U.S. Risk and Treasury Convenience*

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Abstract

We document that, over the past two decades, investors' assessment of U.S. risk has risen relative to other G.7 economies, driven by expectations of greater long-run (permanent) risk. Analytically, we develop a two-country no-arbitrage framework with a rich maturity structure of bonds and convenience yields, alongside equities, which links carry-trade returns, *cross-border* convenience yields and relative country risk across maturities. Informed by our model, we construct novel empirical measures of country risk from bond and equity premia that adjust for *within-country* convenience yields. Taking theory to the data, we find that a perceived increase in U.S. permanent risk is associated with a fall in the convenience yield that Foreign investors attach to long-maturity U.S. Treasuries. Overall, our results suggest that the rise of U.S. risk and fall in long-maturity U.S. Treasuries convenience yields are two sides of the same coin.

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

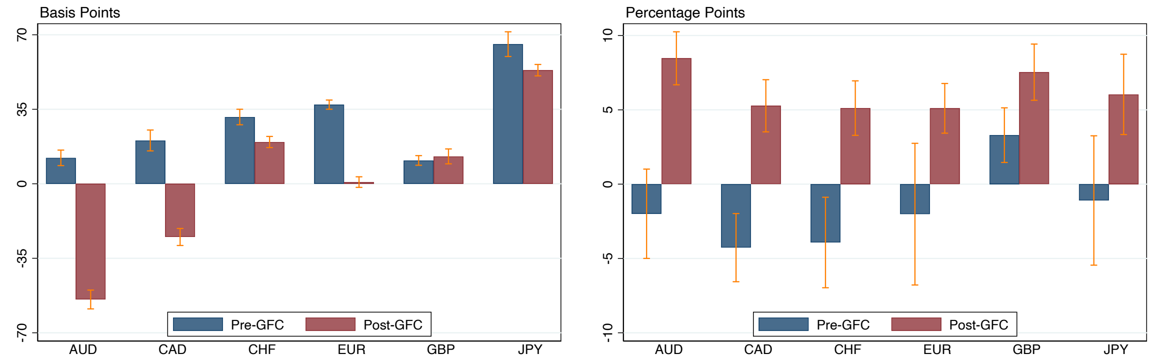
Asymmetries in the international monetary system have been commonplace through history. Today, with a disproportionate share of global debt issuance and trade invoicing denominated in U.S. dollars (e.g., Eichengreen, Hausmann, and Panizza, 2007; Bruno and Shin, 2015; Gopinath, Boz, Casas, Díez, Gourinchas, and Plagborg-Møller, 2020; Maggiori, Neiman, and Schreger, 2020), the U.S. lies at the centre of the global financial system (e.g., Rey, 2015; Miranda-Agrippino and Rey, 2020). As a result of this centrality, the U.S. earns an exorbitant privilege on its external assets (e.g., Gourinchas, Rey, and Govillot, 2010), earns implicit seignorage from abroad due to the relative convenience yield on U.S. Treasuries (e.g., Du, Im, and Schreger, 2018a; Jiang, Krishnamurthy, and Lustig, 2021a), and its currency appreciates in global downturns as investors 'fly to safety' (e.g., Maggiori, 2017; Kekre and Lenel, 2021; Ostry, 2023). While these dimensions of 'specialness' have been studied mostly in isolation, in this paper, we offer a unified treatment of U.S. (dollar) safety in global bond, equity and foreign-exchange markets in both the short-run and long-run.

The extent of U.S. dollar's special role has traditionally been measured by exchange rates, in particular by deviations from uncovered interest parity (UIP), which serve as a barometer of relative risk (e.g., Backus et al., 2001; Engel, 2016). Recently, the specialness of U.S. Treasuries and the dollar have been inferred from deviations in the covered interest parity (CIP), a risk-free arbitrage condition in government bond markets. Consistent with the conventional wisdom, Du et al. (2018a) show that the U.S. Treasury premium on short-maturity bonds, *vis-à-vis* other advanced economies, has tended to be positive, indicating that investors were willing to forego other higher-yield, risk-free government bonds in favour of *convenient* U.S. Treasuries. However, the same risk-free strategy suggests that long-maturity U.S. bonds do not have the same 'shine', and in fact have traded at a discount relative to bonds issued by other G.7 countries, leading to an inverted term structure of CIP deviations. This change is particularly evident when looking at 10-year CIP deviations pre- and post-Global Financial Crisis (GFC) (Figure 1, Panel (a)). The picture is muddied further when we consider that U.S. equity returns have increased substantially post-GFC relative to other G.7 economies (Figure 1, Panel (b)), contributing to suggestions of an end to the U.S.'s exorbitant privilege (Atkeson, Heathcote, and Perri, 2022). Higher U.S. equity returns are consistent with rising U.S. risk (e.g., Farhi and Gourio, 2018).

In this paper, we ask *where* the specialness of the U.S. (dollar) lies? Specifically, we seek to ascertain from the cross-section and time-series of asset prices, within and across countries, whether foreign investors value dollars or U.S. assets, which dollar assets specifically, and at what

Figure 1: U.S. 10-Year CIP Deviations and Equity Returns *vis-à-vis* G.7 Countries

(a) U.S. 10-Year CIP Deviations (b) U.S. Relative Equity Returns



Note. The bars in Panels 1a and 1b reflect the level of average U.S. 10-year CIP deviations, using data from Du et al. (2018a), and equity returns, using data from *MSCI*, respectively, relative to G.7 currencies both pre- and post-GFC. The pre-GFC period ends in 2006:M12; its start is currency-specific (between 1997:M1 and 2000:M2) and is dictated by the availability of CIP data from Du et al. (2018a). The post-GFC period is from 2007:M1 to 2020:M12. Error bars are 68% confidence intervals constructed using Newey and West (1987) standard errors with 4 lags.

maturities? For example, why should U.S. equities deliver higher (expected) returns relative to other G.7 countries while short maturity U.S. bonds are considered particularly safe and liquid? Moreover, why have medium- to long- maturity U.S. bonds been losing their specialness relative to other near-default-free government bonds, a trend that started in the early 2000s?

To answer these questions, consider that standard open-economy models with trade in at least two risk-free bonds denominated in different currencies imply that exchange-rate risk premia can be attributed to differences in stochastic discount factor (SDF) risk across countries (e.g., Backus, Foresi, and Telmer, 2001; Engel, 2014; Lustig, Stathopoulos, and Verdelhan, 2019; Lloyd and Marin, 2020). Combining this with data on asset prices—specifically, the observation that currency excess returns on long-maturity portfolios are small—implies permanent risk across countries—differences in the volatility of permanent innovations to SDFs—should be approximately equal. However, as Figure 1, Panel (b) has demonstrated, this prediction is not verified empirically—in this case, using *ex post* equity returns to proxy for SDF risk. So, how can these predictions and empirical facts be squared?

Our answer to this question lies in the mix of pecuniary returns and non-pecuniary convenience yields, which we show are driven by risk at different horizons.[1] On the theoretical front, we extend the canonical no-arbitrage model (e.g., Backus et al., 2001) to allow for trade

[1]We follow the literature in identifying convenience yields as the residual in the Euler equation for risk-free assets, and take an agnostic view on whether they arise due to liquidity or other factors.

in risk-free bonds of different maturities and a rich term structure of convenience yields. In this class of models, a U.S. investor earns a positive carry-trade return on a portfolio that goes long a Foreign bond (of any maturity), funded by issuing a domestic bond (of commensurate maturity), if the volatility of their SDF is relatively high, as compensation for risk.[2] Allowing for convenience yields, if the U.S. investor experiences higher SDF volatility, they either receive compensation via higher pecuniary carry-trade returns when going long Foreign bonds (greater UIP deviations), or via greater non-pecuniary (convenience) yields on the Foreign bond relative to U.S. Treasuries (a fall in CIP deviations).[3] In this general setting, investors across countries can face different levels of permanent risk, yet carry trade returns on long-maturity assets are near zero because these differences are reflected in convenience yields.

Turning to the data, we construct proxies for total and long-run risk using pricing conditions implied by our framework, and we test the equilibrium relationship between carry trade returns, risk and convenience yields at different maturities. Our main finding is that the inverted term structure of CIP deviations is consistent with rising U.S. risk, relative to other G.7 countries. We show that this is predominantly driven by permanent risk. Specifically, we find that a 1pp increase in investors' perceptions of permanent risk in the U.S. *vis-à-vis* other G.7—measured using a variety of proxies—is associated with a 3-7bp fall in convenience yields. In the context of the near-15pp increase in relative U.S. permanent risk we observe since 2000, this trend explains 45-105bp of the near-3pp decline in convenience yields on long-maturity bonds, *ceteris paribus*—so up to one-third of the observed change. In line with the analysis of Alvarez and Jermann (2005), this is consistent with the higher realised equity premia we observe (Figure 1, Panel (b)). Moreover, our results suggest that the direct and indirect (through convenience yields) effects of higher risk on carry-trade returns offset in the long-run, consistent with a fairly stable level of long-run carry trade returns over our sample. Direct and indirect effects do not offset for carry-trade returns in the short run.

Our findings extend insights on the macro trends identified in Farhi and Gourio (2018) in a number of directions. First, we find that U.S. risk has risen relative to other G.7 countries. In fact, the measure of risk is decreasing in the euro area over the same sample.[4] Moreover, we

[2] If interest rates are low when risk is high, these models are consistent with the Fama (1984) forward-premium or UIP puzzle.

[3] Our paper places a strong emphasis on permanent risk. Permanent risk refers to the volatility of innovations which affects 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. Our focus on permanent risk and, by extension, long maturity CIP deviations is motivated by both the importance of permanent risk in explaining high-yielding risky asset returns (e.g., Alvarez and Jermann, 2005; Hansen and Scheinkman, 2009).

[4] Due to lack of data, the correct measure of risk corresponding to our framework can only be constructed for the euro area and the U.S. However, the relative equity premium of the U.S. vs. other G.7 countries increased too.

find that the rise in U.S. risk is specifically driven by an increase in permanent risk, a finding consistent with the increase in disaster probability suggested in Farhi and Gourio (2018), or the switch to a permanent-innovation regime in Chernov, Lochstoer, and Song (2021).

We emphasise two further contributions we make to the literature. Our measures of country risk and permanent risk build on Hansen and Jagannathan (1991) and are closely related to Alvarez and Jermann (2005). However, relative to these papers we make two improvements. First, whereas the Alvarez and Jermann (2005) bounds posit that the permanent volatility of the SDF is at least as large as the equity-net-bond premium, this is not the case in our extended framework. In a model with a term structure of convenience yields, we show that permanent risk is bounded by the difference between the equity-net-bond premium and the convenience yield earned from holding an infinite maturity bond for a single period. High equity-net-bond premia either reflect high permanent risk, or artificially low bond premia due to the high convenience. Additionally, relative to Alvarez and Jermann (2005), we construct expected equity premia measures using the Gordon (1962) growth model, without using data on realised returns since these are noisy and confound compensation for risk with other factors.[5] In contemporaneous work, Jiang and Richmond (2023) generalise the Hansen and Jagannathan (1991) bounds for total SDF risk to allow for convenience yields.

Finally, we find strong evidence that the transmission of risk to pecuniary and non-pecuniary returns differs across maturities. Specifically, we find that whereas short maturity convenience yields tend to be lower when G.7 country risk rises, long maturity convenience yields tend to fall when U.S. risk rises. Additionally, we extend the analysis of Jiang, Krishnamurthy, and Lustig (2021a) and find that long maturity U.S. treasuries are more valued for the 'dollarness' than short maturity Treasuries—i.e., they have less intrinsic value.

Related Literature. Our paper relates to several interrelated strands of the literature, that focus on convenience yields, the measurement of risk from asset prices, and the predictability of currency returns. Much of 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; Liao and Zhang, 2020; Liu, Schmid, and Yaron, 2020; Diamond and Van Tassel, 2022). Other studies use convenience yields to explain exchange-rate dynamics, predominantly at short horizons (Engel and Wu, 2018; Krishnamurthy and Lustig, 2019; Valchev, 2020; Jiang, Krishnamurthy,

[5]Nonetheless, we show that within our sample, results are qualitatively consistent with using realised equity returns.

and Lustig, 2021a,b; Ostry, 2023). Our paper is distinct in that we specifically study the consequences of macroeconomic risk on convenience yields, but also, how convenience yields can affect other quantities in the macroeconomy and financial markets.

Moreover, we place particular emphasis on the fall in long-maturity U.S. Treasury convenience. Taking the market microstructure view, Du et al. (2022) investigate the drivers of this fall *within* the U.S., measuring *within-country* convenience yields with the swap spread. Their explanation is the post-GFC U.S. fiscal expansion, coupled with tighter financial regulations that moved primary dealers from being net-short to net-long U.S. Treasuries. Relatedly, Augustin, Chernov, Schmid, and Song (2020) attribute 30% of long-maturity CIP deviations to intermediary constraints. Our paper takes a cross-country view, linking the decline in long-maturity convenience to relative risk.

Our paper also draws on a large literature which uses asset prices to ascertain characteristics of SDFs and risk (Hansen and Jagannathan, 1991; Alvarez and Jermann, 2005; Hansen and Scheinkman, 2009; Bakshi and Chabi-Yo, 2012). In closely related, recent contributions, Jiang and Richmond (2023) generalise the Hansen and Jagannathan (1991) bounds to allow for convenience yields and show this can reconcile the high Sharpe ratio on safe bonds.[6] Di Tella, Hébert, Kurlat, and Wang (2023) construct a zero-beta portfolio and find evidence of very high convenience yields. While our paper cannot speak to their finding, what matters in our environment is the relative convenience yield on different assets and across countries.

Finally, we contribute to a classical literature on the exchange-rate 'disconnect' (Meese and Rogoff, 1983; Obstfeld and Rogoff, 2000; Itskhoki and Mukhin, 2021). Many remedies for the failure of UIP, in particular, have been advanced including convenience yields; imperfect financial markets (Jeanne and Rose, 2002; Gabaix and Maggiori, 2015), long-run risk (Colacito and Croce, 2011), disaster risk (Farhi and Gabaix, 2016); and transitory risk (Lloyd and Marin, 2020; Gourinchas, Ray, and Vayanos, 2022), among others. Lilley, Maggiori, Neiman, and Schreger (2019) document a 'reconnect' following the GFC. Relative to these papers, we provide evidence of some 'connection' between exchange-rate dynamics, convenience and relative risk.

Outline. The remainder of this paper is structured as follows. Section 2 outlines the model framework. Section 3 details the main equilibrium relationships which arise. Section 4 outlines

[6] While they find that the Hansen and Jagannathan (1991) bound must be adjusted by the level of the convenience yield on a risk-free bond, scaled by the volatility of the risk free-rate, we find that the adjustment to the Alvarez and Jermann (2005) bound depends on the expected change in the convenience yield of an infinite-maturity bond held to maturity. In both cases, the adjustment is in the same direction since higher convenience yields lower the risk-free rate (Jiang and Richmond, 2023), or the bond premium (in this paper), therefore risky asset returns need not be as high to reflect SDF volatility.

our measures of risk, convenience yields and portfolio returns. Section 5 presents our empirical analysis. Section 6 concludes.

2 A Model of Convenience Yields and Risk

In this section, we introduce the notation and theoretical framework used in this paper. We consider an environment with two countries: Home (i.e., the U.S.) and Foreign (denoted with an asterisk $*$). Representative investors in each country can trade in zero-coupon bonds of varying maturities $k = 1, 2, ..., \infty$ issued in both the Home and Foreign 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 each investor $i = H, F$, asset $j = H, F$ and maturity k, and varies over time. The fact investors can trade in Home and Foreign bonds gives rise to foreign exchange and the possibility to carry trade, from which returns can vary over the term structure. In addition, investors can trade in domestic risky equity.

2.1 Bond Markets

We 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: $R_t \equiv R_t^{(1)} = 1/P_t^{(1)}$. As well as 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 is standard in the literature, we assume these convenience yields satisfy:

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 (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.*

This assumption ensures that, at every point in time t, the representative country-i investor knows the convenience they can earn on each maturity k of asset j. To the extent that the convenience yield arises from a bond's value as collateral, this timing assumption reflects that fact that collateral value will be accounted for in contracts written at time t. Moreover, this assumption clarifies that the convenience from each asset is a 'yield', distinct from the pecuniary return that the bond pays.

We define the nominal exchange rate $\mathcal{E}_t$ to have units of U.S. dollars per unit of Foreign currency. Therefore, an increase in $\mathcal{E}_t$ corresponds to a U.S. (Home) depreciation and a Foreign appreciation.

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

$$e^{-\theta_t^{H,H(k)}} = \mathbb{E}_t\left[M_{t,t+k}R_t^{(k)}\right] \tag{1}$$

$$e^{-\theta_t^{F,F(k)}} = \mathbb{E}_t\left[M^*_{t,t+k}R_t^{(k)*}\right] \tag{2}$$

$$e^{-\theta_t^{H,F(k)}} = \mathbb{E}_t\left[M_{t,t+k}\frac{\mathcal{E}_{t+k}}{\mathcal{E}_t}R_t^{(k)*}\right] \tag{3}$$

$$e^{-\theta_t^{F,H(k)}} = \mathbb{E}_t\left[M^*_{t,t+k}\frac{\mathcal{E}_t}{\mathcal{E}_{t+k}}R_t^{(k)}\right] \tag{4}$$

for all maturities k.[7]

In addition, pecuniary and non-pecuniary returns can be traded prior to a bond's maturity. We define the one-period pecuniary holding return on a k-period Home bond as: $R_{t,t+1}^{(k)} = P_{t+1}^{(k-1)}/P_t^{(k)}$, such that $R_{t,t+k}^{(k)} \equiv R_t^{(k)} = 1/P_t^{(k)}$, for all k. Taking the Home Euler equation for the Home bond (1) as an example, we can write the Euler equation from purchasing a k-maturity bond at time t and selling it at $t+1$ as a $(k-1)$-maturity bond as:

$$e^{-\theta_t^{H,H(k)}} = \mathbb{E}_t\left[M_{t,t+1}R_{t,t+1}^{(k)}e^{-\theta_{t+1}^{H,H(k-1)}}\right] \tag{5}$$

with analogous expressions for the other three trades captured by equations (2)-(4). This equation captures the intuition that when an investor sells a k-period bond prior to maturity, they earn the convenience yield $\theta_t^{i,j(k)}$ at time of purchase t, but forego the yield $\theta_{t+1}^{i,j(k-1)}$ at time of resale $t+1$. It can be shown that $\mathbb{E}_t[e^{\theta_{t,t+1}^{(k)} - \theta_t^{(k)} + \theta_{t+1}^{(k)}}] = 1$, so the expression $\theta_t^{i,j(k)} - \theta_{t+1}^{i,j(k-1)}$ can be interpreted as the (log) one-period convenience holding return on a k period country-j asset for an investor in country i, which is analogous to the pecuniary equivalent $R_{t,t+1}^{(k)}$.

[7]Such convenience yields could be the outcome of an extension to the bond in the utility function formulation used in Valchev (2020) and Jiang and Richmond (2023), but here we remain agnostic on the specific microfoundation.

2.2 Equity Markets

In addition to trading in domestic and foreign bonds, representative investors in each country can trade in domestic risky assets (i.e., equity). We assume these equities satisfy:

Assumption 2 (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, which we normalise to zero.*

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

$$1 = \mathbb{E}_t\left[M_{t,t+1}R^g_{t,t+1}\right] \tag{6}$$

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

2.3 Foreign Exchange

Finally, to close the model, we conjecture an exchange-rate process:

$$\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t} = \frac{M^*_{t,t+1}}{M_{t,t+1}} e^{\theta_t^{F,H(1)} - \theta_t^{H,H(1)}} \tag{8}$$

for all t.[8] Denoting logarithms with lower-case letters (i.e., $x \equiv \log(X)$), we rewrite this process as:

$$\Delta e_{t+1} = m^*_{t,t+1} - m_{t,t+1} + \theta_t^{F,H(1)} - \theta_t^{H,H(1)} \tag{9}$$

where $\Delta e_{t+1} \equiv e_{t+1} - e_t$. This process satisfies Euler equations (1) to (4). Because the exchange-rate process is common to all Euler equations, no-arbitrage implies that Home and Foreign investors agree on the relative convenience yield on Home and Foreign bonds, such that:

$$\theta_t^{F,H(1)} - \theta_t^{H,H(1)} = \theta_t^{F,F(1)} - \theta_t^{H,F(1)} \tag{10}$$

for all t. Trade in long-maturity assets imposes additional restrictions which we discuss in Lemma 2.

[8] We interpret our model to be a form of incomplete markets, induced by missing liquidity markets reflected by convenience yields. The model can be generalised to allow for incomplete markets using stochastic wedges (as in Backus et al., 2001; Lustig and Verdelhan, 2019). In our environment, as in Jiang et al. (2021b), such wedges can interact with convenience yields. However, these interactions do not alter the main results in this paper, so we defer their study to future work.

Before proceeding, we also define the conditional entropy operator for a generic variable X as: $\mathcal{L}_t(X_t) \equiv \mathbb{E}_t[\log X_t] - \log \mathbb{E}_t[X_t]$, where $\mathcal{L}_t(X_t) = \frac{1}{2}\text{var}_t(X_t)$ if X_t is log-normally distributed. Similarly, $\mathcal{L}(X_t)$ denotes the unconditional entropy operator.[9]

3 Convenience, Returns and Risk

In this section, we discuss the model's predictions for the links between convenience, risk and returns in the short and long run.

3.1 Short-Run Convenience, Returns and Risk

We denote by rx^{FX}_{t+1} the *ex post* log return from a one-period carry-trade strategy that goes long the Foreign risk-free bond and short the U.S. risk-free bond:

$$rx^{FX}_{t+1} = \log\left(\frac{R^*_t}{R_t}\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}\right) = r^*_t - r_t + \Delta e_{t+1}. \tag{11}$$

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 can earn a cross-border carry-trade excess return.

Within a classic open-economy setup without convenience yields (i.e., $\theta^{i,j(k)}_t = 0$ 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. In this setting without convenience yields, 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 short-run cross-border carry-trade portfolio that is long the Home bond and short the Foreign bond ($\mathbb{E}_t[rx^{FX}_{t+1}] < 0$) (see, e.g., Backus et al., 2001). The covariance of this return with yields constitutes the forward-premium (or Fama) puzzle (e.g., Fama, 1984). This classic open-economy setup can replicate the Fama puzzle as long as short-term interest rate are procyclical. Verdelhan (2010) achieves this in a model with habits, but as Lloyd and Marin (2020) discuss, at the cost of a counter-factually downward-sloping yield curve.

However, when investors additionally attach convenience yields to short-run bonds, as in

[9] This measure of conditional volatility $\mathcal{L}_t(X)$, often referred to as Theil (1967)'s second conditional entropy measure. It is 0 if $\text{var}_t(X) = 0$, and if X_{t+1} is log-normally distributed, then $\mathcal{L}_t(X_{t+1}) = \frac{1}{2}\text{var}_t(X_{t+1})$. In general, the conditional volatility $\mathcal{L}_t(X_{t+1})$ is 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 ($\frac{\kappa_{3,t}}{3!} + \frac{\kappa_{4,t}}{4!} + ...$).

our setup, the link between the exchange-rate risk premium and relative SDF volatility (risk) takes a different form, as we demonstrate in the proposition below:

Proposition 1 (SDF Volatility, FX Risk and Convenience Yields in the Short Run) *Given $M^*_{t,t+1}$, $M_{t,t+1}$ and the exchange-rate process (9), then relative SDF volatility, the expected excess return on a one-period carry-trade strategy, and one-period convenience yields satisfy:*

$$\mathcal{L}_t(M_{t,t+1}) - \mathcal{L}_t(M^*_{t,t+1}) - \mathbb{E}_t[rx^{FX}_{t+1}] + \theta^{F,H(1)}_t - \theta^{F,F(1)}_t = 0 \tag{12}$$

Proof: Combine expectations of the exchange-rate process (9) and the log-entropy expansions of the $k = 1$-period domestic Euler equations (1) and (2). See Appendix A.1 for full derivation. □

In the model with convenience yields, if the Foreign investor bears greater SDF volatility ($\mathcal{L}_t(M_{t,t+1}) < \mathcal{L}_t(M^*_{t,t+1})$), they earn either greater expected pecuniary returns on a net-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 both. 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 risks across countries can equilibrate.[10] This convenience-based channel may help to explain the apparent failure of UIP in the short-run.

3.2 Transitory and Permanent SDF Risk Decomposition with Convenience

To investigate the properties of risk in the long run, we rely on the Alvarez and Jermann (2005) decomposition of pricing kernels. Following their approach, we decompose the Home pricing kernel Λ_t into two components:

$$\Lambda_t = \Lambda^{\mathbb{P}}_t \Lambda^{\mathbb{T}}_t \tag{13}$$

where $\Lambda^{\mathbb{P}}_t$ is a martingale that captures the 'permanent' component of Λ_t, while $\Lambda^{\mathbb{T}}_t$ reflects the 'transitory' component. We decompose the Foreign pricing kernel analogously.

[10] Proposition 1 is entirely symmetric if the Home investor bears greater risk ($\mathcal{L}_t(M_{t,t+1}) > \mathcal{L}_t(M^*_{t,t+1})$), in which case they earn either greater pecuniary returns from a carry trade net-long the Foreign bond ($\mathbb{E}_t[rx^{FX}_{t+1}] > 0$) or a greater non-pecuniary convenience yield on the Foreign bond relative to the Home bond ($\theta^{H,F(1)}_t - \theta^{H,H(1)}_t > 0$). The latter result can be seen by substituting (10) into Proposition 1.

Under regularity conditions,[11] Alvarez and Jermann (2005) show that the permanent component is defined as $\mathbb{E}_t \Lambda^{\mathbb{P}}_{t+1} = \Lambda^{\mathbb{P}}_t$, where $\Lambda^{\mathbb{P}}_t = \lim_{k\to\infty} \frac{\mathbb{E}_t \Lambda_{t+k}}{\beta^{t+k}}$. 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: $\Lambda^{\mathbb{T}}_t = \lim_{k\to\infty} \frac{\beta^{t+k}}{\mathbb{E}_t[\Lambda_{t+k}]/\Lambda_t}$. This is equivalent to a scaled long-term interest rate.

Employing the conditional volatility measure $\mathcal{L}_t(\cdot)$ to gauge a country's SDF risk, we can derive restrictions on the transitory and permanent components of pricing kernels. While the approach we follow has parallels with Alvarez and Jermann (2005) and Lustig et al. (2019), we show that the presence of convenience yields alters these restrictions in intuitive and meaningful ways.

First, we derive the link between the transitory component of representative investors' SDF, long-term bond returns and long-horizon convenience yields. Under mild regularity conditions analogous to Alvarez and Jermann (2005), the following lemma summarises this:

Lemma 1 (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 change its convenience yield from time t to $t+1$:*

$$\frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t} \equiv M^{\mathbb{T}}_{t,t+1} = \frac{1}{R^{(\infty)}_{t,t+1}} e^{\theta^{H,H(\infty)}_{t+1} - \theta^{H,H(\infty)}_t} \tag{14}$$

There are analogous expressions for the representative Foreign investor.

Proof: See Appendix A.2. □

Absent convenience yields, transitory innovations to pricing kernels—the transitory SDF—reflects the bond term premium only. Instead, with convenience yields on domestic bonds, the transitory SDF is adjusted for changes in long-term convenience. Specifically, when holding an infinite-maturity bond for a single period, the investor earns the convenience yield today on that bond, but sells its convenience yield tomorrow. So, the transitory SDF (14) 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.

[11] 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.

Moreover, in the spirit of Alvarez and Jermann (2005), we can derive bounds on the volatility of the overall and permanent components of SDF risk, in relation to equity premia. This is a key step in our analysis since this will be our preferred measure of overall and permanent SDF risk in our empirical analysis. As the following proposition clarifies, the presence of convenience yields also influences these bounds:

Proposition 2 (Permanent and Total SDF 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:*

$$\mathcal{L}_t(M_{t,t+1}) \geq \mathbb{E}_t \log \left[\frac{R^g_{t,t+1}}{R_t} \right] - \theta_t^{H,H(1)} \tag{15}$$

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

$$\mathcal{L}_t \left(M^{\mathbb{P}}_{t,t+1} \right) \geq \mathbb{E}_t \log \left[\frac{R^g_{t,t+1}}{R_t} \right] - \mathbb{E}_t \left[rx^{(\infty)}_{t+1} \right] - \theta_t^{H,H(\infty)} + \mathbb{E}_t \left[\theta_{t+1}^{H,H(\infty)} \right] \tag{16}$$

There are analogous expressions for the representative Foreign investor.

Proof: See Appendix A.3. □

Inequality (15) in Proposition 2 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 (15) holds for any risky asset, the right-hand side is maximised by using the risk premium on a country's equity. Intuitively, this bound on overall SDF 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) and Lustig et al. (2019)—the additional convenience yield term reflects 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 $\theta_t^{H,H(1)}$.

Inequality (16) in Proposition 2 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 (16)—and the expected one-period return on the asymptotic discount bond, net of changes in the convenience yield on this bond. Again, while inequality (16) holds for any risky asset, the right-hand side is maximised by using the risk premium on the country's equity. Intuitively, this bound on the permanent component of SDF risk reflects that while the riskiest asset in the economy captures both permanent and transitory risk, the premium and convenience yield on a long-maturity bond encodes the transitory (insurable) component of risk only. Relative to a model without convenience, the additional convenience yield terms in (16) reflect that by forgoing the one-period pecuniary return on an infinite maturity bond $\mathbb{E}_t\left[rx_{t+1}^{(\infty)}\right]$, the Home investor also sells the bonds convenience today, but expects to recover its convenience tomorrow when they receive the bond $-\theta_t^{H,H(\infty)} + \mathbb{E}_t\left[\theta_{t+1}^{H,H(\infty)}\right]$.

3.3 Long-Run Convenience, Returns and Risk

While the risk-based explanation, alongside the convenience-based explanation in Proposition 1, may help to reconcile the failure of UIP in the short run, failures to reject UIP over long horizons (e.g., Chinn and Meredith, 2005; Chinn and Quayyum, 2012) mean that risk-based explanations alone imply counterfactually strict restrictions in the long run. Using our SDF decomposition, we can now assess the link between the exchange-rate risk premium and SDF risk in the long run when investors attach convenience yields to bonds. The following proposition presents the unconditional relationship between currency returns, permanent risk and convenience:

Proposition 3 (Long-Run *Unconditional* SDF Volatility, FX Risk and Convenience) *Given $M_{t,t+1}$, $M_{t,t+1}^*$ and the exchange-rate process (9), then unconditional relative permanent SDF volatility, unconditional long-run exchange-rate risk, and unconditional long-run convenience yields satisfy:*

$$\mathcal{L}(M_{t,t+1}^{\mathbb{P}}) - \mathcal{L}(M_{t,t+1}^{\mathbb{P}*}) - \lim_{k\to\infty}\frac{1}{k}\mathbb{E}\left[rx_{t+k}^{FX}\right] + \lim_{k\to\infty}\frac{1}{k}\left\{\mathbb{E}[\theta_t^{F,H(k)}] - \mathbb{E}[\theta_t^{F,F(k)}]\right\} = 0 \tag{17}$$

Proof: See Appendix A.4. □

Absent convenience yields, the classical model suggests that the failure to reject UIP at long horizons ($\lim_{k\to\infty}\frac{1}{k}\mathbb{E}\left[rx_{t+k}^{FX}\right] = 0$) implies that permanent risk should be equalised across countries ($\mathcal{L}(M_{t,t+1}^{\mathbb{P}}) - \mathcal{L}(M_{t,t+1}^{\mathbb{P}*})$), as shown in Lustig et al. (2019). Instead, Equation (17) of Proposition 3 highlights that if UIP holds at long horizons, then differences in permanent risk across countries should reflect differences in the relative convenience derived from Home and Foreign bonds. Specifically, if foreign investors' bear greater permanent SDF risk unconditionally, such that $\mathcal{L}(M_{t,t+1}^{\mathbb{P}}) < \mathcal{L}(M_{t,t+1}^{\mathbb{P}*})$, they must earn a greater (unconditional) non-pecuniary convenience yield on net-long positions in the Home bond (i.e., $\theta^{F,H(\infty)} > \theta^{F,F(\infty)}$), since the pecuniary returns to this cross-border currency position are zero. This characterises a simple general equilibrium condition between convenience yields and macroeconomic risk in the long run.

3.4 SDF Risk and the Term Structure of Convenience and FX Premia

Propositions 1 and 3 individually characterise links between currency returns, convenience yields and SDF risk in the short- and long-run, respectively. However, our setup also imposes restrictions on the term-structure of this relationship, which implies a stronger, conditional long-run general-equilibrium condition.

Intuitively, since no-arbitrage implies that investors' agree on one-period exchange rate movements, as discussed in (10), it also requires that Home and Foreign investors agree on the one-period non-pecuniary relative convenience of holding Home and Foreign bonds. Similarly, since investors taking short- and long-maturity cross-border portfolios agree on the evolution of exchange rates over time, they must also agree on the evolution of relative convenience yields. The following lemma summarises the restrictions this imposes on the term structure of convenience yields in our setup:

Lemma 2 (Term Structure of Convenience Yields) *Given $M_{t,t+1}$, $M^*_{t,t+1}$, the Euler equations (1)-(4), and the exchange-rate process (9), term structure of convenience yields satisfies the following conditions:*

$$\theta_t^{F,H(k)} - \sum_{\tau=0}^{k-1} \theta_{t+\tau}^{F,H(1)} = \theta_t^{H,H(k)} - \sum_{\tau=0}^{k-1} \theta_{t+\tau}^{H,H(1)} \tag{18}$$

for all k and all t. There is an analogous expression for the Home and Foreign investors' convenience yields on Foreign bonds.

Proof: See Appendix A.5. □

In equation (18), 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 deviations from the 'expectations hypothesis' of convenience yields or, otherwise put, 'convenience term premia'. The result in the Lemma states that, for a given bond j, these deviations from the convenience-yield expectations hypothesis on a specific bond (the H-bond above) should be proportional for Home and Foreign investors.

Using this restriction, we can then derive a conditional relationship between relative permanent risk, long-maturity convenience yields, and the expected one-period carry-trade return fashioned using long-maturity Home and Foreign bonds. To do so, we define expected carry-trade return from taking a long position in the Foreign ∞-maturity bond and a short position in the Home ∞-maturity bond for one period as:

$$\mathbb{E}_t[rx_{t+1}^{(\infty),CT}] = \underbrace{\mathbb{E}_t[rx_{t+1}^{FX}]}_{\text{Currency Returns}} + \underbrace{\mathbb{E}_t[rx_{t+1}^{(\infty)*}] - \mathbb{E}_t[rx_{t+1}^{(\infty)}]}_{\text{Difference in Local Bond Returns}} \tag{19}$$

The following proposition summarises the relationship:

Proposition 4 (SDF Volatility, FX Risk and Convenience Yields in the Long Run) *Given $M_{t,t+1}$, $M_{t,t+1}^*$, the exchange-rate process (9) and the restrictions on the term structure of convenience yields imposed by Lemma 2, then relative permanent SDF volatility, the one-period carry-trade returns from long-term bonds, and the change in the relative long-maturity convenience yield satisfy:*

$$\mathcal{L}_t(M_{t,t+1}^{\mathbb{P}}) - \mathcal{L}_t(M_{t,t+1}^{\mathbb{P}*}) - \mathbb{E}_t[rx_{t+1}^{(\infty),CT}] + \left(\theta_t^{F,H(\infty)} - \theta_t^{F,F(\infty)}\right) - \left(\mathbb{E}_t[\theta_{t+1}^{F,H(\infty)}] - \mathbb{E}_t[\theta_{t+1}^{F,F(\infty)}]\right) = 0 \tag{20}$$

Proof: See Appendix A.6. □

Similar to Proposition 3, the relationship documented in Proposition 4 highlights that if pecuniary carry trade returns on long-term bonds are zero ($\mathbb{E}_t[rx_{t+1}^{(\infty),CT}] = 0$), as they are in the data (Lustig et al., 2019), asymmetries in permanent SDF risk across countries must be matched by movements in long-maturity convenience yields. Specifically, if foreign investors' bear greater permanent SDF risk, such that $\mathcal{L}_t(M_{t,t+1}^{\mathbb{P}}) < \mathcal{L}_t(M_{t,t+1}^{\mathbb{P}*})$, equation (20) demonstrates that these Foreign investors must receive a greater non-pecuniary convenience yield from their net-long position in the Home bond at time t, that is, $(\theta_t^{F,H(\infty)} - \theta_t^{F,F(\infty)}) \uparrow$ and be expected to forgo less convenience on the Home bond when they unwind their carry trade position at time $t+1$, i.e., $(\mathbb{E}_t[\theta_{t+1}^{F,H(\infty)}] - \mathbb{E}_t[\theta_{t+1}^{F,F(\infty)}]) \downarrow$, since the pecuniary returns to this carry trade position are zero.

4 Data and Measurement

Before empirically assessing our model predictions, in this section we describe our data and explain how we overcome particular challenges to measuring objects in our model.

We focus our empirical analysis on G.7 countries (currencies): Australia (AUD), Canada (CAD), Switzerland (CHF), euro area/Germany (EUR), Japan (JPY), U.K. (GBP) and U.S. (USD). Our analysis is carried out using an unbalanced panel at a monthly frequency, over the period 1997:01-2020:12.[12] Exchange-rate data is from *Datastream*, and the U.S. is our base country among our sample of advanced economies. In our baseline empirics, we follow a common approach in the literature by using *ex post* changes in exchange rates to proxy for *ex ante* expectations.[13]

We use information on the term structure of interest rates in government bond markets for each of these regions, which are highly liquid, to measure bond yields. Specifically, we use 6-month nominal zero-coupon bond yields in each jurisdiction as our measure of short-term 'safe' interest rates. Like Lustig et al. (2019) and others, we also use 10-year nominal zero-coupon bond yields as our proxy for the long-term, infinite-maturity, bond yield, which amounts to assuming that the yield curve is sufficiently flat beyond the 10-year horizon. 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).

Using these exchange rate and interest rate data, Figure 2 plots the time-series of one-period carry trade returns fashioned using short-term government bonds—the exchange-rate risk premium (rx^{FX}_{t+1}) in Panel 2a and the one-period carry trade return fashioned using long-term bonds ($rx^{(\infty),CT}_{t+1}$) in Panel 2b. Clearly visible in both panels is the large carry-trade loss investors earned during the GFC by taking long-positions in the foreign currency and short positions in the U.S. dollar. Overall, however, there is little evidence of trend movements in pecuniary dollar returns at short- or long-horizons over the past 20 years.

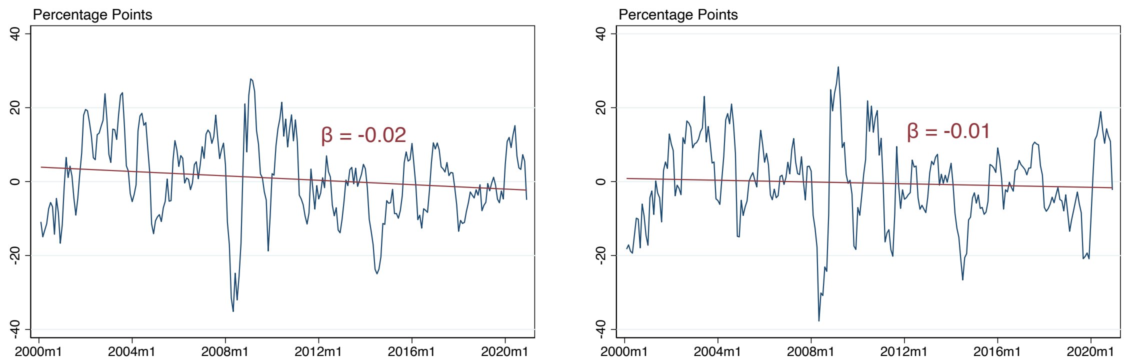
For the the other key objects in our model, namely the term structure of convenience yields across and within countries (i.e., $\theta^{i,j(k)}_t$ for $i \neq j$ and $\theta^{i,i(k)}_t$, respectively), as well as SDF risk, there are specific challenges associated with measurement. We explain how we overcome each, in turn, in the following three sub-sections.

[12]The start and end of our sample is restricted by the convenience yield data from Du et al. (2018a). See Appendix B for details.

[13]We note, however, that this approach is not without loss of generality, especially when there are deviations from full-information rational expectations (see, e.g., Stavrakeva and Tang, 2020; Candian and De Leo, 2023, for discussions).

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

(a) U.S. Short-Run Exchange-Rate Risk Premium (b) U.S. Long-Run Carry Trade Return



Note. Panels 2a and 2b display time series of short-run and long-run one-period U.S. dollar carry-trade returns (rx^{FX}_{t+1} and $rx^{(\infty),CT}_{t+1}$ from equations (11) and (19)), respectively, from 2000:M2 to 2020:M12. In each case, the time series are a cross-sectional average across G.7 currency areas *vis-à-vis* the U.S. The one-period holding period is set to 6 months and the bond maturities are 6 months (short-run) and 10 years (long-run), respectively. *** signifies that the slopes (β) of estimated deterministic trend lines are not equal to zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are chosen to match the CIP-deviation data.

4.1 Measuring Cross-Country U.S. Convenience Yields

We measure the convenience yield a Foreign investor earns from holding a U.S. Treasury, relative to holding their own Foreign bond, (i.e., $\theta_t^{F,H(k)} - \theta_t^{F,F(k)}$) using deviations from Covered Interest Parity (CIP) on government bonds. Our approach to doing so builds on Jiang et al. (2021a), but extends the logic to the term structure of convenience—not only short maturities.

Our measure of the term structure of government-bond CIP deviations, vs. the U.S., is the Treasury basis from Du et al. (2018a). To match our data on bond yields, we use the 6-month CIP deviation to construct our proxy of short-term convenience, and the 10-year CIP deviation to construct our measure of long-term convenience yields.[14] The k-maturity CIP deviation $CIP_t^{(k)}$ is defined as the log return on a covered position that goes long a k-period Foreign bond and short a k-period U.S. Treasury (Home bond):

$$CIP_t^{(k)} = \log\left(\frac{R_t^{(k)*}}{R_t^{(k)}}\frac{F_t^{(k)}}{\mathcal{E}_t}\right) = r_t^{(k)*} - r_t^{(k)} + f_t^{(k)} - e_t \tag{21}$$

where $f_t^{(k)} \equiv \log(F_t^{(k)})$ is the log of the k-period nominal forward exchange rate, with the same units as e_t. When $CIP_{t,k} > 0$, the pecuniary return on a synthetic U.S. dollar-denominated

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

bond, $f_t^{(k)} - e_t + r_t^{(k)*}$, is greater than the pecuniary return on a U.S. Treasury, $r_t^{(k)}$. To the extent that arbitraging CIP deviations is riskless, this implies the non-pecuniary return on the U.S. bond must be greater than that on the Foreign bond.

We map these CIP deviations into the relative convenience yields on U.S. Treasury enjoyed by Foreign investors, which are represented by $\theta_t^{F,H(k)} - \theta_t^{F,F(k)}$ in our model. The main idea underpinning this mapping is that that Foreign investors derive a relative convenience yield from holding U.S. Treasuries both due to the convenience of the Treasury itself and because Treasuries are a claim to U.S. dollars, which may themselves be convenient. These Foreign investors can earn the dollar portion of this convenience by holding synthetic U.S. Treasuries (i.e., by holding Foreign government bonds swapped into dollars using a forward contract). As in Jiang et al. (2021a), we posit that this synthetic Treasury position, in addition to earning $\theta_t^{F,F(k)}$, earns a fraction β_k^* of the relative convenience of the actual Treasury position $\theta_t^{F,H(k)} - \theta_t^{F,F(k)}$. In other words:

$$\mathbb{E}_t \left[M_{t,t+k}^* \frac{\mathcal{E}_t}{\mathcal{E}_{t+k}} \left(\frac{F_t^{(k)}}{\mathcal{E}_t} R_t^{(k)*} \right) \right] = e^{-\theta_t^{F,F(k)} - \beta_k^* (\theta_t^{F,H(k)} - \theta_t^{F,F(k)})} \tag{22}$$

where the term in round brackets on the left-hand side is the synthetic k-period U.S. Treasury pecuniary return.

Comparing equations (21) and (22), and using equation (4), we can derive a linear, maturity-specific, relationship between the convenience yield and the CIP deviation, which is governed by the maturity-specific factor β_k^*:

$$CIP_t^{(k)} = (1 - \beta_k^*)(\theta_t^{F,H(k)} - \theta_t^{F,F(k)}) \tag{23}$$

If $\beta_k^* = 1$, the Foreign investor values the synthetic Treasury exactly the same as a U.S.-issued Treasury bond, which implies that U.S. Treasuries are only convenient due to their currency denomination. In this case, CIP deviations are not informative about the relative convenience yield. If $\beta_k^* < 1$, then there is intrinsic convenience in a U.S. Treasury specifically, beyond its currency denomination. In this case, $CIP_t^{(k)} > 0$ implies that foreigners value Home bonds more than Foreign ones. Moreover, as long as $\beta_k^* \in (0, 1)$, small deviations in the CIP can mask large differences in the relative convenience yield ($CIP_t^{(k)} < \theta_t^{F,H(k)} - \theta_t^{F,F(k)}$).

To estimate this mapping in practice, we follow the steps of Jiang et al. (2021a), but extend the investigation to multiple maturities: specifically 6 months, 1 year and 10 years. In doing so, we provide novel evidence on the relationship between annualised CIP deviations and (annualised) cross-country convenience yields across maturities.

Table 1: Estimated Mapping from CIP to Cross-Country Convenience Yield Across Maturities, β_k^*

Maturity	6-month	1-year	10-year
$\hat{\beta}_k^*$	0.76	0.89	0.85

Note. Table 1 displays the estimated $\hat{\beta}_k^*$ coefficient that maps k-period CIP deviations, defined in equation (21) to the relative convenience yield Foreign investors derive on the k-period Home bond $(\theta^{F,H(k)} - \theta^{F,F(k)})$, according to equation (23), for three maturities k: 6 months, 1 year and 10 years. Details on the estimation procedure can be found in Appendix C.

We present a full explanation of the steps we follow to estimate β_k^*, along with intermediate results, in Appendix C. To do this, we estimate two parameters: (i) a measure of persistence for the k-maturity CIP deviation, $\alpha_1^{(k)}$, which we estimate from a quarterly AR(1) process for each maturity CIP deviation; and (ii) a measure of the marginal effects of innovations to k-maturity CIP deviations on average exchange-rate dynamics of the U.S. dollar vs. the other currencies, $\delta_1^{(k)}$. Our β_k^* estimate then follows from the formula: $\beta_k^* = 1 - \frac{1}{1-\alpha_1^{(k)}} \frac{1}{\delta_1^{(k)}}$.

The resulting estimates for mapping CIP deviations to convenience yields across the term structure are presented in Table 1. The coefficient of 0.89 at the 1-year maturity is comparable to estimate of 0.9 reported in Jiang et al. (2021a). However, our results present novel estimates to suggest that the link between CIP deviations and cross-country convenience varies across the term structure. In particular, our β_k^* estimates are smallest at short maturities, indicating that the greatest intrinsic value from U.S. Treasuries arises at shorter maturities. At longer maturities β_k^* estimates are closer to 1, implying that convenience is linked more to currency, than anything specific about the asset.

Having estimated this mapping, we henceforth define our proxy for cross-country convenience as:

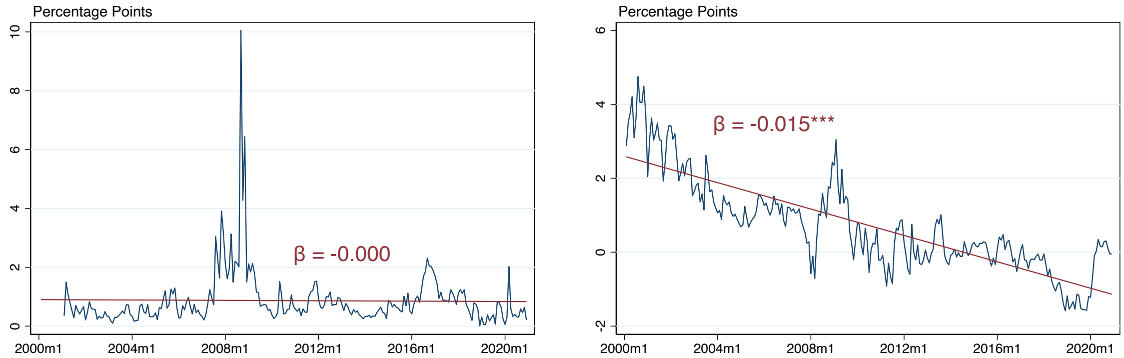
$$\tilde{\theta}_t^{F,H(k)} - \tilde{\theta}_t^{F,F(k)} := \frac{1}{(1-\hat{\beta}_k^*)} CIP_t^{(k)} \tag{24}$$

which follows from equation (23).

Figure 3 plots our resulting annualised measures of cross-country U.S. convenience yields for the 6-month and 10-year tenors, respectively. The time series are constructed by calculating the average cross-country U.S. convenience yield relative to the remaining G.7 currencies. In the figures, a positive value indicates that Foreign investors assign a positive convenience yield to U.S. government bonds relative to Foreign ones and are therefore willing to forgo pecuniary returns (CIP deviations) to hold U.S. Treasuries.

Figure 3: Short- and Long-Run Cross-Country U.S. Treasury Convenience Yields

(a) U.S. 6M Cross-Country CY ($\theta^{F,H(1)} - \theta^{F,F(1)}$) (b) U.S. 10Y Cross-Country CY ($\theta^{F,H(\infty)} - \theta^{F,F(\infty)}$)



Note. Panels 3a and 3b display time series of short-run and long-run cross-country U.S. Treasury convenience yields, respectively ($\theta^{F,H(1)} - \theta^{F,F(1)}$ and $\theta^{F,H(\infty)} - \theta^{F,F(\infty)}$), from 2000:M2 to 2020:M12. In each case, the time series are a cross-sectional average across G.7 currency areas *vis-à-vis* the U.S. The government bond maturities are 6 months (short-run) and 10 years (long-run), respectively. *** signifies that the slopes (β) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are dictated by the underlying CIP-deviation data from Du et al. (2018a).

Comparing figures 3a and 3b reveals notable differences in the evolution of convenience of U.S. bonds at different maturities. While at the 6-month tenor, the average convenience yield is positive and exhibits no significant trend, the 10-year cross-country convenience yield follows a downward trend which started in the early 2000s. Further, since the late 2010s, the long-maturity convenience yield has been negative on average. This implies that foreign investors required compensation to hold a long-term U.S. This is despite the fact that long-run convenience yields were as high as 4pp in the early 2000s.

Looking closer at the series, the largest moves occurred in 2008, during the GFC, when the convenience yield on short-run U.S. Treasuries spiked to nearly 10pp, on average, consistent with there being a 'flight to safety' in the short-run U.S. Treasury market. While long-term U.S. convenience did increase during the GFC, this was far less pronounced.

As such, the convenience yield earned by foreign investors on long-maturity U.S. Treasuries (i.e., the non-pecuniary return) has been significantly eroded over the past two decades.[15] Importantly, about half of the total erosion in long-maturity convenience yields over the past two decades occurred prior to the 2008 GFC, suggesting there is more to this deterioration in

[15]The erosion in U.S. Treasury convenience yields is also present in the cross-section, as shown in Figure 1a, although there is heterogeneity as to the extent. Specifically, it holds for 5 of 6 currency pairs, with the decline in Treasury convenience being largest relative to Australian, Canadian and euro-area (German) government bonds and smaller for Japanese and Swiss government bonds. Interestingly, there is little change in the relative convenience of Treasuries *vis-à-vis* U.K. gilts.

convenience than a post-crisis expansion in Treasury supply (Du et al., 2018a) and tightening of Primary Dealer's constraints, the confluence of which (Du et al., 2022) use to explain the erosion of θ_t^{HH} (see below). Instead, in line with the macroeconomic mechanisms in our model, given little movement in long-run pecuniary dollar returns (Panel 2b), it suggests that the fall in long-maturity Treasury convenience may simply reflect rising long-run risk in the U.S. *vis-à-vis* the rest of the world.

4.2 Measuring Within-Country Convenience Yields

While the focus of our paper is on cross-country convenience, a growing literature has identified a role for convenience within countries that reflects non-pecuniary benefits for domestic investors (e.g., Diamond and Van Tassel, 2022). Moreover, as we emphasise the role of risk, our measures of risk are confounded by these within-country convenience yields (see, e.g., Proposition 2).[16] We account for this in our model through the $\theta_t^{H,H(k)}$ and $\theta_t^{F,F(k)}$ terms, and so require measurable proxies for these quantities.

There are numerous approaches to measuring within-country convenience in the literature. Diamond and Van Tassel (2022), in particular, propose the use of option-implied 'box spreads' to back out risk- and convenience-free rates. However, their approach extends only to the 3-year maturity. To proxy within-country convenience across the term structure, and out to the 10-year maturity specifically, we follow Du et al. (2022) and use interest-rate swap rates at specific maturities to proxy for risk- and convenience-free rates across maturities. As such, we define our proxies for within-country convenience as:

$$\tilde{\theta}_t^{H,H(k)} := r_{irs,t}^{(k)} - r_t^{(k)} \tag{25}$$

$$\tilde{\theta}_t^{F,F(k)} := r_{irs,t}^{*(k)} - r_t^{*(k)} \tag{26}$$

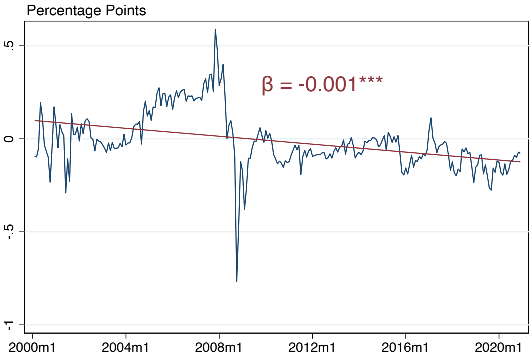
where $r_{irs,t}^{(k)}$ and $r_{irs,t}^{*(k)}$ denote the Home and Foreign log k-period interest-rate-swap rate, respectively. Importantly, due to data limitations, we can only construct Foreign within-country convenience yields for the euro-area. In all, our U.S. ($\tilde{\theta}_t^{H,H(k)}$) and Foreign (euro-area) ($\tilde{\theta}_t^{F,F(k)}$) measures of within-country convenience yields at both the 6-month and 10-year maturities run from 2000:M1 to 2020:M12 (see Appendix B for data sources and details on the calculation).

Panels (a) and (b) of Figure 4 plot our resulting U.S. within-country convenience yield

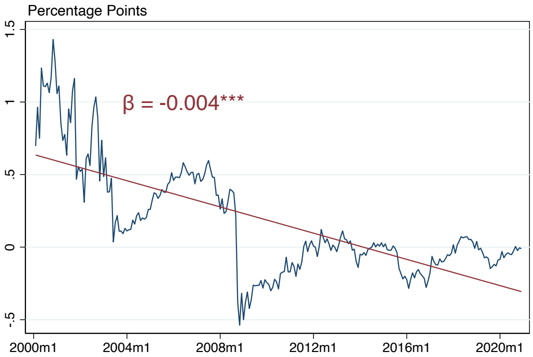
[16] In a related paper, Jiang and Richmond (2023) show that high Sharpe ratios imply either a volatile stochastic discount factor or a high convenience yield. Di Tella et al. (2023) construct zero-beta rates (risk-free, no-convenience yield returns), and identify significantly higher convenience yields.

Figure 4: Short- and Long-Run Within-Country Convenience Yields

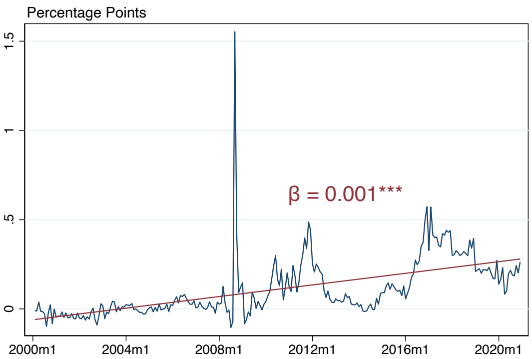
(a) U.S. 6M Within-Country CY ($\theta^{H,H(1)}$)



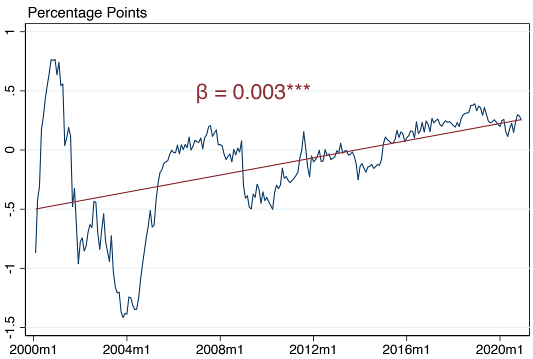
(b) U.S. 10Y Within-Country CY ($\theta^{H,H(\infty)}$)



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



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



Note. Panels 4a and 4b display time series of short-run and long-run within-country U.S. Treasury convenience yields, respectively ($\theta^{H,H(1)}$ and $\theta^{H,H(\infty)}$), from 2000:M2 to 2020:M12. Panels 4c and 4d display time series of short-run and long-run within country German government bond yields ($\theta^{F,F(1)}$ and $\theta^{F,F(\infty)}$), also from 2000:M2 to 2020:M12. In both cases, the yields are calculated as the difference between interest-rate swap rates and the corresponding-maturity zero-coupon government bond yield, as described in equation (25). *** signifies that the slopes (βs) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are chosen to match the CIP-deviation data.

series for the 6-month and 10-year tenors. As with the cross-country measures, there are large differences in dynamics for the short- and long-run tenors. The short-run measure, in panel 4a, hovers around zero. It was rising in the run-up to the GFC, reflecting greater convenience. However, post-GFC, it has turned marginally negative. In contrast, the long-run measure, in panel 4b displays a notable downward trend, albeit less stark than its cross-country counterpart, over the sample period. This suggests that the decline in Foreign investors' perceptions as to the 'specialness' of long-maturity U.S. Treasuries have, to some extent, been matched by domestic investors in the U.S. Interestingly, our within-country convenience yield measure plummets during the GFC, in sharp contrast to our cross-country measure. As pointed out by Du et al. (2022), this is likely the result of a severe tightening in dealer's balance-sheet constraints. Different to these authors, we focus on trends in both cross-country and within-country convenience yields over time, rather than a regime-switch during the GFC, and are particularly concerned with cross-country convenience.

Panels (c) and (d) of Figure 4 plot the analogous euro area within-country convenience yields for the 6-month and 10-year tenors. At both tenors, these euro area convenience yields display the opposite dynamics to the U.S. within-country convenience yields over our sample period, despite the fact that these U.S. and euro area measures are constructed using no common asset prices. Specifically, at both short- and long-maturities, euro area within-country convenience yields have been rising over time, with the trend again more pronounced for the 10-year tenor. As a result, the difference in U.S. and euro area within-country convenience yields ($\theta^{H,H(k)} - \theta^{F,F(k)}$), which will play a role in our measurement of U.S. relative SDF risk (see below), have experienced a pronounced fall over the past 20 years.

Finally, to measure holding-period convenience yields both within and across countries, $\theta_t^{i,j(k)} - \mathbb{E}_t[\theta_{t+1}^{i,j(k)}]$, we adopt the same approach as for exchange rates. Specifically, we proxy *ex ante* holding-period convenience with *ex post* changes (i.e., $\theta_t^{i,j(k)} - \theta_{t+1}^{i,j(k)}$).

4.3 Measuring Relative SDF Risk with Equity Premia

A key contribution of our paper is to tie convenience to a measurable proxy for risk. To measure risk, we adopt numerous approaches. First, we calculate ex-post equity returns using equity price indices for each country in local currency terms from *MSCI*. Each index aggregates information on equity prices for the country's large and mid-cap public firms, thus providing information on the realised return $R_{t,t+1}^{g(*)}$ in our model. This approach mirrors that of Hansen and Jagannathan (1991) and Alvarez and Jermann (2005), amongst others, who also use equity prices to derive measures of total and permanent risk in an economy, using their bounds.[17] Relative to them, we provide a novel adjustment of the risk bounds by accounting for the presence of convenience yields on domestic bonds, as derived in Proposition 2. Importantly, the use of equity index returns helps maximise the right-hand side of each bound.

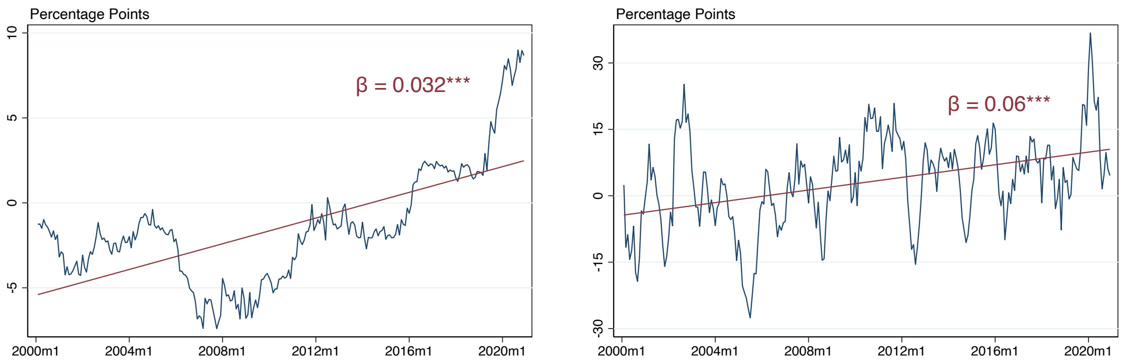
Second, to strengthen our results, we use *ex ante* measures of equity risk premia, as opposed to using realised equity returns that may conflate risk with reward, especially in smaller samples (see Farhi and Gourio, 2018, for a discussion).

To construct *ex ante* measures of equity risk premia, we follow a common approach in

[17] See Jiang and Richmond (2023) for a related generalisation of the Hansen-Jagannathan bounds.

Figure 5: *Ex Ante* Equity Risk Premia and *Ex Post* Equity Returns for the U.S. Relative to Other G.7 Markets

(a) U.S. Relative Expected Equity Risk Premium (b) U.S. Relative Realised Equity Returns



Note. Panels 5a and 5b display time series of U.S. minus Foreign relative *ex ante* equity risk premium constructed estimated via the Gordon growth formula in equation (27) and relative *realised* equity returns, respectively, from 2000:M2 to 2020:M12. The equity risk premia are calculated using dividend-price ratio data for each country, the 10-year nominal government bond yield adjusted with *Consensus Economics* inflation expectations to construct a 'real' risk-free rate, and the average dividend growth in the 10 years prior to a given data to proxy for expected dividend growth. *** signifies that the slopes (β) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are chosen to match the CIP-deviation data.

our baseline and construct equity risk premia using a variant of the Gordon growth model for equities, building on Gordon (1962). Specifically, we proxy the (log) equity risk premium according to the following formula:

$$\mathbb{E}_t \log \left[\frac{\widetilde{R^g_{t,t+1}}}{R_t} \right] := \frac{D_t}{P_t} - r_t + \pi^e_t + g^e_t \tag{27}$$

where D_t/P_t denotes the dividend-price ratio, r_t is the nominal risk-free rate that is converted into real form with a measure of expected inflation π^e_t in a country, and g^e_t represents the expected growth rate of dividends at time t. We use an analogous formula for the foreign country.

We extend Farhi and Gourio (2018) and construct the proxy defined by equation (27) for all G.7 equity markets.[18] To do so, we use data on dividend-price ratios and equity price indices 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).[19] In our baseline, we proxy for

[18] Farhi and Gourio (2018) construct the proxy defined by equation (27) for the U.S., noting that it is preferred as a proxy for risk in equity markets over realised returns, since the latter are especially volatile.

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

expected future dividend growth using the average annual dividend growth in the ten years prior to time t (our findings are similar if we use five- or one-year growth). Finally, we proxy the real risk-free rate using the nominal 10-year zero-coupon bond yield adjusted with inflation expectations from *Consensus Economics*.

As shown in Figure D.3 in Appendix D, our *ex ante* measure of the equity risk premium for the U.S. trends upwards over the past 20 years, as do realised returns, which is consistent with the literature.[20] However, our analysis is centred on cross-country differences in expected equity risk premia. Focusing on Figure 5 highlights a stronger result: the expected equity risk premium in the U.S. *relative* to the other G.7 markets has, on average, increased over the last two decades. The left panel show that relative premia have grown from near -5pp in the early 2000s to close to +10pp around 2020. According to our estimates, this trend increase in U.S. equity risk *vis-à-vis* other G.7 countries is responsible for more than *half* of the rise in realised relative U.S. equity returns, shown in panel 5b, and is particularly marked due to its much lower volatility.

Using these *expected* equity risk premia series, which help maximise the right-hand side of each bound, we measure SDF risk by assuming the lower bounds summarised by expressions (15) and (16) in Proposition 2 hold with equality. As such, we define our empirical proxy for overall risk, $\mathcal{L}_t(M_{t,t+1})$, which we denote by $\tilde{\mathcal{L}}_t(M_{t,t+1})$, as:

$$\tilde{\mathcal{L}}_t(M_{t,t+1}) := \mathbb{E}_t \log \left[\widetilde{\frac{R^g_{t,t+1}}{R_t}} \right] - \tilde{\theta}_t^{H,H(1)}. \tag{28}$$

Similarly, we define our empirical proxy for permanent risk, $\mathcal{L}_t(M^{\mathbb{P}}_{t,t+1})$, which we denote by $\tilde{\mathcal{L}}_t(M^{\mathbb{P}}_{t,t+1})$, as:

$$\tilde{\mathcal{L}}_t \left(M^{\mathbb{P}}_{t,t+1}\right) := \mathbb{E}_t \log \left[\widetilde{\frac{R^g_{t,t+1}}{R_t}} \right] - rx^{(\infty)}_{t+1} - \left(\tilde{\theta}_t^{H,H(\infty)} - \tilde{\theta}_{t+1}^{H,H(\infty)} \right) \tag{29}$$

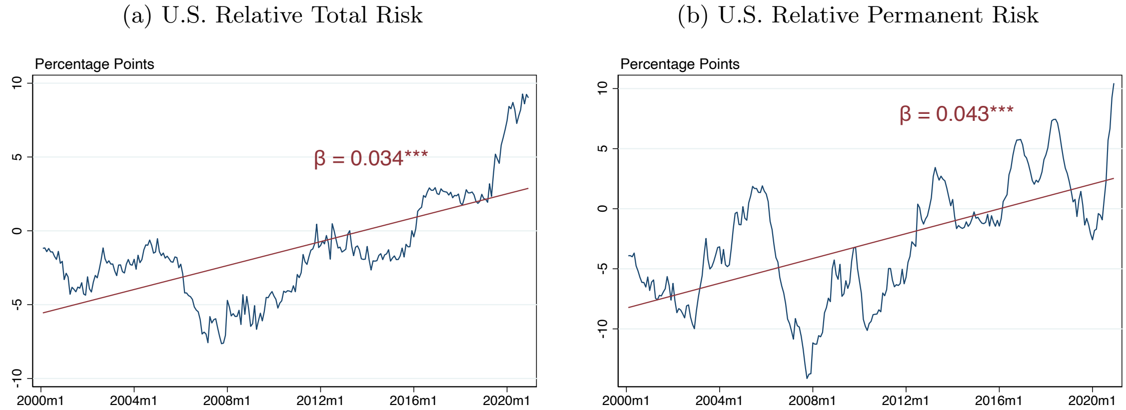
where we have proxied the *ex ante* bond risk premium $\mathbb{E}_t rx^{(\infty)}_{t+1}$ and expected convenience yield $\mathbb{E}_t \tilde{\theta}^{H,H(\infty)}_{t+1}$ with their ex-post realisations ($rx^{(\infty)}_{t+1}$ and $\tilde{\theta}^{H,H(\infty)}_{t+1}$, respectively).[21]

While Figure D.5 in Appendix D shows that U.S. *ex ante* overall risk, and, in particular,

[20] Farhi and Gourio (2018) report a number of alternative *ex ante* measures for the U.S. equity risk premium too. All support their conclusion that perceived risk in U.S. equity markets has risen in past decades. There is also supportive survey data.

[21] For now, we smooth our $rx^{(\infty)}_{t+1}$ using a 1 year trailing moving average of the underlying series. We plan to use a term-structure model moving forward.

Figure 6: *Ex Ante* Total and Permanent Risk for the U.S. Relative to Other G.7 Markets



Note. Panels 6a and 6b display time series of our proxies for U.S. minus Foreign relative *ex ante* total and permanent risk ($\tilde{\mathcal{L}}_t(M_{t,t+1}) - \tilde{\mathcal{L}}_t(M^*_{t,t+1})$ and $\tilde{\mathcal{L}}_t\left(M^{\mathbb{P}}_{t,t+1}\right) - \tilde{\mathcal{L}}_t\left(M^{*\mathbb{P}}_{t,t+1}\right)$), respectively, as calculated according to equations (28) and (29) from 2000:M2 to 2020:M12. In both cases, the time series are a cross-sectional average across G.7 currency areas *vis-à-vis* the U.S. *** signifies that the slopes (βs) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are chosen to match the CIP-deviation data.

permanent risk have risen considerably over the past 20 years, our analysis focuses on cross-country differences in expected risk. Figure 6 highlights the stronger result: U.S. overall risk and permanent risk have risen over the past two decades *relative* to G.7 currencies, on average. In terms of U.S. relative overall risk, Panel 6a highlights that the trend increase is modestly larger than the increase in relative equity premia from Panel 5a, since U.S. 6-month within country convenience yields fell slightly while Euro Area ones—which we use as a proxy for the rest of the G.7, increased slightly (Panels 4a and 4c). Turning now to Panel 6b, that U.S. relative permanent risk has risen more than relative total risk reflects the fact the residual U.S. *transitory* risk:

$$\tilde{\mathcal{L}}_t(M^{SR}_{t,t+1}) := \tilde{\mathcal{L}}_t(M_{t,t+1}) - \tilde{\mathcal{L}}_t\left(M^{\mathbb{P}}_{t,t+1}\right) = rx^{(\infty)}_{t+1} + \left(\tilde{\theta}^{H,H(\infty)}_t - \tilde{\theta}^{H,H(\infty)}_{t+1}\right) - \tilde{\theta}^{H,H(1)}_t, \tag{30}$$

has fallen (modestly) relative to G.7 economies (Figure D.6). The additional increase in U.S. relative permanent risk compared to relative overall risk reflects primarily a trend decrease in the U.S. term premium *vis-à-vis* G.7 economies: $rx^{(\infty)}_{t+1} - rx^{*(\infty)}_{t+1} \downarrow$ (Figure D.7). The mild increase in the relative deviations from the convenience-yield expectations hypothesis between the U.S. and the Euro-Area partially offsets this trend (Figure D.8).

Although our results for the *ex ante* measure are robust to alternatively using realised returns, further work is required to establish the robustness of our *ex ante* measure. In an excellent discussion, Papanikolaou (2018) shows that the expected equity risk premium measure in Farhi and Gourio (2018), closely related to our measure, actually generates insights consistent with the equity premium implied by CFO survey data provided by Duke survey. Moreover, the expected equity risk premium measure is very closely correlated with the long-run uncertainty estimated in Schorfheide, Song, and Yaron (2018), and is even similar to the economic policy uncertainty measure in Baker, Bloom, and Davis (2016).

4.4 Empirical Specification: Combining Theory and Measurement

Using these measures of exchange-rate risk, within- and cross-country convenience yields, and relative SDF risk, we can take the conditional equilibrium results outlined in Propositions 1 (short-run) and 4 (long-run) to the data.

We test the short-run Proposition 1 in Section using the following panel regression specification for the U.S. *vis-à-vis* other G.7 markets:

Testing Proposition 1 (Short-Run SDF Volatility, FX Risk and Convenience)

$$\tilde{\theta}^{F,H(6M)}_{i,t} - \tilde{\theta}^{F,F(6M)}_{i,t} = \beta_0 + \beta_1 \left[\tilde{\mathcal{L}}_t\left(M_{i,t,t+1}\right) - \tilde{\mathcal{L}}_t(M^*_{i,t,t+1})\right] + \beta_2 rx^{FX}_{t+1} + f_i + \varepsilon_{i,t} \tag{31}$$

$$rx^{FX}_{t+1} = \gamma_0 + \gamma_1 \left[\tilde{\mathcal{L}}_t\left(M_{i,t,t+1}\right) - \tilde{\mathcal{L}}_t\left(M^*_{i,t,t+1}\right)\right] + \gamma_2 \left[\tilde{\theta}^{F,H(6M)}_{i,t} - \tilde{\theta}^{F,F(6M)}_{i,t}\right] + f_i + \varepsilon_{i,t} \tag{32}$$

where $\tilde{\theta}^{F,H(6M)}_{i,t} - \tilde{\theta}^{F,F(6M)}_{i,t} = CIP^{(6M)}_t/(1-\beta^*_{6M})$ from (24), $\tilde{\mathcal{L}}_t\left(M_{t,t+1}\right) = \mathbb{E}_t \log\left[\frac{\widetilde{R^g_{t,t+1}}}{R_t}\right] - \tilde{\theta}^{H,H(6M)}_t$ from (28) and $\tilde{\theta}^{H,H(6M)}_t = r^{(6M)}_{irs,t} - r^{(6M)}_t$ from (25).

For regression (31), we expect that higher U.S. risk lowers short-maturity convenience yields ($\beta_1 < 0$) and that higher pecuniary returns on Foreign bonds lower non-pecuniary returns on Foreign bonds—since these are substitutes—and therefore raise them on U.S. bonds $\beta_2 > 0$. Similarly, in regression (32), we expect that higher U.S. risk raises pecuniary returns on Foreign bonds $\gamma_1 > 0$ while higher non-pecuniary returns on Home bonds (lower ones on Foreign bonds) raise pecuniary returns on Foreign bonds $\gamma_2 > 0$.

Given the rise in U.S. relative permanent risk that we document (Figure 6b) along with the fall in long-maturity Treasury convenience (Figure 3b), we are particularly interested in testing the long-run relationship. To do so, in Section 5.2.1, we run the following panel regression specifications for the U.S. *vis-à-vis* other G.7 markets:[22]

Testing Proposition 4 (Long-Run SDF Volatility, FX Risk and Convenience)

$$\tilde{\theta}_{i,t}^{F,H(10Y)} - \tilde{\theta}_{i,t}^{F,F(10Y)} = \beta_0 + \beta_1\left[\tilde{\mathcal{L}}_t\left(M_{i,t,t+1}^{\mathbb{P}}\right) - \tilde{\mathcal{L}}_t(M_{i,t,t+1}^{\mathbb{P}*})\right] + \beta_2 rx_{i,t+1}^{(10Y),CT} + \beta_3\left[\tilde{\theta}_{i,t+1}^{F,H(10Y)} - \tilde{\theta}_{i,t+1}^{F,F(10Y)}\right] + f_i + \varepsilon_{i,t} \tag{33}$$

$$rx_{i,t+1}^{(10Y),CT} = \gamma_0 + \gamma_1\left[\tilde{\mathcal{L}}_t\left(M_{i,t,t+1}^{\mathbb{P}}\right) - \tilde{\mathcal{L}}_t\left(M_{i,t,t+1}^{\mathbb{P}*}\right)\right] + \gamma_2\left[\tilde{\theta}_{i,t}^{F,H(10Y)} - \tilde{\theta}_{i,t}^{F,F(10Y)}\right] + \gamma_3\left[\tilde{\theta}_{i,t+1}^{F,H(10Y)} - \tilde{\theta}_{i,t+1}^{F,F(10Y)}\right] + f_i + \varepsilon_{i,t} \tag{34}$$

where $\tilde{\theta}_{i,t}^{F,H(10Y)} - \tilde{\theta}_{i,t}^{F,F(10Y)} = CIP_t^{(10Y)}/(1-\beta_{10Y}^*)$ from (24), $\tilde{\mathcal{L}}_t\left(M_{t,t+1}^{\mathbb{P}}\right) = \mathbb{E}_t \log\left[\frac{\widetilde{R_{t,t+1}^g}}{R_t}\right] - rx_{t+1}^{(\infty)} - \left(\tilde{\theta}_t^{H,H(\infty)} - \tilde{\theta}_{t+1}^{H,H(\infty)}\right)$ from (29) and $\tilde{\theta}_t^{H,H(10Y)} = r_{irs,t}^{(10Y)} - r_t^{(10Y)}$ from (25).

Focusing on regression (33), our primary interest is the regression coefficient associated with U.S. relative permanent risk β_1. In line with Proposition 4, we expect that $\beta_1 < 0$, that is, if U.S. permanent risk rises relative to other G.7 currency areas, then the relative convenience yield foreign investors derive from holding long-maturity U.S. Treasuries, controlling for this convenience in the future, must fall.[23] We will also decompose the U.S. relative permanent risk measure into its constituents—specifically both (a) U.S. and Foreign permanent risk and (b) relative equity premia, term premia or within-country convenience—and test the effect of these components on cross-border U.S. convenience yields. In addition, we expect that $\beta_2 > 0$ because pecuniary and non-pecuniary returns are substitutes: if U.S. investors' pecuniary long-run carry-trade returns increase, then the non-pecuniary relative convenience yield earned on the Foreign bond should fall $(\tilde{\theta}_{i,t}^{H,F(10Y)} - \tilde{\theta}_{i,t}^{H,H(10Y)}) \downarrow \implies (\tilde{\theta}_{i,t}^{F,H(10Y)} - \tilde{\theta}_{i,t}^{F,F(10Y)}) \uparrow$. Finally, all else equal, we expect that if non-pecuniary returns rise today, they should also rise in the future: $\beta_3 > 0$.

Turning now to regression (34), our primary interest is again the regression coefficient associated with U.S. relative permanent risk γ_1. We expect that $\gamma_1 > 0$, that is, if U.S. permanent risk rises, then the pecuniary return U.S. investors' earn from a cross-border carry trade position must rise as well. We can again decompose this effect into its constituent parts. Next, since

[22] We do not test the unconditional relationship between long-run relative risk, convenience and exchange-rate risk detailed in Proposition 3 since, in practice, the long-run exchange-rate risk premium $\lim_{k\to\infty}\frac{1}{k}\mathbb{E}\left[rx_{t+k}^{FX}\right]$ suffers from having too many overlapping observations and therefore a short sample. For this reason, we focus our empirical analysis on testing Proposition 4.

[23] By the symmetry in our model, this is because the relative convenience yield U.S. investors earn from Foreign government bonds must increase.

pecuniary and non-pecuniary returns are substitutes, we expect that $\beta_2 > 0$ while $\beta_3 < 0$. This is because the one-period holding-period convenience yield on a long-maturity carry trade earns the relative convenience of Foreign bond at time t and sells its convenience at $t+1$. Thus, a lower non-pecuniary return for the U.S. investor on the Foreign bond $(\tilde{\theta}_{i,t}^{H,F(10Y)} - \tilde{\theta}_{i,t}^{H,H(10Y)}) \downarrow$ and $\tilde{\theta}_{i,t+1}^{H,F(10Y)} - \tilde{\theta}_{i,t+1}^{H,H(10Y)}) \uparrow \implies (\tilde{\theta}_{i,t}^{F,H(10Y)} - \tilde{\theta}_{i,t}^{F,F(10Y)}) \uparrow$ and $\tilde{\theta}_{i,t+1}^{H,F(10Y)} - \tilde{\theta}_{i,t+1}^{H,H(10Y)}) \downarrow$ implies they must be compensated with greater pecuniary returns $rx_{i,t+1}^{(10Y),CT} \uparrow$.

5 Empirical Results

In this section, we present our regression results for our tests to Propositions 1 and 4. We discuss how these results help to identify where the 'specialness' of the U.S. (dollar) lies and the extent to which differences in risk can explain cross-country convenience yields across maturities.

5.1 Short-Run Relationships

Table 2 presents our coefficient estimates from regressions (31) (columns (1)-(4)) and (32) (columns (5)-(8)). Our key result is shown consistently across columns (1)-(4), namely: changes in measures relative total risk are robustly negatively associated with short-run cross-country convenience yields. Column (1) shows this for our proxy for relative total risk, which is constructed according to the definition in equation (28) using our measure *ex ante* equity risk premia. Its coefficient indicates that a 1pp increase in that quantity is associated with a 3bp decline in short-horizon convenience yields. Column (2) demonstrates that this finding is robust to separating out the two components of relative total risk—the equity risk premia and short-run within-country convenience yield differences—and column (4) shows that this significant negative association also holds when a measure of *ex post* relative equity returns is used in the proxy for total risk. Additionally, the coefficients in column (3) reveal interesting differences between U.S. and rest-of-the-world risk. In line with our theory, short-run U.S. convenience is negatively associated with U.S. total risk, and negatively with rest-of-the-world total risk. While the coefficients are of similar magnitude, it is the rest-of-the-world coefficient that is statistically significant. This suggests that higher total risk abroad implies leads to an increase in the convenience yield that Foreign investors assign to U.S. bonds relative to their own, consistent with 'flight-to-safety' type dynamics in risk-on events.

While columns (1)-(4) reveal a significant association between relative total risk and cross-country convenience yields in the short run, the coefficients on pecuniary currency returns are

Table 2: Regressions for Short-Run (6-month) Cross-Country Convenience, Total Risk and Currency Returns

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Dep. Var.: $\theta_t^{F,H(1)} - \theta_t^{F,F(1)}$				Dep. Var.: rx_{t+1}^{FX}			
Rel. Tot. Risk	-0.03**				-0.29			
	(0.01)				(0.19)			
Rel. ERP		-0.05**				-0.31*		
		(0.02)				(0.18)		
U.S. ERP			-0.01				-0.27*	
			(0.01)				(0.14)	
R.o.W. ERP			0.01***				-0.01	
			(0.00)				(0.06)	
Rel. Tot. Risk				-0.01*				0.30***
(*Ex Post*)				(0.00)				(0.06)
$\theta_t^{H,H(1)} - \theta_t^{F,F(1)}$		-1.27	-1.06			-0.93	1.17	
		(0.99)	(0.97)			(4.18)	(4.42)	
$\theta_t^{F,H(1)} - \theta_t^{F,F(1)}$					-0.16	-0.25	0.16	0.39
					(0.88)	(0.97)	(1.00)	(0.87)
rx_{t+1}^{FX}	-0.00	-0.00	0.00	0.00				
	(0.00)	(0.00)	(0.01)	(0.01)				
Constant	0.83***	0.64***	0.70***	0.88***	0.37	0.26	2.60	-0.70
	(0.09)	(0.13)	(0.11)	(0.11)	(1.35)	(1.45)	(1.89)	(1.35)
Observations	1,531	1,531	1,531	1,531	1,531	1,531	1,531	1,531
# Countries	6	6	6	6	6	6	6	6
Country FE	YES	YES	YES	YES	YES	YES	YES	YES
Within R^2	0.0184	0.108	0.0714	0.00776	0.0114	0.0119	0.0330	0.103

insignificant and close to zero. That is, changes in pecuniary returns to the U.S. dollar do not appear to be strongly related to non-pecuniary convenience.

Columns (5)-(8) look deeper at the drivers of pecuniary currency returns, reporting the coefficient estimates from the return-predictability regression (32). Interestingly, in this baseline setting, which focuses on 6-month returns over our whole sample, we find that short-run convenience yields today do not provide significant predictive power for short-run currency returns. The coefficient on $\theta_t^{F,H(1)} - \theta_t^{F,F(1)}$ is negative for all specifications in Tables 2. While potentially surprising, this is not inconsistent with the findings in Jiang et al. (2021a), who also find limited evidence of return predictability from convenience yields at the 6-month horizon over long samples. Instead, their results indicate that return predictability is more significant at the 1-year tenor, and for the post-GFC period especially. We verify these findings in Tables D.2 and D.3 in the Appendix.[24] More generally, consistent with our findings in this paper,

[24]In addition to investigating return predictability at different short-horizon maturities, Jiang et al. (2021a) also appeal to lagged adjustment in convenience yields to attain significant results for contemporaneous exchange-rate determination, the focus of Engel and Wu (2018) too, an exercise which is outside of the scope of our study.

Lloyd and Marin (2020) find that return predictability by convenience yields is weak at short horizons but becomes stronger at longer horizons (out to 10-years).

Our coefficient estimates for relative total risk in columns (5)-(8) reveal an interesting but puzzling finding. In contrast to our prediction that that higher U.S. risk raises pecuniary returns on Foreign bonds $\gamma_1 > 0$, our point estimates, using *ex ante* measures of equity risk premia in columns (5)-(7) specifically, are significantly negative. Only with *ex post* measures of relative equity returns (column (8)) do we find a significant and positive coefficient. In many ways, this puzzling result need not be a surprise when considering the fact that it is at short horizons like these where the UIP puzzles bites the most. To investigate this further, we present estimates of the comovement between *ex ante* and *ex post* measures of equity risk and relative interest rates. Table D.1 in the Appendix reports these estimates. Strikingly , we find that the relationship between relative yields and *ex ante* and *ex post* measures has opposing signs. *Ex post* measures are negatively related to relative returns (i.e., procyclical with respect to the business cycle when, intuitively, risk rises in downturns), so can rationalise the UIP puzzle in the spirit of Verdelhan (2010). However, relative yields are strongly positively related to *ex ante* measures of risk, so countercyclical with respect to the business cycle. To the best of our knowledge, we are the first to emphasise this inconsistency of the data. In further work, we plan to assess the extent to which this finding also remains when we consider *ex ante* measures of currency returns.

Our results indicate that relative risk is strongly associated with non-pecuniary convenience in the short-run, with investor perceptions of higher relative U.S. risk associated with reductions in U.S. convenience, and *vice versa*. We also identify a relationship between risk and pecuniary currency returns —albeit dependent on the measure of risk.

5.2 Long-Run Relationships

We now turn to assessing the long-run relationships predicted by our model in Propositions 3 and 4. While we do not take Proposition 3 to the data formally, owing to the limited number of non-overlapping observations in our sample for long-run currency returns $rx^{FX}_{t+\infty}$, we do have a first look at the unconditional relationship predicted by the proposition in Table 3. Here, we plot the coefficient estimates from bivariate regressions of long-run cross-country convenience yields $\theta_t^{F,H(\infty)} - \theta_t^{F,F(\infty)}$ on measures of relative permanent risk.

We run this regression motivated by the wealth of evidence that long-run currency returns are near-zero on average, which implies that the UIP hypothesis cannot be rejected in the long

Table 3: Regressions for Long-Run Cross-Country Convenience, Permanent Risk and Carry Trade Returns

	(1)	(2)	(3)	(4)
		Dep. Var.: $\theta_t^{F,H(\infty)} - \theta_t^{F,F(\infty)}$		
	Eq. Risk Prem.	Eq. Return	Eq. net T.P.	$\mathbb{P}$ Risk
	Ex Ante	*Ex Post*	*Ex Ante*	*Ex Ante*
Rel. Risk Measure	-0.07***	-0.03***	-0.05***	-0.06***
	(0.02)	(0.01)	(0.01)	(0.01)
Observations	1,657	1,657	1,657	1,544
# Countries	6	6	6	6
Country FE	YES	YES	YES	YES
Within R^2	0.0439	0.0594	0.0318	0.0554
Pedroni Panel Coint. Test	-5.43***	-5.36***	-4.86***	-4.46***

run (Chinn and Meredith, 2005; Chinn and Quayyum, 2012). As such, this bivariate regression helps to shed light on the extent to which long-run convenience and permanent risk interact, taking the long-run UIP hypothesis as given. Moreover, to the extent that there may be worries about non-stationarity in the underlying series for permanent risk and long-run convenience that could undermine inference from our subsequent long-run regressions, we carry out panel cointegration and unit-root tests for the series used in these bivariate regressions. While we cannot reject the null hypothesis of a unit root in some of our permanent risk measures, using the Im, Pesaran, and Shin (2003) testing approach, we strongly reject the null hypothesis of no cointegration between permanent risk and long-run convenience yields on a country-by-country basis using the Pedroni (1999, 2004) t test.

For the various measures of risk used, the coefficient estimates in Table 3 are robustly negative. A 1pp increase in relative U.S. permanent risk is associated with a 3-7bp decrease in relative U.S. convenience. In the context of the 15pp increase in relative U.S. permanent risk implied by our baseline measure in Figure 6b, this suggests that this trend explains 45-105bp of the near-3pp decline in long-run U.S. convenience shown in Figure 3b.

These estimates provide initial statistical evidence to support our conclusion that the decline in long-term U.S. convenience and the rise in relative U.S. permanent risk, that we document in Section 4, are two sides of the same coin. However, to test this further, we focus on Proposition 4 in the remainder of this section, analysing the relationships between carry-trade returns on long-term bonds, long-run convenience and permanent risk.

5.2.1 Carry-Trade Returns on Long-Term Bonds

Table 4 presents our baseline results for the equilibrium relationships implied by our model.[25] Columns (1)-(3) focus on the drivers of long-run convenience using regression (33). Consistent with our theory, the coefficient estimates on different measures of relative permanent risk reveal a significantly negative relationship between permanent risk and long-run convenience. Controlling for the other factors that operate in equilibrium, our coefficients indicate that if investor perceptions of U.S. permanent risk rises by 1pp relative to other G.7 currencies, then the relative convenience yield that Foreign investors derive from holding long-maturities U.S. Treasuries falls by 1bp. Moreover, the coefficients in column (2), which separate out the influence of U.S. permanent risk and rest-of-the-world permanent risk, reveal a particularly strong quantitative role for U.S. permanent risk in explaining this—with the coefficient of that roughly three-times the magnitude of the rest-of-the-world coefficient.

In addition, the coefficient estimates in columns (1)-(3) reveal a significant *ceteris paribus* association between carry-trade returns and convenience. The key driver of this, however, is that fact that carry-trade returns from today to tomorrow are strongly associated with changes in convenience yields over the same period, which we control for using forward lags (i.e., $\theta_{t+1}^{F,H(\infty)} - \theta_{t+1}^{F,F(\infty)}$. Interestingly, however, we find that these forward lags on convenience have coefficients that are significantly less than unity. This suggests that there is some decay in the convenience an investor derives from a long-term bond over its holding period, with more convenience earned at time of purchase—potentially because of the role of the long-term bond as a collateral object. We defer a fuller investigation of the horizons over which convenience is earned on long-maturity bonds for future work.

Columns (4)-(6) of Table 4 turn to the drivers of carry-trade returns on long-term bonds, using regression (34). Across all three specifications, we find a robustly positive and significant relationship between the convenience yields Foreign investors earn on U.S. long-term bonds and carry-trade returns. Our point estimates indicate that a 1pp decrease in convenience is associated with a 3.5-4.4pp decrease in carry-trade returns. On their own, this finding extends the results of Engel and Wu (2018) and Jiang et al. (2020) to long maturities, and controlling for risk differentials.

However, through the lens of our model, we can also test the extent to which increases in risk offset this—such that our results are consistent with the fairly stable level of long-run carry trade (pecuniary returns) over our sample. Across specifications, our estimates reveal

[25] We present robustness analysis across sub-samples in Table D.4 in the Appendix.

Table 4: Regressions for Long-Run Cross-Country Convenience, Permanent Risk and Carry Trade Returns

	(1)	(2)	(3)	(4)	(5)	(6)
	Dep. Var.: $\theta_t^{F,H(\infty)}$			Dep. Var.: $rx_{t+1}^{(\infty),CT}$		
Rel. $\mathbb{P}$ Risk	-0.01*			0.18**		
	(0.01)			(0.07)		
U.S. ERP		-0.03***			-0.04	
		(0.01)			(0.13)	
R.o.W. ERP		0.01**			-0.02	
		(0.00)			(0.06)	
Rel. T.P.		-0.03*			-0.35**	
		(0.02)			(0.17)	
Rel. $\mathbb{P}$ Risk			-0.01***			0.30***
(*Ex Post*)			(0.00)			(0.04)
$\theta_t^{H,H(\infty)} - \theta_t^{F,F(\infty)}$		0.20			-6.51	
		(0.20)			(5.49)	
$\theta_t^{F,H(\infty)} - \theta_t^{F,F(\infty)}$				4.34***	3.50***	4.44***
				(0.73)	(0.64)	(0.71)
$\theta_{t+1}^{H,H(\infty)} - \theta_{t+1}^{F,F(\infty)}$		0.12			11.59**	
		(0.23)			(5.72)	
$\theta_{t+1}^{F,H(\infty)} - \theta_{t+1}^{F,F(\infty)}$	0.84***	0.76***	0.83***	-4.44***	-4.60***	-3.91***
	(0.05)	(0.06)	(0.05)	(0.66)	(0.53)	(0.66)
$rx_{t+1}^{(\infty),CT}$	0.03***	0.02***	0.03***			
	(0.00)	(0.00)	(0.00)			
Constant	0.16***	0.36***	0.23***	-0.40	-0.38	-1.87*
	(0.06)	(0.08)	(0.06)	(1.18)	(1.94)	(1.07)
Observations	1,508	1,508	1,508	1,508	1,508	1,508
# Countries	6	6	6	6	6	6
Country FE	YES	YES	YES	YES	YES	YES
Within R^2	0.682	0.706	0.687	0.139	0.185	0.235

that (components) of relative total risk are also robustly and positively associated with carry-trade returns.[26] The coefficients indicate that a 1pp increase in relative U.S. permanent risk is associated with a 0.18-0.30pp increase in carry-trade returns. Hence, higher relative permanent risk leads to lower convenience yields but the effect on carry trade returns is at least partly offset by the direct effect of relative permanent risk on carry trade returns. Such an offsetting phenomenon is not apparent in the short run, where risk seems to drive pecuniary returns in a reinforcing manner directly and indirectly.

[26] In column (5), the negative coefficient on relative bond term premia should be interpreted as yielding a positive relationship, given the definition of permanent risk in equation (29).

6 Conclusion

In this paper, we have shown that extending the basic two-country no-arbitrage framework to allow for a rich term-structure of bonds and convenience yields, alongside equities, can yield new insights into the links between U.S. (dollar) 'specialness' across markets and across maturities. On the theoretical front, we have shown that convenience-yield differentials on long-maturity bonds can admit differences in countries' permanent risk (permanent SDF volatility), even though long-horizon currency returns are near-zero on average. On the empirical front, we have uncovered relationships between convenience yields and risk metrics across horizons. In particular, our key result is that investor perceptions of U.S. permanent risk relative to other G.7 economies, are associated with lower levels of the convenience yields that Foreign investors earn when holding long-term U.S. Treasuries. In short: over long horizons, risk and convenience across countries are two sides of the same coin.

By providing a framework to assess how different elements U.S. (dollar) 'specialness' are connected across markets and horizons, our findings provide novel insights into the nuances in the structure of the International Monetary System. However, our model is not equipped to identify the precise mechanisms through which a rise in perceived relative U.S. risk can influence the extent to which Foreign investors perceive the relative convenience of U.S. assets. Our findings could be consistent with models where higher U.S. risk undermines the possibility to use U.S. assets as collateral internationally. Additionally, our framework does not identify the reasons why relative U.S. permanent risk appears to have risen. Since the dot-com crash in 2001, and more so since the GFC, investors appear to have started pricing in a different amount of permanent risk in the U.S. One possibility suggested by Papanikolaou (2018) is that the rise in fragility may be due to a higher share of intangible capital in recent decades. Alternatively, our findings are also consistent with the 'scarring of beliefs', discussed in Kozlowski, Veldkamp, and Venkateswaran (2019).

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Appendix

A Proofs and Derivations

A.1 Proof to Proposition 1

Taking expectations of the exchange-rate process (9)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 (1) and (2), respectively, we can write:

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

which yields the result by rearranging and using the definition of the *ex post* excess currency return (11). □

A.2 Proof to Lemma 1

Start by rewriting the Home Euler equation for the Home k-period bond (1):

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

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):

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

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

where line 1 is a definition, line 2 uses definition for holding-period returns, line 3 uses the equalities derived prior to this, line 4 rearranges, line 5 multiplies and divides by β^{t+k}, line 6 uses the Alvarez and Jermann (2005) definition 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 (14). □

A.3 Proof to Proposition 2

Equation (6) follows from no-arbitrage, and has the log form:

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

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:

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

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

(15).

Next, to derive a decomposition of total risk, start with the definition of the entropy measure $\mathcal{L}_t(\cdot)$:

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

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 $\mathcal{L}_t\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) = \log \mathbb{E}_t \frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} - \mathbb{E}_t \log \frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}$, line 5 uses the facts that $P_t^{(1)} = 1/R_t^{(1)}$, $\log \mathbb{E}_t \frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} = 0$ and (from Lemma 1) $\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}} = e^{(\theta_{t+1}^{H,H(\infty)} - \theta_t^{H,H(\infty)})} \cdot \frac{1}{R_{t+1}^{(\infty)}}$, line 6 rearranges, line 7 concludes. Rearranging this final expression, and using definitions of SDFs and excess returns, yields the result.

Finally, using the definition of $\mathcal{L}_t(M_{t,t+1})$ from equation (A.1), this final expression can be written as:

$$\begin{aligned}
\mathcal{L}_t(M_{t,t+1}^{\mathbb{P}}) + \mathbb{E}_t[rx_{t+1}^{(\infty)}] - \theta_t^{H,H(1)} - \mathbb{E}_t[\theta_{t+1}^{H,H(\infty)}] + \theta_t^{H,H(\infty)} \geq& \mathbb{E}_t \log\left[\frac{R_{t,t+1}^g}{R_t}\right] - \theta_t^{H,H(1)} \\
\mathcal{L}_t(M_{t,t+1}^{\mathbb{P}}) + \theta_t^{H,H(\infty)} - \mathbb{E}_t[\theta_{t+1}^{H,H(\infty)}] \geq& \mathbb{E}_t \log\left[\frac{R_{t,t+1}^g}{R_t}\right] - \mathbb{E}_t[rx_{t+1}^{(\infty)}]
\end{aligned}$$

where line 1 substitutes in and line 2 rearranges to yield the second expression (16). □

One can also derive the following unconditional bounds on SDF risk:

$$\mathcal{L}(M_{t,t+1}) \geq \mathbb{E} \log\left[\frac{R_{t,t+1}^g}{R_t}\right] - \mathbb{E}[\theta_t^{H,H(1)}]$$

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

which follow from the definition: $\mathbb{E}\left[\mathcal{L}_t(x_{t+k})\right] = \mathcal{L}(x_{t+k}) - \mathcal{L}\left[\mathbb{E}_t(x_{t+k})\right]$. Since, by stationarity of the SDFs, $\lim_{k\to\infty}\frac{1}{k}\left\{\mathcal{L}\left[\mathbb{E}_t(M^{(*)}_{t,t+k}\right]\right\} = 0$, then $\mathcal{L}(M^{(*)}_{t,t+k} = \mathbb{E}\left[\mathcal{L}_t(M^{(*)}_{t,t+k})\right]$. Applying this to (15) yields the first expression and applying this to (16) yields the second expression, when convenience yields are stationary (i.e., $\mathbb{E}\left[\theta^{i,j(k)}_t - \theta^{i,j(k)}_{t+1}\right] = 0$), completing the proof.

A.4 Proof to Proposition 3

Take the limit of the k-period variant of the result in equation (12), Proposition 1:

$$\lim_{k\to\infty}\frac{1}{k}\mathbb{E}_t[rx^{FX}_{t+k}] = \lim_{k\to\infty}\frac{1}{k}\left\{\mathcal{L}_t(M_{t,t+k}) - \mathcal{L}_t(M^*_{t,t+k})\right\} + \lim_{k\to\infty}\frac{1}{k}\left\{\theta^{F,H(k)}_t - \theta^{F,F(k)}_t\right\}$$

Take unconditional expectations of this expression:

$$\begin{aligned}\lim_{k\to\infty}\frac{1}{k}\mathbb{E}[rx^{FX}_{t+k}] &= \lim_{k\to\infty}\frac{1}{k}\mathbb{E}\left\{\mathcal{L}_t(M_{t,t+k}) - \mathcal{L}_t(M^*_{t,t+k})\right\} + \lim_{k\to\infty}\frac{1}{k}\mathbb{E}\left\{\theta^{F,H(k)}_t - \theta^{F,F(k)}_t\right\}\\ &= \lim_{k\to\infty}\frac{1}{k}\mathbb{E}\left\{\mathcal{L}(M_{t,t+k}) - \mathcal{L}[\mathbb{E}_t(M_{t,t+k})] - \mathcal{L}(M^*_{t,t+k}) + \mathcal{L}[\mathbb{E}_t(M^*_{t,t+k})]\right\}\\ &\quad + \lim_{k\to\infty}\frac{1}{k}\mathbb{E}\left\{\theta^{F,H(k)}_t - \theta^{F,F(k)}_t\right\}\end{aligned}$$

where the second line uses $\mathbb{E}[\mathcal{L}_t(x_{t+k})] = \mathcal{L}[x_{t+k}] - \mathcal{L}[\mathbb{E}_t(x_{t+k})]$.

Then we can use two facts to reproduce the result: (i) since the SDFs are stationary, we have that $\lim_{k\to\infty}\frac{1}{k}\{\mathcal{L}[\mathbb{E}_t(M^{(*)}_{t,t+k})]\} = 0$ and (ii) from Alvarez and Jermann (2005), Proposition (6) $\lim_{k\to\infty}\{\mathcal{L}[M^{(*)}_{t,t+k}]\} = \mathcal{L}(M^{\mathbb{P},(*)}_{t,t+1})$. □

A.5 Proof to Lemma 2

Start with the one-period exchange-rate process (9):

$$\Delta e_{t+1} = m^*_{t,t+1} - m_{t,t+1} + \theta^{F,H(1)}_t - \theta^{H,H(1)}_t$$

and note that it must hold period-by-period, such that:

$$\Delta e_{t+\tau} = m^*_{t+\tau-1,t+\tau} - m_{t+\tau-1,t+\tau} + \theta^{F,H(1)}_{t+\tau-1} - \theta^{H,H(1)}_{t+\tau-1}$$

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:

$$\begin{aligned}
\Delta e_{t+k} + \ldots + \Delta e_{t+1} =& m^*_{t+k-1,t+k} - m_{t+k-1,t+k} + \ldots + m^*_{t,t+1} - m_{t,t+1} \\
&+ \theta^{F,H(1)}_{t+k-1} - \theta^{H,H(1)}_{t+k-1} + \ldots + \theta^{F,H(1)}_{t} - \theta^{H,H(1)}_{t} \\
e_{t+k} - e_t =& m^*_{t,t+k} - m_{t,t+k} + \sum_{\tau=0}^{k-1} \left(\theta^{F,H(1)}_{t+\tau} - \theta^{H,H(1)}_{t+\tau} \right)
\end{aligned}$$

Alongside this, consider the k-period analogue of the exchange-rate process (9):

$$e_{t+k} - e_t = m^*_{t,t+k} - m_{t,t+k} + \theta^{F,H(k)}_t - \theta^{H,H(k)}_t$$

Comparing these final two expressions, we see that the following restriction on the term-structure of convenience yields must hold:

$$\sum_{\tau=0}^{k-1} \left(\theta^{F,H(1)}_{t+\tau} - \theta^{H,H(1)}_{t+\tau} \right) + \sum_{\tau=1}^{k} \eta_{t+\tau} = \theta^{F,H(k)}_t - \theta^{H,H(k)}_t$$

This yields the result (18). □

A.6 Proof to Proposition 4

Take the decomposition of total risk from equation (A.1) from Lemma 1:

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

and its Foreign counterpart. Substitute these into the result in equation (12), Proposition 1:

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

Substituting in steps yields:

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

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

where line 2 rearranges and cancels like terms and line 3 uses the definition of $\mathbb{E}_t[rx^{(\infty),CT}_{t+1}]$.

Next, we know from equation (18), Lemma 2:

$$\theta^{HH(\infty)}_t = \theta^{F,H(\infty)}_t - \sum_{\tau=0}^{\infty} \theta^{F,H(1)}_{t+\tau} + \sum_{\tau=0}^{\infty} \theta^{H,H(1)}_{t+\tau}$$

for all t. Therefore:

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

and

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

where line 1 iterates previous expression forward one period, and line 2 takes expectations from time t.

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

$$\begin{aligned}\mathbb{E}_t[rx^{(\infty),CT}_{t+1}] =& \mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) - \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1}) - \theta^{H,H(1)}_t + \theta^{H,H(\infty)}_t - \mathbb{E}_t[\theta^{H,H(\infty)}_{t+1}] \\ &- \theta^{F,F(\infty)}_t + \mathbb{E}_t[\theta^{F,F(\infty)}_{t+1}] + \theta^{F,H(1)}_t \\ =& \mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) - \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1}) - \theta^{H,H(1)}_t \\ &+ \theta^{F,H(\infty)}_t - \mathbb{E}_t\left[\sum_{\tau=0}^{\infty} \theta^{F,H(1)}_{t+\tau}\right] + \mathbb{E}_t\left[\sum_{\tau=0}^{\infty} \theta^{H,H(1)}_{t+\tau}\right] \\ &- \mathbb{E}_t[\theta^{F,H(\infty)}_{t+1}] + \mathbb{E}_t\left[\sum_{\tau=1}^{\infty} \theta^{F,H(1)}_{t+\tau}\right] - \mathbb{E}_t\left[\sum_{\tau=1}^{\infty} \theta^{H,H(1)}_{t+\tau}\right] \\ &- \theta^{F,F(\infty)}_t + \mathbb{E}_t[\theta^{F,F(\infty)}_{t+1}] + \theta^{F,H(1)}_t \\ =& \mathcal{L}_t(M^{\mathbb{P}}_{t,t+1}) - \mathcal{L}_t(M^{\mathbb{P}*}_{t,t+1}) - \theta^{H,H(1)}_t\end{aligned}$$

$$
\begin{aligned}
&+\theta_t^{F,H(\infty)} - \theta_t^{F,H(1)} - \mathbb{E}_t\left[\sum_{\tau=1}^{\infty}\theta_{t+\tau}^{F,H(1)}\right] + \theta_t^{H,H(1)} + \mathbb{E}_t\left[\sum_{\tau=1}^{\infty}\theta_{t+\tau}^{H,H(1)}\right] \\
&- \mathbb{E}_t[\theta_{t+1}^{F,H(\infty)}] + \mathbb{E}_t\left[\sum_{\tau=1}^{\infty}\theta_{t+\tau}^{F,H(1)}\right] - \mathbb{E}_t\left[\sum_{\tau=1}^{\infty}\theta_{t+\tau}^{H,H(1)}\right] \\
&- \theta_t^{F,F(\infty)} + \mathbb{E}_t[\theta_{t+1}^{F,F(\infty)}] + \theta_t^{F,H(1)} \\
=&\mathcal{L}_t(M_{t,t+1}^{\mathbb{P}}) - \mathcal{L}_t(M_{t,t+1}^{\mathbb{P}*}) + \left(\theta_t^{FH(\infty)} - \mathbb{E}_t[\theta_{t+1}^{F,H(\infty)}]\right) - \left(\theta_t^{F,F(\infty)} - \mathbb{E}_t[\theta_{t+1}^{F,F(\infty)}]\right)
\end{aligned}
$$

where line 1 restates the expression, line 2 uses expressions for $\mathbb{E}_t[\theta_t^{HH(\infty)}]$ and $\mathbb{E}_{t+1}[\theta_t^{HH(\infty)}]$, line 3 breaks out the sum from $\tau = 0, ..., \infty$, line 4 cancels terms. This confirms equation (20) in Proposition 4. □

B Data Sources

Convenience Yields. We use convenience yields at 3-month, 1-year and 10-year maturities from Du, Im, and Schreger (2018a) for 6 industrialised countries relative to the U.S.: Australia, Canada, euro area, Japan, Switzerland and U.K. Our benchmark sample begins 1997:01, although the panel is unbalanced as convenience yields are not available from the start of the sample in all jurisdictions. We use the 3-month and 1-year convenience yields to construct 6-month convenience yields by linearly interpolating according to $\theta_{t,6M} = \frac{2}{3}\theta_{t,3M} + \frac{1}{3}\theta_{t,1Y}$. Table B.1 summarises the start dates of the convenience yields in our study. The data runs through 2020:12.

Table B.1: Du, Im, and Schreger (2018a) Convenience Yield 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 at 6-month and 10-year maturities for 7 industrialised countries: U.S., Australia, Canada, euro area, Japan, Switzerland and U.K. Table B.2 summarises the sources of nominal zero-coupon government bond yields for the economies in our study.

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)

Notes: Data ends December 2020.

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, and there are no suitable ways of extending these series back. 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 OIS rates. The start date for these series is October 2001 for the U.S. and October 1999 for the E.U. and they run until the end of our sample. These OIS swap rates are indexed to the overnight fed funds rate and EONIA for the U.S. and E.A. respectively. Second, for the U.S. at a 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. And third, for the E.U. 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).[27] 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 Du et al. (2022), who use data from JP Morgan Markets.[28] Overall, we use our extended interest-rate swap series beginning in 1997, to match the start date of the CIP deviations from Du et al. (2018a).

For the E.A. 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 to one of 1999.

[27] The codes are HQMCB6MT and HQMCB10YRP for the 6-month and 10-year maturities, respectively.

[28] Presumably, as there was no OIS market in 1996, JP Morgan Markets performs a back-fill similar to us.

Equities. Finally, 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 data are used to construct representative ex-post measures of equity 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).[29] We use these data to construct ex-ante measure of equity risk premia as described in the main text. The construction requires a measure of inflation expectations, which we take from *Consensus Economics*. As with our other data, we use the end-of-month observations for all equity-related series.

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

C Estimating the Mapping from CIP to Convenience Yields

We estimate the coefficient β_k^* in equation (23) following the approach of Jiang et al. (2021a). We extend this to multiple maturities, to investigate the convenience of different assets along the term structure. We focus on the 6-month and 10-year CIP deviation, those used in our paper, and the 1-year CIP deviation to match the maturity used by Jiang et al. (2021a).

β_k^* can be estimated from the data using:

$$\beta = 1 - \frac{1}{1-\alpha_1^4}\frac{1}{\delta_1} \in \{0, 1\} \tag{C.1}$$

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. Both are discussed in detail below.

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:

$$\overline{CIP}_{t+3}^{(k)} = \alpha_1 \overline{CIP}_t^{(k)} + \omega_t. \tag{C.2}$$

Table C.1 reports the results.

Table C.1: Autocorrelation in CIP Deviations

Vars	$\overline{CIP}_{t+3}^{(6M)}$	$\overline{CIP}_{t+3}^{(1Y)}$	$\overline{CIP}_{t+3}^{(10Y)}$
$\overline{CIP}_t^{(6M)}$	0.68***		
	(0.05)		
$\overline{CIP}_t^{(1Y)}$		0.80***	
		(0.04)	
$\overline{CIP}_t^{(10Y)}$			0.89***
			(0.03)
Observations	195	195	207
Within R^2	0.47	0.64	0.87

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:

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

The results from this regression are reported in Table C.2.

Table C.2: Constructing CIP Innovations as Residuals to

Vars	$\Delta\overline{CIP}^{(6M)}_{t+3}$	$\Delta\overline{CIP}^{(1Y)}_{t+3}$	$\Delta\overline{CIP}^{(10Y)}_{t+3}$
$\overline{CIP}^{(6M)}_t$	-0.58***		
	(0.06)		
$\overline{r^{*,(6M)}_t - r^{(6M)}_t}$	154		
	(180)		
$\overline{CIP}^{(1Y)}_t$		-0.44***	
		(0.06)	
$\overline{r^{*,(1Y)}_t - r^{(1Y)}_t}$		104	
		(109)	
$\overline{CIP}^{(10Y)}_t$			-0.16***
			(0.04)
$\overline{r^{*,(10Y)}_t - r^{(10Y)}_t}$			-44
			(161)
Observations	198	198	210
R^2	0.29	0.23	0.09

In the second step, we estimate δ_1 by regressing these innovations, $\widetilde{\overline{CIP}}^{(k)}_t \equiv \omega_t$ where ω_t is the residual from regression (C.2), on quarterly exchange rate movements:

$$\Delta\bar{e}_{t+3} = \delta_1 \Delta\widetilde{\overline{CIP}}^{(k)}_t + \varepsilon_t \tag{C.4}$$

Table C.3 reports the results.

Table C.3: CIP Innovations and Exchange Rate Dynamics

Vars	$\Delta\bar{e}_{t+3}$	$\Delta\bar{e}_{t+3}$	$\Delta\bar{e}_{t+3}$
$\Delta\widetilde{\overline{CIP}}^{(6M)}_t$	-5.38***		
	(1.23)		
$\Delta\widetilde{\overline{CIP}}^{(1Y)}_t$		-14.7***	
		(1.97)	
$\Delta\widetilde{\overline{CIP}}^{(10Y)}_{t+3}$			-18.2***
			(2.88)
Observations	198	198	210
Within R^2	0.08	0.20	0.16

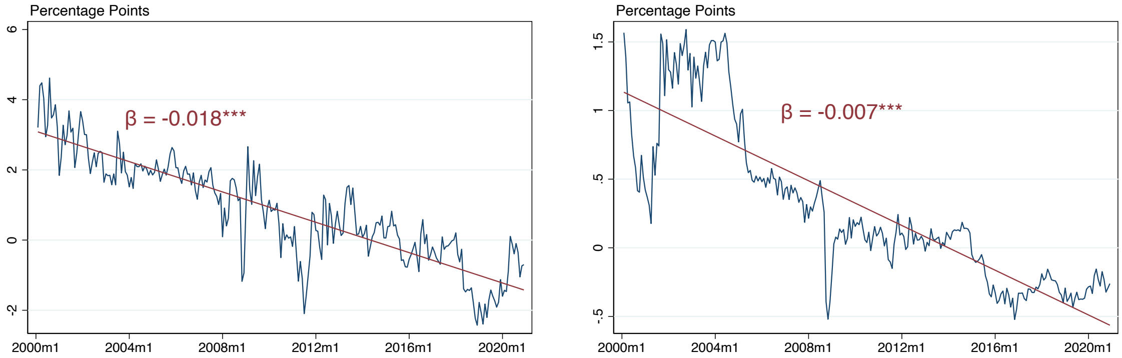
Using these estimates α_1 and δ_1, we find $\beta^{(1Y)} = 0.89$, which is similar to the value found by Jiang et al. (2021a) of $\beta^{(1Y)} = 0.9$. The values found for the 6-month maturity is $\beta^{(6M)} = 0.76$ and for the 10-year maturities is $\beta^{(10Y)} = 0.85$.

D Additional Empirical Exercises

D.1 Additional Raw Data

Figure D.1: Long-Run U.S. Convenience Yields *vis-à-vis* E.A.

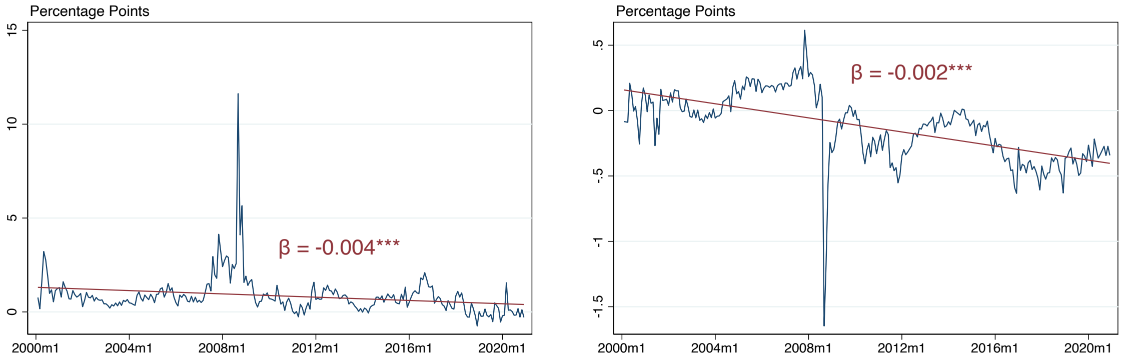
(a) U.S. v E.U 10Y Cross-Country CY $(\theta^{F,H(\infty)})$ (b) U.S. - E.U. 10Y Gov't CY $(\theta^{H,H(\infty)} - \theta^{F,F(\infty)})$



Note. Panels D.1a and D.1b display time series of long-run U.S. minus E.A. cross-country and within-country convenience yields, respectively $(\theta^{F,H(\infty)} - \theta^{F,F(\infty)}$ and $\theta^{H,H(\infty)} - \theta^{F,F(\infty)})$, from 2000:M2 to 2020:M12. The bond maturities are 10 years. *** signifies that the slopes (βs) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are dictated by the underlying CIP-deviation data from Du et al. (2018a).

Figure D.2: Short-Run U.S. Convenience Yields *vis-à-vis* E.A.

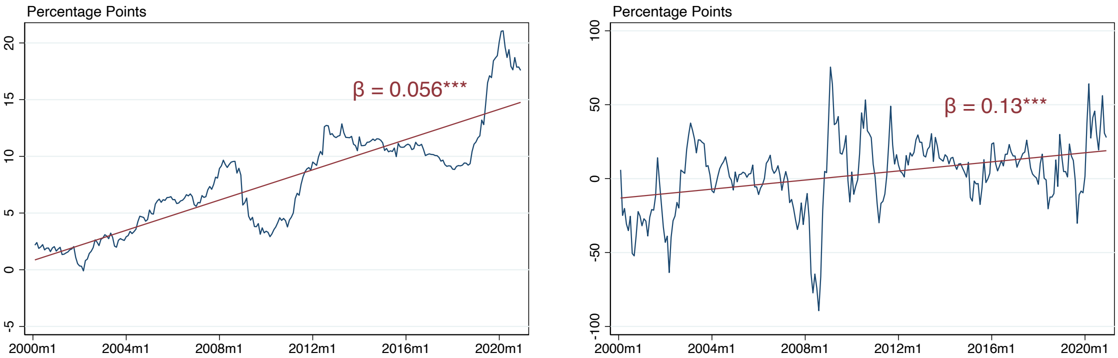
(a) U.S. v E.U. 6M Cross-Country Gov't CY $(\theta^{F,H(1)})$ (b) U.S. - E.U. 6M Gov't CY $(\theta^{H,H(1)} - \theta^{F,F(1)})$



Note. Panels D.2a and D.2b display time series of short-run U.S. minus E.A. cross-country and within-country convenience yields, respectively $(\theta^{F,H(1} - \theta^{F,F(1)}$ and $\theta^{H,H(1)} - \theta^{F,F(1)})$, from 2000:M2 to 2020:M12. The bond maturities are 6 months. *** signifies that the slopes (βs) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are dictated by the underlying CIP-deviation data from Du et al. (2018a).

Figure D.3: U.S. *Ex Ante* Equity Premia and *Ex Post* Equity Returns

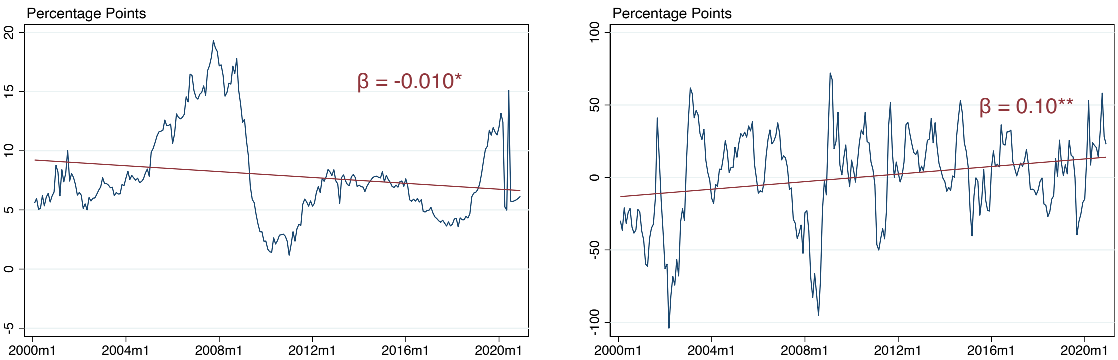
(a) U.S. Expected Equity Premium (b) U.S. Realised Equity Returns



Note. Panels D.3a and D.3b display time series of U.S. *ex ante* equity risk premium constructed estimated via the Gordon growth formula in equation (27) and *realised* equity returns, respectively, from 2000:M2 to 2020:M12. The equity risk premia are calculated using dividend-price ratio data for each country, the 10-year nominal government bond yield adjusted with *Consensus Economics* inflation expectations to construct a 'real' risk-free rate, and the average dividend growth in the 10 years prior to a given data to proxy for expected dividend growth. *** signifies that the slopes (βs) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are chosen to match the CIP-deviation data.

Figure D.4: Euro Area *Ex Ante* Equity Premia and *Ex Post* Equity Returns

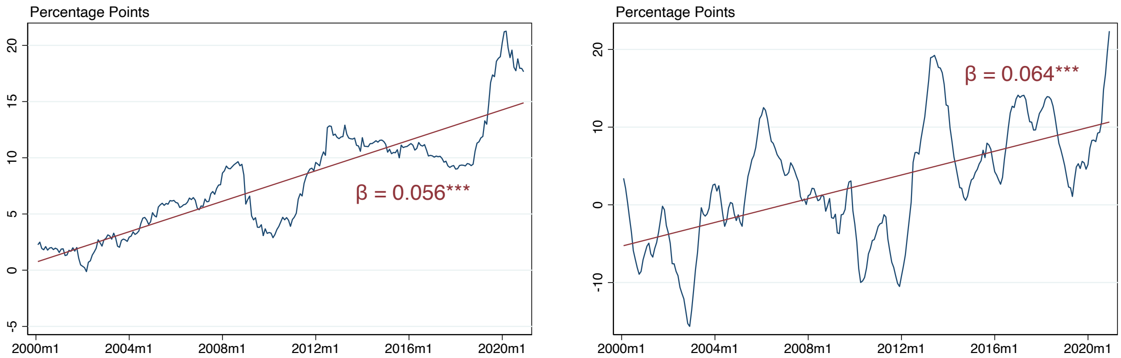
(a) E.A. Expected Equity Premium (b) E.A. Realised Equity Returns



Note. Panels D.4a and D.4b display time series of E.A. *ex ante* equity risk premium constructed estimated via the Gordon growth formula in equation (27) and *realised* equity returns, respectively, from 2000:M2 to 2020:M12. The equity risk premia are calculated using dividend-price ratio data for each country, the 10-year nominal government bond yield adjusted with *Consensus Economics* inflation expectations to construct a 'real' risk-free rate, and the average dividend growth in the 10 years prior to a given data to proxy for expected dividend growth. *** signifies that the slopes (βs) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are chosen to match the CIP-deviation data.

Figure D.5: U.S. *Ex Ante* Overall and Permanent SDF Risk

(a) U.S. Overall SDF Risk (b) U.S. Permanent SDF Risk

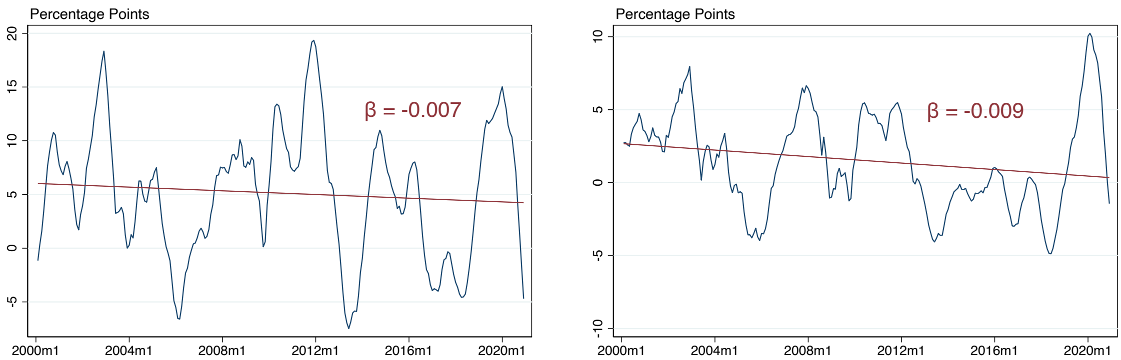


Note. Panels D.5a and D.5b display time series of our proxies for U.S. *ex ante* total and permanent risk ($\tilde{\mathcal{L}}_t(M_{t,t+1})$ and $\tilde{\mathcal{L}}_t\left(M^{\mathbb{P}}_{t,t+1}\right)$), respectively, as calculated according to equations (28) and (29) from 2000:M2 to 2020:M12. *** signifies that the slopes (βs) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are chosen to match the CIP-deviation data.

Figure D.6: U.S. and U.S. Relative to G.7 Transitory Risk

$$rx^{(\infty)}_{t+1} + \left(\tilde{\theta}^{H,H(\infty)}_t - \tilde{\theta}^{H,H(\infty)}_{t+1}\right) - \tilde{\theta}^{H,H(1)}_t$$

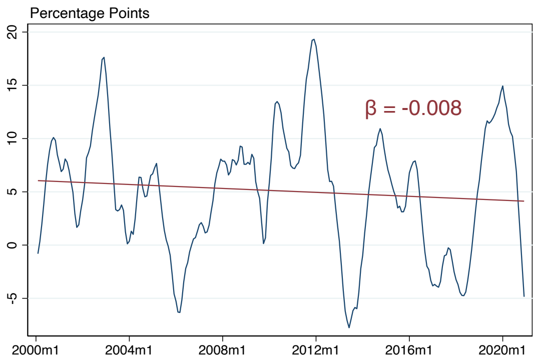
(a) U.S. Transitory Risk (b) U.S. vs. G.7 Transitory Risk



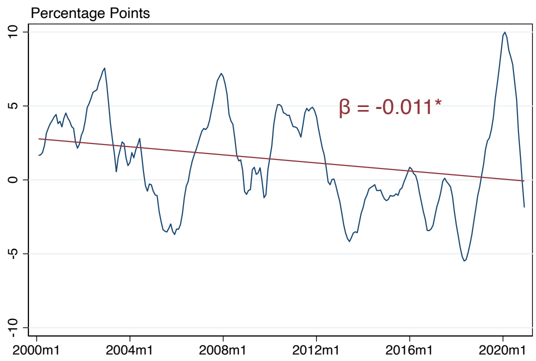
Note. Panels D.6a and D.6b display time series of our proxies for U.S. transitory risk ($\tilde{\mathcal{L}}_t(M^{SR}_{t,t+1})$) and U.S. transitory risk vis-à-vis the average transitory risk across G.7 currency areas, respectively, as calculated according to equations (30) from 2000:M2 to 2020:M12. *** signifies that the slopes (βs) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are chosen to match the CIP-deviation data.

Figure D.7: U.S. and U.S. Relative to G.7 Bond Premium $rx_{t+1}^{(\infty)}$

(a) U.S. Bond Premium



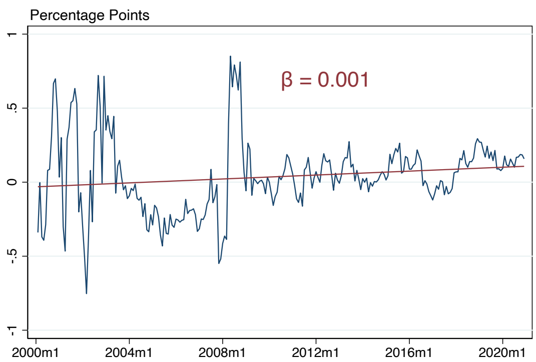
(b) U.S. vs. G.7 Bond Premium



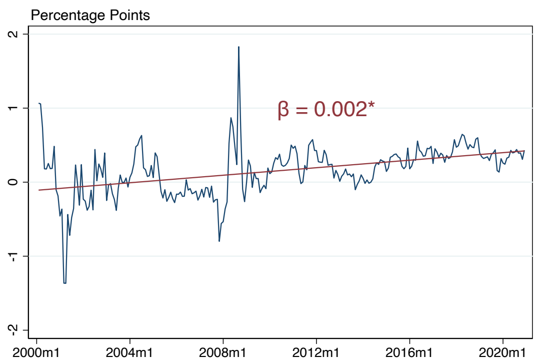
Note. Panels D.7a and D.7b display time series of U.S. bond premium and U.S. bond premium vis-à-vis the average bond premia across G.7 currency areas, respectively, from 2000:M2 to 2020:M12. *** signifies that the slopes (βs) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are chosen to match the CIP-deviation data.

Figure D.8: U.S. and U.S. Relative to E.A. Deviations from Convenience-Yield Expectations Hypothesis: $\left(\tilde{\theta}_t^{H,H(\infty)} - \tilde{\theta}_{t+1}^{H,H(\infty)}\right) - \tilde{\theta}_t^{H,H(1)}$

(a) U.S. Deviation from CY Expectations Hypothesis



(b) U.S. vs. E.A. Deviation from CY Expectations Hypothesis



Note. Panels D.8a and D.8b display time series of U.S. deviations from the convenience-yield expectations and those same deviations relative to those of the G.7 currency areas, respectively, from 2000:M2 to 2020:M12. *** signifies that the slopes (βs) of estimated deterministic trend lines are greater than zero at the 1% significance level based on Newey and West (1987) standard errors with 4 lags. The start and end of the sample are chosen to match the CIP-deviation data.

D.2 Supplementary Regression Results

Table D.1: Bivariate Association Between Relative Risk Measures and Relative Interest Rates

	(1)	(2)	(3)	(4)
	Dep. Var.: Rel. Risk Measure			
	Tot. Risk Ex Ante	ERP Ex Post	Tot. Risk Ex Ante	Eq. Return Ex Post
$r_t^{(6m)} - r_t^{(6m)*}$	1.00***	0.55***	-1.35	-1.04*
	(0.26)	(0.16)	(0.87)	(0.55)
Observations	1,655	2,322	1,655	2,699
# Countries	6	6	6	6
Country FE	YES	YES	YES	YES
Within R^2	0.0565	0.0282	0.0139	0.0110

Table D.2: Regressions for Short-Run (1-year) Cross-Country Convenience, Total Risk and Currency Returns

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Dep. Var.: $\theta_t^{F,H(1)} - \theta_t^{F,F(1)}$				Dep. Var.: rx_{t+1}^{FX}			
Rel. Tot. Risk	-0.03				-0.39***			
	(0.02)				(0.14)			
Rel. ERP		-0.08***				-0.45***		
		(0.03)				(0.11)		
U.S. ERP			-0.03**				-0.36***	
			(0.01)				(0.08)	
R.o.W. ERP			0.02***				0.05	
			(0.01)				(0.06)	
Rel. Tot. Risk				-0.01				0.12***
(*Ex Post*)				(0.01)				(0.04)
$\theta_t^{H,H(1)} - \theta_t^{F,F(1)}$		-2.36***	-2.17***			-2.39	-1.08	
		(0.73)	(0.74)			(3.12)	(3.30)	
$\theta_t^{F,H(1)} - \theta_t^{F,F(1)}$					0.53	0.39	0.44	0.67
					(0.38)	(0.36)	(0.36)	(0.41)
rx_{t+1}^{FX}	0.02	0.01	0.01	0.02				
	(0.01)	(0.01)	(0.01)	(0.02)				
Constant	1.47***	1.11***	1.30***	1.53***	-0.97	-1.18	1.97	-0.72
	(0.16)	(0.13)	(0.17)	(0.17)	(1.24)	(1.33)	(1.52)	(1.34)
Observations	1,435	1,435	1,435	1,435	1,435	1,435	1,435	1,435
# Countries	6	6	6	6	6	6	6	6
Country FE	YES	YES	YES	YES	YES	YES	YES	YES
Within R^2	0.0181	0.135	0.120	0.0156	0.0466	0.0509	0.104	0.0441

Table D.3: Robustness Analysis for Relationship Between Short-Run Cross-Country Convenience, Total Risk and Currency Returns

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dep. Var.:	$\theta_t^{F,H(6m)} - \theta_t^{F,F(6m)}$		$\theta_t^{F,H(1y)} - \theta_t^{F,F(1y)}$		rx_{t+6m}^{FX}		rx_{t+1y}^{FX}	
	Pre-'08	Post-'08	Pre-'08	Post-'08	Pre-'08	Post-'08	Pre-'08	Post-'08
Rel. Tot. Risk	-0.04*	-0.00	-0.04*	-0.01	-0.56**	-0.31	-0.39	-0.44**
	(0.02)	(0.01)	(0.02)	(0.03)	(0.28)	(0.27)	(0.27)	(0.18)
$\theta_t^{F,H(6m)} - \theta_t^{F,F(6m)}$					0.03	2.57		
					(1.60)	(1.83)		
$\theta_t^{F,H(1y)} - \theta_t^{F,F(1y)}$							-2.20***	1.00*
							(0.83)	(0.58)
rx_{t+6m}^{FX}	0.00	0.01						
	(0.01)	(0.00)						
rx_{t+1y}^{FX}			-0.03**	0.03*				
			(0.01)	(0.02)				
Constant	0.60***	0.76***	1.06***	1.67***	1.82	-2.14	4.51***	-3.03*
	(0.08)	(0.07)	(0.16)	(0.19)	(2.53)	(1.81)	(1.60)	(1.66)
Observations	577	882	577	786	577	882	577	786
# Countries	6	6	6	6	6	6	6	6
Country FE	YES	YES	YES	YES	YES	YES	YES	YES
Within R^2	0.0330	0.0176	0.0650	0.0369	0.0224	0.0342	0.0670	0.0846

Table D.4: Robustness Analysis for Relationship Between Long-Run Cross-Country Convenience, Permanent Risk and Carry Trade Returns

	(1)	(2)	(3)	(4)	(5)
	Dep. Var.: $\theta_t^{F,H(\infty)}$			Dep. Var.: $rx_{t+1}^{(\infty),CT}$	
	Ex. Fwd. Lag	Pre-'08	Post-'08	Pre-'08	Post-'08
Rel. $\mathbb{P}$ Risk	-0.06***	-0.03**	0.00	0.35***	0.17
	(0.01)	(0.01)	(0.01)	(0.11)	(0.12)
$\theta_t^{F,H(\infty)} - \theta_t^{F,F(\infty)}$				2.68**	4.79***
				(1.19)	(1.12)
$\theta_{t+1}^{F,H(\infty)} - \theta_{t+1}^{F,F(\infty)}$		0.79***	0.69***	-5.80***	-3.67***
		(0.08)	(0.05)	(0.82)	(0.79)
$rx_{t+1}^{(\infty),CT}$	0.01	0.01**	0.03***		
	(0.01)	(0.00)	(0.01)		
Constant	0.59***	0.38***	0.07	7.84***	-0.98
	(0.14)	(0.12)	(0.08)	(2.46)	(1.44)
Observations	1,544	590	846	590	846
# Countries	6	6	6	6	6
Country FE	YES	YES	YES	YES	YES
Within R^2	0.0645	0.587	0.485	0.208	0.160