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Exchange-Rate Risk and Business Cycles*

Simon P. Lloyd† Emile A. Marin‡

August 5, 2021

Abstract

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1 Introduction

A long-standing literature in international macroeconomics has questioned whether exchange rates can be connected to macroeconomic fundamentals, highlighting an 'exchange rate disconnect' (Meese and Rogoff, 1983; Itskhoki and Mukhin, 2021) that has motivated attempts to 'reconnect' currency moves to fundamentals (e.g. Lilley, Maggiori, Neiman, and Schreger, 2019). In parallel, leading contributions to the asset-pricing literature have taken exchange rate puzzles at face value, assessing the restrictions they impose on the modelling of risk pricing (Backus, Foresi, and Telmer, 2001; Lustig, Stathopoulos, and Verdelhan, 2019). In this paper, building on both approaches, we reconsider the link between exchange rates and the term structure of interest rates both empirically and theoretically. Based on a panel of advanced-economy currencies, we produce novel evidence that cross-country differences in the yield curve slope predict exchange rate dynamics, especially at medium-term horizons. Given the nature of the yield curve slope as a leading indicator of macroeconomic outcomes (e.g. Estrella and Hardouvelis, 1991; Estrella and Mishkin, 1998; Estrella, 2005), we argue that slope differentials across countries drive exchange rate movements in line with asymmetries in business-cycle risk. Using a no-arbitrage framework, we show that, for this to be true, the risks driving stochastic discount factors (SDFs) must be, at least in part, transitory and cyclical. In light of this, we argue that business-cycle risks drive predictable exchange rate movements.

To this end, we first show that cross-country differences in the yield curve slope are a strong predictor of exchange rates, especially at business-cycle horizons. Our starting point is the canonical uncovered interest parity (UIP) regression. UIP predicts that a high interest rate currency should depreciate to equalise exchange rate-adjusted returns, consistent with risk-neutral no-arbitrage. As is well known, the UIP hypothesis is empirically rejected at short to medium horizons: high-yield currencies tend to excessively appreciate (or insufficiently depreciate) due to exchange-rate risk premia (ERRP) (Fama, 1984). But it cannot be rejected at long horizons (e.g. Chinn and Meredith, 2005). By augmenting a UIP regression with cross-country differences in yield curve factors, we show that information in the term structure of interest rates greatly improves explanatory power for exchange rates. Specifically, we document a tent-shaped relationship between the relative yield curve slope and future exchange rate changes. Countries with a relatively steep yield curve tend to depreciate in excess of UIP, with the relationship strongest at the 3 to 5-year horizon.

We then investigate the relationship between the relative slope, bond returns and ERRP across holding periods and bond maturities—extending the empirical analysis in Lustig et al. (2019). Doing so allows us to isolate the specific contribution of the relative yield curve slope to bond risk premia and ERRP in turn. We find that the predictability of exchange rates by the relative yield curve slope predominantly works through ERRP. Once again, we identify a tent-shaped relationship, across holding periods and for a range of bond maturities.

We also extend our specification to account for liquidity yields (Du, Im, and Schreger, 2018), i.e. the non-monetary return that government bonds provide because of their safety, ease of resale, and value as collateral. We find evidence that the term structure of cross-country liquidity yields explains exchange rate fluctuations, extending the results of Engel and Wu (2018) across maturities. Nevertheless, the relationship between the relative yield curve slope and ERRP is robust to this extension, suggesting that business-cycle risk operates through a distinct channel.

Armed with these empirical findings, we investigate the SDF dynamics that rationalise the tent-shaped relationship between the relative yield curve slope and ERRP within a standard no-arbitrage framework. As a first step, we revisit three widely-documented exchange rate puzzles: (i) the failure of UIP at short horizons; (ii) the failure to reject UIP at long horizons; and (iii) the tendency of high-yield currencies to be contemporaneously appreciated, such that UIP holds in 'levels' (Engel, 2016). These empirical regularities imply restrictions on SDFs and interest rates. Together, they suggest that risk driving SDFs must be, in part, transitory and conditionally cyclical. Transitory risk introduces a covariance between today's SDF and expected future SDFs, and drives bond premia (Alvarez and Jermann, 2005). We define risk to be conditionally cyclical when investors, considering shocks up to a given time, expect booms to be followed by busts, dynamics which imply upward-sloping yield curves on average (Piazzesi and Schneider, 2007). We interpret transitory and cyclical dynamics to be indicative of 'business-cycle risk'. In turn, cross-country differences in the slope will reflect asymmetries in business-cycle risk.

We then consider a two-country, two-factor Cox, Ingersoll, and Ross (1985) (CIR) model for interest rates that reflects business-cycle risk. To do this, we derive parametric conditions implied by exchange rate puzzles (i)-(iii), which we show ensure risk driving SDFs is transitory and cyclical. We calibrate the model to satisfy these conditions, and also target moments of bond yields and the Sharpe ratio. The model can quantitatively reproduce the tent-shaped relationship between exchange rate changes and the relative slope, in line with our empirical evidence. Specifically, the model generates this relationship through two pricing factors with different persistence, upon which yields load with opposing sign. Based on our calibration, the model also replicates UIP regression coefficients at short and long horizons.

Through the lens of the model, the relationship between the relative yield curve slope and exchange rates across horizons relies on differences in investors' valuations of returns over the business cycle. A country with a relatively steep yield curve expects comparatively better times ahead. Therefore, investors value returns more highly in the near term, but expect their valuations to decrease over time. Generally, the country with the steeper yield curve will have a relatively low short-term interest rate and will experience a currency depreciation as compensation for exchange-rate risk, consistent with UIP failures at short horizons. However, cross-country return valuations will reverse as investors move along the cycle. Investors that formerly valued returns highly in a (comparative) bust, value them less as they move into a (relative) boom. The relative path of expected future short-term interest rates will reflect these SDF dynamics, and the exchange rate will then begin to appreciate—again consistent with short-horizon UIP failures. The currency of a country with a relatively steep yield curve is therefore expected to depreciate in the near-term. The depreciation will peak at business-cycle horizons, before pressure on the currency to appreciate builds to form the tent-shaped relationship we observe in the data. To reproduce this, two factors are required in the CIR model to generate the cyclical dynamics of future expected short-term interest rates consistent with no-arbitrage at each horizon.

Overall, our paper demonstrates a role for business-cycle risks in driving exchange rate dynamics. These risks are well summarised by the yield curve slope. Therefore, exchange rate fluctuations and, in particular, ERRP can be explained by differences in the term structure of interest rates across countries.

Related literature. Our work is related to a classic literature on the forward-premium puzzle (Hansen and Hodrick, 1980; Fama, 1984), and analysis of UIP across time (Engel, 2016) and horizons (Chinn and Meredith, 2005; Chinn and Quayyum, 2012; Chernov and Creal, 2020). Specifically, our empirical setup builds on Lustig et al. (2019). They show that, for a given one-month holding period, the term structure of carry trade is decreasing. We extend their specification across holding periods to show that the relative slope is a significant predictor of ERRP at holding periods associated with business-cycle horizons, for a range of maturities.

A number of papers show that yield curve factors can significantly predict ERRP, but many focus on horizons shorter than ours (less than 2 years) (Ang and Chen, 2010; Gräb and Kostka, 2018). While Chen and Tsang (2013) also study longer horizons, they only find significance at short ones. We attribute this difference to the fact Chen and Tsang (2013) capture relative yield curve factors by directly estimating Nelson-Siegel decompositions from *relative* interest rate differentials, thus assuming common factor structures across countries. In contrast, we construct proxies for factors using yield curves estimated on a country-by-country basis, allowing factor structures to be country-specific.

We argue that the yield curve reflects business-cycle risk, and show that this can explain time-series variation in ERRP. Colacito, Riddiough, and Sarno (2019) also attribute a role to business cycles in explaining ERRP, but in the cross-section—sorting currencies according to their output gap. Insofar as a high output gap contributes to a steeper yield curve slope, our findings are consistent. However, whilst the output gap is backward-looking, our paper assesses the ability of a forward-looking object (the term structure) in explaining ERRP.

Our theoretical work builds on a prominent literature using no-arbitrage frameworks to derive conditions on SDFs that are consistent with asset prices and asset-pricing puzzles. Alvarez and Jermann (2005) show that the combination of a high equity premium and low term premium requires most SDF volatility to be due to permanent innovations. Lustig et al. (2019) extend this result to show that for long-run UIP to hold between two currencies, countries' SDFs must load symmetrically on permanent innovations. In contrast, we show that differences in the yield curve slope across countries reflect differences in loadings of transitory risk, and has explanatory power for ERRP at short to medium horizons.

Like other papers (e.g. Backus, Foresi, and Telmer, 2001; Lustig, Roussanov, and Verdelhan, 2014; Lustig and Verdelhan, 2019), we use a multi-factor model for interest rates, grounded in Cox et al. (1985), to study currency anomalies. To the best of our knowledge, we are the first to calibrate such a model to match estimated UIP coefficients across short and long horizons.

Recently, Greenwood, Hanson, Stein, and Sunderam (2020) and Gourinchas, Ray, and

Vayanos (2021) relax the assumption of no-arbitrage by considering segmented markets frameworks to explain the relationship between bond premia and ERRP, which we document. This literature is complementary to our work. However, our contribution is to show that the relationship between the relative yield curve slope and ERRP, as well as UIP across horizons, *can* be reconciled within a standard no-arbitrage framework.

The remainder of this paper is structured as follows. Section 2 presents our yield curve-augmented UIP regression. Section 3 documents the empirical relationship between the relative slope and ERRP across holding periods. Section 4 illustrates this relationship within a standard no-arbitrage framework, demonstrating the role of business-cycle risk. Section 5 concludes.

2 Exchange Rates and the Yield Curve Slope

We first summarise results from canonical UIP regressions across horizons augmented with relative yield-curve factors. Although these regressions face some empirical challenges, addressed in Section 3, they do highlight our headline result: the association between exchange rate dynamics and the relative yield curve slope, at business-cycle horizons in particular.

2.1 Canonical UIP Regression

We estimate the following UIP regression for κ-month-ahead exchange rate changes:

$$e_{j,t+\kappa} - e_{j,t} = \beta_{1,\kappa} \left(r^*_{j,t,\kappa} - r_{t,\kappa} \right) + f_{j,k} + u_{j,t+\kappa} \tag{1}$$

where $e_{j,t}$ is the (log) exchange rate of the Foreign country j *vis-à-vis* the Home (base) currency at time t. It is defined as the Foreign price of a unit of base currency such that an increase in $e_{j,t}$ corresponds to a Foreign depreciation. $r^*_{j,t,\kappa}$ is the net κ-period return in the Foreign country and $r_{t,\kappa}$ is the equivalent return in the Home currency. $f_{j,\kappa}$ is a country fixed effect and $u_{j,t+\kappa}$ is the disturbance.

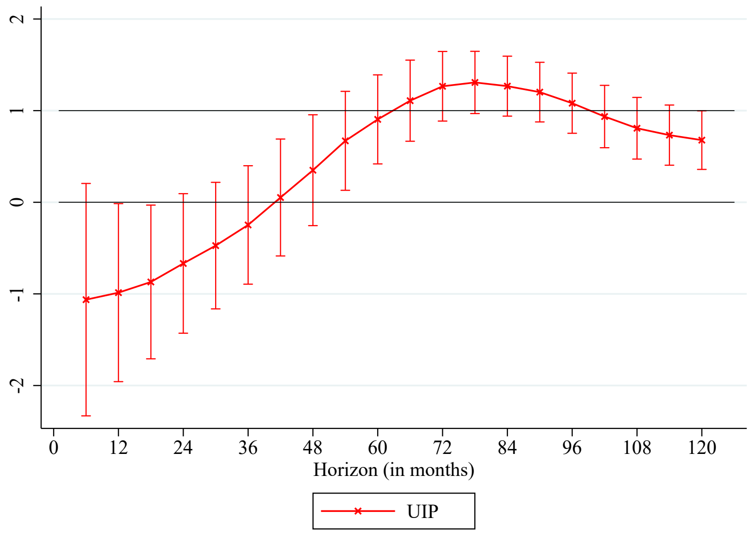
Under the joint assumption of risk neutrality and rational expectations, the null hypothesis of UIP is $\beta_{1,\kappa} = 1$ for all $\kappa > 0$.[1] Empirical rejections of UIP at short to medium horizons—i.e. finding $\hat{\beta}_{1,\kappa} \neq 1$ for small to medium κ—have regularly been used to motivate claims that interest rates do not adequately explain exchange rate dynamics.

Data. We estimate regression (1) using exchange- and interest-rate data for 7 jurisdictions with liquid bond markets: Australia, Canada, Switzerland, euro area, Japan, United Kingdom (UK) and United States (US). The US is the base country among our sample of G7 currencies.[2] To capture the term structure of interest rates in each region, we use nominal zero-coupon government bond yields of 6, 12, 18, ..., 120-month maturities. Yield curves are obtained from

[1] In addition, $f_{j,\kappa} = 0$ for all j and $\kappa > 0$.

[2] The US is our only base currency throughout this paper, as it is well-known that UIP patterns are not fully robust to re-basing.

Figure 1: Estimated coefficients from canonical UIP regression at different horizons



Notes: Red crosses denote $\hat{\beta}_{1,\kappa}$ estimates from regression (1). The horizontal axis denotes the horizon κ in months. Regressions estimated using pooled monthly data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD from 1980:01 to 2017:12, including country fixed effects. 95% confidence intervals, calculated using Driscoll and Kraay (1998) standard errors, are denoted by red bars around point estimates.

a combination of sources, including central banks and Wright (2011) (see Appendix A). Nominal exchange rate data is from *Datastream*. We use end-of-month data from 1980:01 to 2017:12.[3]

Results. Figure 1 plots UIP coefficient estimates $\hat{\beta}_{1,\kappa}$ from regression (1), and these results are tabulated in column (1) of Table 9. The confidence bands around point estimates are derived from Driscoll and Kraay (1998) standard errors, which correct for heteroskedasticity, serial correlation and cross-equation correlation.

The coefficient estimates reinforce that the UIP hypothesis can be rejected at short to medium horizons, but cannot be rejected at longer horizons. At 6 to 36-month tenors, point estimates are negative, indicating that high short-term interest rate currencies tend to appreciate, instead of depreciate. While, at 42 and 48-month horizons point estimates are positive but significantly smaller than unity. Longer-horizon point estimates tend to be positive and close to unity, corroborating with, e.g., Chinn and Meredith (2005) and Chinn and Quayyum (2012).

2.2 Yield Curve-Augmented UIP Regression

To illustrate the link between exchange rates and the yield curve, we augment regression (1) with measures of the relative yield curve slope $S^*_{j,t} - S_t$ and curvature $C^*_{j,t} - C_t$, estimating:

$$e_{j,t+\kappa} - e_{j,t} = \beta_{1,\kappa}\left(r^*_{j,t,\kappa} - r_{t,\kappa}\right) + \beta_{2,\kappa}\left(S^*_{j,t} - S_t\right) + \beta_{3,\kappa}\left(C^*_{j,t} - C_t\right) + f_{j,k} + u_{j,t+\kappa} \qquad (2)$$

[3]As Appendix A documents, our panel of bond yields is unbalanced, with different countries entering the sample at different dates.

for all κ, where $S^*_{j,t}$ ($C^*_{j,t}$) is the slope (curvature) of the Foreign country j yield curve at time t, and S_t (C_t) is the slope (curvature) of the base country yield curve.

Along with the yield curve level, the slope and curvature are known to capture a high degree of variation in bond yields (Litterman and Scheinkman, 1991). We do not include the relative level in regression (2) in order to nest UIP, enabling interpretation of the yield curve's contribution over and above spot-yield differentials. Defining the *ex post* κ-period ERRP for Foreign currency as $rx^{FX}_{j,t,\kappa} \equiv r^*_{j,t,\kappa} - r_{t,\kappa} - (e_{j,t+\kappa} - e_{j,t})$ and combining with equation (2) yields:

$$rx^{FX}_{j,t,\kappa} = (1 - \beta_{1,\kappa})\left(r^*_{j,t,\kappa} - r_{t,\kappa}\right) - \beta_{2,\kappa}\left(S^*_{j,t} - S_t\right) - \beta_{3,\kappa}\left(C^*_{j,t} - C_t\right) - f_{j,\kappa} - u_{j,t+\kappa} \tag{3}$$

Alongside equation (2), $\beta_{2,\kappa}$ can be interpreted as either the average Foreign depreciation (in percent) or the average decrease in the ERRP (in pp) associated with a 1pp increase in the slope of the Foreign yield curve relative to the base country.

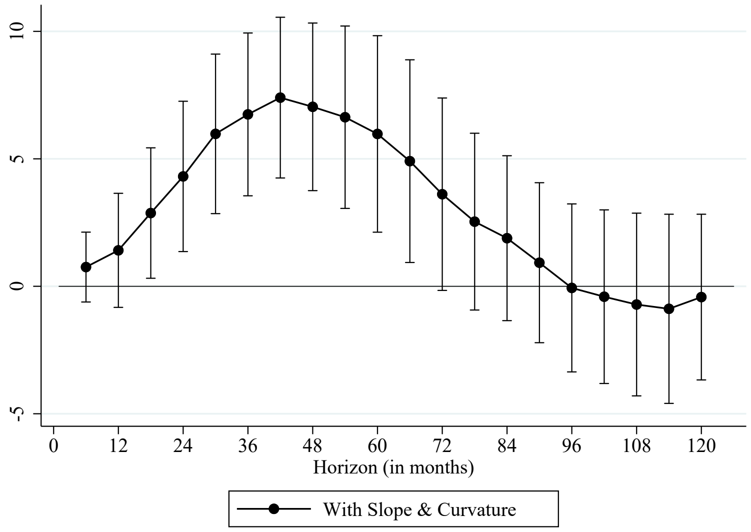
We measure the yield curve slope and curvature in each region with proxies, using the data described in the previous sub-section. We define the slope as the difference between 10-year and 6-month yields, $S^*_{j,t} = y^*_{j,t,10y} - y^*_{j,t,6m}$. Our curvature proxy is a butterfly spread, a function of 6-month, 5 and 10-year yields (Diebold and Rudebusch, 2013): $C^*_{j,t} = 2y^*_{j,t,5y} - (y^*_{j,t,6m} + y^*_{j,t,10y})$. We prefer these measures to principal component estimates of the slope and curvature, which potentially contain look-ahead bias, being defined using weights constructed from information in the whole sample. By construction, our proxies are only based on information available up to time t. Nevertheless, our findings are robust to the use of principal components. Our relative yield-curve proxies are then constructed by taking cross-country differences. Since our proxies are derived from yield curves estimated on a country-by-country basis, we do not assume any symmetry in the factor structure of yield curves across countries.

Results. Figure 2 presents our key result, plotting the relative slope coefficient estimates $\hat{\beta}_{2,\kappa}$. It highlights a tent-shaped relationship across horizons κ between the relative slope and κ-period exchange rate dynamics. Coefficients are insignificantly different from zero at short horizons, but increase in magnitude and significance from short to medium horizons. The $\hat{\beta}_{2,\kappa}$ coefficient peaks at the 3.5-year horizon, quantitatively indicating that a 1pp increase in a country's yield curve slope relative to the US is, on average, associated with a 7.40% exchange rate depreciation over that horizon. At longer horizons—from 6.5-years onwards—the loading on the relative slope is insignificantly different from zero.

Figure 3 complements this by plotting the adjusted R^2 from regression (2) across horizons, together with the comparable figure from the canonical UIP regression (1). The adjusted R^2 from the augmented regression (2) exceeds that of regression (1) at all horizons. But the difference is greatest at 3 to 4-year tenors, indicating that information in the yield curve can account for exchange rate fluctuations over and above spot rate differentials at business-cycle horizons in particular.

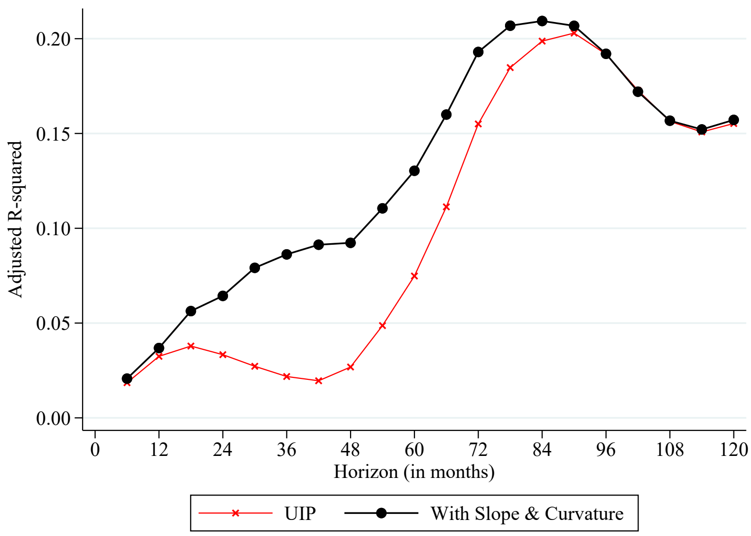
The full results from regression (2), including $\beta_{1,\kappa}$ and $\beta_{3,\kappa}$ estimates, are documented in Table 9 of Appendix B.1. Importantly, the augmentation of the UIP regression with the relative

Figure 2: Estimated relative slope coefficients from augmented UIP regression



Notes: Black circles denote $\hat{\beta}_{2,\kappa}$ point estimates from regression (2). The horizontal axis denotes the horizon κ in months. In regression (2), the slope and curvature in each country are measured using proxies. Regressions are estimated using pooled monthly data from 1980:01 to 2017:12, including country fixed effects. 95% confidence intervals, calculated using Driscoll and Kraay (1998) standard errors, are denoted by black bars around point estimates.

Figure 3: Explanatory power of UIP regression augmented with relative yield curve slope and curvature at different horizons



Notes: Adjusted R^2 from the standard UIP regression (1) of *ex post* exchange rate changes on horizon-specific interest rate differentials (thin, red, crosses) and a yield curve-augmented UIP regression (2) (thick, black, circles), at different horizons κ (horizontal axis, in months). Regressions estimated using pooled end-of-month data for six currencies (AUD, CAD, CHF, EUR, JPY and GBP) *vis-à-vis* the USD from 1980:01 to 2017:12, and include country fixed effects.

yield curve slope and curvature does not significantly alter UIP coefficient estimates. There remains a broadly upward sloping relationship between the UIP coefficient $\hat{\beta}_{1,\kappa}$ and horizons κ. This implies that the contribution of the relative slope can be interpreted over and above spot-yield differentials, as an additional component of ERRP.

2.3 Robustness

In this sub-section, we summarise the robustness of our main empirical finding: that countries with a steeper yield curve tend to experience a subsequent currency depreciation at business-cycle horizons. Further details on the robustness exercises can be found in Appendix B.2.

Predictability of interest rates. The inclusion of interest rates in specification (2) poses a potential challenge, as interest rates are persistent and have a factor structure that is a function of the yield curve slope. To ensure that the relationship between the slope and ERRP is not driven by the predictability of interest rates, we also estimate a simple regression of exchange rate changes on the relative slope and curvature, omitting return differentials. These results, as well as a specification where we include the relative yield curve level alongside slope and curvature as in Chen and Tsang (2013), indicate that the tent-shaped relationship across horizons is robust to these changes.

Long-horizon inference. In long-horizon variants of regressions (1) and (2), the number of non-overlapping observations can be limited. Therefore, size distortions—i.e. the null hypothesis being rejected too often—are a pertinent concern, especially with small samples and persistent regressors (Valkanov, 2003). To carry out more conservative inference, we draw on Moon, Rubia, and Valkanov (2004) who propose the scaling of t-statistics by $1/\sqrt{\kappa}$, showing that these scaled statistics are approximately standard normal when regressors are highly persistent.[4] Our primary result remains significant when using these more conservative t-statistics.

Sub-sample stability. Our main results are robust to splitting the sample into two sub-periods. First, a pre-global financial crisis sample (1980:01-2008:06), which excludes the period in which central banks engaged in unconventional monetary policies. Second, a sample covering the post-crisis period (1990:01-2017:12), in which there was a crash in carry trade around 2008 and a switch in UIP coefficients (Bussière, Chinn, Ferrara, and Heipertz, 2018).

Country-specific regressions. The tent-shaped pattern for the relative slope coefficient is statistically significant for at least three currencies *vis-à-vis* the US dollar. Nevertheless, point estimates exhibit some tent shape for all currencies at short to medium maturities.

[4]Because this is an approximate result, these standard errors are not our preferred metric for inference. Indeed, the scaled t-statistics tend to under-reject the null when regressors are not near-unit root, implying that these confidence bands offer some of the most conservative inference for our regressions.

3 Excess Returns, Risk Premia and the Yield Curve Slope

In this section, we build on the results presented in Section 2 by assessing the association between the relative yield curve slope and different components of government bond returns. To do so, we analyse returns on bonds of maturity κ over different holding periods h. In addition to isolating the contribution of the relative yield curve slope to ERRP and local-currency bond premia, this analysis also reduces the challenges posed by the limited number of non-overlapping observations in regressions (1) and (2) as κ increases.

3.1 Notation

Before presenting our empirical specification, we introduce notation for returns.

Let $P_{t,\kappa}$ denote the price of a κ-maturity zero-coupon bond at time t and $R_{t,\kappa} \geq 1$ denote the gross return on that bond. We distinguish a bond's maturity $\kappa > 0$ from its holding period $h > 0$, where $h \leq \kappa$ and $h = \kappa$ if and only if a bond is held until maturity. The h-month holding period return on a κ-month zero-coupon bond is $HPR^{(\kappa)}_{t,t+h} = P_{t+h,\kappa-h}/P_{t,\kappa}$, i.e. the ratio of the bond's resale price at $t + h$ when its maturity has diminished by h months relative to its time-t price. The (log) excess return on that bond over the holding period h is thus:

$$rx^{(\kappa)}_{t,t+h} = \log\left[\frac{HPR^{(\kappa)}_{t,t+h}}{R_{t,h}}\right] \tag{4}$$

where $R_{t,h}$ is the gross return on an h-month zero-coupon bond at t, i.e. the risk-free rate.

The h-period (log) return on a Foreign bond, expressed in units of US dollars, in excess of the risk-free return in the base currency, $rx^{(\kappa),\$}_{t,t+h}$, can be written:

$$rx^{(\kappa),\$}_{t,t+h} = \log\left[\frac{HPR^{(\kappa)*}_{t,t+h}}{R_{t,h}}\frac{\mathcal{E}_t}{\mathcal{E}_{t+h}}\right] = \log\left[\frac{HPR^{(\kappa)*}_{t,t+h}}{R^*_{t,h}}\right] + \log\left[\frac{R^*_{t,h}}{R_{t,h}}\frac{\mathcal{E}_t}{\mathcal{E}_{t+h}}\right] = rx^{(\kappa)*}_{t,t+h} + rx^{FX}_{t,t+h} \tag{5}$$

where $rx^{(\kappa)*}_{t,t+h}$ represents the (log) local-currency bond return for a Foreign bond and $rx^{FX}_{t,t+h}$ represents the (log) currency excess return.

3.2 Empirical Setup

To study the time series properties of returns, we use the above definitions to estimate the following panel regressions for different holding periods h and bond maturities κ:

$$\mathrm{y}^{(\kappa)}_{j,t,h} = \gamma^{(\kappa)}_{1,h}\left(S^*_{j,t} - S_t\right) + f^{(\kappa)}_{j,h} + \varepsilon^{(\kappa)}_{j,t+h} \tag{6}$$

where $\mathrm{y}^{(\kappa)}_{j,t,h}$ is either the excess return on the Foreign bond in US dollar-terms relative to the US return $rx^{(\kappa),\$}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$ (the dollar-bond return difference), the excess return from Foreign currency $rx^{FX}_{j,t,t+h}$, or the excess return on the Foreign bond in Foreign currency units relative to the US return $rx^{(\kappa)*}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$ (the local currency-bond return difference). $\gamma^{(\kappa)}_{1,h}$ has a similar interpretation to $\beta_{2,\kappa}$ from Section 2, but with opposite sign. When $\mathrm{y}^{(\kappa)}_{j,t,h} = rx^{FX}_{j,t,t+h}$, $\gamma^{(\kappa)}_{1,h}$ can be interpreted like $\beta_{2,\kappa}$, albeit in units of annual excess returns.

Focusing on $h = 1$ and $\kappa = 120$ only, using regression (6), Lustig et al. (2019) show that the relative yield curve slope has an insignificant influence on $rx^{(\kappa),\$}_{t,t+h}$, but opposing effects on $rx^{(\kappa)}_{t,t+h}$ (positive coefficient) and $rx^{FX}_{t,t+h}$ (negative coefficient), which cancel out for the dollar-bond excess return overall. Our empirical framework extends this, assessing the predictability of excess returns with yield curve slope differentials at a range of maturities κ and holding periods h, bridging the gap between our results in Section 2 and those of Lustig et al. (2019).

3.3 Results

The results for regression (6) are presented in Tables 1 and 2.

Importantly, where our regression specification most closely matches Lustig et al. (2019), at short-holding periods $h = 6$ and the longest maturity $\kappa = 120$, our results mirror theirs.[5] The relative slope exerts an insignificant effect on the dollar-bond risk premium difference (Panel A), a positive and significant influence on the local currency-bond risk premium difference $rx^{(120)*}_{j,t,t+6} - rx^{(120)}_{US,t,t+6}$ (Panel C), and a negative and significant influence on the currency risk premium $rx^{FX}_{j,t,t+6}$ (Panel B). The latter two effects are similar in magnitude such that they cancel out for dollar-bond return differences.[6]

Exploring our results at all holding periods h and for all maturities κ, three observations are noteworthy. First, for a given maturity, the loading on the relative slope exhibits a(n inverse) tent shape across holding periods for both the currency risk premium and the relative dollar-bond risk premium. Although significant at shorter holding periods, the relative slope loadings are quantitatively small for local currency-bond premia and are dominated by loadings on currency excess returns in explaining the relative slope's impact on relative dollar-bond risk premia. This supports the findings from our benchmark augmented UIP regression in Section 2. Furthermore, the relative slope exerts its peak influence on dollar-bond and currency excess returns at the 36-month holding period, close to the peak at 42-month horizon from the augmented UIP regression (2).

Second, and related to the first, while the relative yield curve slope does not significantly predict dollar-bond excess return differences at the 6-month holding period for 10-year bonds, the relative slope loading for the same bond maturity is significantly non-zero over longer holding periods. While, in the former case, the influence of the relative slope on currency and local-currency bond returns offset one another (in line with Lustig et al., 2019), our results indicate that the influence of the relative slope on the currency premium dominates over longer holding periods, even for long-term bonds. Nevertheless, for a given holding period, the influence of the relative slope on dollar-bond returns decreases in magnitude with maturity.

[5] Lustig et al. (2019) consider a 1-month holding period, so comparison is not exact.

[6] More generally, the short-horizon local-currency bond return difference predictability confirm results for US bond returns (see, e.g., Fama and Bliss, 1987; Campbell and Shiller, 1991; Cochrane and Piazzesi, 2005).

Table 1: Slope coefficient estimates from pooled regression of excess returns

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
					Holding Periods					
	6m	12m	18m	24m	30m	36m	42m	48m	54m	60m
Panel A: Dependent Variable: $rx^{(\kappa),\$}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$										
12m	-1.74***									
	(0.38)									
18m	-1.63***	-2.17***								
	(0.37)	(0.56)								
24m	-1.52***	-2.09***	-2.75***							
	(0.36)	(0.54)	(0.66)							
30m	-1.42***	-2.02***	-2.71***	-3.00***						
	(0.36)	(0.53)	(0.65)	(0.73)						
36m	-1.32***	-1.94***	-2.66***	-2.99***	-3.32***					
	(0.36)	(0.53)	(0.63)	(0.72)	(0.75)					
42m	-1.21***	-1.86***	-2.60***	-2.97***	-3.32***	-3.38***				
	(0.36)	(0.52)	(0.62)	(0.71)	(0.74)	(0.76)				
48m	-1.11***	-1.77***	-2.54***	-2.94***	-3.31***	-3.39***	-3.06***			
	(0.37)	(0.51)	(0.61)	(0.70)	(0.73)	(0.75)	(0.85)			
54m	-1.00***	-1.68***	-2.46***	-2.90***	-3.28***	-3.38***	-3.07***	-2.53**		
	(0.37)	(0.51)	(0.60)	(0.69)	(0.73)	(0.75)	(0.84)	(1.00)		
60m	-0.90**	-1.59***	-2.39***	-2.85***	-3.25***	-3.36***	-3.07***	-2.54**	-1.95*	
	(0.38)	(0.51)	(0.60)	(0.68)	(0.72)	(0.74)	(0.83)	(0.99)	(1.12)	
66m	-0.80**	-1.50***	-2.31***	-2.79***	-3.21***	-3.34***	-3.05***	-2.54**	-1.95*	-1.53
	(0.39)	(0.51)	(0.59)	(0.68)	(0.71)	(0.73)	(0.82)	(0.98)	(1.11)	(1.23)
72m	-0.71*	-1.42***	-2.24***	-2.74***	-3.17***	-3.31***	-3.03***	-2.53***	-1.95*	-1.53
	(0.39)	(0.50)	(0.58)	(0.67)	(0.71)	(0.73)	(0.82)	(0.97)	(1.10)	(1.22)
78m	-0.61	-1.33***	-2.17***	-2.68***	-3.12***	-3.27***	-3.00***	-2.51***	-1.95*	-1.53
	(0.40)	(0.50)	(0.58)	(0.66)	(0.70)	(0.72)	(0.81)	(0.97)	(1.10)	(1.22)
84m	-0.55	-1.26**	-2.11***	-2.63***	-3.07***	-3.22***	-2.97***	-2.49***	-1.93*	-1.52
	(0.41)	(0.50)	(0.57)	(0.66)	(0.70)	(0.72)	(0.81)	(0.96)	(1.09)	(1.21)
90m	-0.45	-1.19**	-2.04***	-2.57***	-3.02***	-3.18***	-2.93***	-2.46**	-1.91*	-1.51
	(0.41)	(0.50)	(0.57)	(0.66)	(0.69)	(0.71)	(0.80)	(0.95)	(1.08)	(1.20)
96m	-0.37	-1.12**	-1.98***	-2.52***	-2.97***	-3.13***	-2.89***	-2.43**	-1.89*	-1.50
	(0.42)	(0.50)	(0.57)	(0.65)	(0.69)	(0.71)	(0.80)	(0.95)	(1.08)	(1.20)
102m	-0.29	-1.05**	-1.92***	-2.47***	-2.92***	-3.09***	-2.85***	-2.40**	-1.87*	-1.48
	(0.42)	(0.50)	(0.57)	(0.65)	(0.68)	(0.71)	(0.79)	(0.94)	(1.07)	(1.19)
108m	-0.22	-0.99*	-1.86***	-2.42***	-2.87***	-3.04***	-2.81***	-2.36**	-1.84*	-1.46
	(0.43)	(0.51)	(0.56)	(0.65)	(0.68)	(0.70)	(0.79)	(0.94)	(1.06)	(1.19)
114m	-0.15	-0.92*	-1.81***	-2.37***	-2.82***	-2.99***	-2.76***	-2.32**	-1.82*	-1.44
	(0.43)	(0.51)	(0.56)	(0.64)	(0.68)	(0.70)	(0.79)	(0.94)	(1.06)	(1.18)
120m	-0.08	-0.86*	-1.75***	-2.32***	-2.77***	-2.95***	-2.72***	-2.29**	-1.79*	-1.42
	(0.44)	(0.51)	(0.56)	(0.64)	(0.67)	(0.70)	(0.78)	(0.93)	(1.05)	(1.18)
Panel B: Dependent Variable: $rx^{FX}_{j,t,t+h}$										
S^{*}-S	-1.84***	-2.25***	-2.80***	-3.01***	-3.32***	-3.37***	-3.04***	-2.51**	-1.93*	-1.52
	(0.39)	(0.57)	(0.67)	(0.74)	(0.76)	(0.77)	(0.86)	(1.00)	(1.13)	(1.24)
N	2,326	2,290	2,254	2,218	2,182	2,146	2,110	2,074	2,038	2,002

Notes: Coefficient estimates on the relative yield curve slope $S^{*}_t - S_t$ from regressions with the log dollar-bond excess return difference (Panel A) or the h-period log currency excess return (Panel B) as dependent variables. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD for different samples. The log returns and yield curve slopes differentials are annualised. All regressions include country fixed effects. The panels are unbalanced and standard errors (reported in parentheses) are constructed according to the Driscoll and Kraay (1998) methodology. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively.

Table 2: Slope coefficient estimates from pooled regression of excess returns

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
					Holding Periods					
	6m	12m	18m	24m	30m	36m	42m	48m	54m	60m
Panel C: Dependent Variable: $rx^{(\kappa)}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$										
12m	0.09*									
	(0.05)									
18m	0.21**	0.08**								
	(0.10)	(0.04)								
24m	0.31**	0.15**	0.05							
	(0.13)	(0.07)	(0.03)							
30m	0.42**	0.22**	0.09	0.01						
	(0.17)	(0.10)	(0.06)	(0.03)						
36m	0.52***	0.30**	0.14	0.02	0.00					
	(0.20)	(0.13)	(0.09)	(0.05)	(0.02)					
42m	0.62***	0.39**	0.19*	0.04	0.00	-0.01				
	(0.23)	(0.16)	(0.11)	(0.08)	(0.05)	(0.02)				
48m	0.73***	0.48***	0.26*	0.07	0.01	-0.01	-0.01			
	(0.25)	(0.18)	(0.14)	(0.10)	(0.06)	(0.04)	(0.02)			
54m	0.83***	0.57***	0.33**	0.12	0.04	-0.01	-0.02	-0.01		
	(0.28)	(0.20)	(0.15)	(0.11)	(0.08)	(0.06)	(0.04)	(0.02)		
60m	0.94***	0.66***	0.41**	0.16	0.07	0.01	-0.02	-0.02	-0.01	
	(0.30)	(0.21)	(0.17)	(0.13)	(0.10)	(0.07)	(0.05)	(0.03)	(0.02)	
66m	1.03***	0.75***	0.48***	0.22	0.11	0.03	-0.00	-0.02	-0.02	-0.01
	(0.31)	(0.22)	(0.18)	(0.14)	(0.11)	(0.08)	(0.06)	(0.05)	(0.03)	(0.01)
72m	1.13***	0.83***	0.56***	0.27*	0.15	0.07	0.02	-0.01	-0.02	-0.01
	(0.33)	(0.24)	(0.20)	(0.15)	(0.12)	(0.09)	(0.08)	(0.06)	(0.04)	(0.03)
78m	1.23***	0.91***	0.63***	0.33**	0.20	0.11	0.05	0.01	-0.01	-0.01
	(0.34)	(0.25)	(0.21)	(0.17)	(0.13)	(0.10)	(0.09)	(0.07)	(0.06)	(0.04)
84m	1.29***	0.99***	0.69***	0.38**	0.25*	0.15	0.08	0.03	0.01	-0.00
	(0.36)	(0.26)	(0.22)	(0.17)	(0.14)	(0.11)	(0.10)	(0.08)	(0.07)	(0.05)
90m	1.39***	1.06***	0.76***	0.44**	0.30**	0.19	0.12	0.06	0.02	0.01
	(0.37)	(0.27)	(0.23)	(0.18)	(0.15)	(0.12)	(0.11)	(0.09)	(0.08)	(0.06)
96m	1.47***	1.13***	0.81***	0.49**	0.35**	0.24*	0.16	0.09	0.05	0.02
	(0.38)	(0.28)	(0.23)	(0.19)	(0.16)	(0.13)	(0.11)	(0.10)	(0.08)	(0.07)
102m	1.54***	1.19***	0.88***	0.54***	0.40**	0.29**	0.20*	0.12	0.07	0.04
	(0.39)	(0.29)	(0.24)	(0.20)	(0.16)	(0.14)	(0.12)	(0.11)	(0.09)	(0.07)
108m	1.62***	1.26***	0.93***	0.59***	0.45***	0.33**	0.25*	0.15	0.10	0.06
	(0.40)	(0.30)	(0.25)	(0.21)	(0.17)	(0.14)	(0.13)	(0.11)	(0.10)	(0.08)
114m	1.69***	1.32***	0.99***	0.65***	0.50***	0.38**	0.29**	0.19	0.12	0.08
	(0.41)	(0.31)	(0.26)	(0.21)	(0.18)	(0.15)	(0.13)	(0.12)	(0.10)	(0.09)
120m	1.76***	1.38***	1.04***	0.69***	0.55***	0.43***	0.34**	0.23*	0.15	0.10
	(0.42)	(0.32)	(0.27)	(0.22)	(0.18)	(0.15)	(0.14)	(0.12)	(0.11)	(0.09)
N	2,326	2,290	2,254	2,218	2,182	2,146	2,110	2,074	2,038	2,002

Notes: Coefficient estimates on the relative yield curve slope $S^*_t - S_t$ from regressions with the h-period log local currency-bond excess return difference (Panel C) as dependent variable. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD for different samples. The log returns and yield curve slopes differentials are annualised. All regressions include country fixed effects. The panels are unbalanced and standard errors (reported in parentheses) are constructed according to the Driscoll and Kraay (1998) methodology. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively.

Table 3: Robustness of relative slope coefficient estimates from regression (6) for $rx^{FX}_{t,t+h}$

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
					Holding Periods					
	6m	12m	18m	24m	30m	36m	42m	48m	54m	60m
A: Controlling for interest rate differentials										
S^*-S	-0.37	-0.85	-2.00*	-2.66**	-3.57***	-4.11***	-4.26***	-4.13***	-4.02***	-3.90***
	(0.65)	(0.98)	(1.06)	(1.19)	(1.26)	(1.25)	(1.14)	(1.11)	(1.13)	(1.21)
N	2,329	2,293	2,257	2,221	2,185	2,149	2,113	2,077	2,041	2,005
B.i: 1980:01-2008:06 sub-sample										
S^*-S	-2.06***	-2.52***	-3.23***	-3.59***	-4.02***	-4.20***	-4.01***	-3.53***	-2.96**	-2.58*
	(0.41)	(0.63)	(0.75)	(0.83)	(0.85)	(0.84)	(0.90)	(1.04)	(1.17)	(1.32)
N	1,692	1,692	1,692	1,692	1,692	1,692	1,692	1,692	1,692	1,692
B.ii: 1990:01-2017:12 sub-sample										
S^*-S	-1.44***	-1.67***	-2.11***	-2.29***	-2.55***	-2.62***	-2.32***	-1.82*	-1.33	-0.74
	(0.43)	(0.60)	(0.71)	(0.77)	(0.77)	(0.78)	(0.87)	(1.06)	(1.23)	(1.35)
N	1,957	1,921	1,885	1,849	1,813	1,777	1,741	1,705	1,669	1,633

Notes: Coefficient estimates on the relative yield curve slope $S^*_t - S_t$ from regressions with the h-period log currency excess return as dependent variables. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD for different samples. The log returns and yield curve slopes differentials are annualised. All regressions include country fixed effects. The panels are unbalanced and standard errors (reported in parentheses) are constructed according to the Driscoll and Kraay (1998) methodology. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively. Regressions in Panel A also include relative h-period return differentials. Regressions in Panel B.i and B.ii estimated over 1980:01-2008:06 and 1990:01-2017:12 sub-samples, respectively.

Third, for a given holding period, the loading on the relative slope for relative dollar-bond returns is similar across maturities. Insofar as the relative slope influences ERRP, the association is strongest at horizons associated with business cycle movements, specifically 3 to 4-years.

3.4 Robustness

In this sub-section, we briefly summarise robustness analyses for these empirical findings. We focus on the relationship between ERRP and the yield curve slope across horizons.

Controlling for interest rate differentials. We extend regression (6) by including h-period return differentials alongside the relative slope. As Panel A of Table 3 demonstrates, the (inverse) tent-shaped relationship on the relative slope is robust to this addition.

Sub-sample stability. Panel B.i and B.ii of Table 3 demonstrate that the association between the relative slope and ERRP is robust to sub-sample splits. Panel B.i presents a pre-global financial crisis sample (1980:01-2008:06) and Panel B.ii shows results from a sample spanning the period after the crisis (1990:01-2017:12).

Cross-sectional returns. To account for returns in the cross-section, we consider the average returns across maturities κ and holding periods h from a simple investment strategy based on the yield curve slope. Specifically, we consider a strategy that goes long the Foreign bond and short the US bond when the Foreign yield curve is less steep than the US one, and *vice versa*. The results are presented in Appendix B.3. They demonstrate that average returns have a tent-shaped pattern across holding periods, for different maturities, supporting evidence of the yield curve slope's predictive role for returns.

3.5 Accounting for Liquidity Yields

Recent contributions to the literature have emphasised the role for liquidity yields—i.e. non-pecuniary returns, especially for US bonds—in exchange rate determination (see, e.g. Engel and Wu, 2018; Jiang, Krishnamurthy, and Lustig, 2018). In this sub-section, we extend our empirical specification to account for these liquidity yields. We demonstrate that the tent-shaped relationship between the relative slope and ERRP across horizons continues to be robust to this extension, and therefore explains variation in exchange rates independently from liquidity yields. We also analyse the link between the term structure of liquidity yields and ERRP.

Liquidity yield-augmented regression. To do this, we use data on the term structure of liquidity yields from Du et al. (2018).[7] These measure the difference between riskless market rates and government yields at different maturities to quantify the implicit liquidity yield on a government bond, correcting for other frictions in forward markets and sovereign risk. Let $\eta^R_{j,t,\kappa}$ denote the κ-horizon liquidity premium for a κ-horizon US government bond relative to an equivalent-maturity Foreign government bond yield in country j. An increase in $\eta^R_{j,t,\kappa}$ reflects an increase in the relative liquidity of US Treasuries *vis-à-vis* country j.

Although the Du et al. (2018) data is available from 1991:04 for some countries and tenors (e.g. UK), some series begin as late as 1999:01 due to data availability (e.g. euro area). Given these shorter samples, the problem of non-overlapping observations becomes especially pertinent. For this reason, our preferred empirical specification extends regression (6):

$$\mathrm{y}_{j,t,h} = \gamma_{1,h}\left(S^*_{j,t} - S_t\right) + \gamma_{2,h}\eta^R_{j,t,\kappa} + f_{j,h} + \varepsilon_{j,t+h} \tag{7}$$

where the dependent variable $\mathrm{y}_{j,t,h}$ is either the relative dollar-bond return, the currency excess return, or the relative local currency-bond return. The interpretation of $\gamma_{1,h}$ is unchanged relative to equation (6). $\gamma_{2,h}$ can be interpreted as the average influence of a 1pp increase in relative US Treasury convenience. When the currency excess return $rx^{FX}_{t,t+h}$ is the dependent variable, we expect $\gamma_{2,h}$ to be positive, such that an increase in relative US Treasury liquidity is associated with a contemporaneous appreciation of the US dollar (depreciation of Foreign currency) that increases the currency excesses return $rx^{FX}_{t,t+h}$.

Results. The results for the relative dollar-bond excess return are presented in Table 4. Panel A.i documents the estimated coefficient loadings on the relative slope, which are similar to those in Table 1. As before, the slope loading is insignificant for excess returns over short and long holding periods for long-term bonds, consistent with the failure to reject UIP in the long run. At medium holding periods, the influence of the slope is significant, with the coefficient peaking at business-cycle horizons—in this case, 2.5 to 3-years—similar in magnitude to the results presented in Table 1.

[7] Du et al. (2018) show that over 75% of variation in their measure of the 'US Treasury premium' is attributed to liquidity considerations. The data is available for 12, 24, 36, 60, 84 and 120-month tenors only, constraining the maturities we assess in this section.

Table 4: Slope and liquidity yield coefficient estimates from pooled regression of excess returns

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Holding Periods									
	6m	12m	18m	24m	30m	36m	42m	48m	54m	60m
Panel A.i: Dependent Variable: $rx^{(\kappa),\$}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$, Coefficient on $S^* - S$										
12m	-1.74**									
	(0.68)									
24m	-1.29**	-1.84**	-2.34**							
	(0.54)	(0.79)	(0.91)							
36m	-1.06*	-1.72**	-2.27**	-2.27**	-2.29**					
	(0.54)	(0.78)	(0.89)	(1.00)	(1.02)					
60m	-0.52	-1.45*	-2.18***	-2.44**	-2.66***	-2.59**	-2.09*	-1.41	-1.07	
	(0.54)	(0.74)	(0.84)	(0.97)	(1.00)	(1.01)	(1.08)	(1.29)	(1.37)	
84m	-0.06	-1.13	-1.92**	-2.28**	-2.62***	-2.64**	-2.20**	-1.60	-1.32	-1.11
	(0.57)	(0.73)	(0.83)	(0.97)	(1.00)	(1.02)	(1.09)	(1.29)	(1.37)	(1.48)
120m	0.38	-0.73	-1.58*	-1.98**	-2.34**	-2.44**	-2.15**	-1.68	-1.53	-1.34
	(0.61)	(0.73)	(0.80)	(0.92)	(0.94)	(0.95)	(0.96)	(1.11)	(1.16)	(1.26)
Panel A.ii: Dependent Variable: $rx^{(\kappa),\$}_{t,t+h} - rx^{(\kappa)}_{US,t,t+h}$, Coefficient on η^R_κ										
12m	0.03									
	(0.02)									
24m	0.00	0.04	0.06**							
	(0.02)	(0.03)	(0.03)							
36m	0.01	0.04	0.07**	0.12***	0.16***					
	(0.02)	(0.03)	(0.03)	(0.04)	(0.04)					
60m	-0.00	0.03	0.05	0.09**	0.14***	0.17***	0.20***	0.21***	0.21***	
	(0.03)	(0.03)	(0.03)	(0.04)	(0.05)	(0.05)	(0.04)	(0.04)	(0.04)	
84m	0.00	0.03	0.04	0.08**	0.13***	0.16***	0.18***	0.20***	0.20***	0.22***
	(0.03)	(0.03)	(0.03)	(0.04)	(0.05)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)
120m	-0.01	0.02	0.04	0.07*	0.12***	0.16***	0.21***	0.24***	0.27***	0.29***
	(0.02)	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.03)	(0.04)
N	1,733	1,697	1,661	1,625	1,589	1,553	1,517	1,481	1,445	1,409

Notes: Coefficient estimates on the relative yield curve slope $S^*_t - S_t$ (Panel A.i) and cross-country κ-period liquidity yield η^R_κ (Panel A.ii) from regressions with the log dollar-bond excess return difference as dependent variable. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD for different samples within 1991:04-2017:12. The log returns and yield curve slopes differentials are annualised. All regressions include country fixed effects. The panels are unbalanced and standard errors (reported in parentheses) are constructed according to the Driscoll and Kraay (1998) methodology. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively.

Panel A.ii presents the $\gamma_{2,h}$ coefficient estimates for relative liquidity yields. For a given maturity, the coefficient on the relative liquidity yield rises monotonically with respect to holding period, growing in significance. In this case, a higher US Treasury liquidity premium is associated with a higher excess return on a Foreign bond in US dollar terms.

Table 5 focuses on ERRP, from the decomposition of dollar-bond returns into ERRP and local currency-returns. As in the rest of this section, the coefficients indicate that the influence of both of relative slope and relative liquidity yields on dollar-bond excess returns predominantly works through currency excess returns. In contrast, the $\gamma_{2,h}$ loadings (shown in Appendix B.4) for local currency-bond excess returns are negative and relatively small in magnitude.

4 Theory

In this section, we show that the tent-shaped relationship between exchange rates and the relative yield curve slope is consistent with a no-arbitrage framework, driven by cross-country differences in business-cycle risk. As a first step, we revisit empirical regularities for exchange rates—such as UIP at short and long horizons—within a preference-free setting to study the restrictions they impose on SDFs. We then present a two-country Cox, Ingersoll, and Ross

Table 5: Slope and liquidity yield coefficient estimates from pooled regression of excess returns

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Holding Periods									
	6m	12m	18m	24m	30m	36m	42m	48m	54m	60m
Panel B.i: Dependent Variable: $rx^{FX}_{j,t,t+h}$, Coefficient on $S^{*}-S$, when η^{R}_{κ} is additional control										
12m	-1.71**	-2.21**								
	(0.70)	(0.98)								
24m	-1.50***	-1.87**	-2.32**	-2.31**						
	(0.56)	(0.82)	(0.92)	(1.03)						
36m	-1.48***	-1.85**	-2.27**	-2.24**	-2.26**	-2.07**				
	(0.56)	(0.82)	(0.92)	(1.02)	(1.02)	(1.02)				
60m	-1.55***	-2.02**	-2.51***	-2.61**	-2.76***	-2.65**	-2.11*	-1.42	-1.07	-0.78
	(0.57)	(0.80)	(0.91)	(1.02)	(1.03)	(1.03)	(1.09)	(1.30)	(1.37)	(1.47)
84m	-1.59***	-2.12***	-2.61***	-2.77***	-3.00***	-2.93***	-2.41**	-1.76	-1.44	-1.21
	(0.58)	(0.80)	(0.92)	(1.05)	(1.06)	(1.05)	(1.12)	(1.32)	(1.39)	(1.49)
120m	-1.59***	-2.14***	-2.69***	-2.86***	-3.11***	-3.11***	-2.70***	-2.11*	-1.86	-1.62
	(0.57)	(0.79)	(0.91)	(1.03)	(1.02)	(0.99)	(0.99)	(1.14)	(1.20)	(1.30)
Panel B.ii: Dependent Variable: $rx^{FX}_{t,t+h}$, Coefficient on η^{R}_{κ}										
12m	0.03	0.06**								
	(0.02)	(0.03)								
24m	0.02	0.05*	0.06**	0.11***						
	(0.02)	(0.03)	(0.03)	(0.04)						
36m	0.02	0.05*	0.08**	0.13***	0.17***	0.19***				
	(0.02)	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)				
60m	0.02	0.05*	0.07**	0.11***	0.15***	0.18***	0.21***	0.22***	0.22***	0.22***
	(0.02)	(0.03)	(0.03)	(0.04)	(0.05)	(0.05)	(0.04)	(0.04)	(0.04)	(0.03)
84m	0.02	0.06**	0.07**	0.10***	0.15***	0.18***	0.19***	0.21***	0.21***	0.22***
	(0.02)	(0.03)	(0.03)	(0.04)	(0.05)	(0.05)	(0.04)	(0.04)	(0.04)	(0.04)
120m	0.02	0.05*	0.07**	0.11***	0.15***	0.19***	0.23***	0.26***	0.28***	0.30***
	(0.02)	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.03)	(0.03)
N	1,733	1,697	1,661	1,625	1,589	1,553	1,517	1,481	1,445	1,409

Notes: Coefficient estimates on the relative yield curve slope $S^{*}_{t}-S_{t}$ (Panel B.i) and cross-country κ-period liquidity yield η^{R}_{κ} (Panel B.ii) from regressions with the log currency excess return as dependent variable. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD for different samples within 1991:04-2017:12. The log returns and yield curve slopes differentials are annualised. All regressions include country fixed effects. All regressions include country fixed effects. The panels are unbalanced and standard errors (reported in parentheses) are constructed according to the Driscoll and Kraay (1998) methodology. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively. Because currency excess returns are invariant to bond maturity, and depend only on the holding period (unlike the dollar- and local currency-bond returns), we are able to present coefficient estimates on the relative slope and liquidity yield for all holding periods up to, and including, the bond maturity.

(1985) (CIR) model in which we impose parametric restrictions to match the three empirical regularities. We show that this model can generate the tent-shaped relationship, quantitatively aligning with our empirical results.

4.1 Pricing Kernels, Transitory Risk and the Yield Curve Slope

We consider a two-country environment, in which each country—Home (base currency, i.e. US) and Foreign (denoted by an asterisk)—has a representative investor. Throughout, we assume that investors can trade freely in both Home and Foreign risk-free bonds of multiple maturities.

Pricing kernels and stochastic discount factors. The Home nominal pricing kernel V_t represents the marginal value of a currency unit at time t. The nominal SDF $M_{t,t+\kappa}$ represents the growth rate of the pricing kernel between periods t and $t+\kappa$: $M_{t,t+\kappa}=V_{t+\kappa}/V_t$.

The price of a Home zero-coupon bond that promises one currency unit κ periods into the future is given by: $P_{t,\kappa}=\mathbb{E}_t\left[M_{t,t+\kappa}\right]=\mathbb{E}_t\left[M_{t,t+1}P_{t+1,\kappa-1}\right]$, where $M_{t,t+1}$ denotes the one-period SDF and $M_{t,t+\kappa}\equiv\prod_{i=0}^{\kappa-1}M_{t+i,t+i+1}$. Defining the gross return on the Home κ-period zero-coupon bond as $R_{t,\kappa} \equiv 1/P_{t,\kappa} \equiv (1 + r_{t,\kappa}) \geq 1$, this implies:

$$1 = \mathbb{E}_t \left[M_{t,t+\kappa} R_{t,\kappa} \right] \tag{8}$$

which can be rewritten as:

$$\frac{1}{R_{t,\kappa}} = \mathbb{E}_t \left[\prod_{i=0}^{\kappa-1} M_{t+i,t+i+1} \right] \tag{9}$$

Expressions for Foreign returns are analogous, denoted by an asterisk.

Exchange rates. $\mathcal{E}_t$ represents the exchange rate, defined as the Foreign price of a unit of Home currency, such that an increase in $\mathcal{E}_t$ corresponds to a Foreign depreciation. When engaging in cross-border asset trade, the Euler equation for a risk-averse Foreign agent with SDF $M^*_{t,t+\kappa}$ holding a κ-period Home currency-denominated bond is:

$$1 = \mathbb{E}_t \left[M^*_{t,t+\kappa} \frac{\mathcal{E}_{t+\kappa}}{\mathcal{E}_t} R_{t,\kappa} \right] \tag{10}$$

When financial markets are complete, the change in the nominal exchange rate corresponds to the ratio of SDFs:

$$\frac{\mathcal{E}_{t+\kappa}}{\mathcal{E}_t} = \frac{M_{t,t+\kappa}}{M^*_{t,t+\kappa}} \tag{11}$$

for all $\kappa > 0$. This no-arbitrage definition follows from equations (8) and (10).

Currency risk premia. Assuming one-period SDFs, $M_{t,t+1}$ and $M^*_{t,t+1}$, and the exchange rate, $\mathcal{E}_t$, are jointly log-normally distributed, the (log) one-period *ex ante* currency risk premium $\mathbb{E}_t[rx^{FX}_{t,t+1}]$ can be written as the half-difference between the conditional variance of (log) Home and Foreign SDFs:

$$\begin{aligned} \mathbb{E}_t \left[rx^{FX}_{t,t+1} \right] &= r^*_{t,1} - r_{t,1} - \mathbb{E}_t \left[\Delta^1 e_{t+1} \right] \\ &= \frac{1}{2} \left[\text{var}_t \left(m_{t,t+1} \right) - \text{var}_t \left(m^*_{t,t+1} \right) \right] \end{aligned} \tag{12}$$

where the second equality uses the logarithmic expansion of equations (8), its Foreign analog, and (11) when $\kappa = 1$.

Pricing kernel decomposition. To assess the nature of risks driving ERRP predictability, we use the Alvarez and Jermann (2005) decomposition of the pricing kernel V_t into a permanent component $V_t^{\mathbb{P}}$ and a transitory component $V_t^{\mathbb{T}}$:

$$V_t = V_t^{\mathbb{P}} V_t^{\mathbb{T}}, \quad \text{where } V_t^{\mathbb{T}} = \lim_{\kappa \to \infty} \frac{\delta^{t+\kappa}}{P_{t,\kappa}} \tag{13}$$

where the constant δ is chosen to satisfy the regularity condition: $0 < \lim_{\kappa \to \infty} P_{t,\kappa}/\delta^{\kappa} < \infty$ for all t. A pricing kernel V_t is defined as having only transitory innovations if $\lim_{\kappa \to \infty} \frac{\mathbb{E}_{t+1}[V_{t+\kappa}]}{\mathbb{E}_t[V_{t+\kappa}]} = 1$.

So, its permanent component follows a martingale, defined by: $V_t^{\mathbb{P}} = \lim_{\kappa\to\infty} \frac{\mathbb{E}_t[V_{t+\kappa}]}{\delta^{t+\kappa}}$.

Under regularity conditions, Alvarez and Jermann (2005) show that the return on an infinite-maturity bond can be written as a function of transitory innovations to SDFs only: $R_{t,\infty} = \lim_{\kappa\to\infty} R_{t,\kappa} = V_t^{\mathbb{T}}/V_{t+1}^{\mathbb{T}} = \exp(-m_{t,t+1}^{\mathbb{T}})$, where $m_{t,t+1}^{\mathbb{T}}$ denotes the transitory component of the SDF.[8] In contrast, one-period bond returns, defined by equation (8), depend on both transitory and permanent innovations to SDFs.

Yield curve slope and cyclical dynamics. To understand the role of the yield curve slope in capturing business-cycle risks, define the (log) excess return from buying a n-period Home bond at time t for price $P_{t,n} = 1/R_{t,n}$ and selling it at time $t+1$ for $P_{t+1,n-1} = 1/R_{t+1,n-1}$ as $rx_{t,t+1}^{(n)} = p_{t+1,n-1} - p_{t,n} - y_{t,1}$, where $p_{t,n} \equiv \log(P_{t,n})$ and $y_{t,n} \equiv -\frac{1}{n}p_{t,n}$ is the annualised yield on a n-period bond.[9] This excess return can be written as:

$$\mathbb{E}_t\left[rx_{t,t+1}^{(n)}\right] = -\text{cov}_t\left(m_{t,t+1}, \mathbb{E}_{t+1}\sum_{i=1}^{n-1} m_{t+i,t+i+1}\right) - \frac{1}{2}\text{var}_t\left(r_{t+1,n}\right) \tag{14}$$

The covariance term on the right-hand side is the bond risk premium, and is also equal to the covariance between the contemporaneous one-period SDF and the expected price or return on a long-term bond tomorrow, i.e. $p_{t+1,n-1}$ or $r_{t+1,n-1}$. The Foreign excess return is defined analogously.

Two features of the bond risk premium demonstrate that it is driven by business-cycle risk. First, the bond risk premium only captures transitory innovations to investors' SDFs. If the SDF is i.i.d., corresponding to the case of only permanent SDF innovations, the covariance in equation (14) is zero (see Example 1, Alvarez and Jermann, 2005).

Second, the premium reflects cyclicality of risk. It is positive if today's one-period SDF is negatively correlated with expected future marginal utility. That is, if households receive relatively good news about the distant future, they expect to value consumption less at long horizons—i.e. lower $\mathbb{E}_t[m_{t+i,t+i+1}]$ for some $i > 0$—but relatively highly in the near term—i.e. higher $m_{t,t+1}$. Specifically, we define risk to be conditionally cyclical if, conditional on shocks up to time t, investors expect a 'boom' to be followed by a 'bust', or vice versa.

Piazzesi and Schneider (2007) note that, over long enough samples, this risk premium is approximately equal to the yield curve slope, $\mathbb{E}_t[rx_{t,t+1}^{(n)}] \approx S_t$ where $S_t \equiv y_{t,n} - y_{t,1}$, implying that the yield curve will be upward sloping on average if the right-hand side of equation (14) is positive.[10] Taken together, this suggests that the yield curve slope captures business-cycle

[8] They further derive a lower bound for the volatility of $V_t^{\mathbb{P}}$, the martingale, and conclude that this permanent component accounts for the majority of SDF volatility.

[9] The annualised yield $y_{t,n}$ and the log n-period return $r_{t,n}$ have the following relationship: $ny_{t,n} = r_{t,n}$.

[10] To see this, re-write the excess return $rx_{t,t+1}^{(n)}$ as:

$$\begin{aligned} p_{t+1,n-1} - p_{t,n} - y_{t,1} &= ny_{t,n} - (n-1)y_{t+1,n-1} - y_{t,1} \\ &= y_{t,n} - y_{t,1} - (n-1)(y_{t+1,n-1} - y_{t,n}) \end{aligned}$$

Over a long enough sample and with large n, the difference between the average $(n-1)$-period yield and the average n-period yield is zero, implying that $\mathbb{E}_t[rx_{t,t+1}^{(n)}] \approx y_{t,n} - y_{t,1} \equiv S_t$.

risks, with parallels to the literature on recession predictability by the yield curve (Estrella and Hardouvelis, 1991; Estrella and Mishkin, 1998; Estrella, 2005). In turn, the relative yield curve slope $S_t^* - S_t$ reflects asymmetries in business-cycle risk across countries.

4.2 UIP Redux and Business-Cycle Risk

The previous sub-section demonstrated that transitory and conditionally cyclical risks are captured by the yield curve slope. Here, we investigate the characteristics of risk implied by exchange rate dynamics within a preference-free framework. Our starting point is three widely-documented empirical regularities: (i) the appreciation of high-yield currencies in excess of UIP at short horizons (Fama, 1984); (ii) the failure to reject UIP at long horizons (Chinn and Meredith, 2005) and for long-maturity bonds over short holding periods (Lustig et al., 2019); and (iii) the tendency for high interest rate currencies to be contemporaneously appreciated, i.e. UIP holding in 'levels' (Engel, 2016). We derive preference-free conditions corresponding to each of these regularities in turn, which we use in the next sub-section to calibrate a parametric model.

Fama puzzle. The first regularity is the failure of UIP at short horizons, i.e. the finding that $\hat{\beta}_{1,\kappa} < 1$ for small κ, and less than 0 for the smallest κ, in regression (1). By combining equations (1) for $\kappa = 1$ and the expression for ERRP (12), this empirical finding implies that $\text{cov}_t(\mathbb{E}_t[\Delta^1 e_{t+1}], r_{t,1}^* - r_{t,1}) < 0$ or $\text{cov}_t(\mathbb{E}_t[rx_{t,t+1}^{FX}], r_{t,1}^* - r_{t,1}) > 0$. In periods when the Foreign return is relatively high, $r_{t,1}^* > r_{t,1}$, the *ex ante* ERRP $\mathbb{E}_t[rx_{t,t+\kappa}^{FX}]$ is positive, which implies an excess appreciation of the Foreign currency *vis-à-vis* UIP. Considering equation (12), a positive covariance implies the following condition (Verdelhan, 2010):

Preference-free condition 1. *Short-term interest rates are higher (lower) in countries with less (more) volatile SDFs.*

Long-horizon UIP. The second regularity is the failure to statistically reject the UIP hypothesis at long horizons, implying that long-horizon ERRP are approximately zero on average: $\lim_{\kappa\to\infty} \mathbb{E}_t[rx_{t,t+\kappa}^{FX}] \approx 0$. Using the Alvarez and Jermann (2005) decomposition of pricing kernels, Lustig et al. (2019) show that this long-horizon ERRP is proportional to cross-country differences in the variance of permanent innovations to investors' SDFs:

$$\lim_{\kappa\to\infty} \mathbb{E}_t\left[rx_{t,t+\kappa}^{FX}\right] = \frac{1}{2}\left[\text{var}_t\left(\nu_{t+1}^{\mathbb{P}}\right) - \text{var}_t\left(\nu_{t+1}^{\mathbb{P}*}\right)\right] \approx 0 \tag{15}$$

where $\nu_t^{(*)} \equiv \log(V_t^{\mathbb{P}(*)})$. The relative success of long-horizon UIP therefore implies that cross-country differences in permanent SDF volatilities are small.[11]

[11] Alvarez and Jermann (2005) emphasise that to jointly rationalise high equity premia and low bond premia, most SDF volatility must arise from permanent SDF innovations. The contrast in exchange rate markets may be due to differential transmission and risk sharing across countries.

Instead, we focus on ERRP at short to medium horizons that must reflect transitory innovations to SDFs in light of equation (15):[12]

$$\mathbb{E}_t\left[rx^{FX}_{t,t+1}\right] \approx \frac{1}{2}\left[\text{var}_t\left(\nu^{\mathbb{T}}_{t+1}\right) - \text{var}_t\left(\nu^{\mathbb{T}*}_{t+1}\right)\right] \neq 0 \qquad \text{(Short-Horizon ERRP)}$$

$$\mathbb{E}_t\left[rx^{FX}_{t,t+\kappa}\right] = r^*_{t,\kappa} - r_{t,\kappa} - \mathbb{E}_t\left[\Delta^\kappa e_{t+\kappa}\right] \neq 0 \quad \text{for } 1 < \kappa < \infty \qquad \text{(Medium-Horizon ERRP)}$$

Taken together, our findings for short and long-horizon ERRP imply the following condition:

Preference-free condition 2. *ERRP are predominantly driven by transitory risk. Exposure to permanent risk across countries must be approximately symmetric.*

UIP in levels. The third regularity is that high-yield currencies tend to be contemporaneously appreciated, i.e. UIP holds 'in levels'. The relative success of UIP at long horizons is not a sufficient condition for UIP to also hold in levels, unless all term premia are zero. Engel (2016) shows that, to achieve this, the covariance between the current short-term interest rate differential and future one-period exchange rate moves must switch sign across horizons: $\text{cov}_t(\mathbb{E}_t[\Delta^1 e_{t+1}], r^*_{t,1} - r_{t,1}) < 0$, but $\text{cov}_t(\mathbb{E}_t[\Delta^1 e_{t+i}], r^*_{t,1} - r_{t,1}) > 0$ for some $i > 1$. For example, an increase in interest rates will initially be associated with an exchange rate appreciation, before a subsequent depreciation.

Preference-free condition 3. *The covariance of short-term interest rate differentials and future one-period exchange rate changes (and ERRP) must change sign at some horizon.*

Preference-free conditions 1 and 3 rely on the covariance between current interest rates (which are functions of SDFs) and future exchange rate movements (which reflect differences in the conditional mean of future SDFs). Absent additional restrictions, we cannot directly draw a connection to the bond risk premium (14) and the covariance of current and future SDFs. Nevertheless, the cyclicality of risk is a driving factor in reproducing 1 and 3, as we show in a parametric factor model in the next sub-section.

Engel (2016) emphasises that condition 3 evidences multiple factors driving exchange rates. In a regression setting, this implies that interest rate differentials are not sufficient predictors of exchange rate movements, and a second factor will be a significant explanatory variable. Given the argument that the yield curve slope is driven by transitory and cyclical risks, then the relative yield curve slope is a natural candidate driver for this second factor—reflecting cross-country asymmetries in business-cycle risk.

Finally, we note that our empirical finding that the relative yield curve slope covaries negatively with the one-period ERRP can be set against the aforementioned empirical regularities. When combined with the positive covariance between the one-period ERRP and relative short-rates underlying preference-free condition 1, this implies that long-rate differentials must covary

[12] By construction: $\text{var}_t(\nu_{t+1}) = \text{var}_t(\nu^{\mathbb{T}}_{t+1}) + \text{var}_t(\nu^{\mathbb{P}}_{t+1})$.

less strongly with one-period ERRP than short rates.[13] In turn, this implies that long-yield differentials respond less to shocks that drive SDF variation than short-rate differentials.

4.3 Two-Country Cox, Ingersoll and Ross Model

We now present a calibrated two-country CIR model. Under parametric restrictions which we derive, and correspond to preference-free conditions 1-3, we show it can reproduce the tent-shaped relationship across horizons between the relative yield curve slope and ERRP.

The representative Home investor's SDF loads on two country-specific factors $z_{i,t}$ $(i = 1, 2)$:

$$-m_{t,t+1} = \alpha + \chi z_{1,t} + \sqrt{\gamma z_{1,t}} u_{t+1} + \tau z_{2,t} + \sqrt{\delta z_{2,t}} u_{2,t+1} \tag{16}$$

$$z_{i,t+1} = (1 - \phi_i)\theta_i + \phi_i z_{i,t} - \sigma_i \sqrt{z_{i,t}} u_{i,t+1} \quad \text{for } i = 1, 2 \tag{17}$$

We assume that the representative Foreign investor's SDF $m^*_{t,t+1}$ and country-specific pricing factors $z^*_{i,t}$ are defined analogously, and with symmetric loadings ($\alpha^* = \alpha$, $\chi^* = \chi$, $\gamma^* = \gamma$, $\tau^* = \tau$, $\delta^* = \delta$, $\phi^*_i = \phi_i$, $\theta^*_i = \theta_i$ and $\sigma^*_i = \sigma_i$ for $i = 1, 2$).

Assuming log-normality, then equations (16) and (17) can be combined with the expression for the (log) price of an n-period bond, $p_{t,n} = \mathbb{E}_t[m_{t,t+1} + p_{t+1,n-1}] + (1/2)\text{var}_t(m_{t,t+1} + p_{t+1,n-1})$, to write the the (log) bond price as an affine function of pricing factors:

$$p_{t,n} = -\left(A_n + B_n z_{1,t} + C_n z_{2,t}\right) \tag{18}$$

where A_n, B_n and C_n are recursively defined, with $A_n \equiv A_n(\alpha, \phi_1, \phi_2, \theta_1, \theta_2; A_{n-1}, B_{n-1}, C_{n-1})$, $B_n \equiv B_n(\phi_1, \chi, \gamma, \sigma_1; B_{n-1})$ and $C_n \equiv C_n(\phi_2, \tau, \delta, \sigma_2; C_{n-1})$ and initial values of 0 ($A_0 = B_0 = C_0 = 0$). The continuously compounded yield is $y_{t,n} = -\frac{1}{n} p_{t,n}$. Expressions for Foreign bond prices and yields are analogous.

We derive the following parametric counterparts to preference-free conditions 1-3 that the two-country CIR model must satisfy to replicate the three empirical regularities discussed in the previous sub-section.

Lemma (Transitory and Conditionally Cyclical Risks) *We define risk in the two-country CIR model to be transitory and conditionally cyclical if the following three parametric restrictions are met:*

(i) To match short-run UIP failures, we ensure: $\chi - \frac{1}{2}\gamma < 0$ *and* $\chi(\chi - \frac{1}{2}\gamma)var_t(z^*_{1,t} - z_{1,t}) < -\tau(\tau - \frac{1}{2}\delta)var_t(z^*_{2,t} - z_{2,t})$ *for* $\chi, \tau > 0$.

(ii) To match long-run UIP approximately by ruling out permanent innovations to investors' SDFs, we ensure: $(B_n(1 - \phi_1) - \chi)\theta_1 + (C_n(1 - \phi_2) - \tau)\theta_2 \approx 0$ *as* $n \to \infty$.

[13] To see this, note that $S_t = y_{t,L} - r_{t,1}$ for L large. Then $\text{cov}_t(\mathbb{E}_t[rx^{FX}_{t,t+1}], S^*_t - S_t) < 0$, can be written as: $\text{cov}_t(\mathbb{E}_t[rx^{FX}_{t,t+1}], y^*_{t,L} - y_{t,L}) - \text{cov}_t(\mathbb{E}_t[rx^{FX}_{t,t+1}], r^*_{t,1} - r_{t,1}) < 0$. Given also $\text{cov}_t(\mathbb{E}_t[rx^{FX}_{t,t+1}], r^*_{t,1} - r_{t,1}) > 0$ underpinning preference-free condition 1, it must be that $\text{cov}_t(\mathbb{E}_t[rx^{FX}_{t,t+1}], r^*_{t,1} - r_{t,1}) > \text{cov}_t(\mathbb{E}_t[rx^{FX}_{t,t+1}], y^*_{t,L} - y_{t,L}) > 0$.

(iii) To replicate UIP in levels approximately and generate conditional cyclicality, we require that the second factor is sufficiently more persistent than than the first (i.e. $\phi_2 >> \phi_1$) and ensure $\tau - \frac{1}{2}\delta > 0$.

Sketch Proof. Conditions (i) and (iii) ensure that the UIP coefficient is negative at short horizons and positive at long horizons:

$$\beta_h^{UIP} = \frac{\chi \frac{1-\phi_1^h}{1-\phi_1}(B_h)\text{var}_t(z_{1,t}^* - z_{1,t}) + \tau \frac{1-\phi_2^h}{1-\phi_2}(C_h)\text{var}_t(z_{2,t}^* - z_{2,t})}{(B_h)^2\text{var}_t(z_{1,t}^* - z_{1,t}) + (C_h)^2\text{var}_t(z_{2,t}^* - z_{2,t})} \tag{19}$$

Condition (ii) ensures that β_h^{UIP} is close to unity. See Appendix C.1 for a full proof. □

The short-run UIP coefficient (19) is negative if and only if condition (i) holds. In a one-factor version of the model (i.e. $z_{2,t} = 0$), $\chi - \frac{1}{2}\gamma < 0$ would generate short-term interest rates which co-move positively with the SDF, i.e. $\text{cov}_t(m_{t,t+1}, y_{t,1}) > 0$.[14] However, in that model this condition would also generate positive co-movement between long-term interest rates and SDFs, i.e. $\text{cov}_t(m_{t,t+1}, y_{t,n}) > 0$. In turn, this would result in a counterfactually negative yield curve slope on average, and negative long-horizon UIP coefficients.

Within the two-factor CIR model, it is the cyclicality of risk that allows the model to match UIP at short and long horizons. Specifically, the inclusion of a second factor under condition (iii) ensures that long-term interest rates co-move negatively with the SDF, i.e. $\text{cov}_t(m_{t,t+1}, y_{t,n}) < 0$, delivering a positive average yield curve slope.[15] Using equation (16), the bond risk premium can be expressed as:

$$-\text{cov}_t(p_{t+1,n-1}, m_{t,t+1}) = B_{n-1}\sigma_1\sqrt{\gamma}z_{1,t} + C_{n-1}\sigma_2\sqrt{\delta}z_{2,t} \tag{20}$$

If condition (i) holds, implying $B_{n-1} < 0$, the bond risk premium is positive only if the second factor is sufficiently more persistent than the first, i.e. $\phi_2 >> \phi_1$, and $C_{n-1} > 0$. Conditions (i) and (iii) therefore simultaneously generate cyclical SDF dynamics, and result in a sign-switch in UIP coefficients across horizons within the model (19).

Condition (ii) ensures that risk driving SDFs is predominantly transitory, i.e. $\text{var}(m_{t,t+1}) \approx \text{var}(m_{t,t+1}^{\mathbb{T}}) > 0$.[16] This applies Condition 1 in Lustig et al. (2019) to a two-factor CIR model.

We next demonstrate that when imposed jointly, the conditions in the Lemma generate a tent-shaped relationship between expected exchange rate movements and the relative yield curve slope. The following proposition states this for model-implied univariate regressions.

[14] Verdelhan (2010) shows that this is sufficient to match short-run UIP failures.

[15] While two-factor multi-country CIR models have been discussed in the literature, (e.g. Lustig et al., 2019, Appendix IV.C) our paper is (to the best of our knowledge) the first to analyse the implications of such a model and consider the role for the term structure of interest rates in exchange rate determination. Away from our CIR setting, Engel (2016) presents a stylised two-country New-Keynesian model with productivity and liquidity risk. These two sources of risk have parallels to the two factors we consider.

[16] Lustig et al. (2019) derive that long-run UIP holds only if countries are symmetrically exposed to permanent risk. In a model with country-specific factors, this can only be achieved when permanent risk is zero.

Proposition *The two-country CIR model can reproduce a tent-shaped relationship between the expected exchange rate movements* $\mathbb{E}_t[e_{t+\kappa} - e_t]$ *and the relative slope* S_t^R *across horizons* κ *if and only if risk is both transitory and conditionally cyclical.*

Proof. See Appendix C.2. □

The relationship between the relative yield curve slope and exchange rates across horizons summarised in the Proposition relies on differences in investors' valuations of returns over the business cycle. Suppose the Foreign country has a relatively steep yield curve. The representative Foreign investor will value returns more highly in the near term, but expect their valuations to decrease over time, as equation (14) demonstrates when yield curves are upward sloping on average. Since the yield curve is mechanically linked to short-term yields, the Foreign country will generally have a relatively low short-term interest rate and will experience a currency depreciation as compensation for exchange-rate risk, consistent with short-run UIP failures. Therefore, the relationship between the relative slope and exchange rate changes is positive at short horizons.

However, cross-country return valuations will reverse as investors move along the cycle. The Foreign investor, who formerly valued returns highly in a (comparative) bust, value them less as they move into a (relative) boom. These SDF dynamics are reflected in the path of expected relative future short-term interest rates: $\mathbb{E}_t[r^*_{t+i,1} - r_{t+i,1}] = B_1(\mathbb{E}_t[z^*_{1,t+i} - z_{1,t+i}]) + C_1(\mathbb{E}_t[z^*_{2,t+i} - z_{2,t+i}])$ across i, derived from cross-country differences in equation (18) at $n = 1$. For the Foreign country to have relatively low short-term yields: $z^*_{1,t} > z_{1,t}$ as $B_1 < 0$. And due to the higher Foreign slope: $z^*_{2,t} > z_{2,t}$ as $C_1 > 0$. Since $\phi_2 > \phi_1$, the influence of cross-country differences in the second factor on expected relative short-yield differentials will dominate at longer horizons. Consequently, expected future short-term yield differentials will be increasing for small i, as cross-country differences in the first factor matter at short horizons, but dissipate more quickly. For larger i, the value of the differential will be decreasing, as differences in the second factor persist, and the exchange rate will begin to appreciate—again consistent with short-run UIP failures. In sum, the currency of the Foreign country—with the relatively steep yield curve—will depreciate in the near-term. The depreciation will peak at business-cycle horizons, before pressure on the currency to appreciate builds to form the tent-shaped relationship we observe in the data.

There is a close relationship between UIP across horizons and our finding for the relative yield curve slope. The benefit of analysing the relative yield curve slope, as opposed to long-term yield differentials, is that the former is reflects cross-country business-cycle risk, clarifying the underlying drivers of exchange rate dynamics. As a result, business-cycle risk may also explain the switch in the UIP coefficient across horizons.

Numerical calibration. We now calibrate the two-country CIR model to satisfy the conditions in the Lemma, as well as match other empirical moments. We demonstrate that it can quantitatively reproduce the tent-shaped relationship between exchange rate changes and the relative slope that we identify in the data.

We target 11 moments from the data: the mean and variance of the short and long-term interest rates; the autocorrelation of short-term interest rates; the variance of SDFs; two Feller conditions that help ensure $z_{1,t}$ and $z_{2,t}$ remain positive; short and long-horizon UIP coefficients; and the peak coefficient implied by a univariate regression of exchange rate changes on the relative slope across horizons. These calibration targets are summarised in Table 6, and pin down the 11 model parameters. The parameter values we obtain are listed in Table 7.[17]

The calibration identifies key differences between the two pricing factors, whose loadings differ in sign and persistence. Because $\chi - (1/2)\gamma = -0.30 < 0$, bond yields load negatively on the first factor. In contrast, $\tau - (1/2)\delta = 0.80 > 0$, such that bond yields load positively on the second. Consistent with the Lemma, the first factor is less persistent than the second ($\phi_2 = 0.99 > \phi_1 = 0.95$). This ensures that the numerical exercise features transitory and conditionally cyclical risks, which helps to match the three empirical regularities detailed in the previous sub-section. However, in the data, although long-run UIP at the 10-year horizon cannot be statistically rejected, the point estimate is below 1. Therefore, SDF loadings on permanent risk are not perfectly symmetric in our calibration.

The key results from the calibration exercise are plotted in Figure 4. The left-hand chart shows the model-implied UIP coefficient across horizons, alongside empirical estimates (and 95% confidence bands) from equation (1).[18] Between the two calibrated UIP regression coefficients—at $h = 1$ and $h = 120$—the model-implied values broadly lie within the estimated confidence bands. To the best of our knowledge, this is the first multi-country CIR model to quantitatively replicate UIP coefficients across horizons.

The right-hand chart plots the model-implied coefficient from a univariate regression of exchange rate changes across horizons on the relative yield curve slope. Corresponding empirical estimates are presented alongside.[19] Although the model is calibrated to match the empirical estimates at one point only—i.e. $h = 36$ months—the model-implied coefficients have a tent shape across horizons that broadly lie within the estimated confidence bands. As in the data, the relationship between exchange rate changes and the relative slope is small at short horizons, peaks at business-cycle horizons, and becomes small (or even negative) at longer horizons.

We have chosen a symmetric two-country CIR model with country-specific factors for its simplicity and analytical transparency, as opposed to a model with asymmetric loadings on a common factor (see, e.g., Lustig et al., 2014). This comes with two main costs. First, our analysis is conditional, because moments with cross-country terms, e.g. implied regression coefficients, will have an unconditional mean of zero.[20] Second, in the absence of common factors, the cross-country correlation of interest rates is counterfactually zero. However, the mechanisms we discuss, and the role of business-cycle risks for exchange rate fluctuations would

[17] Since we limit our focus to assessing the ability of the model to replicate the empirical regularities we outline in this paper, it is beyond our scope to evaluate the model's performance against other unmatched moments.

[18] For this figure, we omit country fixed effects from the estimation to be consistent with model-implied values.

[19] The empirical estimates documented in this figure come from a univariate pooled OLS regression of exchange rate changes on the relative yield curve slope to be consistent with model-implied values.

[20] For example, consider the conditional ERRP: $\mathbb{E}_t[rx^{FX}_{t,t+1}] = (1/2)(\gamma(z^*_{1,t} - z_{1,t}) + \delta(z^*_{2,t} - z_{2,t}))$. The unconditional ERRP is equal to 0 in the symmetric, country-specific factor setup.

Table 6: Targeted moments for two-country CIR model

Moment	Analytical Expression	Target (Annual)
$\mathbb{E}[r_{t,1}]$	$\alpha + (\chi - \frac{1}{2}\gamma)\theta_1 + (\tau - \frac{1}{2}\delta)\theta_2$	0.40% (4.81%)
$\text{std}(r_{t,1})$	$\sqrt{(\chi - \frac{1}{2}\gamma)^2 \text{var}(z_{1,t}) + (\tau - \frac{1}{2}\delta)^2 \text{var}(z_{2,t})}$	0.31 (3.74)
$\rho(r_{t,1})$	$\frac{(\chi - \frac{1}{2}\gamma)^2 \text{var}(z_{1,t})\phi_1 + (\tau - \frac{1}{2}\delta)^2 \text{var}(z_{2,t})\phi_2}{(\chi - \frac{1}{2}\gamma)^2 \text{var}(z_{1,t}) + (\tau - \frac{1}{2}\delta)^2 \text{var}(z_{2,t})}$	0.99
$\text{std}(m_{t,t+1})$	$\sqrt{\chi^2 \text{var}(z_{1,t}) + \tau^2 \text{var}(z_{2,t}) + \gamma\theta_1 + \delta\theta_2}$	14% (50%, Sharpe ratio)
β_1^{UIP}	$\frac{\chi(\chi - \frac{1}{2}\gamma)\text{var}_t(z_{1,t}^* - z_{1,t}) + \tau(\tau - \frac{1}{2}\delta)\text{var}_t(z_{2,t}^* - z_{2,t})}{(\chi - \frac{1}{2}\gamma)^2 \text{var}_t(z_{1,t}^* - z_{1,t}) + (\tau - \frac{1}{2}\delta)^2 \text{var}_t(z_{2,t}^* - z_{2,t})}$	-0.5 (Short-run (1-month) UIP)
β_{120}^{UIP}	$\frac{\chi \frac{1-\phi_1^{120}}{1-\phi_1}(B_{120})\text{var}_t(z_{1,t}^* - z_{1,t}) + \tau \frac{1-\phi_2^{120}}{1-\phi_2}(C_{120})\text{var}_t(z_{2,t}^* - z_{2,t})}{(B_{120})^2 \text{var}_t(z_{1,t}^* - z_{1,t}) + (C_{120})^2 \text{var}_t(z_{2,t}^* - z_{2,t})}$	0.8 (Long-run (10-year) UIP)
$\mathbb{E}[r_{t,120}]$	$\frac{1}{120}\left[A_{120} + B_{120}\theta_{1,t} + C_{120}\theta_{2,t}\right]$	0.52% (6.29%)
$\text{std}(r_{t,120})$	$\sqrt{\left(\frac{1}{120}B_{120}\right)^2 \text{var}(z_{1,t}) + \left(\frac{1}{120}C_{120}\right)^2 \text{var}(z_{2,t})}$	0.26 (3.14)
$\beta_{36}^{S^R}$	$\frac{\frac{1-\phi_1^{36}}{1-\phi_1}\chi\left[\frac{1}{120}B_{120} - \frac{1}{6}B_6\right]\text{var}_t(z_{1,t}^* - z_{1,t}) + \frac{1-\phi_2^{36}}{1-\phi_2}\tau\left[\frac{1}{120}C_{120} - \frac{1}{6}C_6\right]\text{var}_t(z_{2,t}^* - z_{2,t})}{\left[\frac{1}{120}B_{120} - \frac{1}{6}B_6\right]^2 \text{var}_t(z_{1,t}^* - z_{1,t}) + \left[\frac{1}{120}C_{120} - \frac{1}{6}C_6\right]^2 \text{var}_t(z_{2,t}^* - z_{2,t})}$	2.74 (36-month Slope)
Feller Factor 1	$2(1-\phi_1)\theta_1/\text{var}(z_{1,t})$	20
Feller Factor 2	$2(1-\phi_2)\theta_2/\text{var}(z_{2,t})$	20

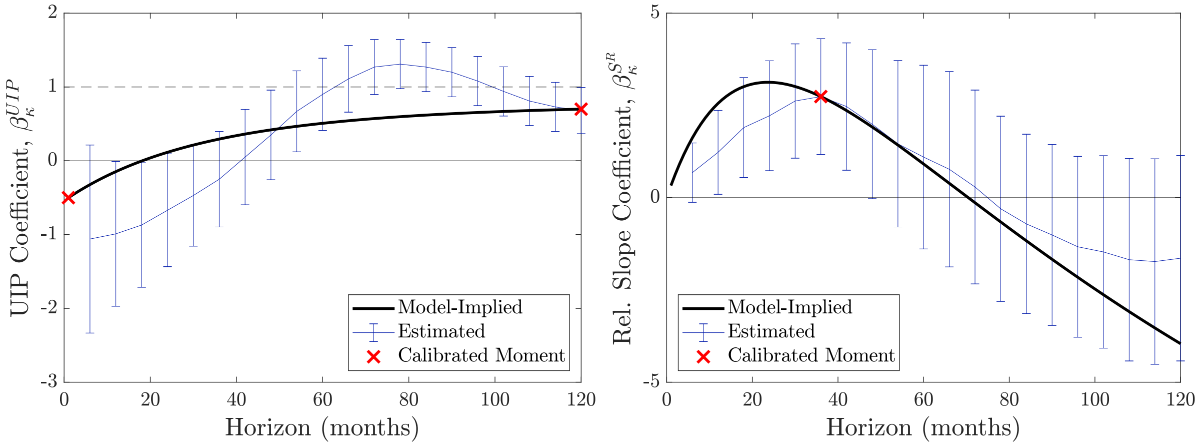
Notes: Monthly calibration with pooled G7 interest rates. Sharpe ratio from Lustig et al. (2014). We differentiate between the conditional variance, $\text{var}_t(z_{i,t}) = \sigma_i^2\theta_i$, and unconditional, $\text{var}(z_{i,t}) = \frac{\sigma_i^2\theta_i}{1-\phi_i^2}$.

Table 7: Calibrated parameters from two-country CIR model

α	χ	γ	τ	δ	
-0.0165	0.971	2.534	1.085	0.567	

θ_1	σ_1	ϕ_1	θ_2	σ_2	ϕ_2
0.0018	0.0239	0.946	0.0263	0.0029	0.993

Figure 4: Implied regression coefficients from two-country CIR model across horizons and comparable estimated coefficients



Notes: Black line denotes model-implied conditional regression coefficients across horizons. Red crosses denote calibration targets for the model. Blue line plots estimated coefficients using pooled monthly data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD from 1980:01 to 2017:12, excluding country fixed effects. 95% confidence intervals, calculated using Driscoll and Kraay (1998) standard errors, are denoted by blue bars around point estimates. Left-hand (right-hand) plot shows coefficients for univariate regression of κ-period exchange rate on return differentials (relative yield curve slope).

carry over to a more complex model with common factors, where our results would then hinge on asymmetries in factors loadings across countries.

5 Conclusion

In this paper, we explore the relationship between the term structure of interest rates and ERRP, both empirically and theoretically. Empirically, our main finding is that a country with a relatively steep yield curve will tend to depreciate at business-cycle horizons, even when controlling for bond liquidity yields. We find a tent-shaped relationship between ERRP and the relative yield curve slope across horizons, which peaks at around 3 to 5 years.

This relationship is consistent with a no-arbitrage framework. For this to be the case, risks driving exchange rates must, at least in part, be transitory and conditionally cyclically, suggesting a role for business-cycle risks. We show that a two-country, two-factor model for interest rates, calibrated to reflect business-cycle risk, can quantitatively match the tent-shaped relationship for the relative slope that we observe in the data, as well as UIP coefficient estimates across horizons. The no-arbitrage framework can also reproduce three other well-documented empirical regularities for exchange rates, namely: the failure of UIP at short horizons, the failure to reject UIP at long horizons, and UIP holding in levels.

Appendix

A Data Sources

We use nominal zero-coupon government bond yields at maturities from 6 months to 10 years for 7 industrialised countries: US, Australia, Canada, euro area, Japan, Switzerland and UK. Our benchmark sample spans 1980:01-2017:12, although the panel of interest rates is unbalanced as bond yields are not available from the start of the sample in all jurisdictions. Table 8 summarises the sources of nominal zero-coupon government bond yields, and the sample availability, for the benchmark economies in our study. In robustness analyses, we also assess results for a broader set of G10 currencies—adding New Zealand, Norway and Sweden—for which zero-coupon government bond yields are available up to 2009:05 from Wright (2011).

Table 8: Yield Curve Data Sources

Country	Sources	Start Date
US	Gürkaynak, Sack, and Wright (2007)	1971:11
Australia	Reserve Bank of Australia	1992:07
Canada	Bank of Canada	1986:01
Euro Area	Bundesbank (German Yields)	1980:01
Japan	Wright (2011) and Bank of England	1986:01
Switzerland	Swiss National Bank	1988:01
UK	Anderson and Sleath (2001)	1975:01

Notes: Data from before 1980:01 are not used in this paper.

Exchange rate data is from *Datastream*, reflecting end-of-month spot rates *vis-à-vis* the US dollar. Liquidity yields are from Du et al. (2018), available at the 1, 2, 3, 5, 7 and 10-year maturities. The earliest liquidity yields are available from 1991:04 for some countries (e.g. UK). The latest liquidity yields are available from 1999:01 (e.g. euro area). For both exchange rates and liquidity yields, we use end-of-month observations.

B Empirical Results

B.1 Full Results from UIP and Yield Curve-Augmented UIP Regressions

Table 9 presents our benchmark results for regressions (1) and (2). Column (1) presents the $\beta_{1,\kappa}$ estimates, at different horizons, from the canonical UIP panel regression using pooled monthly data from 1980:01 to 2017:12. Columns (2)-(4) present the $\beta_{1,\kappa}$, $\beta_{2,\kappa}$ and $\beta_{3,\kappa}$ estimates at different horizons from the yield curve-augmented regression.

B.2 Robustness Results for Yield Curve-Augmented UIP Regression

In this Appendix, we report results for the robustness exercises discussed in Section 2.3.

Predictability of interest rates. Table 10 presents results of regressions of exchange rate changes on: (a) the relative yield curve slope and curvature only; and (b) the relative yield curve level, slope and curvature. These specifications differ from our baseline specification by omitting the interest rate differential. In both cases, the tent-shaped pattern of coefficients on the relative slope remains significant.

In specification (b), we proxy the yield curve level using the difference between 10-year zero-coupon yields $L_{j,t} = i_{j,t,10y}$. This specification replicates that in Chen and Tsang (2013). However, our results differ due to differences in the construction of yield curve factors. Chen and Tsang (2013) capture relative yield curve factors by directly estimating Nelson-Siegel decompositions on *relative* interest rate differentials. To do this, they assume symmetry of factor structures across countries. We, instead, construct proxies for factors using yield curves estimated on a country-by-country basis, and so do not assume such symmetry.

Long-horizon inference. As discussed in Section 2.3, long-horizon forecasting regressions like (1) and (2) can face size distortions, whereby the null hypothesis is rejected too often. Valkanov (2003) demonstrates that this problem is especially pertinent when samples are small and regressors are persistent. Although the Driscoll and Kraay (1998) standard errors used in the main body of the paper are robust to heteroskedasticity and autocorrelation, we assess the robustness of our findings using alternative inference here.

Following Moon et al. (2004), we use scaled t-statistics, whereby standard t-statistics are multiplied by $1/\sqrt{\kappa}$. In the context of long-horizon forecasting regressions like ours, Moon et al. (2004) demonstrate that these scaled t-statistics are approximately standard normal when regressors are sufficiently persistent. However, because the scaled t-statistics can tend to under-reject the null when regressors are not near-integrated, we view these t-statistics as providing more conservative inference than the Driscoll and Kraay (1998) standard errors.

Figure 5 plots the $\beta_{2,\kappa}$ estimates from (2) with 90% confidence intervals implied by the scaled t-statistics of Moon et al. (2004). Relative to table 9, point estimates are unchanged. But the error bands implied by the scaled t-statistics are wider from 12 months onwards. Nevertheless,

Table 9: Coefficient estimates from canonical UIP regression and regression augmented with relative yield curve slope and curvature

Maturity	(1) UIP Regression	(2) Yield Curve Augmented Regression	(3)	(4)
κ	$r^*_\kappa - r_\kappa$	$r^*_\kappa - r_\kappa$	$S^* - S$	$C^* - C$
6-months	-1.06	-0.40	0.75	-0.61
	(0.65)	(1.00)	(0.70)	(0.74)
12-months	-0.99**	-0.22	1.41	-0.82
	(0.50)	(0.82)	(1.14)	(1.09)
18-months	-0.87**	0.29	2.87**	-1.25
	(0.43)	(0.69)	(1.31)	(1.23)
24-months	-0.67*	0.60	4.31***	-2.45
	(0.39)	(0.62)	(1.50)	(1.53)
30-months	-0.47	0.94*	5.98***	-3.67**
	(0.35)	(0.56)	(1.60)	(1.77)
36-months	-0.25	1.11**	6.74***	-4.13**
	(0.33)	(0.52)	(1.63)	(1.74)
42-months	0.05	1.31***	7.40***	-5.11***
	(0.33)	(0.44)	(1.61)	(1.86)
48-months	0.35	1.39***	7.04***	-4.89**
	(0.31)	(0.35)	(1.68)	(2.03)
54-months	0.67**	1.53***	6.63***	-4.51**
	(0.28)	(0.28)	(1.83)	(2.20)
60-months	0.90***	1.60***	5.98***	-3.66
	(0.25)	(0.27)	(1.97)	(2.31)
66-months	1.11***	1.64***	4.91**	-2.06
	(0.23)	(0.26)	(2.03)	(2.37)
72-months	1.27***	1.64***	3.61*	-0.52
	(0.19)	(0.23)	(1.93)	(2.21)
78-months	1.31***	1.55***	2.54	-0.06
	(0.17)	(0.21)	(1.77)	(2.09)
84-months	1.27***	1.42***	1.89	-0.30
	(0.17)	(0.19)	(1.65)	(2.10)
90-months	1.20***	1.28***	0.93	0.32
	(0.17)	(0.18)	(1.60)	(2.07)
96-months	1.08***	1.10***	-0.06	0.90
	(0.17)	(0.16)	(1.68)	(2.24)
102-months	0.94***	0.93***	-0.41	0.63
	(0.17)	(0.16)	(1.74)	(2.25)
108-months	0.81***	0.78***	-0.71	0.25
	(0.17)	(0.16)	(1.83)	(2.31)
114-months	0.73***	0.70***	-0.88	0.20
	(0.17)	(0.16)	(1.89)	(2.50)
120-months	0.68***	0.65***	-0.42	-0.79
	(0.16)	(0.16)	(1.66)	(2.34)

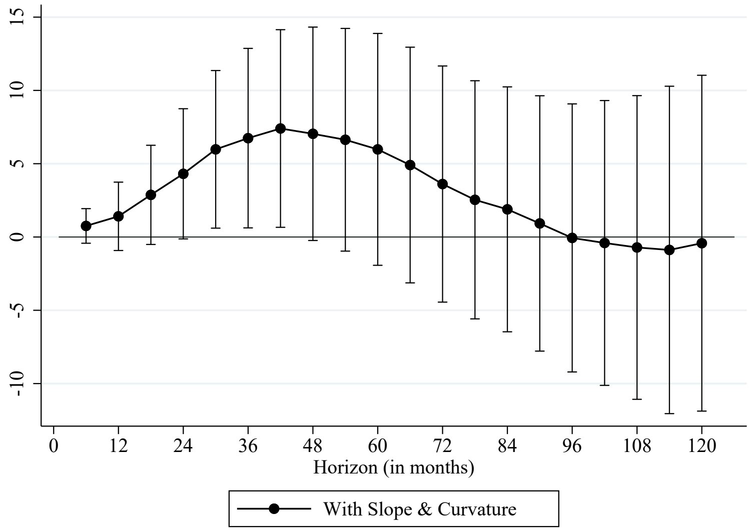
Notes: Column (1) presents coefficient estimates from regression (1)—the canonical UIP regression—a regression of the κ-period exchange rate change $\Delta^\kappa e_{t+\kappa}$ on the κ-period return differential $r^*_{t,\kappa} - r_{t,\kappa}$. Columns (2)-(4) document point estimates from (2)—the augmented regression—using the relative yield curve slope and curvature (measured using proxies) as additional regressors. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD from 1980:01 to 2017:12, including country fixed effects. The panel is unbalanced. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively, using Driscoll and Kraay (1998) standard errors (reported in parentheses).

Table 10: Coefficient estimates from regressions of exchange rate change on relative slope and curvature, and relative level, slope and curvature and regression augmented with relative yield curve slope and curvature

Maturity	(1) Slope & Curvature	(2)	(3) Level, Slope & Curvature	(4)	(5)
κ	$S^* - S$	$C^* - C$	$L^* - L$	$S^* - S$	$C^* - C$
6-months	1.00*	-0.67	-0.22	0.94*	-0.61
	(0.58)	(0.78)	(0.47)	(0.56)	(0.75)
12-months	1.68**	-0.93	-0.20	1.62**	-0.88
	(0.84)	(1.16)	(0.77)	(0.82)	(1.11)
18-months	2.37**	-0.97	0.37	2.46***	-1.04
	(0.98)	(1.33)	(0.99)	(0.93)	(1.25)
24-months	2.97***	-1.56	1.16	3.26***	-1.79
	(1.15)	(1.69)	(1.19)	(1.07)	(1.58)
30-months	3.51***	-1.84	2.27*	4.05***	-2.25
	(1.23)	(1.99)	(1.32)	(1.12)	(1.85)
36-months	3.51***	-1.59	3.22**	4.25***	-2.12
	(1.15)	(1.91)	(1.45)	(1.04)	(1.74)
42-months	3.31**	-1.75	4.35***	4.28***	-2.39
	(1.30)	(1.99)	(1.43)	(1.16)	(1.80)
48-months	2.48	-1.03	5.21***	3.62**	-1.76
	(1.59)	(2.19)	(1.31)	(1.42)	(1.99)
54-months	1.48	-0.03	6.38***	2.87*	-0.91
	(1.87)	(2.40)	(1.18)	(1.66)	(2.18)
60-months	0.57	1.10	7.35***	2.14	0.14
	(2.03)	(2.54)	(1.23)	(1.79)	(2.30)
66-months	-0.56	2.76	8.32***	1.16	1.76
	(2.12)	(2.64)	(1.32)	(1.84)	(2.34)
72-months	-1.71	4.14*	9.11***	0.13	3.13
	(2.05)	(2.49)	(1.29)	(1.73)	(2.15)
78-months	-2.24	4.02*	9.37***	-0.44	3.13
	(1.94)	(2.39)	(1.26)	(1.60)	(2.04)
84-months	-2.22	3.10	9.23***	-0.52	2.34
	(1.88)	(2.43)	(1.26)	(1.55)	(2.10)
90-months	-2.49	3.00	8.95***	-0.92	2.39
	(1.88)	(2.37)	(1.24)	(1.55)	(2.08)
96-months	-2.78	2.93	8.22***	-1.36	2.38
	(1.99)	(2.53)	(1.21)	(1.68)	(2.26)
102-months	-2.52	2.10	7.39***	-1.25	1.59
	(2.06)	(2.53)	(1.26)	(1.76)	(2.28)
108-months	-2.30	1.23	6.57***	-1.20	0.81
	(2.13)	(2.57)	(1.32)	(1.85)	(2.35)
114-months	-2.12	0.77	6.22***	-1.12	0.45
	(2.21)	(2.75)	(1.40)	(1.91)	(2.52)
120-months	-1.34	-0.58	6.07***	-0.44	-0.82
	(1.96)	(2.56)	(1.51)	(1.66)	(2.33)

Notes: Columns (1)-(2) presents coefficient estimates from regression of the κ-period exchange rate change $\Delta^{\kappa} e_{t+\kappa}$ on the relative yield curve slope $S^* - S$ and curvature $C^* - C$. Columns (3)-(5) document point estimates from regression on relative yield curve level $L^* - L$, slope and curvature. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD from 1980:01 to 2017:12, including country fixed effects. The panel is unbalanced. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively, using Driscoll and Kraay (1998) standard errors (reported in parentheses).

Figure 5: Estimated relative slope coefficients from augmented UIP regression using more conservative inference



Notes: Black circles denote $\hat{\beta}_{2,\kappa}$ point estimates from regression (2). The horizontal axis denotes the horizon κ in months. In regression (2), the slope and curvature in each country are measured using proxies. Regressions are estimated using pooled monthly data from 1980:01 to 2017:12, including country fixed effects. 90% confidence intervals, calculated using implied standard errors from scaled t-statistics proposed by Moon et al. (2004) standard errors, are denoted by thick black bars around point estimates.

point estimates are significantly positive according to the more conservative inference from the 2.5 to 4-year horizons, within which the peak of the tent arises.

In addition, Figure 6 plots the $\beta_{1,\kappa}$ and $\beta_{3,\kappa}$ coefficient estimates from (2) alongside the 90% confidence bands implied by the scaled t-statistics. While the overall pattern of $\beta_{1,\kappa}$ coefficients is broadly the same as the canonical UIP regression, the confidence bands with these more conservative t-statistics are wider. The scaled t-statistics also imply that the coefficients on the relative curvature are statistically insignificant at all horizons.

Sub-sample stability. To assess the stability of our results, we estimate regression (2) on two sub-samples. The first, from 1980:01 to 2008:06, is intended to capture the pre-crisis period. The second, from 1990:01 to 2017:12, includes the post-crisis period.

The slope coefficient estimates from different sub-samples are presented in Table 11. For comparison, column (1) includes the relative slope coefficient loadings from our benchmark sample presented in the main body of the paper. Columns (3) and (4) include the estimated loadings over the pre-crisis and predominantly post-crisis samples, respectively. In both cases the coefficient estimates form a tent shape with respect to maturity, peaking at the 4 and 3.5-year horizons, respectively.

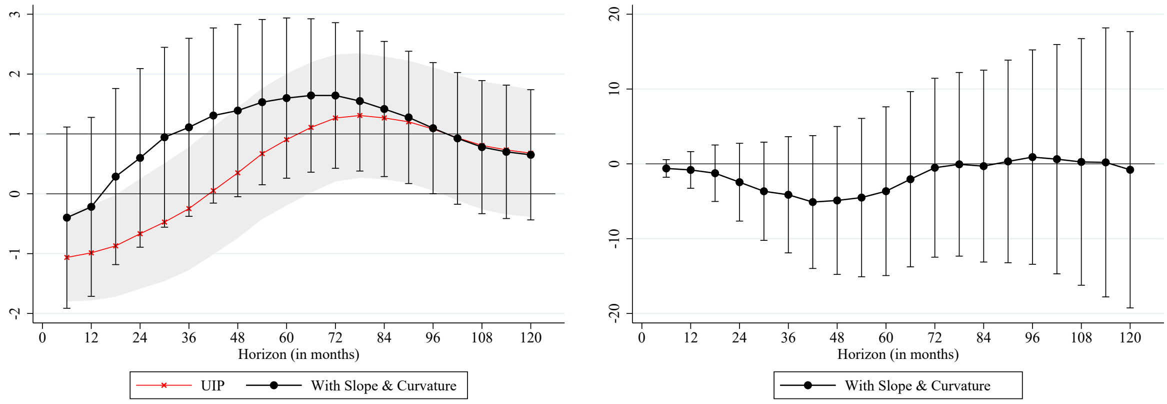
In addition, columns (2) and (5) present two additional robustness exercises. In column (2), we use available G10 currency and yield curve data, adding Sweden, Norway and New

Table 11: Slope coefficient estimates from augmented UIP regression for pooled regression across different samples

Maturity	(1) 1980:01- 2017:12 G7 Currencies	(2) 1980:01- 2008:06 G10 Currencies	(3) 1980:01- 2008:06 G7 Currencies	(4) 1990:01- 2017:12 G7 Currencies	(5) 1980:01- 2017:12 Excl. $C^* - C$ G7 Curr.
6-months	0.75	0.60	0.91	1.18	0.39
	(0.70)	(0.72)	(0.71)	(0.77)	(0.65)
12-months	1.41	1.07	1.47	1.99*	0.86
	(1.14)	(1.18)	(1.23)	(1.19)	(0.98)
18-months	2.87**	2.56*	3.11**	3.10**	2.02*
	(1.31)	(1.34)	(1.32)	(1.46)	(1.06)
24-months	4.31***	4.37***	4.97***	4.33***	2.67**
	(1.50)	(1.53)	(1.47)	(1.64)	(1.19)
30-months	5.98***	6.18***	6.75***	6.39***	3.58***
	(1.60)	(1.63)	(1.54)	(1.68)	(1.26)
36-months	6.74***	7.24***	7.90***	8.00***	4.12***
	(1.63)	(1.68)	(1.52)	(1.57)	(1.25)
42-months	7.40***	8.62***	9.35***	9.01***	4.27***
	(1.61)	(1.66)	(1.53)	(1.50)	(1.14)
48-months	7.04***	9.00***	9.84***	8.82***	4.14***
	(1.68)	(1.67)	(1.67)	(1.69)	(1.11)
54-months	6.63***	8.74***	9.62***	8.72***	4.03***
	(1.83)	(1.78)	(1.93)	(1.98)	(1.13)
60-months	5.98***	8.29***	9.19***	8.29***	3.92***
	(1.97)	(1.96)	(2.18)	(2.24)	(1.21)
66-months	4.91**	7.58***	8.28***	6.82***	3.78***
	(2.03)	(2.01)	(2.23)	(2.19)	(1.25)
72-months	3.61*	6.49***	7.02***	4.81**	3.33***
	(1.93)	(1.83)	(1.97)	(2.14)	(1.18)
78-months	2.54	5.48***	5.74***	3.10	2.50**
	(1.77)	(1.65)	(1.79)	(2.03)	(1.08)
84-months	1.89	4.12**	4.21**	2.25	1.73*
	(1.65)	(1.62)	(1.89)	(1.95)	(1.01)
90-months	0.93	2.55	2.54	1.42	1.09
	(1.60)	(1.61)	(1.91)	(1.92)	(0.97)
96-months	-0.06	1.14	1.22	0.46	0.40
	(1.68)	(1.76)	(2.07)	(2.03)	(0.96)
102-months	-0.41	0.29	0.61	0.13	-0.09
	(1.74)	(1.82)	(2.15)	(2.09)	(1.06)
108-months	-0.71	-0.42	0.05	-0.54	-0.59
	(1.83)	(1.87)	(2.20)	(2.16)	(1.16)
114-months	-0.88	-0.79	0.07	-0.60	-0.78
	(1.89)	(1.91)	(2.25)	(2.28)	(1.20)
120-months	-0.42	-0.42	0.65	-0.07	-0.83
	(1.66)	(1.66)	(2.01)	(2.02)	(1.20)

Notes: Coefficient estimates on the relative yield curve slope $S_t^* - S_t$ from regression (2)—the augmented UIP regression—a regression of the κ-period exchange rate change $\Delta^\kappa e_{t+\kappa}$ on the κ-period return differential $r_{t,\kappa}^* - r_{t,\kappa}$, the relative yield curve slope and the relative yield curve curvature $C_t^* - C_t$. Regressions in columns (1) and (3)-(5) are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD for different samples. Column (2) includes three additional currencies—NOK, NZD and SEK—for zero-coupon government bond yield curve data is available prior to the crisis. All regressions include country fixed effects. The panels are unbalanced and standard errors (reported in parentheses) are constructed according to the Driscoll and Kraay (1998) methodology. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively.

Figure 6: Estimated relative slope coefficients from augmented UIP regression using more conservative inference



Notes: Black circles denote $\hat{\beta}_{1,\kappa}$ (left-hand side) and $\hat{\beta}_{3,\kappa}$ (right-hand side) point estimates from regression (2). The horizontal axis denotes the horizon κ in months. In regression (2), the slope and curvature in each country are measured using proxies. Regressions are estimated using pooled monthly data from 1980:01 to 2017:12, including country fixed effects. 90% confidence intervals, calculated using implied standard errors from scaled t-statistics proposed by Moon et al. (2004) standard errors, are denoted by thick black bars around point estimates.

Zealand to our cross-section of countries, for the pre-crisis period only. In column (5), we drop the relative curvature from regression (2), to demonstrate that the relative slope coefficient is independent on the inclusion of the relative curvature. In both cases, the relative slope loadings continue to follow a tent-shaped pattern with respect to maturity.

Country-specific regressions. Table 12 presents country-specific estimates of the yield curve augmented-UIP regression. Inference is conducted using Newey and West (1987) standard errors, to account for serial correlation. For comparison, column (1) presents the benchmark relative slope coefficient estimates from the panel regression discussed in the main body of the paper. As noted in the main text, although coefficient estimates vary in size and significance across countries, a relative slope coefficient estimates display a tent shape with respect to horizon κ for most of the currencies in our sample (AUD, CHF, EUR, JPY, GBP). A positive tent shape is present for Canada as well, but is insignificant. The peak of the tent realises at 30-42 months for all 6 currency pairs. However, some anomalies arise at long horizons beyond 8 years.

Table 12: Slope coefficient estimates from augmented UIP regression for pooled regression and country-specific regressions

Maturity	(1) Panel	(2) Australia	(3) Canada	(4) Switzerland	(5) Euro area	(6) Japan	(7) United Kingdom
6-months	0.75	2.35	-0.06	-0.08	-0.41	3.04**	0.38
	(0.70)	(1.61)	(1.06)	(1.72)	(1.23)	(1.25)	(1.15)
12-months	1.41	3.66	0.60	1.47	-1.18	4.68**	0.87
	(1.14)	(2.52)	(1.63)	(2.87)	(2.09)	(2.18)	(1.69)
18-months	2.87**	6.45**	1.87	5.50	-1.90	4.90*	2.87
	(1.31)	(2.80)	(1.89)	(3.55)	(2.69)	(2.80)	(2.02)
24-months	4.31***	7.93**	2.17	9.41***	-1.85	5.01	5.46**
	(1.50)	(3.31)	(2.29)	(3.17)	(3.23)	(3.31)	(2.39)
30-months	5.98***	11.52***	2.22	10.14***	-1.06	7.04**	7.87***
	(1.60)	(3.56)	(2.61)	(2.30)	(3.63)	(3.57)	(2.56)
36-months	6.74***	15.93***	2.76	7.84***	0.06	7.09*	9.17***
	(1.63)	(3.36)	(2.68)	(2.70)	(4.05)	(3.86)	(2.49)
42-months	7.40***	18.19***	3.64	8.38**	0.89	6.46*	10.17***
	(1.61)	(3.29)	(2.77)	(3.35)	(4.56)	(3.69)	(2.64)
48-months	7.04***	17.55***	4.05	7.94**	1.93	4.17	9.77***
	(1.68)	(3.85)	(3.03)	(3.72)	(4.52)	(3.44)	(2.73)
54-months	6.63***	16.08***	3.83	7.17*	3.05	3.59	9.14***
	(1.83)	(4.11)	(3.36)	(4.12)	(4.22)	(3.29)	(2.41)
60-months	5.98***	15.22***	3.97	5.36	3.68	3.95	7.81***
	(1.97)	(4.04)	(3.69)	(4.52)	(3.93)	(3.04)	(2.09)
66-months	4.91**	13.17***	2.89	4.33	3.32	3.63	6.24***
	(2.03)	(3.56)	(4.01)	(4.49)	(3.49)	(2.95)	(1.98)
72-months	3.61*	10.16***	1.69	3.38	2.26	2.64	4.85***
	(1.93)	(2.89)	(4.17)	(4.09)	(3.08)	(3.05)	(1.78)
78-months	2.54	7.87***	0.73	3.05	0.98	2.31	3.37**
	(1.77)	(2.96)	(4.10)	(3.49)	(2.69)	(3.05)	(1.54)
84-months	1.89	5.80*	0.68	4.13	-0.81	3.88	2.03
	(1.65)	(3.11)	(4.31)	(2.86)	(2.52)	(2.92)	(1.47)
90-months	0.93	4.61	0.66	3.42	-3.34	5.92**	0.21
	(1.60)	(3.43)	(4.56)	(2.70)	(2.33)	(2.71)	(1.61)
96-months	-0.06	3.24	1.71	2.00	-5.90***	7.38***	-1.38
	(1.68)	(3.84)	(4.78)	(2.77)	(2.23)	(2.80)	(1.73)
102-months	-0.41	3.71	2.72	1.18	-6.51***	8.22***	-2.44
	(1.74)	(4.10)	(4.80)	(2.81)	(2.23)	(2.70)	(1.83)
108-months	-0.71	3.05	3.73	0.03	-6.87***	8.81***	-2.84
	(1.83)	(4.16)	(4.79)	(3.09)	(2.27)	(2.65)	(1.91)
114-months	-0.88	3.21	4.60	0.65	-7.46***	9.96***	-3.62**
	(1.89)	(4.54)	(4.92)	(3.39)	(2.47)	(2.28)	(1.60)
120-months	-0.42	4.45	5.48	1.75	-7.63***	8.63***	-2.29*
	(1.66)	(4.32)	(4.68)	(3.22)	(2.39)	(2.50)	(1.32)

Notes: Coefficient estimates on the relative yield curve slope $S_t^* - S_t$ from regression (2)—the augmented UIP regression—a regression of the κ-period exchange rate change $\Delta^\kappa e_{t+\kappa}$ on the κ-period return differential $r_{t,\kappa}^* - r_{t,\kappa}$, the relative yield curve slope and the relative yield curve curvature $C_t^* - C_t$. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD from 1980:01 to 2017:12. Column (1) presents coefficient estimates from a panel regression of all six countries, including country fixed effects. The panel is unbalanced and standard errors (reported in parentheses) are constructed according to the Driscoll and Kraay (1998) methodology. Columns (2)-(7) report coefficient estimates from country-specific regressions. Newey and West (1987) standard errors (reported in parentheses) are constructed with a maximum lag of 5. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively.

B.3 Robustness Results for Risk Premia

In Table 13, we present the mean return from a simple investment strategy that goes long the Foreign bond and short the US bond when the Foreign yield curve slope is lower than the US yield curve slope, and goes long the US bond and short the Foreign bond when the US yield curve slope is lower than the Foreign yield curve slope. Relative to Lustig et al. (2019), we present the mean dollar-bond return differences for a range of holding periods $h = 6, 12, ..., 60$ and maturities $\kappa = 6, 12, ..., 120$ (in months).

Table 13: Mean Excess Returns from Dynamic Long-Short Bond Portfolios

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
					Holding Periods					
	6m	12m	18m	24m	30m	36m	42m	48m	54m	60m
Dollar-Bond Return Difference: $rx^{(\kappa),\$}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$										
12m	1.95									
18m	1.81	2.48								
24m	1.70	2.38	3.04							
30m	1.60	2.3	2.98	3.3						
36m	1.49	2.21	2.92	3.26	3.30					
42m	1.38	2.12	2.85	3.22	3.27	3.08				
48m	1.26	2.01	2.76	3.16	3.24	3.06	2.9			
54m	1.15	1.91	2.67	3.10	3.20	3.03	2.88	2.57		
60m	1.03	1.81	2.58	3.03	3.15	2.99	2.85	2.55	2.30	
66m	0.93	1.72	2.49	2.95	3.09	2.95	2.82	2.52	2.28	2.35
72m	0.83	1.63	2.40	2.88	3.03	2.89	2.77	2.49	2.25	2.32
78m	0.74	1.55	2.32	2.81	2.96	2.84	2.72	2.45	2.22	2.29
84m	0.67	1.48	2.24	2.74	2.90	2.78	2.67	2.41	2.18	2.26
90m	0.58	1.41	2.17	2.67	2.84	2.72	2.62	2.36	2.14	2.23
96m	0.51	1.35	2.09	2.60	2.78	2.65	2.56	2.31	2.10	2.19
102m	0.45	1.29	2.03	2.54	2.71	2.59	2.50	2.26	2.06	2.16
108m	0.39	1.23	1.96	2.48	2.65	2.53	2.44	2.21	2.02	2.12
114m	0.34	1.18	1.90	2.42	2.59	2.47	2.39	2.16	1.98	2.09
120m	0.29	1.12	1.84	2.36	2.53	2.41	2.33	2.11	1.94	2.05

Notes: Summary return statistics from investment strategies that go long in the Foreign-country bond and short in the US bond when the Foreign yield curve slope is lower than the US yield curve slope, and go long in the US bond and short in the Foreign-country bond when the Foreign yield curve slope is higher than the US yield curve slope. The table reports the mean US dollar-bond excess return difference for different holding periods and different maturities. Returns are annualised and constructed using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD for different country samples spanning 1980:01-2017:12.

B.4 Additional Results for Liquidity Yield-Augmented Regressions

Table 14 shows coefficient estimates from regression (7) for local currency-bond excess returns.

Table 14: Slope and liquidity yield coefficient estimates from pooled regression of excess returns

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
					Holding Periods					
	6m	12m	18m	24m	30m	36m	42m	48m	54m	60m
Panel C.i: Dependent Variable: $rx^{(\kappa)}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$, Coefficient on $S^{*} - S$, when η^{R}_{κ} is additional control										
12m	-0.03									
	(0.05)									
24m	0.21*	0.03	-0.01							
	(0.12)	(0.08)	(0.03)							
36m	0.42**	0.13	0.00	-0.04	-0.02					
	(0.19)	(0.15)	(0.10)	(0.06)	(0.03)					
60m	1.03***	0.57**	0.33	0.18	0.10	0.07	0.04	0.03	0.02	
	(0.31)	(0.25)	(0.20)	(0.15)	(0.11)	(0.08)	(0.05)	(0.04)	(0.02)	
84m	1.52***	0.99***	0.70***	0.50**	0.38**	0.30**	0.23**	0.17*	0.14**	0.11**
	(0.39)	(0.32)	(0.27)	(0.22)	(0.17)	(0.13)	(0.11)	(0.09)	(0.07)	(0.05)
120m	1.97***	1.41***	1.11***	0.89***	0.77***	0.67***	0.56***	0.44***	0.35***	0.28***
	(0.50)	(0.39)	(0.33)	(0.27)	(0.22)	(0.18)	(0.16)	(0.14)	(0.12)	(0.09)
Panel C.ii: Dependent Variable: $rx^{(\kappa)}_{t,t+h} - rx^{(\kappa)}_{US,t,t+h}$, Coefficient on η^{R}_{κ}										
12m	0.00									
	(0.00)									
24m	-0.01	-0.01**	-0.00***							
	(0.01)	(0.00)	(0.00)							
36m	-0.01	-0.01**	-0.01***	-0.01***	-0.00***					
	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)					
60m	-0.02**	-0.03***	-0.02***	-0.02***	-0.02***	-0.01***	-0.01***	-0.00***	-0.00***	
	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	
84m	-0.02	-0.02**	-0.02***	-0.02***	-0.02***	-0.02***	-0.01***	-0.01***	-0.01***	-0.00**
	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
120m	-0.03*	-0.03***	-0.03***	-0.03***	-0.03***	-0.03***	-0.02***	-0.02***	-0.02***	-0.01***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
N	1,733	1,697	1,661	1,625	1,589	1,553	1,517	1,481	1,445	1,409

Notes: Coefficient estimates on the relative yield curve slope $S^{*}_{t} - S_{t}$ (Panel C.i) and cross-country κ-period liquidity yield η^{R}_{κ} (Panel C.ii) from regressions with the log local currency-bond excess return difference as dependent variable. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD for different samples within 1991:04-2017:12. The log returns and yield curve slopes differentials are annualised. All regressions include country fixed effects. All regressions include country fixed effects. The panels are unbalanced and standard errors (reported in parentheses) are constructed according to the Driscoll and Kraay (1998) methodology. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively.

C Derivations for Two-Country Cox, Ingersoll and Ross Model

Bond-pricing recursions. To derive the Home bond price recursions, we guess and verify equation (18). We use the fact that $p_{t,n} = \mathbb{E}_t[m_{t,t+1} + p_{t+1,n-1}] + (1/2)\text{var}_t(m_{t,t+1} + p_{t+1,n-1})$, and combine the guess with equations (16) and (17).

First, consider the one-period bond, $n = 1$:

$$\begin{aligned}
p_{t,1} &= \mathbb{E}_t\left[m_{t,t+1}\right] + \frac{1}{2}\text{var}_t\left(m_{t,t+1}\right) \\
&= -\alpha - \chi z_{1,t} - \tau z_{2,t} + \frac{1}{2}\gamma z_{1,t} + \frac{1}{2}\delta z_{2,t} \\
&= -\alpha - \left(\chi - \frac{1}{2}\gamma\right) z_{1,t} - \left(\tau - \frac{1}{2}\delta\right) z_{2,t}
\end{aligned}$$

where the first line uses the expression for the bond price for $n = 1$, the conditional expectation of equation (16) is used in the second line, and the resulting expression is rearranged to yield the third line. The one-period risk-free yield $y_{t,1}$ is therefore given by:

$$y_{t,1} = \alpha + (\chi - \frac{1}{2}\gamma)z_{1,t} + (\tau - \frac{1}{2}\delta)z_{2,t} \tag{21}$$

Next, consider the general n-period bond price:

$$\begin{aligned}
p_{t,n} =& \mathbb{E}_t\left[m_{t,t+1} + p_{t+1,n-1}\right] + \frac{1}{2}\text{var}_t\left(m_{t,t+1} + p_{t+1,n-1}\right) \\
=& -\alpha - \chi z_{1,t} - \tau z_{2,t} + \mathbb{E}_t\left[-A_{n-1} - B_{n-1}z_{1,t+1} - C_{n-1}z_{2,t+1}\right] \\
&+ \frac{1}{2}\text{var}_t\left(-\sqrt{\gamma z_{1,t}}u_{1,t+1} - \sqrt{\delta z_{2,t}}u_{2,t+1} - B_{n-1}z_{1,t+1} - C_{n-1}z_{2,t+1}\right) \\
=& -\alpha - A_{n-1} - B_{n-1}(1-\phi_1)\theta_1 - C_{n-1}(1-\phi_2)\theta_2 \\
&- \chi z_{1,t} - B_{n-1}\phi_1 z_{1,t} - \tau z_{2,t} - C_{n-1}\phi_2 z_{2,t} \\
&+ \frac{1}{2}\text{var}_t\left(-u_{1,t+1}\left[\sqrt{\gamma} + B_{n-1}\sigma_1\right]\sqrt{z_{1,t}} - u_{2,t+1}\left[\sqrt{\delta} + C_{n-1}\sigma_2\right]\sqrt{z_{2,t}}\right) \\
=& -\alpha - A_{n-1} - B_{n-1}(1-\phi_1)\theta_1 - C_{n-1}(1-\phi_2)\theta_2 \\
&- \chi z_{1,t} - B_{n-1}\phi_1 z_{1,t} - \tau z_{2,t} - C_{n-1}\phi_2 z_{2,t} \\
&+ \frac{1}{2}\left[(B_{n-1}\sigma_1)^2 z_{1,t} + 2B_{n-1}\sigma_1\sqrt{\gamma}z_{2,t} + \gamma z_{1,t} + (C_{n-1}\sigma)^2 z_{2,t} + 2C_{n-1}\sigma_2\sqrt{\delta}z_{2,t} + \delta z_{2,t}\right] \\
=& -\left(\alpha + A_{n-1} + B_{n-1}(1-\phi_1)\theta_1 + C_{n-1}(1-\phi_2)\theta_2\right) \\
&- z_{1,t}\left[\left(\chi - \frac{1}{2}\gamma\right) + B_{n-1}\left(\phi_1 + \sigma_1\sqrt{\gamma}\right) - \frac{1}{2}\left(B_{n-1}\sigma_1\right)^2\right] \\
&- z_{2,t}\left[\left(\tau - \frac{1}{2}\delta\right) + C_{n-1}\left(\phi_2 + \sigma_2\sqrt{\delta}\right) - \frac{1}{2}\left(C_{n-1}\sigma_2\right)^2\right]
\end{aligned}$$

where the second line uses equation (16) and the guess; the third line uses equation (17); the fourth line rearranges the third; and the fifth line collects like terms. The recursions can be seen in the final line:

$$
\begin{aligned}
A_n &\equiv A_n(\alpha, \phi_1, \phi_2, \theta_1, \theta_2; A_{n-1}, B_{n-1}, C_{n-1}) = \alpha + A_{n-1} + B_{n-1}(1-\phi_1)\theta_1 + C_{n-1}(1-\phi_2)\theta_2 \\
B_n &\equiv B_n(\phi_1, \chi, \gamma, \sigma_1; B_{n-1}) = \left(\chi - \frac{1}{2}\gamma\right) + B_{n-1}\left(\phi_1 + \sigma_1\sqrt{\gamma}\right) - \frac{1}{2}\left(B_{n-1}\sigma_1\right)^2 \\
C_n &\equiv C_n(\phi_2, \tau, \delta, \sigma_2; C_{n-1}) = \left(\tau - \frac{1}{2}\delta\right) + C_{n-1}\left(\phi_2 + \sigma_2\sqrt{\delta}\right) - \frac{1}{2}\left(C_{n-1}\sigma_2\right)^2
\end{aligned}
$$

with initial conditions $A_0 = B_0 = C_0 = 0$. So the n-period bond price is:

$$
p_{t,n} = -\left(A_n + B_n z_{1,t} + C_n z_{2,t}\right)
$$

verifying equation (18).

Bond excess returns. The *ex ante* n-period bond excess return is defined as $\mathbb{E}_t[rx^{(n)}_{t,t+1}] = \mathbb{E}_t[p_{t+1,n-1} - p_{t,1} - y_{t,1}]$. This can be written as:

$$
\begin{aligned}
\mathbb{E}_t\left[rx^{(n)}_{t,t+1}\right] =& \mathbb{E}_t\left[p_{t+1,n-1} - p_{t,1} - y_{t,1}\right] \\
=& \mathbb{E}_t\Big[-A_{n-1} + A_n - B_{n-1}z_{1,t+1} + B_n z_{1,t} - C_{n-1}z_{2,t+1} + C_n z_{2,t} \\
& -\alpha - (\chi - \frac{1}{2}\gamma)z_{1,t} - (\tau - \frac{1}{2}\delta)z_{2,t}\Big] \\
=& B_{n-1}(1-\phi_1)\theta_1 + C_{n-1}(1-\phi_2)\theta_2 - B_{n-1}\mathbb{E}_t[z_{1,t+1}] + B_n z_{1,t} \\
& - C_{n-1}\mathbb{E}_t[z_{2,t+1}] + C_n z_{2,t} - (\chi - \frac{1}{2}\gamma)z_{1,t} - (\tau - \frac{1}{2}\delta)z_{2,t} \\
=& \left[-B_{n-1}\phi_1 + B_n - (\chi - \frac{1}{2}\gamma)\right] z_{1,t} + \left[-C_{n-1}\phi_2 + C_n - (\tau - \frac{1}{2}\delta)\right] z_{2,t} \\
=& \left[B_{n-1}\sigma_1\sqrt{\gamma} - \frac{1}{2}(B_{n-1}\sigma_1)^2\right] z_{1,t} + \left[C_{n-1}\sigma_2\sqrt{\delta} - \frac{1}{2}(C_{n-1}\sigma_2)^2\right] z_{2,t} \qquad (22)
\end{aligned}
$$

where line 2 uses equations (18) and (21), line 3 uses the recursion for A_n defined above, line 4 expands the conditional expectation of factors and collects like terms, line 5 uses the recursions for B_n and C_n defined above.

Evaluating the expression above in the limit as $n \to \infty$ yields:

$$
\mathbb{E}_t\left[rx^{(\infty)}_{t,t+1}\right] = \left[B_\infty\sigma_1\sqrt{\gamma} - \frac{1}{2}(B_\infty\sigma_1)^2\right] z_{1,t} + \left[C_\infty\sigma_2\sqrt{\delta} - \frac{1}{2}(C_\infty\sigma_2)^2\right] z_{2,t}
$$

Using the recursions for B_n and C_n, this can be rearranged as:

$$
\mathbb{E}_t\left[rx^{(\infty)}_{t,t+1}\right] = \left[B_\infty(1-\phi_1) - \chi + \frac{1}{2}\gamma\right] z_{1,t} + \left[C_\infty(1-\phi_2) - \tau + \frac{1}{2}\delta\right] z_{2,t}
$$

Moreover, equation (22) can be used to express the *ex ante* bond risk premium:

$$
-\mathrm{cov}_t(p_{t+1,n-1}, m_{t,t+1}) = \mathbb{E}_t\left[rx^{(\infty)}_{t,t+1}\right] + \frac{1}{2}\mathrm{var}_t(r_{n,t+1})
$$

$$=B_{n-1}\sigma_1\sqrt{\gamma}z_{1,t}+C_{n-1}\sigma_2\sqrt{\delta}z_{2,t}$$

recovering equation (20) in the main body.

Yield curve slope. The yield curve slope is defined as the difference between yields on n- and 1-period bonds:

$$S_{t,n}=y_{t,n}-y_{t,1}=\frac{1}{n}\left(A_n+B_nz_{1,t}+C_nz_{2,t}\right)-\alpha-\left(\chi-\frac{1}{2}\gamma\right)z_{1,t}-\left(\tau-\frac{1}{2}\delta\right)z_{2,t}$$

Evaluating this expression in the limit as $n\to\infty$ yields:

$$S_{t,\infty}=B_\infty(1-\phi_1)\theta_1+C_\infty(1-\phi_2)\theta_2-\left(\chi-\frac{1}{2}\gamma\right)z_{1,t}-\left(\tau-\frac{1}{2}\delta\right)z_{2,t}$$

which arises from the recursions for A_n, B_n and C_n, where B_n and C_n have a finite limit and A_n grows linearly.

The approximation of the slope by the bond risk premium $S_{t,\infty}\approx\mathbb{E}_t[rx_{t,t+1}^{(\infty)}]$ is also verified within the CIR model. Over long enough samples, $\mathbb{E}_t[z_{1,t}]=\theta_1$ and $\mathbb{E}_t[z_{2,t}]=\theta_2$, yielding the result.

To calibrate the CIR model at a monthly frequency, we define the slope as the difference between 10-year and 6-month yields:

$$S_t=y_{t,120}-y_{t,6}=\frac{1}{120}\left(A_{120}+B_{120}z_{1,t}+C_{120}z_{2,t}\right)-\frac{1}{6}\left(A_6+B_6z_{1,t}+C_6z_{2,t}\right)\tag{23}$$

Exchange rates. Under complete markets, (log) one-period exchange rate changes are determined as:

$$\mathbb{E}_t[e_{t+1}]-e_t=\mathbb{E}_t[m_{t,t+1}-m_{t,t+1}^*]=\chi(z_{1,t}-z_{1,t}^*)+\tau(z_{2,t}-z_{2,t}^*)$$

The k-step ahead exchange rate change is then given by:

$$\mathbb{E}_t[e_{t+\kappa}]-e_t=\sum_{i=1}^{\kappa}\mathbb{E}_t[\Delta^1e_{t+i}]=\frac{1-\phi_1^k}{1-\phi_1}\chi(z_{1,t}-z_{1,t}^*)+\frac{1-\phi_2^k}{1-\phi_2}\tau(z_{2,t}-z_{2,t}^*)\tag{24}$$

The one-period ERRP can be derived by combining equations (12) and (16):

$$\mathbb{E}_t[rx_{t,t+1}^{FX}]=\frac{1}{2}\gamma(z_{1,t}-z_{1,t}^*)+\frac{1}{2}\delta(z_{2,t}-z_{2,t}^*)$$

Model-implied UIP coefficient. The UIP coefficient is constructed as the scaled conditional covariance of the sum of expected future exchange rate movements and cross-country return differentials across maturities h:

$$\beta_h^{UIP}=\frac{\mathrm{cov}_t(\mathbb{E}_t[e_{t+h}]-e_t,r_{t,h}^*-r_{t,h})}{\mathrm{var}_t(r_{t,h}^*-r_{t,h})}$$

$$= \frac{\chi \frac{1-\phi_1^h}{1-\phi_1} B_h \text{var}_t(z_{1,t}^* - z_{1,t}) + \tau \frac{1-\phi_2^h}{1-\phi_2} C_h \text{var}_t(z_{2,t}^* - z_{2,t})}{B_h^2 \text{var}_t(z_{1,t}^* - z_{1,t}) + C_h^2 \text{var}_t(z_{2,t}^* - z_{2,t})} \tag{25}$$

where line 1 is the definition of the univariate regression coefficient, and line 2 uses equation (24) and the definition for returns.

C.1 Proof to Lemma

Consider the UIP coefficient in equation (25) for a general horizon h.

Condition (i). For $h = 1$: $\frac{1-\phi_1^h}{1-\phi_1} B_h = B_1 = \chi - \frac{1}{2}\gamma$ and $\frac{1-\phi_2^h}{1-\phi_2} C_h = C_1 = \tau - \frac{1}{2}\delta$. Condition (i) follows from requiring the numerator to be negative, since the denominator is strictly positive.

Condition (ii). To rule out permanent innovations to investors' SDFs, we use the expression for the expected (log) bond excess return, equation (22). Lustig et al. (2019) (Appendix IV.C) show that if there are no permanent innovations, then $\lim_{\kappa \to \infty} \mathbb{E}_t[rx_{t,t+1}^{(\kappa)}]$ must equal half the variance of the SDF $\frac{1}{2}\gamma z_{1,t} + \frac{1}{2}\delta z_{2,t}$. Condition (ii) follows from this. If this condition is satisfied, $m_{t,t+1}^{\mathbb{P}}$ is constant.

Condition (iii). For UIP to hold approximately in levels, the UIP coefficient must switch sign over horizons, such that the infinite sum of ERRP is small. To achieve this, in tandem with condition (i), the second term in the numerator of equation (25) must be positive $(\tau - \frac{1}{2}\delta) > 0$ and become sufficiently large, relative to the negative first term, as $h \to \infty$. Note that $\frac{1-\phi_i^h}{1-\phi_i}$ is an increasing function of ϕ_i when $\phi_i < 1$ $(i = 1, 2)$, and $|B_n|$ and $|C_n|$ are also increasing functions of ϕ_1 and ϕ_2, respectively. Since $B_n < 0$ and $C_n > 0$, it follows that the numerator—which is negative for $h = 1$—becomes positive for large h if ϕ_2 is sufficiently larger than ϕ_1. $\square$

C.2 Proof to Proposition

The model-implied coefficient from a univariate regression of the cumulative h-period exchange rate movement and the cross-country slope differential (defined as the difference between 120- and 6-month yields) is given by:

$$\begin{aligned} \beta_h^{SR} &= \frac{\text{cov}_t(\mathbb{E}_t[e_{t+h} - e_t], (y_{t,120}^* - y_{t,6}^*) - (y_{t,120} - y_{t,6}))}{\text{var}_t((y_{t,120}^* - y_{t,6}^*) - (y_{t,120} - y_{t,6}))} \\ &= \frac{\frac{1-\phi_1^h}{1-\phi_1}\chi \left[\frac{1}{120}B_{120} - \frac{1}{6}B_6\right] \text{var}_t(z_{1,t}^* - z_{1,t}) + \frac{1-\phi_2^h}{1-\phi_2}\tau \left[\frac{1}{120}C_{120} - \frac{1}{6}C_6\right] \text{var}_t(z_{2,t}^* - z_{2,t})}{\left[\frac{1}{120}B_{120} - \frac{1}{6}B_6)\right]^2 \text{var}_t(z_{1,t}^* - z_{1,t}) + \left[\frac{1}{120}C_{120} - \frac{1}{6}C_6\right]^2 \text{var}_t(z_{2,t}^* - z_{2,t})} \end{aligned} \tag{26}$$

where line 1 is the definition of the univariate regression coefficient, and line 2 uses equations (23) and (24).

To generate a tent-shaped pattern for coefficients across horizons, we require $\beta_h^{SR} < \beta_{h'}^{SR}$ for some $h' > h$ and $\beta_{h'}^{SR} > \beta_{h''}^{SR}$ for $h'' > h'$. To achieve this, the first term in the numerator, which is associated with the less persistent factor, must be positive such that the term is rising for some $h < h'$. The second term, associated with the more persistent factor, must be negative.

Conditions (i) and (iii) from the Lemma ensure that: $\text{sign}(\chi - \frac{1}{2}\gamma) = -\text{sign}(\tau - \frac{1}{2}\delta)$. This implies: $\text{sign}(\frac{1}{120}B_{120} - \frac{1}{6}B_6) = -\text{sign}(\frac{1}{120}C_{120} - \frac{1}{6}C_6)$. To see this, consider the limit as $\sigma_i \to 0$ $(i = 1, 2)$ and the recursions for B_n and C_n. In this limit, the recursion for B_n collapses to: $B_n = (\chi - \frac{1}{2}\gamma) + \phi B_{n-1} = (\chi - \frac{1}{2}\gamma) + \phi_1(\chi - \frac{1}{2}\gamma) + ... + \phi_1^n B_0 = \frac{1-\phi_1^n}{1-\phi_1}(\chi - \frac{1}{2}\gamma) < 0$. Likewise C_n will collapse to $\frac{1-\phi_2^n}{1-\phi_2}(\tau - \frac{1}{2}\delta) > 0$. These conditions are also satisfied for $\sigma > 0$, as long as σ is not too large.

Furthermore, $\frac{1}{120}B_{120} - \frac{1}{6}B_6 = (\chi - \frac{1}{2}\gamma)\frac{1}{1-\phi_1}[\frac{1-\phi_1^{120}}{120} - \frac{1-\phi_1^6}{6}]$. The term in square brackets is negative for all $\phi_1 < 1$. Therefore the first term in the numerator of β_h^{SR} is positive. By analogy, the second term, involving C_n loadings, is positive.

Since $\chi, \tau > 0$ and $\chi(\chi - \frac{1}{2}\gamma)\text{var}_t(z_{1,t}^* - z_{1,t}) < -\tau(\tau - \frac{1}{2}\delta)\text{var}_t(z_{2,t}^* - z_{2,t})$ from condition (i), the first factor is dominant at short horizons. This implies that β_h^{SR} is increasing for some $h < h'$. However, as $\phi_1 < \phi_2$ from condition (iii), β_κ^{SR} will be decreasing for $h'' > h'$. If (i) or (iii) are not satisfied, the tent cannot be reproduced.

Furthermore, if risk is only permanent, which would violate condition (ii), the slope is constant and has no predictive power for exchange rates. $\square$

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