U.S. Risk and Treasury Convenience*

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Abstract

We document a rise in investors' assessment of U.S. risk relative to other G.7 economies since the late 1990s, driven by higher permanent risk but not reflected in currency returns. Using a two-country framework with trade in a rich maturity structure of bonds which earn convenience yields, alongside risky assets and currencies, we establish an equilibrium relationship between *cross-border* convenience yields, relative country risk and carry-trade returns. Empirically, we identify a cointegrating relationship between relative permanent risk and long-maturity convenience yields. Counterfactual experiments show rising relative permanent risk explains around one-third of declining long-maturity convenience yields over 2002-2006 and 2010-2014.

Key Words: Convenience Yields; Exchange Rates; Long-Run Risk; U.S. Safety, Equity Risk Premium.
JEL Codes: F30; F31; G12.

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1 Introduction

Is the U.S. still a safe haven, and what does this mean for U.S. debt and equity returns, and the U.S. dollar (USD) itself? Notwithstanding its depreciation and increased volatility over the first half of 2025, the USD has generally retained its strength over past decades, with no significant change in currency carry-trade returns on long- (or short-) maturity bonds. Yet, over the same period, U.S. equity returns have risen steadily and have outperformed the rest of the G.7 since the Global Financial Crisis (GFC). Moreover, a significant component of these equity returns were anticipated (as we show in Section 2, Figure 1)—consistent with classical explanations of risk compensation, as opposed to noise or 'good luck'.

This insulation of carry-trade returns from risk is inconsistent with the predictions of a large class of models which rely on complete and integrated financial markets (Backus, Foresi, and Telmer, 2001; Lustig, Stathopoulos, and Verdelhan, 2019). Such models imply the following no-arbitrage relationship across maturities:

$$\text{Carry-trade returns} = \text{Cross-country risk differential} + \text{Complete-markets deviation} \tag{1}$$

If markets are complete, movements in carry-trade returns should precisely mirror changes in cross-country risk differentials. However, over the 1997-2021 period, this is not true in the data and risk is not equalized (Section 2, Figure 2). So, how can theory be reconciled with data?

It is well understood that investors have long been willing to forgo other high-yield, risk-free government bonds in favor of 'convenient' U.S. Treasuries, as measured by deviations from covered interest parity (CIP) (see Du, Im, and Schreger, 2018a). While these 'convenience yields' have remained positive at short maturities (at least until the end of our sample), long-maturity U.S. bonds appear to have progressively 'lost their shine'. Since the early 2000s, CIP deviations on long-maturity U.S. bonds have been falling, and since 2010 U.S. bonds have frequently traded at a discount relative to some G.7 countries (Section 2, Figure 3), giving rise to an inverted term structure of CIP deviations.

In this paper, we build a model in which deviations from complete-market outcomes are attributed entirely to convenience yields. A U.S. investor long in Foreign bonds can be compensated for volatility in their stochastic discount factor (SDF) (i.e., their risk) either via *pecuniary* carry-trade returns or via relatively higher *non-pecuniary* convenience yields on Foreign bonds—reflected by a fall in CIP deviations. Through the lens of our model, the relative stability of carry-trade returns in the data implies that the rise in relative U.S. risk and the fall in convenience yields on long-maturity U.S. Treasuries are two sides of the same coin. Specifically, both reflect a deterioration of investors' perceptions of relative U.S. *permanent* risk, which is not reflected (at least so far) in the dynamics of the USD or of bond premia.[1]

We construct empirical measures of U.S. and G.7 risk to capture the SDF volatility underlying equation (1) using equity and bond returns, building on a tradition in finance (e.g., Hansen and Jagannathan, 1991; Bansal and Lehmann, 1997). However, rather than rely on realized equity returns, we construct valuation-based measures of *ex ante* equity risk premia to remove noise and better isolate the component of returns that compensates for risk (Campbell and Thompson, 2007; Campbell, 2008).[2] A key driver for our measure is expected dividend growth (see also De La'O and Myers, 2021; Atkeson, Heathcote, and Perri, 2024) which we construct using past dividends and survey data for output growth. We show that *ex ante* U.S. equity premia have risen both in absolute terms and *relative* to other G.7 countries, extending facts on macroeconomic trends documented in Farhi and Gourio (2018). Further, we estimate that 37% of the increase in relative equity returns over our sample is attributable to rising *expected* relative equity premia. Since the expected volatility of equity returns has not risen at the same pace, equity premia are the main driver of our risk measures (SDF volatility).

To decompose the increase in U.S. relative risk into transitory and permanent components, we expand the Alvarez and Jermann (2005) SDF decomposition to account for convenience yields. We find that the change in relative risk is almost entirely driven by the permanent component, since changes in relative bond premia (measured as in Adrian et al., 2013) over our sample are modest compared to the rise in U.S. relative equity premia. In particular, U.S. permanent risk has risen by about 5p.p. over the past two decades relative to other G.7 countries. Around this secular increase, our U.S. relative permanent risk measure displays two distinct troughs in which the U.S. was most safe—in 2002 and 2010, following the Dot-com crash and the GFC.[3] Notably, these troughs broadly coincide with the peaks in long-maturity Treasury convenience. Overall, the salience of permanent risk in our findings is supportive of explanations relying on a rise in (perceived) disaster probability (Farhi and Gourio, 2018) or to a switch to a permanent-innovation regime (Chernov, Lochstoer, and Song, 2021).

[1] Permanent risk refers to the volatility of innovations which affect investors' valuations of returns in the long run. In other words, higher permanent risk means a higher volatility in investors' revisions of distant future outcomes. We focus on permanent risk due to our interest in long-maturity CIP deviations and due to the importance of permanent risk in driving high yield asset returns (e.g., Alvarez and Jermann, 2005; Hansen and Scheinkman, 2009).

[2] The fact that 'good luck' cannot explain the high equity premium in the U.S. concurs with earlier work by Constantinides (2002) and Brandt, Cochrane, and Santa-Clara (2006) who argue that 'good luck' can only explain about 30% over a longer sample.

[3] That U.S. permanent risk fell sharply following financial crises is consistent with the realization of a disaster (see, e.g., Caramp and Silva, 2025)

We emphasize three main results. First, using our two-country, no-arbitrage framework, we establish an equilibrium relationship between carry-trade returns on long-maturity bonds, cross-country permanent risk differentials and long-maturity cross-country convenience yields (proxied with long-maturity CIP deviations). We estimate this relationship for 10-year maturity assets over 6-month holding periods. In doing so, we identify a single cointegrating relationship among these variables: specifically an inverse association between permanent-risk differentials and long-maturity CIP deviations, our proxy for convenience yields. In other words, while cross-country risk and long-maturity convenience are individually non-stationary, their linear combination is stationary, reconciling the empirical stability of carry-trade returns on long-maturity bonds with theory. We find that a 1p.p. increase in relative U.S. permanent risk is associated with approximately a 0.5b.p. decline in 10-year CIP deviations in the long run, over and above allowing for a deterministic trend to capture competing explanations for trend variation in long-maturity convenience yields (see, e.g., Jiang, Richmond, and Zhang, 2024, for an explanation based on the stock of U.S. debt). This result is further validated by a placebo test which replaces the measure of permanent risk with a proxy for transitory risk and identifies no such cointegration. We also show that the relationship between permanent risk and long maturity convenience yields is robust to including measures of dollar debt supply, the U.S. VIX and using survey expectations of currency returns. In short, market imperfections captured by convenience yields go a long way to capture the interplay of risk and returns underlying the no-arbitrage relation in equation (1) at long maturities.

Second, we characterize the fuller dynamics of long-maturity convenience yields by estimating an error-correction model, complementing the cointegrating relationship with relative permanent risk with short-run dynamics. All else equal, higher U.S. permanent risk implies lower long-maturity convenience yields both in both the short- and long-run. The model fits the actual data well and, in particular, captures both the cyclical dynamics and the broad decline in long-maturity convenience yields over the sample period. To illustrate the economic significance of these findings, we conduct counterfactual analyses. Using the coefficients estimated over the full sample, we construct counterfactual 10-year CIP deviations for scenarios in which relative permanent risk differentials remain constant. Fixing relative U.S. permanent risk at near its troughs *vis-à-vis* G.7 during the dot-com bubble and GFC, then, over 4-year periods beginning in 2002 and 2010, respectively, we find that increases in relative permanent risk explain around one-third of the decline in long-maturity convenience yields. Specifically, in the four years from 2002, long maturity CIP deviations fell by 35b.p., of which our model accounting for relative permanent risk captures around 89% of this decline. Likewise, in the four years following 2010, our full model captures almost the entirety of the 21b.p. decline in long-maturity CIP deviations. In both cases, using the counterfactual path for risk, we conclude that increases in relative permanent risk explain around one-third of the total decline captured by our model.

Third, we document that the long-maturity equilibrium relationship stands in stark contrast to the link between risk, returns and convenience at short maturities. Short-maturity carry-trade returns are volatile and are significantly associated with short-maturity convenience-yield differentials (Engel and Wu, 2023; Jiang, Krishnamurthy, and Lustig, 2021a). However, we find no significant role for risk differentials—in contrast to the strong empirical association between long-maturity convenience and permanent-risk differences that we uncover. Consistent with this, short-maturity convenience yield differentials are positive, on average, for the entirety of our sample suggesting that, despite uncertainty about relative U.S. permanent risk, the strength of the U.S. dollar and Treasuries remain insulated in the near term.[4]

Related Literature. Our paper contributes to literatures on the predictability of currency returns, the specialness of the dollar and the pricing of risky assets (particularly equity). Our starting point is the risk-based determination of exchange-rate returns (Backus et al., 2001; Verdelhan, 2010; Lustig et al., 2019), extended to account for convenience yields (Engel and Wu, 2023; Jiang et al., 2021a,b, 2024).[5] Our contribution to this literature is to directly measure all elements of the equilibrium relationships in the data.

Our paper investigates asymmetries in the international monetary system which have been widely studied in the literature from several angles. The U.S. has historically benefited from an excess return on its external assets (e.g., Gourinchas, Rey, and Govillot, 2010), seigniorage from abroad due convenience yields (e.g., Du, Im, and Schreger, 2018a; Jiang, Krishnamurthy, and Lustig, 2021a; Jiang, Richmond, and Zhang, 2024), but experiences a currency appreciation in global downturns as investors 'fly to safety' (e.g., Maggiori, 2017; Kekre and Lenel, 2024). Despite this, we emphasize that U.S. risk has been rising at least since the early 2000s (Farhi and Gourio, 2018), which has eroded the return on the net U.S. external portfolio (Atkeson, Heathcote, and Perri, 2025a). Our paper ties together several of these elements: higher risk has come at the cost of lower convenience yields on long-maturity U.S. Treasuries, although the dollar is insulated in the short run.

[4]In Appendix C.2, we present a parsimonious extension of our model which rationalizes why the short-maturity equilirbrium is different, by further allowing for segmentation between markets for short- and long-lived assets (see also Du, Hébert, and Li, 2023). In this framework, risk measured using equities is accurately reflected *only* in the long-maturity carry-trade relationship.

[5]Complementary explanations rely on limits to arbitrage (Jeanne and Rose, 2002; Gabaix and Maggiori, 2015; Vayanos and Vila, 2021; Gourinchas et al., forthcoming), long-run risk (Colacito and Croce, 2011), disaster risk (Farhi and Gabaix, 2016) and transitory risk (Lloyd and Marin, 2020) among others.

The literature on convenience yields focuses on their measurement and drivers, such as liquidity, safety, and limits to arbitrage (Krishnamurthy and Vissing-Jorgensen, 2012; Du, Im, and Schreger, 2018a; Du, Tepper, and Verdelhan, 2018b; Augustin, Chernov, Schmid, and Song, 2024; Liao and Zhang, 2025; Liu, Schmid, and Yaron, 2020; Diamond and Van Tassel, 2022). Other studies use convenience to explain exchange-rate dynamics, predominantly at short horizons (Engel and Wu, 2023; Krishnamurthy and Lustig, 2019; Valchev, 2020; Jiang, Krishnamurthy, and Lustig, 2021a,b). We place particular emphasis on the fall in *long-maturity* U.S. Treasury convenience. Taking the microstructure view, Jermann (2020), Du, Hébert, and Li (2023) and He, Nagel, and Song (2022) investigate the drivers of this fall *within* the U.S. In contrast, our paper draws a link between macroeconomic risk and cross-country convenience yields. In looking for a macro explanation for declining convenience yields, our work relates most closely to Jiang et al. (2024), who attribute the declines to increases in U.S. Treasury supply and show greater sensitivity of long-maturity convenience yields. In our empirical work, we show the role of rising U.S. risk is over-and-above the increase in debt-to-GDP.

Methodologically, we draw on a large literature using asset prices, specifically equity, to ascertain characteristics of SDFs (Hansen and Jagannathan, 1991; Alvarez and Jermann, 2005; Hansen and Scheinkman, 2009; Bakshi and Chabi-Yo, 2012; Chabi-Yo and Colacito, 2019) and macro risk (Farhi and Gourio, 2018; Greenwald et al., 2025). Our risk-based measure further adjusts for noise and 'good luck' using dividend-based pricing (Campbell and Thompson, 2007; Farhi and Gourio, 2018), corrects for expected volatility (Martin, 2017), and accounts for convenience yields. We document an increase in U.S. expected equity premia in line with Atkeson et al. (2025a) and Atkeson et al. (2025b), who argue this is in part explained by diverging paths in total payout ratios. Taking a different approach, Eeckhout (2025) decomposes firm value using profits and attributes 40% of the rise in valuation since 1980 to higher market power.

The remainder of this paper is structured as follows. Section 2 presents key stylized facts on equity, carry-trade returns, and the Treasury basis. Section 3 outlines the theoretical framework and derives the main equilibrium relationships. Section 4 discusses the mapping of theory to the data. Section 5 presents our empirical analysis. Section 6 concludes.

2 Stylized Facts on Equities, Currencies and the Treasury Basis

To motivate our analysis, we present three stylized facts on equity returns, currency returns and the Treasury basis. Later in the paper, we use the data shown here to derive theory-consistent risk measures and estimate no-arbitrage condition (1). Throughout, we focus on G.7 currencies from the late 1990s to 2021:03 (the last date for which the U.S. Treasury basis is available) and defer a fuller discussion of data construction and sources to Section 4.

#1. Rising U.S. Expected Equity Premia. U.S. equities have generally outperformed other G.7 markets over our sample. While part of this may have been unexpected (i.e., 'good luck'), we argue that a large component was anticipated and served as compensation for relatively higher U.S. risk, consistent with the evidence in Farhi and Gourio (2018). To capture this, we follow a large literature on valuation-based pricing (Campbell and Thompson, 2007; Campbell, 2008) and construct expected equity risk premia using restrictions from widely-used steady-state valuation models (e.g., Gordon, 1962).

Expected (log) equity risk premia $R^{eq}_{t,t+1}/R_t$ can be constructed as:[6]

$$\log \mathbb{E}_t \left[\frac{R^{eq}_{t,t+1}}{R_t} \right] = \log \left(1 + \frac{D_t}{P_t} + g^e_t \right) - \log \left(1 + r_t - \pi^e_t \right) \tag{2}$$

where $R^{eq}_{t,t+1}$ denotes the gross return to equities, R_t denotes the gross one-period risk-free rate, D_t/P_t denotes the dividend-price ratio, $r_t \equiv \log(R_t)$, π^e_t is a measure of expected inflation, and g^e_t is the expected net growth rate of dividends at time t.

Figure 1(a) plots the annualized expected (*ex ante*) equity premia for the U.S. *relative* to the rest of the G.7. This has generally increased since 1997 and displays two distinct peaks in U.S. 'risk', in 2004 and 2013, where the difference in premia reaches 10p.p. or higher in annualized terms. Conversely, the U.S. was relatively 'safe' following financial crises (in 2002 and 2009), consistent with a large literature characterizing the U.S. as an insurer during global crises (Gourinchas, Rey, and Govillot, 2010; Maggiori, 2017). Notably, between 2010 and 2013, in the aftermath of the GFC, there was an increase in relative U.S. premia of over 20p.p.

These cross-country expected returns are predominantly driven by variation in expected dividend growth (g^e_t) (see Appendices B.2.1 and B.2.2). We construct this expectation using lagged dividend growth and *Consensus Economics* surveys for next calendar-year GDP growth. A regression of realized dividends on these predictors yields an R^2 of 37% for the U.S., with both lagged dividend growth and GDP expectations being highly significant. Our g^e_t measure is

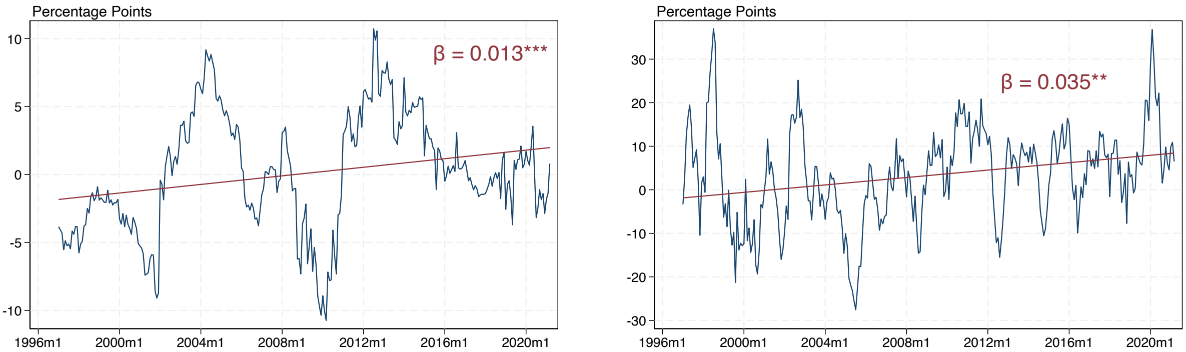
[6] Following Farhi and Gourio (2018), consider the simple static growth model in steady state:

$$\frac{P}{D} = \frac{1}{(1+r^f) + \mathbb{E}[ERP] - (1+g^e)}.$$

Define $\mathbb{E}[R^{eq}] = (1+r^f) + \mathbb{E}[ERP]$, substitute this into the previous expression, and rearrange to yield: $\mathbb{E}[R^{eq}] = 1 + \frac{D}{P} + g^e$. Using $\log \mathbb{E}_t \left[\frac{R^{eq}}{R_t} \right] = \log \mathbb{E}_t [R^{eq}] - \log(1+r^f)$ and $r^f = r - \pi^e$ delivers equation (2).

Figure 1: U.S. vs. G.7 Expected and Realized Equity Risk Premium

(a) U.S. net G.7 average *expected* equity premium (b) U.S. net G.7 average *realized* equity premium



Note. Panels 1a and 1b display time series of our baseline measures of U.S. net average G.7 expected and realized 6-month equity excess returns, respectively. The expected equity excess return in each market is estimated using equation (2), based on the dividend-price ratio, 6-month nominal government bond yields adjusted with *Consensus Economics* next-year inflation expectations, and a weighted average of lagged dividend growth and next-year GDP growth expectations from *Consensus Economics* to proxy for expected dividend growth in each market, as detailed in equation (B.1) in Appendix B.2. The sample runs from 1997:01 to 2021:03. The red trend lines are used for illustration only. *** (**) signifies that the slope (β) from a regression of the variable on a time trend is different from zero at the 1% (5%) significance level based on Newey and West (1987) standard errors with 4 lags.

also strongly correlated with CFO survey expectations for dividend growth (63%), presented in De La'O and Myers (2021) and widely used in the literature. These survey variables, however, are available only for the U.S. from 2003 onwards; as such we rely on our methodology to construct expected dividend growth for all G.7 countries in our sample.[7] Furthermore, in Appendix B.2.3, we show that expected U.S. equity premia have risen systematically on a country-by-country basis, comparing pre- and post-2010 averages.

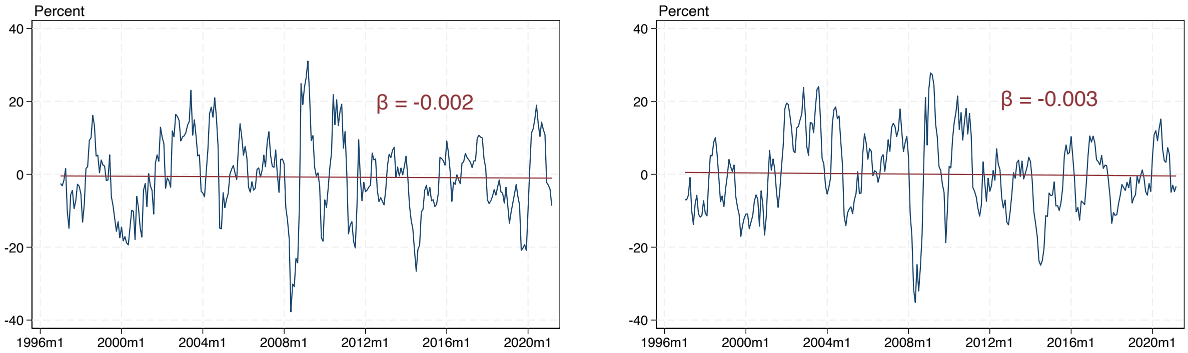
For comparison, Figure 1(b) plots realized (*ex post*) equity returns. Using linear deterministic trends for visualization suggests that about 37% of the observed increase in *realized* equity premia was anticipated over the full sample.

#2. No Significant Change in U.S. Dollar Returns. Next, we look at the pecuniary return investors earn on carry-trade portfolios funded in USD and long Foreign (currency) bonds of various maturities k. The expected one-period carry-trade return on a k-maturity bond, $\mathbb{E}_t[rx_{t+1}^{CT,(k)}]$, is comprised of a currency risk premium and a difference in domestic bond

[7]Within our sample, corporate action and specifically buy-backs are an increasingly important feature of the data and can generate differences in the dynamics of, e.g., aggregate dividends relative to total cash flows to equity holders (Atkeson et al., 2024). Due to lack of data for all G.7 countries, we restrict attention to dividends only, but allow for a deterministic trend to partly capture if total cash flow/price is rising in the U.S. *vis-à-vis* G.7. Additionally, if total payouts were higher, expected returns would be higher further strengthening are results.

Figure 2: U.S. Long- and Short-Maturity Carry-Trade Returns

(a) U.S. 10-Year Carry Trade Returns (b) U.S. 6-Month Carry Trade Returns



Note. Panels 2a and 2b display time series of long-maturity and short-maturity one-period U.S. dollar carry-trade returns, respectively (i.e., $rx_{t+1}^{CT,(\infty)}$ defined in (3) and $rx_{t+1}^{FX} = r_t^* - r_t + \Delta e_{t+1}$, respectively). In each case, the time series reflect the cross-sectional average U.S. dollar returns across the remaining G.7 currencies. The one-period holding period is 6 months and the bond maturities are 10 years (long-maturity) and 6 months (short-maturity), respectively. The sample runs from 1997:01 to 2021:03. The red trend lines are used for illustration only. *** (**) signifies that the slope (β) from a regression of the variable on a time trend is different from zero at the 1% (5%) significance level based on Newey and West (1987) standard errors with 4 lags.

premia:

$$\mathbb{E}_t[rx_{t+1}^{CT,(k)}] := \underbrace{\mathbb{E}_t[rx_{t+1}^{FX}]}_{\text{Currency Returns}} + \underbrace{\mathbb{E}_t[rx_{t+1}^{(k)*}] - \mathbb{E}_t[rx_{t+1}^{(k)}]}_{\text{Difference in Local Bond Returns}} \tag{3}$$

where $rx_{t+1}^{FX} = r_t^* - r_t + \Delta e_{t+1}$, with $e_t = \log(\mathcal{E}_t)$ representing the (log) exchange rate, defined as the number of USD per unit of Foreign currency, and r_t^* representing the Foreign analog of r_t. Similarly, $rx_{t+1}^{(k)(*)} = \log(R_{t,t+1}^{(k)(*)}/R_t^{(*)})$ denotes the one-period log excess holding return on a k-maturity domestic bond, in the U.S. or the Foreign country (denoted with $*$).[8]

Figure 2 plots annualized returns on long- (10-year, Panel (a)) and short-maturity (6-month, Panel (b)) carry-trade returns, over 6-month holding periods. Both are volatile: investors pursuing a carry trade with the USD as the funding currency suffered large losses during the GFC, whether they fashioned their positions using short- or long-maturity bonds. But more generally, both the mean and the variance of short- and long-maturity returns are generally constant over the sample. On a country by country basis, 10-year carry trade returns indeed have not experienced a significant change, pre- vs. post-2010, for all but CAD, but a small decrease is apparent at short maturities, see Appendix B.2.3 demonstrates.

So, if rising U.S. net G.7 expected equity premia over the sample are indicative of increasing relative U.S. risk, but carry-trade returns are flat (or not moving enough to compensate),

[8] $R_{t,t+1}^{(k)(*)}$ is the gross return from purchasing a k-period bond at time t and selling it at $t+1$ as a $k-1$-maturity bond. The exact definition is provided in Section 3.

condition (1) suggests that deviations from the complete and frictionless markets benchmark must be moving to restore no-arbitrage. For this, we turn to the Treasury basis constructed by Du et al. (2018a).

#3. Falling Treasury Basis on Long-Maturity Bonds. Define deviations from the k-maturity covered interest rate parity (CIP) condition as the (annualized) log return on a covered position (using an exchange-rate forward $f_t^{(k)}$), long on a k-period Foreign bond and short on a k-maturity U.S. Treasury (Home bond):

$$CIP_t^{(k)} := \underbrace{r_t^{(k)*} - r_t^{(k)}}_{\text{log interest rate differentials}} + \underbrace{f_t^{(k)} - e_t}_{\text{forward premium}} \tag{4}$$

When $CIP_{t,k} > 0$, the pecuniary return on a synthetic USD-denominated bond, $f_t^{(k)} - e_t + r_t^{(k)*}$, is greater than the return on a k-maturity U.S. Treasury, $r_t^{(k)}$ ($r_t^{(k)*}$ denoting the Foreign analog). Since arbitraging CIP deviations is riskless, this implies foreign investors intrinsically prefer U.S. bonds—i.e., these bonds offer a higher non-pecuniary return than Foreign bond. Note that, while CIP deviations for all maturities are small relative to other premia, they are much larger when mapped to convenience yields (in the utility space)—as Jiang et al. (2021a) show for short maturities—which is the relevant benchmark for no-arbitrage condition (1).

Figure 3(a) documents a persistent decline in 10-year CIP deviations over our sample. Having been as high as 70b.p. in the early 2000s, these long-maturity CIP deviations turned negative post-2016. Put differently, after 2016, foreign investors required excess portfolio returns to hold long-maturity Treasuries. This deterioration in Treasury convenience holds on a country-by-country basis as well, as we show in Appendix B.2.3. As well as trending in the opposite direction to expected equity premia (Figure 1) over our sample, the peaks and troughs of their cyclical fluctuations also broadly correspond.

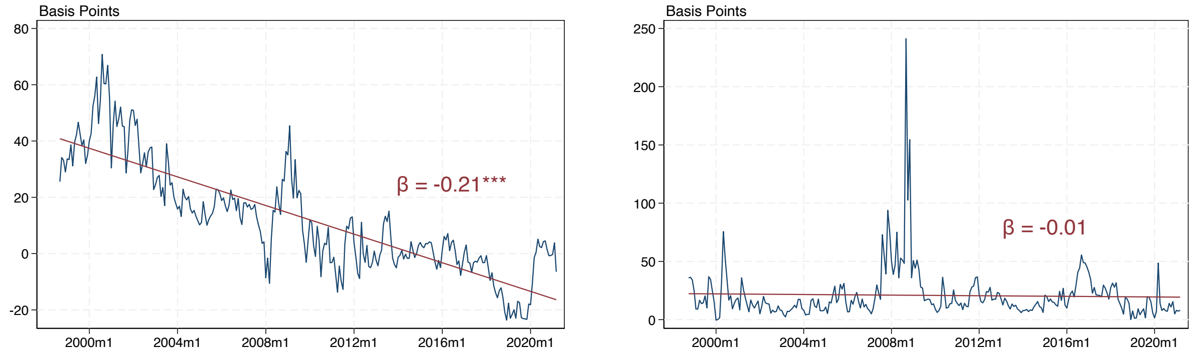
The profile of short-maturity convenience yields (Panel (b)) is sharply different. In contrast to 10-year CIP deviations, 6-month deviations remain positive throughout the sample. They also peaked much higher during the GFC, reaching almost 250b.p.

3 Model of Risk, Returns and Convenience Yields

We next lay out a theoretical framework to study these empirical facts. Consider an environment with two countries: Home (the U.S.) and Foreign (denoted with an asterisk $*$). Representative

Figure 3: Long- and Short-Maturity U.S. Treasury Basis

(a) U.S. 10Y Treasury Basis (b) U.S. 6M Treasury Basis



Note. Panels 3a and 3b display time series of the long-maturity and short-maturity U.S. Treasury basis, respectively. These reflect CIP deviations (defined in (4)), where the government bond maturities are 10 years (long-maturity) and 6 months (short-maturity), respectively. In each case, the time series are a cross-sectional average across G.7 currencies *vis-à-vis* the U.S. The sample runs from 1997:01 to 2021:03. The red trend lines are used for illustration only. *** (**) signifies that the slope (β) from a regression of the variable on a time trend is different from zero at the 1% (5%) significance level based on Newey and West (1987) standard errors with 4 lags.

investors in each country trade in zero-coupon bonds of varying maturities $k = 1, 2, ..., \infty$ issued in both economies. The bonds pay a known return in local currency at maturity and are free from default risk. Investors also earn a non-pecuniary convenience yield from bonds which is specific to Home and Foreign investors $i = H, F$, Home and Foreign assets $j = H, F$, their corresponding maturity k, and varies over time. Investors also trade in risky assets which carry lower convenience, normalized to zero.

Throughout, we do not rely on a log-normal approximation and use entropy operators instead. This is because heteroskedastic models generate non-log-normal returns at longer maturities, even when short-maturity returns are log-normal (see Campbell et al., 1997). For a generic variable X, conditional entropy is $\mathcal{L}_t(X_t) \equiv \mathbb{E}_t[\log X_t] - \log \mathbb{E}_t[X_t]$, where $\mathcal{L}_t(X_t) = \frac{1}{2}\text{var}_t(X_t)$ if X_t is log-normally distributed.[9] Unconditional entropy, denoted by $\mathcal{L}(X_t)$, is similarly defined. Conditional co-entropy, $\mathcal{C}_t(X_t)$, a generalized notion of covariance discussed further in Backus, Boyarchenko, and Chernov (2018), is analogously defined as $\mathcal{C}_t(X_t) = \mathcal{L}_t(X_tY_t) - \mathcal{L}_t(X_t) - \mathcal{L}_t(Y_t)$.

[9]This measure of conditional volatility $\mathcal{L}_t(X)$, often referred to as Theil (1967)'s second conditional entropy measure, is in general equal to one half the second-order cumulant ($\frac{\kappa_{2,t}}{2!} = \frac{\text{var}_t(X_{t+1})}{2}$) plus all higher-order cumulants ($\frac{\kappa_{3,t}}{3!} + \frac{\kappa_{4,t}}{4!} + ...$). If $\text{var}_t(X_t) = 0$, then $\mathcal{L}_t(X_t) = 0$.

3.1 Equilibrium Asset Pricing

3.1.1 Bond Markets

Define $P_t^{(k)}$ as the date-t price of a Home zero-coupon bond of maturity k. The gross pecuniary return on this bond, earned at maturity but known at time t, is: $R_t^{(k)} = 1/P_t^{(k)}$. The 'risk-free rate' R_t is defined for $k = 1$, such that: $R_t \equiv R_t^{(1)} = 1/P_t^{(1)}$. In addition to earning a pecuniary return, investors also earn a non-pecuniary *convenience yield* from holding assets. An investor i purchasing a country-j bond at time t that is held to maturity k earns a convenience yield $\theta_t^{i,j(k)}$, as formalized below.[10]

Assumption 1 (Convenience-Yield Term Structure) *Home and Foreign investors trade in Home and Foreign risk-free bonds of maturity $k = 1, 2, ..., \infty$. The term structure of convenience yields held to maturity (i.e., $\theta_t^{i,j(k)}$ for an investor i purchasing a k-period country-j bond at time t) is observable at time t.*

Let the Home (Foreign) nominal pricing kernel in period t be denoted by Λ_t (Λ_t^*). In turn, define the Home (Foreign) k-period SDF between periods t and $t+k$ as: $M_{t,t+k} \equiv \Lambda_{t+k}/\Lambda_t$ ($M_{t,t+k}^* \equiv \Lambda_{t+k}^*/\Lambda_t^*$). With these SDFs, Euler equations for Home and Foreign agents investing in k-period Home and Foreign risk-free bonds are, respectively, given by:

$$\mathbb{E}_t \left[M_{t,t+k}\right] R_t^{(k)} = e^{-\theta_t^{H,H(k)}} \tag{5}$$

$$\mathbb{E}_t \left[M_{t,t+k}^*\right] R_t^{(k)*} = e^{-\theta_t^{F,F(k)}} \tag{6}$$

$$\mathbb{E}_t \left[M_{t,t+k} \frac{\mathcal{E}_{t+k}}{\mathcal{E}_t}\right] R_t^{(k)*} = e^{-\theta_t^{H,F(k)}} \tag{7}$$

$$\mathbb{E}_t \left[M_{t,t+k}^* \frac{\mathcal{E}_t}{\mathcal{E}_{t+k}}\right] R_t^{(k)} = e^{-\theta_t^{F,H(k)}} \tag{8}$$

for all maturities k. The nominal exchange rate $\mathcal{E}_t$ represents the number of USD per unit of Foreign currency, so an increase in $\mathcal{E}_t$ represents a U.S. depreciation/Foreign appreciation.

[10] To the extent that convenience yields arise from a bond's value as collateral, this timing assumption reflects that collateral value is accounted for in contracts written at time t. Convenience yields are often micro-founded using bond-in-utility formulations as in Valchev (2020), Jiang and Richmond (2023), and Jiang et al. (2021b).

3.1.2 Foreign Exchange

If (5)-(8) hold, any exchange-rate process which admits no arbitrage must satisfy:

$$\mathcal{L}_t(\mathcal{E}_{t+1}/\mathcal{E}_t) = \mathcal{C}_t(M^*_{t,t+1}/M_{t,t+1}, \mathcal{E}_{t+1}/\mathcal{E}_t) + (\theta_t^{H,H(1)} - \theta_t^{H,F(1)}) - (\theta_t^{F,H(1)} - \theta_t^{F,F(1)}) \tag{9}$$

We consider the class of exchange-rate processes given by:

$$\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t} = \frac{M^*_{t,t+1}}{M_{t,t+1}} e^{\eta_{t+1}} \tag{10}$$

where η_{t+1} is the incomplete-markets wedge (see, e.g., Backus et al., 2001; Lustig and Verdelhan, 2019). In the complete-markets benchmark, absent convenience yields, the condition is satisfied by setting $\eta_{t+1} = 0$. With incomplete markets, the wedge η_{t+1} depends on convenience-yield differentials. Combining (5) and (8), and (6) and (7) with (10), respectively, yields:

$$\mathbb{E}_t[\eta_{t+1}] = \mathcal{L}_t(\eta_{t+1}) - \mathcal{C}_t(M_{t,t+1}, e^{\eta_{t+1}}) + \theta_t^{F,H(1)} - \theta_t^{H,H(1)} \tag{11}$$

$$\mathbb{E}_t[\eta_{t+1}] = -\mathcal{L}_t(\eta_{t+1}) - \mathcal{C}_t(M^*_{t,t+1}, e^{\eta_{t+1}}) + \theta_t^{F,F(1)} - \theta_t^{H,F(1)} \tag{12}$$

Ceteris paribus, an increase in the convenience yield enjoyed by Foreign investors $\theta_t^{F,H(1)}$, either leads to a Foreign-currency depreciation (through $\mathbb{E}_t[\eta_{t+1}]$), or must be be compensated by making exchange-rate movements 'safer' for Home investors (see Jiang et al., 2021b). In addition to (11) and (12), there is an analogous restriction on η_{t+1} for each long-maturity bond, but also for each traded risky asset. The following assumption is made for clarity and tractability.

Assumption 2 (Complete Spanning) *We consider the limit where* $\mathcal{L}_t(e^{\eta_{t+1}}) \to 0$ *such that* $\mathbb{E}_t[\eta_{t+1}] = \theta_t^{F,H(1)} - \theta_t^{H,H(1)} = \theta_t^{F,F(1)} - \theta_t^{H,F(1)}$.

Appendix A details an example of a market structure (i.e., a set of SDFs and traded assets) where Assumption 2 arises as a result. Intuitively, this requires trade in additional risky assets which can span both convenience yields and risk, although they themselves carry a lower convenience yield than risk-free bonds. The implied exchange-rate process is unique and can be verified by seeing that the Euler equations for any internationally-traded asset must be equalized up to the convenience yields.[11] While stark, this limiting process delivers clear theoretical predictions which we can take to the data. Exchange rates are given as follows:

[11] The resulting exchange-rate process implies that convenience yields are investor specific—i.e., foreigners assign a symmetrically high (in relative terms) convenience yield to both bonds. A similar assumption is made in Jiang et al. (2021a).

Lemma 1 (FX Process: Complete Spanning and Convenience Yields) *Under Assumptions 1 and 2:* $\Delta e_{t+1} = m^*_{t,t+1} - m_{t,t+1} + \theta_t^{F,H(1)} - \theta_t^{H,H(1)}$.

3.1.3 Equity Markets

In addition to trading in domestic and foreign bonds, investors in each country also trade in a domestic risky asset which we label (leveraged) equity.

Assumption 3 (Equities and Convenience) *Home and Foreign investors also trade in a respective domestic risky asset, with one-period return* $R^g_{t,t+1}$ $(R^{g*}_{t,t+1})$*, on which they derive a baseline level of convenience yield, normalized to zero.*

The returns on risky assets must therefore satisfy the following Euler equations for Home and Foreign investors, respectively, for all t:

$$\mathbb{E}_t \left[M_{t,t+1} R^g_{t,t+1} \right] = 1 \tag{13}$$

$$\mathbb{E}_t \left[M^*_{t,t+1} R^{g*}_{t,t+1} \right] = 1 \tag{14}$$

3.2 Equilibrium Relationships

We next discuss the model's predictions for the links between convenience, risk and returns at short and long maturities.

3.2.1 Short-Maturity Equilibrium Relationships

Define rx^{FX}_{t+1} as *ex post* (log) return from a one-period carry-trade strategy that goes long the Foreign risk-free bond and short the US risk-free bond:

$$rx^{FX}_{t+1} = \log \left(\frac{R^*_t}{R_t} \frac{\mathcal{E}_{t+1}}{\mathcal{E}_t} \right) = r^*_t - r_t + \Delta e_{t+1}. \tag{15}$$

Absent frictions, if investors are risk neutral, then UIP should hold: $\mathbb{E}_t[rx^{FX}_{t+1}] = 0$. Deviations from UIP (i.e., $\mathbb{E}_t[rx^{FX}_{t+1}] \neq 0$) thus reflect an exchange-rate risk premium for which investors earn cross-border carry-trade returns. Within a classic open-economy setup without convenience yields (i.e., $\theta_t^{i,j(k)} = 0$ for all i, j, k), 'risk-based' explanations of UIP failures draw a link between the covariance of investors' SDFs and returns on Foreign-currency portfolios

(Backus et al., 2001). If the Foreign investor bears greater risk—i.e., experiences greater SDF volatility ($\mathcal{L}_t(M_{t,t+1}) < \mathcal{L}_t(M^*_{t,t+1})$)—they earn expected excess returns on a cross-border carry-trade portfolio that is long the one-period Home bond and short the analogous Foreign bond ($\mathbb{E}_t[rx^{FX}_{t+1}] < 0$). The covariance of this return with yields constitutes the forward-premium puzzle (Fama, 1984), for which Verdelhan (2010) provides a risk-based explanation.

When investors attach convenience yields to bonds, the link between the exchange-rate risk premium and relative SDF volatility (i.e., condition (1) for short-maturities) is detailed below:

Proposition 1 (SDF Volatility, FX Risk and Convenience at Short Maturities) *Given $M^*_{t,t+1}$, $M_{t,t+1}$ and the exchange-rate process (10), the expected excess return on a one-period carry-trade strategy, relative SDF volatility, and one-period convenience yields satisfy:*

$$\mathbb{E}_t[rx^{FX}_{t+1}] = \underbrace{\mathcal{L}_t(M_{t,t+1}) - \mathcal{L}_t(M^*_{t,t+1})}_{\text{Risk Differential}} + \underbrace{\theta^{F,H(1)}_t - \theta^{F,F(1)}_t}_{\text{Short-Maturity Convenience}}$$

Proof: Combine expectations of the exchange-rate process (10) and the log-entropy expansions of the $k = 1$-period domestic Euler equations (5) and (6). See Appendix A.2 for full derivation. □

Suppose the Foreign investor faces greater SDF volatility (i.e., $\mathcal{L}_t(M_{t,t+1}) < \mathcal{L}_t(M^*_{t,t+1})$). Then, by no arbitrage, they can be compensated either by greater expected pecuniary returns on a long position in the Home bond ($\mathbb{E}_t[rx^{FX}_{t+1}] < 0$) or greater non-pecuniary convenience yields on the Home bond relative to the Foreign bond ($\theta^{F,H(1)}_t - \theta^{F,F(1)}_t > 0$), or a combination of both.[12] That is, in addition to the traditional open-economy logic that pecuniary currency returns compensate investors for bearing relative risk, Proposition 1 demonstrates that non-pecuniary convenience yields provide a second channel through which asymmetries in SDF volatility across countries can equilibrate.

3.2.2 Transitory and Permanent Risk Decomposition with Convenience

To tie asset returns with risk and convenience at difference horizons, we use the Alvarez and Jermann (2005) decomposition of pricing kernels. We decompose the Home pricing kernel Λ_t

[12] If we relax complete spanning (Assumption 2), (11)-(12) illustrate that the investor could also be compensated by changes in exchange rate cyclicality with respect to SDFs. Note, that under complete spanning, the entropy of SDFs hides a co-entropy operator because $\mathcal{L}_t(M^*_{t+1}) = \mathcal{L}_t(M_{t+1}\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}) = \mathcal{L}_t(M_{t+1}) + \mathcal{L}_t(\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}) + \mathcal{C}_t(M_{t+1}, \frac{\mathcal{E}_{t+1}}{\mathcal{E}_t})$.

into two components:

$$\Lambda_t = \Lambda_t^{\mathbb{P}} \Lambda_t^{\mathbb{T}} \tag{16}$$

where $\Lambda_t^{\mathbb{P}}$ is a martingale that captures the 'permanent' component of Λ_t, while $\Lambda_t^{\mathbb{T}}$ reflects the 'transitory' component. We decompose the Foreign pricing kernel analogously. Under regularity conditions, the permanent component is defined as $\mathbb{E}_t \Lambda_{t+1}^{\mathbb{P}} = \Lambda_t^{\mathbb{P}}$, where $\Lambda_t^{\mathbb{P}} = \lim_{k\to\infty} \frac{\mathbb{E}_t \Lambda_{t+k}}{\beta^{t+k}}$.[13] The permanent measure is unaffected by information at time t that does not lead to revisions of the expected value of Λ in the long run. The transitory component is defined by the residual $\Lambda_t^{\mathbb{T}} = \lim_{k\to\infty} \frac{\beta^{t+k}}{\mathbb{E}_t[\Lambda_{t+k}]/\Lambda_t}$ and is equivalent to a scaled long-term interest rate.

First, we derive the link between the transitory component of representative investors' SDFs, long-maturity bond returns and long-maturity convenience yields:

Lemma 2 (Transitory SDF, Asset Prices and Convenience) *The transitory component of the Home representative investor's SDF is inversely related to the one-period holding return on an infinite-maturity bond (i.e., bond term premium), adjusted for the holding-period convenience yield:*

$$\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}} \equiv M_{t,t+1}^{\mathbb{T}} = \frac{1}{R_{t,t+1}^{(\infty)}} e^{-\theta_{t,t+1}^{H,H(\infty)}} \tag{17}$$

where $\theta_{t,t+1}^{H,H(\infty)} = \lim_{k\to\infty} \frac{e^{\theta_t^{H,H(k)}}}{e^{\theta_{t+1}^{H,H(k-1)}}}$. *An analogous expression exists for the Foreign investor.*

Proof: See Appendix A.3. □

Absent convenience yields, transitory innovations to pricing kernels reflect the bond term premium only. When there are convenience yields on domestic bonds, an investor holding an infinite-maturity bond for a single period earns the convenience yield today on that bond, but forgoes the residual convenience yield tomorrow. Holding-period convenience yields $\theta_{t,t+1}^{i,j(\infty)}$ are defined analogously for any i, j and will play a vital role in our analysis. The transitory SDF (17) therefore reflects both the one-period pecuniary return on the infinite-maturity bond as well as the change in the convenience yield on that infinite-maturity bond.

Second, we derive bounds on the volatility of the overall and permanent components of risk, in relation to equity premia. This is a key step in our analysis since it defines our preferred measure of overall and permanent risk for our empirical analysis. As the following lemma clarifies, the presence of convenience yields influences these bounds:

[13] Specifically, the decomposition assumes: (i) that there is a number β such that $0 < \lim_{k\to\infty} \frac{\mathbb{E}[\Lambda_{t+k}]/\Lambda_t}{\beta^k} < \infty$ for all t; and (ii) for each $t+1$ there is a random variable X_{t+1} such that $\frac{\Lambda_{t+1}}{\beta^{t+1}} \frac{\mathbb{E}_{t+1}[\Lambda_{t+1+k}]/\Lambda_{t+1}}{\beta^k} \leq X_{t+1}$ almost surely, with $\mathbb{E}_t X_{t+1}$ finite for all k.

Lemma 3 (Permanent and Total Risk, Asset Prices and Convenience) *The lower bound for conditional volatility of the Home investor's SDF in the presence of convenience yields is given by:*

$$\mathcal{L}_t(M_{t,t+1}) \geq \mathbb{E}_t \log \left[\frac{R^g_{t,t+1}}{R_t} \right] - \theta_t^{H,H(1)} \tag{18}$$

The lower bound on the conditional volatility of the permanent component of the Home representative investor's SDF is:

$$\mathcal{L}_t \left(M^{\mathbb{P}}_{t,t+1} \right) \geq \mathbb{E}_t \log \left[\frac{R^g_{t,t+1}}{R_t} \right] - \mathbb{E}_t \left[rx^{(\infty)}_{t+1} \right] - \mathbb{E}_t[\theta^{H,H(\infty)}_{t,t+1}] \tag{19}$$

where $rx^{(\infty)}_{t+1} = \log R^{(\infty)}_{t,t+1}/R_t$*. Analogous expressions exists for the Foreign investor.*

Proof: See Appendix A.4. □

Inequality (18) bounds the conditional volatility of the Home investor's overall SDF—the measure of the country's overall riskiness—such that it is at least as large the expected (log) excess return on risky equities, net of the convenience yield the Home investor earns on the Home bond. While inequality (18) holds for any risky asset, the right-hand side is maximized by using the highest risk premium in the economy. Intuitively, this bound on overall risk reflects that the return on the riskiest asset in the economy can be expected to capture all types of risk, both permanent and transitory. Relative to a model without convenience (as in Alvarez and Jermann, 2005; Lustig et al., 2019) the additional convenience-yield term indicates that, by taking a long position in equities, the Home investor not only forgoes the pecuniary return on the safe one-period bond R_t but also the convenience yield on this bond $\theta_t^{H,H(1)}$. However, quantitatively, we find the effects of convenience yields on measures of risk are small relative to the equity premium.

Inequality (19) bounds the permanent component of the Home investor's SDF—the measure of the country's permanent risk—such that it is at least as large as the difference between the expected (log) excess return on risky equities (the first term on the right-hand side of expression (19)) and the expected one-period return on the asymptotic discount bond, net of changes in the convenience yield on this bond. The expression isolates permanent risk because the bond premium and convenience yield on a long-maturity bond encodes the transitory (insurable) component of risk only. Holding-period convenience reflects that, by forgoing the one-period pecuniary return on an infinite maturity bond $\mathbb{E}_t \left[rx^{(\infty)}_{t+1} \right]$, the Home investor also forgoes today's convenience net of what they expect to recover tomorrow, captured by $\mathbb{E}_t \left[\theta^{H,H(\infty)}_{t,t+1} \right]$.

3.2.3 Long-Maturity Equilibrium Relationships

The failure to reject UIP empirically over long horizons (e.g., Chinn and Meredith, 2005; Lustig et al., 2019), along with our first stylized fact on cross-country risk differentials, imply that such explanations impose counterfactually strict restrictions at long maturities. This, alongside our interest in permanent risk specifically, motivates our focus on long maturities. To derive an analog to Proposition 1 for long-maturity terms, we define two new objects. The first is the one-period carry-trade return from a long position in the Foreign ∞-maturity bond funded by a short position in the Home ∞-maturity bond for one period—i.e., the infinite-maturity limit of equation (3), which combines currency and local-currency bond returns. The second pertains to the term structure of convenience yields:

Lemma 4 (Term Structure of Convenience Yields) *Given $M_{t,t+1}$, $M^*_{t,t+1}$, the term structure of convenience yields satisfies:*

$$\theta_t^{H,H(k)} = \theta_t^{F,H(k)} - \sum_{\tau=0}^{\kappa-1} \left\{ \mathbb{E}_t \theta_{t+\tau}^{F,H(1)} - \theta_{t+\tau}^{H,H(1)}] \right\} \tag{20}$$

$$\theta_t^{F,F(k)} = \theta_t^{H,F(k)} + \sum_{\tau=0}^{\kappa-1} \left\{ \mathbb{E}_t \theta_{t+\tau}^{F,F(1)} - \theta_{t+\tau}^{H,F(1)}] \right\} \tag{21}$$

for all k and all t.

Proof: See Appendix A.5. □

Lemma 4 is the limiting case resulting from Assumption 2. In equation (20) and (21), the terms $\theta_t^{i,j(k)} - \sum_{\tau=0}^{k-1} \theta_{t+\tau}^{i,j(1)}$, for investor i and bond j, are analogous to 'convenience term premia'. For a given bond j, these deviations from the convenience-yield expectations hypothesis on a specific bond would be proportional for Home and Foreign investors, extending the symmetry in Assumption 2.[14] Moving beyond the limit of complete spanning, deviations would arise to compensate for variation in the riskiness of convenience yields, e.g., $\text{cov}_t(m_{t+1}^{(*)}, \eta_{t+1})$.

The next proposition details our main theoretical result:

[14]These strict conditions are imposed only because we allow for frictionless trading in every period. If instead we allowed for the potential of only occasionally accessing the market, a richer set of term structures would be accommodated even with complete spanning.

Proposition 2 (SDF Volatility, FX Risk and Convenience Yields at Long Maturities) *Given $M_{t,t+1}$, $M^*_{t,t+1}$, one-period carry-trade returns from long-term bonds are given by:*

$$\mathbb{E}_t[rx^{CT,(\infty)}_{t+1}] = \underbrace{\mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) - \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1})}_{\text{Permanent Risk Differential}} + \underbrace{\mathbb{E}_t[\theta^{F,H(\infty)}_{t,t+1} - \theta^{F,F(\infty)}_{t,t+1}]}_{\text{Long-Maturity Holding-Period Convenience}} \tag{22}$$

Proof: See Appendix A.6. □

Extending equation (1) to long maturities, the proposition above implies that if pecuniary carry-trade returns on long-term bonds are, on average, zero (i.e., $\mathbb{E}_t[rx^{CT,(\infty)}_{t+1}] = 0$), or do not adjust enough, asymmetries in permanent risk across countries must be matched by movements in long-maturity holding-period convenience—i.e., the convenience that a Foreign investor expects to earn from buying a U.S. Treasury, instead of their own domestic bond, today and selling it back tomorrow. Specifically, if Foreign investors experience greater permanent risk, such that $\mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) < \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1})$, equation (22) demonstrates that Foreign investors must receive a greater non-pecuniary convenience yield from their net-long position in the U.S. bond at time t. That is, $\mathbb{E}_t[\theta^{F,H(\infty)}_{t,t+1} - \theta^{F,F(\infty)}_{t,t+1}]$ must increase such that investors earn more convenience on the Home bond at t relative to the convenience they expect to forgo when they unwind their carry-trade position at time $t+1$.

4 Data and Measurement

We now describe the data we use to test our propositions empirically. This requires measures of currency returns, risk and convenience (both within- and across-countries). We focus our analysis on G.7 currencies where financial-market data is sufficiently rich to capture these quantities: Australia (AUD), Canada (CAD), Switzerland (CHF), euro area/Germany (EUR), Japan (JPY), UK (GBP) and US (USD). For our regression-based analysis, we henceforth use a balanced panel at a monthly frequency, over the period 2000:01-2021:03.

We use equation (3) and (15) to construct measures of currency returns, $rx^{CT,(k)}_{t+1}$ and rx^{FX}_{t+1}, respectively, following a common approach in the literature by using *ex post* exchange rates and yields to proxy for *ex ante* expectations for our benchmark measure. Underlying exchange-rate data is from *Datastream*, with the USD as the base currency in our sample. We supplement this with *Consensus Economics* 6-month exchange rate forecasts for robustness. We complement this with data on the term structure of interest rates from government bond markets in each G.7 region. Yield curves are obtained from a combination of sources, including national central banks, Anderson and Sleath (2001), Gürkaynak, Sack, and Wright (2007), and Wright (2011). 6-month nominal zero-coupon bond yields provide our measure of short-term 'safe' interest rates. Like Lustig et al. (2019) and others, 10-year nominal zero-coupon bond yields serve as our proxy for the infinite-maturity bond yield, which amounts to assuming that the yield curve is sufficiently flat beyond the 10-year horizon. Using the full term structure, we construct bond term premia for each G.7 jurisdiction using the Adrian, Crump, and Moench (2013) decomposition of bond yields in our baseline. Data on dividend-price ratios and equity-price indices for each G.7 currency area are from *Global Financial Data.*[15]

4.1 Measuring Risk

We measure *ex ante* equity risk premia, $\log \mathbb{E}_t[R^{eq}_{t,t+1}/R_t]$, using equation (2). To do so, we construct measures of expected 6-month-ahead dividend growth as the fitted value from a regression of future 6-month dividend growth on lagged dividend growth and *Consensus Economics*' next-calendar-year GDP growth survey expectations (see Appendix B.2.1 for more detail). We proxy for the real risk-free rate using the nominal 6-month zero-coupon-bond yield adjusted with next-calendar-year inflation expectations from *Consensus Economics.*

Based on (18), our proxy for total SDF volatility, $TotRisk_t$, is constructed as:

$$TotRisk_t := \log \mathbb{E}_t \left[\frac{R^g_{t,t+1}}{R_t} \right] - \mathcal{L}_t \left[\frac{R^g_{t,t+1}}{R_t} \right] - \theta_t^{H,H(6M)} \tag{23}$$

where the first two terms capture the expected log risk premium $\mathbb{E}_t \log \left[\frac{R^g_{t,t+1}}{R_t} \right]$.

The first term is the log expected return on the (riskiest) portfolio in the economy, which comes from our valuation-based measure, equation (2). However, rather than using a raw equity index, we construct the growth-optimal portfolio (Bansal and Lehmann, 1997; Alvarez and Jermann, 2005), which maximizes the right-hand side of (18) and so acts as the appropriate return.[16] In this case, our raw U.S. risk measure is positive for 91% of our sample and to ensure country series are weakly positive, we set negative entries to zero. The second term ensures that

[15]Specifically, we use the S&P-500 for the US, EuroStoxx-50 for the EA, FTSE-100 for the UK, TOPIX for Japan, S&P/ASX-200 for Australia, S&P/TSX-300 for Canada and SMI for Switzerland.

[16]We calculate the growth-optimal portfolio using $R^g = \phi R^{eq} + (1 - \phi)R$, where $\phi = \log(\mathbb{E}_t[R^{eq}/R])/\sigma^2(\log(R^{eq}/R))$. Using moments from our sample, this yields $\phi = 2.27$ for the U.S. and $\phi^* = 1.66$ for the G.7 average. As such, the portfolio R^g corresponds to levered equity. See Appendix B.5 for further details.

our proxy for total risk is corrected for the expected volatility of returns, $\mathcal{L}_t\left(R^g_{t,t+1}/R_t\right)$.[17] The final term adjusts for within-country convenience, $\theta_t^{H,H(6M)}$, which we calculate using interest-swap spreads at 6-month (6M) maturities, as in Du et al. (2023), see Appendix B.4.

In turn, using equation (19), our proxy for permanent risk, $PermRisk_t$, is calculated as:

$$PermRisk_t := \log \mathbb{E}_t\left[\frac{R^g_{t,t+1}}{R_t}\right] - \mathcal{L}_t\left[\frac{R^g_{t,t+1}}{R_t}\right] - \mathbb{E}_t[rx^{(10Y)}_{t+1}] - \mathbb{E}_t\left[\theta^{H,H(10Y)}_{t,t+1}\right] \tag{24}$$

The first two terms follow from equation (23). The third term accounts for domestic long-maturity bond premia, for which we use the term premia measures constructed using the approach of Adrian et al. (2013) for each G.7 jurisdiction, adjusted for domestic inflation expectations. The final term again corrects for within-country convenience yields, in this case *holding-period* convenience yields on long-maturity bonds, which we calculate using interest-swap spreads at the 10-year (10Y) maturity.

We construct total and permanent risk for each G.7 region. We use U.S. risk for Home and the cross-sectional average of other G.7 jurisdictions for Foreign, to construct total ($TotRisk^*_t$) and permanent ($PermRisk^*_t$) proxies. We define cross-country differences by $\mathcal{D}$, such that $\mathcal{D}TotRisk_t := TotRisk_t - TotRisk^*_t$ and $\mathcal{D}PermRisk_t := PermRisk_t - PermRisk^*_t$. We also construct a proxy for transitory risk using bond premia.

Figure 4 presents our proxies for relative U.S. total ($\mathcal{D}TotRisk_t$, Panel (a)) and permanent ($\mathcal{D}PermRisk_t$, Panel (b)). Similar to expected relative equity premia (Figure 1(a)), relative U.S. total risk has generally increased over the past two decades, albeit with pronounced troughs during the aftermath of the dot-com bubble and the GFC. Importantly, the rise in relative U.S. total risk is attributable almost entirely to a rise in relative U.S. permanent risk (Panel (b)).[18] In turn, this rise in U.S. relative permanent risk is driven almost exclusively by the rise in U.S. relative equity premia (Figure 1a). When looking at U.S. risk in isolation, the trend is also partly attributable to falling real rates, but this effect cancels out in relative terms (see Appendix B.6). As a measure of transitory risk, Figure B.13 in Appendix B.6 plots the difference between total and permanent SDF volatility (Lemma 3). This measure has only risen modestly over our sample, driven by relatively higher U.S. bond premia and a decline in relative deviations from the expectations hypothesis for within-country convenience yields $\mathbb{E}_t\left[\theta^{H,H(10Y)}_{t,t+1}\right] - \theta^{H,H(6M)}_t$.

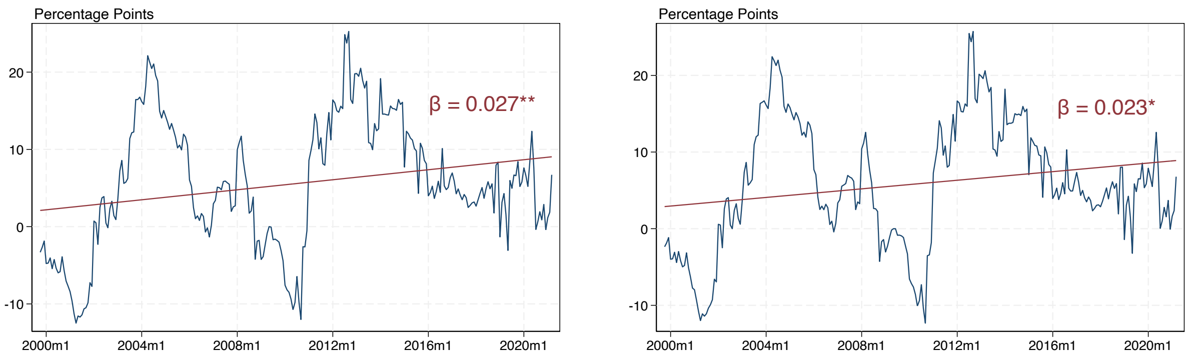
[17] Martin (2017) proxies for this using the squared volatility index, $VIX^2_t/2$, over the relevant holding period. The relevant volatility index for our environment is the 6-month VIX, which is not available for all G.7 markets and is only available for the US post-2008. We therefore construct a proxy for VIX using historical data and a Gaussian approximation based on $VIX^2_t \approx \frac{1}{T-t}\text{var}_t(\log R^g_{t,t+1})$, see Figure B.1 for a comparison.

[18] Figure B.16 shows that U.S. relative permanent risk has also risen systematically on a country-by-country basis, comparing pre- and post-2010 periods.

Figure 4: Proxies for Total and Permanent Risk, U.S. net G.7 Average

(a) U.S. net G.7 Relative Total Risk

(b) U.S. net G.7 Relative Permanent Risk



Note. Panels 4a and 4b display time series of our proxies for U.S. net average G.7 total and permanent risk, respectively. Total risk is calculated according to equation (23); Permanent risk is calculated according to equation (24). ***, **, * signify that the slope (β) from a regression of the variable on a time trend is different from zero at the 1%, 5%, 10% significance levels based on Newey and West (1987) standard errors with 4 lags, respectively. The sample runs from 1999:10 to 2021:03.

The valuation-based pricing we use to calculate expected returns is not without shortcomings. For example, dividends reflect only part of firms' total payouts (Atkeson et al., 2025b; Eeckhout, 2025) and are heavily smoothed by firms. However, as we look to ascertain moments of the SDF as the risk return operator of returns, dividend-based pricing is sensible. Encouragingly, our resulting measure of risk for the U.S. aligns with other measures offered by the literature. As well as the Livingston Survey detailed in Appendix B.2.2, Farhi and Gourio (2018) and Papanikolaou (2018) show valuation based measures are correlated with the long-run uncertainty estimated in Schorfheide, Song, and Yaron (2018), and is even close to the economic policy uncertainty measure in Baker, Bloom, and Davis (2016). Encouragingly, these measures also increase over our sample period.

4.2 Measuring Convenience Yields

To measure the excess convenience a Foreign investor earns on a U.S. Treasury ($\theta_t^{F,H(k)} - \theta_t^{F,F(k)}$), we use data on the Treasury basis from Du et al. (2018a). We construct short-maturity cross-country convenience using 6-month CIP deviations, $CIP_t^{(6M)}$,[19] and the 10-year tenor $CIP_t^{(10Y)}$ for long-maturity cross-country convenience.

However, CIP deviations are still a step removed convenience yields. The literature has

[19]Our 6-month CIP deviation is derived from the Du et al. (2018a) 3-month and 1-year CIP deviation data by linear interpolation, such that: $CIP_t^{(6M)} = \frac{2}{3}CIP_t^{(3M)} + \frac{1}{3}CIP_t^{(1Y)}$. We use the 6-month convenience yield to match the maturity from our zero-coupon bond data.

recognized that convenience may be related to either the 'dollarness' of U.S. Treasuries or the Treasuries themselves, or both, so may not exactly correspond to the Treasury basis. Focusing on short-maturities, Jiang et al. (2021a) propose a method to estimate the relative importance of the two that, in effect, scales-up observable CIP deviations to construct a proxy for the cross-country convenience yields inherent in holding US Treasuries. Consider foreign investors' convenience on synthetic U.S. Treasuries (i.e., Foreign government bonds swapped into dollars using a forward contract). This position earns the convenience of a Foreign bond, $\theta_t^{F,F,(k)}$, plus some maturity-specific fraction β_k^* of relative U.S. Treasury convenience, $\theta_t^{F,H(k)} - \theta_t^{F,F(k)}$, resulting in the following proportionality with CIP deviations:[20] $\theta_t^{F,H(k)} - \theta_t^{F,F(k)} = \frac{1}{1-\beta_k^*} CIP_t^{(k)}$.[21] We proxy convenience yields directly by CIP deviations since scaling does not change the relevant statistical properties of the series as long as β_k^* is constant—so our estimated coefficients implicitly capture the scaling factor. We discuss this in depth in Appendix B.3. Moreover, we adopt a similar strategy for $\theta_t^{H,H}$ and $\theta_t^{F,F}$ using swaps, following Du et al. (2023) (see Appendix B.4).

In addition, Proposition 2 clarifies that *holding-period* convenience—i.e., the *expected* convenience earned on a long-maturity bond position that is unwound a period later—is the relevant quantity in the relationship between currency returns, risk and convenience with long-maturity bonds. Consider a decomposition of convenience over a k-maturity portfolio's lifetime into a flow and a continuation value. For CIP deviations:

$$\theta_t^{F,H(k)} - \theta_t^{F,F(k)} = \omega^{(k)} \left(\theta_t^{F,H(k)} - \theta_t^{F,F(k)} \right) + \mathbb{E}_t[\theta_{t+1}^{F,H(k-1)} - \theta_{t+1}^{F,F(k-1)}]$$

for all $k > 1$. Investors receive a share $\omega^k \in (0,1)$ of their portfolio's convenience-to-maturity, $(\theta_t^{F,H(k)} - \theta_t^{F,F(k)})$ and expect to forgo the convenience on a $k-1$-maturity bond, $(\theta_{t+1}^{F,H(k-1)} - \theta_{t+1}^{F,F(k-1)})$, when selling the bond tomorrow. Taking the limit $k \to \infty$, the holding-period convenience yield $\mathbb{E}_t[\theta_{t,t+1}^{(\infty)}] = \omega^{(\infty)}(\theta_t^{F,H(\infty)} - \theta_t^{F,F(\infty)})$. As such, proxying long-maturity convenience with the 10-year CIP deviation, our long-maturity regression coefficients implicitly capture both

[20] The Euler equation for the synthetic position is:

$$\mathbb{E}_t \left[M_{t,t+k}^* \frac{\mathcal{E}_t}{\mathcal{E}_{t+k}} \left(\frac{F_t^{(k)}}{\mathcal{E}_t} R_t^{(k)*} \right) \right] = e^{-\theta_t^{F,F(k)} - \beta_k^* (\theta_t^{F,H(k)} - \theta_t^{F,F(k)})}$$

where the term in round brackets on the left-hand side is the synthetic k-period U.S. Treasury pecuniary return. Comparing this with equations (4) and (8), results in the linear, maturity-specific, relationship between the cross-country convenience yield and the CIP deviations. If $\beta_k^* = 1$, the Foreign investor values the synthetic Treasury exactly the same as a U.S.-issued Treasury, suggesting convenience arises from currency. If $\beta_k^* < 1$, then there is intrinsic convenience government bonds. As long as $\beta_k^* \in (0,1)$, small CIP deviations correspond to large convenience yields.

[21] Jiang et al. (2021a) identify β_k^* under the assumptions that convenience yields are stationary and independent from risk premia. While we verify these assumptions and reproduce their findings for short-maturity U.S. Treasuries (see Appendix B.3), both assumptions are invalid for long-maturity Treasuries.

the constant weight $\omega^{(k)}$ and $\frac{1}{1-\beta_k^*}$.[22]

A consequence of using the level of CIP deviations to proxy for holding-period convenience is that our proxy inherits the non-stationarity of long-maturity CIP deviations. Alongside this, while carry-trade returns have a constant mean and variance over the sample and are therefore stationary—i.e., integrated of order zero $\mathcal{I}(0)$—our measurement for relative U.S. permanent risk is non-stationary—i.e., integrated of order one $\mathcal{I}(1)$. In turn, these empirical regularities have testable implications for the long-run relationships implied by Proposition 2:

Corollary (Cointegration of Risk and Convenience Yields) *Let permanent risk differentials $\mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) - \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1})$ be $\mathcal{I}(1)$. If one-period long- maturity carry-trade returns $\mathbb{E}_t[rx^{CT,(\infty)}_{t+1}]$ are $\mathcal{I}(0)$, then (22) implies that there exists a unique linear combination of relative permanent risk and holding-period convenience yields $\theta_t^{F,H(\infty)} - \theta_t^{F,F(\infty)}$ that are cointegrated, i.e., $\mathcal{I}(0)$, such that holding-period convenience is $\mathcal{I}(1)$.*

5 Estimating Equilibrium Relationships

To what extent are the decline in long-maturity U.S. convenience yields and the increase in U.S. risk over the past decades connected? We estimate linear relationships between CIP deviations and relative permanent risk to quantify the importance of risk for the long-run evolution of convenience yields. For our baseline, we compare the U.S. to average values across the other G.7 countries and in Appendix B.7 we allow for country-specific long-run relationships.

5.1 Long-Maturity Convenience and Permanent Risk in the Long Run

Our focus is on the relationship between long-maturity carry-trade returns ($rx^{CT(10Y)}_{t+1}$), relative permanent risk ($\mathcal{D}PermRisk_t$), and the excess convenience yield foreigners earn when holding U.S. long-maturity bonds for a single period (proxied with $CIP_t^{(10Y)}$), summarized in Proposition 2. Our core result is that, in line with the Corollary, there is a single long-run cointegrating relationship between these three variables: specifically, a negative association between relative permanent risk and long-maturity CIP deviations.

[22] We treat within-country long-maturity convenience yields ($\theta_t^{H,H(10Y)}$ and $\theta_t^{F,F(10Y)}$) analogously, but specify the most conservative fraction for the holding period flow: $\omega^{(10Y)} = 1/20$, reflecting that a domestic investor earns $1/20^{th}$ of the domestic convenience yield of a 10-year bond over a 6-month holding period. Our results are robust, in fact are stronger, if we assume a larger fraction, since $\theta_t^{H,H(10Y)} - \theta_t^{F,F(10Y)}$ is trending downwards over our sample.

Table 1: Unit Root Test

Variable	ADF Test Statistic	
	Without Trend	With Trend
Panel A: Long-Maturity Variables		
$CIP_t^{(10Y)}$	-1.674	-2.910
$\mathcal{D}PermRisk_t$	-2.570	-2.518
$rx_{t+1}^{CT(10Y)}$	-4.222***	-4.242***
Panel B: Short-Maturity Variables		
$CIP_t^{(6M)}$	-3.442**	-3.444**
$\mathcal{D}TotRisk_t$	-2.575*	-2.522
rx_{t+1}^{FX}	-4.242***	-4.362***

Notes: Augmented Dickey-Fuller (ADF) tests (Dickey and Fuller, 1979), with 6 lags of change in dependent variable. Sample: 2000:01-2021:03. Null hypothesis: series is a random walk (without drift). Alternative hypothesis: series does not include a unit root. *** denotes $p < 0.01$, ** $p < 0.05$ and * $p < 0.10$.

As a first step to reaching this conclusion, we run a battery of stationarity tests for each of our long- and short-maturity variables. We rely on the augmented Dickey-Fuller (ADF) tests (Dickey and Fuller, 1979) with and without a deterministic time trend, including 6 lagged monthly changes of each variable to account for potential autocorrelation. The results are shown in Table 1, Panel A. The null hypothesis of non-stationarity is strongly rejected for carry-trade returns on long-maturity bonds at the 1% level, both with and without a deterministic trend included in the test. In contrast, relative permanent risk is non-stationary over our sample; even allowing for a deterministic trend, and as such the null hypothesis cannot be rejected. Likewise, long-maturity CIP deviations appear non-stationary, regardless of the inclusion of a deterministic trend. These findings contrast with short-maturity carry-trade returns and CIP deviations which appear stationary (Panel B).

The ADF test verifies the assumptions underpinning our Corollary, raising the possibility of cointegration between long-maturity variables in the data. To test this, we rely on the *trace* cointegration (Johansen, 1991) and maximum-eigenvalue tests to investigate the number of long-run relationships between risk, returns and convenience. Table 2 provides strong evidence of there being a *unique* cointegrating relationship between long-maturity variables.

Having established cointegration, Table 3 Panel A identifies the unique long-run relationship as a negative relationship between long-maturity convenience yields and relative permanent risk. We present coefficient estimates from OLS regression, in levels, of long-maturity CIP deviations on relative permanent risk and carry-trade returns on long-maturity bonds. Accounting for a deterministic trend (column (1)), a 1p.p. increase in U.S. permanent risk relative to the

Table 2: Inference on Cointegration

Null Hypothesis	*trace*	5% Crit. Val.	λ_{max}	5% Crit. Val.
$r = 0$	69.50	29.68	57.49	20.97
$r \leq 1$	**12.01**	**15.41**	**7.150**	**14.07**
$r \leq 2$	4.86	3.76	4.86	3.76

Notes: Trace-test *trace* (Johansen, 1991) and maximumum-eigenvalue test λ_{max} statistics for number of cointegrating vectors r characterizing permanent risk, long-maturity convenience yields and carry-trade returns. Sample: 2000:01-2021:03. Null hypothesis: r less than or equal to given quantity. Alternative hypothesis: $r + 1$ cointegrating vectors. Carrying out tests sequentially from $r = 0$, emboldened quantities denote the first set of test statistics where null hypothesis can no longer be rejected.

remaining G.7 is associated with approximtely a 0.5b.p. decline in U.S. long-maturity CIP deviations over the long run. Excluding the deterministic trend (column (2)), the association is stronger, up to almost 1.0b.p.. The deterministic trend serves an important role since it captures other possible secular explanations underpinning the decline in long-maturity convenience yields, which we discuss further in Section 5.4. In contrast, there is no significant relationship between long-maturity carry-trade returns and convenience. Column (3) provides evidence on the relative importance of U.S. risk and Foreign risk. Convenience yields are driven by changes in U.S. permanent risk over and above Foreign risk. Column (4) shows our results are robust to using realized bond premia as opposed to the Adrian et al. (2013) measure for expected premia. Finally, column (5) provides a falsification test for our theory in the form of a placebo test. Replacing permanent-risk differentials with our proxy for relative transitory risk (Figure B.13) indeed yields no relationship with long-maturity convenience yields.

Appendix B.7 confirms the negative relationship between convenience yields and permanent risk differentials in a panel setting, where we estimate separate long-run regression models for the U.S. against each other G.7 countries. The mean-group estimate in Table B.8 suggests that, on average across regressions, a 1p.p. increase in U.S. permanent risk is associated with an approximately 0.3b.p. decline in U.S. long-maturity CIP deviations, and the coefficient is statistically significant.

5.2 Dynamics of Long-Maturity Convenience and Permanent Risk

To characterize the fuller dynamics of the relationship between long-maturity convenience and relative permanent risk, we estimate variants of the following error-correction model, derived from an autoregressive distributed-lag model implied by Proposition 2:

$$\Delta CIP_t^{(10Y)} = \beta_0 + \beta_1 \Delta \mathcal{D}PermRisk_t + \beta_2 \Delta rx_{t+1}^{CT(10Y)} + \gamma \left[CIP_{t-1}^{(10Y)} - \alpha_1 \mathcal{D}PermRisk_{t-1} - \alpha_2 rx_t^{CT(10Y)} \right] + \varepsilon_t \quad (25)$$

where $CIP_t^{(10Y)}$ (and its changes) are measured in basis points, while other variables are measured in percent points. The first-difference terms, pre-multiplied by β_i for $i = 1, 2$, capture the contemporaneous response of long-maturity convenience to changes in relative permanent risk and long-maturity carry-trade returns—i.e., the short-run adjustment. The term pre-multiplied by α_i for $i = 1, 2$ captures the cointegrating relationship implied by the data. The coefficients α_i for $i = 1, 2$ capture the long-run multipliers between long-maturity convenience and relative-permanent risk and carry-trade returns, respectively. In turn, the coefficient γ reflects the degree of adjustment in long-maturity convenience to deviations from the long-run relationship.

We estimate regression (25) using the two-step method of Engle and Granger (1987). The long-run coefficients (α_i), which we discussed in the previous sub-section, are estimated in the first stage and are reported in Panel A of Table 3. The second-stage coefficients are shown in Panel B. The coefficients on disequilibrium adjusted ($\hat{\gamma}$) and permanent risk are statistically significant, so, all else equal, higher permanent risk contributes to lower relative convenience not only in the long run, but also not contemporaneously. Moreover, while there is no significant long-run relationship between long-maturity carry-trade returns and long-maturity U.S. convenience yields, there is some limited scope for adjustment in response to changes in carry-trade returns in the short run: higher U.S. convenience yields are associated with higher currency returns on Foreign bonds, consistent with theory. In the models with deterministic trends, shown in columns (1) and (3), Engle-Granger tests for cointegration, compared to critical values from MacKinnon (2010), reject the null hypothesis of no cointegration.

5.3 Contribution of Permanent Risk to Long-Maturity Convenience

To illustrate the economic significance of our results, using the error-correction model (25) with coefficients estimated in Table 3 Column (1), we carry out pseudo out-of-sample counterfactual exercises. Using the model, we ask: given the path of long-maturity carry-trade returns and the deterministic trend, how would long-maturity CIP deviations have evolved had U.S. permanent risk not increased over time?

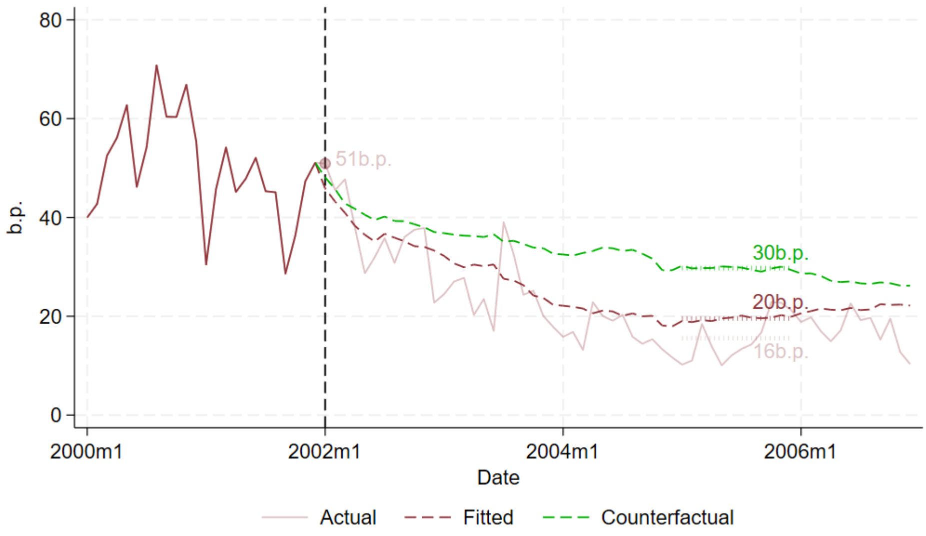
For our first experiment, we calculate the average value of relative permanent risk from

Table 3: Cointegration and Error-Correction Amongst Long-Maturity Variables

	(1)	(2)	(3)	(4)	(5)
		$\mathbb{P}$ Risk			$\mathbb{T}$ Risk Proxy
Panel A: Long-Run Adjustment					
$\mathbb{P}$ Risk Measure		Baseline		Realized Bond-Prem.	
$\mathcal{D}PermRisk_t$	-0.550***	-0.916***		-0.532***	
	(0.150)	(0.316)		(0.190)	
$PermRisk_t$			-0.484***		
			(0.165)		
$PermRisk^*_t$			-0.040		
			(0.372)		
$\mathcal{D}TransRisk_t$					-0.425
					(2.086)
$rx_t^{CT(10Y)}$	0.129	0.203	0.102	0.114	0.140
	(0.109)	(0.203)	(0.118)	(0.118)	(0.129)
Deterministic Trend	-0.204***		-0.197***	-0.202***	-0.214***
	(0.017)		(0.019)	(0.020)	(0.026)
Panel B: Short-Run Adjustment					
$\Delta\mathcal{D}PermRisk_t$	-0.303**	-0.252*		-0.279*	
	(0.139)	(0.144)		(0.142)	
$\Delta PermRisk_t$			-0.316**		
			(0.140)		
$\Delta PermRisk^*_t$			0.008		
			(0.333)		
$\Delta\mathcal{D}TRisk_t$					-3.015**
					(1.318)
$\Delta rx_{t+1}^{CT(10Y)}$	0.101*	0.099	0.092	0.099*	0.076
	(0.059)	(0.061)	(0.060)	(0.059)	(0.061)
Diseq. Adjustment $\hat{\gamma}$	-0.212***	-0.063***	-0.218***	-0.200***	-0.171***
	(0.040)	(0.022)	(0.040)	(0.039)	(0.036)
Adj. R^2 of ECM	0.108	0.036	0.109	0.102	0.103
Engel-Granger Test Stat.	-4.665**	-2.630	-4.777**	-4.523**	-4.128*
Deterministic Trend	Yes	No	Yes	Yes	Yes
# Observations	254	254	254	254	252

Notes: Panel A reports coefficient estimates from regressions of long-maturity CIP deviations (in basis points) on relative permanent risk (in percent points) and long-maturity carry-trade returns (in percent points). Panel B reports coefficient estimates from second-stage Engle and Granger (1987) regressions. Sample: 2000:01-2021:03. Newey and West (1987) standard errors with 6 lags reported, with *** denoting $p < 0.01$, ** $p < 0.05$, and * $p < 0.1$.

Figure 5: Counterfactual Paths for Long-Maturity CIP Deviation 2002:01-2005:12

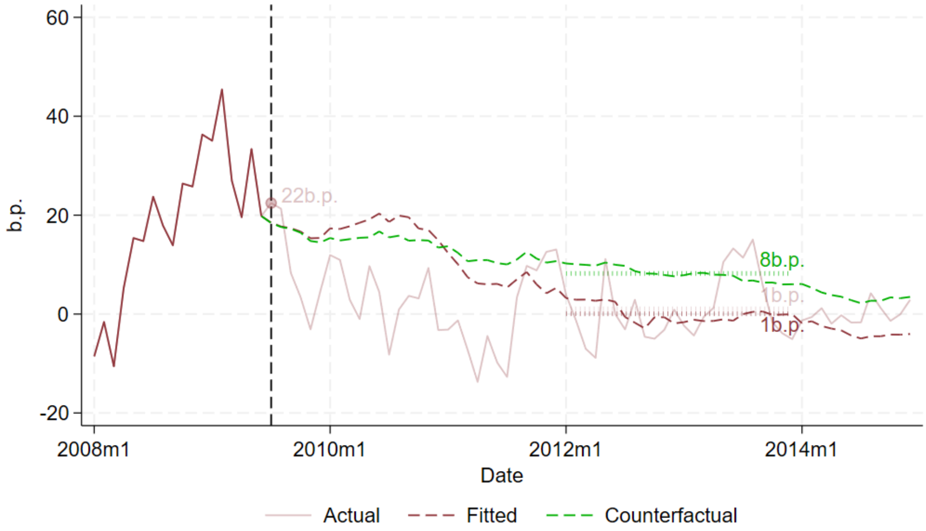


Notes: Counterfactual paths constructed using estimates of error-correction model (25) from Panels B and C Table 3 Column 1. Counterfactuals constructed by setting relative permament risk to average level over period 2000:01-2001:12, from 2002:01 to 2005:12. Other series taken from data.

2000:01 to 2001:12, near its dot-com bubble trough, then fix relative permanent risk to this value from 2002:01 to 2005:12. We then construct a counterfactual path for long-maturity convenience yields over this four-year period. Figure 5 plots this counterfactual path alongside the actual evolution of 10-year CIP deviations and the fitted path from our empirical model. At the starting point for our counterfactual in 2002:01, 10-year convenience yields are 51b.p. The full error-correction model fits the actual data well although it misses many of the high-frequency fluctuations. Most importantly, by 2005 the model predicts a fall in convenience yields to 20b.p., a 31b.p. drop. This compares to a decline in the actual data of 35b.p. to 16b.p. over the same period, such that the empirical model captures around 89% of the secular decline. In contrast, when abstracting from the increase in relative U.S. permanent risk over the 2002-2005 period, CIP deviations decline by much less. Specifically, the counterfactual path for CIP deviations deviates immediately and proceeds to drop by only 21b.p., to 30b.p. That is, through the lens of our model, the rise in U.S. permanent risk vs. the rest of the G.7 in the years following the dot-com bubble (and preceding the GFC) appears to explain around one-third of the decline in long-maturity U.S. Treasury convenience yields.

In the second experiment, we fix relative permanent risk at near its trough around the GFC. Here, we calculate the average value of relative permanent risk from 2008:07 to 2009:06. We then fix relative permanent risk at this level from 2009:07 to end-2013. Figure 6 compares

Figure 6: Counterfactual Paths for Long-Maturity CIP Deviation 2010:02-2016:12:.



Notes: Counterfactual paths constructed using estimates of error-correction model (25) from Panels B and C Table 3, Column (1). Counterfactuals constructed by setting relative permanent risk to 2010:02 trough, from 2010:03 to 2016:12. Other series taken from data.

the actual, fitted and counterfactual paths for 10-year CIP deviations. The full error-correction model captures almost the entirety of the 21b.p. decline in CIP deviations, with both actual and fitted values averaging 1b.p. over the period 2012:01-2013:12. Compared to fitted values, the counterfactual path starts to diverge during 2011, from which point it declines at a slower pace. Over the 2012:01-2013:12 period, the counterfactual path for 10-year CIP deviations is 7b.p. above the fitted values relative to the full model. That is, over this period, permanent risk differentials appear to drive one-third of the decline in convenience yields.

Both experiments lend strong support, at least for long maturity assets, to the prediction of the class of models which suggest that in equilibrium, risk differentials must be compensated by either pecuniary *or* non-pecuniary returns, characterized by variants of (1). In the data, carry-trade (pecuniary) returns do not move enough to compensate for significant movements in relative permanent risk between the U.S. and G.7. As a result, convenience must move to reflect risk differentials. When perceived U.S. risk increases, as is reflected in higher expected returns for risky U.S. assets, cross-country convenience yields (i.e., CIP deviations) on long-maturity assets tend to fall, reflecting a fall in Foreign investors' preferences for U.S. bonds.

5.4 Risk, Dollar Scarcity and Other Measures

Convenience yields have been shown to depend on the supply of dollar assets (Krishnamurthy and Vissing-Jorgensen, 2012; Jiang et al., 2024). We investigate whether the relationship between risk differentials and convenience yields is robust to controlling for debt-to-GDP—a proxy for dollar debt supply. In other words we ask: does the quantity of outstanding bonds still matters when controlling for the level of risk? Table 4 Column (1) shows that, absent our deterministic trend, both permanent risk differentials and U.S. debt-to-GDP matter for convenience yield determination, in the directions predicted by theory. Specifically, a 1p.p. increase in U.S. debt-to-GDP results in a 0.66p.p. decline in the excess convenience yield on U.S. long maturity and the coefficient is significant. However, Column (2) shows that once you control for a deterministic trend, debt supply becomes insignificant while relative permanent risk remains significant with the coefficient largely unchanged. Column (3) confirms this result is robust to using a measure of long-maturity debt specifically.

Column (4) adds the VIX as a control and our headline result for permanent risk is largely unchanged. Nonetheless, the VIX is significant for long-run adjustment: a 1% increase in the VIX corresponds to 0.05b.p. increase in convenience yields—consistent with the idea that foreigners look for U.S. assets during periods of uncertainty, once we have controlled for permanent risk differentials. Column (5) replaces realized carry-trade returns with expected returns constructed using 6-month-ahead survey expectations for exchange rates and Adrian et al. (2013) bond premia. Again, the coefficient for permanent risk differentials is stable and significant. Although, in the long run, higher CIP deviations do correspond to higher expected currency returns, this is almost entirely offset by movements in permanent risk.

More generally, a key result of our analysis is that our measures of risk and relative risk matter for both short-run and long-run (trend) adjustment in convenience yields. As such, our results would not be robust to *replacing* our measures of risk with stationary measures such as the VIX or SVIX (see, e.g., Martin, 2017). However, there are theoretical arguments by which these may not be an appropriate measure in our framework.[23] Instead, our results suggest looking to trending measures of risk such as the Livingston Survey for risk premia (Figure B.6) or measures of policy uncertainty (e.g., Baker et al., 2016).

[23]Specifically, Martin (2017) shows that $\sigma_t(M_{t+1})\text{var}_t^{RN}(R_{t+1}^g) \geq ERP_{t+1} \geq R_t \times SVIX_t^2$ where var_t^{RN} refers to the conditional risk-neutral variance. Our measure relies on the middle argument in the inequality to proxy for the LHS, whereas the $SVIX_t$ appears in the RHS, which could be increasingly biased over time.

Table 4: Cointegration and Error-Correction Amongst Long-Maturity Variables When Controlling for U.S. Treasury Scarcity and VIX, and Using Survey Expectations for FX.

	(1)	(2)	(3)	(4)	(5)
Panel A: Long-Run Adjustment					
$\mathcal{D}PermRisk_t$	-0.502***	-0.602***	-0.610***	-0.549***	-0.497***
	(0.187)	(0.201)	(0.215)	(0.168)	(0.129)
$rx_{t+1}^{CT(10Y)}$	0.158	0.118	0.143	0.129	
	(0.137)	(0.116)	(0.121)	(0.117)	
$\mathbb{E}rx_{t+1}^{CT(10Y)}$					0.523***
					(0.114)
U.S. Debt/GDP	-0.664***	0.381			
	(0.087)	(0.254)			
U.S. Long Debt/GDP			0.266		
			(0.277)		
U.S. VIX				0.052**	
				(2.522)	
Deterministic Trend		-0.306***	-0.267***	-0.201***	-0.193***
		(0.069)	(0.072)	(0.018)	(0.015)
Panel B: Short-Run Adjustment					
$\Delta\mathcal{D}PermRisk_t$	-0.264*	-0.312**	-0.314**	-0.278**	-0.285**
	(0.141)	(0.139)	(0.139)	(0.140)	(0.063)
$\Delta rx_{t+1}^{CT(10Y)}$	0.098	0.096	0.101*	0.075	
	(0.061)	(0.059)	(0.059)	(0.061)	
$\mathbb{E}\Delta rx_{t+1}^{CT(10Y)}$					0.150**
					(0.063)
Δ U.S. Debt/GDP	0.196	0.379			
	(0.568)	(0.555)			
Δ U.S. Long Debt/GDP			-1.007		
			(0.888)		
Δ U.S. VIX				0.0230	
				(1.864)	
Diseq. Adjustment $\hat{\gamma}$	-0.151***	-0.224***	-0.221***	-0.221***	-0.283***
	(0.034)	(0.041)	(0.040)	(0.040)	(0.043)
Adj. R^2 of ECM	0.075	0.112	0.112	0.117	0.158
Engel-Granger Test Stat.	-3.967*	-4.859**	-4.741**	-4.773**	-5.686***
Deterministic Trend	No	Yes	Yes	Yes	Yes
# Observations	254	254	254	254	254

Notes: Panel A reports coefficient estimates from regressions of long-maturity CIP deviations (in basis points) on relative permanent risk (in percent points), long-maturity carry-trade returns (in percent points), U.S. Debt/GDP measures (in percent points) and the U.S. VIX (in percent points). Column (5) uses *Consensus Economics* survey data for 6 month ahead exchange rates. Panel B reports coefficient estimates from second-stage Engle and Granger (1987) regressions. Sample: 2000:01-2021:03. Newey and West (1987) standard errors with 6 lags reported, with *** denoting $p < 0.01$, ** $p < 0.05$, and * $p < 0.1$.

Table 5: Short-Maturity Associations

	(1)	(2)	(3)	(4)
Dep. Var.: $\Delta CIP_t^{(6M)}$				
Δrx_{t+1}^{FX}	0.392*	0.389*	0.391	
	(0.232)	(0.233)	(0.238)	
$\mathbb{E}_t \Delta rx_{t+1}^{FX}$				0.564*
				(0.294)
$\Delta \mathcal{D}TotRisk_t$		0.426		0.539
		(0.834)		(0.819)
$\Delta TotRisk_t$			0.434	
			(0.902)	
$\Delta TotRisk_t^*$			−0.278	
			(1.069)	
Adjusted R^2	0.008	0.008	0.004	0.012

Notes: Coefficient estimates from regressions of short-maturity CIP deviations (in basis points) on relative total risk (in percent points) and short-maturity carry-trade returns (in percent points). Sample: 2000:01-2021:03. Newey and West (1987) standard errors with 6 lags reported, with *** denoting $p < 0.01$, ** $p < 0.05$, and * $p < 0.1$.

5.5 Short-Maturity Convenience and Risk

Finally, we contrast the results for long-maturities with a simple regression of short-maturity CIP deviations on the corresponding carry-trade returns and the measure relative *total* risk, based on Proposition 1. Table 5 reveals striking differences between the relationships for short- and long-maturity maturity portfolios. At short maturities, the relationship between relative risk and CIP deviations is not significantly different from zero: convenience on short-maturity bonds appear insulated from total risk. This is suggestive of further segmentation—i.e., investors active in short-lived and long-lived asset markets appear distinct (see, e.g., Du et al., 2023). On the other hand, higher short-maturity CIP deviations are associated with a contemporaneously appreciated USD, both using realized returns (Columns (1) & (2)) and survey-based expected returns (Column (4)), in line with Jiang et al. (2021a) and Engel and Wu (2023).

Appendix C.2 details an extension of the model where we assume markets are further segmented within countries. Specifically, investors active in markets with long-lived assets such as long-maturity bonds and equities are characterized by SDF process $\{M_{t+1}\}$, whereas investors active in markets with short bonds are characterized by $\{\hat{M}_{t+1}\}$. In this extended environment the long-maturity equilibrium relationship is unaffected by this form of segmentation, but the short-maturity relationship is distorted, rationalizing the findings above.

In sum: over short maturities, cross-border convenience co-moves with carry-trade returns, not with relative risk. Over long maturities, permanent risk differentials matter for convenience and the combination of rising relative permanent risk and declining convenience offset to leave long-maturity carry-trade returns, and therefore the U.S. dollar, largely insulated.

6 Conclusion

In this paper we use asset prices to ascertain whether foreign investors have changed their valuation of USD and dollar assets relative to a panel of other G.7 countries—and if so, which dollar assets specifically and at what maturities. Our analysis is motivated by striking patterns in the data which raise a number of questions. Namely, how can U.S. equities deliver higher (expected) returns relative to other G.7 countries (i.e., command higher risk premia), while short-maturity U.S. bonds continue to be considered particularly safe? Relatedly, why do long-maturity U.S. bonds appear to have lost their 'specialness' relative to other near-default-free government bonds—a trend that started in the early 2000s, while excess returns on carry trades funded in U.S. dollars have remained stable? We find that investor perceptions of U.S. risk have increased, while pecuniary returns to USD have remained insulated. This is consistent with a no-arbitrage framework extended to allow for falling convenience yields. Moreover, because risk pertains to revisions of long-term outcomes (i.e., permanent risk), it is particularly salient in the data for long-maturity bonds.

Our analysis has rich implications for policy and theory. As Atkeson et al. (2025a) argue, rising relative U.S. equity returns have contributed to a deterioration of net foreign income for the U.S., while Jiang et al. (2024) argue that fiscal sustainability has been impacted by the falling (and now negative) long-maturity convenience yields over the same period. We show that these two observations are closely interconnected. Indeed, our measure of risk correlates positively with U.S. debt-to-GDP, as well as indices of economic policy uncertainty. Moreover, our analysis highlights a specific role for rising relative permanent risk (i.e., revisions of expectations for an event expected in the distant future) as an explanatory factor behind the decline in long-maturity convenience yields, consistent with a 'scarring of beliefs' (Kozlowski et al., 2019). The importance of permanent risk in turn raises questions for future research around effective tools for stabilization policy since, traditionally, these typically target shorter business-cycle horizons.

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Appendix

A Model Proofs and Derivations

A.1 Deriving the Exchange-Rate Process

Combining (5) and (7), evaluating at $k = 1$, and expanding yields:

$$\mathbb{E}_t[\Delta e_{t+1}] + \mathcal{L}_t\left(\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}\right) + \mathcal{C}_t\left(M_{t,t+1}, \frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}\right) + r_t^{*(1)} - r_t^{(1)} + \theta_t^{H,F(1)} - \theta_t^{H,H(1)} = 0 \tag{A.1}$$

Then, using (6) and (8):

$$-\mathbb{E}_t[\Delta e_{t+1}] + \mathcal{L}_t\left(\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}\right) + \mathcal{C}_t\left(M^*_{t,t+1}, \frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right) + r_t^{(1)} - r_t^{*(1)} + \theta_t^{F,H(1)} - \theta_t^{F,F(1)} = 0 \tag{A.2}$$

Combining the above yields (9).

A.2 Proof to Proposition 1

Taking expectations of the exchange-rate process (10), and substituting in for $M_{t,t+1}$ and $M^*_{t,t+1}$ using the log-entropy expansions of the $k = 1$-period domestic Euler equations (5) and (6), respectively, we can write:

$$\mathbb{E}_t[\Delta e_{t+1}] = -r_t^* - \theta_t^{F,F(1)} - \mathcal{L}_t(M^*_{t,t+1}) + r_t + \theta_t^{H,H(1)} + \mathcal{L}_t(M_{t,t+1}) + \mathbb{E}_t[\eta_{t+1}]$$

Using Assumption 2, applying the definition of the *ex post* excess currency return (15), and rearranging yields the result. □

A.3 Proof to Lemma 2

Start by rewriting the Home Euler equation for the Home k-period bond (5):

$$\begin{aligned}
e^{-\theta_t^{H,H(k)}} =& \mathbb{E}_t\left[M_{t,t+k}R_t^{(k)}\right] \\
e^{-\theta_t^{H,H(k)}} =& \mathbb{E}_t\left[\frac{\Lambda_{t+k}}{\Lambda_t}\frac{1}{P_t^{(k)}}\right] \\
\Rightarrow \; P_t^{(k)} =& \mathbb{E}_t\left[\frac{\Lambda_{t+k}}{\Lambda_t}\right]e^{\theta_t^{H,H(k)}}
\end{aligned}$$

so, also $P_{t+1}^{(k-1)} = \mathbb{E}_{t+1}\left[\frac{\Lambda_{t+k}}{\Lambda_{t+1}}\right] e^{\theta_{t+1}^{H,H(k-1)}}$.

Now, solve for the one-period holding return on a long-term bond, following similar steps to Alvarez and Jermann (2005) in their proof to Proposition 2(i):

$$
\begin{aligned}
R_{t,t+1}^{(\infty)} &\equiv \lim_{k\to\infty} R_{t,t+1}^{(k)} \\
&= \lim_{k\to\infty} \frac{P_{t+1}^{(k-1)}}{P_t^{(k)}} \\
&= \lim_{k\to\infty} \frac{e^{\theta_{t+1}^{H,H(k-1)}} \cdot \mathbb{E}_{t+1}\left[\frac{\Lambda_{t+k}}{\Lambda_{t+1}}\right]}{e^{\theta_t^{H,H(k)}} \cdot \mathbb{E}_t\left[\frac{\Lambda_{t+k}}{\Lambda_t}\right]} \\
&= \frac{\lim_{k\to\infty} e^{\theta_{t+1}^{H,H(k-1)}} \cdot \frac{\mathbb{E}_{t+1}[\Lambda_{t+k}]/\beta^{t+k}}{\Lambda_{t+1}}}{\lim_{k\to\infty} e^{\theta_t^{H,H(k)}} \cdot \frac{\mathbb{E}_t[\Lambda_{t+k}]/\beta^{t+k}}{\Lambda_t}} \\
&= e^{-\theta_{t,t+1}^{H,H(\infty)}} \frac{\lim_{k\to\infty} \frac{\mathbb{E}_{t+1}[\Lambda_{t+k}]/\beta^{t+k}}{\Lambda_{t+1}}}{\lim_{k\to\infty} \frac{\mathbb{E}_t[\Lambda_{t+k}]/\beta^{t+k}}{\Lambda_t}} \\
&= e^{-\theta_{t,t+1}^{H,H(\infty)}} \frac{\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_{t+1}}}{\frac{\Lambda_t^{\mathbb{P}}}{\Lambda_t}} \\
&= e^{-\theta_{t,t+1}^{H,H(\infty)}} \frac{\Lambda_t^{\mathbb{T}}}{\Lambda_{t+1}^{\mathbb{T}}},
\end{aligned}
$$

where line 1 is a definition, line 2 uses definition for holding-period returns, line 3 uses bond-pricing and Euler equations, line 4 rearranges and multiplies and divides by β^{t+k}, line 5 defines convenience holding-period return, line 6 uses the ådefinition of $\Lambda^{\mathbb{P}}$, and line 7 uses the definition of the pricing kernel $\Lambda = \Lambda^{\mathbb{P}}\Lambda^{\mathbb{T}}$. Rearranging this final expression yields the result in equation (17). □

A.4 Proof to Lemma 3

Taking logs of (13):

$$\log \mathbb{E}_t\left[M_{t,t+1}R_{t,t+1}^g\right] = 0$$

By concavity of the log, we have that:

$$
\begin{aligned}
\log \mathbb{E}_t\left[M_{t,t+1}R_{t,t+1}^g\right] &= 0 \geq \mathbb{E}_t \log\left[M_{t,t+1}R_{t,t+1}^g\right] \\
\Rightarrow -\mathbb{E}_t \log M_{t,t+1} &\geq \mathbb{E}_t \log[R_{t,t+1}^g]
\end{aligned}
$$

Using this in the definition of the entropy measure, we can derive:

$$\begin{aligned}
\mathcal{L}_t(M_{t,t+1}) &= \log \mathbb{E}_t[M_{t,t+1}] - \mathbb{E}_t \log M_{t,t+1} \\
\mathcal{L}_t(M_{t,t+1}) &\geq \log \mathbb{E}_t[M_{t,t+1}] + \mathbb{E}_t \log[R^g_{t,t+1}] \\
\mathcal{L}_t(M_{t,t+1}) &\geq \log \left[\frac{1}{R_t^{(1)}} e^{-\theta_t^{H,H(1)}}\right] + \mathbb{E}_t \log[R^g_{t,t+1}] \\
\mathcal{L}_t(M_{t,t+1}) &\geq \mathbb{E}_t \log \left[\frac{R^g_{t,t+1}}{R_t}\right] - \theta_t^{H,H(1)}
\end{aligned}$$

where line 1 is the definition of entropy, line 2 follows from the inequality derived above, line 3 uses the Home Euler for a Home one-period bond (5), and line 4 rearranges. This verifies (18).

Next, decompose total risk starting with the definition of conditional entropy $\mathcal{L}_t(\cdot)$:

$$\begin{aligned}
\mathcal{L}_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) &= \log \mathbb{E}_t\left[\frac{\Lambda_{t+1}}{\Lambda_t}\right] - \mathbb{E}_t \log \frac{\Lambda_{t+1}}{\Lambda_t} \\
&= \log\left(P_t^{(1)} e^{-\theta_t^{H,H(1)}}\right) - \mathbb{E}_t \log \frac{\Lambda^{\mathbb{P}}_{t+1}\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{P}}_t \Lambda^{\mathbb{T}}_t} \\
&= \log\left(P_t^{(1)} e^{-\theta_t^{H,H(1)}}\right) - \mathbb{E}_t \log \frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t} - \mathbb{E}_t \log \frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t} \\
&= \log\left(P_t^{(1)} e^{-\theta_t^{H,H(1)}}\right) + \mathcal{L}_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) - \log \mathbb{E}_t \frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t} - \mathbb{E}_t \log \frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t} \\
&= \log\left(\frac{1}{R_t^{(1)}} e^{-\theta_t^{H,H(1)}}\right) + \mathcal{L}_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) - \mathbb{E}_t \log\left(e^{-\theta_{t,t+1}^{H,H(\infty)}} \cdot \frac{1}{R^{(\infty)}_{t,t+1}}\right) \\
&= -\log R_t^{(1)} - \theta_t^{H,H(1)} + \mathcal{L}_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) + \mathbb{E}_t \theta_{t,t+1}^{H,H(\infty)} + \mathbb{E}_t \log R^{(\infty)}_{t,t+1} \\
&= \mathbb{E}_t \log \frac{R^{(\infty)}_{t,t+1}}{R_t^{(1)}} + \mathcal{L}_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) - \theta_t^{H,H(1)} + \mathbb{E}_t \theta_{t,t+1}^{H,H(\infty)}
\end{aligned} \tag{A.3}$$

where line 1 is a definition, line 2 uses the pricing expression and the definition of the pricing kernel, line 3 separates the second term, line 4 uses the definition $\mathcal{L}_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) = \log \mathbb{E}_t \frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t} - \mathbb{E}_t \log \frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}$, line 5 uses the facts that $P_t^{(1)} = 1/R_t^{(1)}$, $\log \mathbb{E}_t \frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t} = 0$ and (from Lemma 2) $\frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t} = e^{-\theta_{t,t+1}^{H,H(\infty)}} \cdot \frac{1}{R^{(\infty)}_{t+1}}$, line 6 rearranges, line 7 concludes. Rearranging this final expression, and using definitions of SDFs and excess returns, yields the result.

Finally, this can be written as:

$$\mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) + \mathbb{E}_t[rx^{(\infty)}_{t+1}] - \theta_t^{H,H(1)} + \mathbb{E}_t[\theta_{t,t+1}^{H,H(\infty)}] \geq \mathbb{E}_t \log\left[\frac{R^g_{t,t+1}}{R_t}\right] - \theta_t^{H,H(1)}$$

$$\mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) + \mathbb{E}_t[\theta^{H,H(\infty)}_{t,t+1}] \geq \mathbb{E}_t \log\left[\frac{R^g_{t,t+1}}{R_t}\right] - \mathbb{E}_t[rx^{(\infty)}_{t+1}]$$

where line 1 substitutes and line 2 rearranges to yield condition (19). □

A.5 Proof to Lemma 4

Start with the one-period exchange-rate process (10) in logs:

$$\mathbb{E}_t[\Delta e_{t+1}] = \mathbb{E}_t[m^*_{t,t+1}] - \mathbb{E}_t[m_{t,t+1}] + \mathbb{E}_t[\eta_{t+1}]$$

and note that, since there is trade in assets in every period, it follows that:

$$\mathbb{E}_t[\Delta e_{t+\tau}] = \mathbb{E}_t[m^*_{t+\tau-1,t+\tau}] - \mathbb{E}_t[m_{t+\tau-1,t+\tau}] + \mathbb{E}_t[\eta_{t+\tau}]$$

for all τ. Noting that $m^{(*)}_{t+\tau-1,t+\tau} \equiv \lambda_{t+\tau} - \lambda_{t+\tau-1}$, where $\lambda_{t+\tau} \equiv \log(\Lambda_{t+\tau})$, we can sum the exchange-rate process k-periods forward:

$$\begin{aligned}\mathbb{E}_t[\Delta e_{t+k} + \ldots + \Delta e_{t+1}] =& \mathbb{E}_t[m^*_{t+k-1,t+k} - m_{t+k-1,t+k} + \ldots + m^*_{t,t+1} - m_{t,t+1}] \\ &+ \mathbb{E}_t[\eta_{t+k} + \cdots + \eta_{t+1}]\end{aligned}$$

Alongside this, consider the k-period analog of the exchange-rate process (10):

$$e_{t+k} - e_t = m^*_{t,t+k} - m_{t,t+k} + \eta_{t+k}$$

Comparing these final two expressions, we see that the following restriction on the term-structure of incomplete-market wedges—which implies a term structure of convenience yields—must hold:

$$\sum_{\tau=0}^{k-1} \mathbb{E}_t[\eta_{t+\tau+1}] = \mathbb{E}_t[\eta_{t+k}], \tag{A.4}$$

Expanding (A.4) and using (11) and (12) :

$$\theta^{F,H(k)}_t - \theta^{H,H(k)}_t = \sum_{\tau=0}^{\kappa-1}\left\{\mathbb{E}_t[\theta^{F,H(1)}_{t+\tau}] - \mathbb{E}_t[\theta^{H,H(1)}_{t+\tau}]\right\} \tag{A.5}$$

$$\theta^{F,F(k)}_t - \theta^{H,F(k)}_t = \sum_{\tau=0}^{\kappa-1}\left\{\mathbb{E}_t[\theta^{F,F(1)}_{t+\tau}] - \mathbb{E}_t[\theta^{H,F(1)}_{t+\tau}]\right\} \tag{A.6}$$

which correspond to (20) and (21). □

A.6 Proof to Proposition 2

Take the decomposition of total risk from equation (A.3) from Lemma 3:

$$\mathcal{L}_t(M_{t,t+1}) = \mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) + \mathbb{E}_t[rx^{(\infty)}_{t+1}] - \theta^{H,H(1)}_t + \mathbb{E}_t[\theta^{H,H(\infty)}_{t,t+1}]$$

and its Foreign counterpart. Substituting this into Proposition 1 yields:

$$\begin{aligned}\mathbb{E}_t[rx^{FX}_{t+1}] =& \mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) + \mathbb{E}_t[rx^{(\infty)}_{t+1}] - \theta^{H,H(1)}_t + \mathbb{E}_t[\theta^{H,H(\infty)}_{t,t+1}] \\ & - \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1}) - \mathbb{E}_t[rx^{(\infty)*}_{t+1}] + \theta^{F,F(1)}_t - \mathbb{E}_t[\theta^{F,F(\infty)}_{t,t+1}] + \theta^{F,H(1)}_t - \theta^{F,F(1)}_t \\ \mathbb{E}_t[rx^{CT,(\infty)}_{t+1}] =& \mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) - \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1}) - \theta^{H,H(1)}_t + \mathbb{E}_t[\theta^{H,H(\infty)}_{t,t+1}] \\ & - \mathbb{E}_t[\theta^{F,F(\infty)}_{t,t+1}] + \theta^{F,H(1)}_t\end{aligned}$$

where line 2 rearranges, cancels like terms, and uses the definition of $\mathbb{E}_t[rx^{CT,(\infty)}_{t+1}]$.

Next, we know from equation (20), Lemma 4:

$$\mathbb{E}_t[\theta^{H,H(\infty)}_t] = \mathbb{E}_t[\theta^{F,H(\infty)}_t] - \sum_{\tau=0}^{\infty}\left\{\mathbb{E}_t[\theta^{F,H(1)}_{t+\tau}] + \mathbb{E}_t[\theta^{H,H(1)}_{t+\tau}]\right\}$$

for all t. Now, calculate $\mathbb{E}_t[\theta^{H,H(\infty)}_{t,t+1}] = \lim_{k\to\infty}\{\theta^{H,H(\kappa)}_t - \mathbb{E}_t[\theta^{H,H(\kappa-1)}_{t+1}]\}$:

$$\begin{aligned}\mathbb{E}_t[\theta^{H,H(\infty)}_{t,t+1}] = \mathbb{E}_t[\theta^{F,H(\infty)}_t] - \sum_{\tau=0}^{\infty}\left\{\mathbb{E}_t[\theta^{F,H(1)}_{t+\tau}] + \mathbb{E}_t[\theta^{H,H(1)}_{t+\tau}]\right\} \\ - \mathbb{E}_t\left[\mathbb{E}_{t+1}[\theta^{F,H(\infty)}_{t+1}] - \sum_{\tau=1}^{\infty}\left\{\mathbb{E}_{t+1}[\theta^{F,H(1)}_{t+\tau}] + \mathbb{E}_t[\theta^{H,H(1)}_{t+\tau}]\right\}\right] \\ = \mathbb{E}_t[\theta^{F,H(\infty)}_{t,t+1}] - (\mathbb{E}_t[\theta^{F,H(1)}_t] - \mathbb{E}_t[\theta^{H,H(1)}_t])\end{aligned}$$

Now, substitute $\mathbb{E}_t[\theta^{H,H(\infty)}_{t,t+1}]$ into the expression for $\mathbb{E}_t[rx^{CT,(\infty)}_{t+1}]$:

$$\begin{aligned}\mathbb{E}_t[rx^{CT,(\infty)}_{t+1}] =& \mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) - \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1}) - \theta^{H,H(1)}_t + \mathbb{E}_t[\theta^{H,H(\infty)}_{t,t+1}] - \mathbb{E}_t[\theta^{F,F(\infty)}_{t,t+1}] + \theta^{F,H(1)}_t \\ =& \mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) - \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1}) - \theta^{H,H(1)}_t + \mathbb{E}_t[\theta^{F,H(\infty)}_{t,t+1}] - (\theta^{F,H(1)}_t - \theta^{H,H(1)}_t) - \mathbb{E}_t[\theta^{F,H(\infty)}_{t+1}] \\ & - \mathbb{E}_t[\theta^{F,F(\infty)}_{t,t+1}] + \theta^{F,H(1)}_t \\ =& \mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) - \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1}) + \mathbb{E}_t[\theta^{F,H(\infty)}_{t,t+1}] - \mathbb{E}_t[\theta^{F,F(\infty)}_{t,t+1}]\end{aligned}$$

where line 1 restates the expression, line 2 uses expressions for $\mathbb{E}_t[\theta_t^{HH(\infty)}]$ and $\mathbb{E}_{t+1}[\theta_t^{HH(\infty)}]$ and line 3 rearranges to yield equation (22) in Proposition 4. $\square$

B Data and Empirics

B.1 Data Sources

CIP Deviations. We use CIP deviations at 3-month, 1-year and 10-year maturities from Du, Im, and Schreger (2018a) for 6 industrialized countries relative to the U.S.: Australia, Canada, euro area, Japan, Switzerland and U.K. The sample begins 1997:01 and ends 2021:03, although the panel is unbalanced as convenience yields are not available from the start of the sample in all jurisdictions. Owing to the availability of swap-spread data used to measure within-country convenience yields, as discussed below, our sample begins in 2000:01. We use the 3-month and 1-year convenience yields to construct 6-month convenience yields by linearly interpolating according to $CIP_t^{(6M)} = \frac{2}{3}CIP_t^{(3M)} + \frac{1}{3}CIP_t^{(1Y)}$. Table B.1 summarizes the start dates of the convenience yields. The data runs through 2021:03, which defines the end-point of our sample.

Table B.1: Du, Im, and Schreger (2018a) CIP Data Start Dates

Country	6-month Start Date	10-year Start Date
Australia	1997:10	1997:03
Canada	2001:02	2000:02
Euro Area	1999:01	1999:01
Japan	1997:06	1997:06
Switzerland	1998:09	1998:09
U.K.	1997:01	1997:01

Notes: The 6-month start date is the max{3-month, 1-year} start date.

Bond Yields. We use nominal zero-coupon government bond yields for maturities ranging from 6 months to 10 years (at 6-month intervals) for the G.7 economies. Table B.2 summarizes the sources of nominal zero-coupon government bond yields for the economies in our study. We also use these yields to construct G.7 bond premia using the approach of Adrian et al. (2013).

Table B.2: Yield Curve Data Sources

Country	Sources
U.S.	Gürkaynak, Sack, and Wright (2007)
Australia	Reserve Bank of Australia
Canada	Bank of Canada
Euro Area	Bundesbank (German Yields)
Japan	Wright (2011) and Bank of England
Switzerland	Swiss National Bank
U.K.	Anderson and Sleath (2001)

Exchange Rates. Exchange rate data is from *Datastream*, reflecting end-of-month spot rates *vis-à-vis* the U.S. dollar.

Swap Rates and Extensions. We use interest-rate swap rate data for two jurisdictions—the U.S. and the E.A.—and for two maturities—6 months and 10 years. The swap market only became sufficiently developed and liquid for the other G.7 currency areas around 2007. Thus, we do not consider swap rates for other G.7 currencies.

We begin by collecting the following data on interest-rate swap rates from *Datastream*. First, at the 6-month maturity, we collect overnight indexed swap (OIS) rates. The start date for these series is 2001:10 for the U.S. and 1999:10 for the euro area and they run until the end of our sample. These OIS swap rates are indexed to the effective Fed Funds rate and EONIA for the U.S. and euro area, respectively. Second, for the U.S. at the 10-year maturity, we start with interest-rate swaps rates that are indexed to LIBOR. This series becomes available in June 2003 and runs until the end of our sample. Third, for the euro area and at a 9-year maturity—which we use as a proxy for the 10-year due to data availability—we again use OIS rates, which become available in August 2005.

For the U.S. case, we then extend the 6-month and 10-year interest-rate swap rate series by back-filling them with 6-month and 10-year risk- and convenience-free AAA corporate bond rates, respectively, from the FRED database, à la Krishnamurthy and Vissing-Jorgensen (2012).[24] The correlations between the interest-rate swap and AAA corporate series are 89% and 82% for the 6-month and 10-year maturities, respectively. After accounting for a modest level-effect (the variances between the series are similar), our extended U.S. interest-rate swap series closely tracks that from **?**, who use data from JP Morgan Markets.[25] Overall, we use our extended U.S. interest-rate swap series beginning in 1997:01, to match the start date of the CIP deviations from Du et al. (2018a), although, in practice, data beginning only in 1999:10 is used due to availability of euro area data (see below).

For the euro area case, we extend the 9-year OIS rates by back-filling them with the 6-month OIS rates, since these display a correlation of 91%. Again, we account for a level effect (the variances are again comparable). In all, this allows us to extend our long-maturity series from a start date of 2005:08 to one of 1999:10.

[24]The codes are HQMCB6MT and HQMCB10YRP for the 6-month and 10-year maturities, respectively.
[25]Presumably, as there was no OIS market in 1996, JP Morgan Markets performs a back-fill similar to us.

Equities. We use two main sources related to equities. First, we obtain equity price indices of large and mid-cap sized firms for the U.S., Australia, Canada, Euro Area, Japan, Switzerland and the U.K. from *MSCI*. These equity returns are used to construct representative *ex post* measures of aggregate risk in each jurisdiction. Second, we collect data on dividend-price ratios and equity prices for each G.7 currency area from *Global Financial Data*. Specifically, we collect these data for the S&P-500 (U.S.), EuroStoxx-50 (E.A.), FTSE-100 (U.K.), TOPIX (Japan), S&P/ASX-200 (Australia), S&P/TSX-300 (Canada) and SMI (Switzerland).[26] We use these data to construct *ex ante* measures of equity risk premia as described in the main text. As with our other data, we use the end-of-month observations for all equity-related series.

Expectations. The construction of our equity risk premia measures requires a measure of future GDP growth and inflation expectations, which we proxy using next year's expectations using data from *Consensus Economics*.

VIX. To measure $\mathcal{L}(R^g_{t,t+1}/R_t) = \frac{T-t}{2} VIX_t^2$ (Martin (2017), result 3) for a 6-month holding period, we need the 6-month VIX. However, the 6-month VIX is available only from 2005:01 for the euro area (VSTOXX) and from 2008:02 for the U.S. (S&P500) and is unavailable for other G.7 jurisdictions. The standard 1-month VIX, which could be used as a proxy, is only available from the start of our sample for the U.S., with data for the majority of other G.7 jurisdictions available only after the global financial crisis. Therefore, we use the Gaussian approximation outlined in Martin (2017), result 4: $VIX_t^2 \approx \frac{1}{T-t}\text{var}_t(\log R^g_{t,t+1})$ (see Figure B.1 for details this construction), such that $\mathcal{L}(R^g_{t,t+1}/R_t) \approx \frac{1}{2}\text{var}_t(\log R^g_{t,t+1})$. For the U.S., the correlation between $\frac{1}{2}VIX_t^2$ and $\frac{1}{2}\text{var}_t(\log R^g_{t,t+1})$ is 58%, with comparable magnitudes.

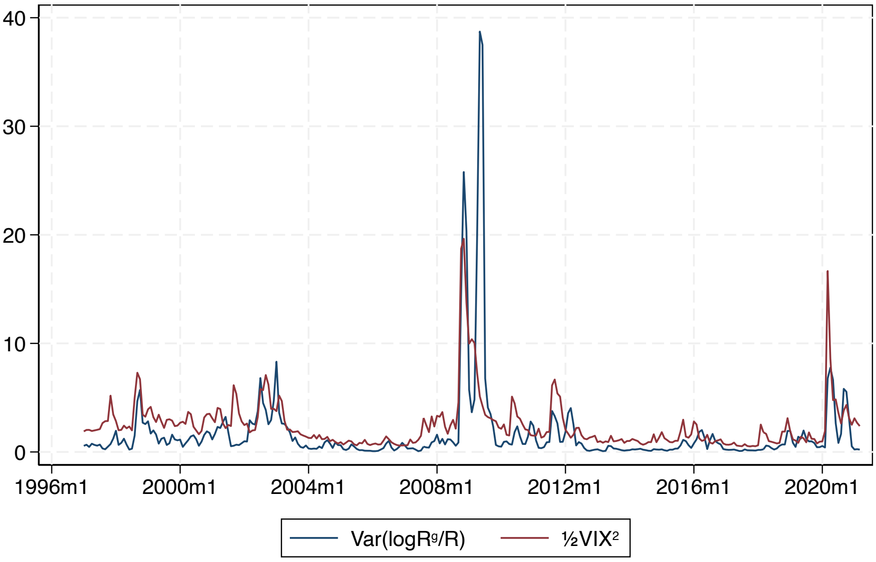
Government Debt-to-GDP. We source quarterly U.S. general government total and long-term nominal debt (from debt securities) as a percentage of GDP from the *World Bank Quarterly Public Sector Debt Database*. We linearly interpolate from quarterly to monthly frequency.

B.2 Further Details on Stylized facts from Section 2

Here, we discuss in detail the construction of our ex-ante equity risk premia measure and provide a comparison with the other measures in the literature (Appendix B.2.1 and B.2.2). We also offer cross-sectional (U.S. vs. each of the other G.7 economies) evidence on equity premia, dollar

[26] This data is not available for the greater set of large- and mid-cap public firms in *MSCI*.

Figure B.1: Comparison of $\frac{1}{2}VIX^2$ and $\frac{1}{2}var_t(\log R^g_{t,t+1})$



Notes: Time series of $\frac{1}{2}VIX^2$ and $\frac{1}{2}\text{var}_t(\log R^g_{t,t+1}/R_t)$ for the U.S. To construct $\text{var}_t(\log R^g_{t,t+1}/R_t)$, we use daily equity price data to construct levered 6-month equity excess returns $(\log R^g_{t,t+1}/R_t)$, and then take the variance within rolling 6-month windows to construct $\text{var}_t(\log R^g_{t,t+1}/R_t)$. Dividing by 2 (as shown in Martin (2017)) gives our proxy for $\mathcal{L}(R^g_{t,t+1}/R_t)$.

returns and Treasury convenience (Appendix B.2.3).

B.2.1 Estimating Expected Dividend Growth

In our baseline, we proxy for expected-future dividend growth using a weighted average of past dividend growth and next-year GDP growth expectations from *Consensus Economics*. Specifically, we regress 6-month (one-period) future dividend growth g_t on the previous year's dividend growth g_{t-1} and next year's GDP growth expectations $g_t^{GDP,e}$ from *Consensus Economics* country-by-country and use the predicted value as our proxy for g_t^e:[27]

$$g_t^e := \hat{g}_t = \hat{\alpha} + \hat{\beta} g_{t-1} + \hat{\beta}^{GDP} g_t^{GDP,e}. \tag{B.1}$$

The results from estimating regression (B.1) for the U.S. are presented in Table B.3. The results highlight the strong fit of the expectations regression, with an R^2 of 37%, and with both lagged dividend growth and GDP growth expectations being highly significant. Combining past dividend growth with forward-looking survey data on GDP allows our measure to capture both

[27] Atkeson et al. (2025b) argue that public firms' valuations since the 1970s may reflect an increasing cash flow in the form of buybacks—implying that the total payout to price ratio may be significantly higher and more volatile. Due to lack of availability of data outside the US, we abstract from this.

trends in risk and cyclical dynamics. We estimate analogous regressions for each country in our sample.

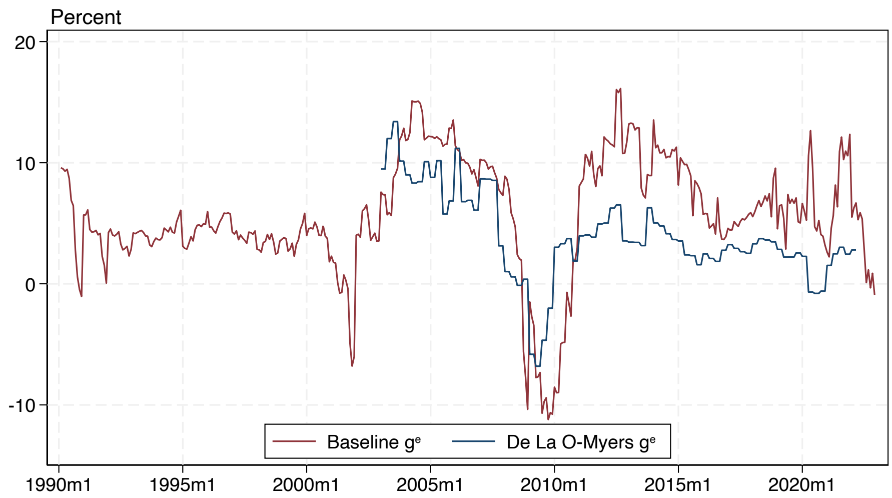
Table B.3: U.S. (S&P 500) Future Dividend Growth Regressions

Dep. Var.: g_t	
g_{t-1}	0.61***
	(0.12)
$g_t^{GDP,e}$	3.45***
	(1.01)
Adjusted R^2	0.37
N	394

Notes: Coefficient estimates from estimating regression (B.1) for the U.S. Newey and West (1987) standard errors with 4 lags reported, with *** denoting $p < 0.01$, ** $p < 0.05$, and * $p < 0.1$.

Encouragingly, Figure B.2 shows that our measure of expected U.S. dividend growth constructed from B.1 co-moves strongly with CFO survey data from De La'O and Myers (2021) on dividend growth expectations over the same horizon.

Figure B.2: Comparison of Baseline g^e Measure with De La'O and Myers (2021) Survey Data



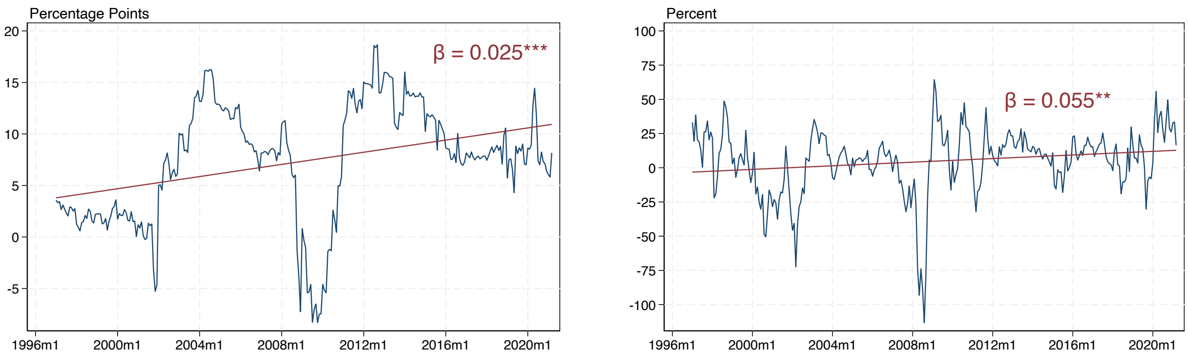
Notes: Figure B.2 presents time series of our baseline U.S. expected 6-month-ahead dividend growth (g^e) measure with analogous measure based on CFO survey data from De La'O and Myers (2021). The two series have a correlation of 0.63 beginning in January 2003, when the latter series becomes available.

B.2.2 Decomposing Expected Equity Risk Premia

Figure B.3 contrasts expected U.S. equity premia obtained using equation (2) and measuring expected dividend growth using (B.1) (Panel a), with realized equity premia (Panel b). For illustration, the slope of the trend line in *ex ante* premia is roughly 45% that of *ex post* premia lending support to the view that a significant share of increased U.S. equity premia reflects compensation for risk.

Figure B.3: U.S. Expected and Realized Equity Excess Returns

(a) U.S. Expected Equity Excess Returns (b) U.S. Realized Equity Excess Returns



Note. Panels B.3a and B.3b display time series of our baseline measures of U.S. expected and realized 6-month equity excess returns, respectively. The expected equity excess return is estimated with the Gordon growth formula in equation (2), with expected dividend growth calculated according to equation (B.1). *** (**) signifies that the slope (β) of the estimated deterministic trend line is different from zero at the 1% (5%) significance level based on Newey and West (1987) standard errors with 4 lags. The sample runs from 1997:01 to 2021:03 to match data on CIP deviations.

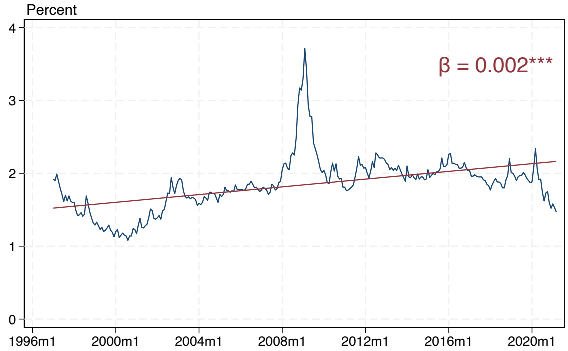
Figure B.4 offers an empirical decomposition of our estimate for *ex ante* U.S. equity premia in the four objects on the right-hand-side of (2). The trend in U.S. equity premia is predominantly driven by positive expected dividend growth (Panel b), combined with falling short-term rates (Panel c). The dividend-price ratio (panel a) also, but at a slow pace, over the sample.

Figure B.5 shows the result of the same exercise, this time looking at the equity premia in relative (U.S. versus the rest of the G.7) terms (as shown in Figure 1(a)). There are a few important differences. Starting from Panel (c), there is no trend in short rate differentials, although expected short-term inflation declines slightly more in the U.S. Most of the action is instead driven by expected dividend growth (panel B) which is clearly stronger in the U.S. relative to the rest of the G.7. Indeed, this trend seems to be specific to the U.S, with the rest of the G.7 expected dividend growth instead falling over our sample. Additionally, the dividend-price ratio is falling in relative terms.

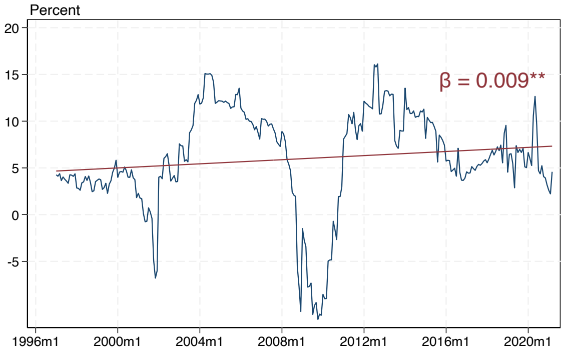
Finally, Figure B.6 contrasts our expected risk premium measure with data from the Livingston Survey (only available bi-annually) and, interestingly, there is substantial co-movement both in the short-run and the trend.

Figure B.4: Decomposing U.S. Equity Premia: $\log \mathbb{E}_t[\frac{R^{eq}_{t,t+1}}{R_t}] \approx \frac{D_t}{P_t} + g^e_t - (r_t - \pi^e_t)$

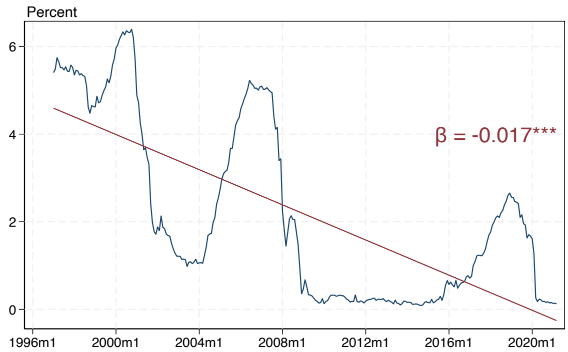
(a) U.S. Dividend/Price Ratio (D/P)



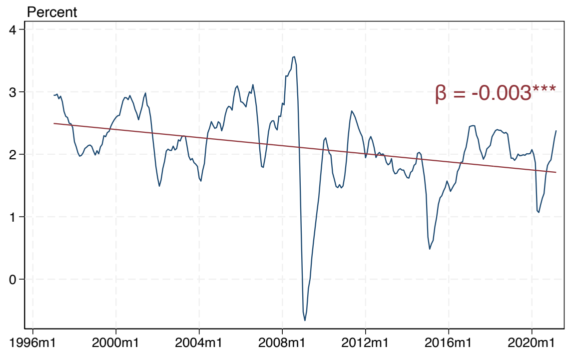
(b) U.S. Expected Dividend Growth (g^e)



(c) U.S. short-term rate (r)



(d) U.S. expected short-term inflation (π^e)



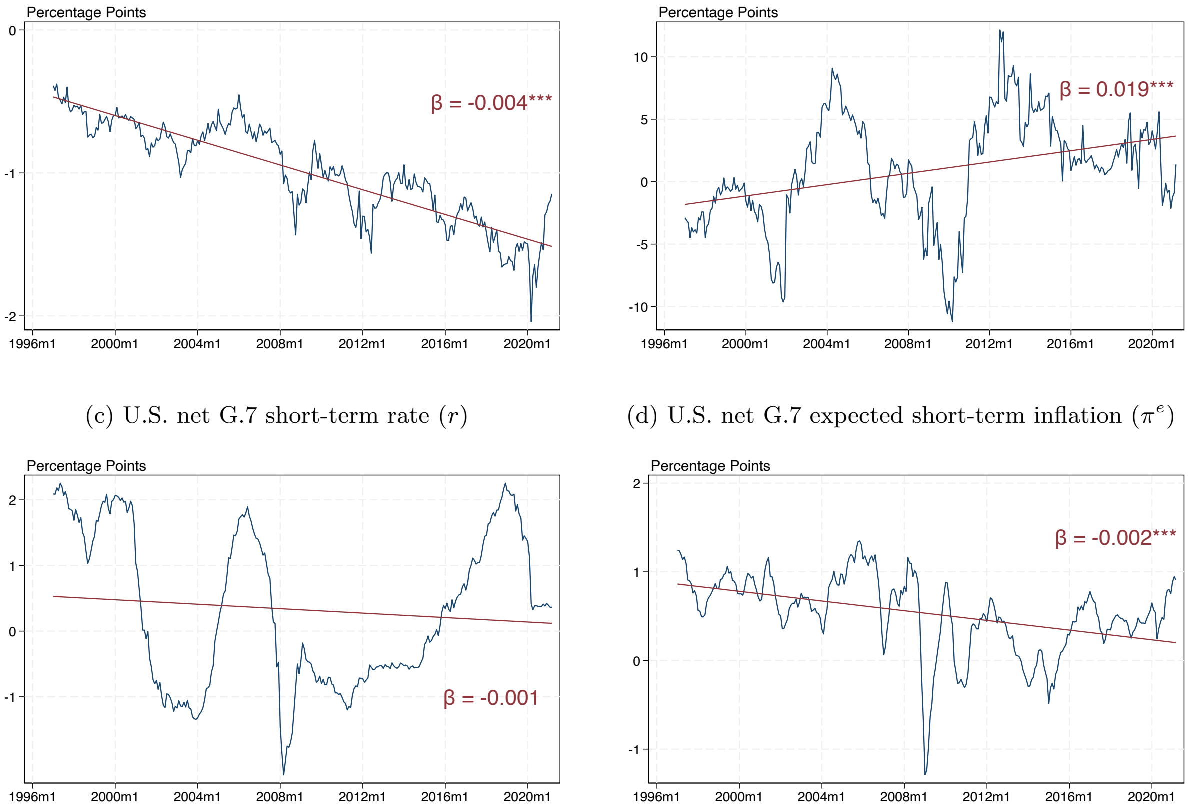
Note. Panels of Figure B.4 provide a decomposition of our baseline measure of the U.S. log equity risk premium (shown in Figure B.3a) into the sub-components of the Gordon growth formula in equation (2): U.S. dividend-price ratio (Panel B.4a), U.S. expected dividend growth (Panel B.4b), U.S. short-maturity (6-month) interest rate (Panel B.4c) and U.S. next-year inflation expectations (Panel B.4d). *** (**) signifies that the slopes (βs) of the estimated deterministic trend lines are different from zero at the 1% (5%) significance level based on Newey and West (1987) standard errors with 4 lags. The sample runs from 1997:01 to 2021:03.

B.2.3 Equity Premia, Dollar Returns and Treasury Convenience in Cross-Section

Figures B.7-B.9 offer a cross-sectional analysis of our stylized facts, comparing bilaterally the U.S. with each of the other G.7 countries, pre- and post-2010. Panel (a) in B.7 shows that, post-2010, expected equity premia in the U.S. are greater than equity premia in 5 of the other 6 G.7 countries, whereas pre-2010, this was the case for only 1 of 6. A similar picture plays out for realized equity premia in Panel (b).

Figure B.8 plots carry-trade returns. The two panels show that these returns tend to be negative in the post-2010 period. However, for 10-year maturities the difference with the pre-2010 period is statistically insignificant for all economies, except Canada. At 6-month

Figure B.5: Decomposing U.S. net G.7 Equity Premia: $\log \mathbb{E}_t[\frac{R^{eq}_{t,t+1}}{R_t}] \approx \frac{D_t}{P_t} + g^e_t - (r_t - \pi^e_t)$



(a) U.S. net G.7 Dividend/Price Ratio (D/P)

(b) U.S. net G.7 Expected Dividend Growth (g^e)

(c) U.S. net G.7 short-term rate (r)

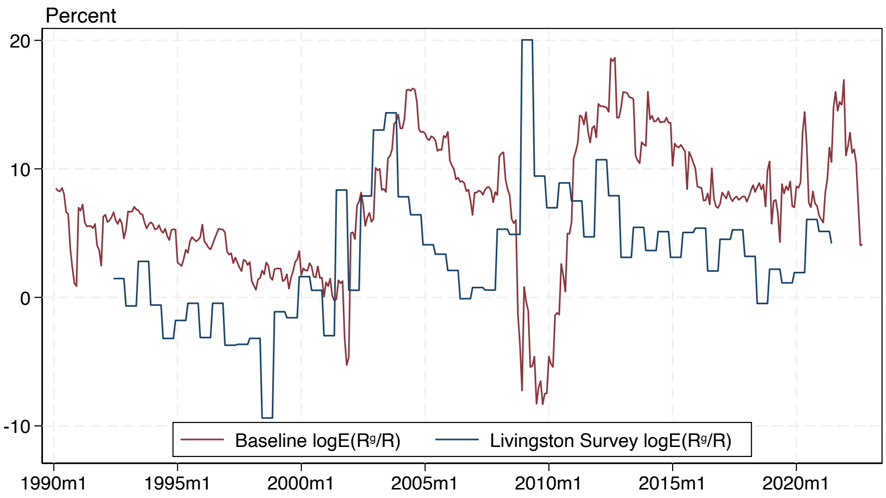
(d) U.S. net G.7 expected short-term inflation (π^e)

Note. Decomposition of our baseline measure of the U.S. net average G.7 log equity risk premium (shown in Figure 1a) into the sub-components of the Gordon growth formula in equation (2): U.S. net average G.7 dividend-price ratio (Panel B.5a), U.S. net average G.7 expected dividend growth (Panel B.5b), U.S. net average G.7 short-maturity (6-month) interest rate (Panel B.5c) and U.S. net average G.7 next-year inflation expectations (Panel B.5d). *** signifies that the slopes (βs) of the estimated deterministic trend lines are different from zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The sample runs from 1997:01 to 2021:03.

maturities, the difference is significant also for Australia and the euro area. This suggests that, at short maturities, some change in dollar risk is detectable, bilaterally, in the currency market. As shown in the text, nonetheless, there is no systematic evidence of it when averaging across the countries in the sample.

Figure B.9 plots CIP deviations. Panel (a) suggests a strong and significant deterioration of U.S. convenience at long maturities relative to Australia, Canada the euro area and more marginal, the UK and Switzerland. No such pattern emerges at short maturities. Even when differences between periods are significant, they are quite small.

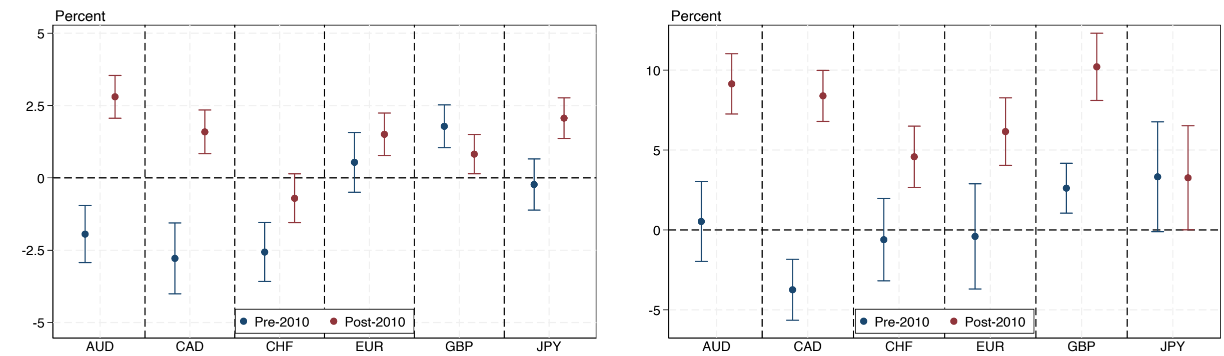
Figure B.6: Comparison of Baseline ERP Measure with Livingston Survey



Notes: Figure B.6 presents time series of our baseline U.S. log equity risk premium ($\log \mathbb{E}_t[\frac{R^{eq}_{t,t+1}}{R_t}]$) measure, as shown in Figure B.3(a), with a measure constructed from the Livingston Survey, available only bi-annually.

Figure B.7: U.S. vs. G.7 Expected and Realized equity premium Pre- and Post-2010

(a) U.S. net G.7 Expected Equity Premium (b) U.S. net G.7 Realized Equity Premium



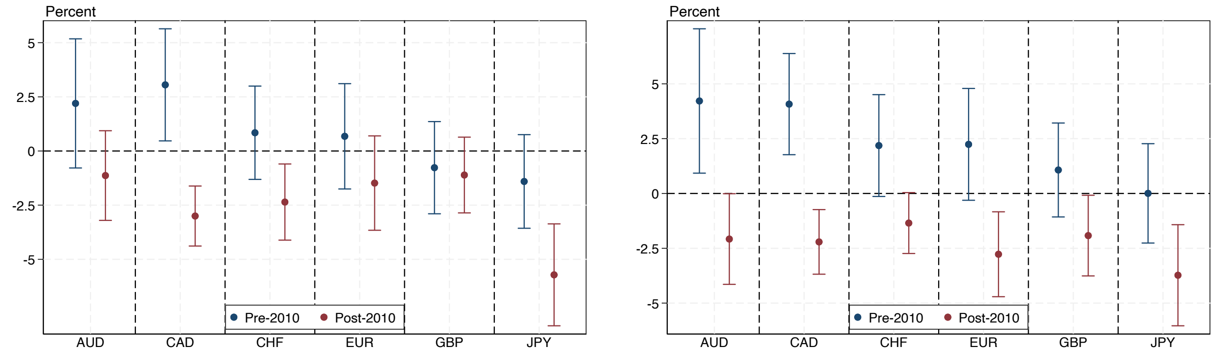
Note. The bars in Panels B.7a and B.7b reflect average 6-month expected and realized equity excess returns, respectively, for the U.S. relative to other G.7 markets, both pre- and post-2010. The post-2010 period is from 2011:01 to 2021:03 (the end-date is determined by the availability of CIP data from Du et al., 2018a). The start-date for the pre-2010 period is market-specific (ranging from 1997:01 to 2000:02), once again dictated by the availability of CIP data. Error bars are 68% confidence intervals constructed using Newey and West (1987) standard errors with 4 lags.

B.3 Estimating a Mapping from CIP to Cross-Country Convenience Yields

We estimate the coefficient β^*_k in $\theta^{F,H(k)}_t - \theta^{F,F(k)}_t = \frac{1}{1-\beta^*_k} CIP^{(k)}_t$ following the approach of Jiang et al. (2021a). We focus on the 6-month CIP deviation, our measure of short-maturity convenience, as well as the 1-year CIP deviation to match the maturity used by Jiang et al. (2021a). As we discuss below, we cannot extend the same methodology to 10-year CIP deviations since

Figure B.8: U.S. vs. G.7 long- and short-maturity carry trade returns Pre- and Post-2010

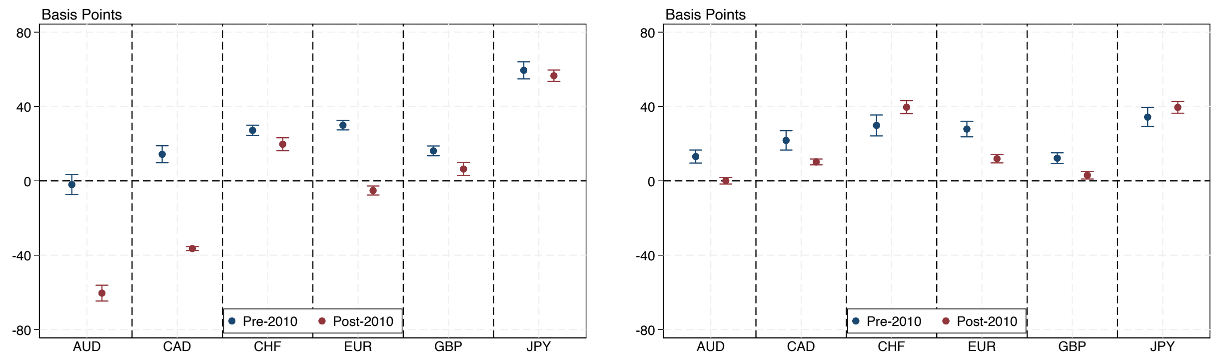
(a) U.S. 10-Year Carry Trade Returns (b) U.S. 6-Month Carry Trade Returns



Note. The bars in Panels B.8a and B.8b reflect average 6-month carry trade returns from 10-year and 6-month government bonds, respectively, for the U.S. relative to other G.7 markets, both pre- and post-2010. The post-2010 period is from 2011:01 to 2021:03 (the end-date is determined by the availability of CIP data from Du et al., 2018a). The start-date for the pre-2010 period is market-specific (ranging from 1997:01 to 2000:02), once again dictated by the availability of CIP data. Error bars are 68% confidence intervals constructed using Newey and West (1987) standard errors with 4 lags.

Figure B.9: U.S. vs. G.7 long- and short-maturity CIP deviations Pre- and Post-2010

(a) U.S. 10-Year CIP Deviations (b) U.S. 6-Month CIP Deviations



Note. The bars in Panels B.9a and B.9b reflect average CIP deviations on 10-year and 6-month government bonds, respectively, for the U.S. relative to other G.7 markets, both pre- and post-2010. The post-2010 period is from 2011:01 to 2021:03 (the end-date is determined by the availability of CIP data from Du et al., 2018a). The start-date for the pre-2010 period is market-specific (ranging from 1997:01 to 2000:02), once again dictated by the availability of CIP data. Error bars are 68% confidence intervals constructed using Newey and West (1987) standard errors with 4 lags.

they violate the required assumption of stationarity and independence of CIP from exchange rate premia.

The coefficient of interest, β_k^*, can be estimated from the data using:

$$\beta_k^* = 1 - \frac{1}{1-[\alpha_1^{(k)}]^4}\frac{1}{\delta_1^{(k)}} \tag{B.2}$$

where α_1 is the persistence of an AR(1) process for the average U.S. CIP deviation across currencies and δ_1 is the marginal effect of innovations to CIP deviations on exchange rate movements. The intuition for this formula comes from solving Proposition 1 forward for the exchange rate level:

$$e_t = \sum_{h=0}^{\infty} \theta_{t+h}^{F,H(k)} - \theta_{t+h}^{F,F(k)} + \sum_{h=0}^{\infty} r_{t+h}^{(k)*} - r_{t+h}^{(k)} + \sum_{h=0}^{\infty} \mathcal{L}_t(M_{t+h,t+h+k}) - \mathcal{L}_t(M^*_{t+h,t+h+k}) + \bar{e}, \tag{B.3}$$

where $\theta_{t+h}^{F,H(k)} - \theta_{t+h}^{F,F(k)} = CIP_{t+h}^{(k)}/(1-\beta_k^*)$. That is, the future path of convenience yields must move 1 for 1 with exchange rates, which equation (B.2) ensures. Importantly, since $\theta_{t+h}^{F,H(10Y)} - \theta_{t+h}^{F,F(10Y)}$ is non-stationary, its infinite sum is not well defined, which precludes extending the mapping from CIP to convenience yields at long maturities. As discussed in the main text, however, this is not an issue for our error correction model as long as as the scale factor is linear and constant.

Nonetheless, we show below that there are important differences along the term structure of CIP deviations. The coefficient α_1 is estimated from a quarterly AR(1) process for the average U.S. CIP deviation across the six remaining G.7 currencies $\overline{CIP}_t^{(k)}$, for a maturity k:

$$\overline{CIP}_{t+3}^{(k)} = \alpha_1 \overline{CIP}_t^{(k)} + \nu_t. \tag{B.4}$$

Table B.4 reports the results.

Table B.4: Autocorrelation in CIP Deviations

Vars	$\overline{CIP}_{t+3}^{(6M)}$	$\overline{CIP}_{t+3}^{(1Y)}$
$\overline{CIP}_t^{(6M)}$	0.68***	
	(0.05)	
$\overline{CIP}_t^{(1Y)}$		0.80***
		(0.04)
Observations	230	230
Within R^2	0.46	0.64

The estimation of δ_1 is in two steps. First, we construct quarterly innovations to the CIP deviation, $\Delta\widetilde{\overline{CIP}}_t^{(k)}$ as the residual from estimating:

$$\Delta\overline{CIP}_{t+3}^{(k)} = \gamma_0 + \gamma_1 \overline{CIP}_t^{(k)} + \gamma_2 (\overline{r_t^{*,(k)} - r_t^{(k)}}) + \omega_t \tag{B.5}$$

The results from this regression are reported in Table B.5.

In the second step, we estimate δ_1 by regressing quarterly exchange rate movements on

Table B.5: Constructing CIP Innovations as Residuals to

Vars	$\Delta\overline{CIP}^{(6M)}_{t+3}$	$\Delta\overline{CIP}^{(1Y)}_{t+3}$
$\overline{CIP}^{(6M)}_t$	-0.59***	
	(0.06)	
$\overline{r^{*,(6M)}_t - r^{(6M)}_t}$	227	
	(138)	
$\overline{CIP}^{(1Y)}_t$		-0.44***
		(0.05)
$\overline{r^{*,(1Y)}_t - r^{(1Y)}_t}$		170
		(83)
Observations	230	230
R^2	0.30	0.23

these innovations, $\Delta\widetilde{\overline{CIP}}^{(k)}_t \equiv \omega_t$ where ω_t is the residual from regression (B.5):

$$\Delta\bar{e}_{t+3} = \delta_1 \Delta\widetilde{\overline{CIP}}^{(k)}_t + \varepsilon_t \tag{B.6}$$

Table B.6 reports the results.

Table B.6: CIP Innovations and Exchange Rate Dynamics

Vars	$\Delta\bar{e}_{t+3}$	$\Delta\bar{e}_{t+3}$
$\Delta\widetilde{\overline{CIP}}^{(6M)}_t$	-5.71***	
	(1.15)	
$\Delta\widetilde{\overline{CIP}}^{(1Y)}_t$		-13.8***
		(1.85)
Observations	230	230
Within R^2	0.10	0.19

Using these estimates α_1 and δ_1, we find $\beta^{(1Y)} = 0.88$, which is similar to the value found by Jiang et al. (2021a) of $\beta^{(1Y)} = 0.9$. The value found for the 6-month maturity is $\beta^{(6M)} = 0.77$, which suggests that longer-maturity Treasuries carry lower convenience above their currency denomination than longer-maturity ones, consistent with our findings that long bonds are more sensitive to risk differentials.

B.4 Within-Country Convenience Yields.

We measure within-country convenience yields according to:

$$\theta_t^{H,H(k)} := r_{irs,t}^{(k)} - r_t^{(k)} \tag{B.7}$$

$$\theta_t^{F,F(k)} := r_{irs,t}^{*(k)} - r_t^{*(k)} \tag{B.8}$$

where $r_{irs,t}^{(k)}$ and $r_{irs,t}^{*(k)}$ denote the Home and Foreign log k-period interest-rate-swap rate, respectively. Note that, due to data limitations, we can only construct Foreign within-country convenience yields for the euro-area.

Figure B.10 plots U.S. within-country convenience yield for the 6-month (panel (a)) and 10-year (panel (b)) tenors. Short maturity yields were rising in the build-up the GFC, but have since averaged just below zero. In contrast, for long maturities, there is a clear downward trend post 2000, albeit less pronounced than its cross-country counterpart. Therefore, the fading 'specialness' of U.S. treasuries to Foreign investors is, to an extent, matched by domestic investors in the U.S. Panels (c) and (d) plot within-country convenience yields for the euro area. For both tenors, convenience yields display the opposite dynamics to the U.S. and have been rising over time, with the trend is again more pronounced for the 10-year tenor.

B.5 Growth Optimal Portfolio

In our baseline, we rely on the frequently used approximation for portfolio returns:

$$R_{t+1}^{g} = \phi(1 + r_{t+1}^{eq}) + (1 - \phi)(1 + r_t) \tag{B.9}$$

The excess log return can be expressed as:

$$r_{t+1}^{g} - r_t = \log(1 + \phi(e^{r_{t+1}^{eq} - r_t} - 1) \tag{B.10}$$

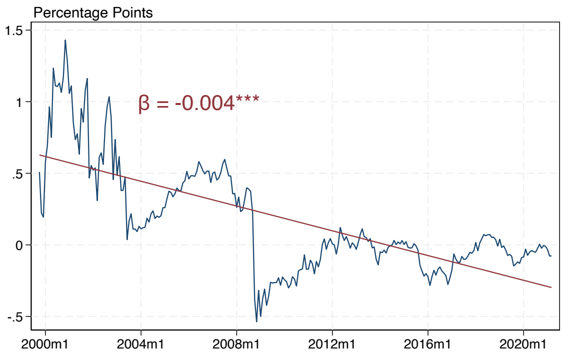
To solve for the *maximizing* portfolio , consider a first order approximation of the form $f(x) = \log(1+\phi(e^x - 1)$ around $x = 0$ yields $\phi x + \frac{1}{2}(\phi - \phi^2)x^2 + \overline{o}(x)^2$. Applying this to excess portfolio returns:

$$r_{t+1}^{g} - r_t = \phi(r_{t+1}^{g} - r_t) + \frac{1}{2}(\phi - \phi^2)(r_{t+1}^{g} - r_t)^2 + \overline{o}(r_{t+1}^{g} - r_t)^2 \tag{B.11}$$

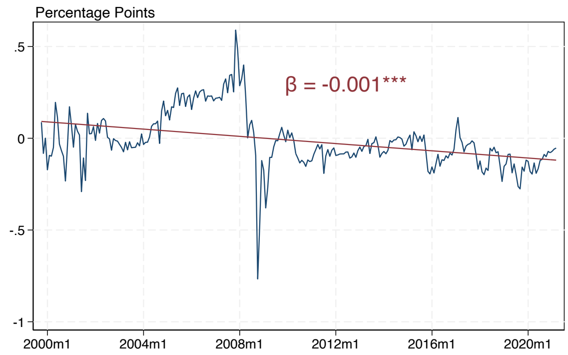
and we approximate assuming $\overline{o}(r_{t+1}^{g} - r_t)^2 \to 0$.

Figure B.10: Long- and Short-Maturity Within-Country U.S. and E.A. Convenience Yields

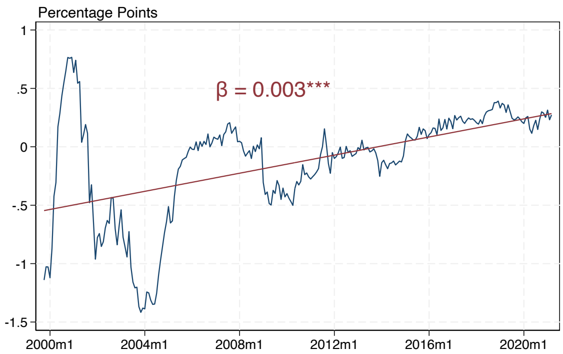
(a) U.S. 10Y Within-Country CY ($\theta^{H,H(10Y)}$)



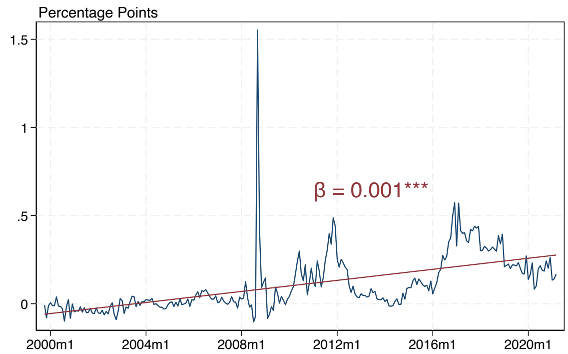
(b) U.S. 6M Within-Country CY ($\theta^{H,H(6M)}$)



(c) E.A. 10Y Within-Country CY ($\theta^{F,F(10Y)}$)



(d) E.A. 6M Within-Country CY ($\theta^{F,F(6M)}$)



Note. Panels B.10a, B.10b, B.10c and B.10d display time series of long-maturity and short-maturity within-country U.S. Treasury and German Bund convenience yields, respectively. The convenience yields are calculated as the difference between interest-rate swap rates and the corresponding-maturity zero-coupon government bond yield, as described in equation (B.7) and (B.8). *** signifies that the slopes (βs) of the estimated deterministic trend lines are different from zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The sample runs from 1999:10 to 2021:03 based on the availability of data used to estimate within-country convenience yields and CIP deviations.

Under the approximation of log-normal returns, assuming CRRA utility with coefficient γ, it can be shown that the investor maximizes:

$$\max \mathbb{E}_t[r^g_{t+1}] + \frac{1}{2}(1-\gamma)var_t(r^g_{t+1})$$

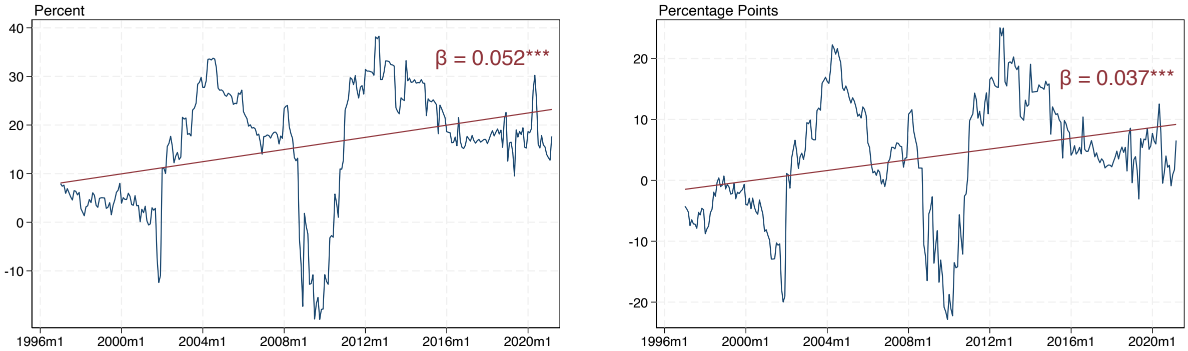
Plugging (B.11) into the maximization, solving for ϕ yields:

$$\phi = \frac{\mathbb{E}_t[r^{eq}_{t+1}] - r_t + \frac{1}{2}\sigma_t(r^{eq}_{t+1})^2}{\gamma\sigma_t(r^{eq}_{t+1})^2} \tag{B.12}$$

The growth optimal portfolio which maximizes the log excess return then also coincides with the maximization of an investor with log-utility. Figure B.11 plots expected returns to the growth optimal portfolio which inherits most of the dynamics of expected equity premia (see Figure B.3).

Figure B.11: U.S. Growth-Optimal Portfolio Excess Return, Absolute and Relative to G.7 Average

(a) U.S. Growth-Optimal Excess Return (b) U.S. net G.7 Growth-Optimal Excess Return



Note. Panels B.11a and B.11b display time series of our baseline measures of U.S. growth-optimal portfolio excess return, in absolute terms and net G.7 average, respectively. The growth-optimal portfolio is constructed by leveraging up the equity holdings by taking short positions in the risk-free rate, according to (B.9) and (B.10). *** signifies that the slope (β) of the estimated deterministic trend line is different from zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The sample runs from 1997:01 to 2021:03.

B.6 Further Details for Risk Measures

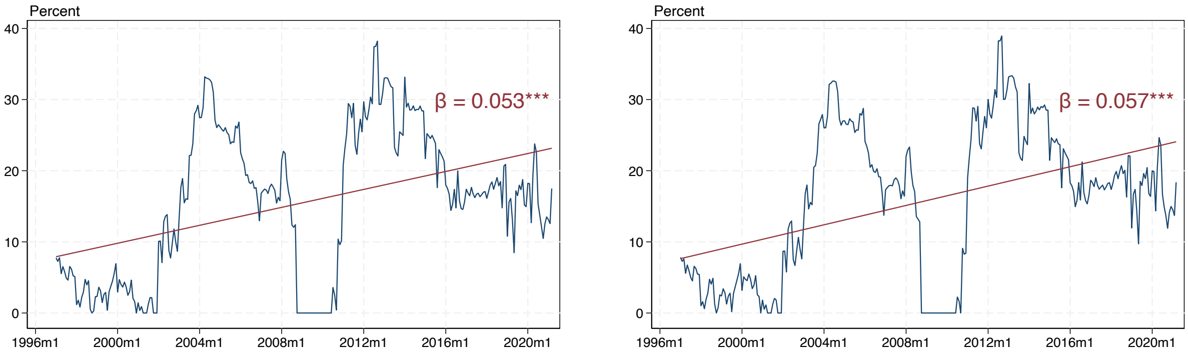
Measures for U.S. SDF Volatility. Figure B.12 presents our measures of U.S. total and permanent risk. As with the relative risk measures discussed in the main text, U.S. permanent risk explains the entire rise in U.S. total risk over our sample. All risk measures are truncated at zero. Figure B.13 shows our proxy for transitory risk, constructed as the difference between total risk and permanent risk.[28]

Decomposing U.S. Absolute and Relative Permanent Risk. Figures B.14 and B.15 decompose our measures of U.S. absolute and relative permanent risk into their constituents. In both cases, the rise in U.S. (relative) permanent risk is driven mostly by the rise in the U.S. growth-optimal (levered equity) risk premium. While the U.S. real bond premium declines over the sample (Figures B.14(c)), reinforcing the rise in U.S. levered equity premia for U.S. absolute permament risk, relative U.S. bond premia instead increase (B.15(c)), partially offsetting the increase in relative levered equity premia for U.S. relative permanent risk. In both cases,

[28]Note that this residual is not itself a volatility, since permanent and transitory components are correlated. As such, it is not truncated at zero. The residual coincides with the log inverse transitory SDF (Lemma 2) net of the convenience adjusted risk-free rate.

Figure B.12: U.S. Absolute Total and Permanent Risk

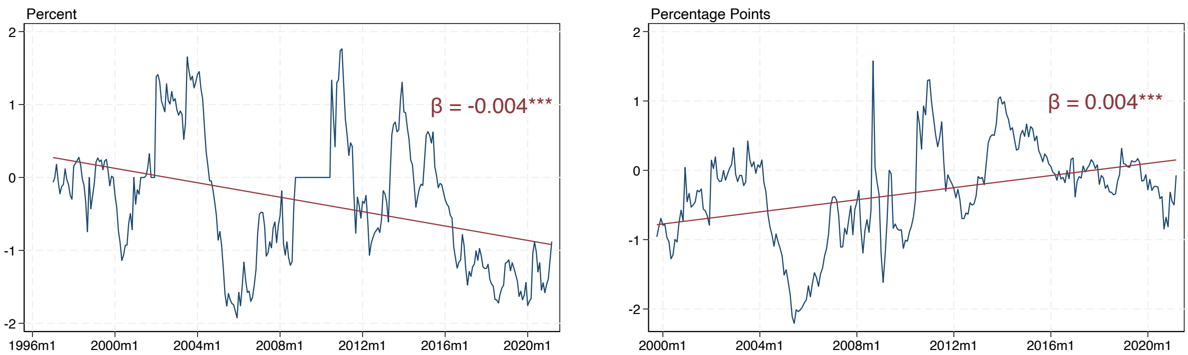
(a) U.S. Total Risk (b) U.S. Permanent Risk



Note. Panels B.12a and B.12b display time series of our baseline proxies for U.S. total and permanent risk, respectively. Total risk is calculated according to equation (23); Permanent risk is calculated according to equation (24). *** (**) signifies that the slope (β) of the estimated deterministic trend line is different from zero at the 1% (5%) significance level based on Newey and West (1987) standard errors with 4 lags. The sample runs from 1997:01 to 2021:03.

Figure B.13: Proxies for Absolute and Relative Transitory Risk

(a) U.S. Transitory Risk (b) U.S. net G.7 Relative Transitory Risk

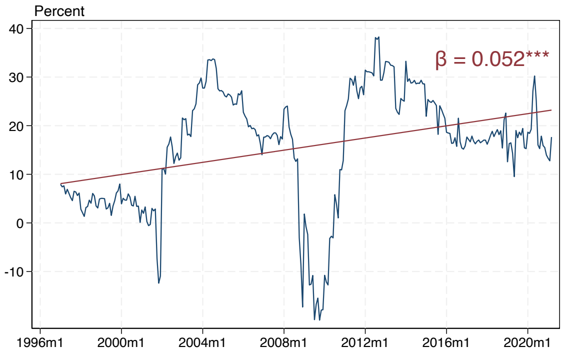


Note. Panels B.13a and B.13b display time series of our proxies for U.S. absolute and net average G.7 transitory risk, respectively. Transitory risk is calculated as the residual of total risk (equation (23)) and permanent risk (equation (24)). ***, **, * signify that the slope (β) from a regression of the variable on a time trend is different from zero at the 1%, 5%, 10% significance levels based on Newey and West (1987) standard errors with 4 lags, respectively. The sample runs from 1999:10 to 2021:03.

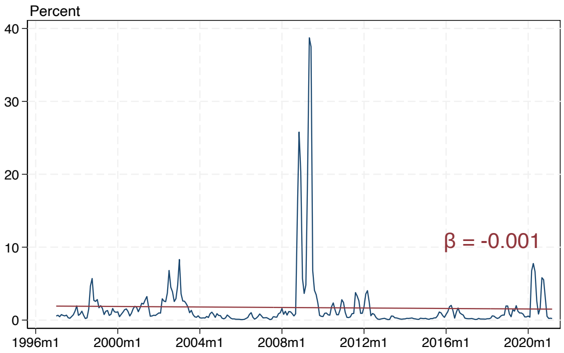
while U.S. (relative) within-country convenience yields decline (Panel c), which increases U.S. (relative) permanent risk, the magnitude of this decline is small. In addition, the expected volatility of equity returns (Panel d), both in absolute and relative terms, are not trending over our sample.

Figure B.14: Decomposing U.S. Permanent Risk $PermRisk_t$

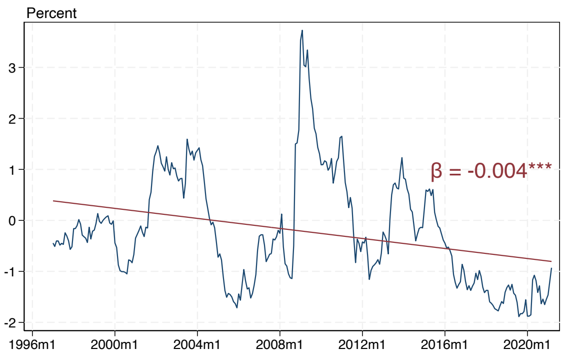
(a) U.S. Growth Optimal Premium ($\log \mathbb{E}[R^g]/R$)



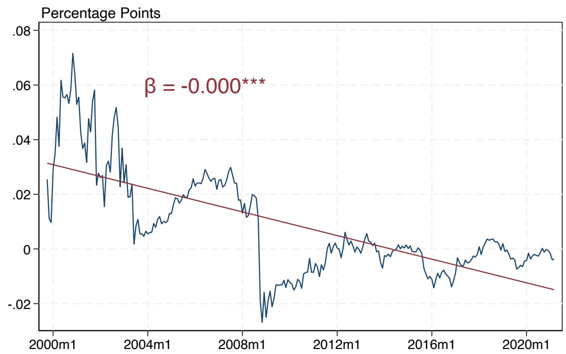
(b) U.S. R^g/R Volatility Adjustment ($\mathcal{L}(R^g/R)$)



(c) U.S. real bond premium ($\mathbb{E}(rx^{(10Y)})$)



(d) U.S. Within-Country CY HPR ($\omega^{(10Y)}\theta^{H,H(10Y)}$)

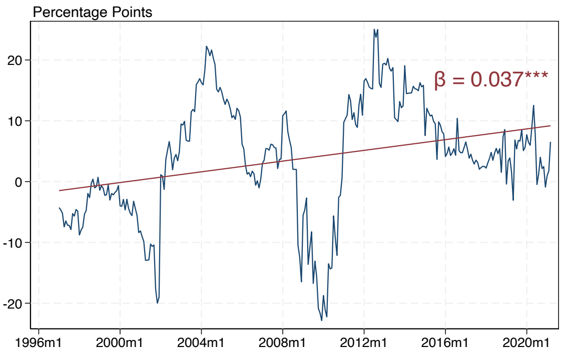


Note. Panels of Figure B.14 provide a decomposition of our baseline measure of the U.S. permanent risk (shown in Figure B.12) into its sub-components, as given by equation (24): U.S. log growth optimal risk premium (Panel B.14a), U.S. risk-premium volatility adjustment (Panel B.14b), U.S. expected real bond premium (Panel B.14c) and U.S. long-maturity within-country convenience yield holding-period return (Panel B.14d). *** (**) signifies that the slopes (β) of the estimated deterministic trend lines are different from zero at the 1% (5%) significance level based on Newey and West (1987) standard errors with 4 lags. The sample runs from 1997:01 to 2021:03.

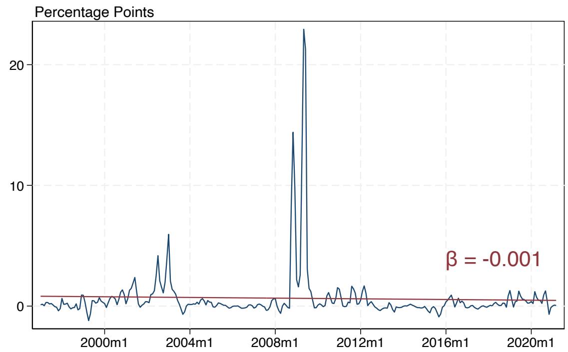
Relative Risk in the Cross-Section. Figure B.16 presents a cross-sectional analysis of our relative risk measures, comparing pre- and post-2010 averages of the U.S. with each of the other G.7 countries. Panel (a) details total risk and Panel (b) details permanent risk. In both cases, U.S. risk rises relative to all other G.7 countries from pre-2010 to post-2010, except for the U.K. (GBP). This highlights that our central result from the main text holds across currency areas.

Figure B.15: Decomposing U.S. net G.7 Permanent Risk $\mathcal{D}PermRisk_t$

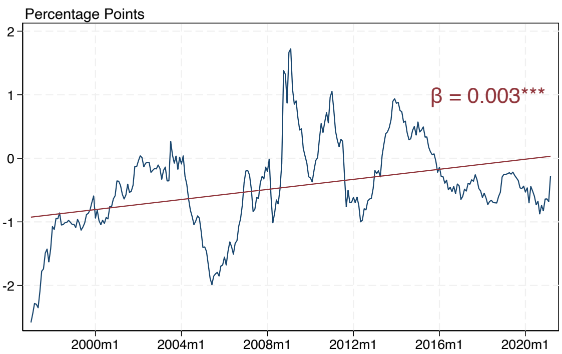
(a) U.S. net G.7 Growth Optimal Prem. $(\log \mathbb{E}[R^g]/R)$



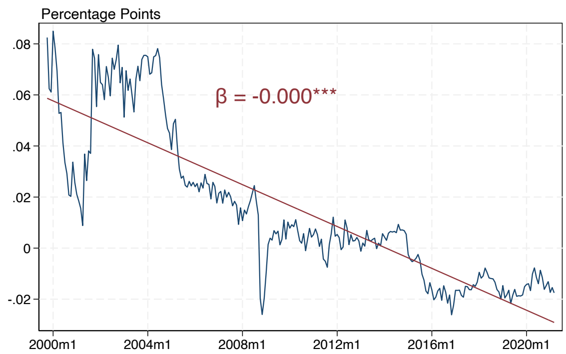
(b) U.S. net G.7 R^g/R Volatility Adjustment $(\mathcal{L}(R^g/R))$



(c) U.S. net G.7 real bond premium $(\mathbb{E}(rx^{(10Y)}))$

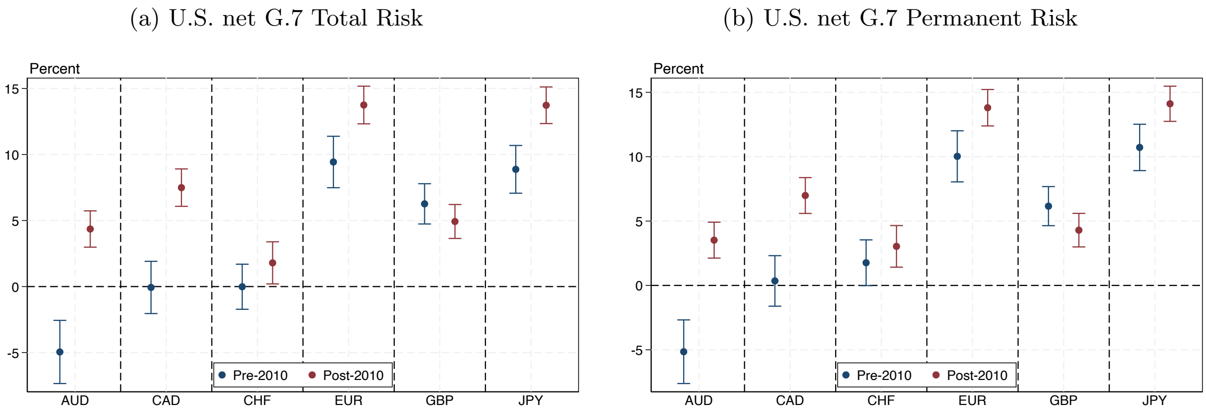


(d) U.S. net G.7 within-country CY HPR $(\omega^{(10Y)}\theta^{H,H(10Y)})$



Note. Panels of Figure B.15 provide a decomposition of our baseline measure of the U.S. net average G.7 permanent risk (shown in Figure 4b) into its sub-components, as given by equation (24): U.S. net average G.7 log growth-optimal risk premium (Panel B.15a), U.S. net average G.7 risk-premium volatility adjustment (Panel B.15b), U.S. net average G.7 real bond premium (Panel B.15c) and U.S. net average G.7 long-maturity within-country convenience yield holding-period return (Panel B.15d). *** (**) signifies that the slopes (βs) of the estimated deterministic trend lines are different from zero at the 1% (5%) significance level based on Newey and West (1987) standard errors with 4 lags. The sample runs from 1999:10 to 2021:03 based on the availability of data used to estimate within-country convenience yields and CIP deviations.

Figure B.16: U.S. vs. G.7 Relative Total and Permanent Risk Pre- and Post-2010



Note. Panels B.16a and B.16b reflect our baseline proxies for average total and permanent risk, respectively, for the U.S. relative to other G.7 markets, both pre- and post-2010. Total risk is calculated according to equation (23); Permanent risk is calculated according to equation (24). The post-2010 period is from 2011:01 to 2021:03 (the end-date is determined by the availability of CIP data from Du et al., 2018a). The start-date for the pre-2010 period is 2000:M1, due to data used to estimate within-country convenience yields . Error bars are 68% confidence intervals constructed using Newey and West (1987) standard errors with 4 lags.

Table B.7: Panel Unit-Root Test Results for Variables

Panel A: Long-Maturity Variables			
	$CIP_t^{(10Y)}$	$\mathcal{D}PermRisk_t$	$rx_{t+1}^{CT(10Y)}$
Test Stat.	-2.36^{*}	-1.97	-8.91^{***}
Panel B: Short-Maturity Variables			
	$CIP_t^{(6M)}$	$\mathcal{D}TotRisk_t$	rx_{t+1}^{FX}
Test Stat.	-4.46^{***}	-1.92	-8.17^{***}

Notes: Table displays the W[t-bar] statistics from Im et al. (2003) panel unit-root tests. All tests include a deterministic time trend, the convenience yield tests include 6 lags. Null hypothesis: all panels include unit root. Alternative hypothesis: at least one panel does not include a unit root.

B.7 Supplementary Empirical Results

In this appendix, we demonstrate the robustness of our main results by analysing the full panel of G.7 variables, *vis-à-vis* the U.S., as opposed to only their cross-sectional average. Table B.7 presents unit-root test results using the Im, Pesaran, and Shin (2003) test. Here, the null hypothesis is that all panels contain a unit root, while the alternative implies that some panels are stationary. As in the main body, we include a deterministic trend in our unit-root tests.

The panel results corroborate those for the G.7 average. The null hypothesis that relative permanent risk contains a unit root in all panels cannot be rejected, even with a deterministic time trend included. The test statistic for long-maturity CIP deviations is marginal too; only at the 10% level can the same null be rejected. In contrast, the null for long- and short-maturity currency returns, as well as short-maturity CIP deviations, is strongly rejected, suggesting that at least one panel does not contain a unit root for these variables.

To analyze the association between long-maturity CIP deviations, carry-trade returns and relative permanent risk in the panel dimension, we estimate a panel error-correction model which allows for country-specific long-run relationships between variables. Table B.8 presents mean-group estimates from this specification, alongside Westerlund (2007) panel cointegration test statistics. The mean-group coefficient estimates reveal similar patterns to our benchmark results. There is a significant negative association between long-maturity convenience yields and permanent risk (*vis-à-vis* the U.S.), even in the presence of a deterministic trend. In contrast, there is no long-run relationship between long-maturity convenience and carry-trade returns. The cointegration test statistics also reveal that the null hypothesis of no cointegration among the panels is strongly rejected, in favor of the alternative that at least one of the cross sectional-units is cointegrated (for the Gt and Ga tests) and that all panels are cointegrated (for the Pt and Pa tests).

Table B.8: Cointegration and Error-Correction Amongst Long-Maturity Variables in a Panel Setup

Mean-Group Coefficient Estimates	
Panel A: Long-Run Adjustment	
$\mathcal{D}PermRisk_t$	-0.284**
	(0.122)
$rx_t^{CT(10Y)}$	-0.082
	(0.128)
Deterministic Trend	-0.191***
	(0.039)
Panel B: Short-Run Adjustment	
$\Delta\mathcal{D}PermRisk_t$	-0.130***
	(0.035)
$\Delta rx_t^{CT(10Y)}$	0.100***
	(0.035)
Diseq. Adjustment $\hat{\gamma}$	-0.152***
	(0.041)
Cointegration Tests	
Westerlund Coint. Test Gt	-5.217***
Westerlund Coint. Test Ga	-6.549***
Westerlund Coint. Test Pt	-3.903***
Westerlund Coint. Test Pa	-5.223***

Note. Upper-segment of table displays mean-group coefficient estimates for error-correction model estimated across panel. Sample: 2000:01-2021:03. *** denoting $p < 0.01$, ** $p < 0.05$, and * $p < 0.1$. Lower segment displays Westerlund (2007) panel cointegration test statistics. For all, the null hypothesis is of no cointegration among the panels. For Gt and Ga tests, alternative is that at least one of the cross-sectional units is cointegrated. For Pt and Pa tests, alternative is that all panels are cointegrated. All tests include a deterministic time trend.

C Model Extensions

C.1 Market Structure: Complete Spanning

In this section, we provide a simple example market structure supporting the exchange-rate process proposed in Lemma 1. We build on the discrete-time Cox et al. (1985) model by Sun (1992), see also Backus et al. (2001); Lustig and Verdelhan (2019). Suppose $m_{t,t+1}, m^*_{t,t+1}$ are formed using power utility over consumption:

$$m_{t,t+1} = \beta - \gamma\sqrt{z_t}u_{t+1}, \quad m^*_{t,t+1} = \beta - \gamma\sqrt{z^*_t}u^*_{t+1}$$

where z_t, z^*_t follow AR(1) processes. Moreover, we conjecture the incomplete markets wedge follows:

$$\eta_{t+1} = \alpha + \beta_0 z_t + \beta_1\sqrt{z_t}u_{t+1} + \beta_1\sqrt{z_t}\epsilon_{t+1}$$

where u_{t+1}, ϵ_{t+1} are standard normal innovations reflecting spanned and unspanned risk respectively, as in Lustig and Verdelhan (2019).

Suppose further that there are risky assets: i) one with return $\tilde{r}_{t+1} = \sqrt{z_t}u_{t+1}$ that carries convenience yields $\beta^\theta\theta_t^{H,H(1)}$ for Home investors and $\beta^\theta\theta_t^{F,H(1)}$ for Foreign with $\beta^\theta < 1$ and ii) another with return $\tilde{r}^*_{t+1} = \sqrt{z^*_t}u^*_{t+1}$ that carries convenience yields $\beta^\theta\theta_t^{H,F(1)}$ for Home investors and $\beta^\theta\theta_t^{F,F(1)}$ for Foreign with $\beta^\theta < 1$, such that it is less convenient. These assets are traded by both Home and Foreign investors.

Consider the arbitrage conditions implied by Home and Foreign investors purchasing $1/\beta^\theta$ units of asset (i):

$$\begin{aligned} \mathbb{E}_t\left[M_{t,t+1}\frac{1}{\beta^\theta}\tilde{R}_{t,t+1}\right] &= e^{-\theta_t^{H,H(1)}} \\ \mathbb{E}_t\left[M^*_{t,t+1}\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}^{-1}\frac{1}{\beta^\theta}\tilde{R}_{t,t+1}\right] &= e^{-\theta_t^{F,H(1)}} \end{aligned}$$

equating and using (10) implies:

$$-\mathbb{E}_t[\eta_{t+1}] + \frac{1}{2}\text{var}_t(\eta_{t+1}) - \text{cov}_t(\eta_{t+1}, m_{t+1}) - \text{cov}_t(\eta_{t+1}, \tilde{r}_{t+1}) = \theta_t^{H,H(1)} - \theta_t^{F,H(1)}$$

Since there is still trade in risk-free assets, comparing with the analogous restriction (11) imposes:

$$\text{cov}_t(\eta_{t+1}, \tilde{r}_{t+1}) = \beta_1 z_t = 0$$

where the second equality follows from the processes defined for η_{t+1} and $\tilde{r}_{t+1}$.

In turn, Home and Foreign investors purchasing $1/\beta^{\theta}$ units of asset (ii):

$$\mathbb{E}_t\left[M_{t,t+1}\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}\frac{1}{\beta^{\theta}}\tilde{R}^*_{t,t+1}\right] = e^{-\theta_t^{H,F(1)}},$$
$$\mathbb{E}_t\left[M^*_{t,t+1}\frac{1}{\beta^{\theta}}\tilde{R}^*_{t,t+1}\right] = e^{-\theta_t^{F,F(1)}}$$

equating:

$$\mathbb{E}_t[\eta_{t+1}] + \frac{1}{2}\text{var}_t(\eta_{t+1}) + \text{cov}_t(\eta_{t+1}, m^*_{t+1}) + \text{cov}_t(\eta_{t+1}, \tilde{r}^*_{t+1}) = \theta_t^{F,F(1)} - \theta_t^{H,F(1)}$$

Comparing with the analogous restriction from trade in risk-free assets (10) imposes:

$$\text{cov}_t(\eta_{t+1}, \tilde{r}^*_{t+1}) = \beta_1 z^*_t = 0$$

So, $\beta_1 = \beta^*_1 = 0$ which implies $\text{cov}_t(\eta_{t+1}, m_{t,t+1}) = \text{cov}_t(\eta_{t+1}, m^*_{t,t+1}) = 0$.

Conjecture that $\theta_t^{F,F(1)} - \theta_t^{H,F(1)} = \theta_t^{F,H(1)} - \theta_t^{H,H(1)}$. Then:

$$\mathbb{E}_t[\eta_{t+1}] - \frac{1}{2}\text{var}_t(\eta_{t+1}) = \mathbb{E}_t[\eta_{t+1}] + \frac{1}{2}\text{var}_t(\eta_{t+1})$$

which is only possible for $\text{var}_t(\eta_{t+1}) = 0$. Using (11)-(12) again implies $\alpha = \theta_t^{F,H(1)} - \theta_t^{H,H(1)} = \theta^{F,F(1)} - \theta^{H,F(1)}$ and $\beta_0 = 0$, yielding the process from Lemma 1.

C.2 Further Segmentation and the Short-Run Disconnect

We assume "short" investors price short debt and "long" investors price all long debt and equities. This segmentation can be due to regulation, taxation or more generally a preferred habitat, over and above the bond- specific convenience yields.[29] Define the safe investor's $\hat{M}$ and the risky investor's SDF is related as follows $M = \hat{M}e^{\Delta^{R/S}}$. Consequently:

$$\mathcal{L}(M) = \mathcal{L}(\hat{M}) + \frac{1}{2}\sigma^{2\ R/S}_{\Delta}, \tag{C.1}$$

where we assume log-normality, for simplicity, for the $\Delta^{R/S}$-wedge and $\mathbb{E}[\Delta^{R/S}] = 0$. More generally, the wedge can be allowed to be heteroskedastic and the derivations below follow

[29]This layer of segmentation is over and above that implied by convenience-yields which lie at the centre of our focus. In contrast to convenience yields, this segmentation is common across all long-lived assets, both bonds and equity.

accordingly. Consistent with a high risk premium, *risky* investors have a higher SDF volatility. We abstract from segmentation across currencies (Gabaix and Maggiori, 2015; Itskhoki and Mukhin, 2021) since it affects both short and long maturity conditions equally.

The extended framework we present below shows that using equities measures the SDF volatility of *long* investors. Instead, the equilibrium carry trade returns with short maturity assets reflect the risk differencials for *short* investors—so this condition is mismeasured. Turning to carry trade returns on portfolios formed with long maturity bonds, the relevant SDF volatility switches to that of the "long" investors which is correctly measured by equities.[30]

The relevant set of Euler equations for safe assets is as follows:

$$\mathbb{E}_t\left[\hat{M}_{t,t+1}\right]R_t = e^{-\theta_t^{H,H(1)}} \tag{C.2}$$
$$\mathbb{E}_t\left[\hat{M}^*_{t,t+1}\right]R^*_t = e^{-\theta_t^{F,F(1)}} \tag{C.3}$$
$$\mathbb{E}_t\left[\hat{M}_{t,t+1}\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}\right]R^*_t = e^{-\theta_t^{H,F(1)}} \tag{C.4}$$
$$\mathbb{E}_t\left[\hat{M}^*_{t,t+1}\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right]R_t = e^{-\theta_t^{F,H(1)}} \tag{C.5}$$

The exchange rate process is given by:

$$\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t} = \frac{\hat{M}^*_{t,t+1}}{\hat{M}_{t,t+1}}e^{\theta^{F,H(1)}-\theta^{H,H(1)}} \tag{C.6}$$

and $\theta^{F,H(1)} - \theta^{H,H(1)} = \theta^{F,F(1)} - \theta^{H,F(1)}$. The process can be verified by plugging $\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}$ back into the Euler equations.

On the other hand, for *long* investors characterized by $M_{t,t+1}$, they invest in risky portfolios (equities):

$$\mathbb{E}_t\left[M_{t,t+1}R^g_{t,t+1}\right] = 1 \tag{C.7}$$
$$\mathbb{E}_t\left[M^*_{t,t+1}R^{g*}_{t,t+1}\right] = 1 \tag{C.8}$$

and long maturity bonds:

$$\mathbb{E}_t\left[M_{t,t+k}\right]R_t^{(k)} = e^{-\theta_t^{H,H(k)}} \tag{C.9}$$
$$\mathbb{E}_t\left[M^*_{t,t+k}\right]R_t^{(k)*} = e^{-\theta_t^{F,F(k)}} \tag{C.10}$$

[30] Such frictions between asset markets can also explain the divergence between the Treasury basis and AAA or Refcorp spreads.

$$\mathbb{E}_t\left[M_{t,t+k}\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}\right]R_t^{(k)*}=e^{-\theta_t^{H,F(k)}} \tag{C.11}$$

$$\mathbb{E}_t\left[M^*_{t,t+k}\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right]R_t^{(k)}=e^{-\theta_t^{F,H(k)}} \tag{C.12}$$

Following Alvarez and Jermann (2005), using the Eulers for long-maturity bonds held to maturity, we can calculate:

$$R^{(\infty)}_{t,t+1}=\mathbb{E}_t[M^{\mathbb{T}\ -1}_{t,t+1}]e^{-\theta^{H,H(\infty)}_{t,t+1}},\quad R^{(\infty)\ *}_{t,t+1}=\mathbb{E}_t[M^{*\mathbb{T}\ -1}_{t,t+1}]e^{-\theta^{F,F(\infty)}_{t,t+1}} \tag{C.13}$$

Intuitively, to derive this the short bonds were never used, so the holding period return only captures the "long" investor.

Using the Euler for risky assets:

$$\mathcal{L}(M_{t,t+1})\geq \log\mathbb{E}_t[M_{t+1}]+\mathbb{E}_t\log R^g_{t+1} \tag{C.14}$$

Moreover, we can express:

$$\begin{aligned}\mathcal{L}(M_{t,t+1})&=\log\mathbb{E}_t[M_{t,t+1}]-\mathbb{E}_t[\log M^{\mathbb{T}}_{t,t+1}M^{\mathbb{P}}_{t,t+1}]\\&=\log\mathbb{E}_t[M_{t,t+1}]+\mathcal{L}(M^{\mathbb{P}}_{t,t+1})-\mathbb{E}\left[\log\left(\tfrac{1}{R^{(\infty)}_{t,t+1}}e^{-\theta^{H,H(\infty)}_{t,t+1}}\right)\right]\end{aligned} \tag{C.15}$$

where the second equality uses $M=M^{\mathbb{P}}M^{\mathbb{T}}$. Combining with (C.14), yields our measurement equation for the "risky" SDF:

$$\mathcal{L}(M^{\mathbb{P}}_{t,t+1})\geq\mathbb{E}\log\tfrac{R^g}{R^\infty}-\mathbb{E}_t\theta^{H,H(\infty)}_{t+1} \tag{C.16}$$

Similar derivations apply abroad.

Derivations of Short-Maturity Relationship. Next we derive the short-run carry trade relationship. Taking logs of (C.6) and expectations:

$$\mathbb{E}_t[\Delta e_{t+1}]=\mathbb{E}_t[\hat{m}^*_{t,t+1}]-\mathbb{E}_t[\hat{m}_{t,t+1}]+\theta_t^{F,H(1)}-\theta_t^{H,H(1)} \tag{C.17}$$

Subbing out using the short-run ("safe") Eulers yields Proposition 1:

$$\underbrace{\mathbb{E}_t[\Delta e_{t+1}]+r^*_t-r_t}_{\mathbb{E}_t[rx^{FX}_{t+1}]}=\mathcal{L}(M_{t,t+1})-\mathcal{L}(M^*_{t,t+1})-\frac{1}{2}[\sigma^2_\Delta-\sigma^{*2}_\Delta]+\theta^{F,H(1)}-\theta^{F,F(1)} \tag{C.18}$$

where we use (C.1). Intuitively, we have no measure of the "safe" SDF volatility, so the short-run equilibrium relationship has an error term $-\frac{1}{2}[\sigma_\Delta^2 - \sigma_\Delta^{*2}]$.

Derivations of Long-Maturity Relationship. To derive the long run relationship, we use (C.15):

$$\mathcal{L}(M_{t,t+1}) = \log \underbrace{E_t[\hat{M}_{t,t+1}]}_{\frac{1}{R_t}e^{-\theta_t^{H,H(1)}}} + \frac{1}{2}\sigma_\Delta^2 + \mathcal{L}(M_{t,t+1}^{\mathbb{P}}) - \mathbb{E}\left[\log\left(\frac{1}{R_{t,t+1}^{(\infty)}}e^{-\theta_{t,t+1}^{H,H(\infty)}}\right)\right] \tag{C.19}$$

Subbing this into Proposition 1 yields:

$$\mathbb{E}[rx_{t+1}^{CT(\infty)}] = \mathcal{L}(M_{t,t+1}^{\mathbb{P}}) - \mathcal{L}(M_{t,t+1}^{\mathbb{P}*}) + \mathbb{E}_t[\theta_{t,t+1}^{H,H(\infty)}] - \mathbb{E}_t[\theta_{t,t+1}^{F,F(\infty)}] + \theta^{F,H(1)} - \theta^{H,H(1)} \tag{C.20}$$

where $\mathbb{E}[rx_{t+1}^{CT(\infty)}] = \mathbb{E}_t[rx_{t+1}^{FX}] - (\mathbb{E}_t[rx_{t+1}^{(\infty)}] - \mathbb{E}_t[rx_{t+1}^{(\infty)*}])$. Finally, we use the following relationship for the term structure of convenience yields (re-derived in the next subsection):

$$\mathbb{E}_t[\theta_{t,t+1}^{H,H(\infty)}] = \mathbb{E}_t[\theta_{t,t+1}^{F,H(\infty)}] - \theta_t^{F,H(1)} - \theta_t^{H,H(1)}$$

Subbing this in above yields the relationship in Proposition 2:

$$\mathbb{E}[rx_{t+1}^{CT(\infty)}] = \mathcal{L}(M_{t,t+1}^{\mathbb{P}}) - \mathcal{L}(M_{t,t+1}^{\mathbb{P}*}) + \mathbb{E}_t[\theta_{t,t+1}^{F,H(\infty)}] - \mathbb{E}_t[\theta_{t,t+1}^{F,F(\infty)}] \tag{C.21}$$

Critically, in contrast to Proposition 1, the relevant SDF volatility is correctly measured using equities and long bonds.

Derivation of Term Structure. The results are unchanged relative to the model without segmentation. Taking logs of (C.6) and expectations, evaluated for the long-maturity Eulers:

$$\mathbb{E}_t[\Delta e_{t+k}] = \mathbb{E}_t[m_{t,t+k}^*] - \mathbb{E}_t[m_{t,t+k}] + \theta_t^{F,H(k)} - \theta_t^{H,H(k)} \tag{C.22}$$

This can be rewritten as:

$$\mathbb{E}_t[\Delta e_{t+k}] = \mathbb{E}_t[\hat{m}_{t,t+k}^{R/S*} + \Delta_{t,t+k}^*] - \mathbb{E}_t[\hat{m}_{t,t+k} + \Delta_{t,t+k}^{R/S}] + \theta_t^{F,H(k)} - \theta_t^{H,H(k)} \tag{C.23}$$

Summing the 1-period relationship forward, assuming it holds period by period, and using $\mathbb{E}[\Delta^{R/S(*)}] = 0$, yields:

$$\theta_t^{F,H(k)} - \theta_t^{H,H(k)} = \sum_{i=0}^{k-1} \mathbb{E}_t[\theta_{t+i}^{F,H(1)} - \theta_{t+i}^{H,H(1)}] \tag{C.24}$$

Taking $\theta_t^{F,H(k)} - \theta_t^{H,H(k)} - \mathbb{E}_t[\theta_{t+1}^{F,H(k-1)}] - \mathbb{E}_t[\theta_{t+1}^{H,H(k-1)}]$ and taking the limit as $k \to \infty$:

$$\mathbb{E}_t[\theta_{t,t+1}^{F,H(\infty)}] - \mathbb{E}_t[\theta_{t,t+1}^{H,H(\infty)}] = \theta_t^{F,H(1)} - \theta_t^{H,H(1)} \tag{C.25}$$