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EXCHANGE RATE RISK AND BUSINESS CYCLES

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Exchange Rate Risk and Business Cycles*

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Abstract

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JEL Codes: E43, F31.

Key Words: Exchange rates; Risk premia; Uncovered interest parity; Yield curves.

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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), giving rise to exchange rate risk premia (ERRP) that complicate forecasting the level and volatility of exchange rates. But the UIP hypothesis cannot be rejected at long horizons (e.g. Chinn and Meredith, 2005; Engel, 2016). These two pieces of evidence together set the stage for our analysis of the 'UIP puzzle'.

In this paper, we show 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 show this both by estimating UIP regressions across different maturities and by analysing excess returns over different holding periods. We interpret our findings through the lens of a standard no-arbitrage model and the decomposition of the stochastic discount factor (SDF) into transitory and permanent components proposed by Alvarez and Jermann (2005). The term structure of interest rates—specifically, the risk-free government bond yield curve—is not only a key component of a risk-neutral arbitrage relationship characterising exchange rate movements, but also prices transitory innovations to investors' SDFs. Our results suggest that exchange rate movements in excess UIP systematically reallocate returns intertemporally, to investors who value them most highly.

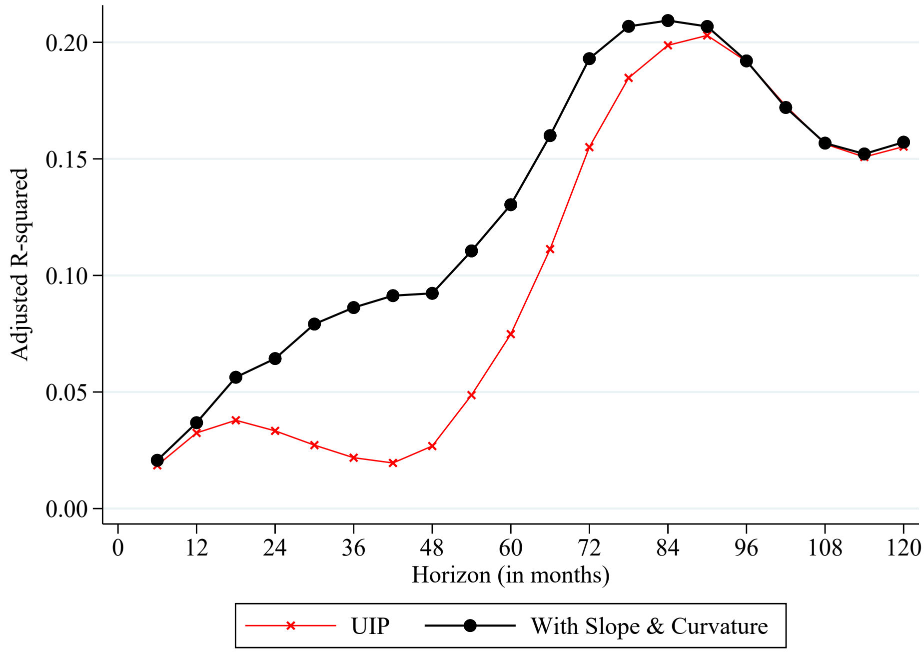
To motivate our analysis, Figure 1 plots the adjusted R^2 of a canonical panel UIP regression—a regression of the κ-period-ahead exchange rate change on κ-maturity cross-country yield differentials—together with the adjusted R^2 from the same regression extended to include measures of the cross-country relative yield curve slope and curvature. As shown in the figure, we find strong evidence that information in the yield curve can account for exchange rate fluctuations over and above the spot rate differential. The adjusted R^2 of our augmented regression is around treble that of the canonical UIP regression at business cycle horizons (3 to 4 years). 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]

Building on this, we consider a flexible empirical specification which allows bond holding periods and maturities to differ. We regress excess returns—for different maturity bonds over varying holding periods—on relative yield curve slopes. We find that the relative slope is a significant predictor of dollar-bond excess return differences, at holding periods associated with business cycle horizons (around 3 years), even with long-maturity bonds (up to 10 years).[2]

[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.

[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-period holding-period return regressions for 10-year bonds in Lustig et al. (2019).

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.

Nonetheless, the term structure of carry trade is decreasing with maturity at every holding period, consistent with the empirical findings in Lustig, Stathopoulos, and Verdelhan (2019).[3]

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 (building on the decomposition proposed in Alvarez and Jermann, 2005) in explaining ERRP, consistent with differences in the pricing of business cycle risk across countries, as 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. UIP failures can be understood with reference to risk-averse investors' current and future valuation of returns, summarised by the path of relative SDFs, and serve to reallocate returns intertemporally. For instance, an expected foreign exchange appreciation (domestic depreciation) increases the expected return from a foreign bond in domestic currency, while a subsequent depreciation lowers future expected returns. Empirically, we find that movements in the ERRP systematically reallocate

[3]The decreasing term structure of carry trade returns is rationalised by Gourinchas, Ray, and Vayanos (2019) in a framework with segmented markets and constrained arbitrageurs.

returns to investors with a relatively high valuation of returns, thus lowering the risk for home investors holding foreign assets. Since in periods where valuations of returns are high, yields on domestic risk-free bonds are low, this mechanism is consistent with the UIP puzzle.

We derive the relationship between ERRP and the relative yield curve slope. When yield curves are upward sloping on average, nearer-term return valuations—reflected by nearer-term SDFs—are high relative to longer-horizon valuations reflecting business cycle risk.[4] In our two-country setting, relative yield curve slopes influence exchange rate dynamics because they capture investors' *relative* desire to reallocate returns intertemporally—i.e. in response to transitory innovations to their SDF. ERRP, arising from an excess exchange rate depreciation, reallocate returns intertemporally to the country with a relatively high near-term valuation only in so far as it also faces a declining path of future valuations (reflected by a relatively steep slope). In this case, domestic investors are worse-off both relative to foreigners, and over time (in a business cycle sense). In contrast, permanent innovations to SDFs 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 relative yield curve slopes capture business cycle risk as a key determinant of exchange rate predictability. To gain insight on the relationship between ERRP and the relative yield curve slope, we construct two stylised examples, in which asset markets are complete and investors' pricing kernels follow independent mean-reverting first and second-order autoregressive processes, respectively. In both examples, the relationship is positive and in the latter example, we show that the relationship can have a tent shape with respect to the horizon, mirroring our empirical results.

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, through the lens of our theory, we show that, in contrast to business cycle risk, liquidity yields appear to reflect permanent innovations to SDFs. A key novelty of our work is to assess the differential influence of liquidity yields at different horizons. 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 is an important factor for understanding the cross-sectional dimension of UIP failures, business cycle risk which is our main focus, reflects the time-series dimension of UIP failures.

Finally, we unveil a novel empirical fact about the joint dynamics of ERRP and the yield curve consistent with the idea that business cycle risk is a driver of both. We find that while conventional short-horizon UIP failures arise when yield curves are upward sloping, the UIP coefficient switches sign when the foreign yield curve is inverted, but the sign of the relative

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

slope remains unchanged. Notably, this result holds whether or not we include the 2007-9 global financial crisis. This finding suggests that the 'New Fama Puzzle' highlighted by Bussière, Chinn, Ferrara, and Heipertz (2018)—that high-yield currencies experienced a depreciation in excess of UIP after the global financial crisis—is a recurrent phenomenon that can be related to yield curve inversions. Our findings are consistent with a model of rare disasters (Gabaix, 2012; Farhi and Gabaix, 2016).

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).[5] Building on this, 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,[6] highlighting the role of business cycle risk in explaining exchange rate dynamics.

Closely related to this paper, Lustig et al. (2019) investigate the term structure of carry trade returns. They show that whilst carry trade portfolios are profitable with short-maturity Treasuries, the returns to carry trade, for a fixed one-month holding period, are monotonically decreasing in the maturity of the asset as currency risk premia are offset by local-currency bond premia. Complementing this finding, we show that for given maturity bonds, the predictability of currency risk premia dominates as holding periods increase. Both our paper and Lustig et al. (2019) build on Alvarez and Jermann (2005), who propose a decomposition of the SDF into a permanent and a transitory component.

Our paper provides novel insights about the relevance of business cycle risk, priced into the yield curve, in explaining *time series* variation in ERRP. In a related paper, Colacito, Riddiough, and Sarno (2019) show that sorting currencies according to their output gap delivers positive excess portfolio returns, indicative of a role for business cycles in explaining *cross-sectional* variation in ERRP. Insofar as a high output gap contributes to a steeper yield curve slope, our finding that a relatively steep yield curve is associated with a contemporaneously appreciated currency (depreciating going forward) is consistent with the findings in Colacito et al. (2019) for excess portfolio returns. Furthermore, our paper highlights the differing implications of transitory (business cycle) and permanent SDF innovations for variation in ERRP.

Our empirical framework also relates closely to a largely atheoroetic literature linking the term structure of interest rates to exchange rates. Ang and Chen (2010), Chen and Tsang (2013) and Gräb and Kostka (2018) show that yield curve factors can significantly predict exchange rates, predominantly focusing on short horizons (less than 2 years). We show the role of the term structure in explaining medium-horizon ERRP fluctuations and, within our theoretical setup, directly attribute this to transitory SDF innovations.

[5] Engel's findings correspond to those in Chinn and Meredith (2005) if the expectations hypothesis of interest rates holds and interest rate differentials are characterised by a first-order Markov process.

[6] Hassan and Mano (2019) decompose the failure of the UIP into cross-currency, between-time-and-currency, and cross-time components. While they show that exchange rate predictability relies predominantly on the cross-time component, carry trade returns are driven by the cross-currency component and particularly permanent, or highly persistent, differences across currencies.

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) by investigating the relationship between the term structure of liquidity yields and exchange rates. We show that, unlike business cycle risk contained in bond premia, liquidity yields also explain permanent innovations to exchange rates and thus contribute to cross-sectional differences across currencies.

An emerging literature argues that representative-agent no-arbitrage models struggle to deliver the equilibrium pricing kernels required to reconcile exchange rate dynamics consistent with observed UIP deviations as well as price other assets in the economy. Greenwood, Hanson, Stein, and Sunderam (2019) and Gourinchas et al. (2019) present two-country general equilibrium models based on segmented markets which reconcile the joint dynamics of exchange rate and bond risk premia presented in Lustig et al. (2019).[7] Indeed, in sections 4 and 5 of our paper, we present additional results, relating to liquidity and yield curve inversions respectively, which largely provide support to this argument. However, in section 3 we show that, at least qualitatively, simple pricing kernels which arise naturally in representative agent economies are consistent with our main finding—the tent-shape relationship between relative yield curve slopes and ERRP.

Bussière et al. (2018) and Stavrakeva and Tang (2019), *inter alia*, document a reversal in the direction of UIP failures following the global financial crisis, with high yield currencies depreciating rather than appreciating. We show that reversals in the joint dynamics of exchange and interest rates are associated with yield curve inversions. In 5.2 we argue that prominent consumption-based asset pricing models that reconcile the UIP puzzle rely on either external habits as in Verdelhan (2010), building on Campbell and Cochrane (1999) or Epstein and Zin (1989) preferences and stochastic volatility as in Bansal and Shaliastovich (2013) struggle to jointly account for ERRP and excess bond premia. Our findings point towards an explanation based on currency crashes due to rare disasters (Gabaix, 2012; Farhi and Gabaix, 2016).

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. Section 5 assesses how exchange rate dynamics change around yield curve inversions, and section 6 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 through the lens of a preference-free setting.

[7] Notably, in a segmented markets explanation, the pricing kernel for bonds and exchange rates does not necessarily coincide with the pricing kernel on other assets—such as equity—allowing for the possibility of specific bond and foreign exchange dynamics.

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:

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

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

which, 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, imply:[8]

$$1 = \mathbb{E}_t \left[M_{t,t+\kappa} R_{t,\kappa} \right] \tag{3}$$

$$1 = \mathbb{E}_t \left[M^*_{t,t+\kappa} R^*_{t,\kappa} \right] \tag{4}$$

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:

$$1 = \mathbb{E}_t \left[M_{t,t+\kappa} \frac{\mathcal{E}_{t+\kappa}}{\mathcal{E}_t} R^*_{t,\kappa} \right] \tag{5}$$

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:[9]

$$\mathbb{E}_t \left[\Delta^\kappa e_{t+\kappa} \right] + \frac{1}{2} \mathrm{var}_t \left(\Delta^\kappa e_{t+\kappa} \right) = \left(i_{t,\kappa} - i^*_{t,\kappa} \right) - \mathrm{cov}_t \left(m_{t,t+\kappa}, \Delta^\kappa e_{t+\kappa} \right) \tag{6}$$

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 $m_{t,t+\kappa} \equiv \log(M_{t,t+\kappa})$. This expression indicates that expected κ-period exchange rate changes should be proportional to

[8] Throughout the paper we only consider nominal values, which are what we observe in the 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.

[9] 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.

κ-period interest differentials, corrected for the covariance between investors' SDF and exchange rate dynamics.[10] Were investors risk neutral and absent financial frictions, this covariance term would drop away and the expression predicts that returns on Home and Foreign assets should be equated in expectation, correcting for exchange rate valuation effects. In equilibrium, the ERRP is the average of the ERRP demanded by Home investors on Foreign bonds and Foreign investors on Home bonds.

2.2 Canonical UIP Regression

Motivated by (6), a large empirical literature has tested UIP by regressing *ex post* exchange rate changes on interest rate differentials. Following Chinn and Meredith (2005) and Chinn and Quayyum (2012), our benchmark empirical framework for testing UIP at different horizions builds on this canonical regression. Using panel data for a cross-section of countries j over time t, we estimate a sequence of regressions for each κ-month horizon under assumptions of risk neutrality and rational expectations:

$$e_{j,t+\kappa} - e_{j,t} = \beta_{1,\kappa}\left(i_{j,t,\kappa} - i^*_{t,\kappa}\right) + f_{j,\kappa} + u_{j,t+\kappa} \tag{7}$$

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 $\kappa > 0$.[11] Empirical rejections of UIP at short to medium horizons—i.e. finding $\hat{\beta}_{1,\kappa} \neq 1$ in regression (7) for small to medium κ—have regularly been used to motivate claims that interest rates do not adequately explain exchange rate dynamics and the ERRP.

Data To estimate our regressions, we use exchange rate and interest rate data for six jurisdictions with liquid bond markets: Australia (AU), Canada (CA), Switzerland (CH), the euro area (EA), Japan (JP) and the United Kingdom (UK). Additionally, the United States (US) 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), which we detail in Appendix A. Our nominal exchange rate data is from *Datastream*, measuring the value of domestic currency price per unit of US dollar.

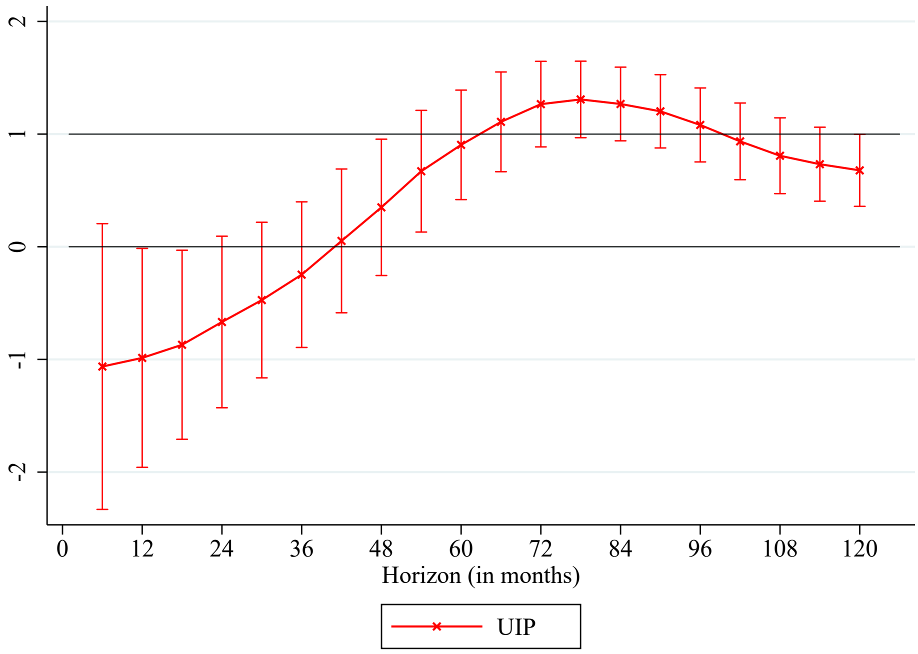
We explore exchange rate dynamics at horizons that match the available maturities of government bond yields, from 6 to 120 months. We use end-of-month data for the period 1980:01 to 2017:12.[12]

[10]The second left-hand side term, $\frac{1}{2}\text{var}_t(\Delta^\kappa e_{t+\kappa})$, is a Jensen's inequality term.

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

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

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



Notes: Red crosses denote $\hat{\beta}_{1,\kappa}$ estimates from regression (7). 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.

In the paper's main body, regressions are estimated using available data over this benchmark sample, unless otherwise stated. Importantly, however, 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 (e.g. Brunnermeier, Nagel, and Pedersen, 2009; Jordà and Taylor, 2012; Ca' Zorzi and Marin, 2018) and a switch in UIP coefficients (Bussière et al., 2018). This robustness is an important indicator of the pervasiveness of our results.

Results Figure 2 plots UIP coefficient $\beta_{1,\kappa}$ estimates—i.e. loadings on the κ-period interest rate differential $i_{t,\kappa} - i^*_{t,\kappa}$ in (7)—from a panel regression with country fixed effects over our benchmark sample.[13] 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, while the null hypothesis cannot be rejected at longer

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

horizons. The $\beta_{1,\kappa}$ point estimates are significantly below unity at short to medium horizons, consistent with the UIP puzzle. At the 6 to 36-month horizons, the point estimate is negative, indicating that high short-term interest rate currencies tend to appreciate, instead of depreciate, in line with Fama (1984). While, at the 42 and 48-month horizons point estimates are positive but significantly smaller than unity. Longer-horizon point estimates of $\beta_{1,\kappa}$ from regression (7) tend to be positive and close to unity, corroborating with, *inter alia*, Chinn and Meredith (2005) and Chinn and Quayyum (2012).[14]

2.3 Interpreting the UIP Puzzle

To interpret failures of UIP, we build on the preference-free environment from section 2.1. We first define the κ-period *ex post* ERRP at time t as:[15]

$$\lambda_{t,\kappa} \equiv i^*_{t,\kappa} - i_{t,\kappa} + \Delta^\kappa e_{t+\kappa} \tag{8}$$

By assuming that the Foreign agent undertakes a similar optimisation when purchasing Home assets to the Home investors purchasing Foreign—equation (5)—standard empirical methods provide evidence on the following *ex post* equilibrium ERRP (Engel, 2014):

$$\lambda_{t,\kappa} = -\text{cov}_t\left(\frac{m_{t,t+\kappa} + m^*_{t,t+\kappa}}{2}, \Delta^\kappa e_{t+\kappa}\right) \tag{9}$$

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.

Our point estimates, consistent with the literature, show that at short and medium horizons, where $\hat{\beta}_{1,\kappa} < 1$, high interest rate currencies tend to depreciate insufficiently (and sometimes appreciate) relative to the UIP benchmark, implying $\text{cov}_t(\lambda_{t,\kappa}, i_{t,\kappa} - i^*_{t,\kappa}) < 0$. 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. The opposite is true in periods where $i_{t,\kappa} < i^*_{t,\kappa}$.[16]

Cast in a general equilibrium framework, these dynamics imply that the ERRP acts to lower the risk for Home agents when holding Foreign bonds, since their effective return is relatively high when investors value returns most. A low Home interest rate $i_{t,\kappa} < i^*_{t,\kappa}$, via the domestic bond-pricing Euler equations (3) and (4), is associated with a relatively high valuation of returns in $t+\kappa$ by Home investors, $(\mathbb{E}_t[m_{t,t+\kappa}] > \mathbb{E}_t[m^*_{t,t+\kappa}])$.

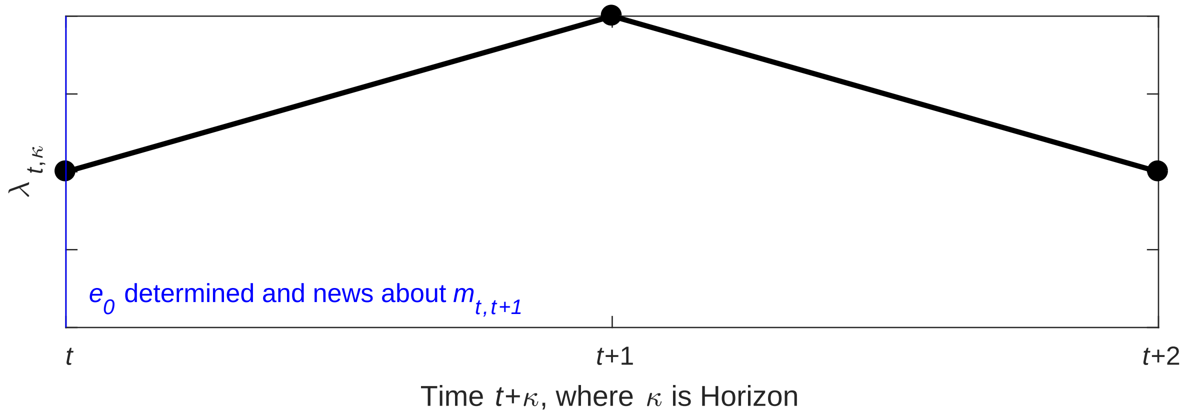
For illustration, suppose that, while $i_{t,2} = i^*_{t,2}$, some disturbance results in a drop of the Home short term interest rate $i_{t,1} < i^*_{t,1}$. As already mentioned, by the Euler equations (3) and (4), it must be the case that $\mathbb{E}_t[m_{t,t+1}] > \mathbb{E}_t[m^*_{t,t+1}]$, while $\mathbb{E}_t[m_{t,t+2}] = \mathbb{E}_t[m^*_{t,t+2}]$—note

[14]The broadly upward sloping relationship between $\beta_{1,\kappa}$ coefficient estimates and horizon κ is also apparent in country-specific regressions, reported in Appendix B.1.

[15]The *ex ante* ERRP is defined analogously, with the expectations operator: $\tilde{\lambda}_{t,\kappa} \equiv i^*_{t,\kappa} - i_{t,\kappa} + \mathbb{E}_t\left[\Delta^\kappa e_{t+\kappa}\right]$.

[16]To see this, note that (7) can be substituted into the definition of the ERRP (9) to yield $\lambda_{t,\kappa} = (\beta_{1,\kappa} - 1)(i_{t,\kappa} - i^*_{t,\kappa})$, where other terms in the regression have been suppressed for simplicity.

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 $i_{t,2} = i^*_{t,2}$



that this could arise from an anticipated transitory monetary policy or shock to consumption.[17] Under either scenario, Home investors will value one unit of Home currency at time $t+1$, more than Foreign investors value a unit of Foreign currency over the same horizon. Now, from the data, we know that $\text{cov}_t(\lambda_{t,\kappa}, i_{t,\kappa} - i^*_{t,\kappa}) < 0$: in our example, the ERRP implied by this observation first depreciates the domestic currency at $t+1$, then appreciates it from time $t+1$ to $t+2$—i.e. $\mathcal{E}_0 = \mathcal{E}_2 < \mathcal{E}_1$. The effect is illustrated in Figure 3, with a tent-like shape. From the vantage point of the Home investor, who holds both Home and Foreign bonds (both Euler equations hold), the depreciation increases the $t+1$ return from the Foreign bond in Home-currency terms, reallocating resources intertemporally—both from t to $t+1$ (over which horizon the Home currency is expected to depreciate), and from $t+2$ to $t+1$ (by virtue of the currency appreciation). In light of this interpretation of the empirical evidence, the residents in the low (high) yield country, who have a relatively high (low) nearer-term valuation of returns, can expect an excess depreciation (appreciation) relative to UIP, such that the ERRP reallocates returns to investors who value them the most.

However, the spot yield differential only captures part of the incentive for intertemporal reallocation of returns. Since the shock we consider is transitory, and as we show in Section 3.2 only transitory shocks are relevant for our mechanism, the equilibrium depreciation at $t+1$ depends on the path of future valuations $\{m^{(*)}_{t+\kappa-1,t+\kappa}\}_{\kappa=1,2,\ldots,\infty}$ which govern the losses from a subsequent expected appreciation, and is captured by the term structure of interest rates. The role of the term structure in ERRP determination is the central message of this paper and we develop this relationship in section 3.

Importantly, this mechanism does not hinge on the assumed degree of financial market completeness. Under complete markets (CM), the exchange rate is uniquely given by $\mathbb{E}[e_{t+\kappa}] -$

[17] 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.

$e_t = m^*_{t,t+\kappa} - m_{t,t+\kappa}$, so the ERRP collapses to:

$$\lambda^{CM}_{t,\kappa} = \frac{1}{2}[\text{var}_t(m_{t,t+\kappa}) - \text{var}_t(m^*_{t,t+\kappa})] \tag{10}$$

The CM paradigm is a useful benchmark because it provides a determinate model for exchange rate dynamics without the need to specify the nature of shock processes affecting the economy.[18] Time variation in the risk premium therefore requires volatility to be stochastic, an issue we revisit in section 5.2. 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, Foresi, and Telmer, 2001). Another, more general, admissible exchange rate process under market incompleteness can be written as $\mathbb{E}[e_{t+\kappa}] - e_t = m^*_{t,t+\kappa} - m_{t,t+\kappa} + \eta_{t,t+\kappa}$, where $\eta_{t,t+\kappa}$ encapsulates non-trade risk (Lustig and Verdelhan, 2019). In this case the risk premium is given by:[19]

$$\lambda^{IM}_{t,\kappa} = \frac{1}{2}[\text{var}_t(m_{t,t+\kappa}) - \text{var}(m^*_{t,t+\kappa}) + \text{cov}_t(m_{t,t+\kappa} + m^*_{t,t+\kappa}, \eta_{t,t+\kappa})] \tag{11}$$

While more realistic, this formulation is less tractable and requires assumptions on the covariance on non-traded risk and the SDFs.

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 domestic yield curves because they contain information on transitory variation in SDFs.

3.1 Yield Curve-Augmented UIP Regression

To study how information in the yield curve, over and above spot interest rate differentials, matters for exchange rate determination, we augment regression (7) 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:

$$e_{j,t+\kappa} - e_{j,t} = \beta_{1,\kappa}\left(i_{j,t,\kappa} - i^*_{t,\kappa}\right) + \beta_{2,\kappa}(S_{j,t} - S^*_t) + \beta_{3,\kappa}(C_{j,t} - C^*_t) + f_{j,\kappa} + u_{j,t+\kappa} \tag{12}$$

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.

We use the yield curve slope and curvature in our extended regression to capture information in the yield curve in a parsimonious way. Along with the yield curve level, the yield curve slope

[18] Note that this is an assumption on cross-country markets and does not impose any restrictions on the time series of either Home or Foreign SDFs.

[19] Lustig and Verdelhan (2019) discuss the limits of this approach, subject to the additional assumption that $M_{t,t+\kappa}, M^*_{t,t+\kappa}$ are themselves not affected by market incompleteness.

and curvature are known to capture a high degree of variation in bond yields (Litterman and Scheinkman, 1991). We do not include the relative yield curve level in our baseline regression for two reasons.[20] First, we interpret the spot rate differential—the relevant quantity for UIP—as a proxy for the level.[21] Second, by nesting the canonical UIP regression (7) within our regression framework, the $\beta_{2,\kappa}$ and $\beta_{3,\kappa}$ coefficients have a dual interpretation in terms of exchange rate changes *and* ERRP. Combining (8) with (12), the *ex post* κ-period ERRP can be expressed as:

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

Comparing this with (12), 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 measure the yield curve slope as the difference between 10-year and 6-month yields—$i_{j,t,10y}$ and $i_{j,t,6m}$, respectively—such that $S_{j,t} = i_{j,t,10y} - i_{j,t,6m}$. Our curvature proxy a butterfly spread, a function of 6-month, 5 and 10-year yields (Diebold and Rudebusch, 2013), where $i_{j,t,5y}$ is the 5-year yield, such that $C_{j,t} = 2i_{j,t,5y} - (i_{j,t,6m} + i_{j,t,10y})$. We prefer these measures to principal component estimates of the yield curve slope and curvature. A principal component measure potentially contains 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 definition of yield curve slope and curvature.

Results Our benchmark results for regression (12) 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 (7). Driscoll and Kraay (1998) standard errors are reported in parentheses.

Two 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 reflects a reasonably low correlation between spot yield differentials, and the relative slope and curvature, respectively.[22] Importantly, this implies that the additional contribution of relative yield curve slope and curvature can be interpreted over and above the role for spot interest rate differentials, as an additional component of the ERRP.

[20] The exclusion of the yield curve level marks an important difference between our empirical framework and that of Chen and Tsang (2013).

[21] The yield curve level is often proxied by a specific bond yield—often the 10-year rate.

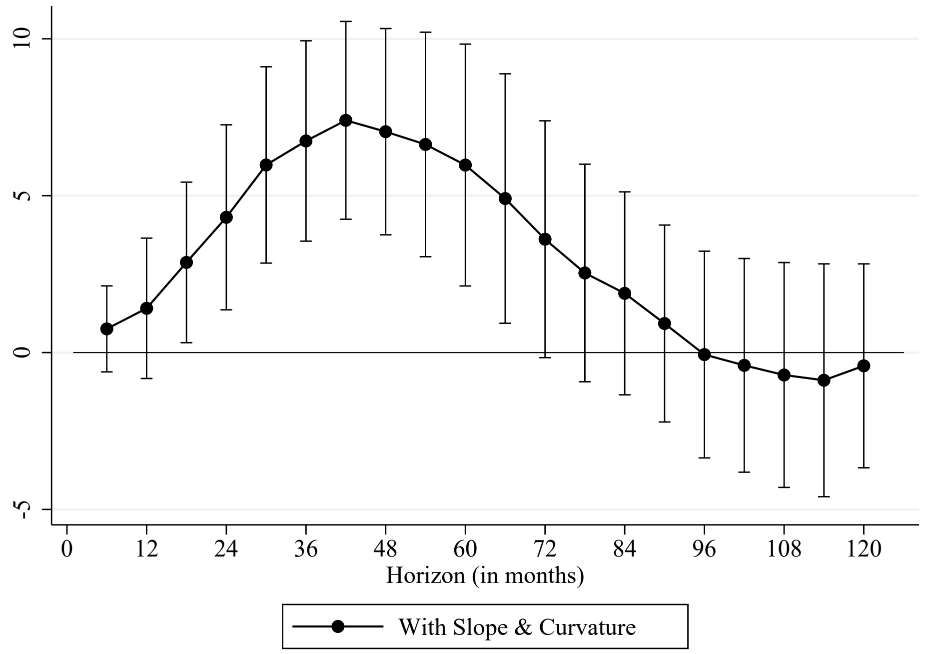
[22] To a large extent, this low correlation is driven by the approximate orthogonality of each country's yield curve level, slope and curvature proxies. This is the case because the proxies approximate factors constructed using principal components analysis, which are orthogonal by definition. Although the regressions do not include the relative level of the yield curve, the κ-period interest rate differential is likely to include similar information. Therefore, it is somewhat unsurprising that the inclusion of the *relative* yield curve slope and curvature only minimally affect point estimates of $\beta_{1,\kappa}$.

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

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

Notes: Column (1) presents coefficient estimates from regression (7)—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 (12)—the augmented regression—using the relative yield curve slope and curvature (measured using proxies) as additional regressors. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD from 1980:01 to 2017:12, including country fixed effects. The panel is unbalanced. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively, using Driscoll and Kraay (1998) standard errors (reported in parentheses).

Figure 4: Estimated relative slope coefficients from augmented UIP regression



Notes: Black circles denote $\hat{\beta}_{2,\kappa}$ point estimates from regression (12). The horizontal axis denotes the horizon κ in months. In regression (12), 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.

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 $\beta_{2,\kappa}$ coefficient estimates with respect to the horizon 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. In figure 4, 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 relative 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 yield curve slope is insignificantly different from zero.

In Appendix B.2, we explore the robustness of this benchmark result. First, we assess the significance of our coefficient estimates using more conservative standard errors. In the main body of the paper, we report Driscoll and Kraay (1998) standard errors, which are robust to heteroskedasticity and autocorrelation—an important consideration given that the error terms in forecasting regressions like (7) are moving average processes of order $\kappa - 1$ by construction. However, for the longer horizon variants of (7), where the number of non-overlapping observations can be limited, size distortions—i.e. the null hypothesis being rejected too often—are a pertinent concern, especially with small samples and persistent regressors (Valkanov, 2003). To carry out more conservative inference, we draw on Moon, Rubia, and Valkanov (2004), who propose the scaling of t-statistics by $1/\sqrt{\kappa}$, showing that these scaled t-statistics are *approximately* standard normal when regressors are highly persistent.[23] Using the more conservative scaled t-statistics, our primary result—the tent-shaped pattern for $\beta_{2,\kappa}$ estimates with respect to horizon κ—remains significant. However, the contribution of the relative curve is insignificant at all horizons. Second, we demonstrate that our findings are robust to sample period choice. In particular, the coefficients on the relative slope remain significant and tent-shaped with respect to maturity when the sample is ended in 2008:06, to omit the crisis and post-crisis period. Third, we show that the tent-shaped pattern for the loading on the relative slope broadly holds at a country level, by separately estimating (7) and (12) for each of our six currencies *vis-à-vis* the USD.

3.2 Interpreting the Role of the Yield Curve

These results imply that the relative yield curve slope plays an economically and statistically significant role in explaining future exchange rate movements at business cycle horizons especially. To interpret these findings, we argue that the depreciation associated with a having a relatively steep slope is consistent with the low-risk asset dynamics for the exchange rate we described in section 2.3, reflecting a business cycle risk component of ERRP.

Yield Curve Slopes A first key building block for this interpretation is an understanding of yield curve slopes within a closed economy setting. For a Home agent investing in an n-period Home bond, the relevant Euler equation (3) can be rewritten as:

$$\frac{1}{R_{t,n}} = \mathbb{E}_t \left[\prod_{i=0}^{n-1} M_{t+i,t+i+1} \right] \tag{13}$$

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,[24] then Piazzesi and Schneider (2007) show that

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

Here, the covariance term on the right-hand side is the risk premium on bonds. The variance term reflects Jensen's inequality. 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 time $t+1$ to $t+n$. The risk premium is positive if today's one-period SDF is

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

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

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:[25]

$$S_t \approx -\text{cov}_t\left(m_{t,t+1}, \mathbb{E}_{t+1}\sum_{i=1}^{n-1} m_{t+i,t+i+1}\right) \tag{15}$$

where the right-hand-most Jensen's inequality term in (14) has been suppressed. As a result, the yield curve will be upward sloping on average if the right-hand side of (14) is positive. In turn, the fact we empirically see 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 time $t+1$ to $t+n$ is indeed negative.

Extending (15) to our general international asset pricing framework indicates that the relative cross-country yield curve slope, $S^R \equiv S - S^*$, can be interpreted as the difference between investors' SDF autocovariance.

ERRP and Transitory Risk To assess and interpret the driver of the relationship between the relative yield curve slope and exchange rate dynamics, we consider the following decomposition of pricing kernels V_t, proposed in Alvarez and Jermann (2005):

$$V_t = V_t^{\mathbb{T}} V_t^{\mathbb{P}}$$

for all t, 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.[26] 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, Lustig et al. (2019) show that the cross-country difference in permanent SDF volatility must be

[25] 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 $\mathbb{E}_t\left[rx_{t+1,n}\right] \approx y_{t,n} - y_{t,1} \equiv S_t$.

[26] 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 $V_t^{\mathbb{P}} = \lim_{\kappa\to\infty} \frac{P_{t,\kappa}}{\delta^{t+\kappa}} V_t = \lim_{\kappa\to\infty} \frac{\mathbb{E}_t\left[V_{t+\kappa}\right]}{\delta^{t+\kappa}}$.

zero. In the limit, the ERRP is given by:

$$\lim_{\kappa \to \infty} \mathbb{E}_t[\lambda_{t+\kappa}] = \frac{1}{2}\left[\text{var}_t(\nu^{\mathbb{P}}_{t+1}) - \text{var}_t(\nu^{\mathbb{P}*}_{t+1})\right] \tag{16}$$

where $\nu_t^{(*)} \equiv \log(V_t^{\mathbb{P}(*)})$.[27] To reconcile empirical UIP deviations with theory at both short and long horizons under the complete markets benchmark (10), we require that high-rate currencies have relatively less volatile transitory pricing kernels ($\text{var}_t(\nu_t^{\mathbb{T}(*)})$), while the volatilities of the permanent components are similar. This is a key motivation for our main empirical exercise since transitory movements in SDFs, in contrast to permanent ones, are captured by the term structure of risk free yields as we illustrate in the examples below.

Exchange rate dynamics appear to characterise an equilibrium adjustment described in section 2.3, reallocating returns intertemporally to investors with the highest valuation of returns. If innovations were permanent, investors have no incentive to reallocate consumption intertemporally since the relative valuation over time is unchanged. Consequently, the mechanism we describe is specific to transitory innovations, which we interpret as business cycle fluctuations. Transitory innovations to the pricing kernel are relevant for exchange rate dynamics and carry most explanatory power at 3 to 5-year horizons. Consequently, the future path of SDFs, as captured by the term structure of the yield curves, is key to understanding exchange rate dynamics. We now present analytical examples explicitly relating variation in ERRP and the cross-country yield curve slope differential.

Example 1 (First-Order Autrogressive 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 (10) can be written*

$$\lambda_{t,\kappa} = \frac{1}{2}\rho_\nu^{2(\kappa-1)}\left[\text{var}_t\left(\nu_{t+1}\right) - \text{var}_t\left(\nu_{t+1}^*\right)\right] \tag{17}$$

and, using (15), the relative yield curve slope can (approximately) be written as:

$$S_t^R \equiv S_t - S_t^* = \left(1 - \rho_\nu^{(n-1)}\right)\left[\text{var}_t\left(\nu_{t+1}\right) - \text{var}_t\left(\nu_{t+1}^*\right)\right] \tag{18}$$

Combining (17) and (18) yields the following relationship between the empirical ERRP and the relative yield curve slope:

$$\lambda_{t,\kappa} = \frac{1}{2}\frac{\rho_\nu^{2(\kappa-1)}}{1-\rho_\nu^{(n-1)}}S_t^R. \tag{19}$$

The AR(1) structure delivers analytical simplicity and $\rho_\nu \in (0,1)$ ensures a positive yield curve slope, a salient empirical feature on average. Figure 5 demonstrates the strength of the relationship between ERRP $\lambda_{t,\kappa}$, across horizons κ, and the relative slope S^R for an example where $\rho_\nu = 0.9$. This parameterisation delivers a simulated first-order SDF autocorrelation

[27] 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.

of around -0.02, close to the lower bound identified in Chrétien (2012). In this case, and all instances where $\rho_\nu \in (0,1)$ such that yield curves slope upward for this AR(1) example, there is a positive relationship between the ERRP and the relative slope, declining as κ increases.[28] 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 can reflect business cycle risk.

Example 2 (Second-Order Autrogressive 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 (10) can be written*

$$\lambda_{t,\kappa} = \frac{1}{2}\psi_{\kappa-1}^2 \left[\mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{var}_t\left(\nu_{t+1}^*\right)\right] \tag{20}$$

and, using (15), the relative yield curve slope can (approximately) be written as:

$$S_t^R \equiv S_t - S_t^* = \left(1 - \psi_{n-1}\right)\left[\mathrm{var}_t\left(\nu_{t+1}\right) - \mathrm{var}_t\left(\nu_{t+1}^*\right)\right] \tag{21}$$

Combining (20) and (21) yields the following relationship between the empirical ERRP and the relative yield curve slope:

$$\lambda_{t,\kappa} = \frac{1}{2}\frac{\psi_{\kappa-1}^2}{1-\psi_{n-1}} S_t^R. \tag{22}$$

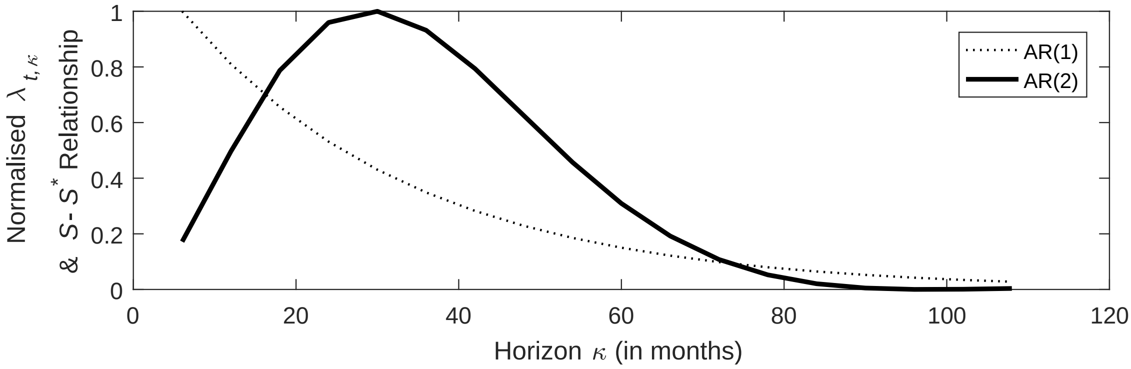
This example is able to capture both the positive sign of the relationship between ERRP and the relative slope, and the tent-shaped relationship with respect to horizon κ under certain restrictions on the parameters ψ_i.[29]

To demonstrate 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 equations (19) and (22), respectively. To highlight horizon-variation in the relationship, we normalise both lines by the maximum factor across horizons, such that the peak normalised relationship is unity. The parameters are chosen to ensure the autoregressive processes are mean-reverting and, thus, capture the transitory component of the pricing kernel. 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

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

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

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 $\rho_2 = -0.75$.

long horizons.[30]

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, as reflected by a negative autocovariance of consumption.[31] The representative investor wants to reallocate returns to the short run. We have shown both empirically, and independently in a theoretical example, that the difference in the incentive to reallocate returns intertemporally across countries, as captured by the relative slope, is a strong predictor of exchange rate movements at numerous horizons.

3.3 Recasting the Slope: Holding Period Returns

Our empirical analysis in section 3.1 has the advantage of nesting the benchmark UIP regression and, in turn, allows simultaneous interpretation of the relationship between the relative yield curve slope and exchange rate dynamics, as well as with the ERRP. However, two pertinent concerns around the empirical framework remain, which we address in this subsection.

The first concern is conceptual. In section 3.1, we assume that investors hold Home and Foreign bonds to maturity, investigating returns to each bonds' maturity. In reality, investors may not hold bonds to maturity, with bond holding periods h falling short of bond maturity κ. As a consequence, in a risk-neutral environment there would be a whole set of no-abitrage conditions, for bonds of different maturities over varying holding periods.

[30]In this example, the first-order autocorrelation of the SDF is positive, but medium-horizon SDF autocorrelations turn negative as the tent-shaped relationship turns downward. Although Chrétien (2012) shows empirical evidence that the first-order autocorrelation of SDFs is small but negative, this finding might be driven by the permanent component of SDFs. In our application, the yield curve slope captures only the transitory SDF component.

[31]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.

The second concern is empirical. Because regression (12) is based on *ex post* exchange rate outcomes, the sample of available observations diminishes as bond maturity κ increases. Moreover, the number of available non-overlapping observations also diminishes, posing additional challenges for inference. Although our benchmark analyses relies on standard errors that are robust to the serial correlation that overlapping observations result in, these two concerns can be further addressed by assessing the returns on bonds of a given maturity κ over different holding periods h. Doing this both acts as an additional robustness exercise to underpin our benchmark results, but also allows us to investigate whether price dynamics originate from bond permia or the currency premium, extending the findings in Lustig et al. (2019).[32]

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-period holding period return on a κ-period 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 time $t+h$ when its maturity has diminished by h periods relative to its time t price. The (log) excess return on that bond over that holding period h is thus

$$rx^{(\kappa)}_{t,t+h} = \log\left[\frac{HPR^{(\kappa)}_{t,t+h}}{R_{t,h}}\right] \tag{23}$$

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

The h-period (log) return on a domestic κ-period 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:

$$rx^{(\kappa),\$}_{t,t+h} = \log\left[\frac{HPR^{(\kappa)}_{t,t+h}}{R^*_{t,h}}\frac{\mathcal{E}_t}{\mathcal{E}_{t+h}}\right] = \log\left[\frac{HPR^{(\kappa)}_{t,t+h}}{R_{t,h}}\right] + \log\left[\frac{R_{t,h}}{R^*_{t,h}}\frac{\mathcal{E}_t}{\mathcal{E}_{t+h}}\right] \equiv rx^{(\kappa)}_{t,t+h} + rx^{FX}_{t,t+h} \tag{24}$$

where the rx^{FX}_{t+h} following the last equality represents the (log) currency excess return, defined relative to (8) as $rx^{FX}_{t,t+h} \equiv -\lambda_{t,h}$. In addition, the limiting relationship between currency excess returns and the per period return on infinite-maturity bonds, as shown in Alvarez and Jermann (2005) and Lustig et al. (2019), is:

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

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

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

[32] 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.

where the dependent variable $y_{j,t,h}$ 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 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 $rx^{FX}_{t,t+h} = -\lambda^{H}_{t,h}$, then the interpretation of $\gamma_{1,h}$ is reversed: $-\gamma_{1,h}$ can be interpreted in a similar way to $\beta_{2,\kappa}$ when $y_{j,t,h} = rx^{FX}_{k,t,t+h}$.

Equation (24) also indicates that the dollar-bond excess return is equal to the sum of the local currency-bond and currency excess returns, $rx^{(\kappa),\$}_{t,t+h} = rx^{(\kappa)}_{t,t+h} + rx^{FX}_{t,t+h}$. 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.[33] 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 $rx^{(120),\$}_{j,t,t+6} - rx^{(120)}_{US,t,t+6}$.[34]

More generally, exploring our results at all holding periods h and for all maturities κ, two observations are noteworthy.[35]

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 dollar-bond risk premium difference. Although significant at shorter holding periods, the relative slope loadings are quantitatively small for local currency-bond risk premia differences and are dominated by its loadings on currency excess returns in explaining the relative slope's impact on dollar-bond risk premia differences. This supports the findings from our benchmark augmented UIP regression in section 3.1. A steep domestic yield curve is associated with a higher ERRP for Home investors and, in turn, a higher excess return on Home bonds in US dollar units, dominating the influence of

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

[34]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).

[35]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.

the relative slope on local currency-bond excess returns differences. Through the lens of our theory, a steep yield curve today signals relatively low distant-future valuations but, via (15), relatively high nearer-term valuations. Reflecting this, the ERRP adjusts to transfer returns intertemporally towards periods in which they are valued relatively highly.

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. The loading on the relative slope is largest in magnitude at the 36-month horizon, for both $rx^{FX}_{j,t,t+h}$ and $rx^{(\kappa),\$}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$ for all bonds with maturity of 42-months or more. The fact the relative slope loading is similar across bond maturities for a given holding period is telling. It 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.

Second, 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 for the same bond maturity is significantly non-zero over longer holding periods. While, in the former case, the influence of the relative slope on currency and local-currency bond returns offset one another (in line with Lustig et al., 2019), our results indicate that the influence of the relative slope on the currency premium dominates over longer holding periods (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 with maturity.

4 Accounting for Liquidity Yields

In recent work, Engel and Wu (2018) show evidence that the liquidity yield on government bonds is a significant driver of exchange rates. The liquidity yield on a bond is the non-pecuniary return that government bonds provide because of their safety, ease of sale and value of collateral. Consistent with the predictions of a two-country New Keynesian model, they find that more liquid currencies tend to appreciate contemporaneously focusing on 1-year interest rates. In light of this, an important robustness test for our headline result is that the relative yield curve slope continues to significantly influence exchange rate dynamics when accounting for liquidity yields, insofar as liquidity yields influence domestic yield curves.

In this section, we extend our empirical specification to account for liquidity yields. The analysis serves two purposes. First, we demonstrate that the tent-shaped relationship between relative slope and ERRP across horizons is robust to the inclusion of liquidity yields. Second, we build on the results of Engel and Wu (2018) by studying the influence of the term structure of liquidity yields, not just a single horizon. Our results suggest that liquidity yields most strongly influence ERRP at long horizons, indicating that they, in part, capture permanent innovations to SDFs relevant for exchange rate dynamics and thus cross-sectional ERRP. In this sense, liquidity yields operate independently from the business cycle risk we attribute to relative yield curve slopes operating via transitory SDF innovations.

4.1 Liquidity Yield-Augmented Regression

To account for liquidity yields in our empirical setup, we use data on the term structure of liquidity yields from Du et al. (2018), as in Engel and Wu (2018).[36] These measure the difference between riskless market rates the government bond rate at different maturities to quantify the implicit liquidity yield on a government bond, correcting for other frictions in forward markets or sovereign default 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 $\eta^R_{j,t,\kappa}$ thus 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 in this section.[37] For this reason, our preferred empirical specification extends on the excess return regressions in section 3.3, rather than the augmented UIP regression in section 3.1.[38]

Our benchmark regression is therefore:

$$\mathrm{y}_{j,t,h} = \gamma_{1,h} \left(S_{j,t} - S^*_t\right) + \gamma_{2,h}\eta^R_{j,t,\kappa} + f_{j,h} + \varepsilon_{j,t+h} \tag{26}$$

where the dependent variable $\mathrm{y}_{j,t,h}$ is either the excess return on the Home bond in USD relative to the US $rx^{(\kappa),\$}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$ (the 'dollar-bond return' difference), the excess return from Home currency $rx^{FX}_{j,t,t+h}$, or the excess return on the Home bond in Home currency units relative to the US $rx^{(\kappa)}_{j,t,t+h} - rx^{(\kappa)}_{US,t,t+h}$ (the 'local currency-bond return' difference). Here, 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 $rx^{FX}_{t,t+h}$ is the dependent variable, a positive $\gamma_{2,h}$ is hypothesised, with an increase in relative US Treasury liquidity associated with a contemporaneous appreciation of the USD (depreciation of Home currency) that lowers the ERRP $\lambda_{t,\kappa}$ (increases currency excesses return $rx^{FX}_{t,t+h}$). The interpretation of $\gamma_{1,h}$ is unchanged relative to section 3.3.

Results The results for the dollar-bond excess return difference 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 holding periods for long-term bonds, as well as for 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.

[36]The Du et al. (2018) data is available for 12, 24, 36, 60, 84 and 120-month tenors only, constraining the maturities we assess in this section.

[37]For instance, a 10-year forecasting regression over a 20-year sample has zero non-overlapping observations.

[38]Nevertheless, Appendix B.4 presents extended UIP regression results, with liquidity yields and relative yield curve factors as additional regressors. These corroborate the results from excess return regressions reported in the main body of the paper.

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: $rx^{(\kappa),\$}_{t,t+h}$, Coefficient on η^R_κ										
12m	0.03									
	(0.02)									
24m	0.00	0.04	0.06**							
	(0.02)	(0.03)	(0.03)							
36m	0.01	0.04	0.07**	0.12***	0.16***					
	(0.02)	(0.03)	(0.03)	(0.04)	(0.04)					
60m	-0.00	0.03	0.05	0.09**	0.14***	0.17***	0.20***	0.21***	0.21***	
	(0.03)	(0.03)	(0.03)	(0.04)	(0.05)	(0.05)	(0.04)	(0.04)	(0.04)	
84m	0.00	0.03	0.04	0.08**	0.13***	0.16***	0.18***	0.20***	0.20***	0.22***
	(0.03)	(0.03)	(0.03)	(0.04)	(0.05)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)
120m	-0.01	0.02	0.04	0.07*	0.12***	0.16***	0.21***	0.24***	0.27***	0.29***
	(0.02)	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.03)	(0.04)
N	1,733	1,697	1,661	1,625	1,589	1,553	1,517	1,481	1,445	1,409

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

Panel A.ii documents a pattern of monotonically increasing $\gamma_{2,h}$ coefficients for liquidity yields. As the holding period increases, the loading on the liquidity yield becomes more positive and more statistically significant. 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 currency 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 returns predominantly works through currency excess returns, with the sign of coefficients for the former matching those of the latter.

Panel B.i of table 5 demonstrates that the relative slope loadings for currency excess returns are negative and tent-shaped with respect to holding period. As before, an increase in the Home country's relative yield curve slope is associated with a subsequent depreciation, that occurs most strongly at business cycle horizons. In contrast, the relative slope loadings for local currency-bond excess return differences are positive and quantitatively small.

The $\gamma_{2,h}$ coefficient estimates for currency excess returns are presented in panel B.ii of table 5. For a given maturity, the coefficient on the relative liquidity yield rises monotonically with

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.

respect to holding period, growing in significance. This suggests that the liquidity yields may contain some information about permanent SDF variations, extending the results in Lustig et al. (2019). In contrast, the $\gamma_{2,h}$ loadings for local currency-bond excess returns are negative and comparably small in magnitude.

4.2 Interpreting the Role of Liquidity Yields

We briefly interpret our results through the lens of our preference-free framework, in which we draw on Jiang et al. (2018) to model liquidity yields. The theory highlights a key difference between liquidity yields and relative yield curve slopes in their contribution to currency premia: 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.

Extending (5) and (4), let the Home and Foreign (US) representative investors' Euler equa-

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: $rx^{(\kappa)}_{t,t+h}$, 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.

tions for US bonds be given by,

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

$$e^{\xi^*_{t,\kappa}} = \mathbb{E}_t[\hat{M}^*_{t,t+\kappa}R^*_{t,\kappa}] \tag{28}$$

where $\xi_{t,\kappa}$ denotes the liquidity yield of US bonds for Home investors and ξ^* denotes the liquidity yield of US bonds for the Foreign representative investor. For simplicity, the liquidity yield of Home bonds is normalised to zero. In comparison to (5) and (4), the SDFs in (27) and (28) have the following relation: $M^{(*)}_{t,t+\kappa} \equiv \hat{M}^{(*)}_{t,t+\kappa}e^{-\xi^{(*)}_{t,\kappa}}$.[39]

Jiang et al. (2018) conjecture the exchange rate process follows $\Delta^\kappa e_{t+\kappa} = m_{t,t+\kappa} - m^*_{t,t+\kappa} + (\xi^*_{t,\kappa} - \xi_{t,\kappa})$, which cannot hold under complete markets. Consequently, we interpret liquidity yields as a form of non-traded risk, as in Lustig and Verdelhan (2019) who show that under

[39] In Appendix C.4, we show that this can be derived by decomposing the SDF into a pecuniary and a liquidity component, and we show that liquidity yields confound the liquidity need of the investor, the liquidity offered by the asset and the asset-investor specific liquidity yield.

incomplete markets:

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

such that exchange rates (or, conversely, the Foreign SDF) are uniquely determined for a given wedge $\eta_{t,t+\kappa}$. Liquidity has a natural interpretation as a form of non-traded risk since it corresponds to missing insurance markets, such that $\eta_{t,t+\kappa} \propto \xi^*_{t,\kappa} - \xi_{t,\kappa}$. If there existed perfect liquidity insurance, bonds would not carry a liquidity premium.[40] The risk premium can then be expressed as:

$$\mathbb{E}_t[\lambda_{t,\kappa}] = \frac{1}{2}[\mathrm{var}_t(m_{t,t+\kappa}) - \mathrm{var}_t(m^*_{t,t+\kappa})] + \mathbb{E}_t[\xi^*_{t,\kappa} - \xi_{t,\kappa}]. \tag{29}$$

Thus liquidity yields can contribute to exchange rate predictability, as in table 5.[41]

The mechanism discussed in section 2.3 generalises to the case with liquidity considerations. Notice that the risk premium reflects the difference in liquidity yields from the same asset. It therefore captures investors' liquidity needs and asset-investor specific liquidity. If the foreign economy is liquidity constrained such that the liquidity yield differential is high $\xi^*_{t,\kappa} > \xi_{t,\kappa}$, ERRP result in an expected appreciation of foreign exchange rates since foreign investors are willing to forego pecuniary returns in favour of liquidity. In levels, this is consistent with Jiang et al. (2018) who show that highly liquid countries experience a contemporaneous appreciation.

Deriving (16) with an additional term for liquidity yield differentials results in:

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

If $\lambda^{(\infty)}_{t,1}$ tends to zero, our empirical results suggest that drift in the relative liquidity yield may contribute to cross-country differences in the variation of the permanent component of pricing kernels:

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

5 Yield Curve Inversions and Exchange Rate Dynamics

So far in this paper, we have focused on the role of the term structure in capturing business cycle risks relevant for exchange rate dynamics. In this section, we study how the term structure also captures changes in patterns of risk—i.e. reversals in UIP dynamics—which we find coincide with yield curve inversions.

5.1 Testing the Influence of Yield Curve Inversions

To test the influence of yield curve inversions, on short to medium-horizon UIP dynamics especially, we alter our benchmark yield curve-augmented UIP regression specification. Consistent

[40] Formally, if there is a traded asset that yields returns $\hat{\xi}$ in the form of liquidity services, by the definition of non-traded risk then $\mathrm{cov}_t(\xi_{t,\kappa}, \eta_{t,t+\kappa}) = 0$.

[41] ERRP are determined by the difference in liquidity yields which matters because US investors' liquidity yields lowers their domestic yields.

with our findings in section 3, we continue to include the relative yield curve slope and curvature in our regression. Second, to account for the effect of yield curve inversions, we define an indicator variable $\mathbb{1}_{j,t} \equiv \mathbb{1}(S_{j,t} < 0)$ which is set equal to unity in months when the domestic yield curve is inverted. Because we use the domestic, country-j, yield curve as the indicator, our indicator variable varies across countries, providing useful cross-sectional variation to identify interaction effects. The general updated regression is:

$$\begin{aligned} e_{j,t+\kappa} - e_{j,t} = & \beta_{1,\kappa}\left(i_{j,t,\kappa} - i^*_{t,\kappa}\right) + \delta_\kappa\left[\left(i_{j,t,\kappa} - i^*_{t,\kappa}\right) \times \mathbb{1}_{j,t}\right] + \beta_{2,\kappa}\left(S_{j,t} - S^*_t\right) \\ & + \beta_{3,\kappa}\left(C_{j,t} - C^*_t\right) + \vartheta_\kappa \mathbb{1}_{j,t} + f_{j,\kappa} + u_{j,t+\kappa} \end{aligned} \tag{30}$$

In the main body of the paper, we report results using an indicator variable that is set to unity during months of a domestic yield curve inversion only, and zero otherwise. However, in an Appendix, we demonstrate that our findings are robust to extending the definition of the dummy variable indicator such that it is equal to unity for an additional 1 to 2 years after the yield curve ceases to be inverted, capturing the real economic effects of a downturn. To ensure that we capture the same set of global macroeconomic events for all countries, we shorten our benchmark sample to one that is common for all our G7 currencies, 1992:07-2017:12, although our results are robust to trimming the sample at 2008:06 to exclude the global financial crisis and post-crisis periods.[42]

Figure 6 presents coefficient estimates for $\beta_{1,\kappa}$ (no inversion) and $\beta_{1,\kappa} + \delta_\kappa$ (with inversion) for all horizons. At short horizons, the UIP coefficient, and the sign of ERRP, is significantly different in periods where the domestic yield curve is inverted. For instance, at the 6-month horizon, the implied UIP coefficient outside of inversion periods is -1.55, while during periods of domestic yield curve inversion the implied coefficient is 7.71, implying a more than proportional relationship between exchange rates and interest differentials. Consistent with the 'New Fama Puzzle', our results indicate that the sign of short-horizon UIP coefficients flips in periods following a yield curve inversion. In contrast, at medium and longer-horizons, the differences between the two periods are insignificantly different.

Table 7 documents relative slope and curvature coefficient estimates, in addition to $\beta_{1,\kappa}$ and δ_κ estimates from the extended regression. Even when accounting for the additional interaction, the loadings on the relative slope continue to exhibit a tent shape with respect to maturity, peaking at the 3.5 to 4-year horizon.

5.2 Understanding Yield Curve Inversions

Our empirical contribution is to show that an inversion in the yield curve is associated with a statistically significant 'sign reversal' in the UIP regression coefficient at short horizons. Theoretical models reconcile UIP by generating $\text{cov}_t(\lambda_{t,\kappa}, i_{t,\kappa} - i^*_{t,\kappa}) < 0$, but are unable to deliver this covariance conditional on an upward sloping yield curve, and by extension, cannot explain the reversal in the sign of this covariance observed conditional on an inversion.

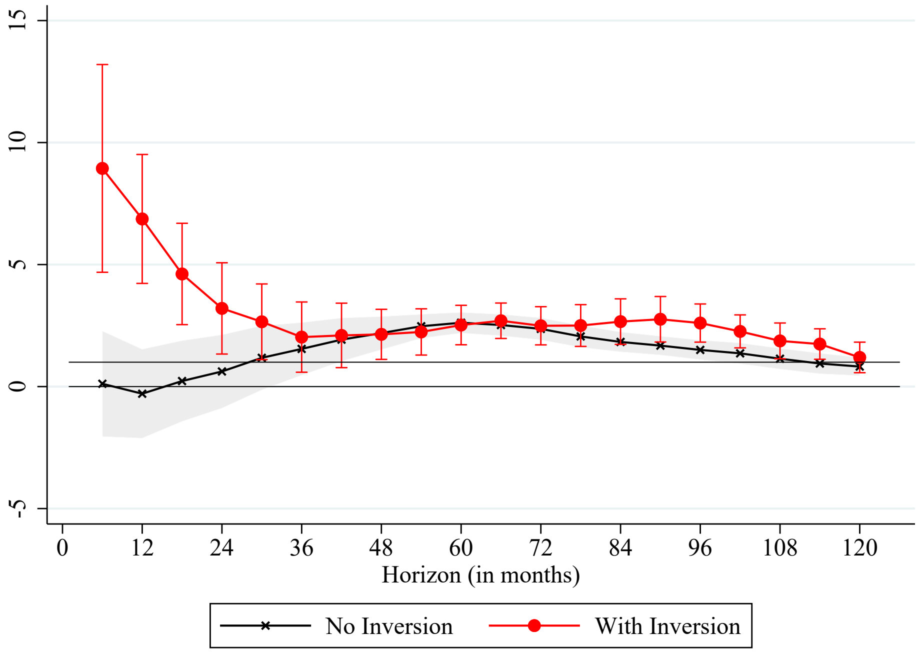
[42] We also show that extending the sample to 1980:01 is associated with significant changes in exchange rate dynamics, although differences are not as stark compared to the balanced panel sample that starts in 1992:07.

Table 7: Coefficient estimates from UIP regressions with additional yield curve inversion interactions

	(1)	(2)	(3)	(4)	(5)	(6)
	UIP Regression		Augmented Regression			
	$i_\kappa - i^*_\kappa$	$\mathbb{1} \times (i_\kappa - i^*_\kappa)$	$i_\kappa - i^*_\kappa$	$\mathbb{1} \times (i_\kappa - i^*_\kappa)$	$S - S^*$	$C - C^*$
6-months	-1.55**	9.26***	0.11	8.83***	1.51**	-1.25
	(0.71)	(1.86)	(1.11)	(1.91)	(0.71)	(1.01)
12-months	-1.71***	7.29***	-0.29	7.16***	2.22*	-1.21
	(0.60)	(1.03)	(0.93)	(1.09)	(1.23)	(1.70)
18-months	-1.50***	4.39***	0.23	4.39***	3.82**	-1.79
	(0.56)	(0.73)	(0.85)	(0.78)	(1.52)	(2.02)
24-months	-1.19**	2.60***	0.62	2.59***	5.25***	-3.02
	(0.53)	(0.68)	(0.77)	(0.74)	(1.76)	(2.29)
30-months	-0.85*	1.59***	1.18*	1.48**	7.27***	-5.21**
	(0.47)	(0.56)	(0.67)	(0.63)	(1.85)	(2.50)
36-months	-0.52	0.75	1.54***	0.48	8.73***	-7.58***
	(0.39)	(0.53)	(0.55)	(0.62)	(1.66)	(2.22)
42-months	-0.11	0.58	1.92***	0.17	9.94***	-10.07***
	(0.40)	(0.53)	(0.45)	(0.60)	(1.55)	(2.15)
48-months	0.37	0.33	2.19***	-0.05	9.94***	-10.38***
	(0.41)	(0.46)	(0.35)	(0.49)	(1.66)	(2.32)
54-months	0.83**	0.09	2.47***	-0.23	9.97***	-10.25***
	(0.40)	(0.43)	(0.25)	(0.46)	(1.86)	(2.42)
60-months	1.16***	0.14	2.62***	-0.10	9.71***	-9.61***
	(0.37)	(0.38)	(0.22)	(0.41)	(2.04)	(2.61)
66-months	1.39***	0.23	2.52***	0.17	8.06***	-6.28**
	(0.32)	(0.32)	(0.23)	(0.33)	(2.19)	(2.74)
72-months	1.58***	-0.01	2.36***	0.13	5.81***	-2.01
	(0.25)	(0.32)	(0.23)	(0.32)	(2.14)	(2.67)
78-months	1.55***	0.26	2.06***	0.45	3.92*	0.15
	(0.22)	(0.36)	(0.23)	(0.37)	(2.06)	(2.70)
84-months	1.50***	0.65	1.83***	0.83*	2.98	0.65
	(0.21)	(0.43)	(0.21)	(0.45)	(2.00)	(2.84)
90-months	1.47***	0.90*	1.68***	1.08**	2.24	1.17
	(0.20)	(0.46)	(0.20)	(0.49)	(1.95)	(2.95)
96-months	1.38***	0.86**	1.50***	1.10**	1.30	3.03
	(0.21)	(0.41)	(0.21)	(0.46)	(2.00)	(3.13)
102-months	1.26***	0.66*	1.36***	0.90**	1.40	3.05
	(0.22)	(0.36)	(0.22)	(0.41)	(1.97)	(3.01)
108-months	1.08***	0.49	1.14***	0.73*	0.75	4.03
	(0.22)	(0.34)	(0.22)	(0.40)	(1.99)	(3.00)
114-months	0.89***	0.55**	0.94***	0.80**	0.66	4.35
	(0.22)	(0.27)	(0.22)	(0.33)	(2.07)	(3.31)
120-months	0.77***	0.21	0.82***	0.37	1.32	3.26
	(0.19)	(0.24)	(0.19)	(0.28)	(1.80)	(3.24)

Notes: Columns (1) and (2) present estimates from a regression the κ-period exchange rate change on the κ-period interest rate differential $i_{j,\kappa} - i^*_\kappa$, a dummy variable indicator for country-j yield curve inversions $\mathbb{1}_j$ and an interaction between the two $\mathbb{1}_j \times (i_{j,\kappa} - i^*_\kappa)$. Columns (3)-(6) present estimates from the same regression with two additional regressors, the relative yield curve slope $S_j - S^*$ and curvature $C_j - C^*$. The inversion indicator $\mathbb{1}_j$ is set to 1 in months where the country-j yield curve slope is negative, and zero otherwise. 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 1992:07 to 2017:12. All regressions include country fixed effects. The panel is balanced and standard errors (reported in parentheses) are constructed according to the Driscoll and Kraay (1998) methodology. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively.

Figure 6: Estimated coefficients from UIP regression augmented with domestic yield curve inversion indicator interactions



Notes: Black crosses denote $\hat{\beta}_{1,\kappa}$ estimates and red dots denote $\hat{\beta}_{1,\kappa} + \hat{\delta}_{\kappa}$ estimates from regression (30). 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 US from 1992:07 to 2017:12 (common sample for all), including country fixed effects. Yield curve inversion indicator is set equal to 1 when the domestic yield curve slope $S_{j,t}$ is less than 0. 95% confidence intervals, calculated using Driscoll and Kraay (1998) standard errors are denoted by gray shaded area and red bars, respectively, around point estimates.

A sizeable literature explains UIP failures with external habits, building on Campbell and Cochrane (1999) or Epstein and Zin (1989) preferences, and stochastic volatility. Verdelhan (2010) generates the desired covariance between the ERRP and the interest rate differential using a consumption based model with external habits at the expense of a downward sloping yield curve.[43] Bansal and Shaliastovich (2013) use Epstein and Zin (1989) preferences and can generate both empirically consistent UIP deviations and an upward sloping term structure using idiosyncratic (country) shocks to real (consumption) and nominal (inflation) volatilty. However, for the ERRP and the cross-country yield differential, only total uncertainty matters and the model implies that the high-yield currency appreciates in excess of UIP, independently of the slope of the yield curve.

A large and growing literature on disaster risk and currency crashes attributes the excess

[43] Wachter (2006) show that countercyclical interest rates are needed to produce an upward sloping yield curve in this environment. However, this framework (as in Verdelhan (2010)) requires procyclical rates to deliver empirically consistent UIP deviations. Grasso and Natoli (2018) extend Wachter (2006) and present a model that delivers yield curve inversions following a temporary rise in consumption uncertainty but the model then produces UIP dynamics that we find to hold only conditional on an upwards yield curve.

volatility to non-Gaussian (jump) shocks to the SDF, see Gabaix (2012) and Farhi et al. (2015).[44] In particular, Farhi and Gabaix (2016) consider a domestic country-j SDF given by:

$$M_{j,t,t+1} = \left(\frac{C_{j,t+1}}{C_{j,t}}\right)^{1-\gamma} \times \begin{cases} 1 & \text{if no disaster at } t+1 \text{ with probability } 1-p_t \\ B_{t+1}^{-\gamma} F_{j,t+1} & \text{if disaster at } t+1 \text{ with probability } p_t \end{cases} \tag{31}$$

while during a disaster event, the world SDF is scaled only by $B_{t+1}^{-\gamma}$. The factor B_{t+1} denotes the size of the disaster and is common across countries as is probability of disasters, while $F_{j,t+1}$ is idiosyncratic and captures the sensitivity of the domestic currency to global disasters.[45]

Countries whose domestic currency depreciates heavily during disasters feature a higher interest rate. Farhi et al. (2015) investigate this model in reduced form, and define disaster vulnerability by $p_t(J_{j,t} - 1) = p_t(B_{t+1}^{-\gamma} F_{j,t+1} - 1)$. In the limit of small time intervals,[46]

$$\begin{aligned} i_{j,t} &= g_{j,t} - p_t \mathbb{E}_t[J_{j,t} - 1], \\ \lambda_t &= \text{cov}_t(\epsilon_t, \epsilon_t - \epsilon_t^*) + p_t \mathbb{E}_t[J_{j,t} - J_{j,t}^*] \end{aligned}$$

where $g_{j,t}$ is the growth rate of consumption in country j and ϵ_t is an innovation to consumption growth. For the risk premium, the first term relates to our standard expression, while the second term is the difference in vulnerabilities. If the probability of a disaster rises (or a disaster occurs) the high yield currency will tend to depreciate further, as we observed during the global financial crisis. Note that this is consistent with the mechanism outlined in section 2.3: investors in disaster-vulnerable countries experience a large depreciation to compensate for a large consumption shock. Outside of disasters, an appreciation reallocates consumption the the (uncertain) future.

Gabaix (2012) shows, in a closed economy context, that an upward sloping nominal yield curve will arise if disasters are associated with a jump in inflation, akin to a supply-driven shock. Piazzesi and Schneider (2007) rely on the same mechanism, but model inflation using Gaussian risk. The mechanism is that long-term bond returns are eroded by inflation and therefore bond premia must be rising with maturity. Our findings suggest that while yield curves are upward sloping during normal times, when the probability of disaster p_t is low, yield curve inversions arise at the onset of crises, when the probability of a disaster is high. Consequently, a combination of Gaussian inflation risk as in Piazzesi and Schneider (2007) and deflation associated with disaster events—consistent with a demand-driven shock—would deliver the conditional UIP deviations we document.[47] Were disaster events to be associated with

[44] This related to a large, earlier, empirical literature on Peso events. While Peso events focus on a small sample issue—i.e. that there are large rare depreciations out of sample—disaster risk focuses on risk premia which arise even if Peso events occur in sample.

[45] In Farhi and Gabaix (2016), the factor $F_{j,t+1}$ multiplies the yield in traded goods (world numéraire) from investing a single non-traded good. Their concept of an exchange rate coincides with traditional notions in the limit where the consumption basket consists only of non-tradable goods.

[46] Discrete time equations are presented in Farhi and Gabaix (2016).

[47] Consistent with this, Cujean and Hasler (2017) show that whilst an increase in nominal uncertainty increases bond premia (due to non-neutral inflation), an increase in real uncertainty (the component of σ_j relating to B_{t+1}) lowers bond premia.

inflation—i.e. via supply-driven shocks—this mechanism would be weaker. Indeed, Appendix B.5 shows that our results are somewhat sensitive to sample inclusion, with results for the UIP switch weakening when the sample start date is extended back 1980:01, which may be attributable to changes in the relationship between inflation and disasters over time.

6 Conclusion

In this paper, we explore the extent to which information in the term structure of interest rates can explain ERRP, over and above spot rate differentials. Our main empirical 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. This result arises in both a long-horizon panel UIP regression and a regression of excess returns over varying holding periods.

To interpret our findings, we begin with a reprisal of UIP cast in an equilibrium asset pricing setup, and we argue that ERRP lower the risk for investors holding foreign bonds. In particular, an expected depreciation systematically reallocates consumption intertemporally to investors with a relatively high valuation of returns. The relative slope of the yield curve across countries contains information about the future path of SDFs and captures the relative desire for intertemporal reallocation. By decomposing the SDF into a transitory and a permanent component, we show formally that ERRP variation predictable by the relative slope arises as compensation for transitory risk. By means of two analytical examples for SDF processes, we derive both the sign and the horizon-variation of the relationship between the relative slope of the yield curve and ERRP.

Our findings are robust to the inclusion of liquidity yields which we show operate through a distinct channel. In contrast to business cycle risk captured by the relative slope, liquidity yields contribute to cross-sectional differences across currencies and capture permanent differences in SDFs.

Finally, we show that yield curve inversions coincide with periods of reversals in the joint dynamics of exchange rates and interest rates. While the yield curve is upward sloping, standard ERRP dynamics suggest that high yield currencies experience an excess appreciation at short horizons. In contrast, following yield curve inversions, high yield currencies tend to experience an excess depreciation—even if we exclude the global financial crisis from our sample. While we show that this poses a challenge for many consumption asset pricing models, it can be reconciled in a model of rare disasters.

Appendix

A Data Sources

We use nominal zero-coupon government bond yields at maturities from 6 months out to ten years for 7 different industrialised countries in our benchmark sample: 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 nominal zero-coupon bond yields are not available from the start of the sample in all jurisdictions. Table 8 summarises the sources of nominal zero-coupon government bond yields, and the sample availability, for the benchmark economies in our study. In robustness analyses, we also assess results for a broader set of G10 currencies—adding New Zealand, Norway and Sweden—for which zero-coupon government bond yields are available to 2009:05 from Wright (2011).

Table 8: Yield Curve Data Sources

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

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

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

B Empirical Results

B.1 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 9, 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 document the robustness of our benchmark results in section 3.1, along a number of dimensions.

Conservative Inference As discussed in section 3.1, long-horizon forecasting regressions like (7) and (12) 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 in this Appendix.

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 in the paper's main body.

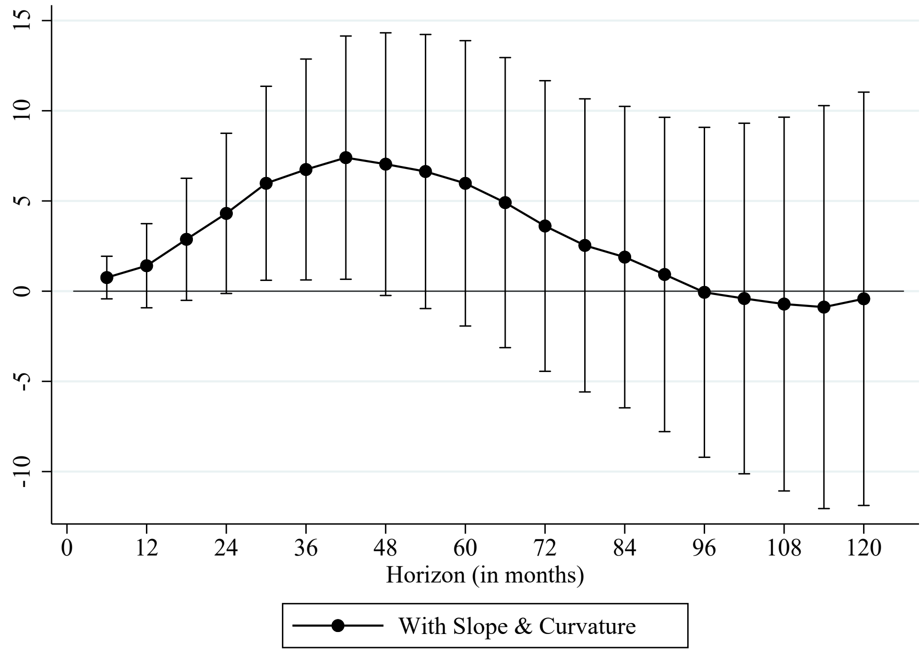
Importantly, our headline result—the tent-shaped pattern of the coefficient loading on the relative yield curve slope with respect to the horizon κ—is robust to the use of this more conservative inference. To demonstrate this, figure 7 plots the $\beta_{2,\kappa}$ estimates from (12) with 90% confidence intervals implied by these scaled t-statistics. Relative to table 1, point estimates are unchanged. But the error bands implied by the scaled t-statistics are wider. Nevertheless,

Table 9: 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 (7)—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.

Figure 7: Estimated relative slope coefficients from augmented UIP regression using more conservative inference



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

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

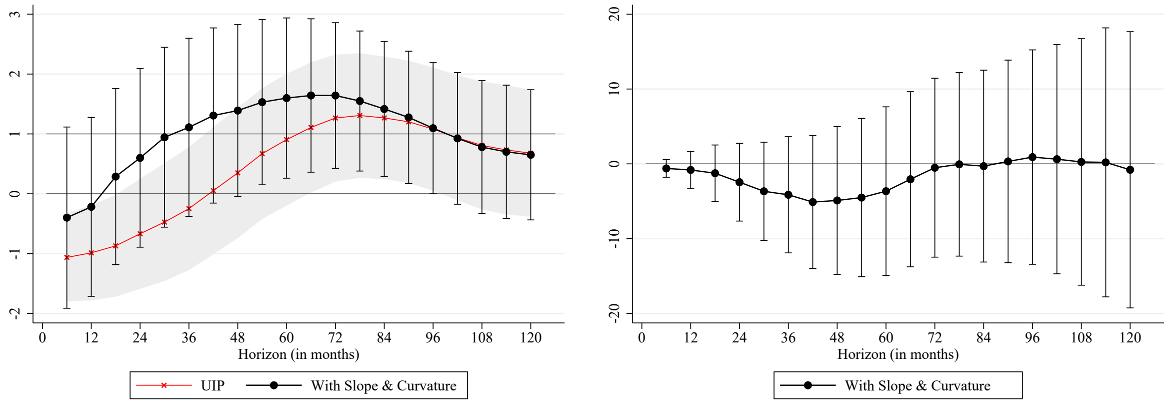
In addition, figure 8 plots the $\beta_{1,\kappa}$ and $\beta_{3,\kappa}$ coefficient estimates from (12) alongside the 90% confidence bands implied by the scaled t-statistics. While the overall pattern of $\beta_{1,\kappa}$ coefficients is broadly the same as the canonical UIP regression, the confidence bands with these more conservative t-statistics are wider. Importantly, the scaled t-statistics also imply that the coefficients on the relative yield curve curvature are statistically insignificant at all horizons.

Sub-sample Stability To assess the stability of our results, we estimate regression (12) 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.[48]

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

[48]We cannot run long-horizon regressions over a purely post-crisis sample because our regressions rely on *ex post* exchange rate changes.

Figure 8: 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 (12). The horizontal axis denotes the horizon κ in months. In regression (12), 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.

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.[49] In column (5), we drop the relative curvature from regression (12), 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. Although coefficient estimates vary in size and significance across countries, a relative slope coefficient estimates display a tent shape with respect to horizon κ for 4 of the 6 currencies in our sample (AUD, CHF, EUR, GBP).

[49]We do not have complete zero-coupon bond yield curves for these countries post-crisis, hence we omit them from our results in the main body of the paper.

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 (12)—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 (12)—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.

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

In this appendix, 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. The excess return regressions presented in the main body of the paper are our preferred empirical specification in the case of liquidity yields due to data constraints that heighten worries around the pervasiveness of non-overlapping observations for inference in longer-horizon UIP regressions. Nevertheless, in this appendix, we demonstrate that point estimates from an extended

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 (7) (UIP regression), (12) (yield curve-augmented UIP regression) and (32) (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.

UIP regression deliver similar results to those in section 4.1.

Using the definition 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 an extended UIP regression:

$$e_{j,t+\kappa} - e_{j,t} = \beta_{1,\kappa}\left(i_{j,t,\kappa} - i^*_{t,\kappa}\right) + \beta_{2,\kappa}(S_{j,t} - S^*_t) + \beta_{3,\kappa}(C_{j,t} - C^*_t) + \beta_{4,\kappa}\eta^R_{j,t,\kappa} + u_{j,t+\kappa} \quad (32)$$

Unlike Engel and Wu (2018), we assess the importance of the relative liquidity yield at a range of horizons, rather than just the 1-year tenor. Doing so provides additional novel insights, while also demonstrating the independent importance of yield curve factors and business cycle risk. 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 (32).

The results are presented in table 13, comparing the liquidity yield *and* yield curve-augmented regression (7) with the baseline UIP regression (7) and the yield curve-augmented regression (12) over a common sample.

The most important observation, consistent with the central claim of our paper that the relative slope captures a business cycle risk component of ERRP, is that the coefficient on the relative yield curve slope is robust to the additional inclusion of liquidity yields as a regressor in (32). 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.

Consistent with the findings in Engel and Wu (2018), the inclusion of the relative liquidity yield substantially improves the fit for exchange rates. At all horizons, the adjusted R^2 of the regression (32) exceeds that of (7) and (12). This provides additional evidence to complement the results in Engel and Wu (2018), showing that the relevance of convenience yields is important at all horizons. In addition, the largest increase in $\overline{R}^2$ from liquidity yields comes at the 10-year horizon, suggesting that liquidity yields most strongly influence longer-horizon exchange rate dynamics. The coefficient estimates in column (10) support this. Consistent with the hypothesis that a more liquid currency should appreciate contemporaneously, depreciating subsequently, our results suggest that this phenomenon especially powerful a medium to long horizons. In particular, $\beta_{4,\kappa}$ estimates are significantly negative from the 2-year horizon and beyond, growing in magnitude with respect to tenor. The negative coefficients can be interpreted in the following way: 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.

B.5 Yield Curve Inversion Interactions

In this Appendix, we document the robustness of our results in section 5. For reference figure 9 plots country-by-country time series of the 6-month interest rate differential and 6-month exchange rate change, *vis-à-vis* the USD, for the 6 currencies in our benchmark sample. In turn, domestic yield curve inversions are indicated by shaded areas in the figure. In some cases, the GBP especially, there is visible evidence of changes in short-horizon exchange rate dynamics during inversion periods—e.g. there is a broadly negative relationship between UK-US interest rate differentials and the exchange rate during the mid-2000s, which turns positive during the global financial crisis.

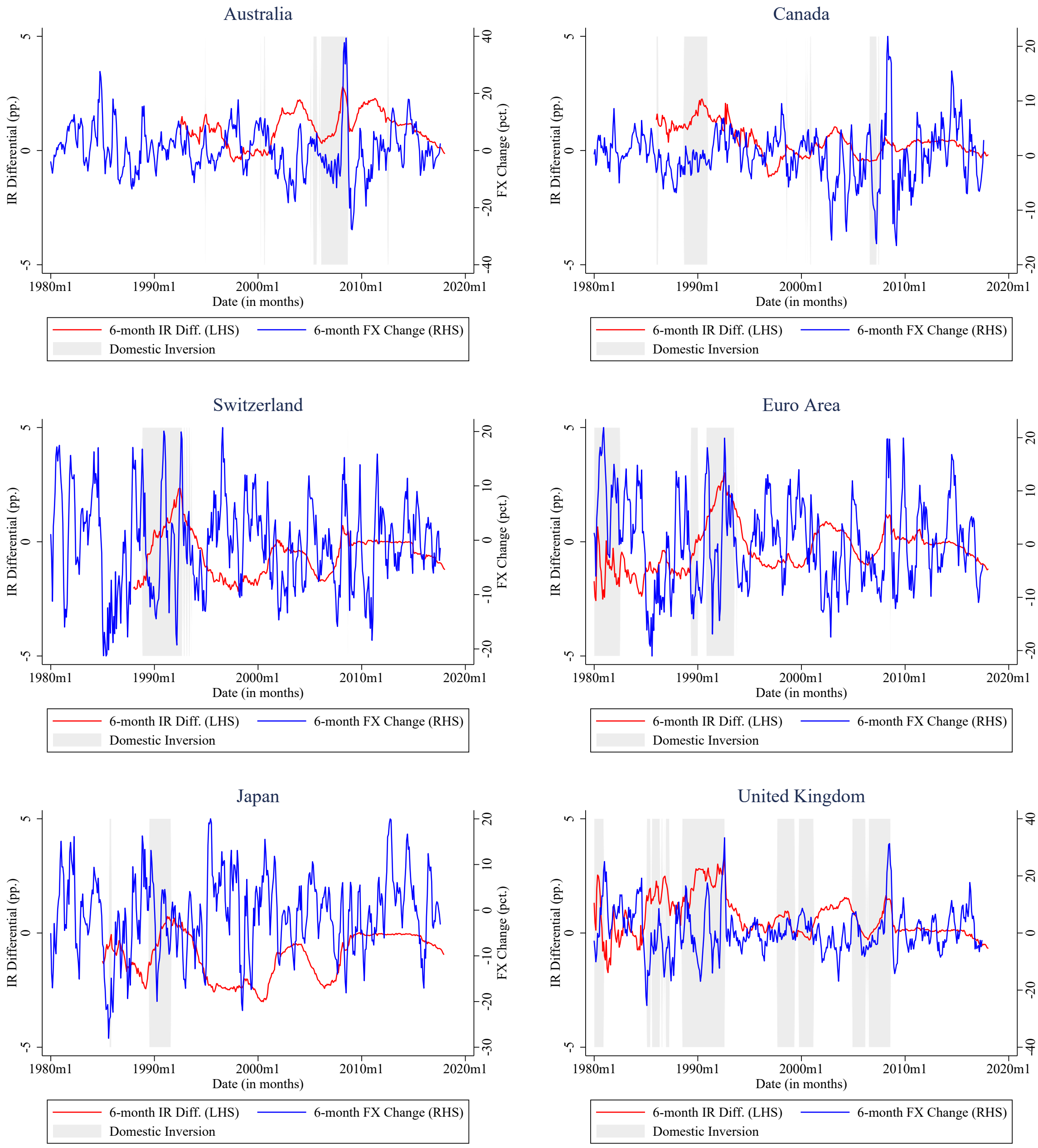
Persistence of Yield Curve Inversion Indicator In the main body of the paper, our benchmark inversion indicator $\mathbb{1}_{j,t}$ is set to unity in months where the domestic (country-j) yield curve is inverted, and zero otherwise. Insofar, as the inversion captures a change in expectations of future growth that are subsequently realised, then the change in exchange rate dynamics around inversions may be expected to persist, even after the inversion has ended.

In tables 14 and 15 we present robustness analyses in which the inversion indicator $\mathbb{1}_{j,t}$ is set to unity during a domestic inversion and for 1 and 2 years after it ends, respectively. The two tables demonstrate that our baseline result—a reversal in short-horizon UIP dynamics—is robust to this redefinition of the inversion indicator.

Additional Inversion Interactions with Relative Yield Curve Slope and Curvature In our benchmark analysis in section 5, we study only the interaction between the yield curve inversion and the UIP coefficient. Our benchmark regression controls for the relative yield curve slope and curvature, but does not additionally include their interaction with the inversion indicator.

In table 16, we demonstrate that our headline result—a reversal in short-horizon UIP dynamics—is broadly robust to adding these additional interactions to the regression. The

Figure 9: Time series of 6-month interest rate differentials and exchange rate changes alongside domestic yield curve inversions



Notes: Red lines denote 6-month zero-coupon bond yield differential *vis-à-vis* the US in percentage points. Blue lines denote 6-month exchange rate changes *vis-à-vis* the USD in percent. The gray shaded areas denote periods in which the domestic yield curve was inverted.

Table 14: Coefficient estimates from UIP regressions with additional yield curve inversion interactions, with inversion indicator persisting for 1-year after end of each inversion

	(1)	(2)	(3)	(4)	(5)	(6)
	UIP Regression		Augmented Regression			
	$i_\kappa - i_\kappa^*$	$\mathbb{1} \times (i_\kappa - i_\kappa^*)$	$i_\kappa - i_\kappa^*$	$\mathbb{1} \times (i_\kappa - i_\kappa^*)$	$S - S^*$	$C - C^*$
6-months	-1.63**	8.42***	-0.10	7.95***	1.37*	-1.15
	(0.72)	(1.89)	(1.12)	(1.98)	(0.74)	(1.05)
12-months	-1.69***	6.09***	-0.43	5.91***	2.00	-1.19
	(0.60)	(1.20)	(0.95)	(1.33)	(1.28)	(1.80)
18-months	-1.49***	3.60***	0.14	3.48***	3.63**	-1.89
	(0.57)	(0.88)	(0.88)	(0.97)	(1.63)	(2.12)
24-months	-1.18**	2.17***	0.50	2.02**	4.94***	-3.13
	(0.53)	(0.77)	(0.79)	(0.88)	(1.86)	(2.41)
30-months	-0.84*	1.31*	1.08	1.06	6.95***	-5.37**
	(0.46)	(0.69)	(0.68)	(0.80)	(1.93)	(2.60)
36-months	-0.54	0.85	1.47***	0.49	8.52***	-7.64***
	(0.38)	(0.60)	(0.56)	(0.66)	(1.71)	(2.25)
42-months	-0.14	0.79	1.86***	0.33	9.78***	-10.10***
	(0.39)	(0.58)	(0.45)	(0.60)	(1.59)	(2.16)
48-months	0.36	0.39	2.17***	-0.06	9.85***	-10.52***
	(0.41)	(0.52)	(0.34)	(0.50)	(1.68)	(2.32)
54-months	0.81**	0.19	2.45***	-0.22	9.88***	-10.40***
	(0.40)	(0.46)	(0.25)	(0.44)	(1.88)	(2.37)
60-months	1.14***	0.29	2.57***	-0.07	9.56***	-9.82***
	(0.36)	(0.39)	(0.22)	(0.38)	(2.08)	(2.57)
66-months	1.38***	0.18	2.49***	-0.03	7.95***	-6.73**
	(0.31)	(0.37)	(0.22)	(0.36)	(2.23)	(2.73)
72-months	1.56***	0.01	2.35***	-0.03	5.79***	-2.39
	(0.24)	(0.37)	(0.23)	(0.34)	(2.17)	(2.62)
78-months	1.54***	0.24	2.05***	0.26	3.95*	-0.22
	(0.22)	(0.38)	(0.22)	(0.36)	(2.07)	(2.62)
84-months	1.50***	0.47	1.84***	0.51	3.04	0.12
	(0.21)	(0.44)	(0.20)	(0.44)	(1.99)	(2.70)
90-months	1.48***	0.58	1.70***	0.63	2.39	0.50
	(0.20)	(0.44)	(0.20)	(0.45)	(1.96)	(2.85)
96-months	1.40***	0.56	1.54***	0.67	1.61	2.25
	(0.21)	(0.37)	(0.20)	(0.40)	(1.99)	(2.91)
102-months	1.28***	0.38	1.41***	0.48	1.89	2.03
	(0.22)	(0.32)	(0.22)	(0.36)	(1.91)	(2.76)
108-months	1.09***	0.23	1.17***	0.34	1.24	2.69
	(0.22)	(0.31)	(0.22)	(0.36)	(1.99)	(2.81)
114-months	0.90***	0.38	0.97***	0.55*	0.60	4.42
	(0.22)	(0.25)	(0.21)	(0.31)	(2.11)	(3.41)
120-months	0.77***	0.09	0.83***	0.18	1.38	3.28
	(0.20)	(0.21)	(0.20)	(0.25)	(1.83)	(3.42)

Notes: Columns (1) and (2) present estimates from a regression the κ-period exchange rate change on the κ-period interest rate differential $i_{j,\kappa} - i_\kappa^*$, a dummy variable indicator for country-j yield curve inversions $\mathbb{1}_j$ and an interaction between the two $\mathbb{1}_j \times (i_{j,\kappa} - i_\kappa^*)$. Columns (3)-(6) present estimates from the same regression with two additional regressors, the relative yield curve slope $S_j - S^*$ and curvature $C_j - C^*$. The inversion indicator $\mathbb{1}_j$ is set to 1 in months where the country-j yield curve slope is negative and remains 1 for 1-year after the end of each inversion, and zero otherwise. 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 1992:07 to 2017:12. All regressions include country fixed effects. The panel is balanced 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 15: Coefficient estimates from UIP regressions with additional yield curve inversion interactions, with inversion indicator persisting for 2-years after end of each inversion

	(1)	(2)	(3)	(4)	(5)	(6)
	UIP Regression		Augmented Regression			
	$i_\kappa - i^*_\kappa$	$\mathbb{1} \times (i_\kappa - i^*_\kappa)$	$i_\kappa - i^*_\kappa$	$\mathbb{1} \times (i_\kappa - i^*_\kappa)$	$S - S^*$	$C - C^*$
6-months	-1.59**	8.37***	-0.37	7.92***	1.14	-1.08
	(0.72)	(1.87)	(1.14)	(1.97)	(0.76)	(1.04)
12-months	-1.65***	5.77***	-0.59	5.58***	1.70	-1.15
	(0.60)	(1.21)	(0.97)	(1.34)	(1.32)	(1.77)
18-months	-1.46***	3.39***	-0.01	3.22***	3.25*	-1.85
	(0.56)	(0.86)	(0.90)	(0.96)	(1.68)	(2.10)
24-months	-1.17**	2.19***	0.35	1.98**	4.49**	-3.03
	(0.51)	(0.74)	(0.79)	(0.83)	(1.87)	(2.35)
30-months	-0.85*	1.52**	0.94	1.21*	6.50***	-5.23**
	(0.44)	(0.66)	(0.67)	(0.72)	(1.91)	(2.50)
36-months	-0.54	0.92	1.40**	0.50	8.27***	-7.62***
	(0.37)	(0.59)	(0.55)	(0.60)	(1.69)	(2.18)
42-months	-0.13	0.75	1.82***	0.23	9.60***	-10.21***
	(0.39)	(0.56)	(0.45)	(0.53)	(1.58)	(2.10)
48-months	0.37	0.34	2.15***	-0.17	9.73***	-10.67***
	(0.41)	(0.50)	(0.34)	(0.45)	(1.68)	(2.24)
54-months	0.81**	0.25	2.42***	-0.24	9.72***	-10.50***
	(0.39)	(0.43)	(0.24)	(0.37)	(1.89)	(2.29)
60-months	1.13***	0.33	2.54***	-0.12	9.39***	-9.99***
	(0.36)	(0.40)	(0.21)	(0.34)	(2.09)	(2.48)
66-months	1.38***	0.11	2.49***	-0.24	7.86***	-7.00***
	(0.31)	(0.41)	(0.22)	(0.34)	(2.24)	(2.63)
72-months	1.56***	0.01	2.35***	-0.23	5.78***	-2.67
	(0.24)	(0.43)	(0.23)	(0.36)	(2.19)	(2.58)
78-months	1.54***	0.14	2.05***	-0.01	3.97*	-0.64
	(0.22)	(0.43)	(0.22)	(0.39)	(2.11)	(2.58)
84-months	1.50***	0.23	1.84***	0.13	3.12	-0.46
	(0.21)	(0.43)	(0.21)	(0.43)	(2.05)	(2.68)
90-months	1.48***	0.32	1.72***	0.24	2.58	-0.13
	(0.21)	(0.41)	(0.21)	(0.43)	(2.03)	(2.78)
96-months	1.39***	0.22	1.56***	0.18	1.95	1.27
	(0.22)	(0.36)	(0.22)	(0.40)	(2.06)	(2.78)
102-months	1.26***	0.07	1.41***	0.02	2.24	0.69
	(0.23)	(0.32)	(0.23)	(0.36)	(2.03)	(2.66)
108-months	1.07***	0.05	1.16***	0.04	1.46	1.76
	(0.24)	(0.29)	(0.23)	(0.33)	(2.09)	(2.70)
114-months	0.90***	0.05	0.97***	0.07	0.95	3.08
	(0.23)	(0.26)	(0.22)	(0.30)	(2.23)	(3.35)
120-months	0.76***	-0.14	0.84***	-0.20	1.80	2.18
	(0.22)	(0.19)	(0.21)	(0.21)	(1.94)	(3.34)

Notes: Columns (1) and (2) present estimates from a regression the κ-period exchange rate change on the κ-period interest rate differential $i_{j,\kappa} - i^*_\kappa$, a dummy variable indicator for country-j yield curve inversions $\mathbb{1}_j$ and an interaction between the two $\mathbb{1}_j \times (i_{j,\kappa} - i^*_\kappa)$. Columns (3)-(6) present estimates from the same regression with two additional regressors, the relative yield curve slope $S_j - S^*$ and curvature $C_j - C^*$. The inversion indicator $\mathbb{1}_j$ is set to 1 in months where the country-j yield curve slope is negative and remains 1 for 2-years after the end of each inversion, and zero otherwise. 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 1992:07 to 2017:12. All regressions include country fixed effects. The panel is balanced 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.

coefficients on the slope-inversion interaction reveal that, while the tent shape relationship between exchange rates and the relative slope persists, it is weaker during inversion periods.

Sample Stability In table 17, we assess the robustness of our findings to sample period. In particular, column (6) shows that when the 1992:07-2017:12 is shortened to exclude the global financial crisis, ending in 2008:06, the interaction coefficient remains significant and of opposite sign at short horizons.

Columns (2) and (4) demonstrate significant differences in short-horizon UIP coefficients when the sample is extended to begin in 1980:01, although differences are quantitatively smaller. This weakening in the extent of the UIP switch in the 1980s may be linked to a time-varying relationship between disasters and inflation dynamics.

Table 16: Coefficient estimates from UIP regressions with additional yield curve inversion interactions

	(1) $i_\kappa - i^*_\kappa$	(2) $\mathbb{1} \times (i_\kappa - i^*_\kappa)$	(3) $S - S^*$	(4) $\mathbb{1} \times (S - S^*)$	(5) $C - C^*$	(6) $\mathbb{1} \times (C - C^*)$
6-months	0.45	4.70	1.96***	-5.70***	-1.74*	5.69***
	(1.08)	(3.83)	(0.71)	(1.71)	(1.02)	(2.10)
12-months	0.02	4.36**	2.99**	-9.02***	-2.18	11.69***
	(0.90)	(1.98)	(1.19)	(2.56)	(1.72)	(2.80)
18-months	0.45	3.09**	4.56***	-8.04***	-2.89	12.95***
	(0.83)	(1.45)	(1.49)	(3.03)	(2.08)	(3.14)
24-months	0.83	2.03*	6.05***	-7.77**	-4.39*	15.37***
	(0.76)	(1.12)	(1.75)	(3.19)	(2.35)	(4.53)
30-months	1.33**	0.98	7.98***	-6.98**	-6.47**	13.50***
	(0.67)	(0.80)	(1.82)	(2.96)	(2.56)	(4.79)
36-months	1.64***	-0.32	9.38***	-7.32**	-8.63***	10.98**
	(0.55)	(0.77)	(1.60)	(2.86)	(2.23)	(4.92)
42-months	1.96***	-0.61	10.43***	-6.24*	-10.73***	7.02
	(0.45)	(0.74)	(1.53)	(3.30)	(2.27)	(5.74)
48-months	2.18***	-1.11*	10.46***	-7.76**	-10.79***	5.53
	(0.36)	(0.58)	(1.68)	(3.19)	(2.49)	(5.57)
54-months	2.45***	-1.43***	10.70***	-10.61***	-10.86***	8.36
	(0.27)	(0.49)	(1.90)	(3.33)	(2.70)	(5.63)
60-months	2.59***	-1.11**	10.44***	-10.55***	-10.27***	9.00
	(0.23)	(0.44)	(2.10)	(3.29)	(2.92)	(5.68)
66-months	2.50***	-0.57	8.74***	-9.43***	-7.00**	9.25*
	(0.23)	(0.36)	(2.26)	(3.04)	(3.01)	(5.57)
72-months	2.31***	-0.45	6.24***	-6.53**	-2.17	3.85
	(0.23)	(0.31)	(2.21)	(2.82)	(2.96)	(5.77)
78-months	1.98***	-0.01	4.16*	-4.08	0.51	-1.27
	(0.23)	(0.34)	(2.13)	(2.68)	(2.98)	(5.95)
84-months	1.71***	0.32	3.26	-4.09	1.37	-4.63
	(0.21)	(0.37)	(2.06)	(2.60)	(3.07)	(5.76)
90-months	1.55***	0.56	2.68	-5.01*	1.81	-4.46
	(0.19)	(0.43)	(2.01)	(2.59)	(3.19)	(5.64)
96-months	1.40***	0.71*	1.86	-5.64**	3.22	-0.78
	(0.20)	(0.43)	(2.11)	(2.81)	(3.48)	(5.39)
102-months	1.29***	0.65*	2.02	-5.86*	2.83	2.73
	(0.22)	(0.38)	(2.05)	(3.20)	(3.24)	(5.85)
108-months	1.08***	0.50	1.67	-8.91***	3.17	8.52
	(0.21)	(0.35)	(2.08)	(3.15)	(3.17)	(6.09)
114-months	0.91***	0.68**	1.67	-9.26***	3.02	11.52*
	(0.20)	(0.32)	(2.11)	(2.78)	(3.44)	(5.99)
120-months	0.81***	0.42	1.95	-6.64***	2.20	10.24*
	(0.19)	(0.27)	(1.83)	(2.34)	(3.18)	(5.69)

Notes: Coefficient estimates from a regression the κ-period exchange rate change on the κ-period interest rate differential $i_{j,\kappa} - i^*_\kappa$, the relative yield curve slope $S_j - S^*$, the relative yield curve curvature $C_j - C^*$, a dummy variable indicator for country-j yield curve inversions $\mathbb{1}_j$ and interactions between the inversion indicator and other regressors. The inversion indicator $\mathbb{1}_j$ is set to 1 in months where the country-j yield curve slope is negative, and zero otherwise. 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 1992:07 to 2017:12. All regressions include country fixed effects. The panel is balanced 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 17: Coefficient estimates from UIP regressions with additional yield curve inversion interaction

Sample	(1)	(2)	(3)	(4)	(5)	(6)
	1980:01-2017:12		1980:01-2008:06		1992:07-2008:06	
	$i_\kappa - i^*_\kappa$	$\mathbb{1} \times (i_\kappa - i^*_\kappa)$	$i_\kappa - i^*_\kappa$	$\mathbb{1} \times (i_\kappa - i^*_\kappa)$	$i_\kappa - i^*_\kappa$	$\mathbb{1} \times (i_\kappa - i^*_\kappa)$
6-months	-1.87***	1.76*	-2.77***	2.21**	-2.93***	10.08***
	(0.58)	(1.05)	(0.62)	(0.94)	(0.77)	(1.58)
12-months	-1.59***	1.66**	-2.50***	2.30***	-3.28***	9.32***
	(0.50)	(0.78)	(0.52)	(0.74)	(0.59)	(0.88)
18-months	-1.21**	0.86	-2.09***	1.45**	-3.00***	6.15***
	(0.48)	(0.67)	(0.50)	(0.65)	(0.57)	(0.77)
24-months	-0.83*	0.25	-1.66***	0.79	-2.62***	4.18***
	(0.45)	(0.57)	(0.45)	(0.54)	(0.51)	(0.66)
30-months	-0.53	-0.08	-1.32***	0.42	-2.16***	2.98***
	(0.41)	(0.52)	(0.40)	(0.49)	(0.40)	(0.54)
36-months	-0.22	-0.25	-0.98***	0.20	-1.76***	1.95***
	(0.38)	(0.53)	(0.36)	(0.50)	(0.34)	(0.55)
42-months	0.08	-0.24	-0.66*	0.20	-1.31***	1.69***
	(0.37)	(0.50)	(0.35)	(0.46)	(0.38)	(0.56)
48-months	0.39	-0.25	-0.30	0.16	-0.73*	1.30**
	(0.36)	(0.47)	(0.34)	(0.43)	(0.43)	(0.52)
54-months	0.72**	-0.24	0.14	0.12	-0.07	0.82*
	(0.33)	(0.42)	(0.32)	(0.39)	(0.45)	(0.49)
60-months	0.94***	-0.12	0.45	0.17	0.43	0.60
	(0.30)	(0.37)	(0.31)	(0.35)	(0.44)	(0.47)
66-months	1.14***	-0.11	0.75***	0.12	0.81**	0.55
	(0.27)	(0.32)	(0.29)	(0.31)	(0.38)	(0.40)
72-months	1.32***	-0.20	1.03***	0.00	1.16***	0.28
	(0.23)	(0.25)	(0.25)	(0.24)	(0.31)	(0.39)
78-months	1.37***	-0.24	1.18***	-0.12	1.29***	0.31
	(0.20)	(0.20)	(0.22)	(0.19)	(0.29)	(0.40)
84-months	1.32***	-0.26	1.23***	-0.20	1.42***	0.57
	(0.18)	(0.17)	(0.20)	(0.16)	(0.28)	(0.47)
90-months	1.25***	-0.27*	1.21***	-0.23	1.49***	0.79
	(0.17)	(0.15)	(0.18)	(0.15)	(0.25)	(0.49)
96-months	1.10***	-0.21	1.08***	-0.20	1.42***	0.78*
	(0.17)	(0.14)	(0.18)	(0.14)	(0.24)	(0.44)
102-months	0.91***	-0.07	0.90***	-0.07	1.29***	0.59
	(0.18)	(0.14)	(0.18)	(0.14)	(0.23)	(0.38)
108-months	0.75***	0.04	0.75***	0.03	1.10***	0.44
	(0.18)	(0.13)	(0.18)	(0.13)	(0.22)	(0.36)
114-months	0.65***	0.13	0.65***	0.13	0.90***	0.54**
	(0.17)	(0.12)	(0.17)	(0.12)	(0.22)	(0.27)
120-months	0.63***	0.05	0.63***	0.05	0.77***	0.21
	(0.17)	(0.12)	(0.17)	(0.12)	(0.19)	(0.24)

Notes: Coefficient estimates from a regression the κ-period exchange rate change on the κ-period interest rate differential $i_{j,\kappa} - i^*_\kappa$, a dummy variable indicator for country-j yield curve inversions $\mathbb{1}_j$ and interactions between the two. The inversion indicator $\mathbb{1}_j$ is set to 1 in months where the country-j yield curve slope is negative, and zero otherwise. Regressions are estimated using pooled end-of-month data for six currencies—AUD, CAD, CHF, EUR, JPY and GBP—*vis-à-vis* the USD. All regressions include country fixed effects. Coefficients in columns (1) and (2) are estimated on a 1980:01-2017:12 sample, an unbalanced panel. Coefficients in columns (3) and (4) are estimated on a 1980:01-2008:06 sample, an unbalanced panel. Coefficients in columns (5) and (6) are estimated on a 1992:07-2008:06 sample, a balanced panel. Standard errors (reported in parentheses) are constructed according to the Driscoll and Kraay (1998) methodology. *, ** and *** denote statistically significant point estimates at 10%, 5% and 1% significance levels, respectively.

C Additional Derivations

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

Use the definition of the κ-period *ex post* ERRP from the perspective of the Home agent (8). From (6), this is equal to $\lambda^H_{t,\kappa} = -\text{cov}_t\left(m_{t,t+\kappa}, \Delta^\kappa e_{t+\kappa}\right)$.

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

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

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 (9) 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 (10) 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

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

This expression captures the intuition that, under complete markets, the ERRP is determined by the relative variance of countries' pricing kernels.

Second, the expression for the slope of a given yield curve (14), can be re-expressed given the AR(1) (log) pricing kernel (C.2.1).[50] 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. The expression captures the intuition that the slope of the yield curve reflects the autocovariance of a representative investor's intertemporal consumption valuation, albeit simplified with an AR(1) specification of pricing kernels.

[50] In this derivation, we ignore the Jensen's inequality term for $-\frac{1}{2}\mathrm{var}_t\left(p_{t+1,n-1}\right)$ in (14).

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_t^R

$$S_t^R \equiv S_t - S_t^* = \left(1 - \rho_\nu^{(n-1)}\right) \text{var}_t\left(\nu_{t+1}\right) - \left(1 - \rho_\nu^{*(n-1)}\right) \text{var}_t\left(\nu_{t+1}^*\right) \tag{C.2.5}$$

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

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

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:

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

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 (10), the *ex post* κ-period ERRP under complete markets can be written as:

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

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 final expression captures the intuition that, under complete markets, the ERRP is determined by the relative variance of countries' pricing kernels.

The expression for the slope of a given yield curve (14), 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:

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

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 expression captures the intuition that the slope of the yield curve reflects the autocovariance of a representative investor's intertemporal consumption valuation with an AR(2) specification of pricing kernels.

The relative slope under symmetry is therefore

$$S_t^R \equiv S_t - S_t^* = (1 - \psi_{n-1}) \left[\text{var}_t\left(\nu_{t+1}\right) - \text{var}_t\left(\nu_{t+1}^*\right)\right] \tag{C.3.6}$$

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:

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

where, when $\psi_{n-1} \in (0, 1)$ and $\kappa, n > 1$, $\psi_{\kappa-1}^2/(1 - \psi_{n-1}) > 0$, 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.4 Further Derivations for Section 4.2

Consider a decomposition of the Home (Foreign) SDF $M_{t,t+\kappa}^{(*)}$ into a valuation of pecuniary returns $M_{t,t+\kappa}^{r(*)}$ and a valuation for liquidity $M_{t,t+\kappa}^{\ell(*)}$:

$$M_{t,t+\kappa}^{(*)} = M_{t,t+\kappa}^{r(*)} M_{t,t+\kappa}^{\ell(*)}, \tag{33}$$

where $M_{t,t+\kappa}^{(*)} = M_{t,t+\kappa}^{r(*)}$ if there is no role for liquidity. $M_{t,t+\kappa}^{\ell(*)}$ captures the price of liquidity at time t, which is both investor-specific and asset-specific. In a similar way that we consider return valuation to vary across time due to a stochastic sequence of consumption or wealth, we can assume the investor receives a stochastic endowment of liquidity. First, the liquidity component of the pricing kernel is higher in periods when the investor is liquidity constrained. Second, the pricing kernel of a given investor can differ across assets now as they vary in the liquidity services they provide. For example, a Home representative investor prices Foreign (US) bonds according to $M_t^{\ell(\$)}$ which is not necessarily equal to that when the same investor prices domestic bonds M_t^{ℓ}. However, the SDF can be rewritten such that it applies to any asset by separating the asset-specific liquidity component—i.e. for a Home investor purchasing Home and US bonds, respectively:

$$e^{\xi_{t,\kappa}} = \mathbb{E}_t[\hat{M}_{t,t+\kappa} R_{t,\kappa}], \qquad e^{\xi_{t,\kappa}^{\$}} = \mathbb{E}_t[\hat{M}_{t,t+\kappa} \frac{\mathcal{E}_{t+\kappa}}{\mathcal{E}_t} R_{t,\kappa}^{\$}]$$

which imply

$$\mathbb{E}_t[\Delta^\kappa e_{t+\kappa}] + \frac{1}{2}\text{var}_t(\Delta e_{t+\kappa}) = (i_{t,\kappa} - i_{t,\kappa}^{\$}) - \text{cov}_t(m_{t,t+\kappa}, \Delta^\kappa e_{t+\kappa}) - (\xi_{t,\kappa} - \xi_{t,\kappa}^{\$})$$

which, for a given SDF, requires either a Home expected appreciation (expected depreciation of dollar) or a fall in the US yield at a range of horizons (maturities) κ—in line with our empirical evidence. Note that this is consistent with a contemporaneous appreciation of the US dollar.

References

Alvarez, F. and U. J. Jermann (2005): "Using Asset Prices to Measure the Persistence of the Marginal Utility of Wealth," *Econometrica*, 73, 1977–2016.

Anderson, N. and J. Sleath (2001): "New estimates of the UK real and nominal yield curves," Bank of England Staff Working Paper 126, Bank of England.

Ang, A. and J. S. Chen (2010): "Yield Curve Predictors of Foreign Exchange Returns," Working paper, Columbia University.

Backus, D., N. Boyarchenko, and M. Chernov (2018): "Term structures of asset prices and returns," *Journal of Financial Economics*, 129, 1–23.

Backus, D. K., S. Foresi, and C. I. Telmer (2001): "Affine Term Structure Models and the Forward Premium Anomaly," *Journal of Finance*, 56, 279–304.

Bansal, R. and I. Shaliastovich (2013): "A Long-Run Risks Explanation of Predictability Puzzles in Bond and Currency Markets," *Review of Financial Studies*, 59, 1481–1509.

Benigno, G., P. Benigno, and S. Nisticò (2012): "Risk, Monetary Policy, and the Exchange Rate," *NBER Macroeconomics Annual*, 26, 247–309.

Brunnermeier, M. K., S. Nagel, and L. H. Pedersen (2009): "Carry Trades and Currency Crashes," in *NBER Macroeconomics Annual 2008, Volume 23*, National Bureau of Economic Research, Inc, NBER Chapters, 313–347.

Bussière, M., M. D. Chinn, L. Ferrara, and J. Heipertz (2018): "The New Fama Puzzle," NBER Working Papers 24342, National Bureau of Economic Research, Inc.

Ca' Zorzi, M. and E. A. Marin (2018): "Exchange rate premia in imperfect financial markets," European Central Bank and University of Cambridge.

Campbell, J. and R. Shiller (1991): "Yield Spreads and Interest Rate Movements: A Bird's Eye View," *Review of Economic Studies*, 58, 495–514.

Campbell, J. Y. and J. Cochrane (1999): "Force of Habit: A Consumption-Based Explanation of Aggregate Stock Market Behavior," *Journal of Political Economy*, 107, 205–251.

Chen, Y.-c. and K. P. Tsang (2013): "What Does the Yield Curve Tell Us about Exchange Rate Predictability?" *The Review of Economics and Statistics*, 95, 185–205.

Chinn, M. D. and G. Meredith (2005): "Testing Uncovered Interest Parity at Short and Long Horizons during the Post-Bretton Woods Era," NBER Working Papers 11077, National Bureau of Economic Research, Inc.

Chinn, M. D. and S. Quayyum (2012): "Long Horizon Uncovered Interest Parity Re-Assessed," NBER Working Papers 18482, National Bureau of Economic Research, Inc.

Chrétien, S. (2012): "Bounds on the autocorrelation of admissible stochastic discount factors," *Journal of Banking & Finance*, 36, 1943–1962.

Cochrane, J. H. and M. Piazzesi (2005): "Bond Risk Premia," *American Economic Review*, 95, 138–160.

Colacito, R., S. J. Riddiough, and L. Sarno (2019): "Business Cycles and Currency Returns," NBER Working Papers 26299, National Bureau of Economic Research, Inc.

Cujean, J. and M. Hasler (2017): "Why Does Return Predictability Concentrate in Bad Times?" *Journal of Finance*, 72, 2717–2758.

Diebold, F. X. and G. D. Rudebusch (2013): *Yield Curve Modeling and Forecasting: The Dynamic Nelson-Siegel Approach*, Princeton University Press.

Driscoll, J. C. and A. C. Kraay (1998): "Consistent Covariance Matrix Estimation With Spatially Dependent Panel Data," *The Review of Economics and Statistics*, 80, 549–560.

Du, W., J. Im, and J. Schreger (2018): "The U.S. Treasury Premium," *Journal of International Economics*, 112, 167–181.

Engel, C. (2014): *Exchange Rates and Interest Parity*, Elsevier, vol. 4 of *Handbook of International Economics*, chap. 0, 453–522.

——— (2016): "Exchange Rates, Interest Rates, and the Risk Premium," *American Economic Review*, 106, 436–474.

Engel, C. and S. P. Y. Wu (2018): "Liquidity and Exchange Rates: An Empirical Investigation," NBER Working Papers 25397, National Bureau of Economic Research, Inc.

Epstein, L. G. and S. E. Zin (1989): "Substitution, Risk Aversion, and the Temporal Behavior of Consumption and Asset Returns: A Theoretical Framework," *Econometrica*, 57, 937–969.

Fama, E. and R. R. Bliss (1987): "The Information in Long-Maturity Forward Rates," *American Economic Review*, 77, 680–92.

Fama, E. F. (1984): "Forward and spot exchange rates," *Journal of Monetary Economics*, 14, 319–338.

Farhi, E., S. Fraiberger, X. Gabaix, R. Ranciere, and A. Verdelhan (2015): "Crash Risk in Currency Markets," Working Paper 20948, Harvard University OpenScholar.

Farhi, E. and X. Gabaix (2016): "Rare Disasters and Exchange Rates," *The Quarterly Journal of Economics*, 131, 1–52.

Gabaix, X. (2012): "Variable Rare Disasters: An Exactly Solved Framework for Ten Puzzles in Macro-Finance," *The Quarterly Journal of Economics*, 127, 645–700.

Gourinchas, P.-O., W. Ray, and D. Vayanos (2019): "A Preferred-Habitat Model of Term Premia and Currency Risk," Mimeo.

Gräb, J. and T. Kostka (2018): "Predicting risk premia in short-term interest rates and exchange rates," Working Paper Series 2131, European Central Bank.

Grasso, A. and F. Natoli (2018): "Consumption volatility risk and the inversion of the yield curve," Working Paper Series 2141, European Central Bank.

Greenwood, R., S. G. Hanson, J. C. Stein, and A. Sunderam (2019): "A Quantity-Driven Theory of Term Premiums and Exchange Rates," Working paper, Harvard Business School.

Gürkaynak, R. S., B. Sack, and J. H. Wright (2007): "The U.S. Treasury yield curve: 1961 to the present," *Journal of Monetary Economics*, 54, 2291–2304.

Hansen, L. and R. Hodrick (1980): "Forward Exchange Rates as Optimal Predictors of Future Spot Rates: An Econometric Analysis," *Journal of Political Economy*, 88, 829–53.

Hassan, T. A. and R. C. Mano (2019): "Forward and Spot Exchange Rates in a Multi-Currency World," *The Quarterly Journal of Economics*, 134, 397–450.

Jiang, Z., A. Krishnamurthy, and H. Lustig (2018): "Foreign Safe Asset Demand for US Treasurys and the Dollar," *AEA Papers and Proceedings*, 108, 537–541.

Jordà, O. and A. M. Taylor (2012): "The carry trade and fundamentals: Nothing to fear but FEER itself," *Journal of International Economics*, 88, 74–90.

Litterman, R. B. and J. Scheinkman (1991): "Common Factors Affecting Bond Returns," *The Journal of Fixed Income*, 1, 54–61.

Lustig, H., A. Stathopoulos, and A. Verdelhan (2019): "The Term Structure of Currency Carry Trade Risk Premia," *American Economic Review*, Forthcoming.

Lustig, H. and A. Verdelhan (2019): "Does Incomplete Spanning in International Financial Markets Help to Explain Exchange Rates?" *American Economic Review*, forthcoming.

Moon, R., A. Rubia, and R. Valkanov (2004): "Long-Horizon Regressions When the Predictor is Slowly Varying," Working paper, University of California, San Diego.

Newey, W. K. and K. D. West (1987): "A Simple, Positive Semi-definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix," *Econometrica*, 55, 703–08.

Piazzesi, M. and M. Schneider (2007): "Equilibrium Yield Curves," in *NBER Macroeconomics Annual 2006, Volume 21*, National Bureau of Economic Research, Inc, NBER Chapters, 389–472.

Stavrakeva, V. and J. Tang (2019): "The Dollar During the Great Recession: US Monetary Policy Signaling and The Flight To Safety," Working paper, London Business School.

Valkanov, R. (2003): "Long-horizon regressions: theoretical results and applications," *Journal of Financial Economics*, 68, 201–232.

Verdelhan, A. (2010): "A Habit-Based Explanation of the Exchange Rate Risk Premium," *Journal of Finance*, 65, 123–146.

Wachter, J. A. (2006): "A consumption-based model of the term structure of interest rates," *Journal of Financial Economics*, 79, 365–399.

Wright, J. H. (2011): "Term Premia and Inflation Uncertainty: Empirical Evidence from an International Panel Dataset," *American Economic Review*, 101, 1514–1534.