BANK OF ENGLAND

Staff Working Paper No. 872

Exchange rate risk and business cycles

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June 2020

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Abstract

We show that currencies with a steeper yield curve tend to depreciate at business cycle horizons, in violation of uncovered interest parity. The yield curve adds no explanatory power over and above spot yield differentials in explaining exchange rates at longer horizons. Analysing bond holding period returns, we identify a tent-shaped relationship between the exchange rate risk premium and the relative slope across horizons. We derive this relationship analytically within an asset pricing framework and show it is driven by differences in transitory innovations to investors' stochastic discount factor, captured by the relative yield curve slope and consistent with business cycle risk. Our mechanism is robust to the inclusion of liquidity yields, which instead contribute to explaining cross-sectional differences across currencies and reflect permanent innovations to investors' stochastic discount factor.

Key words: Business cycle risk, exchange rates, risk premia, stochastic discount factor, uncovered interest parity, yield curves.

JEL classification: E43, F31, G12.

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The views expressed in this paper are those of the authors, and not necessarily those of the Bank of England or its committees. We are especially grateful to Giancarlo Corsetti for many helpful discussions. We also thank an anonymous referee, Gianluca Benigno, Ambrogio Cesa-Bianchi, Luca Dedola, Pierre-Olivier Gourinchas, Matteo Maggiori, Ilaria Piatti (discussant), Katrin Rabitsch and Lucio Sarno 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 and University of Cambridge 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'.

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ISSN 1749-9135 (on-line)

1 Introduction

Uncovered interest parity (UIP) predicts that, under risk neutrality, a high interest rate currency should depreciate to equalise exchange rate-adjusted returns on assets. 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) which reflect differences in the conditional volatilities of stochastic discount factors (SDFs), as shown in Backus, Foresi, and Telmer (2001). Since the UIP hypothesis cannot be rejected at long horizons (e.g. Chinn and Meredith, 2005; Engel, 2016), ERRP must reflect differences in the volatilities of the transitory component of SDFs, as defined in Alvarez and Jermann (2005). Based on these these two pieces of evidence, which together set the stage for our analysis of the 'UIP puzzle', in this paper we explore the relationship between the term structure of interest rates and ERRP. We argue that ERRP arise to compensate investors for business cycle risk across currencies, captured by relative yield curve slopes.

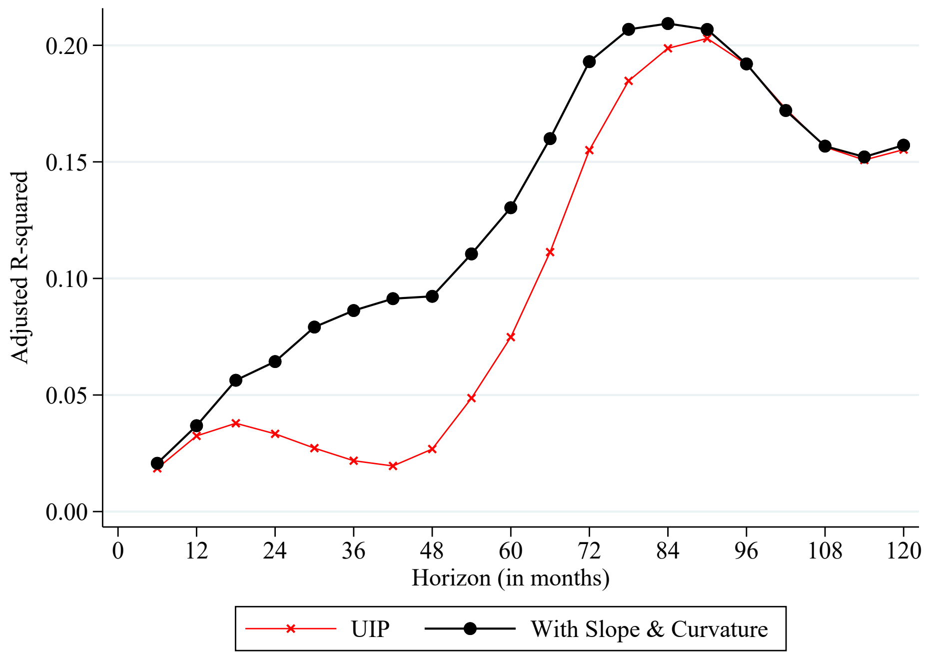
To this end, we begin by showing that information in the yield curve, over and above spot interest rate differentials, greatly improves explanatory power for exchange rates, specifically at business cycle horizons. We do so both by estimating UIP regressions across different maturities and by analysing excess returns over different holding periods. We then interpret our findings within a standard no-arbitrage model and decompose the SDF into transitory and permanent components. In an analytical example, we show that the relative yield curve slope prices transitory innovations to investors SDFs and we derive the relationship between ERRP and the relative slope. Our results suggest that exchange rate movements in excess of UIP systematically reallocate returns intertemporally to investors who value them most highly.

Figure 1 plots the adjusted R^2 of a canonical panel UIP regression—a regression of the κ-period-ahead exchange rate change on κ-maturity cross-country yield differentials—together with the adjusted R^2 from the same regression extended to include measures of the cross-country relative yield curve slope and curvature. We find strong evidence that information in the yield curve can account for exchange rate fluctuations over and above the spot rate differential at business cycle horizons (3 to 4 years). At these horizons, the adjusted R^2 of our augmented regression is around treble that of the canonical UIP regression. Coefficient estimates indicate that a country with a steeper yield curve tends to experience a depreciation over time, and the relationship exhibits a *tent shape* with respect to the horizon: rising from zero at short horizons, achieving a peak around medium, business cycle, horizons and falling to zero at longer horizons.[1] We demonstrate the robustness of these results in a number of ways, including by distinguishing between the predictability of exchange and interest rates, using conservative inference, assessing sub-sample stability and looking at country pairs.

Building on this, we consider a flexible empirical specification which allows bond holding periods and maturities to differ. Decomposing excess returns on bonds in dollar terms into an ERRP and a local-currency bond premium (Lustig, Stathopoulos, and Verdelhan, 2019), we

[1] Throughout, the exchange rate is defined as the domestic price of foreign currency, so an increase or a positive coefficient denotes a domestic depreciation.

Figure 1: Explanatory power of UIP regression augmented with relative yield curve slope and curvature at different horizons



Notes: Plot of the adjusted R^2 from the standard UIP regression of *ex post* exchange rate changes on horizon-specific interest rate differentials (thin, red, crosses) and a UIP regression augmented with the relative yield curve slope and curvature (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.

regress each component—for different maturity bonds over varying holding periods—on relative yield curve slopes. We find that the relative slope is a significant predictor of both ERRP and dollar-bond risk premia at holding periods associated with business cycle horizons (around 3 years), even with long-maturity bonds (up to 10 years).[2] Nonetheless, the term structure of carry trade is decreasing with maturity at every holding period, consistent with the empirical findings in Lustig et al. (2019).

Taken together, our results indicate that the term structure of interest rates explains significant variation in ERRP at medium horizons. We argue that these findings point to an important role for transitory innovations to investors' SDFs, which we later show are captured by the relative yield curve slope. In a standard two-country asset pricing setup, ERRP arise as equilibrium outcomes, necessary to compensate risk-averse investors for macroeconomic risks. For instance, an expected foreign exchange appreciation (domestic depreciation) increases the expected return from a foreign bond in domestic currency, while a subsequent depreciation low-

[2] These holding-period regressions also help to assuage worries around the limited number of non-overlapping observations in long-horizon UIP regressions. They can be interpreted as a hybrid regression specification, in between the long-horizon regressions of Chinn and Meredith (2005) and the one-month holding-period return regressions for 10-year bonds in Lustig et al. (2019).

ers future expected returns. Taking the correlation between ERRP and spot yield differentials from the data and through the lens of this analytical framework, UIP failures appear to systematically reallocate returns to investors with a relatively high valuation of current returns, lowering the risk for home investors holding foreign assets.

We then explain the relationship between ERRP and the relative yield curve slope. When yield curves are upward sloping on average, nearer-term return valuations—captured by nearer-term SDFs—are high relative to longer-horizon valuations reflecting business cycle risk.[3] In our two-country setting, both the current relative valuation of returns and the future path of relative valuations matter for ERRP. The relative yield curve slope influences exchange rate dynamics because it captures investors' *relative* desire to reallocate returns intertemporally—i.e. in response to transitory innovations to their SDF. In contrast, permanent SDF innovations do not result in a desire for intertemporal reallocation of returns and consequently are neither captured by bond premia, nor require exchange rates to reallocate returns across time. Thus, we argue that the relationship between ERRP and the relative yield curve slope is driven by the need to compensate investors for business cycle risk. To derive the relationship between ERRP and the relative slope analytically, we set up two stylised examples, where we assume complete asset markets and mean-reverting first or second-order autoregressive processes for investors' pricing kernels. In both, the relationship between ERRP and the relative slope is positive and in the second case, we show that the relationship can have a tent shape with respect to the horizon, mirroring our empirical results.

Finally, we extend our empirical specification to account for liquidity yields—the non-monetary return that government bonds provide because of their safety, ease of resale, and value as collateral—highlighted by Engel and Wu (2018) as an important predictor of exchange rate movements. The extension serves two purposes. First, using data on US Treasury premia from Du, Im, and Schreger (2018), we show that our main results regarding relative yield curve factors and business cycle risk are robust to the inclusion of cross-country liquidity yields as additional regressors. Second, within our theory, we show that, in contrast to business cycle risk, liquidity yields appear to reflect permanent innovations to SDFs. We find evidence that the addition of horizon-specific cross-country liquidity yields increases the explanatory power of our yield curve-augmented regression at medium to long horizons, suggesting that liquidity yields reflect permanent innovations to SDFs. So, while liquidity yields are an important factor for understanding the cross-sectional dimension of UIP failures, business cycle risk—our main focus—reflects the time-series dimension of UIP failures.

Related Literature Our work is related to a classic literature on the forward premium puzzle rooted in Hansen and Hodrick (1980) and Fama (1984), and analysis of the UIP across time (Engel, 2016) and horizons (e.g. Chinn and Meredith, 2005; Chinn and Quayyum, 2012). 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.

[3] Wachter (2006) and Piazzesi and Schneider (2007) discuss this negative intertemporal correlation of SDFs in a closed economy setup.

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 horizons. 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 symmetry of 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 differ across countries.

Our empirical specification with holding period returns builds on Lustig et al. (2019). They show that, for a given one-month holding period, the term structure of carry trade is decreasing, implying that differences in permanent SDF innovations across countries are small. By extending the specification across holding periods, we show that the relative slope is a significant predictor of ERRP at holding periods associated with business cycle horizons, for any given maturity, thus emphasising differences in transitory SDF innovations as drivers of ERRP. We interpret this as evidence that business cycle risk, priced into the yield curve, can explain time series variation in ERRP. Colacito, Riddiough, and Sarno (2019) also attribute a role to business cycles in explaining ERRP, albeit in the cross-section, by 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.

We further contribute to a literature studying the role of liquidity for exchange rate dynamics (see, e.g. Engel and Wu, 2018; Jiang, Krishnamurthy, and Lustig, 2018). A key novelty of our empirical setup is to investigate the relationship between the term structure of liquidity yields and exchange rates at different horizons, extending the results of Engel and Wu (2018) who study the 1-year tenor one.

In the remainder of this paper, Section 2 reprises empirical evidence on UIP at different horizons and introduces our two-country preference-free theoretical environment. Section 3 highlights the role of the relative yield curve slope in explaining ERRP and business cycle risk. Section 4 extends the analysis to account for liquidity yields, and Section 5 concludes.

2 UIP Puzzle Redux

In this section, we first define notation, then summarise the empirical performance of UIP at different horizons, interpreting UIP failures within a preference-free setting.

2.1 Notation

We set up an environment in which there are two countries—Home and Foreign (the latter denoted by an asterisk)—each with a representative investor. We maintain two key assumptions throughout. First, all investors are risk averse. Second, bonds in each country are priced by domestic agents. The Home and Foreign SDFs spanning the period t to $t + \kappa$ are denoted by

$M_{t,t+\kappa} \leq 1$ and $M^*_{t,t+\kappa} \leq 1$, respectively. We assume that these SDFs satisfy Euler equations for Home and Foreign κ-period risk-free zero-coupon bonds, with prices $P_{t,\kappa} \leq 1$ and $P^*_{t,\kappa} \leq 1$, respectively: $P_{t,\kappa} = \mathbb{E}_t\left[M_{t,t+1}P_{t+1,\kappa-1}\right]$ and $P^*_{t,\kappa} = \mathbb{E}_t\left[M^*_{t,t+1}P^*_{t+1,\kappa-1}\right]$. By forward iteration using $M^{(*)}_{t,t+\kappa} \equiv \prod_{i=0}^{\kappa-1} M^{(*)}_{t+i,t+i+1}$ and the law of iterated expectations, this implies:[4]

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

$$1 = \mathbb{E}_t\left[M^*_{t,t+\kappa}R^*_{t,\kappa}\right] \tag{2}$$

where $R^{(*)}_{t,\kappa} \equiv 1/P^{(*)}_{t,\kappa} \equiv (1 + i^{(*)}_{t,\kappa}) \geq 1$ is the gross return on the Home (Foreign) κ-period zero-coupon bond. Additionally, it is useful to define the pricing kernels $V^{(*)}_t \geq 0$ that comprise the SDF as $M^{(*)}_{t,t+\kappa} \equiv V^{(*)}_{t+\kappa}/V^{(*)}_t$.

When engaging in cross-border asset trade, a risk-averse Home agent with κ-period SDF $M_{t,t+\kappa}$ prices risk-free κ-period Foreign currency-denominated assets according to:

$$1 = \mathbb{E}_t\left[M_{t,t+\kappa}\frac{\mathcal{E}_{t+\kappa}}{\mathcal{E}_t}R^*_{t,\kappa}\right] \tag{3}$$

where $\mathcal{E}_t$ is the exchange rate, defined as the Home price of a unit of Foreign currency such that an increase in $\mathcal{E}_t$ corresponds to a Home depreciation. Assuming $\mathcal{E}_t$ and $M_{t,t+\kappa}$ are jointly log-normally distributed, international no-arbitrage requires that the exchange rate satisfies:[5]

$$\mathbb{E}_t\left[\Delta^\kappa e_{t+\kappa}\right] + \frac{1}{2}\text{var}_t\left(\Delta^\kappa e_{t+\kappa}\right) = \left(i_{t,\kappa} - i^*_{t,\kappa}\right) - \text{cov}_t\left(m_{t,t+\kappa}, \Delta^\kappa e_{t+\kappa}\right) \tag{4}$$

where $e_t \equiv \log(\mathcal{E}_t)$, $\Delta^\kappa e_{t+\kappa} \equiv e_{t+\kappa} - e_t$, $i^{(*)}_{t,\kappa} \equiv \log(R^{(*)}_{t,\kappa})$, and $m_{t,t+\kappa} \equiv \log(M_{t,t+\kappa})$. Expected κ-period exchange rate changes should be proportional to κ-period interest differentials, corrected for the covariance between investors' SDF and exchange rates. If investors were risk neutral, and absent financial frictions, the covariance term would drop away and exchange rate-adjusted returns should be equated in expectation.

Standard empirical methods provide evidence on the average ERRP demanded by Home investors on Foreign bonds and Foreign investors on Home bonds (Engel, 2014), given by:

$$\lambda_{t,\kappa} = -\text{cov}_t\left(\frac{m_{t,t+\kappa} + m^*_{t,t+\kappa}}{2}, \Delta^\kappa e_{t+\kappa}\right) \tag{5}$$

This equilibrium ERRP reflects the covariance of the cross-country average SDF for the period t to $t+\kappa$ with corresponding-horizon exchange rate dynamics.

[4] Throughout the paper we only consider nominal values, consistent with our data. Since the SDF is itself nominal, the examples and intuition we present should be interpreted in terms of utility units. If prices are fixed, movements in valuation are then entirely driven by changes to consumption growth.

[5] The assumption of log-normality is often relaxed in recent literature, which instead employs a measure of entropy $\mathcal{L}(\cdot)$ instead of variance var$(\cdot)$. This is defined according to $\mathcal{L}_t(X_{t+1}) = \log \mathbb{E}_t[X_{t+1}] - \mathbb{E}_t \log[X_{t+1}]$ (see Backus, Boyarchenko, and Chernov, 2018). For our purposes, the assumption of log-normality yields analytical results parsimoniously.

2.2 Canonical UIP Regression

Motivated by (4) under the joint assumption of risk neutrality and rational expectations, we estimate a sequence of regressions for each κ-month horizon using panel data for a cross-section of countries j over time t:

$$e_{j,t+\kappa} - e_{j,t} = \beta_{1,\kappa}\left(i_{j,t,\kappa} - i^*_{t,\kappa}\right) + f_{j,\kappa} + u_{j,t+\kappa} \tag{6}$$

where $e_{j,t}$ is the (log) exchange rate of country j *vis-à-vis* the base currency at time t, $i_{j,t,\kappa}$ is the net κ-period return in country j at time t, $i^*_{t,\kappa}$ is the equivalent return in the base currency, $f_{j,\kappa}$ is a country fixed effect, and $u_{j,t+\kappa}$ is the disturbance.

Under the null hypothesis of UIP, $\beta_{1,\kappa} = 1$ for all $\kappa > 0$.[6] Empirical rejections of UIP at short to medium horizons—i.e. finding $\hat{\beta}_{1,\kappa} \neq 1$ in regression (6) for small to medium κ—have regularly been used to motivate claims that interest rates do not adequately explain exchange rate dynamics.

Data To estimate our regressions, we use exchange rate and interest rate data for six jurisdictions with liquid bond markets: Australia, Canada, Switzerland, the euro area, Japan and the United Kingdom. Additionally, the United States acts as the base country, such that our benchmark sample covers G7 currencies. To capture the term structure of interest rates in each country, we use nominal zero-coupon government bond yield data of the following maturities: 6, 12, 18, ..., 120 months. Nominal zero-coupon government bond yield curves are obtained from a combination of sources, including central banks and Wright (2011), detailed in Appendix A. Our nominal exchange rate data is from *Datastream*, measuring the value of domestic currency price per unit of US dollar. We use end-of-month data from 1980:01 to 2017:12.[7]

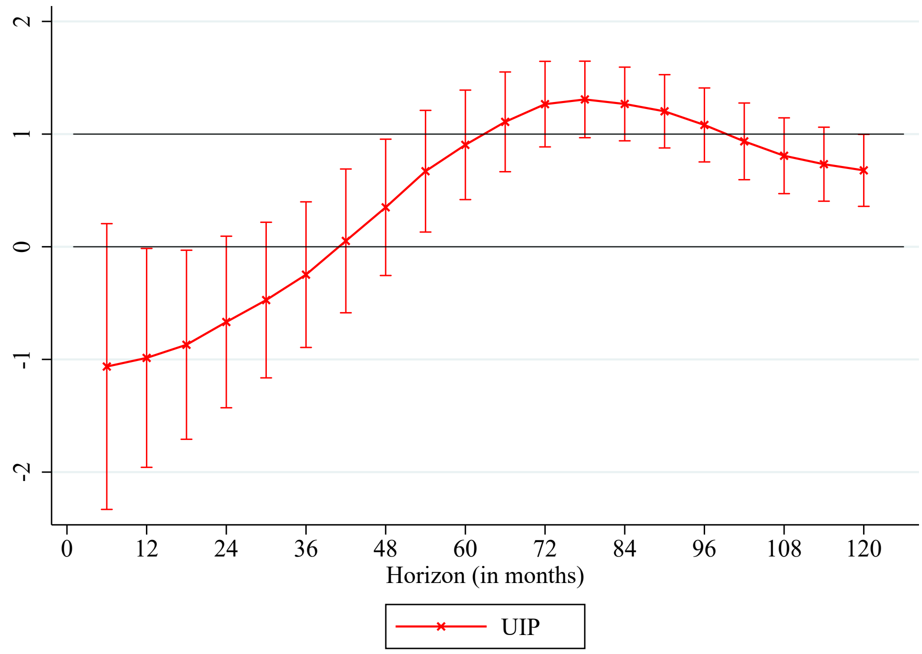
Results Figure 2 plots UIP coefficient $\beta_{1,\kappa}$ estimates from a panel regression with country fixed effects over our benchmark sample.[8] The confidence bands around these point estimates are derived from Driscoll and Kraay (1998) standard errors, which correct for heteroskedasticity, serial correlation and cross-equation correlation. The coefficient estimates in figure 2 reinforce the view that the UIP hypothesis can be rejected at short to medium horizons, but cannot be rejected at longer horizons. At 6 to 36-month horizons, point estimates are negative, indicating that high short-term interest rate currencies tend to appreciate, instead of depreciate, in line with Fama (1984). 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, *inter alia*, Chinn and Meredith (2005) and Chinn and Quayyum (2012).

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

[7] As Appendix A documents, our panel of nominal zero-coupon government bond yields is unbalanced, with different countries entering the sample at different dates.

[8] The same results are tabulated in column (1) of table 1, and the adjusted R^2 of each regression is plotted in figure 1.

Figure 2: Estimated coefficients from canonical UIP regression at different horizons



Notes: Red crosses denote $\hat{\beta}_{1,\kappa}$ estimates from regression (6). 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.

2.3 ERRP and Business Cycle Risk

To interpret UIP failures and motivate the role of business cycle risk and the yield curve, we build on the preference-free environment from Section 2.1. We first define the κ-period *ex post* ERRP at time t as:

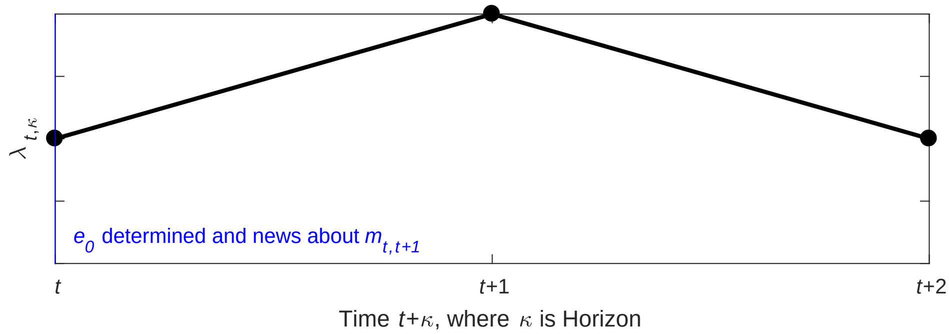
$$\lambda_{t,\kappa} \equiv i^*_{t,\kappa} - i_{t,\kappa} + \Delta^\kappa e_{t+\kappa} \tag{7}$$

Substituting (6) into (7), implies that $\text{cov}_t(\lambda_{t,\kappa}, i_{t,\kappa} - i^*_{t,\kappa}) < 0$ at short to medium horizons where $\hat{\beta}_{1,\kappa} < 1$. In periods when the Home interest rate is relatively high $i_{t,\kappa} > i^*_{t,\kappa}$, the *ex post* ERRP on Foreign currency is negative $\lambda_{t,\kappa} < 0$ resulting in an excess appreciation, and *vice versa* when $i_{t,\kappa} < i^*_{t,\kappa}$.

In our asset pricing framework, a low Home interest rate $i_{t,\kappa} < i^*_{t,\kappa}$, via the domestic bond-pricing Euler equations (1) and (2), is associated with a relatively high valuation of returns in $t+\kappa$ by Home investors, $\mathbb{E}_t[m_{t,t+\kappa}] > \mathbb{E}_t[m^*_{t,t+\kappa}]$. Conditional on this, our estimated covariance implies that the ERRP $\lambda_{t,\kappa} > 0$, making Foreign bonds more attractive since their effective return is relatively high when investors value returns most.

Consider a three-period illustrative example, where at time t investors trade in 1 and 2-period bonds that mature at time $t+1$ and $t+2$, respectively. While the time-t return on 2-period bonds are equal across countries $i_{t,2} = i^*_{t,2}$, consider a drop in the Home 1-period

Figure 3: Illustrative path of ERRP $\lambda_{t,\kappa}$ around a transitory exchange rate depreciation for representative Home investor, with $i_{t,1} < i^*_{t,1}$ and $i_{t,2} = i^*_{t,2}$



interest rate $i_{t,1} < i^*_{t,1}$, which could, for example, arise from an increase in the probability of a low state of consumption at $t+1$ compensated by a commensurate change at $t+2$.[9] By (1) and (2), it must be the case that $\mathbb{E}_t[m_{t,t+1}] > \mathbb{E}_t[m^*_{t,t+1}]$, while $\mathbb{E}_t[m_{t,t+2}] = \mathbb{E}_t[m^*_{t,t+2}]$. Home investors will value one unit of Home currency at $t+1$ more than Foreign investors value a unit of Foreign currency over the same horizon. Given that $\text{cov}_t(\lambda_{t,\kappa}, i_{t,\kappa} - i^*_{t,\kappa}) < 0$ in the data, the ERRP will be higher between t and $t+1$ than between t and $t+2$, as illustrated in figure 3. From the vantage point of the Home investor who holds both Home and Foreign bonds, the ERRP reallocate returns intertemporally—both from t to $t+1$ (over which horizon the Home currency is expected to depreciate in excess of UIP), and from $t+2$ to $t+1$ (by virtue of the subsequent currency appreciation).

However, the 1-period yield differential only captures part of the incentive for intertemporal reallocation of returns. The subsequent appreciation between $t+1$ and $t+2$ is desirable in our illustrative example because we have equalised long-rates, i.e. $i_{t,2} = i^*_{t,2}$. In general, an investor with a relatively high long-term valuation of returns, reflected by $i_{t,2} < i^*_{t,2}$, would benefit less from a high ERRP between t and $t+1$ because the subsequent appreciation from $t+1$ to $t+2$ would be relatively costly. In an infinite-horizon setting, the equilibrium depreciation at $t+1$ therefore depends on the path of future valuations $\{m^{(*)}_{t+\kappa-1,t+\kappa}\}_{\kappa=1,2,\ldots,\infty}$ which governs the losses from a subsequent expected appreciation, and is captured by the term structure of interest rates. The ERRP, which reallocates returns to investors in periods where they value them relatively highly, compensates them for business cycle risk. The resulting role of the term structure in ERRP determination is the central message of this paper and we further develop this relationship in Section 3.

This mechanism does not hinge on the assumed degree of financial market completeness.

[9] This could arise from an unaccommodated demand shock or monetary policy shocks at t. For example, Benigno, Benigno, and Nisticò (2012) show that to a first order in a large general equilibrium model, keeping the real rate fixed and under flexible prices, $\mathbb{E}_t[m_{t,t+1}] = -\pi_t$ where π_t is the inflation rate and is determined as the sum of an inflation target and monetary policy shocks.

The ERRP can be written as:

$$\lambda_{t,\kappa} = \frac{1}{2}[\mathrm{var}_t(m_{t,t+\kappa}) - \mathrm{var}_t(m^*_{t,t+\kappa})] + \eta_{t+\kappa} \tag{8}$$

where the first, difference in variance, term reflects a 'complete markets' ERRP, while $\eta_{t+\kappa} = e_{t+\kappa} - e_t - (m^*_{t,t+\kappa} - m_{t,t+\kappa})$ captures non-traded risk in an incomplete markets framework (Lustig and Verdelhan, 2019). In Section 3, we focus on the complete markets benchmark because the exchange rate is uniquely determined, providing a useful benchmark which does not require a specification of the nature of shock processes affecting the economy.[10]

3 UIP and the Yield Curve

In this section, we demonstrate that currencies with relatively steep yield curves tend to depreciate most strongly at business cycle horizons. We extend our preference-free setup to interpret this finding for ERRP, attributing it to business cycle risk that is captured in yield curves because they contain information on transitory variation in SDFs.

3.1 Yield Curve-Augmented UIP Regression

We augment regression (6) with measures of the relative yield curve slope $S_{j,t} - S^*_t$ and relative yield curve curvature $C_{j,t} - C^*_t$. For all κ, we estimate the extended panel regression:

$$e_{j,t+\kappa} - e_{j,t} = \beta_{1,\kappa}\left(i_{j,t,\kappa} - i^*_{t,\kappa}\right) + \beta_{2,\kappa}(S_{j,t} - S^*_t) + \beta_{3,\kappa}(C_{j,t} - C^*_t) + f_{j,\kappa} + u_{j,t+\kappa} \tag{9}$$

where $S_{j,t}$ ($C_{j,t}$) is the slope (curvature) of the country j yield curve at time t and S^*_t (C^*_t) is the slope (curvature) of the base country yield curve.

Along with the yield curve level, the slope and curvature are known to capture a high degree of variation in bond yields (Litterman and Scheinkman, 1991). We do not include the relative level in our baseline regression in order to nest UIP, enabling interpretation of the yield curve's contribution over and above spot yield differentials. Combining (7) with (9), the *ex post* κ-period ERRP can be expressed as:

$$\lambda_{j,t,\kappa} = (\beta_{1,\kappa} - 1)\left(i_{j,t,\kappa} - i^*_{t,\kappa}\right) + \beta_{2,\kappa}(S_{j,t} - S^*_t) + \beta_{3,\kappa}(C_{j,t} - C^*_t) + f_{j,\kappa} + u_{j,t+\kappa} \tag{10}$$

Comparing this with (9), then $\beta_{2,\kappa}$ can be interpreted as *either* the average domestic depreciation (in percent) *or* the average increase in the ERRP (in pp) associated with a 1pp increase in the slope of the domestic yield curve relative to the US (base) country.

To measure the yield curve slope and curvature, we use proxies. We define the slope as the difference between 10-year and 6-month yields, $S_{j,t} = i_{j,t,10y} - i_{j,t,6m}$. Our curvature proxy is a butterfly spread, a function of 6-month, 5 and 10-year yields (Diebold and Rudebusch, 2013),

[10] If markets are incomplete, for any given pair of Home and Foreign SDFs, the exchange rate process is not uniquely determined, although the complete markets outcome above remains an admissible equilibrium (Backus et al., 2001).

$C_{j,t} = 2i_{j,t,5y} - (i_{j,t,6m} + i_{j,t,10y})$. We prefer these measures to principal component estimates of the yield curve slope and curvature, as principal component measures potentially contain look-ahead bias, being defined using weights constructed from information in the whole sample. By construction, our slope proxy is only based on information available up to time t. Nevertheless, our findings are robust to the use of principal components. Our relative yield curve proxies are then constructed by taking cross-country differences. Since our proxies are derived from yield curves estimated on a country-by-country basis, we do not assume any symmetry in the factor structure of yield curves across countries.

Results Our benchmark results for regression (9) are documented in table 1. Columns (2)-(4) present the $\beta_{1,\kappa}$, $\beta_{2,\kappa}$ and $\beta_{3,\kappa}$ estimates at different horizons from the panel regression using pooled monthly data from 1980:01 to 2017:12. For comparison, column (1) includes the $\beta_{1,\kappa}$ estimates from the canonical UIP regression (6). In addition, column (5) documents the implied $\hat{\beta}_{1,\kappa} - 1$ estimate from the augmented regression, which represents the association between the *ex post* ERRP and interest rate differential. Driscoll and Kraay (1998) standard errors are reported in parentheses.

Three observations are particularly noteworthy. First, the broadly upward sloping relationship between the UIP coefficient $\beta_{1,\kappa}$ and horizon κ is robust to the augmentation of the UIP regression with the relative yield curve slope and curvature. This implies that the additional contribution of relative slope and curvature can be interpreted over and above the role for spot interest rate differentials, as an additional component of the ERRP.

Second, and most importantly, our point estimates of $\beta_{2,\kappa}$ reveal a tent-shaped relationship with respect to horizon κ between the relative yield curve slope and κ-period exchange rate dynamics. Figure 4 shows this visually, plotting $\beta_{2,\kappa}$ with respect to κ. Coefficients on the slope differential are insignificantly different from zero at short horizons, but increase in sign and significance from short to medium horizons. The $\hat{\beta}_{2,\kappa}$ coefficient peaks at the 3.5-year horizon, quantitatively indicating that a one percentage point increase in a country's yield curve slope relative to the US is, on average, associated with a 7.40% exchange rate depreciation over that horizon. At longer horizons—from 6.5-years onwards—the loading on the relative slope is insignificantly different from zero.

Third, focusing on column (5) of table 1, our results show that, after controlling for relative yield curve factors, interest rate differentials provide no significant explanatory power for *ex post* ERRP at short to medium horizons—from 6 to 48-months.

3.1.1 Robustness

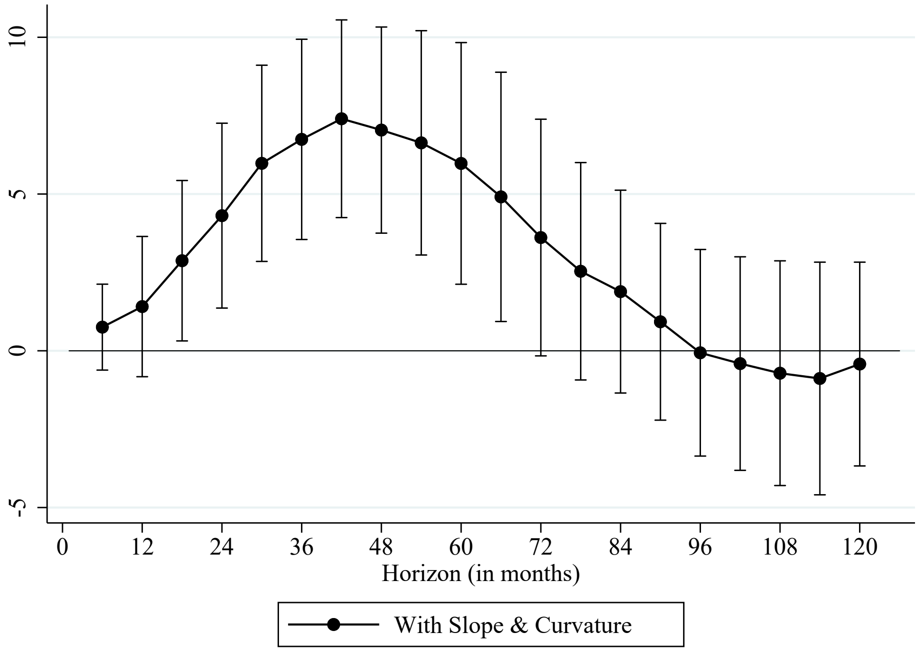
In this sub-section, we summarise the robustness of our main empirical finding: that countries with a steeper yield curve tend to experience a subsequent currency depreciation at business cycle horizons. Further details on the robustness exercises can be found in Appendix B.2.

Table 1: Coefficient estimates from canonical UIP regression and regression augmented with relative yield curve slope and curvature

Maturity	(1) UIP Regression	(2)	(3) Yield Curve Augmented Regression	(4)	(5)
κ	$i_\kappa - i^*_\kappa$	$i_\kappa - i^*_\kappa$	$S - S^*$	$C - C^*$	*Implied* $\hat{\beta}_{1,\kappa} - 1$
6-months	-1.06	-0.40	0.75	-0.61	*-1.40*
	(0.65)	(1.00)	(0.70)	(0.74)	
12-months	-0.99**	-0.22	1.41	-0.82	*-1.22*
	(0.50)	(0.82)	(1.14)	(1.09)	
18-months	-0.87**	0.29	2.87**	-1.25	*-0.71*
	(0.43)	(0.69)	(1.31)	(1.23)	
24-months	-0.67*	0.60	4.31***	-2.45	*-0.40*
	(0.39)	(0.62)	(1.50)	(1.53)	
30-months	-0.47	0.94*	5.98***	-3.67**	*-0.06*
	(0.35)	(0.56)	(1.60)	(1.77)	
36-months	-0.25	1.11**	6.74***	-4.13**	*0.11*
	(0.33)	(0.52)	(1.63)	(1.74)	
42-months	0.05	1.31***	7.40***	-5.11***	*0.31*
	(0.33)	(0.44)	(1.61)	(1.86)	
48-months	0.35	1.39***	7.04***	-4.89**	*0.39*
	(0.31)	(0.35)	(1.68)	(2.03)	
54-months	0.67**	1.53***	6.63***	-4.51**	*0.53**
	(0.28)	(0.28)	(1.83)	(2.20)	
60-months	0.90***	1.60***	5.98***	-3.66	*0.60***
	(0.25)	(0.27)	(1.97)	(2.31)	
66-months	1.11***	1.64***	4.91**	-2.06	*0.64***
	(0.23)	(0.26)	(2.03)	(2.37)	
72-months	1.27***	1.64***	3.61*	-0.52	*0.64****
	(0.19)	(0.23)	(1.93)	(2.21)	
78-months	1.31***	1.55***	2.54	-0.06	*0.55****
	(0.17)	(0.21)	(1.77)	(2.09)	
84-months	1.27***	1.42***	1.89	-0.30	*0.42***
	(0.17)	(0.19)	(1.65)	(2.10)	
90-months	1.20***	1.28***	0.93	0.32	*0.28*
	(0.17)	(0.18)	(1.60)	(2.07)	
96-months	1.08***	1.10***	-0.06	0.90	*0.10*
	(0.17)	(0.16)	(1.68)	(2.24)	
102-months	0.94***	0.93***	-0.41	0.63	*-0.07*
	(0.17)	(0.16)	(1.74)	(2.25)	
108-months	0.81***	0.78***	-0.71	0.25	*-0.22*
	(0.17)	(0.16)	(1.83)	(2.31)	
114-months	0.73***	0.70***	-0.88	0.20	*-0.30**
	(0.17)	(0.16)	(1.89)	(2.50)	
120-months	0.68***	0.65***	-0.42	-0.79	*-0.35***
	(0.16)	(0.16)	(1.66)	(2.34)	

Notes: Column (1) presents coefficient estimates from regression (6)—the canonical UIP regression—a regression of the κ-period exchange rate change $\Delta^\kappa e_{t+\kappa}$ on the κ-period interest rate differential $i_{t,\kappa} - i^*_{t,\kappa}$. Columns (2)-(4) document point estimates from (9)—the augmented regression—using the relative yield curve slope and curvature (measured using proxies) as additional regressors. Column (5) documents the implied $\hat{\beta}_{1,\kappa} - 1$ estimates, and associated statistical significance; this corresponds to the coefficient on interest rate differentials that arises from regression of the unannualised κ-period *ex post* ERRP $\lambda_{t,\kappa}$ on interest rate differentials and relative yield curve factors. 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 4: Estimated relative slope coefficients from augmented UIP regression



Notes: Black circles denote $\hat{\beta}_{2,\kappa}$ point estimates from regression (9). The horizontal axis denotes the horizon κ in months. In regression (9), the slope and curvature in each country are measured using proxies. Regressions are estimated using pooled monthly data from 1980:01 to 2017:12, including country fixed effects. 95% confidence intervals, calculated using Driscoll and Kraay (1998) standard errors, are denoted by thick black bars around point estimates.

Predictability of interest rates The inclusion of interest rate differentials in our preferred specification (9) poses a potential challenge, as interest rates are persistent and have a factor structure that, in turn, is a function of the yield curve slope (Litterman and Scheinkman, 1991). To ensure that the relationship between the slope and ERRP is not driven by the predictability of interest rates, we also estimate a simple regression of exchange rate changes on the relative slope and curvature, omitting interest rate differentials. The results are presented in table 9 of Appendix B.2, alongside a specification where we include the relative yield curve level, slope and curvature as in Chen and Tsang (2013). Qualitatively, we continue to find a significant tent-shaped relationship between exchange rate changes and the relative slope across horizons, with a similarly timed peak.

Long-horizon inference In long-horizon variants of (6) and (9), the number of non-overlapping observations can be limited and so size distortions—i.e. the null hypothesis being rejected too often—are a pertinent concern, especially with small samples and persistent regressors (Valkanov, 2003). To carry out more conservative inference than implied by Driscoll and Kraay (1998) standard errors, we draw on Moon, Rubia, and Valkanov (2004), who propose the scaling of t-statistics by $1/\sqrt{\kappa}$, showing that these scaled t-statistics are *approximately* standard normal when regressors are highly persistent.[11] Using the more conservative scaled t-statistics, our primary result remains significant, as figure 6 in Appendix B.2 demonstrates.

Sub-sample stability Our main results are robust to splitting the sample into two sub-periods. First, a pre-global financial crisis period, spanning 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, spanning 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). In both, presented in table 10 of Appendix B.2, the coefficients on the relative slope remain significant and tent-shaped with respect to maturity.

Country-specific regressions Table 11 of Appendix B.2 presents country-specific estimates of regression (9), using Newey and West (1987) standard errors. The tent-shaped pattern for the loading on the relative slope broadly holds at a currency level *vis-à-vis* the USD.

3.2 The Role of the Yield Curve

Our results so far imply that the relative yield curve slope plays an economically and statistically significant role in explaining future exchange rate movements at business cycle horizons. In this sub-section, we tie the yield curve slope to transitory SDF innovations which drive ERRP.

Yield Curve Slopes Our first key building block for this interpretation is the relationship between the yield curve slope and the autocovariance of SDFs. For a Home agent investing in an n-period Home bond, the relevant Euler equation (1) can be rewritten as:

$$\frac{1}{R_{t,n}} = \mathbb{E}_t \left[\prod_{i=0}^{n-1} M_{t+i,t+i+1} \right] \tag{11}$$

Defining the (log) excess return from buying an n-period 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 $rx_{t+1,n} = p_{t+1,n-1} - p_{t,n} - y_{t,1}$, 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 an n-period bond,[12] then Piazzesi and Schneider (2007) show that

$$\mathbb{E}_t \left[rx_{t+1,n} \right] = -\text{cov}_t \left(m_{t,t+1}, \mathbb{E}_{t+1} \sum_{i=1}^{n-1} m_{t+i,t+i+1} \right) - \frac{1}{2} \text{var}_t \left(p_{t+1,n-1} \right) \tag{12}$$

Here, the covariance term on the right-hand side is the risk premium on bonds. It implies that the risk premium on an n-period domestic bond is given by the covariance of today's one-period SDF with the sum of all future one-period SDFs from $t+1$ to $t+n$. The risk

[11] 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 the most conservative inference for our regressions.

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

premium is positive if today's one-period SDF is *negatively correlated* with expected changes in future marginal utility. That is, if households receive good news about the distant future and, as a consequence, value consumption less at this horizon—i.e. lower $\mathbb{E}_t\left[m_{t+i,t+i+1}\right]$ for some $i > 0$—they will value consumption relatively highly in the near-term—i.e. higher $m_{t,t+1}$.

Piazzesi and Schneider (2007) further note that, over long enough samples, the average excess return on an n-period bond is approximately equal to the average yield curve slope, defined as the spread between the n-period yield and the short rate, $S_t \equiv y_{t,n} - y_{t,1}$, so:[13]

$$S_t \approx -\text{cov}_t\left(m_{t,t+1}, \mathbb{E}_{t+1}\sum_{i=1}^{n-1} m_{t+i,t+i+1}\right) \tag{13}$$

where the right-hand-most Jensen's inequality term in (12) has been suppressed. As a result, the yield curve will be upward sloping on average if the right-hand side of (13) is positive. In turn, the fact yield curves slope upwards on average indicates that the covariance of today's one-period SDF with the sum of all future one-period SDFs from $t+1$ to $t+n$ is indeed negative.

ERRP and Transitory Risk Next, consider the following decomposition of pricing kernels V_t, proposed in Alvarez and Jermann (2005): $V_t = V_t^{\mathbb{T}} V_t^{\mathbb{P}}$, where $V_t^{\mathbb{T}}$ is a component with only transitory innovations and $V_t^{\mathbb{P}}$ is a component with permanent innovations which follows a martingale.[14] A variable X is defined as having only transitory innovations if $\lim_{\kappa\to\infty} \frac{\mathbb{E}_{t+1}[X_{t+\kappa}]}{\mathbb{E}_t[X_{t+\kappa}]} = 1$. We interpret transitory innovations to the pricing kernel as business cycle risk. Although Alvarez and Jermann (2005) show that most SDF volatility is attributable to permanent innovations using data on domestic yields, Lustig et al. (2019) show that the cross-country difference in permanent SDF volatility must be close to zero. In the limit, the ERRP is given by:

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

where $\nu_t^{(*)} \equiv \log(V_t^{\mathbb{P}(*)})$.[15]

To reconcile UIP deviations seen in the data with theory at both short and long horizons under (8), we require that high-yield currencies have relatively less volatile transitory pricing kernels $\text{var}_t(\nu_t^{\mathbb{T}(*)})$, while the volatilities of the permanent components are similar. Our main empirical exercise exploits the fact that transitory components of the SDFs, in contrast to permanent ones, are captured by the term structure of risk-free yields. We next present two

[13] To see this, re-write the excess return $rx_{t+1,n}$ as

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

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

[14] Formally $V_t^{\mathbb{T}} = \lim_{\kappa\to\infty} \frac{\delta^{t+\kappa}}{P_{t,\kappa}}$ where δ is a constant chosen to satisfy $0 < \lim_{\kappa\to\infty} \frac{P_{t,\kappa}}{\delta^\kappa} < \infty$ for all t. Given this $V_t^{\mathbb{P}} = \lim_{\kappa\to\infty} \frac{P_{t,\kappa}}{\delta^{t+\kappa}} V_t = \lim_{\kappa\to\infty} \frac{\mathbb{E}_t\left[V_{t+\kappa}\right]}{\delta^{t+\kappa}}$.

[15] Lustig et al. (2019) derive this as the conditional dollar term premium on an infinite-maturity bond, but in the limit the two risk premia are equivalent, as we discuss in Section 3.3.

analytical examples explicitly deriving the relationship between ERRP and the cross-country yield curve slope differential.

Example 1 (First-Order Autoregressive Pricing Kernel) *Let the (log) Home (Foreign) pricing kernel* $\nu_t^{(*)}$ *follow an AR(1) process* $\nu_t^{(*)} = \rho_\nu^{(*)}\nu_{t-1}^{(*)} + \varepsilon_{\nu,t}^{(*)}$*, where* $\varepsilon_{\nu,t}^{(*)} \sim \mathcal{N}\left(0, \sigma_\nu^{(*)}\right)$*,* $\rho_\nu = \rho_\nu^* \in (0,1)$ *and* $\sigma_\nu^{(*)} > 0$*. Under complete markets, the ERRP (8) can be written*

$$\lambda_{t,\kappa} = \frac{1}{2}\rho_\nu^{2(\kappa-1)}\left[\mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{var}_t\left(\nu_{t+1}^*\right)\right] \tag{15}$$

and, using (13), the relative yield curve slope can (approximately) be written as:

$$S_t^R \equiv S_t - S_t^* = \left(1 - \rho_\nu^{(n-1)}\right)\left[\mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{var}_t\left(\nu_{t+1}^*\right)\right] \tag{16}$$

Combining (15) and (16) yields the following relationship between the empirical ERRP and the relative yield curve slope:

$$\lambda_{t,\kappa} = \frac{1}{2}\frac{\rho_\nu^{2(\kappa-1)}}{1-\rho_\nu^{(n-1)}}S_t^R. \tag{17}$$

Subject to the AR(1) structure, $\rho_\nu \in (0,1)$ delivers a positive yield curve slope, a salient empirical feature. Figure 5 illustrates the relationship between ERRP $\lambda_{t,\kappa}$ and the relative slope S^R across horizons κ where $\rho_\nu = 0.9$. This parameterisation delivers a simulated first-order SDF autocorrelation of around -0.02, close to the lower bound identified in Chrétien (2012). In this case, and all instances where yield curves slope upward in this AR(1) example, there is a positive relationship between the ERRP and the relative slope, declining as κ increases.[16] In the case of $\rho_\nu^{(*)} = 1$, the SDF is a martingale and only has permanent innovations, $\nu_t = \nu_t^{\mathbb{P}}$ and the yield curve in each country is flat. In the case of $\rho_\nu^{(*)} = 0$, the pricing kernel is i.i.d. and contains no information on ERRP. Consequently, consistent with our reasoning, the explanatory power of the relative yield curve slope on the ERRP originates from predictable transitory innovations to pricing kernels which reflect business cycle risk.

Example 2 (Second-Order Autoregressive Pricing Kernel) *Let the (log) Home (Foreign) pricing kernel* $\nu_t^{(*)}$ *follow an AR(2) process* $\nu_t^{(*)} = \rho_{1,\nu}^{(*)}\nu_{t-1}^{(*)} + \rho_{2,\nu}^{(*)}\nu_{t-2}^{(*)} + \varepsilon_{\nu,t}^{(*)}$*, where* $\varepsilon_{\nu,t}^{(*)} \sim \mathcal{N}\left(0, \sigma_\nu^{(*)}\right)$*,* $\rho_{1,\nu} = \rho_{1,\nu}^*$*,* $\rho_{2,\nu} = \rho_{2,\nu}^*$ *and* $\sigma_\nu^{(*)} > 0$*. Let* ψ_i *denote the coefficients that result from the conversion of an AR(2) process to an MA(*∞*) using the Wold decomposition theorem. Under complete markets, the ERRP (8) can be written*

$$\lambda_{t,\kappa} = \frac{1}{2}\psi_{\kappa-1}^2\left[\mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{var}_t\left(\nu_{t+1}^*\right)\right] \tag{18}$$

and, using (13), the relative yield curve slope can (approximately) be written as:

$$S_t^R \equiv S_t - S_t^* = \left(1 - \psi_{n-1}\right)\left[\mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{var}_t\left(\nu_{t+1}^*\right)\right] \tag{19}$$

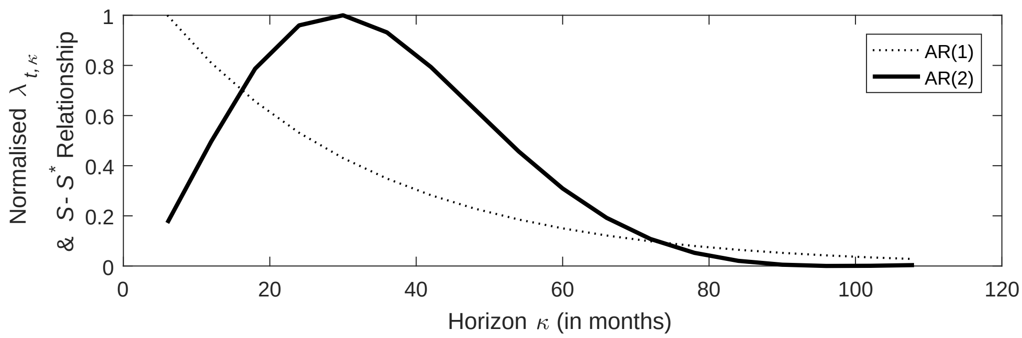
[16] See Appendix C.2 for a full derivation of Example 1.

Combining (18) *and* (19) *yields the following relationship between the empirical ERRP and the relative yield curve slope:*

$$\lambda_{t,\kappa} = \frac{1}{2} \frac{\psi_{\kappa-1}^2}{1 - \psi_{n-1}} S_t^R. \tag{20}$$

This example is able to capture both the positive sign of the relationship between ERRP and the relative slope, and the tent-shaped relationship with respect to horizon κ under certain restrictions on the parameters ψ_i.[17] To illustrate the tent-shaped relationship, figure 5 plots the relative size of coefficients linking the ERRP $\lambda_{t,\kappa}$ and the relative slope S_t^R across horizons in (17) and (20), respectively. We normalise both lines by the maximum factor across horizons, such that the peak normalised relationship is unity. In the AR(2) case, the strength of the relationship between the ERRP and the relative slope increases over short horizons, because with $\rho_{1,\nu} > 1$ innovations to the pricing kernel at time t have larger effects on subsequent pricing kernels than the contemporaneous one. Thereafter, as the horizon grows, a $\rho_{2,\nu} < 0$ ensures that the second-order term begins to dominate, ensuring the relationship approaches zero at long horizons.

Figure 5: Relationship between exchange rate risk premium $\lambda_{t,\kappa}$ and relative yield curve slope $S_t - S_t^*$ in examples 1 and 2



Notes: Plot of implied relationship between $\lambda_{t,\kappa}$ and S_t^R in examples 1 and 2, normalised such that the peak relationship is unity. For the AR(1) process (Example 1), the persistence parameter calibration is $\rho_\nu = 0.9$. For the AR(2) process (Example 2), the calibration is: $\rho_1 = 1.7$ and $\rho_2 = -0.75$.

Intuitively, if the yield curve of a country is upward sloping on average, the valuation of returns in the short run is high relative to the future.[18] The representative investor wants to reallocate returns to the short run, and this drives equilibrium ERRP determination. Examples 1 and 2 show analytically that the difference in the incentive to reallocate returns, as captured by the relative slope, is a driver of ERRP, consistent with our empirical findings.

[17] See Appendix C.3 for a full derivation of Example 2 and discussion of required parameter restrictions for a tent-shaped relationship.

[18] Piazzesi and Schneider (2007) argue that the upward sloping yield curve can also suggest bad news about future inflation, particularly in the early 1980s. However, the covariance between consumption growth and inflation driving this interpretation in insignificant in recent sub-samples. Additionally, yield curve on inflation-indexed securities also tend to slope upwards, on average.

3.3 Recasting the Slope: Holding Period Returns

In this sub-section, we assess returns on bonds of maturity κ over different holding periods h, extending the empirical specification in Lustig et al. (2019). Doing so allows us to isolate the contribution of the relative yield curve slope to ERRP and local-currency bond premia, sharpening our results. This analysis also has a secondary benefit, providing additional evidence to support the robustness of our results in Section 3.1, by helping to reduce the challenges posed by there being a limited number of non-overlapping observations as bond maturity κ increases in regression (9).[19]

Additional Notation 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 that holding period h is thus

$$rx^{(\kappa)}_{t,t+h} = \log\left[\frac{HPR^{(\kappa)}_{t,t+h}}{R_{t,h}}\right] \tag{21}$$

where, from Section 2.1, $R_{t,h}$ is the gross return on an h-month zero-coupon bond at t, i.e. the risk-free rate.

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

$$rx^{(\kappa),\$}_{t,t+h} = \log\left[\frac{HPR^{(\kappa)}_{t,t+h}}{R^*_{t,h}}\frac{\mathcal{E}_t}{\mathcal{E}_{t+h}}\right] = \log\left[\frac{HPR^{(\kappa)}_{t,t+h}}{R_{t,h}}\right] + \log\left[\frac{R_{t,h}}{R^*_{t,h}}\frac{\mathcal{E}_t}{\mathcal{E}_{t+h}}\right] \equiv rx^{(\kappa)}_{t,t+h} + rx^{FX}_{t,t+h} \tag{22}$$

where the rx^{FX}_{t+h} following the last equality represents the (log) currency excess return. In addition, the limiting relationship between currency excess returns and the per period return on infinite-maturity bonds, as shown in Lustig et al. (2019), is:

$$\lim_{\kappa\to\infty} -\frac{1}{\kappa}\mathbb{E}_t[rx^{FX}_{t,t+\kappa}] = \lim_{\kappa\to\infty} \frac{1}{\kappa}\mathbb{E}_t[\lambda_{t,\kappa}] = \lambda^{(\infty)}_{t,1} \tag{23}$$

Empirical Setup Using the above definitions, we estimate the following panel regressions for different holding periods h and bond maturities κ:

$$\mathrm{y}_{j,t,h} = \gamma_{1,h}\,(S_{j,t} - S^*_t) + f_{j,h} + \varepsilon_{j,t+h} \tag{24}$$

where the dependent variable $\mathrm{y}_{j,t,h}$ is either the excess return on the Home bond in USD 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

[19] Given that our dependent variable remains an *ex post* return, non-overlapping observations are not completely removed. But the share of non-overlapping observations in each sample does increase, even when assessing 5-year holding period returns on 10-year maturity bonds.

Home currency $rx^{FX}_{j,t,t+h}$, or the excess return on the Home bond in Home 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).

Compared to the results in Section 3.1, the $\gamma_{1,h}$ coefficients have a slightly different interpretation to $\beta_{2,\kappa}$. $\beta_{2,\kappa}$ can be interpreted as the average increase in the ERRP for Home investors associated with a 1pp increase in the slope of the Home yield curve relative to the US (base) country. Given that $rx^{FX}_{t,t+h}$ and $\lambda_{t,h}$ are inversely related up to some annualisation, then the interpretation of $\gamma_{1,h}$ is reversed: $-\gamma_{1,h}$ can be interpreted in a similar way to $\beta_{2,\kappa}$ when $y_{j,t,h} = rx^{FX}_{k,t,t+h}$, albeit in units of annual excess returns.

Focusing on $h = 1$ and $\kappa = 120$ only, 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 on 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 3.1 and those of Lustig et al. (2019).

Results The results are presented in table 2 and 3. Importantly, where our regression specification most closely matches Lustig et al. (2019), at short-holding periods $h = 6$ and long maturity $\kappa = 120$, our results mirror theirs.[20] The slope exerts an insignificant effect on the dollar-bond risk premium difference, a positive and significant influence on the local currency-bond risk premium difference $rx^{(120)}_{j,t,t+6} - rx^{(120)}_{US,t,t+6}$, and a negative and significant influence on the currency risk premium $rx^{FX}_{j,t,t+6}$, with the latter two effects similar in magnitude such that they cancel out for $rx^{(120),\$}_{j,t,t+6} - rx^{(120)}_{US,t,t+6}$.[21]

Exploring our results at all holding periods h and for all maturities κ, three observations are noteworthy.[22] First, for a given maturity, the loading on the relative slope exhibits a tent shape across holding periods for both the currency risk premium and the relative dollar-bond risk premium. Although significant at shorter holding periods, the relative slope loadings are quantitatively small for relative local currency-bond risk premia and are dominated by its loadings on currency excess returns in explaining the relative slope's impact on relative dollar-bond risk premia. This supports the findings from our benchmark augmented UIP regression in Section 3.1. Furthermore, the relative slope exerts its peak influence on dollar-bond and currency excess returns at the 36-month holding period, close to the 42-month horizon its influence peaks in the augmented UIP regression in Section 3.1.

Second, and related to the first, while the relative yield curve slope does not significantly predict dollar-bond excess return differences at the 6-month holding period for 10-year bonds, the relative slope loading for the same bond maturity is significantly non-zero over longer holding

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

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

[22]In Appendix B.3, we present average returns across maturities κ and holding periods h from dynamic investment strategies that involve going long the Home bond and short the US bond when the Home yield curve slope is lower than the US yield curve slope, and *vice versa*.

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 A: Dependent Variable: $rx^{(\kappa),\$}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$, Coefficient on $S - S^{*}$										
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 3: 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.

periods. While, in the former case, the influence of the relative slope on currency and local-currency bond returns offset one another (in line with Lustig et al., 2019), our results indicate that the influence of the relative slope on the currency premium dominates over longer holding periods (with a tent shape), even for long-term bonds. Nonetheless, for a given holding period, the influence of the relative slope on dollar-bond returns decreases in magnitude with maturity.

Third, for a given holding period, the loading on the relative slope for relative dollar bond returns is similar across bond maturities. This indicates that, insofar as the relative yield curve slope reflects ERRP, its influence is strongest at a 3 to 4-year horizon, supporting our interpretation the relative slope as an indicator of business cycle risk.

4 Accounting for Liquidity Yields

In this section, we extend our empirical specification to account for liquidity yields. We demonstrate that the tent-shaped relationship between relative slope and ERRP across horizons is robust to this extension and analyse the influence of the whole term structure of liquidity yields on ERRP.

4.1 Liquidity Yield-Augmented Regression

We use data on the term structure of liquidity yields from Du et al. (2018).[23] 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 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 on the excess return regressions in Section 3.3, although we present extended UIP regressions in Appendix B.4. Our benchmark regression is therefore:

$$\mathrm{y}_{j,t,h} = \gamma_{1,h}\left(S_{j,t} - S^*_t\right) + \gamma_{2,h}\eta^R_{j,t,\kappa} + f_{j,h} + \varepsilon_{j,t+h} \tag{25}$$

where the dependent variable $\mathrm{y}_{j,t,h}$ is either the relative dollar-bond return, the currency excess return, or the relative local currency-bond return as in (24). Here, the interpretation of $\gamma_{1,h}$ is unchanged relative to Section 3.3. The $\gamma_{2,h}$ estimate can be interpreted as the average influence of 1pp increase in relative US Treasury convenience. When the currency excess return $rx^{FX}_{t,t+h}$

[23] 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.

is the dependent variable, we expect $\gamma_{2,h}$ to be positive. An increase in relative US Treasury liquidity is associated with a contemporaneous appreciation of the USD (depreciation of Home currency) that lowers the ERRP $\lambda_{t,\kappa}$ (increases 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 2. As before, the slope loading is insignificant for excess returns over short and long holding periods for long-term bonds, consistent with evidence of UIP holding 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.

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

Tables 5 and 6 decompose these findings into the influence on ERRP and local currency-bond excess returns, respectively. As in Section 3.3, a comparison of the two tables indicates that the influence of both of relative slope and relative liquidity yields on dollar-bond excess

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.

returns predominantly works through currency excess returns. In contrast, the $\gamma_{2,h}$ loadings for local currency-bond excess returns are negative and relatively small in magnitude. As we show in the next sub-section, this suggests that the liquidity yields may contain some information about permanent SDF variations, extending the results in Lustig et al. (2019).

4.2 The Role of Liquidity Yields

To interpret our results through the lens of our framework, we draw a link between liquidity yields, as in Jiang et al. (2018), and non-traded risk, as in Lustig and Verdelhan (2019). We highlight a key difference between liquidity yields and relative yield curve slopes in their contribution to ERRP. Tying the theory with our empirical results, we show that liquidity yields capture permanent innovations to SDFs and influence long-horizon (cross-country) exchange rate differences, while relative yield curve slopes reflect business cycle risks captured in transitory SDF innovations.

Consider, Home and Foreign (US) representative investors taking positions in Home and Foreign bonds. From their perspectives, the Euler equations for their, respective, κ-period

Table 6: 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.

domestic bonds are (1) and (2). To allow for incomplete markets, we follow Backus et al. (2001) and introduce a wedge $\eta_{t+\kappa}$, as in (8), between (log) exchange rate changes and SDFs $\Delta^{\kappa} e_{t+\kappa} = \eta_{t+\kappa} + m^{*}_{t,t+\kappa} - m_{t,t+\kappa}$. Generally under incomplete markets, given the Home SDF $M_{t,t+\kappa}$ and the exchange rate process $\mathcal{E}_{t+\kappa}/\mathcal{E}_t$, the Foreign SDF $M^{*}_{t,t+\kappa}$ is not uniquely determined. Under the conditions laid out in Lustig and Verdelhan (2019, Proposition 1), a unique SDF exists in the space of traded assets such that each wedge $\eta_{t+\kappa}$ defines an SDF:

$$\frac{\mathcal{E}_{t+\kappa}}{\mathcal{E}_t} M_{t,t+\kappa} = M^{*}_{t,t+\kappa} e^{\eta_{t+\kappa}} \tag{26}$$

In our framework, the wedge $\eta_{t+\kappa}$ reflects variation in the cross-country difference between investor-specific liquidity needs, over and above what is captured by domestic yields. Using (26), consider both the representative Home investor's Euler equation for κ-period Foreign assets and the Foreign investor's Euler equation for Home assets, respectively:

$$\mathbb{E}_t\left[M_{t,t+\kappa}\frac{\mathcal{E}_{t+\kappa}}{\mathcal{E}_t}R^*_{t,\kappa}\right] = \mathbb{E}_t\left[M^*_{t,t+\kappa}e^{\eta_{t+\kappa}}R^*_{t,\kappa}\right] = 1 \tag{27}$$

$$\mathbb{E}_t\left[M^*_{t,t+\kappa}\frac{\mathcal{E}_t}{\mathcal{E}_{t+\kappa}}R_{t,\kappa}\right] = \mathbb{E}_t\left[M_{t,t+\kappa}e^{-\eta_{t+\kappa}}R_{t,\kappa}\right] = 1 \tag{28}$$

If $\eta_{t+\kappa} > 0$, the representative Home investor derives relatively more liquidity from Foreign assets than the Foreign investor, for any given valuation of domestic liquidity. Or conversely, the Foreign investor derives relative less liquidity from Home assets than the Home investor. In Appendix C.4 we extend Jiang et al. (2018) by explicitly modelling market incompleteness as in Lustig and Verdelhan (2019). Differentiating between investor-specific liquidity needs and asset-specific liquidity yields, we show the market incompleteness result from differences in investor-specific liquidity across countries.

Combining the four Euler equations (1), (2), (27) and (28), delivers the ERRP under incomplete markets as in (8). Here the wedge $\eta_{t+\kappa}$ reflects non-traded liquidity risk. The mechanism discussed in Section 2.3 generalises to the case with incomplete markets due to liquidity considerations. Consider the case where the Home economy is liquidity-constrained and values liquidity relatively more than Foreign households, $\eta_{t+\kappa} > 0$, for given domestic yields. Home investors are willing to forego pecuniary returns in favour of liquidity and, under no-arbitrage, ERRP result in an expected depreciation of Foreign currency. In levels, this reasoning is consistent with Jiang et al. (2018) and Engel and Wu (2018) and delivers a contemporaneous appreciation of highly liquid currencies.

Given this, we derive an analytical relationship between liquidity yields on long-term bonds and permanent SDF innovations. We take the limit of the expectation of (8), define $\eta_{t,\infty} = \lim_{\kappa\to\infty}\mathbb{E}_t[\eta_{t+\kappa}]$ to derive:

$$\lambda^{(\infty)}_{t,1} = \frac{1}{2}[\text{var}_t(\nu^{\mathbb{P}}_{t+1}) - \text{var}_t(\nu^{\mathbb{P}*}_{t+1})] + \eta_{t,\infty}$$

This is an equation in three unknowns, the ERRP $\lambda^{(\infty)}_{t,1}$, differences in permanent SDPs and the wedge $\eta_{t,\infty}$. We observe proxies for two of these. First, $\lambda^{(\infty)}_{t,1}$ is approximated by 6-month holding period returns from 10-year maturity bonds. Second, the measure of relative liquidity yields from Du et al. (2018) confounds traded and non-traded liquidity. However, from (8), we know that this measure predicts ERRP at longer-horizons from panel B.ii of table 5, so the non-traded component must be non-zero. With $\lambda^{(\infty)}_{t,1}$ tending to zero as in Lustig et al. (2019), then our evidence implies the following link between non-traded liquidity and cross-country differences in the variation of the permanent component of pricing kernels:

$$\frac{1}{2}[\text{var}_t(\nu^{\mathbb{P}}_{t+1}) - \text{var}_t(\nu^{\mathbb{P}*}_{t+1})] = -\eta_{t,\infty}$$

Persistent differences in investor-specific liquidity needs across countries are reflected in permanent SDF innovations and contribute to cross-sectional differences across currencies.

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 experience a depreciation in excess of UIP at business cycle horizons—3 to 5 years, especially. We attain this result using both long-horizon panel UIP regressions and regressions of excess returns over varying holding periods.

Theoretically, we derive the relationship between ERRP and the relative yield curve slope, and show this is driven by cross-country differences in the volatilities of transitory SDF innovations. If the yield curve of a country is relatively steep, the valuation of returns in the short run is relatively high. Investors would like to reallocate returns to the near term and, under no-arbitrage, a positive ERRP arises to deliver a depreciation in excess of UIP. Exchange rate dynamics compensate investors for business cycle risk.

Our findings are robust to the inclusion of liquidity yields which we show operate through a distinct channel. While the relative slope explains variation in ERRP arising to compensate investors for business cycle risk, liquidity yields contribute to cross-sectional differences across currencies and are associated with differences in the volatilities of permanent SDF innovations.

This paper has focused on the medium and long-run relationship between bond yields and ERRP. In related and ongoing work (Corsetti, Lloyd, and Marin, 2020), we investigate the short-run relationship around yield curve inversions, expanding our notion of business cycle risk to include time-varying disaster risk (e.g. Gourio, 2012; Gabaix, 2012).

Appendix

A Data Sources

We use nominal zero-coupon government bond yields at maturities from 6 months out to ten years for 7 industrialised countries: the United States, Australia, Canada, the euro area, Japan, Switzerland and the United Kingdom. Our benchmark sample begins in 1980:01 and ends in 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 7 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 to 2009:05 from Wright (2011).

Table 7: 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 Canonical UIP Regression

In Section 2, we document horizon-variation in the UIP condition, corroborating results in Chinn and Meredith (2005) and Chinn and Quayyum (2012). The results in figure 2 are derived from a panel regression of six currencies *vis-à-vis* the US dollar.

In table 8, we document that the broad upward sloping relationship between the UIP coefficient $\beta_{1,\kappa}$ and the horizon κ is robust when regressions are estimated on a country-by-country basis too. Column (1) of the table reprises the panel coefficient estimates of figure 2, with standard errors (reported in parentheses) constructed using the Driscoll and Kraay (1998) methodology. Columns (2)-(7) reports coefficient estimates for country-specific regressions. For each of these individual country regressions, we report Newey and West (1987) standard errors with five lags to account for serial correlation.

For all six currencies, short-horizon $\beta_{1,\kappa}$ coefficient estimates are negative out to, at least, the 24-month horizon. The coefficient rise with the horizon to be significantly above zero at longer tenors and, in most cases, close to unity.

B.2 Yield Curve-Augmented UIP Regression

In this Appendix, we present the results for the robustness exercises discussed in Section 3.1.

Predictability of interest rates Table 9 presents the results of regressions of exchange rate changes on (a) 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 (b), we proxy the yield curve level using the difference between 10-year zero-coupon yields $L_{j,t} = i_{j,t,10y}$. This specification replicates that in Chen and Tsang (2013). However, our results differ due to differences in the construction of yield curve factors. Chen and Tsang (2013) capture relative yield curve factors by directly estimating Nelson-Siegel decompositions on *relative* interest rate differentials. To do this, they assume symmetry of factor structures across countries. We, instead, construct proxies for factors using yield curves estimated on a country-by-country basis, and so do not assume such symmetry.

Long-horizon inference As discussed in Section 3.1, long-horizon forecasting regressions like (6) and (9) 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 when regressors are persistent. Although the Driscoll and Kraay (1998) standard errors used in the panel regressions in the main body of the paper are robust to heteroskedasticity and autocorrelation, we assess the robustness of our findings using alternative inference here.

Table 8: Coefficient estimates from canonical 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	-1.06	-0.75	-0.08	-1.28	-0.84	-1.57*	-1.23
	(0.65)	(1.92)	(0.58)	(0.92)	(0.92)	(0.83)	(1.13)
12-months	-0.99**	-1.44	-0.00	-1.26*	-0.64	-1.37**	-0.99
	(0.50)	(1.20)	(0.59)	(0.72)	(0.73)	(0.69)	(0.89)
18-months	-0.87**	-1.91**	-0.13	-1.08*	-0.46	-1.02	-0.92
	(0.43)	(0.79)	(0.60)	(0.60)	(0.69)	(0.67)	(0.72)
24-months	-0.67*	-1.62**	-0.08	-1.05**	-0.22	-0.68	-0.78
	(0.39)	(0.69)	(0.58)	(0.52)	(0.65)	(0.65)	(0.68)
30-months	-0.47	-1.29*	0.09	-1.19***	-0.10	-0.22	-0.64
	(0.35)	(0.68)	(0.56)	(0.44)	(0.64)	(0.62)	(0.64)
36-months	-0.25	-0.98	0.33	-1.26***	0.01	0.17	-0.23
	(0.33)	(0.72)	(0.54)	(0.38)	(0.62)	(0.61)	(0.55)
42-months	0.05	-0.34	0.54	-1.01**	0.23	0.49	0.07
	(0.33)	(0.76)	(0.51)	(0.41)	(0.57)	(0.57)	(0.56)
48-months	0.35	0.55	0.84*	-0.62	0.43	0.70	0.26
	(0.31)	(0.75)	(0.50)	(0.39)	(0.51)	(0.53)	(0.55)
54-months	0.67**	1.43**	1.06**	-0.25	0.66	0.89*	0.62
	(0.28)	(0.68)	(0.51)	(0.37)	(0.45)	(0.50)	(0.49)
60-months	0.90***	2.30***	1.18**	-0.02	0.87**	1.00**	0.75*
	(0.25)	(0.57)	(0.48)	(0.36)	(0.40)	(0.48)	(0.43)
66-months	1.11***	2.92***	1.43***	0.27	1.09***	1.05**	0.78**
	(0.23)	(0.46)	(0.42)	(0.34)	(0.37)	(0.46)	(0.38)
72-months	1.27***	3.20***	1.56***	0.51*	1.25***	1.04**	0.95***
	(0.19)	(0.39)	(0.36)	(0.29)	(0.33)	(0.44)	(0.32)
78-months	1.31***	3.07***	1.56***	0.69***	1.34***	0.92**	1.02***
	(0.17)	(0.37)	(0.36)	(0.25)	(0.30)	(0.41)	(0.31)
84-months	1.27***	2.88***	1.52***	0.75***	1.35***	0.81**	0.94***
	(0.17)	(0.34)	(0.38)	(0.22)	(0.29)	(0.39)	(0.27)
90-months	1.20***	2.63***	1.50***	0.74***	1.35***	0.72**	0.82***
	(0.17)	(0.31)	(0.41)	(0.26)	(0.28)	(0.36)	(0.25)
96-months	1.08***	2.15***	1.42***	0.57*	1.28***	0.69*	0.69***
	(0.17)	(0.33)	(0.43)	(0.31)	(0.26)	(0.36)	(0.24)
102-months	0.94***	1.74***	1.35***	0.38	1.15***	0.64*	0.54**
	(0.17)	(0.39)	(0.45)	(0.36)	(0.24)	(0.37)	(0.21)
108-months	0.81***	1.56***	1.25***	0.15	1.04***	0.59	0.40**
	(0.17)	(0.38)	(0.46)	(0.39)	(0.22)	(0.37)	(0.20)
114-months	0.73***	1.45***	1.15**	0.02	0.94***	0.68*	0.30
	(0.17)	(0.37)	(0.47)	(0.37)	(0.22)	(0.37)	(0.19)
120-months	0.68***	1.40***	1.23***	-0.12	0.85***	0.78**	0.17
	(0.16)	(0.34)	(0.47)	(0.34)	(0.21)	(0.35)	(0.20)

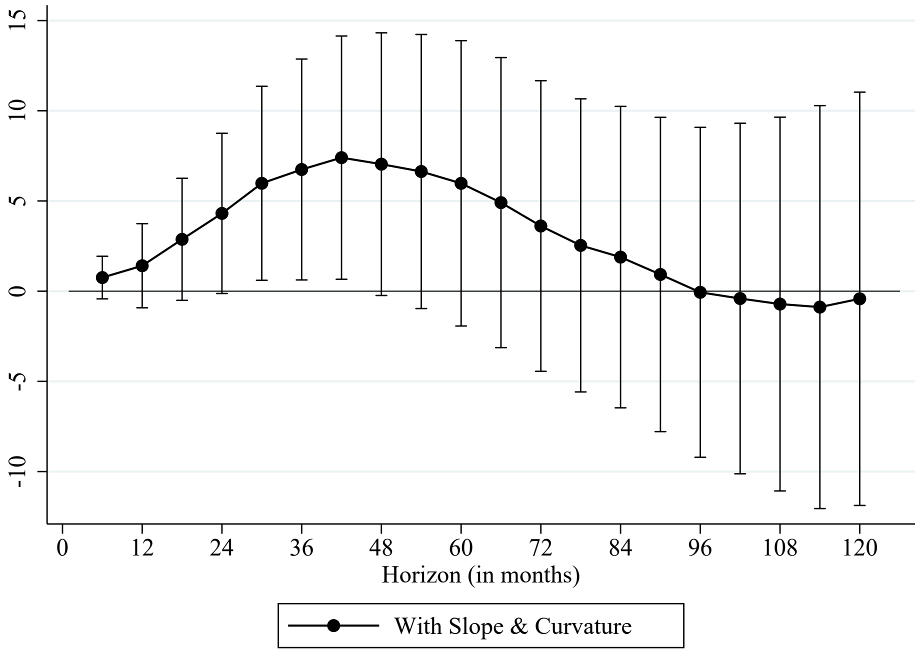
Notes: Coefficient estimates from regression (6)—the canonical UIP regression—a regression of the κ-period exchange rate change $\Delta^{\kappa} e_{t+\kappa}$ on the κ-period interest rate differential $i_{t,\kappa} - i^*_{t,\kappa}$. Regressions are estimated using pooled end-of-month data for six currencies—AUD, 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.

Table 9: Coefficient estimates from regressions of exchange rate change on relative slope and curvature, and relative level, slope and curvature and regression augmented with relative yield curve slope and curvature

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

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

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



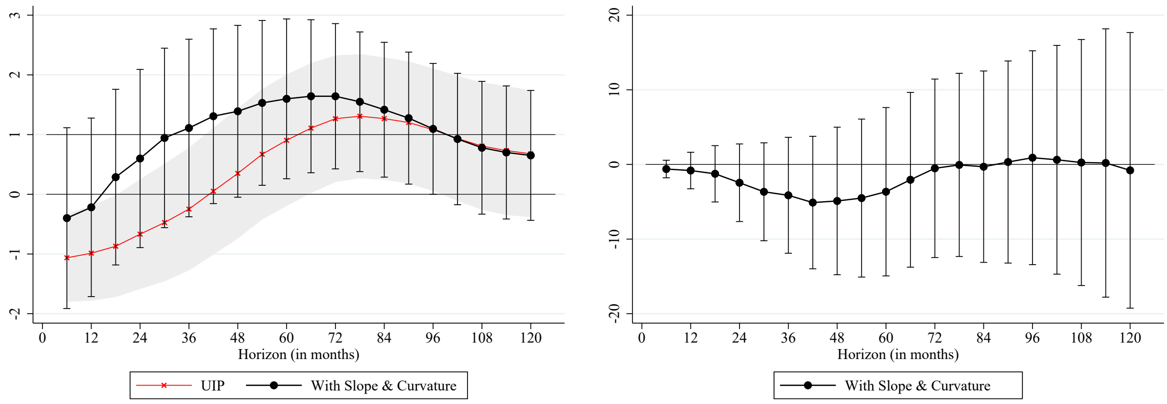
Notes: Black circles denote $\hat{\beta}_{2,\kappa}$ point estimates from regression (9). The horizontal axis denotes the horizon κ in months. In regression (9), the slope and curvature in each country are measured using proxies. Regressions are estimated using pooled monthly data from 1980:01 to 2017:12, 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.

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 (9) with 90% confidence intervals implied by the scaled t-statistics of Moon et al. (2004). Relative to table 1, 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 4-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 (9) 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.

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 (9). The horizontal axis denotes the horizon κ in months. In regression (9), 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.

Sub-sample stability To assess the stability of our results, we estimate regression (9) 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 10. For comparison, column (1) includes the relative slope coefficient loadings from our benchmark sample presented in the main body of the paper. Columns (3) and (4) include the estimated loadings over the pre-crisis and predominantly post-crisis samples, respectively. In both cases the coefficient estimates form a tent shape with respect to maturity, peaking at the 4 and 3.5-year horizons, respectively.

In addition, columns (2) and (5) present two additional robustness exercises. In column (2), we use available G10 currency and yield curve data, adding Sweden, Norway and New Zealand to our cross-section of countries, for the pre-crisis period only. In column (5), we drop the relative curvature from regression (9), to demonstrate that the relative slope coefficient is independent on the inclusion of the relative curvature. In both cases, the relative slope loadings continue to follow a tent-shaped pattern with respect to maturity.

Country-specific regressions Table 11 presents country-specific estimates of the yield curve augmented-UIP regression. Inference is conducted using Newey and West (1987) standard errors, to account for serial correlation. For comparison, column (1) presents the benchmark relative slope coefficient estimates from the panel regression discussed in the main body of the paper. As noted in the main text, although coefficient estimates vary in size and significance across countries, a relative slope coefficient estimates display a tent shape with respect to horizon

Table 10: Slope coefficient estimates from augmented UIP regression for pooled regression across different samples

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

Notes: Coefficient estimates on the relative yield curve slope $S_t - S_t^*$ from regression (9)—the augmented UIP regression—a regression of the κ-period exchange rate change $\Delta^\kappa e_{t+\kappa}$ on the κ-period interest rate differential $i_{t,\kappa} - i^*_{t,\kappa}$, 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.

Table 11: 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 (9)—the augmented UIP regression—a regression of the κ-period exchange rate change $\Delta^\kappa e_{t+\kappa}$ on the κ-period interest rate differential $i_{t,\kappa} - i^*_{t,\kappa}$, 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.

Table 12: 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 Home-country bond and short in the US bond when the Home yield curve slope is lower than the US yield curve slope, and go long in the US bond and short in the Home-country bond when the Home 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.

κ for 5 of the 6 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.

B.3 Dynamic Portfolio Returns

In table 12, we present the mean return from a simple investment strategy that goes long the Home bond and short the US bond when the Home yield curve slope is lower than the US yield curve slope, and goes long the US bond and short the Home 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).

At the $h = 6$ holding period and $\kappa = 120$ maturity, most closely corresponding to Lustig et al. (2019), the mean dollar-bond return difference is insignificantly different from zero due to offsetting currency and local currency bond returns. But, away from this point, table 12 demonstrates that dollar-bond return differences are non-zero and, for given hold periods, have a tent-shaped pattern across maturities, supporting evidence of the yield curve slope's predictive role for returns due to business cycle risk.

B.4 Liquidity Yield-Augmented Regressions

Here, we demonstrate that our results regarding liquidity yields, presented in Section 4.1 using excess return regressions, also hold true when extending the UIP regression. Using the definition

Table 13: Coefficient estimates from extended UIP regression, with relative yield curve slope and curvature and horizon-specific liquidity yield as additional regressors

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	UIP Regression		Yld. Curve-Augmented				Liq. Yld. & Yld. Curve-Augmented				
Mat.	$i_\kappa - i^*_\kappa$	$\overline{R}^2$	$i_\kappa - i^*_\kappa$	$S-S^*$	$C-C^*$	$\overline{R}^2$	$i_\kappa - i^*_\kappa$	$S-S^*$	$C-C^*$	η_κ	$\overline{R}^2$
12m	-0.78	0.012	1.61	4.35**	-4.11*	0.037	1.67	4.25**	-3.66*	-0.05	0.045
	(0.76)		(1.29)	(1.79)	(2.12)		(1.31)	(1.77)	(2.05)	(0.03)	
24m	-0.59	0.016	0.73	4.59**	-4.51**	0.039	0.59	3.63**	-2.89	-0.09**	0.057
	(0.53)		(0.75)	(1.82)	(2.18)		(0.77)	(1.74)	(2.12)	(0.04)	
36m	-0.13	0.011	1.49***	7.78***	-7.97***	0.066	1.33**	5.77***	-4.26**	-0.17***	0.123
	(0.40)		(0.57)	(1.85)	(2.24)		(0.61)	(1.64)	(2.00)	(0.04)	
60m	1.25***	0.097	2.30***	7.76***	-7.37***	0.155	2.06***	5.99***	-3.40	-0.20***	0.207
	(0.32)		(0.21)	(2.01)	(2.47)		(0.26)	(1.95)	(2.58)	(0.03)	
84m	1.35***	0.161	1.60***	2.45	1.17	0.185	1.32***	1.53	4.21	-0.23***	0.235
	(0.21)		(0.19)	(2.02)	(2.73)		(0.19)	(1.92)	(2.71)	(0.06)	
120m	0.59***	0.169	0.64***	2.81	-1.97	0.179	-0.04	1.19	1.64	-0.56***	0.420
	(0.22)		(0.22)	(2.27)	(3.21)		(0.19)	(1.60)	(2.56)	(0.07)	

Notes: Coefficient estimates and adjusted R^2 ($\overline{R}^2$) from regression (6) (UIP regression), (9) (yield curve-augmented UIP regression) and (29) (liquidity yield and yield curve-augmented UIP regression). Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD over a common sample (1991:04 to 2017:12), defined by the availability of liquidity yields. Regressions estimated using panel data for 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. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively.

of the κ-horizon liquidity premium for a κ-horizon US government bond relative to an equivalent-maturity government bond yield in country j, $\eta^R_{j,t,\kappa}$, we estimate a sequence of extended UIP regressions for each κ:

$$e_{j,t+\kappa} - e_{j,t} = \beta_{1,\kappa}\left(i_{j,t,\kappa} - i^*_{t,\kappa}\right) + \beta_{2,\kappa}(S_{j,t} - S^*_t) + \beta_{3,\kappa}(C_{j,t} - C^*_t) + \beta_{4,\kappa}\eta^R_{j,t,\kappa} + u_{j,t+\kappa} \quad (29)$$

The central hypothesis in Engel and Wu (2018) is that because liquidity is attractive to investors, an increase in a country's relative liquidity yield should *ceteris paribus* appreciate a currency today and, this, result in an expected depreciation in the future. Given the definition of $\eta^R_{j,t,\kappa}$ as the relative liquidity of US Treasuries *vis-à-vis* other countries, this implies a hypothesised $\beta_{4,\kappa} < 0$ in regression (29).

Table 13 presents results, comparing the liquidity yield *and* yield curve-augmented regression (6) with the baseline UIP regression (6) and the yield curve-augmented regression (9) over a common sample. The coefficient on the relative yield curve slope is robust to the additional inclusion of liquidity yields as a regressor in (29). This is seen by comparing columns (4) and (8) of table 13. The loading on the relative slope remains tent-shaped with respect to maturity, peaking here at the 5-year tenor and declining to insignificant values at the 7 and 10-year tenors.

The inclusion of the relative liquidity yield substantially improves the fit for exchange rates. At all horizons, the adjusted R^2 of the regression (29) exceeds that of (6) and (9). In addition, the largest increase in $\overline{R}^2$ from liquidity yields comes at the 10-year horizon and the coefficient estimates in column (10) support this. Our results suggest that this phenomenon especially powerful a medium to long horizons, with$\beta_{4,\kappa}$ estimates significantly negative from the 2-year horizon and beyond, growing in magnitude with respect to tenor. An increase in $\eta^R_{j,t,\kappa}$ represents higher perceived relative liquidity for US Treasuries, placing contemporaneous appreciation pressure on the dollar and *vice versa* for country-j currency.

C Additional Derivations

C.1 Derivation of Exchange Rate Risk Premia $\lambda_{t,\kappa}$

From (4), the κ-period *ex post* ERRP from the perspective of the Home agent is $\lambda^H_{t,\kappa} = -\text{cov}_t\left(m_{t,t+\kappa}, \Delta^\kappa e_{t+\kappa}\right)$.

For the Foreign agent, with SDF $M^*_{t,\kappa}$, an analogous cross-border no-arbitrage condition can be attained, satisfying

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

Assuming $\mathcal{E}_t$ and $M^*_{t,\kappa}$ are jointly log-normally distributed, international no-arbitrage requires

$$\mathbb{E}_t\left[\Delta^\kappa e_{t+\kappa}\right] - \frac{1}{2}\text{var}_t\left(-\Delta^\kappa e_{t+\kappa}\right) = \left(i_{t,\kappa} - i^*_{t,\kappa}\right) + \text{cov}_t\left(m^*_{t,t+\kappa}, -\Delta^\kappa e_{t+\kappa}\right)$$

So, from the representative Foreign agent's perspective, the κ-period *ex post* excess return from engaging in international asset trade is defined as $\lambda^F_{t,\kappa} = -\text{cov}_t\left(m^*_{t,t+\kappa}, -\Delta^\kappa e_{t+\kappa}\right)$.

Engel (2014) emphasises that standard empirical models do not measure $\lambda^H_{t,\kappa}$ or $\lambda^F_{t,\kappa}$, but instead provide more direct evidence on

$$\begin{aligned}
\lambda_{t,\kappa} \equiv \frac{\lambda^H_{t,\kappa} - \lambda^F_{t,\kappa}}{2} &= \frac{1}{2}\left[-\text{cov}_t\left(m_{t,t+\kappa}, \Delta^\kappa e_{t+\kappa}\right) + \text{cov}_t\left(m^*_{t,t+\kappa}, -\Delta^\kappa e_{t+\kappa}\right)\right] \\
&= \frac{1}{2}\left[-\text{cov}_t\left(m_{t,t+\kappa}, \Delta^\kappa e_{t+\kappa}\right) - \text{cov}_t\left(m^*_{t,t+\kappa}, \Delta^\kappa e_{t+\kappa}\right)\right] \\
&= -\frac{1}{2}\text{cov}_t\left(m_{t,t+\kappa} + m^*_{t,t+\kappa}, \Delta^\kappa e_{t+\kappa}\right) \\
&= -\text{cov}_t\left(\frac{m_{t,t+\kappa} + m^*_{t,t+\kappa}}{2}, \Delta^\kappa e_{t+\kappa}\right) \qquad \text{(C.1.1)}
\end{aligned}$$

replicating (5) in the main body.

C.2 Derivations for Example 1

In Example 1, we specify that the (log) pricing kernel of the Home (Foreign) agent $\nu^{(*)}_t \equiv \log V^{(*)}_t$, where $m^{(*)}_{t,t+\kappa} \equiv \nu^{(*)}_{t+\kappa} - \nu^{(*)}_t$, follows a mean-zero first-order autoregressive process, with persistence parameter $\rho^{(*)}_\nu \in (0,1)$:

$$\nu^{(*)}_t = \rho^{(*)}_\nu \nu^{(*)}_{t-1} + \varepsilon^{(*)}_{\nu,t}, \qquad \varepsilon^{(*)}_{\nu,t} \sim \mathcal{N}\left(0, \sigma^{(*)}_\nu\right) \qquad \text{(C.2.1)}$$

where $\sigma^{(*)}_\nu > 0$. Note that this pricing kernel is stationary, implying that $\nu^{(*)}_t$ contains some transitory component $\nu^{(*)^{\mathbb{T}}}_t \equiv \log V^{(*)^{\mathbb{T}}}_t$.

To derive our result—an analytical relationship between the ERRP $\lambda_{t,\kappa}$ and the relative cross-country yield curve slope S^R_t—we use two ingredients. First, by using the specific functional form for the (log) pricing kernel (C.2.1), the *ex post* κ-period ERRP under complete markets (8) can be written as

$$\begin{aligned}
\lambda_{t,\kappa} &= \frac{1}{2}\left[\mathrm{var}_t\left(m_{t,t+\kappa}\right) - \mathrm{var}_t\left(m^*_{t,t+\kappa}\right)\right] \\
&= \frac{1}{2}\left[\mathrm{var}_t\left(\nu_{t+\kappa} - \nu_t\right) - \mathrm{var}_t\left(\nu^*_{t+\kappa} - \nu^*_t\right)\right] \\
&= \frac{1}{2}\left[\mathrm{var}_t\left(\nu_{t+\kappa}\right) - \mathrm{var}_t\left(\nu^*_{t+\kappa}\right)\right] \\
&= \frac{1}{2}\left[\mathrm{var}_t\left(\rho_\nu^{(\kappa-1)}\nu_{t+1} + \sum_{i=0}^{\kappa-1}\rho_\nu^i\varepsilon_{\nu,t+\kappa-i}\right) - \mathrm{var}_t\left(\rho_\nu^{*(\kappa-1)}\nu^*_{t+1} + \sum_{i=0}^{\kappa-1}\rho_\nu^{*i}\varepsilon^*_{\nu,t+\kappa-i}\right)\right] \\
&= \frac{1}{2}\left[\rho_\nu^{2(\kappa-1)}\mathrm{var}_t\left(\nu_{t+1}\right) - \rho_\nu^{*2(\kappa-1)}\mathrm{var}_t\left(\nu^*_{t+1}\right)\right] \qquad \text{(C.2.2)}
\end{aligned}$$

where line 2 uses the definition of the log SDF and the log pricing kernels $m_{t,t+\kappa} \equiv \nu_{t+\kappa} - \nu_t$, line 3 conditions on information available at time t, line 4 uses a backward iteration of $\nu^{(*)}_{t+\kappa}$ in terms of $\nu^{(*)}_{t+1}$, and line 5 expands this by conditioning on information at time t. In addition, if $\rho_\nu = \rho^*_\nu$, then (C.2.2) can be rewritten as

$$\lambda_{t,\kappa} = \frac{1}{2}\rho_\nu^{2(\kappa-1)}\left[\mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{var}_t\left(\nu^*_{t+1}\right)\right] \qquad \text{(C.2.3)}$$

Second, the expression for the slope of a given yield curve (13), can be re-expressed given the AR(1) (log) pricing kernel (C.2.1). For the Home country the yield curve slope S_t can approximately be expressed as:

$$\begin{aligned}
S_t \approx \mathbb{E}_t\left[rx_{t+1,n}\right] &= -\mathrm{cov}_t\left(m_{t,t+1}, \mathbb{E}_{t+1}\sum_{i=1}^{n-1} m_{t+i,t+i+1}\right) \\
&= -\mathrm{cov}_t\left(\nu_{t+1} - \nu_t, \mathbb{E}_{t+1}\left[\nu_{t+n} - \nu_{t+1}\right]\right) \\
&= \mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{cov}_t\left(\nu_{t+1}, \mathbb{E}_{t+1}\left[\nu_{t+n}\right]\right) \\
&= \mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{cov}_t\left(\nu_{t+1}, \mathbb{E}_{t+1}\left[\rho_\nu^{(n-1)}\nu_{t+1} + \sum_{i=0}^{n-1}\rho_\nu^i\varepsilon_{\nu,t+n-i}\right]\right) \\
&= \left(1 - \rho_\nu^{(n-1)}\right)\mathrm{var}_t\left(\nu_{t+1}\right) \qquad \text{(C.2.4)}
\end{aligned}$$

where line 2 uses the definition of the log SDF and the log pricing kernels $m_{t,t+\kappa} \equiv \nu_{t+\kappa} - \nu_t$, line 3 conditions on information available at time $t+1$ to break-up the expectation and information available at time t to simplify the covariance, line 4 uses a backward iteration of ν_{t+n} in terms of ν_{t+1}, and line 5 expands this and simplifies the resulting expression.

An analogous expression to (C.2.4) can be derived for the Foreign representative investors, and together these yield the following expression for the relative yield curve slope S^R_t up to Jensen's inequality terms:

$$S^R_t \equiv S_t - S^*_t = \left(1 - \rho_\nu^{(n-1)}\right)\mathrm{var}_t\left(\nu_{t+1}\right) - \left(1 - \rho_\nu^{*(n-1)}\right)\mathrm{var}_t\left(\nu^*_{t+1}\right) \qquad \text{(C.2.5)}$$

which, when $\rho_\nu = \rho_\nu^*$, can be written as

$$S_t^R \equiv S_t - S_t^* = \left(1 - \rho_\nu^{(n-1)}\right)\left[\mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{var}_t\left(\nu_{t+1}^*\right)\right] \tag{C.2.6}$$

Comparing the expression for the *ex post* ERRP under symmetry (C.2.3) and the expression for the relative cross-country yield curve slope (C.2.6), the two have the following analytical relationship:

$$\lambda_{t,\kappa} = \frac{1}{2}\frac{\rho_\nu^{2(\kappa-1)}}{1-\rho_\nu^{(n-1)}} S_t^R \tag{C.2.7}$$

where, when $\rho_\nu \in (0,1)$ and $\kappa, n > 1$, $\rho_\nu^{2(\kappa-1)}/(1-\rho_\nu^{(n-1)}) > 0$, such that

$$\frac{\partial \lambda_{t,\kappa}}{\partial S_t^R} > 0$$

implying that a steeper Home yield curve is associated with a Home exchange rate depreciation over time and, thus, an increase in the *ex post* ERRP on Foreign currency. □

C.3 Derivations for Example 2

In Example 2, we specify that the (log) pricing kernel of the Home (Foreign) agent $\nu_t^{(*)}$ follows a mean-zero second-order autoregressive process:

$$\nu_t^{(*)} = \rho_{1,\nu}^{(*)}\nu_{t-1}^{(*)} + \rho_{2,\nu}^{(*)}\nu_{t-2}^{(*)} + \varepsilon_{\nu,t}^{(*)}, \qquad \varepsilon_{\nu,t}^{(*)} \sim \mathcal{N}\left(0, \sigma_\nu^{(*)}\right) \tag{C.3.1}$$

where $\sigma_\nu^{(*)} > 0$.

For simplicity, we impose that $\rho_{1,\nu} = \rho_{1,\nu}^{(*)}$ and $\rho_{2,\nu} = \rho_{2,\nu}^{(*)}$. Defining L as the lag operator, (C.3.1) can be rewritten as

$$\rho(L)\nu_t^{(*)} \equiv \left(1 - \rho_{1,\nu}L - \rho_{2,\nu}L^2\right)\nu_t^{(*)} = \varepsilon_{\nu,t}^{(*)} \tag{C.3.2}$$

which we define to be stationary, such that $\nu_t^{(*)}$ contains some transitory component—i.e. the roots of the characteristic equation, $C(x) = 1 - \rho_{1,\nu}x - \rho_{2,\nu}x^2 = 0$, lie outside of the unit circle.

Using the Wold decomposition theorem, (C.3.2) can be written as $\nu_t = \sum_{i=0}^{\infty}\psi_i\varepsilon_{\nu,t-i}^{(*)} \equiv \psi(L)\varepsilon_{\nu,t}^{(*)}$, where, for an AR(2) process, $\psi_1 = \rho_{1,\nu}$, $\psi_2 = \rho_{1,\nu}\psi_1 + \rho_{2,\nu}$ and $\psi_i = \rho_{1,\nu}\psi_{i-1} + \rho_{2,\nu}\psi_{i-2}$ for $i \geq 3$. For a given lag ℓ, then

$$\nu_t = \psi_\ell\nu_{t-\ell} + \sum_{i=0}^{\ell-1}\psi_i\varepsilon_{\nu,t-i} \tag{C.3.3}$$

Combining (C.3.3) with (8), the *ex post* κ-period ERRP under complete markets can be written as:

$$
\begin{aligned}
\lambda_{t,\kappa} &= \frac{1}{2}\left[\mathrm{var}_t\left(m_{t,t+\kappa}\right) - \mathrm{var}_t\left(m^*_{t,t+\kappa}\right)\right] \\
&= \frac{1}{2}\left[\mathrm{var}_t\left(\nu_{t+\kappa} - \nu_t\right) - \mathrm{var}_t\left(\nu^*_{t+\kappa} - \nu^*_t\right)\right] \\
&= \frac{1}{2}\left[\mathrm{var}_t\left(\nu_{t+\kappa}\right) - \mathrm{var}_t\left(\nu^*_{t+\kappa}\right)\right] \\
&= \frac{1}{2}\left[\mathrm{var}_t\left(\psi_{\kappa-1}\nu_{t+1} + \sum_{i=0}^{\kappa-2}\psi_i\varepsilon_{\nu,t+\kappa-i}\right) - \mathrm{var}_t\left(\psi_{\kappa-1}\nu^*_{t+1} + \sum_{i=0}^{\kappa-2}\psi_i\varepsilon^*_{\nu,t+\kappa-i}\right)\right] \\
&= \frac{1}{2}\psi^2_{\kappa-1}\left[\mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{var}_t\left(\nu^*_{t+1}\right)\right] \qquad \text{(C.3.4)}
\end{aligned}
$$

where line 2 uses the definition of the log SDF and the log pricing kernels $m_{t,t+\kappa} \equiv \nu_{t+\kappa} - \nu_t$, line 3 conditions on information available at time t, line 4 uses a backward iteration of $\nu^{(*)}_{t+\kappa}$ in terms of $\nu^{(*)}_{t+1}$, and line 5 expands this by conditioning on information at time t.

The expression for the slope of a given yield curve (13), can be re-expressed given the AR(2) (log) pricing kernel (C.3.3). For the Home country the yield curve slope S_t can approximately be expressed as:

$$
\begin{aligned}
S_t \approx \mathbb{E}_t\left[rx_{t+1,n}\right] &= -\mathrm{cov}_t\left(m_{t,t+1}, \mathbb{E}_{t+1}\sum_{i=1}^{n-1} m_{t+i,t+i+1}\right) \\
&= -\mathrm{cov}_t\left(\nu_{t+1} - \nu_t, \mathbb{E}_{t+1}\left[\nu_{t+n} - \nu_{t+1}\right]\right) \\
&= \mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{cov}_t\left(\nu_{t+1}, \mathbb{E}_{t+1}\left[\nu_{t+n}\right]\right) \\
&= \mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{cov}_t\left(\nu_{t+1}, \mathbb{E}_{t+1}\left[\psi_{n-1}\nu_{t+1} + \sum_{i=0}^{n-2}\psi_i\varepsilon_{\nu,t+n-i}\right]\right) \\
&= \left(1 - \psi_{n-1}\right)\mathrm{var}_t\left(\nu_{t+1}\right) \qquad \text{(C.3.5)}
\end{aligned}
$$

where line 2 uses the definition of the log SDF and the log pricing kernels $m_{t,t+\kappa} \equiv \nu_{t+\kappa} - \nu_t$, line 3 conditions on information available at time $t+1$ to break-up the expectation and information available at time t to simplify the covariance, line 4 uses a backward iteration of ν_{t+n} in terms of ν_{t+1}, and line 5 expands this and simplifies the resulting expression. The relative slope under symmetry, and up to a Jensen's inequality term, is therefore

$$
S^R_t \equiv S_t - S^*_t = \left(1 - \psi_{n-1}\right)\left[\mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{var}_t\left(\nu^*_{t+1}\right)\right] \qquad \text{(C.3.6)}
$$

Comparing the expression for the *ex post* ERRP under symmetry (C.3.4) and the relative cross-country yield curve slope (C.3.6), the two have the following analytical relationship:

$$
\lambda_{t,\kappa} = \frac{1}{2}\frac{\psi^2_{\kappa-1}}{1 - \psi_{n-1}} S^R_t \qquad \text{(C.3.7)}
$$

where, when $\psi_{n-1} \in (0,1)$ and $\kappa, n > 1$, $\psi^2_{\kappa-1}/(1-\psi_{n-1}) > 0$, such that

$$\frac{\partial \lambda_{t,\kappa}}{\partial S^R_t} > 0$$

implying that a steeper Home yield curve is associated with a Home exchange rate depreciation over time and, thus, an increase in the *ex post* ERRP on Foreign currency. □

C.4 Derivations for Section 4.2

In this Appendix we draw a link between liquidity yields and market incompleteness. In order to derive a relationship between cross-country differences in permanent innovations to SDFs and ERRP, we apply the approach in Backus et al. (2001) and Lustig and Verdelhan (2019) to study liquidity yields. This contrasts with Jiang et al. (2018), who note that liquidity yields cannot drive exchange rates in complete markets, but do not specify the form of market incompleteness.

Let representative Home (Foreign) investors derive a liquidity yield, denoted by $\xi^{H,H}_{t+\kappa}$ ($\xi^{F,F}_{t+\kappa}$), from holding their domestic κ-period bonds such that domestic Euler equations are:

$$1 = \mathbb{E}_t \left[M_{t,t+\kappa} R_{t,\kappa} e^{\xi^{H,H}_{t+\kappa}} \right] \tag{C.4.1}$$

$$1 = \mathbb{E}_t \left[M^*_{t,t+\kappa} R^*_{t,\kappa} e^{\xi^{F,F}_{t+\kappa}} \right] \tag{C.4.2}$$

Furthermore, suppose that liquidity yields confound an investor-specific liquidity need and an asset-specific liquidity yield. We denote investor-specific liquidity needs by $e^{\xi^{H,\cdot}_{t+\kappa}}$ and $e^{\xi^{F,\cdot}_{t+\kappa}}$ for the representative Home and Foreign investor, respectively. We denote asset-specific liquidity yields by $e^{\xi^{\cdot,H}_{t+\kappa}}$ and $e^{\xi^{\cdot,F}_{t+\kappa}}$ for the Home and Foreign κ-period asset, respectively. Therefore $\xi^{H,H}_{t+\kappa} = \xi^{H,\cdot}_{t+\kappa} + \xi^{\cdot,H}_{t+\kappa}$ and $\xi^{F,F}_{t+\kappa} = \xi^{F,\cdot}_{t+\kappa} + \xi^{\cdot,F}_{t+\kappa}$.

Additionally, let representative Home (Foreign) investors derive a liquidity yield, denoted by $\xi^{H,F}_{t+\kappa}$ ($\xi^{F,H}_{t+\kappa}$), from holding foreign κ-period bonds such that cross-country Euler equations for κ-period assets are, respectively:

$$1 = \mathbb{E}_t \left[M_{t,t+\kappa} \frac{\mathcal{E}_{t+\kappa}}{\mathcal{E}_t} R^*_{t,\kappa} e^{\xi^{H,F}_{t+\kappa}} \right] \tag{C.4.3}$$

$$1 = \mathbb{E}_t \left[M^*_{t,t+\kappa} \frac{\mathcal{E}_t}{\mathcal{E}_{t+\kappa}} R_{t,\kappa} e^{\xi^{F,H}_{t+\kappa}} \right] \tag{C.4.4}$$

Again, suppose that these liquidity yields are similarly decomposed into investor- and asset-specific components, such that $\xi^{H,F}_{t+\kappa} = \xi^{H,\cdot}_{t+\kappa} + \xi^{\cdot,F}_{t+\kappa}$ and $\xi^{F,H}_{t+\kappa} = \xi^{F,\cdot}_{t+\kappa} + \xi^{\cdot,H}_{t+\kappa}$.

An exchange rate process satisfying (C.4.2) and (C.4.3) is given, in logs, by $\Delta^\kappa e_{t+\kappa} = m^*_{t,t+\kappa} - m_{t,t+\kappa} + \xi^{F,F}_{t+\kappa} - \xi^{H,F}_{t+\kappa}$. We further require $\xi^{F,F}_{t+\kappa} - \xi^{H,F}_{t+\kappa} = \xi^{F,H}_{t+\kappa} - \xi^{H,H}_{t+\kappa}$ to also satisfy (C.4.1) and (C.4.4). Using the liquidity yield decompositions, then $\xi^{F,F}_{t+\kappa} - \xi^{H,F}_{t+\kappa} = \xi^{F,H}_{t+\kappa} - \xi^{H,H}_{t+\kappa} = \xi^{F,\cdot} - \xi^{H,\cdot}$, such that the exchange rate process can be mapped to (26), with $\eta_{t+\kappa} = (\xi^{F,\cdot}_{t+\kappa} - \xi^{H,\cdot}_{t+\kappa})$.[24] Non-traded risk reflects cross-country differences in investor-specific liquidity needs.

[24] Additionally allowing for investor-asset-specific liquidity does not alter our results.

In Section 4.2, we emphasise variation in non-traded risk over and above what influences domestic yields. This can be achieved when variation in investor-specific liquidity needs $\xi_{t+\kappa}^{H,\cdot}$ ($\xi_{t+\kappa}^{F,\cdot}$) does not influence $R_{t,\kappa}$ ($R_{t,\kappa}^{*}$). In the simplest case, using a log expansion of (C.4.1), Home investor-specific liquidity needs will not influence Home yields if they are offset by their covariance with Home asset-specific liquidity yields and the Home SDF:

$$\mathbb{E}_t\left[\xi_{t+\kappa}^{H,\cdot}\right] + \frac{1}{2}\mathrm{var}_t\left(\xi_{t+\kappa}^{H,\cdot}\right) = -\mathrm{cov}_t\left(\xi_{t+\kappa}^{H,\cdot}, \xi_{t+\kappa}^{\cdot,H}\right) - \mathrm{cov}_t\left(m_{t,t+\kappa}, \xi_{t+\kappa}^{H,\cdot}\right)$$

An analogous condition for $\xi_{t+\kappa}^{F,\cdot}$ is implied by a log expansion of (C.4.2). These conditions resemble those for η in Lustig and Verdelhan (2019, Proposition 1), except for the fact $\xi^{H,\cdot}$ and $\xi^{F,\cdot}$ are country-specific, while η is a cross-country term.

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