "Global Markets Daily: Are Treasuries Cheap Enough as a Risk Hedge Yet?" Goldman Sachs Economics Research, June 12, 2025.

In a panel regression of U.S. subjective expected bond excess returns onto betas, we find an economically and statistically significantly positive relationship, consistent with the evidence in Figure 1, Panel B. The magnitude and significance of the relationship is unchanged when we include a time trend, maturity and time fixed effects, and control for the term spread or the principal components of yields. The relationship is also similar for each bond maturity, with or without controlling for a time trend. A higher bond return variance is also associated with higher subjective excess returns, but leaves the coefficient on bond-stock betas unchanged and offers only limited additional explanatory power. When we use one-year rolling estimates of betas, the relationship between betas and expected bond excess returns is also strongly positive and significant, though somewhat smaller.

Next, using the U.S. maturity-by-month panel, we document that subjective excess returns predict future realized excess returns, even though betas do not. In other words, the component of subjective expected returns orthogonal to betas strongly predicts future realized returns with a coefficient not statistically distinguishable from the full information rational expectations (FIRE) hypothesis. In our sample period, the slow-moving component associated with bond-stock betas does not predict realized excess returns, but the higher-frequency residual in expected bond excess returns does. This suggests that professional forecasters have information about the bond market, but our sample may be too short to detect a significant relationship between realized bond excess returns and slow-moving betas.

We show that these new facts carry over to an international panel of twelve developed economies from September 1989 to December 2024, where subjective expected bond 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, already experiencing negative bond-stock betas in the mid-1990s, when U.S. betas were still positive. Matching this different pattern in betas, subjective bond excess returns turned negative earlier in Japan than in the U.S.

In a panel regression for the twelve developed countries, we estimate a strong and positive relationship between betas and subjective expected excess returns. A unit increase in beta is associated with an 8 percentage point higher expected excess return. This strong positive relationship holds when controlling for a time trend or country and time fixed effects, and the term spread, long-term, and short-term interest rates. Bond return variance enters positively but not significantly for the international panel. The estimated relationship between subjective excess returns and betas is positive and indistinguishable from the U.S. coefficient for

all countries in our sample, and generally precisely estimated.[4] Similarly to the U.S., subjective excess returns in the international panel have predictive power for realized bond excess returns, so forecasters appear to be informed. The relationship between bond-stock betas and realized bond excess returns is positive but imprecisely estimated, consistent with the idea that realized bond excess returns are too noisy over our sample to detect a relationship with slow-moving betas.

We then use our data to shed light on the channels of Quantitative Easing (QE). Over the last 25 years, central banks around the world have engaged in large-scale asset purchases with the purpose of lowering long-term bond yields, when traditional short-term policy rates were constrained by the zero lower bound. However, the channels by which QE moves long-term bond yields remain unclear (see Tenreyro and Wazzi (2025) for a review). First, QE could reduce the price that investors require for holding bond market risk, the so-called "portfolio balance" channel (Tobin (1958, 1969); Vayanos and Vila (2021)). Second, state-contingent promises of future QE could change the risks that bonds are exposed to (Haddad et al. (2025)). Finally, QE could leave bond risk premia unchanged under the "signaling channel", if QE simply acts as a credible signal about future short rates. Using our international panel, we provide new evidence on the drivers of bond risk premia during QE. We find that during QE, the relationship between betas and subjective excess returns weakens. Analogously, we find that while bond return volatility is associated with higher subjective expected excess returns during non-QE periods, this relationship also weakens during QE. Finally, we find that periods of QE are often associated with changes in bond-stock betas, though the strength and even direction can vary depending on macroeconomic conditions. These results matter for policy, suggesting that QE should have different effects on long-term bond yields depending on whether bonds are risky or safe, as measured by bond-stock betas.

Finally, our estimated relationship between bond-stock betas and expected bond excess returns can be used to decompose 10-year U.S. Treasury bond yields into risk-neutral and term premium components, which are relevant to monetary and fiscal policy (Caballero and Simsek, 2022; Du et al., 2020). The long decline in U.S. Treasury yields since the 1980s has generated an active debate about the extent to which the 'natural' rate has declined, due to a secular decline in economic growth and lower inflation (Bauer and Rudebusch, 2020). The U.S. zero-coupon 10-year Treasury yield declined by 7.28 percentage points over the thirty year period from January 1986 to January 2016, the end of the first zero-lower-bound period.

[4]Only a small number of countries with substantial credit risk, Italy and Spain, and some countries with shorter samples, such as Norway and Switzerland, have insignificant relationships between betas and expected bond excess returns.

Since then, the 10-year Treasury rate has increased substantially by 2.27 percentage points. Our estimates suggest that about half the decline from 1986 to 2016 and half the increase since 2016 can be attributed to term premia due to the changing riskiness of Treasury bonds. The increasing risk of the real (i.e., inflation-indexed) component of bond returns accounts for most of the recent increase in estimated term premia. While our estimated peak-to-trough magnitudes are similar to conventional term premium estimates (Adrian et al., 2013; Kim and Wright, 2011), we estimate an earlier switch to negative term premia during the early 2000s. If this early decline in long-term yields was not correctly attributed to falling term premia, it may have contributed to an overly accommodative stance of monetary policy in the run-up to the global financial crisis.

1.1 Literature Review

This paper speaks to the long-standing question in asset pricing on what constitutes risk and should be priced in the cross-section of financial assets, providing new evidence on the integration of bond and stock markets. 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. Ross (1976)'s arbitrage pricing theory generalizes this intuition by allowing a small number of priced factors rather than a single market factor. Empirically, the literature—focused largely on U.S. equities—finds mixed support. Early tests by Black et al. (1972) and Fama and MacBeth (1973) document a positive but too-flat relation between beta and returns, while later work identifies substantial cross-sectional deviations (see Fama and French (2004); Frazzini and Pedersen (2014)). Beyond equities, Fama and French (1993) shows that bond and stock returns load on distinct term and default factors.

We also contribute to the literature on the term structure of interest rates. While the expectations hypothesis of interest rates states that expected bond excess returns are constant and long-term bond yields reflect only expected future short-term rates (Cochrane, 2005, Ch. 19), there is substantial evidence that bond risk premia vary strongly over time (Fama and Bliss, 1987; Campbell and Shiller, 1991; Cochrane and Piazzesi, 2005; Adrian et al., 2013, e.g.). Affine term structure models, building on Vasicek (1977) and Cox et al. (1985), allow for time-varying bond risk premia driven by a small set of state variables—typically bond yields themselves—but are largely silent on links to the stock market. 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. In contrast, segmented-markets approaches, such as Vayanos and Vila (2021), attribute bond risk premia to arbitrageur constraints and maturity-specific demand rather than full market integration.

More recently, a growing literature emphasizes that bond risks—as measured by bond–stock betas—vary substantially over time. While U.S. Treasury bonds comoved positively with the stock market in earlier decades, their comovement turned negative around the turn of the millennium, with bonds acting as hedges against equity risk (Baele et al. (2010), Campbell et al. (2017)). Building on this evidence, Campbell et al. (2020) show that bond–stock betas are closely linked to inflation cyclicality, with time-varying risk premia providing a quantitatively important amplification mechanism. Pflueger (2025) shows that turning bonds into risky assets, as they were in the 1980s, requires both: (a) volatile supply-type shocks; and (b) an aggressive Federal Reserve response that prioritizes fighting inflation over supporting employment. Neither element alone is sufficient. In standard consumption-based models where the stock market approximates the maximum Sharpe ratio portfolio, time-varying bond–stock betas are tightly connected to term premia (Campbell et al. (2026)). Despite this strong theoretical link, tests based on realized bond excess returns find little relation between betas and returns. We contribute by providing new evidence that betas are an important determinant of professional forecasts of expected bond excess returns.

In our use of survey expectations, we build on tools from the recent literature investigating the role of expectations in asset prices. Following Piazzesi et al. (2015) and Nagel and Xu (2023), we define subjective expected excess returns as those that would obtain if professional forecasts of bond yields at a one-year horizon were realized. While in an infinite sample with rational expectations, the gap between subjective expectation and realized bond excess returns would merely be noise, several authors have pointed out that it can be substantial and systematic in practice (Cieslak, 2018). Such errors may arise from slow but rational learning (Farmer et al., 2024) or behavioral biases (Shue et al., 2025). Our measure of subjective expected bond excess returns automatically adjusts for expectations errors due to both small samples and behavioral biases, allowing us to study how bond risks relate to subjective expected bond risk premia.

Beyond bonds, a large literature shows that survey expectations contain useful information about asset prices across markets. In equities, earnings expectations are priced (Bordalo et al., 2024; Delao and Myers, 2021). In foreign exchange, survey forecasts predict long-horizon currency returns (Kremens et al., 2025). Couts et al. (2026) show that institutional investors' long-run return expectations are consistent with their expectations for covariance, with a median forecast horizon of ten years. Different from them, we consider actual bond-stock betas and show that they are linked to one-year subjective excess returns.

Within fixed income, Pesch et al. (2024) links bond return volatility to subjective excess returns, while Rogers (2026) uses derivatives to construct short-horizon expectations, and Molavi et al. (2025b) finds little evidence of time-varying term premia for short-term bonds.

We contribute to this literature by showing that betas are a key determinant of subjective risk premia for long-term bonds.

Finally, this paper contributes to the literature on the channels of quantitative easing (QE) and the secular decline in long-term bond yields. QE may affect long-term yields through several mechanisms: signaling lower expected future short rates, removing duration risk from private investors' balance sheets (the "portfolio balance" channel), and committing to state-contingent interventions in financial markets (Krishnamurthy et al., 2013; Haddad et al., 2025; Tenreyro and Wazzi, 2025). Using our international panel, we find evidence consistent with the portfolio balance and state-contingent policy channels, as QE is associated with a decline in bond risk premia and a lower price of risk for bond–stock betas. More broadly, a large literature documents a secular decline in long-term yields and term premia (Laubach and Williams, 2003, 2016; Bianchi et al., 2022; Hillenbrand, 2025). Leveraging our estimated link between bond–stock betas and bond risk premia, we show that the implied term premium declined substantially over time. The magnitude of this decline is comparable to standard estimates based on affine term structure models (Adrian et al., 2013; Kim and Wright, 2011), although our estimates suggest that the decline began earlier.

2. U.S. Panel Evidence

This section presents the main results for the U.S. We describe and plot data for expected excess returns and bond-stock betas, and estimate their relationship in a month-by-maturity panel regression. Realized bond excess returns are regressed onto expected bond excess returns, decomposed into the fitted component from bond-stock beta and a residual.

2.1 U.S. Data and Summary Statistics

The data for our U.S. sample come from three 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, various Treasury yields, and corporate bond yields. Participating institutions include large investment banks and other financial institutions, with between 30 and 50 institutions participating in each survey. Anecdotally, interest rate forecasts at major financial institutions are constructed by fixed income forecasting divisions closely connected with fixed income trading, so the BCFF interest rate expectations are likely to be relevant for bond pricing. Our sample is monthly, starting in January 1988, when a consistent set of Treasury yields 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.

Subjective Expected Excess Returns on Bonds Following the data construction of Piazzesi et al. (2015) and Nagel and Xu (2023), we compute the subjective expected bond excess return or risk premium for an n-year coupon bond by combining the current 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:

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

$y^{par}_{n,t}$ denotes the current n-year government bond yield, $\tilde{E}_t y^{par}_{n-1,t+12}$ denotes the forecast for the $(n-1)$-year yield in 12 months, and $y_{1,t}$ is the one-year zero-coupon yield.[5] All yields are in annualized percent, so the expected excess return is also in annualized percentage units. Because the forecasts are made for coupon Treasury bonds and Treasury bonds tend to sell close to par, we approximate the duration of the n-year bond selling at par in month as in (Campbell, 2017, pp. 236–237):

$$dur_{n,t} = \left(1 - (1 + y^{par}_{n,t})^{(-n)}\right) / \left(1 - (1 + y^{par}_{n,t})^{(-1)}\right), \tag{2}$$

where $y^{par}_{n,t}$ is the n-year average par yield in the last week of month t from Gürkaynak et al. (2007), for $n = 3, 6, 11$. Since Blue Chip contains forecasts for the two-, 5-, and 10-year bond yields, we can construct subjective expected excess returns for the 3-, 6-, and 11-year government bonds. We proxy for the 12-month ahead forecast with 4-quarter forecasts, averaged across forecasters. 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 of February in expression (1). We average the realized par yield over the last week of each month to reduce day-to-day noise.

This procedure gives us a monthly series for U.S. expected bond excess returns over a sample from January 1988 through March 2024. Because we have monthly data and three different bond maturities ($n = 3, 6, 11$), this leaves us with roughly 1300 observations, though of course we must account for joint variation across maturities and time in our regressions.

Realized Bond Excess Returns We also construct realized twelve-month bond excess returns for an n-year par bond:

$$xr_{n,t\to t+12} = dur_{n,t}\, y^{par}_{n,t} - (dur_{n,t} - 1)\, y^{par}_{n-1,t+12} - y_{1,t}. \tag{3}$$

[5] For the one-year zero coupon yield we use the market yield on U.S. Treasury Securities at one year constant maturity downloaded from Fred (ticker `DGS1`).

Realized bond excess returns on n-year zero-coupon bonds are computed analogously, using zero-coupon bond yields from Gürkaynak et al. (2007) and setting the bond duration equal to the bond maturity.

Summary statistics in Table I, Panel A, show that in contrast to realized bond excess returns, subjective 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 subjective expected excess returns on the three-year bond ranges from -1.2% to 1.5%, with a standard deviation of 0.7%. The moving average of subjective expected excess returns on the eleven-year exhibits more variation, ranging from -5.5% to 4.7%, with a standard deviation of 2.7%.

Our measures of subjective expected excess returns are plotted in Figure 2, together with a 12-month moving average. While there is substantial variation month-to-month, the moving average shows clear patterns over time and across maturities.[6]

Bond Betas The other key variable in our analysis is the CAPM-beta of Treasury bonds with respect to the stock market, which we estimate using daily data. Daily U.S. stock market excess returns are obtained from Ken French's website. The Treasury bond excess return from day d to $d+1$ is computed using nominal Treasury par yields from Gürkaynak et al. (2007) as

$$xr_{n,d+1} = dur_{n,d} y^{par}_{n,d} - \left(dur_{n,d} - \frac{1}{252} \right) y^{par}_{n,d+1} - y_{rfr,d}, \tag{4}$$

where the daily risk-free rate $y_{rfr,d}$ is from Ken French's website, and the duration of the n-year par bond on day d is computed analogously to equation (2).[7]

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 daily data:

$$xr_{n,d} = a_{n,t} + \beta^{w}_{n,t} xr^{eq}_{d} + \varepsilon_{d}, \tag{5}$$

where d denotes the trading day and runs from the first trading day of month $t-w$ to the last trading day of month t. Our baseline estimate of $\hat{\beta}^{w}_{n,t}$ uses a rolling window of $w = 60$ months (a five-year rolling beta) but we also report estimates based on a shorter window of

[6] 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. In Tables IV and V, where expected excess returns are on the right-hand side, we use the twelve-month moving average to mitigate attenuation bias.

[7] Par yields are in annualized percent, the risk-free rate is the daily T-Bill return in percent, and duration is measured in years.

$w = 12$ months (a one-year rolling beta).

The monthly series of estimated betas are shown in Figure 3, and summary statistics are provided in Panel B of Table I. Across maturities, the overall time-series patterns are broadly similar, although there are notable cross-sectional differences. During periods of positive betas—such as at the beginning of the sample—the eleven-year bond exhibits a substantially larger beta than the shorter maturities. At its peak, the five-year rolling beta for the eleven-year bond reaches 0.4, compared with a peak of 0.1 for the three-year bond.

Figures 2 and 3 suggest a strong association between betas and subjective expected returns, both over time and across maturities. Subjective expected excess returns were positive until the turn of the millennium, then turned negative after 2000. They fluctuated around zero during the GFC and its aftermath. Several more years of negative expected bond excess returns during the 2010s and early 2020s were then followed by a return to positive expected bond excess returns in 2023 and early 2024. There is also substantial variation across bond maturities, with expected excess returns on eleven-year bonds more positive during the 1980s and 1990s, but also more negative during the 2000s.

The parallels to bond-stock betas in Figure 3 are visually apparent. Betas were positive before the turn of the millennium, but then became negative thereafter. The change in bond stock betas is most apparent in the eleven-year bond, but still visible in the smoother beta for three-year maturity bonds. The parallels also extend to the increase in subjective expected returns in recent years. Betas turned positive around the same time as subjective expected excess returns, with both patterns more pronounced in long-maturity bonds than shorter-maturity bonds. These figures provide a first hint that the rise in bond-stock correlations after 2023 may be linked to the substantial increase in long-term bond yields during the post-pandemic period.

In our regressions below, these rolling betas are technically generated regressors. However, as Panel B of Table I shows, they are very precisely estimated, so any concerns about statistical inference with generated regressors should be minimal.[8] The five-year rolling beta is somewhat more precisely estimated than the one-year, so we use it as the baseline in discussing our results, but we report all of our main results for both betas.

Alternative Beta Estimates Our baseline bond–stock beta is estimated using daily returns over a five-year rolling window. In addition to this baseline measure and the alternative specification based on a one-year rolling window, we also consider several alternative

[8]For betas computed over five-year rolling windows, the average Huber-White robust standard errors are 0.008 for the three-year Treasury bond, 0.014 for the 6-year Treasury bond, and 0.022 for the 11-year Treasury bond. For comparison, the bond-stock betas exhibit quite large variation over time. The standard deviations of the point estimates of these betas are 0.05 for the three-year Treasury bond, 0.11 for the 6-year Treasury bond, and 0.19 for the 11-year Treasury bond.

Table I: Summary Statistics: U.S. Panel

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

	Subjective XR			Subjective XR (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	1.02	2.08	2.84
Std	0.84	1.76	3.11	0.71	1.46	2.70	2.43	4.75	7.24
Min	-2.03	-4.23	-8.18	-1.21	-3.17	-5.53	-7.05	-13.21	-22.11
Max	2.17	4.42	8.16	1.55	3.06	4.67	6.75	12.77	18.42

Panel (B): Bond-Stock Betas

	5-y rolling β			1-y rolling β		
	3Y	6Y	11Y	3Y	6Y	11Y
Average	-0.01	-0.01	-0.02	-0.00	-0.00	-0.01
Std	0.05	0.11	0.19	0.07	0.14	0.23
Min	-0.09	-0.13	-0.27	-0.14	-0.22	-0.36
Max	0.13	0.25	0.39	0.21	0.42	0.67
Average S.E.	0.008	0.014	0.022	0.013	0.025	0.039

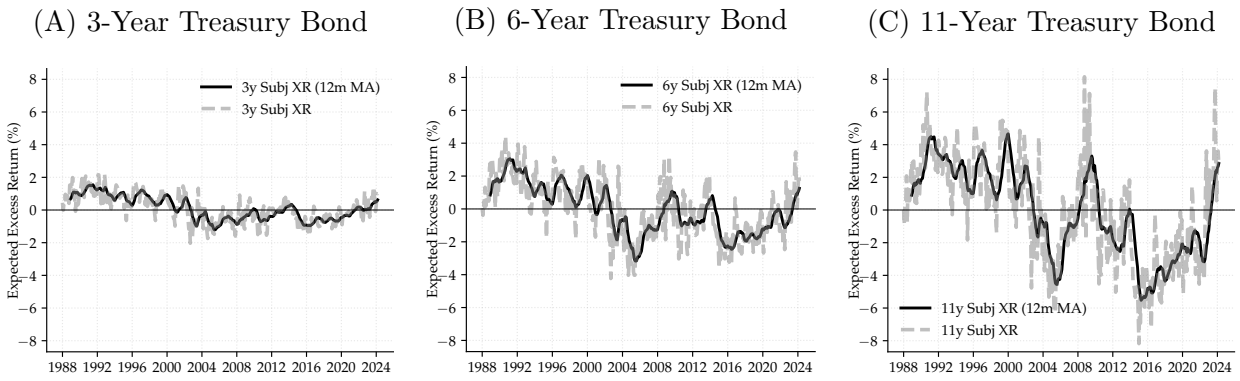
Note: Panel A reports the summary statistics for the expected excess returns (columns (1) to (3)), 12-month moving average (MA) expected excess returns (columns (4) to (6)), and realized excess returns (columns (7) to (9)) for bonds with three-, 6-, and 11-year maturities. Subjective expected bond excess returns are computed according to Equation 1. We use 4-quarter par bond yield forecasts from BCFF. Panel B reports the summary statistics for the rolling estimates of Treasury bond–stock betas for par bonds with three-year, 6-, and 11-year maturities. Bond–stock betas are computed as the regression coefficient of daily par bond excess returns on stock excess returns over a backward-looking rolling window, as defined in Equation 5. We use either 60-month rolling windows (five-year rolling β) or 12-month rolling windows (one-year rolling β). We resample the data to a monthly frequency by taking the estimate from the last trading day of each month. The table reports the average estimate, standard deviation, minimum and maximum values, and the average HW-robust standard errors. The sample runs from January 1988 to March 2024.

estimates of bond–stock betas. First, we estimate rolling regressions using daily data and apply the Dimson (1979) adjustment to account for non-synchronous trading. The Dimson beta is obtained by including two lags and two leads of the market return in the regression and summing the corresponding coefficients.[9] Second, we estimate the five-year rolling betas using monthly rather than daily returns. Finally, we construct bond–stock betas using exponentially decaying weights with a half-life of three years. The resulting beta estimates are shown in Appendix Figure A.2.

Other Predictors We also include additional predictors of bond excess returns. In particular, we construct a measure of bond return variance using the rolling variance of daily bond excess returns, computed over either a 60-month or a 12-month window. We further include controls for the shape of the yield curve. Specifically, we compute the level,

[9] Specifically, 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 $\beta_n^{Dimson} = \sum_{k=-2}^{2} \beta_{n,k}$.

Figure 2: Subjective Expected Excess Returns: U.S. Treasury Bonds



Note: This figure plots one-year subjective expected excess return for U.S. Treasury par bonds of three- (Panel A), 6- (Panel B), and 11-year (Panel C) maturities. Subjective expected excess returns are computed according to equation (1). We use 4-quarter par bond yield forecasts from the Blue Chip Financial Forecasts. The dashed gray line shows monthly subjective expected excess returns, and the black line shows the 12-month moving average. The sample is monthly from January 1988 to March 2024.

slope, and curvature factors as the first three principal components of zero-coupon nominal bond yields with maturities of one, two, 5, 7, 10, 15, and 20 years. Next, we define the term spread as the difference between the n-year par yield and the one-year zero-coupon yield. 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 subjective expected excess returns than for realized ex-post bond excess returns (Nagel and Xu (2023)). In our specification, we use the maturity-specific term spread: for the n year expected excess return, we use the n-year term spread. Finally, we 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.

2.2 U.S. Expected Excess Returns vs. Betas

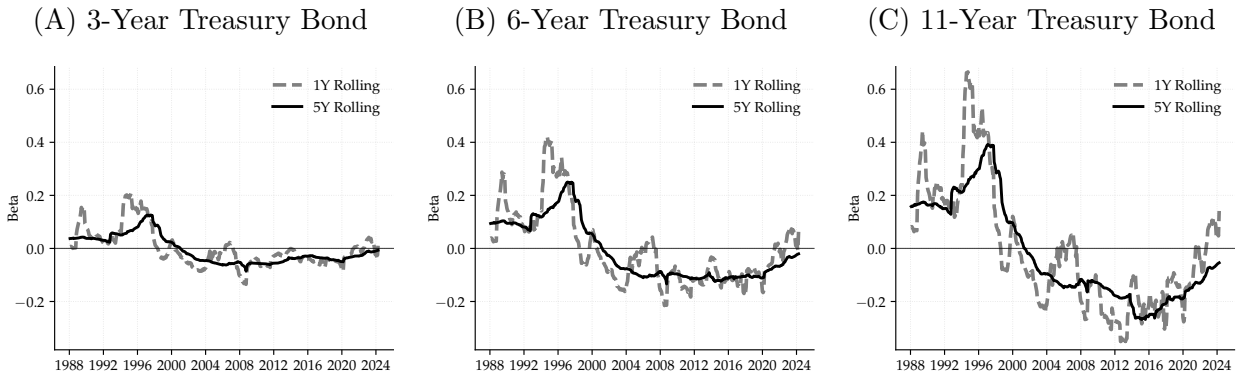
We next test whether subjective expected excess returns reflect bond risk as captured by bond–stock betas.[10] We evaluate the CAPM for government bonds by estimating the following regression:

$$\tilde{E}_t xr_{n,t\to t+12} = \alpha + \lambda \hat{\beta}^w_{n,t} + X'_{n,t}\delta + \eta_{n,t}, \tag{6}$$

where $\hat{\beta}^w_{n,t}$ denotes the estimated rolling beta available at time t, and $X_{n,t}$ is a vector of control variables. We estimate regression (6) both in a stacked panel, including bonds of all

[10]We sometimes refer to *expected excess return* as shorthand for subjective expected bond excess return and *beta* as a shorthand for bond-stock beta.

Figure 3: Estimated Stock Market Betas for U.S. Treasury Bonds



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 daily par bond excess returns onto stock market excess returns over a backward-looking rolling window according to equation (5). The dashed gray line uses 12-month rolling windows, and the black line uses 60-month rolling windows. 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.

three available maturities, and separately for each maturity using only time series variation. According to the CAPM, the slope λ is predicted to be positive and equal to the equity premium, and the intercept α is predicted to equal zero.

We are careful to ensure that the information in betas and other yield curve-based controls is known at the time the forecasts used to construct our subjective expected returns expectations are made. For concreteness, consider the March observation 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.

Table II shows the results for the stacked panel, where the unit of observation is bond maturity $n \in \{3, 6, 11\}$ by month t. Panel A uses our baseline betas estimated over five-year rolling backward-looking windows, while Panel B uses the noisier but faster-updating one-year rolling beta. 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 Panel A shows our baseline result. A unit increase in beta is associated with an increase in subjective expected excess returns of 10.9 percentage points per year. This magnitude is large and statistically indistinguishable from conventional estimates of the equity premium, which range from 6 to 8 percent per year. The R^2 of the regression is high at 43%, corresponding to a correlation coefficient of 66%. In addition, the constant, or the expected alpha, in Column (1) is close to zero and statistically insignificant, indicating that forecasters did not expect bonds to earn a premium above and beyond compensation for exposure to the stock market. As a simple back-of-the-envelope calculation, the peak-to-trough decline in the beta of the eleven-year bond from roughly 0.4 to -0.3 can hence explain a decline in the expected bond excess return from $0.4 \times 10.9 = 4.26\%$ to $-0.3 \times 10.9 = -2.91\%$. For comparison, the moving average of expected excess returns in Figure 2 Panel C declined from roughly 4.7% to -5.5%.

We next show that the relationship between betas and subjective expected excess returns is unchanged economically and statistically when we include a host of controls. Given that both subjective expected return and betas drift downward over our sample, it is important to control for a time trend. Column (2) shows that this leaves our baseline result unchanged. Column (3) includes maturity fixed effects to allow for the possibility that some bond maturities may persistently have higher or lower expected excess returns, for example, due to different liquidity. Column (4) adds bonds' return variance, estimated as the variance of daily bond excess returns over a rolling five-year window. The estimated coefficient on the beta in this specification remains similar, and the R-squared increases slightly. The coefficient on bond variance, equal to 5.7, is also positive and statistically significant. This result is consistent with the idea that investors require higher expected excess returns during periods of elevated bond return volatility, though the relationship with betas remains unchanged.

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 one-year par bond yield. Column (6) controls for the first three principal components of the yield curve: level, slope, and curvature. Throughout Panel A, the R^2 only increases modestly from column (1) to column (6), emphasizing the substantial explanatory power of betas for subjective expected excess

returns.

Column (7) flexibly controls for both time-series patterns and maturity-specific characteristics by including time and maturity fixed effects. Because time fixed effects absorb aggregate variation common to all maturities, this specification exploits only cross-maturity variation within a given month. The estimated coefficient in Column (7), equal to 11.7, remains statistically significant and similar in magnitude to those reported in the other columns. This finding indicates that the relationship between bond–stock betas and subjective expected excess returns holds both over time and across maturities.

The results in Panel B show that the one-year rolling beta has similarly high explanatory power for expected bond excess returns, though the coefficients and R^2s are somewhat attenuated compared to Panel A. This is what one would expect if using the shorter one-year rolling window leads to measurement error and hence attenuation bias. However, the magnitude of the relationship is still strong, implying that a one-unit increase in beta is associated with a 4 to 8 percentage point increase in expected excess returns. The results are again economically and statistically similar when including a time trend, maturity, and time fixed effects, and term structure controls.

Time Series Table III reports results from estimating the same specification separately for each maturity. Panel A presents estimates based on the five-year bond–stock beta, while Panel B reports results using the one-year bond–stock beta. The table shows that the relationship between the bond–stock beta and subjective expected returns is both statistically and economically significant for each maturity. The estimated coefficients are similar across maturities. Using the five-year bond–stock beta, without any additional controls, the estimated $\hat{\lambda}$ is approximately 11.0 for the three-year bond, 9.9 for the 6-year bond, and 11.2 for the 11-year bond.

Robustness Table A1 in the Appendix reports results obtained using alternative measures of bond–stock betas. In Column (3), we use betas estimated from daily returns over a five-year rolling window with the Dimson (1979) correction for non-synchronous trading, while Column (5) uses betas estimated from monthly returns over a five-year window. In both cases, the estimated coefficients and the R^2 are broadly unchanged relative to the baseline specification. In Column (6), we construct betas using monthly data with exponentially decaying weights with a half-life of three years. In this case, the estimated coefficient is larger (12.47), and the explanatory power of the regression increases, with the R^2 rising to 47%.

Table A2 in the Appendix reports results from specifications that control for a range of alternative variables, including realized bond return volatility and realized stock return

volatility estimated over different rolling windows, as well as the VIX. We also show that our results are distinct from, and incremental to, the expected excess returns implied by biased interest rate forecasts. Specifically, we control for the expected excess return implied by the slow-learning model of Farmer et al. (2024), which has been shown to explain the majority of the forecast errors in policy rate expectations in recent decades, including forecasters' failure to anticipate the timing of liftoff from the zero lower bound. The estimated coefficient on the bond–stock beta is virtually unchanged when controlling for any of these variables.[11]

Overall, our results point to a robust relationship between expected excess returns and betas. This relationship holds both over time and across maturities and remains stable across a wide range of controls and fixed effects.

2.3 U.S. Realized Excess Return Predictability

One might wonder whether the relationship between betas and expected bond excess returns carries over to realized excess returns. Under the full information rational expectations (FIRE) benchmark, betas should have the same relationship with subjective expected returns and realized returns in a sufficiently long sample. In practice, however, our sample is finite, and interest rate forecasts may be biased due to behavioral biases (Bordalo et al., 2020; Shue et al., 2025; Molavi et al., 2025a) or imperfect knowledge of the parameters describing the interest rate process (Farmer et al., 2024; Molavi et al., 2024; Bauer et al., 2024).

We estimate predictive regressions of the form:

$$xr^{par}_{n,t\to t+12} = b_0 + b_1\, \tilde{E}xr_{n,t\to t+12} + b_2\, (\hat{\lambda} \times \hat{\beta}^{w}_{n,t}) + X'_{n,t}\delta + \varepsilon_{n,t+12}. \tag{7}$$

Here, $\hat{\beta}^{w}_{n,t}$ denotes the fitted value from the regression of expected excess returns on bond–stock betas, as specified in equation (6). We use the fitted value solely to ensure that the magnitude of the coefficient is directly comparable to that of subjective expected excess returns.

Under the FIRE benchmark, we would expect to find that the subjective expected excess return $\tilde{E}xr_{n,t\to t+12}$ and the scaled beta $(\hat{\lambda} \times \hat{\beta}^{w}_{n,t})$ should both individually enter with unit coefficients, i.e. $b_1 = b_2 = 1$. However, if forecasters have no information above and beyond

[11] Tables A3–A4 interact bond–stock betas with state variables capturing whether the price of risk varies with macro-financial conditions. Each column considers a different state variable, including valuation ratios (P/E and P/D), the VIX, and measures of market and bond return volatility. Using five-year rolling bond–stock betas, the results suggest that the premium increases when past realized market volatility—measured using five-year rolling volatility—is high (Column (5) of Table A3). The interaction coefficient is positive and statistically significant. This finding is consistent with the idea that periods of high market volatility are associated with higher risk premia. For example, when betas are negative, the estimates imply that the hedging premium of bonds becomes larger in magnitude when stock market volatility is high, consistent with the notion that hedging is particularly valuable in high-risk environments.

Table II: Expected Excess Returns and Stock Market Betas: U.S. Panel

Panel (A): Five-Year Rolling Beta

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Bond-Stock Beta 5YR	10.87***	10.12***	10.87***	11.24***	11.09***	8.38***	11.69***
	(1.84)	(2.41)	(1.84)	(1.80)	(1.75)	(1.87)	(1.89)
Bond Variance				5.73**			
				(2.87)			
Term Spread					0.32		
					(0.24)		
Level						0.47***	
						(0.18)	
Slope						-0.26	
						(0.17)	
Curvature						0.14	
						(0.17)	
Constant	0.16						
	(0.20)						
Time Trend		✓					
Maturity FE			✓	✓	✓	✓	✓
Time FE							✓
R-squared	0.43	0.43	0.43	0.44	0.44	0.47	0.86
Observations	1305	1305	1305	1305	1305	1305	1305

Panel (B): One-Year Rolling Beta

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Bond-Stock Beta 1YR	7.14***	5.11***	7.13***	7.42***	7.76***	4.20***	7.78***
	(1.55)	(1.61)	(1.53)	(1.56)	(1.67)	(1.40)	(1.71)
Bond Variance				5.27***			
				(1.95)			
Term Spread					0.49**		
					(0.23)		
Level						0.81***	
						(0.19)	
Slope						-0.33*	
						(0.17)	
Curvature						0.07	
						(0.16)	
Constant	0.02						
	(0.25)						
Time Trend		✓					
Maturity FE			✓	✓	✓	✓	✓
Time FE							✓
R-squared	0.29	0.34	0.29	0.34	0.33	0.41	0.82
Observations	1305	1305	1305	1305	1305	1305	1305

Note: The table reports the regression estimates for equation (6):

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

Panel (A) presents the results for five-year bond betas, while Panel (B) reports the results for one-year bond betas. 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 III: Expected Excess Returns and Stock Market Betas: U.S., Individual Maturities

Panel (A): Five-Year Rolling Beta

	3-year			6-year			11-year		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Beta 5YR	10.99***	9.23***	7.31***	9.89***	8.40***	5.89**	11.20***	10.22***	6.08***
	(1.89)	(2.42)	(2.10)	(1.94)	(3.06)	(2.54)	(1.86)	(3.53)	(2.22)
Level			0.28**			0.59**			1.16***
			(0.12)			(0.29)			(0.43)
Slope			-0.07			-0.26			-0.39
			(0.07)			(0.17)			(0.30)
Curvature			0.10			0.18			0.07
			(0.08)			(0.16)			(0.31)
Constant	0.17*	0.39*	0.15*	0.11	0.46	0.05	0.18	0.55	0.07
	(0.09)	(0.23)	(0.08)	(0.20)	(0.60)	(0.17)	(0.34)	(1.06)	(0.30)
Time Trend		✓			✓			✓	
R-squared	0.44	0.45	0.52	0.38	0.38	0.46	0.44	0.45	0.50
Observations	435	435	435	435	435	435	435	435	435

Panel (B): 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 (6) estimated maturity by maturity:

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

Panel (A) presents the results for five-year betas, while Panel (B) reports the results for one-year betas. Columns (2), (5), and (8) include a time trend. 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.

the CAPM, the beta should be the only driver of expected bond excess returns, the regressors would be collinear, and b_1 and b_2 might not be identified in a multivariate regression. We hence first introduce the regressors individually and then combine them in a multivariate regression.

Table IV reports the results. To increase power, we again stack different bond maturities in a panel, where the unit of observation is bond maturity ($n \in \{3, 6, 11\}$) by month. Column (1) reports a univariate regression of realized excess returns onto subjective expected returns,

finding a coefficient of 0.61. This coefficient is statistically significantly different from zero, but not significantly different from one. Expected bond excess returns hence predict realized bond excess returns with a large magnitude that does not reject that expected bond excess returns are roughly rational. While the R-squared in these columns may appear small at roughly 8%, recall that this is similar to the R-squared in typical stock return predictability regressions and can lead to a substantial improvement in investor Sharpe ratios (Campbell and Thompson, 2008). Overall, forecasters appear to have some useful information, but realized returns are at best a very noisy measure of ex-ante expected returns.

We next examine whether subjective expected returns are predictive due to their correlation with betas or due to the part of subjective expected return that is orthogonal to beta. In column (2), we regress realized excess returns on the beta (using the five-year rolling estimate). Consistent with Figure 1 in the introduction, the coefficient b_2 (in column (2)) is equal to 0.19 and is not statistically different from zero. However, the standard error is large, and we cannot reject that b_2 equals one. Hence, while forecasts are informative about future bond excess returns, the picture looks quite different for the component driven by betas. This could simply be due to noise in short samples, if realized bond yields have a substantial volatility relative to expectations, and bond-stock betas move slowly. If interest rate expectations further contain systematic bias, the empirical relationship between realized excess returns and betas could be further weakened relative to ex ante expectations. Either way, the results in Table IV reiterate the usefulness of using subjective expected excess returns to uncover the relationship with betas.

When we include both subjective expected returns and scaled betas in column (3), the two predictors enter with opposite signs, and the coefficient on expected bond excess returns strengthens further to 0.95. Controlling for the risk of Treasury bonds, a one percentage point increase in expected bond excess returns predicts a 95 bps increase in the realized bond excess return, i.e., almost one-for-one. The R^2 of the regression strengthens to 10.3%. A potential explanation for this result is that forecasters have higher-frequency information about bond returns, e.g., information about unconventional monetary policy or supply-and-demand imbalances, while betas are lower-frequency drivers of excess returns. Within our sample, subjective expected returns appear to better predict the first, higher-frequency component.

Column (4) includes the variance of bond returns, which enters with a positive coefficient, although it is not statistically significant. The inclusion of this control has little effect on the estimated coefficient on expected excess returns. The coefficient on the scaled beta remains of similar magnitude but is estimated less precisely.

The remaining columns of Table IV introduce term-structure controls and show that

Table IV: Realized and Expected Excess Returns: U.S. Panel

Panel (A): Realized and Expected Excess Returns

	(1)	(2)	(3)	(4)	(5)	(6)
Subjective Excess Returns	0.61***		0.95***	0.91***	0.74***	0.53**
	(0.23)		(0.22)	(0.21)	(0.22)	(0.23)
$\lambda\times$ Beta		0.19	-0.77**	-0.67	-0.46	-1.30**
		(0.38)	(0.36)	(0.42)	(0.46)	(0.63)
Bond Variance				11.68		
				(18.33)		
Term Spread					1.45**	
					(0.74)	
Level						1.99**
						(0.84)
Slope						-1.10*
						(0.65)
Curvature						0.73
						(0.57)
Maturity FE	✓	✓	✓	✓	✓	✓
R-squared	0.08	0.02	0.10	0.11	0.15	0.22
Observations	1269	1269	1269	1269	1269	1269

Panel (B): Realized and Expected Excess Returns (MA)

	(1)	(2)	(3)	(4)	(5)	(6)
Subj. Excess Returns (MA)	0.84***		1.85***	1.80***	1.55***	1.18***
	(0.28)		(0.33)	(0.32)	(0.35)	(0.33)
$\lambda\times$ Beta		0.19	-1.69***	-1.59***	-1.32**	-2.00***
		(0.38)	(0.45)	(0.53)	(0.59)	(0.73)
Bond Variance				6.98		
				(18.62)		
Term Spread					0.85	
					(0.74)	
Level						2.08**
						(0.91)
Slope						-0.74
						(0.67)
Curvature						0.57
						(0.56)
Maturity FE	✓	✓	✓	✓	✓	✓
R-squared	0.10	0.02	0.18	0.18	0.19	0.26
Observations	1236	1269	1236	1236	1236	1236

Note: The table reports the estimates of equation (7):

$$xr^{par}_{n,t\to t+12} = b_0 + b_1\,\tilde{E}xr_{n,t\to t+12} + b_2\,(\hat{\lambda}\times\hat{\beta}^{w}_{n,t}) + X'_{n,t}\delta + \varepsilon_{n,t+12}.$$

All independent variables are lagged by twelve months. $\lambda \times Beta$ is the fitted value from a regression of expected excess returns on five-year betas, as specified in equation (6). The beta is thus rescaled so that its magnitude is comparable to the expected excess return. Panel A uses expected excess returns, while Panel B uses a 12-month moving average of expected excess returns. 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.

the predictive power of expected bond excess returns remains economically and statistically significant when both subjective expected returns and the bond–stock beta are included (columns (5) and (6)). This suggests that survey forecasts contain information about future bond excess returns that is not fully captured by the term structure, highlighting the benefits of using survey forecasts directly rather than relying on term-structure models in which expected returns and yields are typically driven by a limited number of factors.

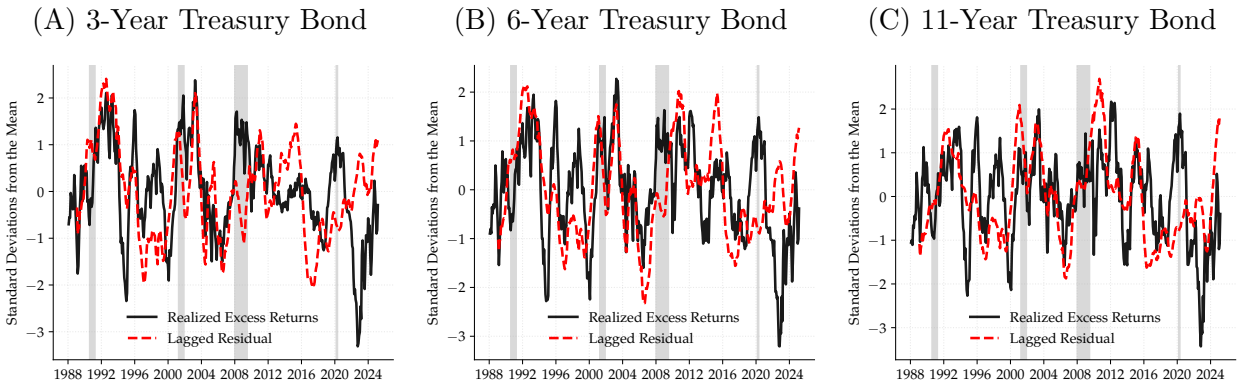
Table IV, Panel B repeats the analysis using the 12-month moving average of subjective expected returns, as plotted in Figure 2 to reduce noise. As one would expect if smoothing reduces the month-to-month noise in subjective expected returns, the coefficient b_1 increases in all specifications, and the R^2 increases as well. The coefficients on subjective expected returns in columns (1) and (6) (with all controls) are now statistically indistinguishable from one. The results for betas are unchanged. Overall, the results indicate that the predictability of future realized excess returns comes primarily from the component of expected excess returns that is orthogonal to beta.

Table V reports results from regressions estimated separately by maturity. The table confirms the findings from the panel regression. Panel (A), column (1), (5), and (9), report coefficients from a simple predictive regression of future realized bond excess returns on subjective expected returns. The estimated coefficient is approximately 0.6 for the three, 6, and 11-year maturities. When using the smoothed moving-average measure, the estimated coefficient increases to approximately 0.9, and the explanatory power rises substantially. In this case, the R^2 equals 14% for the three-year maturity, 17% for the six-year maturity, and 17% for the eleven-year maturity, indicating that subjective expected returns contain substantial predictive power for future realized bond excess returns.

Consistent with the results from the panel regression, the beta is not statistically significant when included on its own, but it enters with a negative and statistically significant coefficient when included jointly with expected excess returns. Finally, when we include term structure controls, the coefficient on expected excess returns remains broadly unchanged relative to the baseline specification without controls, although it is estimated with lower precision and is not statistically significant in most specifications.

To shed more light on the relationship between realized and expected excess returns, we plot the corresponding time series in Figure 4. As shown in Tables IV and V, expected excess returns are particularly informative for predicting future realized returns once we control for the bond–stock beta. We therefore construct residual expected excess returns, $\eta_{n,t}$, from equation (6), using the maturity-specific regressions based on the moving-average expected excess returns. Figure 4 plots the lagged residual series $\eta_{n,t}$ together with the realized excess returns over the subsequent twelve months. Both series are standardized

Figure 4: Realized Excess Returns and Lagged Residual Expected Excess Returns



Note: The figure plots realized excess returns together with the lagged, residualized expected excess returns from equation (6), where expected excess returns are smoothed using a 12-month moving average. Both series are standardized to facilitate comparison within and across panels. Realized excess returns are shown in black, while the residualized expected excess returns are shown in dashed red. Shaded regions indicate NBER recession periods. The sample spans January 1988 to March 2024.

to facilitate comparison within and across panels. Realized excess returns appear in black, while the residual expected excess returns are shown in red. [12]

The figure shows that the relationship between realized and expected excess returns is consistent throughout the sample. Forecasters correctly foresaw high returns in the early 1990s, early 2000s, and in the aftermath of the Global Financial Crisis. They correctly foresaw low returns in the late 1980s, mid 2000s, and mid 2010s.

3. International Panel Evidence

We now present the data and evidence for the international panel of twelve developed countries: Canada, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, the U.K., and the U.S.

3.1 International Data and Summary Statistics

We estimate subjective expected excess returns following equation (1). To obtain bond yield forecasts for the international panel, we turn to Consensus Economics. For each month, we use the average forecast of the ten-year yield twelve months ahead. This is the only bond maturity consistently reported across countries in Consensus Economics, giving a month-by-country panel. Consensus Economics reports the exact survey date, which typically falls

[12] Appendix Figure A.3 instead plots realized returns alongside the lagged moving-average expected excess returns, without residualizing them with respect to the bond–stock beta.

Table V: Realized and Expected Excess Returns: U.S., Individual Maturities

Panel (A): Realized and Expected Excess Returns

	3-year				6-year				11-year			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Subjective XR	0.64*		0.99**	0.29	0.67***		1.01***	0.37	0.58**		0.92***	0.34
	(0.37)		(0.43)	(0.49)	(0.25)		(0.26)	(0.36)	(0.24)		(0.23)	(0.25)
$\lambda\times$ Beta		0.23	-0.77	-1.76**		0.14	-0.88*	-2.18*		0.20	-0.74**	-1.72*
		(0.50)	(0.54)	(0.85)		(0.47)	(0.49)	(1.15)		(0.35)	(0.34)	(1.03)
Level				1.38**				2.81**				4.15**
				(0.65)				(1.35)				(2.02)
Slope				-0.22				-0.94				-2.17**
				(0.31)				(0.64)				(1.02)
Curvature				0.31				0.77				0.97
				(0.26)				(0.56)				(0.91)
Constant	0.99**	1.02**	1.04**	1.16***	2.17***	2.13***	2.18***	2.03***	3.02***	2.95***	3.04***	2.77***
	(0.42)	(0.45)	(0.43)	(0.39)	(0.74)	(0.77)	(0.71)	(0.64)	(1.03)	(1.07)	(0.98)	(0.93)
R-squared	0.05	0.00	0.07	0.23	0.06	0.00	0.09	0.26	0.06	0.00	0.08	0.26
Observations	423	423	423	423	423	423	423	423	423	423	423	423

Panel (B): Realized Expected Excess Returns (MA)

	3-year				6-year				11-year			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Subjective XR (MA)	0.93**		2.19***	0.85	0.96***		2.03***	0.93	0.79***		1.77***	0.72**
	(0.45)		(0.63)	(1.04)	(0.31)		(0.40)	(0.75)	(0.29)		(0.34)	(0.36)
$\lambda\times$ Beta		0.23	-1.96**	-2.38**		0.14	-1.90***	-2.84**		0.20	-1.61***	-2.14**
		(0.50)	(0.82)	(0.96)		(0.47)	(0.65)	(1.24)		(0.35)	(0.42)	(1.07)
Level				1.49**				3.03**				4.33*
				(0.73)				(1.46)				(2.27)
Slope				-0.15				-0.66				-1.80
				(0.38)				(0.80)				(1.10)
Curvature				0.26				0.60				0.80
				(0.24)				(0.52)				(0.90)
Constant	1.01**	1.02**	1.08***	1.26***	2.23***	2.13***	2.21***	2.18***	3.04***	2.95***	3.04***	2.92***
	(0.43)	(0.45)	(0.42)	(0.38)	(0.75)	(0.77)	(0.68)	(0.63)	(1.03)	(1.07)	(0.92)	(0.92)
R-squared	0.07	0.00	0.14	0.28	0.09	0.00	0.17	0.30	0.09	0.00	0.17	0.29
Observations	412	423	412	412	412	423	412	412	412	423	412	412

Note: The table reports the estimates of equation (7):

$$xr^{par}_{n,t\to t+12} = b_0 + b_1\,\tilde{E}xr_{n,t\to t+12} + b_2\,(\hat{\lambda}\times\hat{\beta}^{w}_{n,t}) + X'_{n,t}\delta + \varepsilon_{n,t+12},$$

estimated maturity by maturity. All independent variables are lagged by twelve months. $\lambda \times Beta$ is the fitted value from a regression of the expected excess return on the five-year bond-stock beta, as specified in equation (6). The beta is therefore rescaled so that its magnitude is comparable to the expected excess return. Panel A uses expected excess returns, while Panel B uses moving average expected excess returns. The moving average expected excess return is the average expected excess return over the past year. The sample consists of monthly data from January 1988 to March 2024. Driscoll-Kraay standard errors with 27 lags are reported in parentheses.

on the second or third Monday of each month.[13] The Consensus Economics sample begins

[13]The definition of the Consensus Economics forecasts is largely comparable to that of the Blue Chip forecasts described above. Consistent with this, for the U.S. the two series are strongly correlated, with a correlation of 0.99 (see Appendix Figure A.4). Minor differences may arise because: (i) Consensus Economics elicits forecasts of the yield twelve months ahead, whereas Blue Chip reports forecasts of the average yield in the quarter four quarters ahead; (ii) survey dates may differ slightly within the month; and (iii) the sets of surveyed economists 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 month-end ten-year realized government bond yields, one-year government bond yields, T-bill rates, and total stock return indexes from the Global Financial Database (GFD).[14] For all countries except the U.S., we use the 10-year bond yield as a substitute for the 11-year bond yield in equation (1). 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.[15] To proxy the risk-free rate in equation (1), we use the one-year government bond yield. For a small subset of countries, the one-year government bond yield is not available over the full sample. In those cases, we instead use the three-month T-bill rate.[16]

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. We compute realized returns on the ten-year government bond using equation (4), based on one-month holding-period returns rather than daily returns due to data availability in the international sample. To compute monthly excess returns in the beta estimation, we use the T-bill rate as the risk-free rate. We then estimate equation (5) separately for each country. As only monthly data are available, bond–stock betas are estimated using a five-year rolling window (60 months), since shorter rolling windows would lead to substantial estimation noise. 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.[17]

International expected bond excess returns and rolling betas are plotted in Figure 5, and summary statistics are reported in Table VI. In each panel of Figure 5, the dashed gray line (left-hand y-axis) represents subjective expected excess returns, while the solid black line (right-hand y-axis) represents the beta for a specific country. The figure illustrates the usefulness of bringing in data from other countries. First, a close correlation between expected bond excess returns is visible consistently across countries. For example, the uni-

[14]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 downloaded from Datastream.

[15]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.

[16]The one-year rate is not available for Canada until 2007, for the Netherlands until 1999, and for Norway until 2003.

[17]Summary statistics for betas with respect to U.S. stock excess returns are reported in Appendix Table A7.

Table VI: Summary Statistics: International Panel

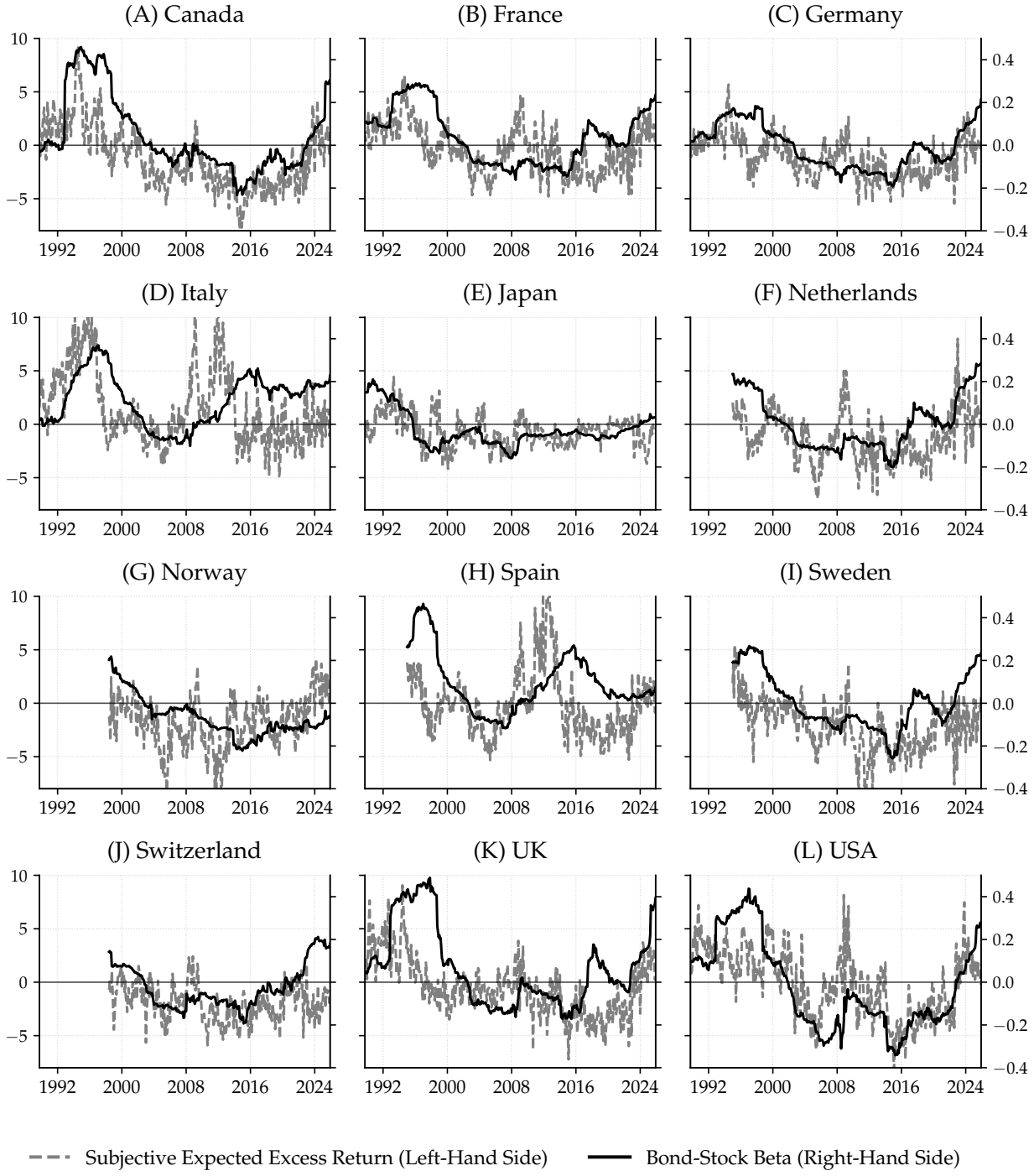
Panel (A): Subjective Expected Excess Returns

	Pooled	Canada	France	Germany	Italy	Japan	Netherlands	Norway	Spain	Sweden	Switzerland	UK	USA
Average	-0.73	-1.23	-0.30	-1.10	1.50	-0.67	-1.21	-1.76	-0.07	-2.14	-1.89	-0.51	-0.04
Std	2.70	2.92	2.21	1.86	3.40	1.43	2.36	2.36	3.06	2.35	1.68	2.62	3.12
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	4823	435	435	435	435	435	372	331	372	372	331	435	435
First Year	1989	1989	1989	1989	1989	1989	1994	1998	1994	1994	1998	1989	1989
Panel (B): Subjective Expected Excess Returns (Moving Average)													
Average	-0.72	-1.22	-0.30	-1.10	1.53	-0.67	-1.21	-1.80	-0.04	-2.09	-1.88	-0.51	-0.03
Std	2.35	2.61	1.82	1.51	3.00	1.10	1.92	1.80	2.69	1.95	1.21	2.27	2.71
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	1.94	2.06	1.90	1.67	2.62	1.61	2.06	0.95	2.74	2.42	1.67	1.54	1.91
Std	6.95	6.30	6.82	6.55	9.68	4.88	6.63	5.89	8.45	7.36	5.07	7.17	6.76
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	-19.87
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	15.54
Panel (D): Bond-Stock Betas													
Average	0.02	0.04	0.04	-0.01	0.12	-0.03	0.00	-0.07	0.09	0.00	-0.02	0.07	-0.02
Std	0.15	0.18	0.13	0.10	0.12	0.08	0.12	0.09	0.14	0.13	0.10	0.18	0.21
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	0.21	0.37	0.21	0.29	0.22	0.46	0.27	0.21	0.49	0.44
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

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), and bond-stock betas (Panel D) for different countries. Subjective expected bond excess returns are computed according to Equation 1. We use four-quarter par bond yield forecasts from Consensus Economics. 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 5. 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, the average HW-robust standard errors, and the number of observations. The sample is monthly and spans September 1989 to December 2024.

variate correlation 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 this was matched by a correspondingly earlier switch to negative expected bond excess returns.

Figure 5: Expected Excess Returns and Betas by Country



Note: The figure reports estimated subjective expected excess returns and betas for each country. The sample runs from September 1989 through December 2024. We compute subjective 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 subjective 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 five-year (60 months) rolling window. Consensus Economics data are not available since the start of the sample for a number of countries: the Netherlands, Norway, Spain, Sweden, and Switzerland. We use the same x-axis and y-axis for all panels.

3.2 International Expected Excess Returns vs. Betas

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

$$\tilde{E}_t xr_{c,t\to t+12} = \alpha + \lambda \hat{\beta}_{c,t} + X'_{c,t}\delta + \eta_{c,t}, \tag{8}$$

where c indexes countries, $E_t xr_{c,t\to t+12}$ denotes subjective expected excess return for the eleven-year bond in country c over a twelve-month forecast horizon, $\beta_{c,t}$ is the five-year rolling beta estimated using information available at time t, and $X_{c,t}$ is a vector of control variables. The specification mirrors equation (6), with the cross-sectional dimension given by countries rather than maturities. As before, the CAPM predicts a positive coefficient λ equal to the equity premium, and a zero intercept α.

Table VII shows that the estimated relationship between bond-stock betas and expected bond excess returns is indeed 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 8.3. Its magnitude is comparable to the U.S. results reported in Table II and on the order of the equity premium. The standard deviation of beta in the sample is 0.15, while the standard deviation of subjective expected excess returns is 2.70. A one-standard-deviation increase in the beta is associated with a 2.25=0.15 $\times$ 8.3 percentage point increase in the subjective expected excess return, or slightly less than one standard deviation. The estimated constant in column (1) is statistically significantly negative, different from what the CAPM predicts. However, we need to keep in mind that in the international panel, we approximate the 11-year yield in the expected excess return equation (1) by a 10-year yield. If the actual par yield $y_{n,t}^{par}$ is under-estimated by 10 bps and the bond duration equals 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, even when the actual CAPM alpha equals zero.[18]

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 a linear time trend, which has little effect on the results, yielding a slightly smaller coefficient of 7.1. In Column (3), we additionally include country fixed effects, which absorb time–invariant country–specific characteristics, such as persistent differences in risk premia. The estimated coefficient is slightly smaller and equal to 7.7. With country fixed effects, identification comes from within–country variation over time, indicating that changes in beta within a country

[18] On average, the U.S. 11-year par yield exceeds the U.S. 10-year par yield by 7 bps, but the difference may be larger in the international sample where bonds tend to be riskier.

are associated with changes in subjective expected excess returns. In Column (4), we add the rolling variance of bond excess returns as a control, where the variance is estimated over the same window and using the same data frequency as bond-stock betas. One might expect that higher bond return variance goes along with higher bond risk premia if the marginal bond investor is primarily invested in bonds. In the international panel, we again find a positive coefficient on the bond return variance, similar to the U.S. regression in Table II, though it is smaller and not statistically distinguishable from zero in the international panel. The coefficient on bond-stock betas and the R-squared are again little changed.

Columns (5) and (6) add term structure controls, showing that this again leaves the coefficient on betas unchanged. In Column (5), we add the term spread. Column (6) separately controls for the one-year and ten-year bond yields, since level, slope and curvature are not available for our international panel. The coefficient on beta remains unchanged, while the coefficient on the term spread is close to zero. This suggests that the term spread does not contain information beyond the beta in explaining subjective expected excess returns. A likely reason why the term spread does not explain expected bond excess returns controlling for bond-stock betas is that variation in the term spread conflates the expected path of interest rates with compensation for risk.[19] By contrast, subjective expected bond excess returns adjust for the expected path of interest rates, providing a more direct measure of bond risk premia than the term spread.

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 time fixed effects absorb common time-series variation across countries. 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 subjective 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, though it remains highly statistically significant and about half as large as the equity premium. In this specification, a large share of the variation is absorbed by fixed effects, so maybe 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

[19] A long literature has found that the term spread's ability to predict bond excess returns is not as strong internationally as in the U.S. The term spread appears to predict bond excess returns better in countries and time periods when interest rates are better approximated by a random walk, which would lead the expected path component to drop out of the term spread (Mankiw et al., 1987; Hardouvelis, 1994).

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

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Bond-Stock Beta	8.32***	7.07***	7.71***	6.62***	7.75***	3.92***	5.75***	3.81***
	(1.66)	(1.66)	(1.58)	(1.62)	(1.56)	(1.46)	(1.90)	(1.16)
Bond Variance				0.15				
				(0.11)				
Term Spread					0.18			
					(0.17)			
10-Year Rate						0.70***		
						(0.21)		
1-Year Rate						-0.20		
						(0.18)		
Constant	-0.90***							
	(0.26)							
Time Trend		✓						
Country FE			✓	✓	✓	✓		✓
Time FE							✓	✓
R-squared	0.21	0.25	0.28	0.29	0.29	0.45	0.55	0.62
Observations	4822	4822	4822	4822	4822	4822	4822	4822

Note: The table reports the results of equation (8):

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

estimated using our international panel. The bond-stock beta is estimated using monthly data over a five-year rolling window (60 months) and using domestic stock market excess returns. Expected yields are from Consensus Economics, while bond yields, T-bill rates, and stock returns are from the Global Financial Database (GFD). The term spread is the difference between the ten-year government bond yield and the one-year government bond yield. 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.

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 A5 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 the international panel regression 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 (8) 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 6, 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 exceeds 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, where it remains positive but is not always statistically significant.

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 subjective expected returns, as can be seen in Figure 5. For both countries, betas and subjective expected excess returns display a very strong comovement from 1990 to 2014. After 2014, this relationship weakens: the beta remains positive, while subjective expected excess returns turn negative. This break coincides with the launch of the ECB's quantitative easing (QE) program in January 2015, through the Public Sector Purchase Programme. One interpretation is that QE substantially lowered sovereign spreads and term premia in these countries, leading to persistently low expected returns. We explore the relationship between expected excess returns and QE further in Section 3.4.

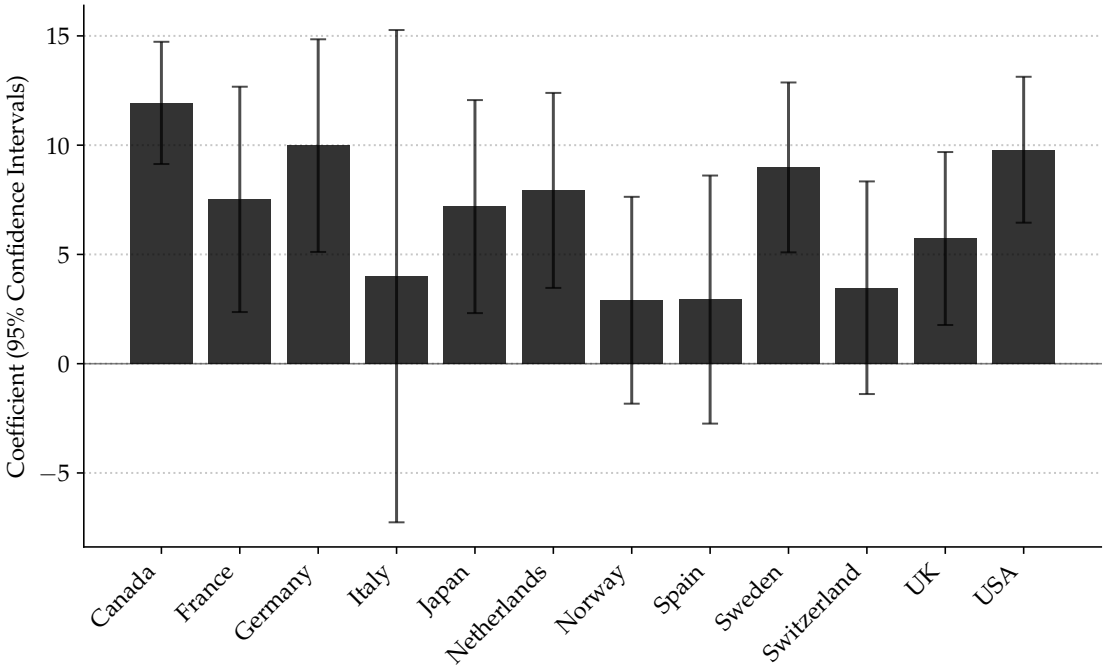
In contrast, Norway and Switzerland display a weaker relationship between expected bond excess returns and bond-stock betas simply because their data start in 1998, when betas were already zero. Lacking any positive betas over their shorter samples, it is unsurprising that the relationship is estimated to be weaker for those two countries.

Figure A.6 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.

3.3 International Realized Excess Return Predictability

Having established a robust relationship between the beta and subjective expected excess returns in the international context, we now extend the analysis of Section 2.3 to the international setting. We examine whether the relationship between the beta and expectations also translates into predictability for realized excess returns. To this end, we estimate predictive

Figure 6: Country-by-Country Expected Excess Returns onto Beta



Note: The figure plots the estimated coefficient from regression (8), 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 VI.

regressions of the following form:

$$xr^{par}_{c,t\to t+12} = b_0 + b_1\, \tilde{E}_t xr_{c,t\to t+12} + b_2\, (\hat{\lambda} \times \hat{\beta}^w_{c,t}) + X'_{c,t}\delta + \varepsilon_{c,t+12}. \tag{9}$$

Here, $\hat{\beta}^w_{c,t}$ denotes the fitted value from the regression of subjective expected excess returns on betas, as specified in equation (8). The regression is estimated in a panel setting with countries as the cross-sectional dimension.

Table VIII reports the results. Panel A uses subjective expected excess returns as the independent variable, while Panel B replaces this measure with its twelve–month moving average. In Panel A, Columns (1) and (2) report a simple predictive regression of realized excess returns on subjective expected excess returns without and with country fixed effects. The estimated univariate coefficient is statistically significantly different from zero and around 0.55 in both cases. In the case with country fixed effects, the estimated coefficient is statistically different from zero, but not statistically different from one, as one would expect under FIRE. While the point estimate is hence somewhat smaller than one would expect under FIRE, we cannot necessarily reject FIRE with statistical significance. Overall,

it appears that surveys are informed about international realized bond excess returns.

The next few columns decompose expected bond excess returns into a component due to bond-stock betas and the remaining variation in expected bond excess returns. In Column (3), we replace subjective expected excess returns with the scaled beta. The estimated coefficient equals 0.76, but is not precisely estimated, and we cannot reject the null that it is equal to zero. Nevertheless, the magnitude of the coefficient is similar to that obtained when using subjective expected excess returns directly, suggesting that there may be a positive relationship, though the data on realized international bond excess returns is to noisy to reliably estimate this relationship.

In Column (4), we estimate a regression that includes both subjective expected excess returns and the scaled bond–stock beta. If expected bond excess returns were fully driven by bond-stock betas, the two regressors in this specification would be collinear, and the coefficients could not be estimated. We find that the coefficient on subjective expected excess returns remains very similar to the univariate regression in column (1), while the coefficient on the bond–stock beta declines further and remains statistically insignificant. In Column (5), we further control for the term spread, which enters positively and significantly as in Campbell and Shiller (1991), but leaves the other coefficients unchanged.

Panel B of Table VIII reports the results when subjective expected excess returns are replaced by their twelve–month moving average, finding an even larger magnitude for the coefficient on subjective expected excess returns. The estimated coefficient increases to 0.72, close to the FIRE baseline of one. Overall, international surveys are informative for future bond excess returns, and we cannot reject that forecasters are rational. However, the most significant predictability of realized bond excess returns arises from the component in expected bond excess returns that is orthogonal to betas. Similar to the U.S., one possible interpretation is that forecasters have information about international bond returns, but because betas move at low frequency, the sample is too short to detect a significant relationship between betas and necessarily noisy realized bond excess returns.

3.4 Quantitative Easing in the International Panel

So far, we have seen that bond risk premia are closely linked to their betas with respect to the aggregate stock market. In this section, we build on this baseline result to shed light on the channels of Quantitative Easing (QE) using our international panel.

While QE is now a standard tool of central banks, its transmission mechanisms remain debated (Tenreyro and Wazzi, 2025). First, under the signaling channel, QE may act by credibly committing the central bank to lower future short rates. In that case, long-term yields fall through the expectations hypothesis, with no effect on risk premia. Second,

Table VIII: Realized Returns, Expected Excess Returns, and Betas: International Evidence

Panel (A): Realized and Expected Excess Returns

	(1)	(2)	(3)	(4)	(5)
Subjective Excess Returns	0.53***	0.57***		0.53***	0.46**
	(0.20)	(0.21)		(0.18)	(0.19)
$\lambda\times$ Beta			0.76	0.23	0.33
			(0.65)	(0.65)	(0.57)
Term Spread					1.91***
					(0.43)
Constant	2.52***				
	(0.93)				
Country FE		✓	✓	✓	✓
R-squared	0.045	0.052	0.021	0.053	0.132
Observations	4677	4677	4677	4677	4677

Panel (B): Realized and Expected Excess Returns (MA)

	(1)	(2)	(3)	(4)	(5)
Subj. Excess Returns (MA)	0.72***	0.81***		0.85***	0.69**
	(0.26)	(0.28)		(0.28)	(0.29)
$\lambda\times$ Beta			0.76	-0.17	0.04
			(0.65)	(0.67)	(0.61)
Term Spread					1.79***
					(0.41)
Constant	2.65***				
	(0.90)				
Country FE		✓	✓	✓	✓
R-squared	0.063	0.072	0.021	0.073	0.141
Observations	4677	4677	4677	4677	4677

Note: The table reports the results of equation (9) estimated using our international panel:

$$xr^{par}_{c,t\to t+12} = b_0 + b_1\,\tilde{E}_t xr_{c,t\to t+12} + b_2\,(\hat{\lambda}\times\hat{\beta}^{w}_{c,t}) + X'_{c,t}\delta + \varepsilon_{c,t+12},$$

All independent variables are lagged by twelve months. $\lambda \times Beta$ is the fitted value from a regression of the expected excess return on the five-year bond-stock beta, as specified in equation (6). The beta is therefore rescaled so that its magnitude is comparable to the expected excess return. Panel A uses expected excess returns, while Panel B uses moving average expected excess returns. The moving average expected excess return is the average expected excess return over the past year. 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 portfolio balance channel predicts that by removing duration risk from private investors' balance sheets, QE lowers the price of bond risk and thus the compensation investors require to hold long-duration bonds, especially when bond returns are more volatile (Tobin, 1958, 1969; Vayanos and Vila, 2021). Under a broader interpretation of this channel, QE may also reduce the price of bond–stock beta risk by raising investor wealth and lowering aggregate risk aversion, as in intermediary-based models (Kekre et al., 2024), and similar to habit-based frameworks (Campbell et al., 2020). Third, QE may affect bond market liquidity, thereby lowering bond yields and risk premia by reducing market stress. Finally, QE may be perceived as state-contingent if investors expect similar interventions in future crises (Haddad et al., 2025). Under this channel, QE need not change the price of bond–stock beta. Instead, it changes bond betas themselves, and therefore bond risk premia. Distinguishing among these channels is important because they imply that QE is more effective under different market conditions.

The international panel is advantageous because different countries initiated QE at different times, allowing us to exploit both cross-country and time-series variation. We construct an indicator variable $\mathrm{QE}_{c,t}$ that equals one during periods in which QE is active in country c at time t. QE was implemented in Italy, France, Germany, the Netherlands, Spain, the U.S., the U.K., Japan, Canada, and Sweden. The remaining countries, Norway and Switzerland, are included in the panel with their QE dummies set to zero.[20] We define QE periods based on central bank announcement dates. Specifically, the start date corresponds to the initial QE announcement, while the end date is given by the announcement that QE would be terminated or that quantitative tightening (QT) would begin. If multiple programs overlap, we consider QE to be active as long as at least one program remains in place.[21] Of course, even in our international sample, the variation in QE announcement dates is limited, and QE implementation has varied from country to country, so future QE episodes may differ from the historical experience.

We test for a direct effect of QE on expected bond excess returns and whether QE changes the prices of beta and volatility risk through panel regressions of the form:

$$\tilde{E}_t xr_{c,t\to t+12} = \alpha + \alpha_{QE}\mathrm{QE}_{c,t} + \lambda\hat{\beta}_{c,t} + \lambda_{QE}\left(\mathrm{QE}_{c,t}\times\hat{\beta}_{c,t}\right) + \mu\hat{\sigma}^2_{c,t} + \mu_{QE}\left(\mathrm{QE}_{c,t}\times\hat{\sigma}^2_{c,t}\right) + \eta_{c,t}, \tag{10}$$

where c indexes countries, and $\hat{\sigma}^2_{c,t}$ is the rolling bond return variance over a five-year moving

[20] While the Swiss National Bank had a program purchasing foreign currency, it did not have a program purchasing Swiss government bonds. We therefore set the Switzerland QE dummy to zero.

[21] For example, the ECB's Asset Purchase Programme (APP) and Pandemic Emergency Purchase Programme (PEPP).

window. As before, $\hat{\beta}_{c,t}$ denotes the five-year rolling bond–stock beta. The four different channels have distinct predictions for the coefficients in this regression. The signaling channel predicts that QE has no effect on risk premia, i.e., $\alpha_{QE} = 0$, $\lambda_{QE} = 0$, and $\mu_{QE} = 0$. The portfolio balance channel predicts $\lambda_{QE} < 0$, $\mu_{QE} < 0$ and $\alpha_{QE} = 0$. Under the liquidity channel, we would expect $\alpha_{QE} < 0$ with $\lambda_{QE} \approx \mu_{QE} \approx 0$. Finally, under the state-contingent QE channel, we would expect $\alpha_{QE} < 0$ in a univariate regression on the QE dummy, but this should disappear once $\hat{\beta}_{c,t}$ and $\hat{\sigma}^2_{c,t}$ are controlled for. Of course, this classification is highly simplified, and the channels sometimes overlap, with the differentiation between liquidity and portfolio balance channels somewhat fluid.

The results are displayed in Table IX. For all specifications, we use the twelve-month moving average of subjective expected returns to reduce noise.[22] In column (1), we regress expected excess returns on the QE dummy and obtain a negative and statistically significant coefficient of -1.2. This estimate implies that QE reduced expected excess returns by 1.2 percentage points per year. However, this result may simply reflect the fact that QE was enacted in a period of generally low bond risk premia. This is visible in the small and insignificant dummy coefficient in column (2), which includes country and time fixed effects.

In column (3), we extend the regression to include beta and its interaction with the QE dummy. The base coefficient on beta, λ, is similarly large and significant as in our baseline regressions at 9.4. In addition, column (2) shows a large and negative interaction coefficient between beta and the QE dummy. To assess whether this relationship effectively disappears, we test whether the implied slope during QE, $\lambda + \lambda_{QE}$, is equal to zero using a Wald test of the linear restriction $\lambda + \lambda_{QE} = 0$. We find that we cannot reject this hypothesis (p-value= 0.758), implying that the pricing of beta is statistically indistinguishable from zero during QE periods. The results for λ and λ_{QE} in column (3) carry over to column (4), which additionally includes country and time fixed effects. Overall, we find a lower price of risk for bond-stock betas during QE periods, in line with the broader portfolio balance channel.

Next, we study how the price of bond return variance changes during QE. Column (5), shows that an increase in bond return volatility is associated with higher expected bond excess returns. The coefficient μ is equal to 0.6 and is statistically significant. In addition, the interaction coefficient between bond return variance and the QE dummy is strongly negative and significant, fully offsetting the positive effect of bond return variance during non-QE periods. Consistent with this pattern, we cannot reject that the implied slope during QE is equal to zero (p-value=0.911). Column (6) adds country and time fixed effects and finds similar results. However, different from column (4), the QE dummy enters with a positive and significant coefficient in column (6).

[22] Results using raw subjective expected returns are similar and are reported in Appendix Table A9.

Column (7) provides our full and preferred specification, including beta, bond return variance, and their interactions with QE. The price of bond-stock beta risk during non-QE periods remains statistically significant and economically large at 7.1, and the coefficient on bond return variance is also positive and highly statistically significant. The interaction coefficients, on the other hand, are both negative and similar in magnitude to the baseline coefficients, indicating a substantial reduction in the price of bond-stock beta risk and bond return variance risk during QE periods. The coefficient on the QE dummy, α_{QE}, is small and statistically indistinguishable from zero, as one would expect if a combination of changing prices and quantities of risk explains QE's effect on bond risk premia.

Column (8) reports results for the same specification with both country and time fixed effects. The estimates for the prices of risk and their interactions are qualitatively and quantitatively similar, though the direct QE dummy now enters positively and significantly. One possibility is that in this relatively rich specification with bond-stock betas and bond return variances, the difference in timing of QE across countries leaves us with little variation to identify α_{QE}. For this reason, we prefer the estimates in column (7). Overall, our results indicate that a combination of the changing price of risk – as predicted by the portfolio balance channel – combined with changing quantities of risk – as predicted by the state-contingent policy channel – can account for bond risk premia during QE periods.

3.4.1 QE and Betas

The results discussed so far suggest that QE may weaken the relationship between bond–stock betas (and bond return variance) and subjective excess bond returns. A natural question is whether QE also affects the bond–stock betas themselves. In this subsection, we provide exploratory evidence on the evolution of betas during QE periods. Broadly, we find that bond-stock betas changed substantially during QE periods, with both negative and positive bond-stock betas compressed towards zero. Combined with the reduction in the price of risk documented in the previous subsection, this further implies that QE would tend to reduce bond risk premia more strongly for countries whose bonds are risky with positive bond-stock betas than in countries whose bonds are hedges.

Figure 7 visualizes bond–stock betas for the main countries that have experienced QE, with QE periods indicated by shaded areas. In what follows, we place particular emphasis on the first QE announcement, which—consistent with the findings of Haddad et al. (2025)—appears to be associated with the most pronounced movements in financial markets.

Panel A of Figure 7 plots the U.S. bond–stock beta. We report three estimates based on 5-year rolling windows: one constructed using daily returns (black line), one using daily returns with the Dimson (1979) correction (green), and one based on monthly returns (red).

Table IX: Quantitative Easing: International Panel

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
QE	-1.18***	0.16	-0.92***	-0.32	1.17**	1.92***	0.20	1.37***
	(0.39)	(0.32)	(0.21)	(0.40)	(0.52)	(0.39)	(0.53)	(0.47)
Beta			9.42***	5.88***			7.05***	4.45***
			(1.61)	(1.74)			(1.53)	(1.43)
QE × Beta			-8.85***	-6.89***			-6.41***	-3.17
			(2.22)	(2.66)			(2.37)	(2.78)
Bond Variance					0.58***	0.31***	0.28***	0.23***
					(0.10)	(0.10)	(0.09)	(0.09)
QE × Bond Variance					-0.58***	-0.49***	-0.29**	-0.42***
					(0.13)	(0.14)	(0.13)	(0.11)
Constant	-0.52		-0.77***		-2.89***		-1.86***	
	(0.39)		(0.23)		(0.53)		(0.45)	
Country FE		✓		✓		✓		✓
Time FE		✓		✓		✓		✓
R-squared	0.03	0.62	0.35	0.66	0.27	0.66	0.39	0.68
Observations	4822	4822	4822	4822	4822	4822	4822	4822

Note: The table reports the results of equation (10) estimated using our international panel:

$$\tilde{E}_t xr_{c,t\to t+12} = \alpha + \alpha_{QE}\mathrm{QE}_{c,t} + \lambda\hat{\beta}_{c,t} + \lambda_{QE}\left(\mathrm{QE}_{c,t}\times\hat{\beta}_{c,t}\right) + \mu\hat{\sigma}^2_{c,t} + \mu_{QE}\left(\mathrm{QE}_{c,t}\times\hat{\sigma}^2_{c,t}\right) + \eta_{c,t}$$

The beta is estimated using monthly data over a five-year rolling window (60 months), based on excess returns on the domestic stock market. Similarly, bond variance is estimated as the rolling variance of monthly bond excess returns over the same five-year window. Expected yields are from Consensus Economics, while bond yields, T-bill rates, and stock returns are from the Global Financial Database (GFD). The term spread is the difference between the ten-year government bond yield and the one-year government bond yield. 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.

All three series display an increase around the QE announcement, though the magnitude differs somewhat across measures. The change appears particularly pronounced for the beta estimated using monthly returns, while the beta based on daily returns shows only a modest increase. The Dimson-corrected beta lies between these two. These patterns suggest that returns during QE were more positively correlated across asset classes than normally, which could occur if QE news raise equity and bond prices simultaneously.

We now turn to the Euro area in Panel B, which provides the most striking evidence of changing betas during QE. The figure plots bond–stock betas for the five Euro area countries in our sample: Spain (green) and Italy (red) exhibit markedly higher betas than France (dashed), the Netherlands (gray), and Germany (black).

Before turning to QE, it is useful to describe the broader evolution of betas in the Euro area. In the 1990s, there was substantial cross-country heterogeneity in bond-stock betas. However, as the introduction of the euro approached in the late 1990s, these differences narrowed, and betas converged. By the early 2000s, bond–stock betas are virtually identical across countries. This convergence likely reflected the convergence of monetary policy and the perceived mutualization of sovereign credit risk. With investors treating Euro area sovereign bonds as near-perfect substitutes, their comovement with equity returns became effectively identical.

This convergence in Euro area betas reversed sharply following the global financial crisis. Starting in 2008, betas began to diverge, with the divergence intensifying during the sovereign debt crisis. Over this period, the bond–stock betas of core countries (Germany, the Netherlands, and France) become increasingly negative, reaching values around -0.2 by the end of 2014. In contrast, the betas of peripheral countries (Italy and Spain) turn positive, rising to approximately 0.3 over the same period. The divergence is consistent with flight-to-safety dynamics. During periods of equity market declines or adverse macroeconomic news, investors reallocated away from peripheral bonds toward core sovereign debt. ECB unconventional policies—including the Long-Term Refinancing Operations (LTROs), the Securities Markets Programme (SMP), and the "whatever it takes" speech associated with the Outright Monetary Transactions (OMT) framework—did not materially alter this divergence in betas, even though they did affect the *levels* of bond yields. According to our benchmark results, this divergence in bond-stock betas across core and periphery has direct pricing implications—it implies that core bonds commanded a negative risk premium, while peripheral bonds required a positive premium.

Finally, a sharp reversal in Euro area bond-stock betas occurred in 2015, following the announcement of the Public Sector Purchase Programme (PSPP). At this point, peripheral bond–stock betas began to decline, while core betas increased. In other words, the ECB's QE announcement appeared to halt the divergence in bond–stock betas. This pattern is consistent with the interpretation in Haddad et al. (2025), whereby central bank purchases provide support to bond prices in adverse states of the world. It is also consistent with Leombroni et al. (2021), who document that between 2009 and 2014 the response of sovereign yields to ECB policy was highly heterogeneous across countries, while the announcement of QE restored a more uniform transmission mechanism. While the precise channel driving bond-stock betas is not crucial for our core results, one potential transmission mechanism is that ECB QE entailed fiscal redistribution across euro-area countries (Chien et al., 2025). In the Euro area context, it appears that QE reduced the extent to which peripheral bonds behaved like risky assets in bad states.

Panels (C) and (D) report results for the U.K. and Japan, showing patterns consistent with the U.S. For the U.K., the bond–stock beta increased sharply at the onset of QE in March 2009. For Japan, the compression of the bond–stock beta to zero during QE is more gradual, but still visible. Moreover, QE in Japan partly took the form of yield curve control, which mechanically dampens bond yield fluctuations and, in turn, tends to compress the bond–stock beta toward zero. We complement the descriptive evidence with a more formal panel analysis relating bond–stock betas to QE episodes. The results are displayed in Appendix Table A10. Consistent with the patterns discussed above, QE is associated with a compression of bond–stock betas toward zero.

Overall, the evidence suggests that the effects of QE on bond-stock betas are substantial but vary across countries, depending critically on the institutional and macroeconomic context. For this reason, future QE programs may have different implications for bond risks than past ones. To the extent that QE affects bond risks in either direction, this illustrates the importance of understanding the link between bond-stock betas and bond risk premia, as we do in our benchmark results.

4. Implications for Long-Term Bond Premia

Our findings have implications for the decomposition of long-term bond yields into a term premium and expected future short rates, thereby shedding light on the decades-long decline in U.S. long-term bond yields. Define the risk-neutral n-year zero-coupon yield as the yield that would hold under the expectations hypothesis (EH)

$$y_{n,t}^{rn} = \frac{1}{n}\tilde{E}_t \sum_{i=0}^{n-1} y_{1,t+12\times i}. \tag{11}$$

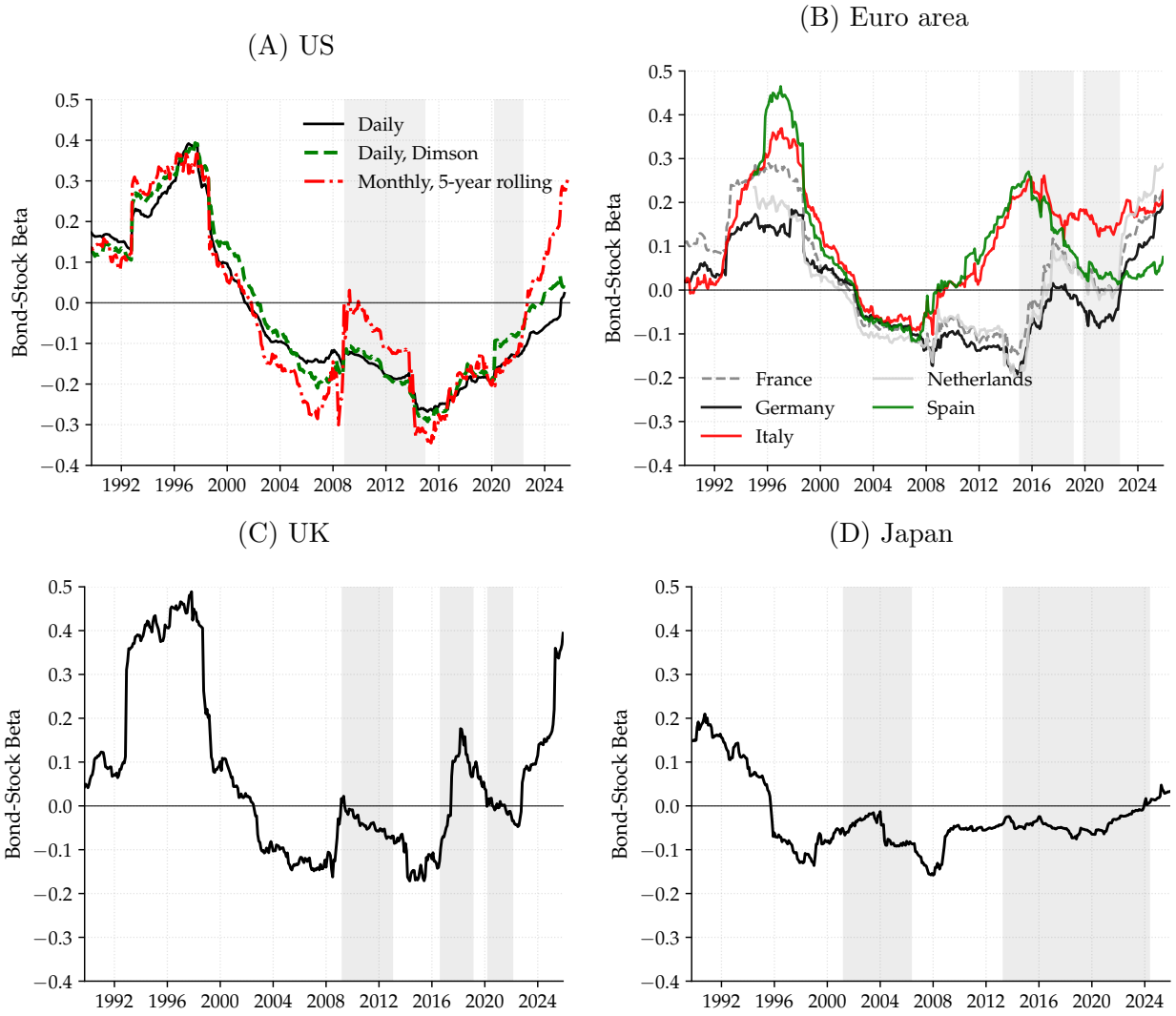
This is the long-term bond yield that should prevail if investors simply price the expected path of future short rates. We make the simplifying assumption that the one-year zero-coupon yield does not contain a risk premium, which is likely to be a good approximation in practice.

We define the term premium component of the n-year zero-coupon bond yield as the difference between the actual yield and its risk-neutral counterpart. This can be written in terms of the average expected excess return over the lifetime of the bond:

$$y_{n,t}^{rp} \equiv y_{n,t} - y_{n,t}^{rn} = \frac{1}{n}\sum_{i=0}^{n-1} \tilde{E}_t xr_{n-i,t+12\times(i+1)}. \tag{12}$$

We next construct estimates of the expected beta at different forecast horizons. We use a

Figure 7: Bond-Stock Betas and QE

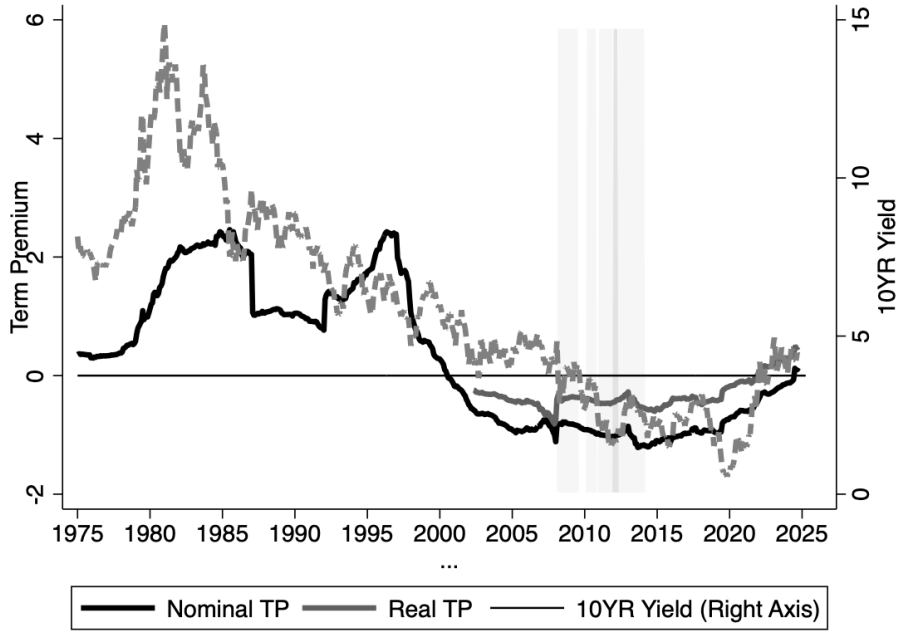


Note: The figures plot the bond-stock betas for different countries, with the shaded area representing QE Periods. Bond-stock betas are estimated using monthly bond and stock excess returns from the Global Financial Database. We use a five-year (60 months) rolling window. The sample runs from 1989 to 2024. For the U.S., in Panel A, we show three estimates based on 5-year rolling windows: one constructed using daily returns (black line), one using daily returns with the Dimson (1979) correction (green), and one based on monthly returns (red). Panel B plot the bond-stock betas for the five Euro area countries in our sample: Spain (green) and Italy (red) exhibit markedly higher betas than France (dashed), the Netherlands (gray), and Germany (black). Panel C shows the bond-stock beta for the U.K., and Panel D for Japan.

simple econometric procedure that exploits their empirical persistence. Realized betas for n-year zero-coupon bonds are estimated from daily bond and stock returns over five-year rolling windows, analogously to the par bond-stock betas used in our main analysis. We then estimate the persistence of these betas through a panel regression of the form:

$$\hat{\beta}_{n,t+12} = b + \rho \hat{\beta}_{n,t} + \varepsilon_{t+12}, \tag{13}$$

Figure 8: Implied U.S. Ten-Year Term Premium



Note: This figure reports the estimated zero-coupon nominal Treasury bond term premia estimated via equation (14) for U.S. ten-year nominal and real bonds from January 1986 to June 2025. The ten-year nominal par yield from Gürkaynak et al. (2007) is shown for comparison on the right y-axis. Quantitative Easing program dates for QE1, QE2, Operation Twist (MEP), QE3 are shaded in gray.

using a monthly panel of bond-stock betas with bond maturities $n = 1, 2, ..., 10$. The regression yields an autoregressive coefficient of $\hat{\rho} = 0.96$ with a Driscoll-Kraay standard error with 27 lags of 0.05 and a constant of $\hat{b} = -0.002$ with a standard error of 0.005. Because the constant is tiny and not statistically different from zero, we set it to zero going forward. Plugging in our baseline estimate gives the following estimate for the n-year zero-coupon term premium:

$$\hat{y}^{rp}_{n,t} = \frac{\hat{\lambda}}{n} \sum_{i=0}^{n-1} \rho^i \hat{\beta}_{n-i,t}. \tag{14}$$

The expression for the risk premium on a real or inflation-indexed bond is constructed analogously, by simply using bond-stock betas inflation-indexed bonds, re-estimating (13).

Figure 8 shows estimated term premia for nominal bonds of different bond maturities, and compares them against the zero-coupon 10-year Treasury yield. Our estimates imply a substantial decline in the estimated ten-year term premium from 2.31% annualized in January 1986 to -1.06% in January 2016, when the Fed lifted off from the zero-lower-bound.

For comparison, over the same time period, the 10-year nominal par yield declined from 9.30% to 2.02%, so the estimated term premium accounts for about half the decline in 10-year yields over these three decades. Our estimates that the nominal term premium then increased by 1.21 percentage points until the end of the sample in June 2025, with the increase in the real term premium equal to 1.03 percentage points. At the same time, the 10-year nominal yield increased by 2.27 percentage points, so the term premium accounts again for half of this more recent increase, with most of the term premium component attributable to the increased real bond risks.

While the estimated overall decline in the 10-year term premium is similar to conventional estimates by Kim and Wright (2011) or Adrian et al. (2013), we differ in terms of the timing of the estimated decline in term premia. Our estimated nominal 10-year term premium already turns negative in 2001, whereas the estimated term premia by Kim and Wright (2011) and Adrian et al. (2013) remain positive until 2011 and 2015, respectively.

Our estimates provide a different interpretation of monetary policy before the Global Financial Crisis. During this period, low long-term yields were sometimes interpreted as a downward-trend in the natural rate or r-star Laubach and Williams (2016), or as the result of a "savings glut" from international investors (Caballero et al., 2008). While those factors likely also played a role, our estimates indicate that U.S. Treasury bonds' improved hedging value also contributed substantially to the decline in long-term bond yields prior to the financial crisis. This matters for monetary policy because – different from a low natural rate – a negative term premium does s not necessarily call for lower short-term policy rates (Caballero and Farhi, 2018).

5. Conclusion

This paper shows that bond risks, as measured by bond–stock betas, are a central determinant of expected bond excess returns. Using survey forecasts to isolate ex-ante expectations, we document a strong and stable relationship between bond–stock betas and subjective expected returns in both U.S. and international data. The estimated market price of risk is economically large and comparable to the equity premium, consistent with a CAPM-style pricing relation for bonds. We also show that subjective expected returns predict realized returns with a coefficient that is indistinguishable from full information rational expectations (FIRE). However, the relationship between realized bond excess returns and betas is weak, potentially reflecting both small-sample issues and bias in interest rate expectations in finite samples.

Our findings provide a new lens on the effects of quantitative easing and the secular decline in long-term yields. By compressing bond risk premia, central bank asset purchases

are more effective at lowering long-term yields in countries with initially riskier government bonds. Our estimated link between bond risk premia and bond-stock betas attributes roughly half of the secular decline in interest rates since the 1980s to term premia. Different from standard estimates, our estimates suggest that term premia already turned negative during the early 2000s, prior to the global financial crisis.

Our findings open several avenues for future research. First, they suggest that term premium estimates should explicitly incorporate time-varying bond–stock comovements. Second, reconciling our results with sources of higher-frequency variation in expected returns—such as monetary policy, supply–demand imbalances, and intermediary constraints—remains an important challenge. Third, our findings raise the question under which market conditions the CAPM fails for the cross-section of stocks, but describes risk premia for government bonds.

References

Adrian, T., R. K. Crump, and E. Moench (2013): "Pricing the term structure with linear regressions," *Journal of Financial Economics*, 110, 110–138.

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.

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

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

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

Caballero, R. J. and E. Farhi (2018): "The Safety Trap," *Review of Economic Studies*, 85, 223–274.

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.

Caballero, R. J. and A. Simsek (2022): "Monetary policy with opinionated markets," *American Economic Review*, 112, 2353–2392.

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.

Campbell, J. Y. and S. B. Thompson (2008): "Predicting excess stock returns out of sample: Can anything beat the historical average?" *Review of Financial Studies*, 21, 1509–1531.

Chien, Y.-L., Z. Jiang, M. Leombroni, and H. Lustig (2025): "What Does It Take? Quantifying Cross-Country Transfers in the Eurozone," NBER Working Paper 34311, National Bureau of Economic Research, available at SSRN: https://ssrn.com/abstract=5568702.

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

Cochrane, J. H. (2005): *Asset Pricing*, Princeton University Press, revised ed.

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*, this version: March 2026.

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

Delao, R. and S. Myers (2021): "Subjective cash flow and discount rate expectations," *The 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.

Du, W., C. Pflueger, and J. Schreger (2020): "Sovereign Debt Portfolios, Bond Risks, and the Credibility of Monetary Policy," *Journal of Finance*, 75, 3091–3137.

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.

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.

Hardouvelis, G. A. (1994): "The Term Structure Spread and Future Changes in Long and Short Rates in the G7 Countries," *Journal of Monetary Economics*, 33, 255–283.

Hillenbrand, S. (2025): "The Secular Decline in Interest Rates," *Review of Financial Studies*.

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.

Kim, D. H. and J. H. Wright (2011): "An arbitrage-free three-factor term structure model and the recent behavior of long-term yields and distant-horizon forward rates," .

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

Krishnamurthy, A., A. Vissing-Jorgensen, et al. (2013): "The ins and outs of LSAPs," in *Kansas City Federal Reserve Symposium*, Federal Reserve Bank of Kansas City Kansas City, 57–111.

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.

Leombroni, M., A. Vedolin, G. Venter, and P. Whelan (2021): "Central bank communication and the yield curve," *Journal of Financial Economics*, 141, 860–880.

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.

Mankiw, N. G., J. A. Miron, and D. N. Weil (1987): "The Adjustment of Expectations to a Change in Regime: A Study of the Founding of the Federal Reserve," *American Economic Review*, 77, 358–374.

Molavi, P., A. Tahbaz-Salehi, and A. Vedolin (2024): "Model Complexity, Expectations, and Asset Prices," *Review of Economic Studies*, 91, 2462–2507.

——— (2025a): "Expectations and the Term Structure of Interest Rates," *Working paper, Northwestern and Boston University*.

Molavi, P., A. Vedolin, and A. Tahbaz-Salehi (2025b): "Expectations and the Term Structure of Interest Rates," *Working Paper, Boston University and Northwestern University*.

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

Pesch, D., K. Whelan, and M. 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*.

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

Ross, S. A. (1976): "The Arbitrage Theory of Capital Asset Pricing," *Journal of Economic Theory*, 13, 341–360.

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*.

Tenreyro, S. and M. Wazzi (2025): "Unconventional Policy Tools at the Fed: Lessons from Theory and Practice," *Working Paper, London School of Economics & University of Oxford. Verkkolähteet*.

Tobin, J. (1958): "Liquidity preference as behavior towards risk," *Review of Economic Studies*, 25, 65–86.

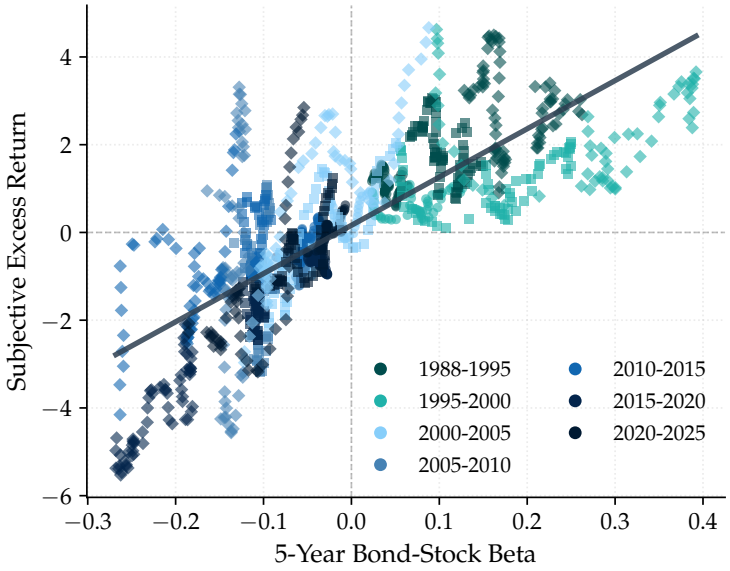
——— (1969): "A general equilibrium approach to monetary theory," *Journal of Money, Credit and Banking*, 1, 15–29.

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

Vayanos, D. and J.-L. Vila (2021): "A preferred-habitat model of the term structure of interest rates," *Econometrica*, 89, 77–112.

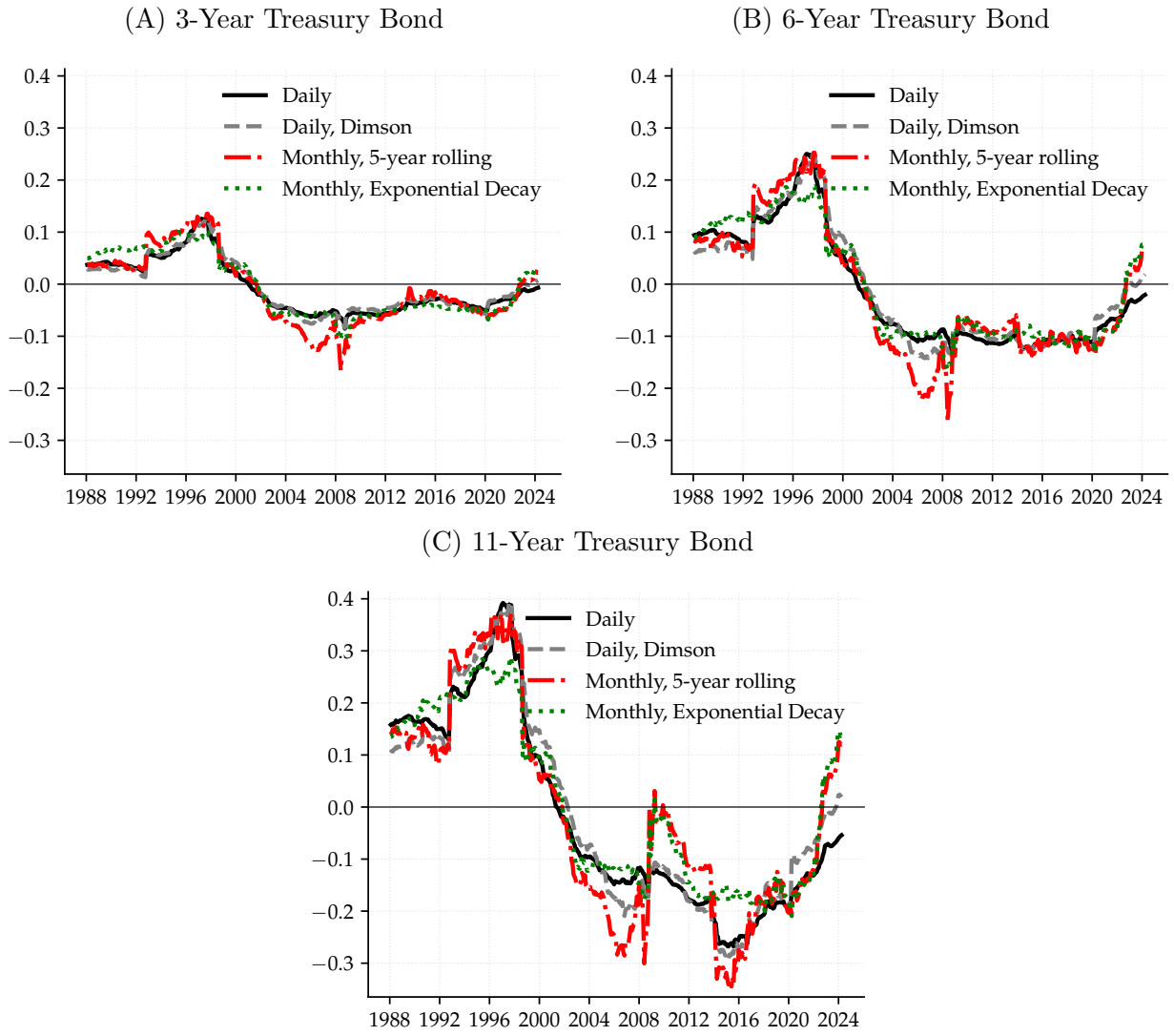
A. Additional Tables and Figures

Figure A.1: Subjective Excess Returns (MA) and Stock Market Betas



Note: This figure plots subjective excess returns on U.S. Treasury bonds (12 months moving average) against their five-year stock-market betas. The figure pools bonds of different maturities, which are distinguished by marker shape: three-year maturities (circles), six-year maturities (squares), and eleven-year maturities (diamonds). The sample spans January 1988 to March 2024.

Figure A.2: Estimated Betas



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 daily par bond excess returns on stock market excess returns. The dashed gray line reports betas estimated using 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 monthly returns instead of daily returns. The green dotted line reports betas estimated 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 A1: Subjective Excess Returns and Stock Market Betas, Alternative Betas

	(1)	(2)	(3)	(4)	(5)	(6)
Beta 5YR	10.87***					
	(1.84)					
Beta 1YR		7.14***				
		(1.55)				
Beta 5YR (Dimson)			10.59***			
			(1.82)			
Beta 1YR (Dimson)				7.13***		
				(1.24)		
Beta 5YR (Monthly)					9.99***	
					(1.62)	
Beta HL 3Y (Monthly)						12.47***
						(1.70)
Constant	0.16	0.02	0.10	0.01	0.16	0.06
	(0.20)	(0.25)	(0.20)	(0.23)	(0.20)	(0.17)
R-squared	0.43	0.29	0.42	0.34	0.44	0.47
Observations	1305	1305	1305	1305	1305	1305

Note: The table reports regression estimates of Equation (6) using alternative measures of beta. The dependent variable is the subjective excess return. Columns (1) and (2) use betas estimated from daily returns over five-year and one-year rolling windows, respectively. Columns (3) and (4) use the corresponding Dimson betas, 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 five-year rolling beta estimated from monthly returns. 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. The sample consists of monthly observations from January 1988 to March 2024.

Table A2: Subjective Excess Returns and Stock Market Betas, Controlling for Standard Deviation

Panel (A): Five-Year Rolling Beta

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Beta 5YR	10.87***	10.97***	12.90***	11.06***	13.09***	10.43***	11.72***
	(1.84)	(1.87)	(1.67)	(1.94)	(1.86)	(2.02)	(2.09)
Bond Var. (5Y)		1.01			0.73	0.39	
		(1.08)			(1.28)	(1.24)	
Market Vol. (5Y)			1.75***		1.60***	2.05***	
			(0.50)		(0.60)	(0.72)	
VIX				0.03	0.02	0.02	
				(0.02)	(0.02)	(0.02)	
Level						0.63***	
						(0.19)	
Slope						-0.04	
						(0.17)	
Curvature						0.29***	
						(0.11)	
FNS Excess Return							-0.01
							(0.04)
Constant	0.16	0.03	-1.72***	-0.48	-2.01***	-2.44***	0.07
	(0.20)	(0.11)	(0.54)	(0.45)	(0.50)	(0.52)	(0.18)
R-squared	0.43	0.43	0.47	0.44	0.47	0.53	0.47
Observations	1305	1305	1305	1233	1233	1233	435

Panel (B): One-Year Rolling Beta

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Beta 1YR	7.14***	7.27***	8.34***	7.87***	8.47***	4.81***	8.12***
	(1.55)	(1.60)	(1.46)	(1.57)	(1.66)	(1.41)	(1.61)
Bond Var. (1Y)		2.05			0.99	0.38	
		(1.31)			(1.26)	(1.43)	
Market Vol. (1Y)			1.16***		0.84**	1.33***	
			(0.35)		(0.39)	(0.34)	
VIX				0.06***	0.03	-0.00	
				(0.02)	(0.02)	(0.02)	
Level						1.07***	
						(0.15)	
Slope						-0.24	
						(0.15)	
Curvature						0.03	
						(0.14)	
FNS Excess Return							-0.00
							(0.04)
Constant	0.02	-0.25	-1.20**	-1.20**	-1.47***	-1.33***	-0.08
	(0.25)	(0.24)	(0.47)	(0.53)	(0.52)	(0.43)	(0.23)
R-squared	0.29	0.31	0.35	0.33	0.35	0.49	0.36
Observations	1305	1305	1305	1233	1233	1233	435

Note: The table reports the regression estimates for Equation 6, controlling for the realized volatility of bond and stock. Panel A presents the results for 5-year betas, while Panel B reports the results for 1-year betas. Driscoll–Kraay standard errors with 27 lags are reported in parentheses. The sample consists of monthly data from January 1988 to March 2024.

Table A3: Subjective Excess Returns and Stock Market Betas, Interacted with State Variables

Panel (A): Five-Year Rolling Beta

	(1)	(2)	(3)	(4)	(5)
Beta 5Y	12.03**	6.56	14.26***	12.49***	7.98***
	(5.23)	(9.56)	(4.50)	(2.15)	(1.74)
Beta × P/E	-0.05				
	(0.23)				
Beta × P/D		0.08			
		(0.19)			
Beta × VIX			-0.18		
			(0.27)		
Beta × Market Var. (1Y)				-1.33	
				(1.43)	
Beta × Market Var. (5Y)					6.05***
					(2.11)
P/E Ratio	0.01				
	(0.02)				
P/D Ratio		-0.01			
		(0.02)			
VIX			0.02		
			(0.02)		
Market Var. (1Y)				0.20*	
				(0.12)	
Market Var. (5Y)					1.12***
					(0.24)
Constant	-0.11	0.90	-0.32	-0.16	-0.97***
	(0.58)	(0.98)	(0.47)	(0.25)	(0.30)
R-squared	0.43	0.43	0.44	0.46	0.49
Observations	1305	1305	1233	1305	1305

Note: Panel (A) reports panel regressions of subjective excess bond returns on 5-year rolling bond–stock betas and their interactions with state variables capturing whether the price of risk varies with macro-financial conditions. Each column considers a different state variable, including valuation ratios (P/E and P/D), the VIX, and measures of market and bond return volatility at both the one-year and five-year horizons.
All specifications include maturity fixed effects. Standard errors are computed using Driscoll–Kraay corrections with 27 lags. The sample consists of monthly data from January 1988 to March 2024.

Table A4: Subjective Excess Returns and Stock Market Betas, Interacted with State Variables (continued)

Panel (B): One-Year Rolling Beta

	(1)	(2)	(3)	(4)	(5)
Beta 1Y	4.66	2.81	3.24	7.85***	6.78***
	(4.46)	(7.03)	(3.94)	(1.95)	(2.29)
Beta × P/E	0.14				
	(0.20)				
Beta × P/D		0.09			
		(0.14)			
Beta × VIX			0.28		
			(0.22)		
Beta × Market Var. (1Y)				0.15	
				(1.69)	
Beta × Market Var. (5Y)					1.29
					(1.49)
P/E Ratio	0.03				
	(0.03)				
P/D Ratio		-0.01			
		(0.02)			
VIX			0.08***		
			(0.03)		
Market Var. (1Y)				0.36**	
				(0.15)	
Market Var. (5Y)					0.50
					(0.32)
Constant	-0.67	0.56	-1.48***	-0.44	-0.55
	(0.69)	(1.23)	(0.56)	(0.31)	(0.51)
R-squared	0.30	0.30	0.34	0.34	0.31
Observations	1305	1305	1233	1305	1305

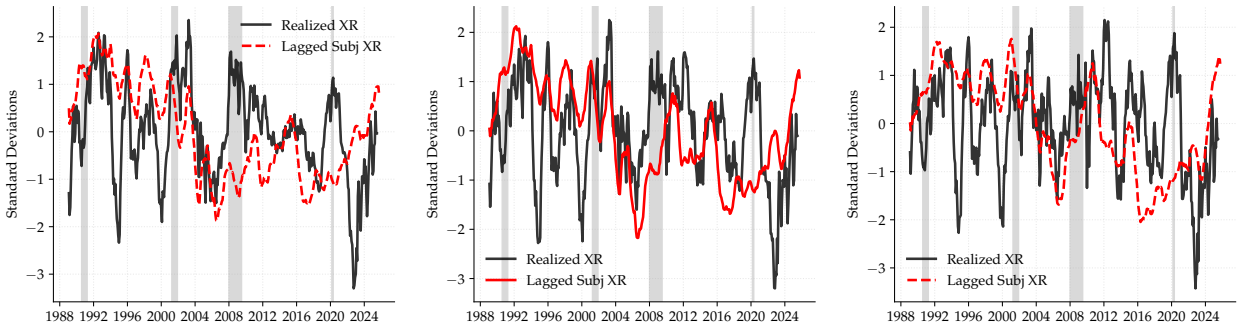
Note: Panel (B) reports panel regressions of subjective excess bond returns on 1-year rolling bond–stock betas and their interactions with state variables capturing whether the price of risk varies with macro-financial conditions. Each column considers a different state variable, including valuation ratios (P/E and P/D), the VIX, and measures of market and bond return volatility at both the one-year and five-year horizons.
All specifications include maturity fixed effects. Standard errors are computed using Driscoll–Kraay corrections with 27 lags. The sample consists of monthly data from January 1988 to March 2024.

Figure A.3: Realized Excess Returns and Lagged Subjective Excess Returns

(A) 3-Year Treasury Bond Maturity

(B) 6-Year Treasury Bond Maturity

(C) 11-Year Treasury Bond Maturity



Note: The figure plots realized excess returns together with the lagged, residualized subjective excess returns from Equation 6, where subjective excess returns are smoothed using a 12-month moving average. Both series are standardized to facilitate comparison within and across panels. Realized excess returns are shown in blue, while the residualized subjective excess returns are shown in teal. Shaded regions indicate NBER recession periods. The sample spans January 1988 to March 2024.

Table A5: Subjective Excess Returns and Bond-Stock (US) Market Betas: International Panel

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Beta 5YR (US)	7.94***	6.77***	7.28***	6.23***	7.35***	3.71**	5.98***	3.58**
	(1.58)	(1.63)	(1.45)	(1.37)	(1.46)	(1.53)	(2.26)	(1.53)
Bond Variance				0.17				
				(0.11)				
Term Spread					0.21			
					(0.16)			
10-Year Rate						0.71***		
						(0.21)		
1-Year Rate						-0.21		
						(0.17)		
Constant	-0.89***							
	(0.25)							
Time Trend		✓						
Country FE			✓	✓	✓	✓		✓
Time FE							✓	✓
R-squared	0.21	0.25	0.29	0.30	0.29	0.45	0.55	0.62
Observations	4822	4822	4822	4822	4822	4822	4822	4822

Note: This table is analogous to Table VII in the main paper, but it uses bond betas with respect to U.S. stock excess returns rather than domestic stock excess returns. The data are monthly and span September 1989 to December 2024.

Table A6: Subjective Excess Returns and Stock Market Betas: International Panel (Subset)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Beta 5YR	8.20***	6.12***	8.55***	8.55***	7.09***	2.43	2.63*
	(1.51)	(1.62)	(1.49)	(1.49)	(1.53)	(1.49)	(1.58)
Term Spread				-0.03			
				(0.09)			
10-Year Rate					0.02		
					(0.09)		
T-bill Rate					0.14		
					(0.09)		
Constant	-1.22***	-0.07					
	(0.19)	(0.45)					
Time Trend		✓					
Country FE			✓	✓	✓		✓
Time FE						✓	✓
R-squared	0.27	0.31	0.43	0.43	0.46	0.74	0.75
Observations	3065	3065	3065	3065	3065	3065	3065

Note: The table excludes France, Italy, Netherlands and Spain

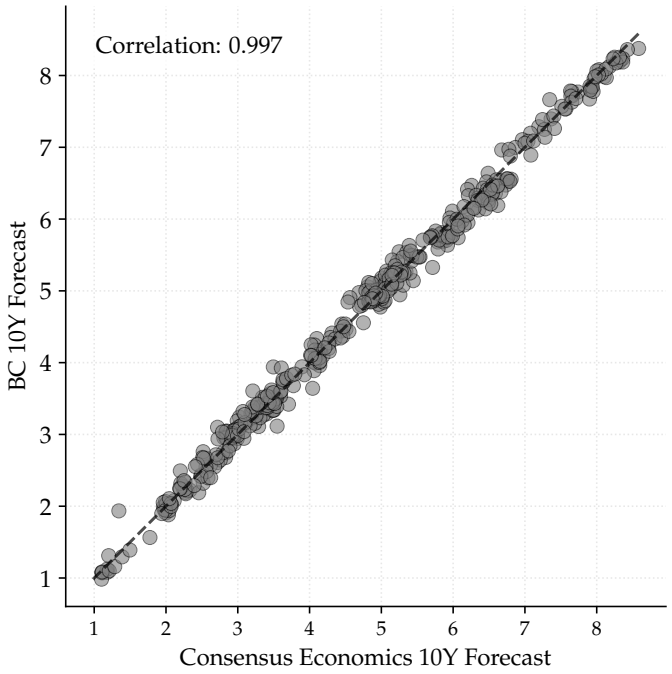
Table A7: Summary Statistics: International Stock-Bond β

Panel (B): β w.r.t. US Stock Returns

	Pooled	Canada	France	Germany	Italy	Japan	Netherlands	Norway	Spain	Sweden	Switzerland	UK	USA
Average	0.02	0.04	0.03	-0.01	0.10	0.02	-0.01	-0.03	0.10	0.00	-0.03	0.02	-0.02
Std	0.15	0.17	0.13	0.13	0.13	0.09	0.13	0.13	0.16	0.19	0.10	0.16	0.21
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	0.32	0.27	0.39	0.31	0.32	0.29	0.55	0.45	0.22	0.39	0.44
Average S.E.	0.06	0.07	0.07	0.06	0.08	0.04	0.06	0.06	0.07	0.07	0.05	0.08	0.08

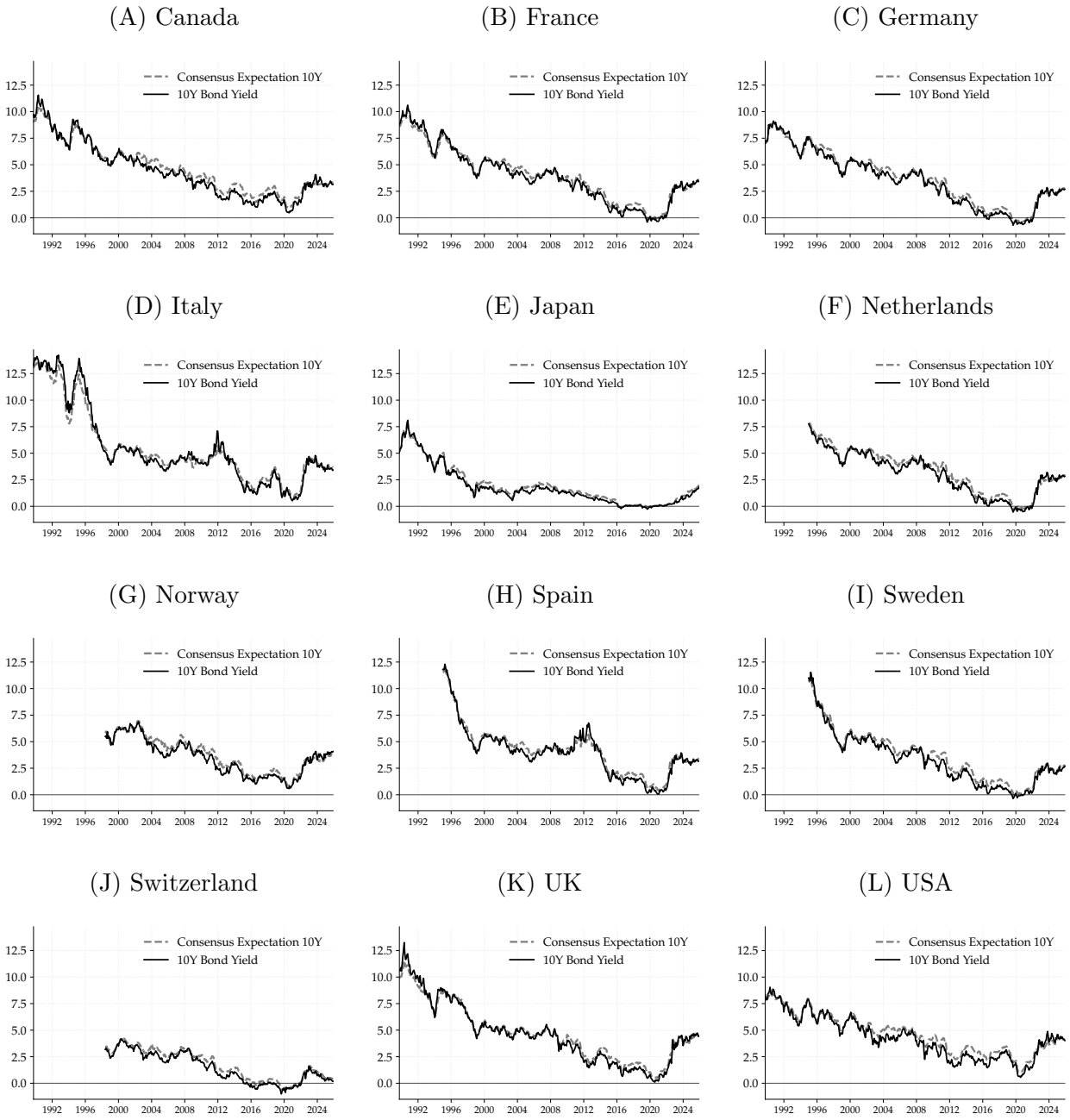
Note: This table reports the summary statistics for the rolling estimates of Government bond–stock betas for par bonds with ten-year maturities. Bond–stock betas are computed as the regression coefficient of monthly par bond excess returns on stock excess returns over a backward-looking rolling window, as defined in equation 5. We use 60-month rolling windows (5-year rolling β). Par bond excess returns are constructed using 10-year par yields from GFD database. Stock market excess returns are also computed using the GFD. The table uses the US stock returns. The table reports the average estimate, standard deviation, minimum and maximum values, the average HW-robust standard errors and the number of observations. The sample is monthly from September 1989 to December 2024.

Figure A.4: Consensus Economics vs Blue Chip for the US



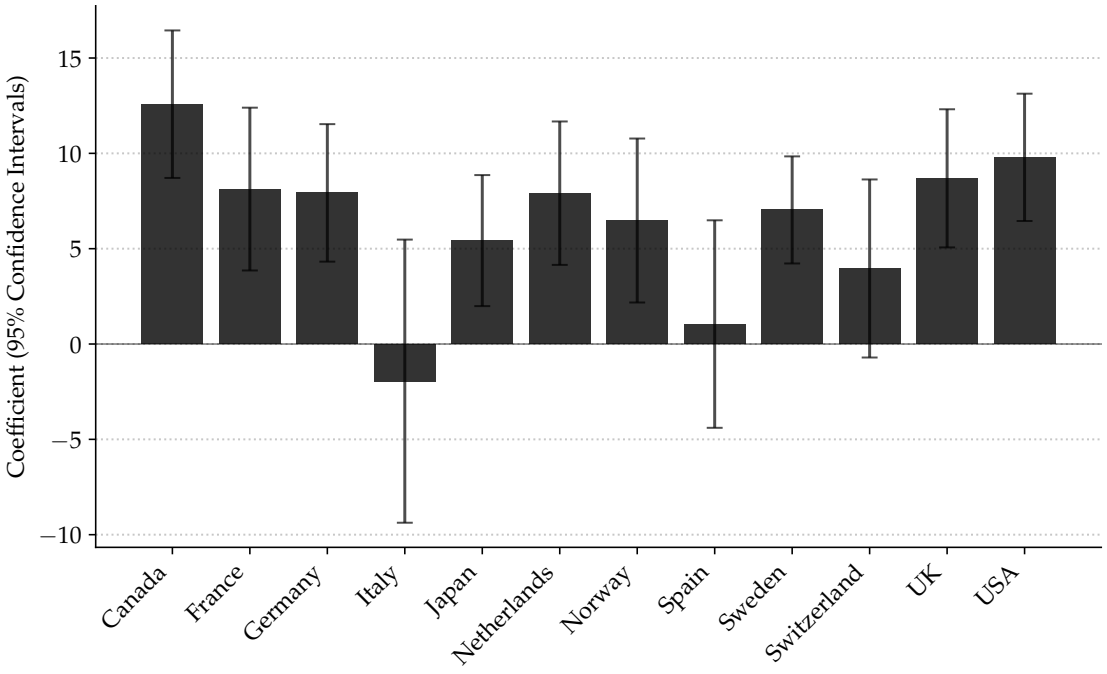
Note: The figure plots Consensus Economics forecasts of the 10-year government bond yield on the x-axis and the corresponding Blue Chip forecasts on the y-axis. Sample: September 1989 to June 2024.

Figure A.5: Consensus Economic Forecasts and Actual 10-year Yield



Note: The figure plots the realized 10-year government bond yield for each country together with the consensus economists' average forecast of the 10-year yield one year ahead. The realized yield is measured on the survey date, that is, at the time the forecasts are collected.

Figure A.6: Subjective Returns on Beta, using US stock returns



Note: The figure plots the estimated coefficient from regression (8), estimated separately for each country. This figure is analogous to Figure 6 in the main paper but uses bond betas with respect to U.S. stock excess returns instead of bond betas with respect to domestic market stock excess returns.

Table A8: Quantitative Easing (QE) Policy Periods Across Central Banks

Country	Central Bank	Regime	Start	End	Links
Italy	ECB	QE (APP)	2015–01–15	2018–12–19	Start – End
Italy	ECB	QE (Restart)	2019–11–01	2022–06–30	Start – End
Italy	ECB	QE (PEPP)	2020–03–26	2022–03–31	Start – End
France	ECB	QE (APP)	2015–01–15	2018–12–19	Start – End
France	ECB	QE (Restart)	2019–11–01	2022–06–30	Start – End
France	ECB	QE (PEPP)	2020–03–26	2022–03–31	Start – End
Germany	ECB	QE (APP)	2015–01–15	2018–12–19	Start – End
Germany	ECB	QE (Restart)	2019–11–01	2022–06–30	Start – End
Germany	ECB	QE (PEPP)	2020–03–26	2022–03–31	Start – End
Spain	ECB	QE (APP)	2015–01–15	2018–12–19	Start – End
Spain	ECB	QE (Restart)	2019–11–01	2022–06–30	Start – End
Spain	ECB	QE (PEPP)	2020–03–26	2022–03–31	Start – End
USA	Federal Reserve	QE1–QE3	2008–11–25	2014–10–29	Start – End
USA	Federal Reserve	QE (COVID)	2020–03–15	2022–03–15	Start – End
UK	Bank of England	QE	2009–03–05	2012–11–01	Start – End
UK	Bank of England	QE (Brexit)	2016–08–04	2018–12–01	Start – End
UK	Bank of England	QE (COVID)	2020–03–19	2021–12–01	Start – End
Japan	Bank of Japan	QE	2001–03–19	2006–03–09	Start – End
Japan	Bank of Japan	QQE	2013–04–04	2024–03–19	Start – End
Canada	Bank of Canada	QE	2020–04–01	2021–10–27	Start – End
Sweden	Riksbank	QE	2015–02–12	2021–11–01	Start – End

Notes: The table reports QE programme periods across major central banks. Hyperlinks point to official policy decision announcements. Norway is excluded as it did not implement QE.

Table A9: International and QE, Unsmoothed Subjective Excess Returns

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
QE	-1.17***	0.15	-0.92***	-0.24	0.99*	1.81***	0.15	1.43***
	(0.39)	(0.33)	(0.25)	(0.42)	(0.55)	(0.42)	(0.53)	(0.50)
Beta			8.78***	4.64***			6.62***	3.51**
			(1.88)	(1.70)			(1.85)	(1.50)
QE × Beta			-7.47***	-5.74**			-5.11**	-1.96
			(2.43)	(2.78)			(2.43)	(2.86)
Bond Variance					0.54***	0.24**	0.26***	0.18**
					(0.11)	(0.10)	(0.10)	(0.08)
QE × Bond Variance					-0.53***	-0.45***	-0.28**	-0.41***
					(0.14)	(0.13)	(0.13)	(0.10)
Constant	-0.53		-0.77***		-2.72***		-1.76***	
	(0.39)		(0.25)		(0.55)		(0.46)	
Country FE		✓		✓		✓		✓
Time FE		✓		✓		✓		✓
R-squared	0.03	0.49	0.24	0.50	0.18	0.51	0.26	0.52
Observations	4822	4822	4822	4822	4822	4822	4822	4822

Note: The table reports the results of equation (10) estimated using our international panel:

$$\tilde{E}_t xr_{c,t\to t+12} = \alpha + \alpha_{QE}\mathrm{QE}_{c,t} + \lambda\hat{\beta}_{c,t} + \lambda_{QE}\left(\mathrm{QE}_{c,t}\times\hat{\beta}_{c,t}\right) + \mu\hat{\sigma}^2_{c,t} + \mu_{QE}\left(\mathrm{QE}_{c,t}\times\hat{\sigma}^2_{c,t}\right) + \eta_{c,t}.$$

The bond–stock beta is estimated using monthly data over a five-year rolling window (60 months), based on excess returns on the domestic stock market. Similarly, bond variance is estimated as the rolling variance of monthly bond excess returns over the same five-year window. Expected yields are from Consensus Economics, while bond yields, T-bill rates, and stock returns are from the Global Financial Database (GFD). The term-spread is the difference between the 10-year government bond yield and the 1-year government bond yield. 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.

Table A10: QE and Bond-Stock Beta

	$\tilde{\beta}$	$\tilde{\beta}$	$\lvert\tilde{\beta}\rvert$	$\lvert\tilde{\rho}\rvert$	$\tilde{\sigma}_{bond}$	$\tilde{\sigma}_{mkt}$
	(1)	(2)	(3)	(4)	(5)	(6)
QE	0.03	-0.10	-0.14***	-0.10**	-0.15	0.08
	(0.08)	(0.14)	(0.05)	(0.05)	(0.10)	(0.13)
QE × ($\tilde{\beta} < 0$)		0.25**				
		(0.12)				
$\tilde{\beta} < 0$		-0.66***				
		(0.13)				
Constant	0.01	0.37***	0.23***	0.26***	0.18*	0.57***
	(0.07)	(0.11)	(0.06)	(0.07)	(0.09)	(0.21)
Lagged Value	✓	✓	✓	✓	✓	✓
R-squared	0.84	0.87	0.63	0.58	0.93	0.77
Observations	4677	4677	4677	4677	4677	4677

Note: Table A10 reports the results from the following regression:

$$Y_{c,t} = \alpha + b_1 \mathrm{QE}_{c,t} + b_2 Y_{c,t-12} + \eta_{c,t}, \tag{15}$$

where $Y_{c,t}$ denotes a set of dependent variables, and $Y_{c,t-12}$ is the corresponding value lagged by 12 months. Controlling for the lagged dependent variable is important, as it allows us to isolate the effect of QE from slow-moving dynamics in bond–stock betas. For instance, betas are on average negative in the post-2000 period but were positive on average prior to 2000; failing to control for persistence would confound these low-frequency trends with the effect of QE.

Column (1) regresses the level of $\tilde{\beta}$ on the QE indicator, where the tilde denotes that the variable is standardized (i.e., divided by its standard deviation). Column (2) allows the effect of QE to differ depending on whether the beta is positive or negative. Column (3) instead uses the absolute value of $\tilde{\beta}$. Finally, Columns (4) to (6) examine the effects of QE on the absolute value of the correlation, $\tilde{\rho}$, and on the standard deviations of bond and equity returns.