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Reference Details

1996	Cambridge Working Papers in Economics
2019/22	Cambridge-INET Working Paper Series
Published	2 December 2019
Revised	18 October 2022
Key Words	Business-cycle risk, Exchange rates, Risk premia, Stochastic discount factor, Uncovered interest parity, Yield curves
JEL Codes	E43, F31, G12
Websites	www.econ.cam.ac.uk/cwpe www.inet.econ.cam.ac.uk/working-papers

Exchange-Rate Risk and Business Cycles*

Simon P. Lloyd[†] Emile A. Marin[‡]

October 17, 2021

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We show that currencies with a steeper yield curve depreciate at business-cycle horizons. We identify a tent-shaped relationship between exchange-rate risk premia (ERRP) and the relative yield curve slope across horizons that peaks at 3-5 years and is robust to a number of controls, including liquidity yields. Within a no-arbitrage framework, ERRP reflect investors' changing return valuations over the business cycle. We calibrate a two-country, two-factor model of interest rates, where exchange rates are driven by business-cycle—transitory and cyclical—risk. The model quantitatively reproduces the tent-shaped relationship, as well as variation in uncovered interest parity coefficients across horizons.

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*We are especially grateful to Giancarlo Corsetti for many helpful discussions. We also thank anonymous referees, Gianluca Benigno, Ambrogio Cesa-Bianchi, Mikhail Chernov, Luca Dedola, Pierre-Olivier Gourinchas, Matteo Maggiori, Thomas Nitschka (discussant), Ilaria Piatti (discussant), Katrin Rabitsch, Lucio Sarno and Adrien Verdelhan as well as presentation attendees at the Bank of Canada, Bank of England, Centre for Central Banking Studies, BdF-BoE International Macroeconomics Workshop, CRETE 2019, European Economic Association Annual Conference 2019, Federal Reserve Bank of New York, Money, Macro and Finance Annual Conference 2019, the Royal Economic Society Annual Conference 2019, University of Nottingham, University of Cambridge, the 10th Banca d'Italia-European Central Bank Workshop on Exchange Rates and the 37th International Conference of the French Finance Association for useful comments. The paper was previously presented with the title "Uncovered Interest Parity and the Yield Curve: The Long and the Short of It". Marin acknowledges the Keynes Fund at the University of Cambridge (Project JHUK) for financial support on this project. The views expressed in this paper are those of the authors, and not necessarily those of the Bank of England.

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1 Introduction

A long-standing literature in international macroeconomics questions that the exchange rate can be explained by macroeconomic fundamentals, highlighting an 'exchange rate disconnect' puzzle (Meese and Rogoff, 1983; Itskhoki and Mukhin, 2021). While this puzzle challenges open-economy macroeconomics (see Lilley, Maggiori, Neiman, and Schreger, 2019, for a recent attempt to reconcile theory and evidence), leading contributions to the asset-pricing literature have taken the disconnect and other exchange rate puzzles at face value, imposing them as empirically-based restrictions on risk-pricing models (Backus, Foresi, and Telmer, 2001; Lustig, Stathopoulos, and Verdelhan, 2019).

In this paper, we draw on both literatures and reconsider, both empirically and theoretically, the link between exchange rates and the term structure of interest rates. 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), this is evidence that exchange rate movements are driven by the business cycle risk reflected in the slope differentials across countries. We show that, for this to be true in a standard no-arbitrage theoretical framework, the risks driving stochastic discount factors (SDFs) are, at least in part, transitory and cyclical. We specify a two-country, two-factor model of the term structure of interest rates, and derive restrictions that isolate these sources of risk and are implied by by well-documented exchange rate puzzles. We show that, subject to the discipline of our empirically-motivated restrictions, the model can quantitatively reproduce the relationship between the relative slope and exchange rates that we unveil in the data.

Our main results are as follows. First, we 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 is not 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.

Second, we show that the predictability of exchange rates by the relative yield curve slope predominantly works through ERRP. To do so, we 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). This allows us to isolate the specific contribution of the relative yield curve slope to bond risk premia and ERRP in turn. Generalising our first result, we identify a tent-shaped relationship across holding periods and for a range of bond maturities.

Third, we show that the relationship between the relative yield curve slope and ERRP is robust to extending 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 does explain exchange rate fluctuations, extending the results of Engel and Wu (2018) across maturities. Nevertheless, the fact that our main result—namely, the relationship between the relative yield curve slope and ERRP—is unaffected by this extension suggests that business-cycle risk operates through a distinct channel.

Fourth, we show that a standard no-arbitrage framework can predict the tent-shaped relationship between the relative yield curve slope and ERRP once, realistically, the risks driving SDFs are specified to be, at least 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 expect shocks generating economic booms to be followed by busts—a dynamic that implies upward-sloping yield curves on average (Piazzesi and Schneider, 2007). Since, in line with this definition, transitory and cyclical SDF dynamics are indicative of 'business-cycle risk', cross-country differences in the slope will reflect asymmetries in business-cycle risk, and these asymmetries drive predictable ERRP movements.

Finally, we specify a two-country version of the Cox, Ingersoll, and Ross (1985) (CIR) model for interest rates. We calibrate the model by deriving parametric conditions implied by three widely-documented exchange rate puzzles: (i) the failure to reject UIP at long horizons; (ii) the failure of UIP at short horizons; and (iii) the tendency of high-yield currencies to be contemporaneously appreciated, such that UIP holds in 'levels' (Engel, 2016). We show that the restrictions implied by these empirical regularities ensure that risk driving SDFs in the model is transitory and cyclical. Based on this calibration, 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.

Through the lens of the model, the relationship between the relative yield curve slope and exchange rates across horizons relies on investors' changing valuations of returns over the business cycle. A relatively steep yield curve indicates comparatively better times ahead. Therefore, residents in a country with a steep yield curve expect their valuation of returns to be higher in the near term than further into the future—i.e. they expect their valuations to decrease going forward. We show that, in general, the country with the steeper yield curve will simultaneously have a relatively low short-term interest rate and experience a currency depreciation, consistent with UIP failures at short horizons. However, as the economy moves along the cycle, and relative cross-country return valuations by investors are expected to fall, investors expect the interest rate to begin rise and the currency to appreciate—again consistent with short-horizon UIP failures. In other words, as the economy moves from a bust into a boom, the relative path of expected future short-term interest-rate differentials reverses, reflecting the anticipated SDF dynamics. As a result, the currency of a country with a relatively steep yield curve is expected to first depreciate—the depreciation peaking at business-cycle horizons—and then appreciate, consistent with the tent-shaped relationship we unveil in the data.

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). Our analysis is focused on a cross-time component of UIP failures, which Hassan and Mano (2019) show is an important component of exchange rate predictability. 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

To motivate our analysis, we first estimate canonical UIP regressions across horizons augmented with relative yield-curve factors. Section 3 goes a step further, analysing bond and exchange-rate risk premia. Both empirical approaches highlight our headline result: exchange rate dynamics are predictable by 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 continuously compounded 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.

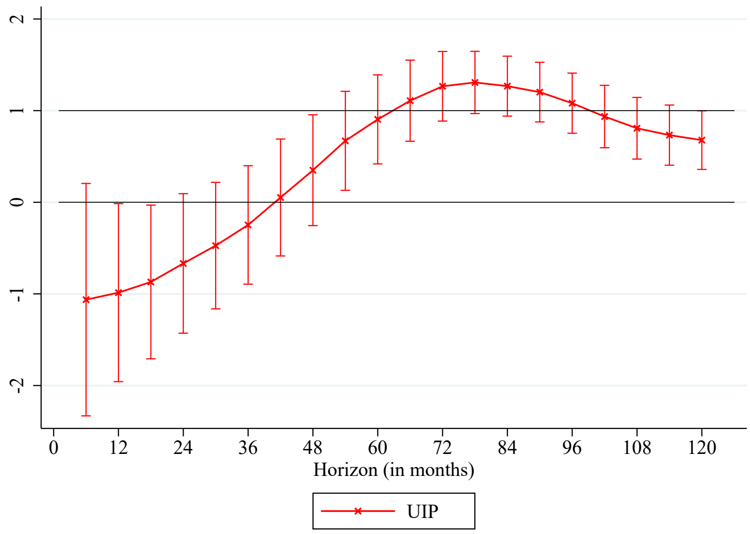
Under the joint assumption of risk neutrality and rational expectations, the null hypothesis of UIP is $\beta_{1,\kappa} = 1$ for all $\kappa > 0$.[1] Empirical rejections of UIP at short to medium horizons—i.e. finding $\hat{\beta}_{1,\kappa} \neq 1$ for small to medium κ—have regularly been used to 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

[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.

government bond yields of 6, 12, 18, ..., 120-month maturities. Yield curves are obtained from 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 10. 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).

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

2.2 Yield Curve-Augmented UIP Regression

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

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

for all κ, where $S^*_{j,t}$ is the slope of the Foreign country j yield curve at time t, and S_t is the slope of the base-country yield curve. In robustness analysis in Section 2.3, we further extend equation (2) to account for the relative yield curve curvature $C^*_{j,t} - C_t$.

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 slope'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:

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

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 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, $S^*_{j,t} = y^*_{j,t,10y} - y^*_{j,t,6m}$. In robustness analysis, we proxy the curvature using 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 proxies 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 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 $\hat{\beta}_{2,\kappa}$ from equation (2). 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 4.27% exchange rate depreciation over that horizon. Compared to the 6-month horizon, this point estimate is significantly different too. At longer horizons—from 7-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 $\beta_{1,\kappa}$ estimates, are documented in Table 9 of Appendix B.1. Augmentation of the UIP regression with the relative yield curve slope does not significantly alter UIP coefficient estimates, as confidence bands from regressions (1) and (2) overlap. 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. However, the $\beta_{1,\kappa}$ estimates from regression (2) are larger, and standard errors suggest it is harder to reject the hypothesis that $\beta_{1,\kappa} = 1$ in the augmented regression.

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.

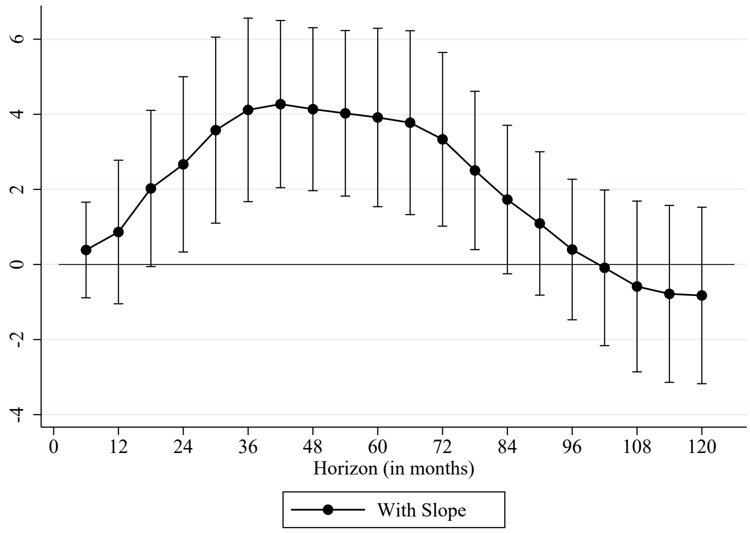
Relative curvature. Adding a proxy for the relative yield curve curvature to regression (2) does not significantly alter conclusions around the relative yield curve slope. There remains a tent-shaped relationship on relative slope coefficients across horizons. The relative curvature coefficient has a negative tent-shaped relationship across horizons. But this finding is not robustly significant,[4] and the marginal explanatory power from the relative curvature is comparatively small, justifying our focus on the relative yield curve slope.

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, and correlation with, interest rates, we also estimate a simple regression of exchange rate changes on the relative slope, 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

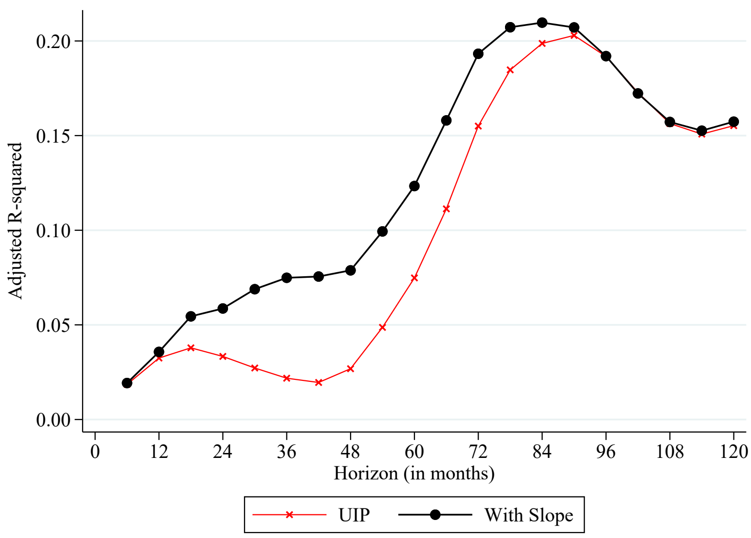
[4] Using more conservative standard errors, described in the subsequent 'Long-horizon inference' paragraph, we do not find a significant relationship between exchange rate changes and the relative curvature across horizons.

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. 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 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 relative slope-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.

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.[5] 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.

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.

[5] 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.

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:

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

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:

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

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 coincide.[6] 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.[7]

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

[7] 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.

Exploring our results at all holding periods h and for all maturities κ, three observations are noteworthy. First, the relative slope drives dollar-bond risk premia predominantly through its influence on currency risk premia. 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 (Panel B) and the relative dollar-bond risk premium (Panel A). At short horizons, the relative slope also influences local currency-bond premia, but the coefficients are quantitatively small. So, for dollar-bond premia overall, the relationship between the relative slope and the currency risk premium dominates, supporting the findings from our benchmark augmented UIP regression in Section 2.

Second, regardless of maturity, the relationship between the relative slope and relative dollar-bond returns is strongest at around the 3-year horizon (Panel A)—close to the peak at 42-month horizon from the augmented UIP regression (2). In other words, the tent peaks at business-cycle horizons. The same is true for currency risk premia (Panel B).

Third, for a given holding period, the relationship between the relative yield curve slope and dollar-bond premia declines across maturities. In other words, there is a decreasing term structure for carry trade returns. This is evidenced by the declining magnitude of coefficients down the columns in Panel A. Lustig et al. (2019) find this for the short holding period (i.e. $h = 6$). We show that this decreasing term structure extends up to the 5-year holding period.

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.

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.

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).[8] 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):

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

[8] 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.

where the dependent variable $y_{j,t,h}$ 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.

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 hold-

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.

ing 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 investigate our empirical results within a theoretical setting. We take the following approach. Building on the asset-pricing literature, we discipline a standard no-arbitrage model– using restrictions implied by well-documented exchange rate puzzles. Imposing these restrictions, we show that the tent-shaped relationship between exchange rates and the relative yield curve slope is driven by cross-country differences in business-cycle risk. We present a two-country Cox, Ingersoll, and Ross (1985) (CIR) model and derive three key results. First, we show that the relationship between the relative slope and exchange rates is driven by transitory risk. Second, we show that a one-factor model with no permanent innovations cannot simultaneously account for short-run UIP failures and positive bond premia. A two-factor model resolves this and reproduces the tent-shaped relationship between the relative slope and exchange rates if interest rates are 'conditionally cyclical'. Third, we calibrate the model to match moments of bond and equity markets, and show that it quantitatively aligns 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.

4.1.1 Notation

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 $\kappa > 0$, which follows from equations (8) and (10). The general no-arbitrage framework we consider is defined by equations (8), its foreign counterpart, (10) and (11) for all maturities.

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 $\kappa = 1$.

4.1.2 Decomposing the Pricing Kernel

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

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

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: $V^{\mathbb{P}}_t = \lim_{\kappa\to\infty} \frac{\mathbb{E}_t[V_{t+\kappa}]}{\delta^{t+\kappa}}$.

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: $R_{t,\infty} = \lim_{\kappa\to\infty} R_{t,\kappa} = V^{\mathbb{T}}_t/V^{\mathbb{T}}_{t+1} = \exp(-m^{\mathbb{T}}_{t,t+1})$, where $m^{\mathbb{T}}_{t,t+1}$ denotes the transitory component of the SDF. In contrast, one-period bond returns, defined by equation (8), depend on both transitory and permanent innovations to SDFs.

Using equations (12) and (13), Lustig et al. (2019) show that the long-horizon ERRP is proportional to cross-country differences in the variance of permanent innovations to investors' SDFs:

$$\lim_{\kappa\to\infty} \mathbb{E}_t\left[rx^{FX}_{t,t+\kappa}\right] = \frac{1}{2}\left[\text{var}_t\left(\nu^{\mathbb{P}}_{t+1}\right) - \text{var}_t\left(\nu^{\mathbb{P}*}_{t+1}\right)\right] \tag{14}$$

where $\nu^{(*)}_t \equiv \log(V^{\mathbb{P}(*)}_t)$. The relative success of long-horizon UIP requires equation (14) is approximately zero, therefore implying that cross-country differences in permanent SDF volatilities are small.[9]

Instead, our focus is on short to medium horizons, where UIP can be rejected and ERRP are therefore non-zero. At short horizons, ERRP must also reflect transitory SDF innovations.

[9] 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.

At $\kappa = 1$ ERRP can be written as:

$$\mathbb{E}_t\left[rx_{t,t+1}^{FX}\right] \approx \frac{1}{2}\left[\mathrm{var}_t\left(\nu_{t+1}^{\mathbb{T}}\right) - \mathrm{var}_t\left(\nu_{t+1}^{\mathbb{T}*}\right)\right] \neq 0 \tag{15}$$

where we have used equation (14) and the fact that $\mathrm{var}_t(\nu_{t+1}) = \mathrm{var}_t(\nu_{t+1}^{\mathbb{T}})$ when $\mathrm{var}_t(\nu_{t+1}^{\mathbb{P}}) = 0$.For medium horizons, $\kappa > 1$, ERRP are non-zero, but cannot be expressed simply as a function of variances of SDF innovations as in equations (14) and (15).

4.1.3 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.[10] This excess return can be written as:

$$\mathbb{E}_t\left[rx_{t,t+1}^{(n)}\right] = -\mathrm{cov}_t\left(m_{t,t+1}, \mathbb{E}_{t+1}\sum_{i=1}^{n-1} m_{t+i,t+i+1}\right) - \frac{1}{2}\mathrm{var}_t\left(r_{t+1,n}\right) \tag{16}$$

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.

To appreciate the role of business-cycle risk as a driver of the bond risk premium, consider the following. 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 (16) 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 longer horizons—i.e. lower $\mathbb{E}_t[m_{t+i,t+i+1}]$ for some $i > 0$—relative to their valuation in the near term $m_{t,t+1}$. 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 (16) is positive.[11] Taken together, these results suggest that the yield curve slope captures business-cycle risks, with parallels to the literature on recession predictability by the yield curve (Estrella

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

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

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

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.

We have shown that the relationship between the relative slope and ERRP is driven by transitory and cyclical factors, which we define as 'business-cycle risk'. The discussion so far has been presented in a preference-free setting. To derive this link in closed form, we next turn to a parametric model of the term structure of interest rates and exchange rates.

4.2 Two-Country Cox, Ingersoll and Ross Model

We use a calibrated two-country CIR model, imposing parametric restrictions that allow the model to capture business-cycle risk. We show that these restrictions can be mapped to three widely-documented empirical regularities, namely: (i) 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); (ii) the appreciation of high-yield currencies in excess of UIP at short horizons (Fama, 1984); and (iii) the tendency for high interest rate currencies to be contemporaneously appreciated, i.e. UIP holding in 'levels' (Engel, 2016). Applying these restrictions, we show that the model can reproduce the tent-shaped relationship across horizons between the relative yield curve slope and ERRP.

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

$$-m_{t,t+1} = \alpha + \chi z_{1,t} + \sqrt{\gamma z_{1,t}} u_{t+1} + \tau z_{2,t} + \sqrt{\delta z_{2,t}} u_{2,t+1} \tag{17}$$

$$z_{i,t+1} = (1 - \phi_i)\theta_i + \phi_i z_{i,t} - \sigma_i \sqrt{z_{i,t}} u_{i,t+1} \quad \text{for } i = 1, 2 \tag{18}$$

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 $i = 1, 2$).[12]

Assuming log-normality, then equations (17) and (18) 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:

$$p_{t,n} = -\left(A_n + B_n z_{1,t} + C_n z_{2,t}\right) \tag{19}$$

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.

Since we emphasise the importance of transitory and cyclical risk in explaining the relation-

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 $\mathbb{E}_t[rx_{t,t+1}^{(n)}] \approx y_{t,n} - y_{t,1} \equiv S_t$.

[12] We discuss the consequences of using country-specific factors, with no cross-country correlation, at the end of this sub-section when evaluating our findings.

ship between exchange rate movements and the term structure of interest rates, we use equation (13) to decompose the SDF in equation (17) into its transitory $m^{\mathbb{T}}_{t,t+1}$ and permanent $m^{\mathbb{P}}_{t,t+1}$ components in logarithmic form.

$$\begin{aligned} m^{\mathbb{T}}_{t,t+1} =& \ln\beta + B_\infty \left[(\phi_1-1)(z_{1,t}-\theta_1) - \sigma_1\sqrt{z_{1,t}}u_{1,t+1}\right] \\ &+ C_\infty \left[(\phi_2-1)(z_{2,t}-\theta_2) - \sigma_2\sqrt{z_{2,t}}u_{2,t+1}\right] \end{aligned} \tag{20}$$

$$\begin{aligned} m^{\mathbb{P}}_{t,t+1} =& -\ln\beta - \alpha - \chi z_{1,t} - \sqrt{\gamma z_{1,t}}u_{1,t+1} - \tau z_{2,t} - \sqrt{\delta z_{2,t}}u_{2,t+1} \\ &- B_\infty \left[(\phi_1-1)(z_{1,t}-\theta_1) - \sigma_1\sqrt{z_{1,t}}u_{1,t+1}\right] \\ &- C_\infty \left[(\phi_2-1)(z_{2,t}-\theta_2) - \sigma_2\sqrt{z_{2,t}}u_{2,t+1}\right] \end{aligned} \tag{21}$$

4.2.1 Deriving Model Restrictions

We discipline our model using three well-documented exchange rate puzzles. In this sub-section, we first briefly describe these puzzles and explain their mapping to the CIR framework—leaving the derivations that underpin this mapping to Appendix C. We then summarise the restrictions in a Lemma and show that they the generate transitory and cyclical risk within the model.

Long-horizon UIP and permanent innovations. As discussed above, the fact UIP cannot be rejected at long horizons implies that equation (14) is approximately zero. This rules out permanent innovations to investors' SDFs as drivers of ERRP. To eliminate these permanent drivers from the model, we set the infinite-maturity bond excess return

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

to equal half the variance of the SDF: $\mathbb{E}_t[rx^{(\infty)}_{t,t+1}] = \frac{1}{2}\gamma z_{1,t} + \frac{1}{2}\delta z_{2,t}$. Assuming this restriction holds factor-by-factor, permanent innovations are eliminated if: $B_\infty = \sqrt{\gamma}/\sigma_1$ and $C_\infty = \sqrt{\delta}/\sigma_2$. So, using equation (21): $m^{\mathbb{P}}_{t,t+1} = -\alpha - \chi_1\theta_1 - \tau\theta_2$ such that $\text{var}_t(m^{\mathbb{P}}_{t,t+1}) = 0$. Because loadings are symmetric across countries, long-run UIP deviations are ruled out, and only transitory risk drives exchange rate dynamics.

Short-horizon UIP and the need for a second factor. To replicate the short-horizon failure of UIP alongside matching long-horizon UIP, a second factor is necessary in our model. To see this, consider a one-factor model variant, i.e. $z_{2,t} = 0$ and $C_n = 0$ in the above equations. In the one-factor model with no permanent innovations, the infinite-maturity bond excess return is: $\mathbb{E}_t[rx^{(\infty)}_{t,t+1}] = \frac{1}{2}\gamma z_{1,t}$. For the short-run UIP coefficient to be negative, we require $\text{cov}_t(\mathbb{E}_t[\Delta^1 e_{t+1}], r^*_{t,1} - r_{t,1}) < 0$. In Appendix C, we show that this requires $\chi - \frac{1}{2}\gamma < 0$ in the one-factor model, which implies $B_n < 0$ for all n. This contradicts the restriction required to replicate long-run UIP—namely $B_\infty > 0$—therefore necessitating our use of a second factor.

UIP in levels and conditional cyclicality. To discipline the second factor, we also draw on evidence in Engel (2016) that high-yield currencies tend to be contemporaneously appreciated, i.e. UIP holds in 'levels'. This requires the covariance between the current short-term interest rate differential and future one-period exchange rate moves to switch sign across horizons: $\text{cov}_t(\mathbb{E}_t[\Delta^1 e_{t+1}], r^*_{t,1} - r_{t,1}) < 0$ for UIP to fail in the short run, but $\text{cov}_t(\mathbb{E}_t[\Delta^1 e_{t+i}], r^*_{t,1} - r_{t,1}) > 0$ for some $i > 1$, helping to ensure UIP holds in the long run. To replicate this, in tandem with the failure of short-horizon UIP, we show in Appendix C that the the second factor must have a positive loading and be relatively persistent.

The Lemma summarises the parametric restrictions that reproduce these three empirical regularities, and shows that they generate transitory and cyclical risk within the model.

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) $(B_n(1-\phi_1) - \chi)\theta_1 + (C_n(1-\phi_2) - \tau)\theta_2 = 0$ *as* $n \to \infty$. *This rules out permanent innovations to investors' SDFs, so ensures the model matches long-run UIP.*

(ii) $\chi - \frac{1}{2}\gamma < 0$ *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* $\chi, \tau > 0$. *This ensures the model replicates short-run failures of UIP.*

(iii) The second factor is sufficiently more persistent than than the first, $\phi_2 >> \phi_1$, *and* $\tau - \frac{1}{2}\delta > 0$. *This ensures the model approximately replicate UIP in levels and generates conditional cyclicality.*

Sketch Proof. Condition (i) ensures that β^{UIP}_h is unity. Conditions (ii) and (iii) ensure that the UIP coefficient is negative at short horizons and positive at long horizons:

$$\beta^{UIP}_h = \frac{\chi \frac{1-\phi_1^h}{1-\phi_1}(B_h)\text{var}_t(z^*_{1,t} - z_{1,t}) + \tau \frac{1-\phi_2^h}{1-\phi_2}(C_h)\text{var}_t(z^*_{2,t} - z_{2,t})}{(B_h)^2\text{var}_t(z^*_{1,t} - z_{1,t}) + (C_h)^2\text{var}_t(z^*_{2,t} - z_{2,t})} \tag{22}$$

See Appendix C.1 for a full proof. □

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.[13] Using equation (17), the bond risk premium can be expressed as:

$$-\text{cov}_t(p_{t+1,n-1}, m_{t,t+1}) = B_{n-1}\sigma_1\sqrt{\gamma}z_{1,t} + C_{n-1}\sigma_2\sqrt{\delta}z_{2,t} \tag{23}$$

If condition (ii) 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 (ii)

[13] 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.

and (iii) therefore simultaneously generate cyclical SDF dynamics, and result in a sign-switch in UIP coefficients across horizons, given by equation (22), within the model.

4.2.2 Exchange Rates and the Relative Slope

We now return to the main result of the paper. In the following Proposition, we 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 Proposition states this for model-implied univariate regressions.

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 conditions (i)-(iii) in the Lemma hold.*

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 (16) 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 (19) 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.

4.2.3 Numerical Calibration

We now calibrate the two-country CIR model to satisfy the conditions in the Lemma, as well as match other empirical moments for bonds and equities. 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,[14] and pin down the 11 free model parameters. The parameter values we obtain are listed in Table 7.[15]

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.

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).[16] 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.[17] 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) or multiple correlated factors. This comes with two main costs. First, our analysis is conditional, because moments with cross-country terms,

[14] The calibration differs from the restrictions summarised in the Lemma in one small way. Although long-run UIP at the 10-year horizon cannot be statistically rejected, we find the point estimate to be below 1, so target a long-run UIP coefficient of 0.8 (i.e. condition (i) holds approximately, not exactly). Assuming long-run UIP still held exactly would still deliver the tent-shaped coefficient on the relative slope, as per the Proposition.

[15] 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.

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

[17] 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.

Table 6: Targeted moments for two-country CIR model

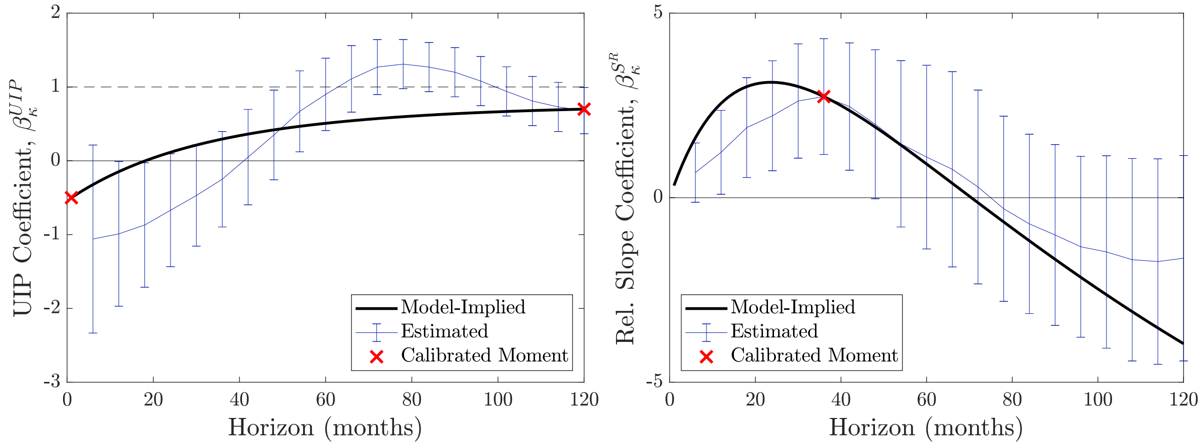
Moment	Analytical Expression	Target (Annual)
$\mathbb{E}[r_{t,1}]$	$\alpha + (\chi - \frac{1}{2}\gamma)\theta_1 + (\tau - \frac{1}{2}\delta)\theta_2$	0.40% (4.81%)
$\text{std}(r_{t,1})$	$\sqrt{(\chi - \frac{1}{2}\gamma)^2 \text{var}(z_{1,t}) + (\tau - \frac{1}{2}\delta)^2 \text{var}(z_{2,t})}$	0.31 (3.74)
$\rho(r_{t,1})$	$\frac{(\chi-\frac{1}{2}\gamma)^2\text{var}(z_{1,t})\phi_1+(\tau-\frac{1}{2}\delta)^2\text{var}(z_{2,t})\phi_2}{(\chi-\frac{1}{2}\gamma)^2\text{var}(z_{1,t})+(\tau-\frac{1}{2}\delta)^2\text{var}(z_{2,t})}$	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).

e.g. implied regression coefficients, will have an unconditional mean of zero.[18] 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 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 reproduce other well-documented exchange rate puzzles.

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

Supplementary Appendix: For Internet Publication

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)-(3) present the $\beta_{1,\kappa}$ and $\beta_{2,\kappa}$ estimates at different horizons from the slope-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.

Relative curvature. Table 10 presents results from an extended variant of regression (2) which also includes the relative yield curve curvature $C^*_{j,t} - C_t$. Column (3) indicates that there remains a tent-shaped relationship on relative slope coefficients across horizons. Although the relative curvature coefficients in column (4) appear to have a significant negative tent-shaped relationship, this relationship is not robust. Using more conservative standard errors, we do not find the relative curvature coefficients to be statistically different from zero.

In addition, Figure 5 demonstrates that the marginal explanatory power from the relative curvature is small in comparison to that of the relative slope. It plots the adjusted R^2 from regression (1), regression (2) and regression (2) with the relative curvature. The increase in explanatory power from the relative slope substantially exceeds that of the relative curvature across medium-term horizons, further justifying our focus on the relative yield curve slope.

Predictability of interest rates. Table 11 presents results of regressions of exchange rate changes on: (a) the relative yield curve slope; (b) the relative yield curve slope and curvature; 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 (c), 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

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

	(1) UIP Regression	(2) Augmented Regression	(3)
	$r^*_\kappa - r_\kappa$	$r^*_\kappa - r_\kappa$	$S^* - S$
6-months	-1.06	-0.53	0.39
	(0.65)	(1.03)	(0.65)
12-months	-0.99**	-0.37	0.86
	(0.50)	(0.83)	(0.98)
18-months	-0.87**	0.10	2.02*
	(0.43)	(0.68)	(1.06)
24-months	-0.67*	0.30	2.67**
	(0.39)	(0.62)	(1.19)
30-months	-0.47	0.57	3.58***
	(0.35)	(0.57)	(1.26)
36-months	-0.25	0.76	4.12***
	(0.33)	(0.50)	(1.25)
42-months	0.05	0.95**	4.27***
	(0.33)	(0.42)	(1.14)
48-months	0.35	1.10***	4.14***
	(0.31)	(0.34)	(1.11)
54-months	0.67**	1.31***	4.03***
	(0.28)	(0.28)	(1.13)
60-months	0.90***	1.45***	3.92***
	(0.25)	(0.26)	(1.21)
66-months	1.11***	1.57***	3.78***
	(0.23)	(0.26)	(1.25)
72-months	1.27***	1.63***	3.33***
	(0.19)	(0.23)	(1.18)
78-months	1.31***	1.55***	2.50**
	(0.17)	(0.21)	(1.08)
84-months	1.27***	1.41***	1.73*
	(0.17)	(0.19)	(1.01)
90-months	1.20***	1.28***	1.09
	(0.17)	(0.18)	(0.97)
96-months	1.08***	1.11***	0.40
	(0.17)	(0.17)	(0.96)
102-months	0.94***	0.93***	-0.09
	(0.17)	(0.17)	(1.06)
108-months	0.81***	0.78***	-0.59
	(0.17)	(0.16)	(1.16)
114-months	0.73***	0.70***	-0.78
	(0.17)	(0.16)	(1.20)
120-months	0.68***	0.65***	-0.83
	(0.16)	(0.16)	(1.20)

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)-(3) document point estimates from (2)—the augmented regression—using the relative yield curve slope (measured using a proxy) as an additional regressor. 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 canonical UIP regression and regression augmented with relative yield curve slope and curvature

	(1)	(2)	(3)	(4)
Maturity	UIP Regression	Yield Curve Augmented Regression		
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)

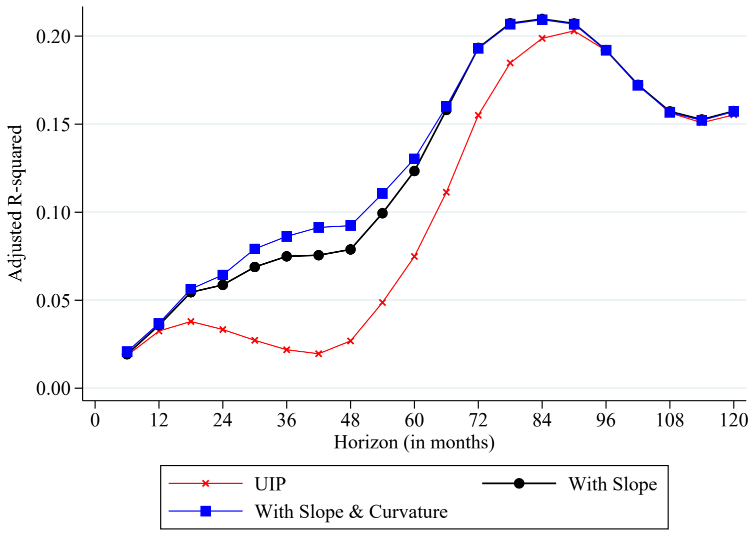
Notes: Column (1) presents coefficient estimates from regression (1)—the canonical UIP regression—a regression of the κ-period exchange rate change $\Delta^\kappa e_{t+\kappa}$ on the κ-period return differential $r^*_{t,\kappa} - r_{t,\kappa}$. Columns (2)-(4) document point estimates from an extended regression (2) 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 11: Coefficient estimates from regressions of exchange rate change on relative slope, on relative slope and curvature, and on relative level, slope and curvature and regression augmented with relative yield curve slope and curvature

Maturity	(1)	(2)	(3)	(4)	(5)	(6)
	Slope	Slope & Curvature		Level, Slope & Curvature		
κ	$S^{*}-S$	$S^{*}-S$	$C^{*}-C$	$L^{*}-L$	$S^{*}-S$	$C^{*}-C$
6-months	0.66*	1.00*	-0.67	-0.22	0.94*	-0.61
	(0.40)	(0.58)	(0.78)	(0.47)	(0.56)	(0.75)
12-months	1.19**	1.68**	-0.93	-0.20	1.62**	-0.88
	(0.57)	(0.84)	(1.16)	(0.77)	(0.82)	(1.11)
18-months	1.84***	2.37**	-0.97	0.37	2.46***	-1.04
	(0.67)	(0.98)	(1.33)	(0.99)	(0.93)	(1.25)
24-months	2.15***	2.97***	-1.56	1.16	3.26***	-1.79
	(0.74)	(1.15)	(1.69)	(1.19)	(1.07)	(1.58)
30-months	2.54***	3.51***	-1.84	2.27*	4.05***	-2.25
	(0.76)	(1.23)	(1.99)	(1.32)	(1.12)	(1.85)
36-months	2.65***	3.51***	-1.59	3.22**	4.25***	-2.12
	(0.78)	(1.15)	(1.91)	(1.45)	(1.04)	(1.74)
42-months	2.38***	3.31**	-1.75	4.35***	4.28***	-2.39
	(0.85)	(1.30)	(1.99)	(1.43)	(1.16)	(1.80)
48-months	1.89*	2.48	-1.03	5.21***	3.62**	-1.76
	(1.00)	(1.59)	(2.19)	(1.31)	(1.42)	(1.99)
54-months	1.36	1.48	-0.03	6.38***	2.87*	-0.91
	(1.11)	(1.87)	(2.40)	(1.18)	(1.66)	(2.18)
60-months	0.98	0.57	1.10	7.35***	2.14	0.14
	(1.22)	(2.03)	(2.54)	(1.23)	(1.79)	(2.30)
66-months	0.64	-0.56	2.76	8.32***	1.16	1.76
	(1.29)	(2.12)	(2.64)	(1.32)	(1.84)	(2.34)
72-months	0.15	-1.71	4.14*	9.11***	0.13	3.13
	(1.28)	(2.05)	(2.49)	(1.29)	(1.73)	(2.15)
78-months	-0.45	-2.24	4.02*	9.37***	-0.44	3.13
	(1.22)	(1.94)	(2.39)	(1.26)	(1.60)	(2.04)
84-months	-0.88	-2.22	3.10	9.23***	-0.52	2.34
	(1.16)	(1.88)	(2.43)	(1.26)	(1.55)	(2.10)
90-months	-1.23	-2.49	3.00	8.95***	-0.92	2.39
	(1.16)	(1.88)	(2.37)	(1.24)	(1.55)	(2.08)
96-months	-1.61	-2.78	2.93	8.22***	-1.36	2.38
	(1.15)	(1.99)	(2.53)	(1.21)	(1.68)	(2.26)
102-months	-1.83	-2.52	2.10	7.39***	-1.25	1.59
	(1.22)	(2.06)	(2.53)	(1.26)	(1.76)	(2.28)
108-months	-2.12*	-2.30	1.23	6.57***	-1.20	0.81
	(1.27)	(2.13)	(2.57)	(1.32)	(1.85)	(2.35)
114-months	-2.20*	-2.12	0.77	6.22***	-1.12	0.45
	(1.29)	(2.21)	(2.75)	(1.40)	(1.91)	(2.52)
120-months	-2.15	-1.34	-0.58	6.07***	-0.44	-0.82
	(1.31)	(1.96)	(2.56)	(1.51)	(1.66)	(2.33)

Notes: Column (1) presents coefficient estimates from regression of the κ-period exchange rate change $\Delta^{\kappa} e_{t+\kappa}$ on the relative yield curve slope $S^{*}-S$. Columns (2)-(3) present coefficient estimates from a regression with the relative yield curve slope *and* curvature $C^{*}-C$ as regressors. Columns (4)-(6) 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: Explanatory power of UIP regression and augmented variants 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), a relative slope-augmented UIP regression (2) (thick, black, circles) and a relative slope and curvature-augmented UIP regression (thin, blue, squares) 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.

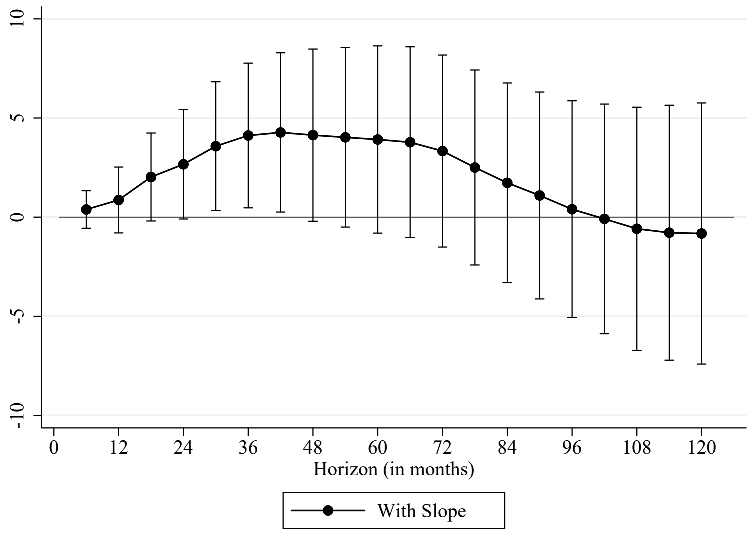
the main body of the paper are robust to heteroskedasticity and autocorrelation, we assess the robustness of our findings using alternative inference here.

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 6 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 10, point estimates are unchanged. But the error bands implied by the scaled t-statistics are wider from 12 months onwards. Nevertheless, point estimates are significantly positive according to the more conservative inference from the 2.5 to 3.5-year horizons, within which the peak of the tent arises.

In addition, Figure 7 plots the $\beta_{1,\kappa}$ and $\beta_{3,\kappa}$ coefficient estimates from an extended variant of regression (2), which includes the relative curvature, 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, justifying our focus on the relative slope in this paper.

Figure 6: 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. 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.

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 12. For comparison, column (1) includes the relative slope coefficient loadings from our benchmark sample presented in the main body of the paper. Columns (2) 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, column (3) presents an additional robustness exercises, where we use available G10 currency and yield curve data, adding Sweden, Norway and New Zealand to our cross-section of countries, for the pre-crisis period only. The relative slope loadings continue to follow a tent-shaped pattern with respect to maturity.

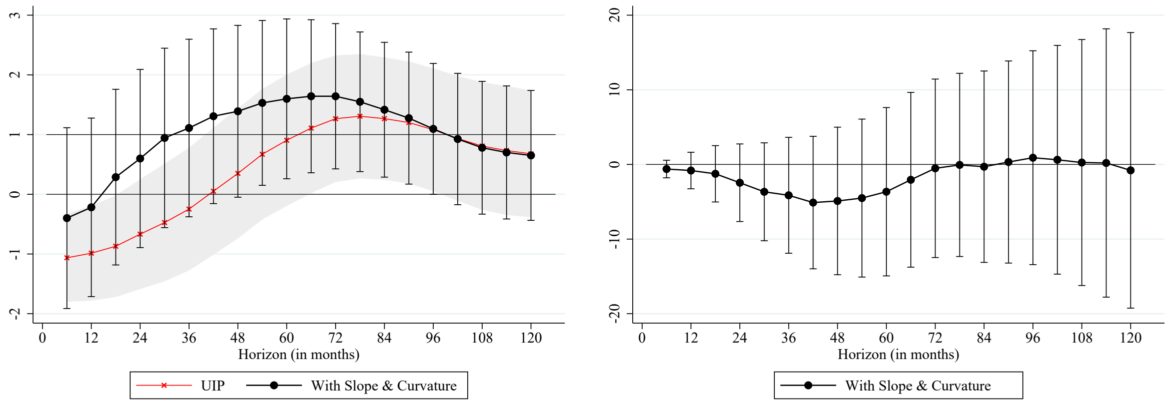
Country-specific regressions. Table 13 presents country-specific estimates of the yield curve augmented-UIP regression. Inference is conducted using Newey and West (1987) standard errors, to account for serial correlation. For comparison, column (1) presents the benchmark relative slope coefficient estimates 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 hori-

Table 12: 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 G7 Currencies	(3) 1980:01-2008:06 G10 Currencies	(4) 1990:01-2017:12 G7 Currencies
6-months	0.39	0.28	0.37	0.41
	(0.65)	(0.67)	(0.60)	(0.58)
12-months	0.86	0.54	0.71	0.36
	(0.98)	(1.04)	(0.89)	(0.75)
18-months	2.02*	1.80	1.94*	0.94
	(1.06)	(1.13)	(0.99)	(0.88)
24-months	2.67**	2.51*	2.78**	1.36
	(1.19)	(1.28)	(1.12)	(1.01)
30-months	3.58***	3.47**	3.68***	2.31**
	(1.26)	(1.36)	(1.18)	(1.09)
36-months	4.12***	4.09***	4.23***	2.88**
	(1.25)	(1.33)	(1.16)	(1.11)
42-months	4.27***	4.33***	4.35***	3.18***
	(1.14)	(1.21)	(1.06)	(1.03)
48-months	4.14***	4.23***	4.18***	3.58***
	(1.11)	(1.20)	(1.09)	(1.03)
54-months	4.03***	4.23***	4.13***	3.89***
	(1.13)	(1.24)	(1.16)	(1.13)
60-months	3.92***	4.27***	4.25***	3.78***
	(1.21)	(1.34)	(1.27)	(1.28)
66-months	3.78***	4.27***	4.23***	3.60***
	(1.25)	(1.37)	(1.31)	(1.25)
72-months	3.33***	3.88***	3.85***	3.21**
	(1.18)	(1.29)	(1.25)	(1.26)
78-months	2.50**	3.11***	2.97***	2.23*
	(1.08)	(1.15)	(1.11)	(1.16)
84-months	1.73*	2.21**	1.98*	1.26
	(1.01)	(1.05)	(1.03)	(1.06)
90-months	1.09	1.48	1.31	0.54
	(0.97)	(1.01)	(1.07)	(1.03)
96-months	0.40	0.69	0.75	-0.29
	(0.96)	(0.99)	(1.09)	(1.01)
102-months	-0.09	0.10	0.49	-0.93
	(1.06)	(1.08)	(1.21)	(1.09)
108-months	-0.59	-0.47	0.29	-1.57
	(1.16)	(1.17)	(1.34)	(1.18)
114-months	-0.78	-0.73	0.30	-1.87
	(1.20)	(1.21)	(1.37)	(1.20)
120-months	-0.83	-0.83	0.36	-1.94*
	(1.20)	(1.20)	(1.34)	(1.17)

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 7: 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.

zon 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 13: 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 14, 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 14: 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 15 shows coefficient estimates from regression (7) for local currency-bond excess returns.

Table 15: 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 (19). 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 (17) and (18).

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 (17) 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:

$$y_{t,1} = \alpha + (\chi - \frac{1}{2}\gamma)z_{1,t} + (\tau - \frac{1}{2}\delta)z_{2,t} \tag{24}$$

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 (17) and the guess; the third line uses equation (18); the fourth line rearranges the third; and the fifth line collects like terms. The recursions can be seen in the final line:

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

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 (19).

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:

$$\begin{aligned}
\mathbb{E}_t\left[rx^{(n)}_{t,t+1}\right] =& \mathbb{E}_t\left[p_{t+1,n-1} - p_{t,1} - y_{t,1}\right] \\
=& \mathbb{E}_t\Big[-A_{n-1} + A_n - B_{n-1}z_{1,t+1} + B_n z_{1,t} - C_{n-1}z_{2,t+1} + C_n z_{2,t} \\
& -\alpha - (\chi - \frac{1}{2}\gamma)z_{1,t} - (\tau - \frac{1}{2}\delta)z_{2,t}\Big] \\
=& B_{n-1}(1-\phi_1)\theta_1 + C_{n-1}(1-\phi_2)\theta_2 - B_{n-1}\mathbb{E}_t[z_{1,t+1}] + B_n z_{1,t} \\
& - C_{n-1}\mathbb{E}_t[z_{2,t+1}] + C_n z_{2,t} - (\chi - \frac{1}{2}\gamma)z_{1,t} - (\tau - \frac{1}{2}\delta)z_{2,t} \\
=& \left[-B_{n-1}\phi_1 + B_n - (\chi - \frac{1}{2}\gamma)\right] z_{1,t} + \left[-C_{n-1}\phi_2 + C_n - (\tau - \frac{1}{2}\delta)\right] z_{2,t} \\
=& \left[B_{n-1}\sigma_1\sqrt{\gamma} - \frac{1}{2}(B_{n-1}\sigma_1)^2\right] z_{1,t} + \left[C_{n-1}\sigma_2\sqrt{\delta} - \frac{1}{2}(C_{n-1}\sigma_2)^2\right] z_{2,t} \qquad (25)
\end{aligned}$$

where line 2 uses equations (19) and (24), 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 (25) can be used to express the *ex ante* bond risk premium:

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

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

recovering equation (23) in the main body.

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

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

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 $S_{t,\infty} \approx \mathbb{E}_t[rx^{(\infty)}_{t,t+1}]$ 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:

$$S_t = y_{t,120} - y_{t,6} = \frac{1}{120}\left(A_{120} + B_{120}z_{1,t} + C_{120}z_{2,t}\right) - \frac{1}{6}\left(A_6 + B_6 z_{1,t} + C_6 z_{2,t}\right) \qquad (26)$$

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:

$$\mathbb{E}_t[e_{t+\kappa}] - e_t = \sum_{i=1}^{\kappa} \mathbb{E}_t[\Delta^1 e_{t+i}] = \frac{1-\phi_1^k}{1-\phi_1}\chi(z_{1,t} - z^*_{1,t}) + \frac{1-\phi_2^k}{1-\phi_2}\tau(z_{2,t} - z^*_{2,t}) \qquad (27)$$

The one-period ERRP can be derived by combining equations (12) and (17):

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

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:

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

$$= \frac{\chi \frac{1-\phi_1^h}{1-\phi_1} B_h \text{var}_t(z_{1,t}^* - z_{1,t}) + \tau \frac{1-\phi_2^h}{1-\phi_2} C_h \text{var}_t(z_{2,t}^* - z_{2,t})}{B_h^2 \text{var}_t(z_{1,t}^* - z_{1,t}) + C_h^2 \text{var}_t(z_{2,t}^* - z_{2,t})} \tag{28}$$

where line 1 is the definition of the univariate regression coefficient, and line 2 uses equation (27) and the definition for returns.

C.1 Proof to Lemma

Consider the UIP coefficient in equation (28) for a general horizon h.

Condition (i). To rule out permanent innovations to investors' SDFs, we use the expression for the expected (log) bond excess return, equation (25). 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 (ii). 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 (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 (28) 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. □

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:

$$\begin{aligned} \beta_h^{SR} &= \frac{\text{cov}_t(\mathbb{E}_t[e_{t+h} - e_t], (y_{t,120}^* - y_{t,6}^*) - (y_{t,120} - y_{t,6}))}{\text{var}_t((y_{t,120}^* - y_{t,6}^*) - (y_{t,120} - y_{t,6}))} \\ &= \frac{\frac{1-\phi_1^h}{1-\phi_1}\chi\left[\frac{1}{120}B_{120} - \frac{1}{6}B_6\right]\text{var}_t(z_{1,t}^* - z_{1,t}) + \frac{1-\phi_2^h}{1-\phi_2}\tau\left[\frac{1}{120}C_{120} - \frac{1}{6}C_6\right]\text{var}_t(z_{2,t}^* - z_{2,t})}{\left[\frac{1}{120}B_{120} - \frac{1}{6}B_6)\right]^2\text{var}_t(z_{1,t}^* - z_{1,t}) + \left[\frac{1}{120}C_{120} - \frac{1}{6}C_6\right]^2\text{var}_t(z_{2,t}^* - z_{2,t})} \end{aligned} \tag{29}$$

where line 1 is the definition of the univariate regression coefficient, and line 2 uses equations (26) and (27).

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. □

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