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Abstract

We study the properties of foreign exchange risk premia that can explain the forward bias puzzle – the tendency of high-interest rate currencies to appreciate rather than depreciate. These risk premia arise endogenously from imposing the no-arbitrage condition on the relation between the term structure of interest rates and exchange rates, and they compensate for both currency risk and interest rate risk. In our empirical analysis, we estimate risk premia using an affine multi-currency term structure model and find that model-implied risk premia yield unbiased predictions for exchange rate excess returns. While interest rate risk affects the level of risk premia, the time-variation in excess returns is almost entirely driven by currency risk. Furthermore, risk premia are (i) closely related to global risk aversion, (ii) countercyclical to the state of the economy, and (iii) tightly linked to traditional exchange rate fundamentals.

JEL classification: F31; E43; G10.

Keywords: term structure; exchange rates; forward bias; predictability.

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1 Introduction

Uncovered interest rate parity (UIP) postulates that the expected exchange rate change must equal the interest rate differential or (because covered interest parity holds) the forward premium. UIP also forms the economic foundation for the forward unbiasedness hypothesis (FUH) stating that the forward exchange rate should be an unbiased predictor of the future spot rate. The empirical observation that there is a negative association between forward premia and subsequent exchange rate returns, first noted in Hansen and Hodrick (1980), Bilson (1981), and Fama (1984), implies a rejection of UIP and the FUH. This stylized fact is often termed the 'forward bias puzzle'. A large literature has argued that risk premia must be at the heart of this observation.

In this paper, we re-examine the relationship between the term structure of interest rates and exchange rates by deriving the relation between forward and spot exchange rates from the principle of no-arbitrage without assuming risk neutrality. This setting implies that the forward exchange rate is the sum of the expected spot rate plus a time-varying risk premium which compensates both for currency risk and interest rate risk. Thus, our first contribution of the paper is to show that, in contrast to the FUH, forward rates are generally biased predictors of future spot exchange rates under no-arbitrage. As a consequence, expected spot rate changes are also not only determined by the forward premium as postulated by UIP, but additionally comprise a time-varying risk premium. We refer to these general, model-free relations that extend the conventional FUH and UIP in that they are free of risk preferences and consistent with no-arbitrage as the 'risk-adjusted FUH' (RA-FUH) and as 'risk-adjusted UIP' (RA-UIP).

To work with the RA-UIP condition empirically, we put structure on the international financial market with a model for interest rate risk and currency risk. We use an affine multi-economy term structure model that relates two countries' pricing kernels such that arbitrage-free pricing is ensured. We employ latent factors to model the uncertainty underlying the international economy for two reasons. First, this approach gives us maximum flexibility with respect to the statistical framework even with a relatively small number of factors. Second, we do not have to rely on exogenous observable variables driving the economy which are available only at low frequencies.[1] The design of our model follows the pioneering work of Backus, Foresi, and Telmer (2001) but is more general in that it accounts for interest rate risk arising from fluctuations in the bond market over multiple periods. It also accommodates the findings of Brennan and Xia (2006) and extends their work in that we do not approximate the risk premium but derive the term structure of foreign exchange risk premia in closed form. This allows us to not only jointly match the term structures of interest rates but simultaneously the term structure of foreign exchange risk premia in the estimation procedure. Using a data set that comprises daily observations for six major US dollar exchange rates over the last 20 years, we generate model-implied exchange rate expectations and risk premia for horizons ranging from 1 day to 4 years.

Our second contribution is a battery of empirical results which relate to the properties of the model-implied risk premia. The empirical results suggest that our model is capable of identifying time-varying risk premia and that observed exchange rate behavior complies with RA-UIP. We find that model expectations and risk premia fulfill the two conditions established by Fama (1984) such that the omission of the risk premium in conventional UIP tests results in a forward bias. We then show that our RA-UIP model - in contrast to UIP - generates unbiased predictions for exchange rate excess returns. This implies that accounting for risk premia can be sufficient to resolve the forward bias puzzle without additionally requiring departures from rational expectations. We also perform a variety of predictive ability tests which, on the one hand, complement evidence that excess returns are predictable, and, on the other hand, further confirm that the RA-UIP model fits the data substantially better than UIP and also better than a random walk. Finally, we decompose the risk premium, and show that although there is a compensation for interest rate risk, deviations from UIP and hence foreign exchange excess returns can almost entirely be explained by the premium for currency risk.

[1]Such economic variables are typically available at quarterly or at best at monthly frequency. In our context this is not feasible, as we are also interested in short horizons such as 1 day or 1 week, and our model estimation is hence based on daily data. However, as discussed below, we relate the model-implied risk premia to observable economic variables later in the paper to refine our understanding of the drivers of the latent factors.

We also provide empirical evidence that risk premia are closely linked to economic variables that proxy for global risk, the US business cycle, and traditional exchange rate fundamentals. The results suggest that expected excess returns reflect flight-to-quality and flight-to-liquidity considerations. Expected excess returns also depend on macroeconomic variables (e.g. output growth, money supply growth, consumption growth) in a way that risk premia in dollar exchange rates are countercyclical to the US economy. Moreover, a large part of expected excess returns can be explained by fundamentals deemed relevant in traditional exchange rate models.

Related Literature in More Detail There is a large literature documenting deviations from UIP and studying the forward bias puzzle, starting from Hansen and Hodrick (1980), Bilson (1981), and Fama (1984). Our paper offers a risk-based explanation by contributing to the literature that investigates the puzzle in the light of interest rate risk and no-arbitrage.[2]

Earlier papers that study the link between interest rates and exchange rates with term structure factor models include Nielsen and Saá-Requejo (1993), Saá-Requejo (1994), Bakshi and Chen (1997), and Bansal (1997). A pioneering paper is Backus, Foresi, and Telmer (2001), who adapt modern (affine) term structure theory to a multi-economy setting. They establish important theoretical relations that must hold in the absence of arbitrage between the pricing kernels and the exchange rate driving the international economy. In their discrete-time one-period setting, they can replicate the puzzle un-

[2]There are many other papers that try to shed light on the puzzle from other angles than relating the term structure of interest rates of two countries and their exchange rate. Explanations that build on risk premium arguments - based, among others, on equilibrium models or consumption-based asset pricing - include Frankel and Engel (1984), Domowitz and Hakkio (1985), Hodrick (1987), Cumby (1988), Mark (1988), Backus, Gregory, and Telmer (1993), Bekaert and Hodrick (1993), Bansal, Gallant, Hussey, and Tauchen (1995), Bekaert (1996), Bekaert, Hodrick, and Marshall (1997), Lustig and Verdelhan (2007), Brunnermeier, Nagel, and Pedersen (2008), Farhi and Gabaix (2008), Jurek (2008), Lustig, Roussanov, and Verdelhan (2008), Verdelhan (2008), Bansal and Shaliastovich (2009), and Farhi, Fraiberger, Gabaix, Ranciere, and Verdelhan (2009). Other recent papers look at the puzzle, for instance, in the context of incomplete information processing, e.g. Bacchetta and van Wincoop (2009), differences in developed versus emerging markets, e.g. Bansal and Dahlquist (2000) and Frankel and Poonawala (2007), and the profitability and economic value of currency speculation, e.g. Burnside, Eichenbaum, Kleshchelski, and Rebelo (2006), and Della Corte, Sarno, and Tsiakas (2009).

der the following two alternative specifications: either, there is a common-idiosyncratic factor structure and interest rates take on negative values with positive probabilities, or, global factors and state variables have asymmetric effects on state prices in different countries. Motivated by the latter, related empirical studies, Dewachter and Maes (2001), Ahn (2004), Inci and Lu (2004), Mosburger and Schneider (2005), and Anderson, Hammond, and Ramezani (2009) elaborate on the effects of local versus global factors in an international economy.[3] Brandt and Santa-Clara (2002) and Anderson, Hammond, and Ramezani (2009) extend affine multi-country term structure models to account for market incompleteness to investigate exchange rate excess volatility.

Brennan and Xia (2006) investigate the relations between the foreign exchange risk premium, exchange rate volatility, and the volatilities of the pricing kernels for the underlying currencies, under the assumption of integrated capital markets. The continuous-time model proposed by Brennan and Xia (2006) jointly determines the term structure of interest rates and an approximation of the risk premium in a no-arbitrage setting. Their analysis suggests that the volatility of exchange rates is associated with the estimated volatility of the relevant pricing kernels, and risk premia are significantly related to both the estimated volatility of the pricing kernels and the volatility of exchange rates. The estimated risk premia mostly satisfy the Fama (1984) necessary conditions for explaining the forward bias puzzle, although the puzzle remains in several cases.

Our modeling setup follows Backus, Foresi, and Telmer (2001) and Brennan and Xia (2006) and extends their work in that our model incorporates interest rate risk and allows to derive the term structure of foreign exchange risk premia in closed form. These theoretical extensions enable us to simultaneously estimate the term structures of interest rates as well as the term structure of foreign exchange risk premia. Our results reveal that the model can match the empirical properties of foreign exchange risk premia more accurately than previous research. In particular, we match the pattern of excess return predictability and explain (a large fraction of) the forward bias puzzle.

[3] Another recent related article is Leippold and Wu (2007). Instead of using an affine model, they propose a class of multi-currency quadratic models with a factor structure in the pricing kernel of each economy.

Additionally, we provide evidence that expected excess returns are (i) related to global risk aversion consistent with the the flight-to-quality and flight-to-liquidity arguments in Lustig, Roussanov, and Verdelhan (2008) and Brunnermeier, Nagel, and Pedersen (2008), (ii) countercyclical to the state of the US economy in line with e.g. Lustig and Verdelhan (2007) and De Santis and Fornari (2008), and (iii) driven by traditional exchange rate fundamentals relevant in monetary models as in Engel and West (2005).

The remainder of the paper is set out as follows. Section 2 discusses the link between interest rates and exchange rates in light of previous literature and elaborates the relationship between forward and expected spot rates implied by no-arbitrage. We describe the empirical model, the estimation procedure and the criteria applied to evaluate RA-UIP in Section 3, and we present the results in Section 4. Section 5 presents empirical evidence that financial and macroeconomic variables are important drivers of the foreign exchange risk premium. Section 6 concludes. Appendices A to D provide some technical details related to the derivation of RA-UIP and RA-FUH, the econometric model, the estimation procedure, and the bootstrap techniques applied.

2 Exchange Rates, Interest Rates and No-Arbitrage Conditions

This section defines the basic relationships linking exchange rates and interest rates, and shows the implications of imposing the no-arbitrage condition in this context. This results in the risk-aversion variants of UIP and FUH, which are shown to imply intuitive properties for the foreign exchange risk premium.

2.1 Covered Interest Rate Parity

We start from the convention that the spot exchange rate S_t is expressed as units of domestic currency per unit of foreign currency. We denote by $F_{t,T}$ the forward exchange rate at time t for an exchange of currencies at time $T > t$. The time t prices of zero bonds with maturity T denominated in domestic and foreign currency are $p_{t,T}$ and $p^{\star}_{t,T}$, respectively. With the US dollar (USD) as the domestic currency and the British pound (GBP) as the foreign currency, we have that 1 USD denominated in GBP and put into a GBP account at time t will be worth $\frac{1}{S_t p^\star_{t,T}}$ GBP at maturity T. 1 USD put into a US account at time t will be worth $1/p_{t,T}$ USD at T. Entering a forward contract at time t to exchange USD for GBP at time T will yield $\frac{1}{p_{t,T} F_{t,T}}$ GBP at maturity. In the absence of arbitrage we may infer the Covered Interest Rate Parity (CIP) relation

$$F_{t,T} = \frac{p^\star_{t,T} S_t}{p_{t,T}}. \tag{1}$$

Let $y_{t,T}$ and $y^\star_{t,T}$ denote the T-period zero-yields of $p_{t,T}$ and $p^\star_{t,T}$, i.e. $y_{t,T} \equiv -\log p_{t,T}$ and $y^\star_{t,T} \equiv -\log p^\star_{t,T}$. The CIP condition can then be written as

$$\begin{aligned} F_{t,T} &= S_t \exp[y_{t,T} - y^\star_{t,T}] \\ f_{t,T} - s_t &= y_{t,T} - y^\star_{t,T} \end{aligned} \tag{2}$$

where $s_t \equiv \log S_t$ and $f_{t,T} \equiv \log F_{t,T}$.

2.2 Uncovered Interest Parity and Forward Unbiasedness Hypothesis

Assuming risk-neutrality and rational expectations, Uncovered Interest Rate Parity (UIP) postulates that

$$\mathbb{E}^{\mathbb{P}}_t[s_T] - s_t = y_{t,T} - y^\star_{t,T}, \tag{3}$$

where $\mathbb{E}^{\mathbb{P}}_t$ denotes the expectation conditional on time t information under the physical probability measure. UIP also forms the economic foundation for the Forward Unbiasedness Hypothesis (FUH): comparing eqs. (2) and (3) yields that the forward rate should be an unbiased predictor of the future spot exchange rate

$$\mathbb{E}^{\mathbb{P}}_t[s_T] = f_{t,T}. \tag{4}$$

Empirical tests of UIP are usually performed by estimating the 'Fama regression' (Fama, 1984)

$$\Delta s_{t,T} = \alpha + \beta(y_{t,T} - y^{\star}_{t,T}) + \eta_{t,T}, \tag{5}$$

where $\Delta s_{t,T} = s_T - s_t$. The null hypothesis that the yield differential (or equivalently the forward premium $f_{t,T} - s_t$) is an unbiased predictor of future exchange rate changes holds if $\alpha = 0$, $\beta = 1$, and $\eta_{t,T}$ is serially uncorrelated. If UIP holds, this also implies that excess returns are unpredictable. Excess returns, defined as the return of a long forward position in the foreign currency, $rx_{t,T} \equiv s_T - f_{t,T}$, can be represented in the following regression by subtracting the yield differential from both sides in eq. (5)

$$rx_{t,T} = \alpha + \underbrace{(\beta - 1)}_{\gamma}(y_{t,T} - y^{\star}_{t,T}) + \eta_{t,T}. \tag{6}$$

As eq. (6) contains the same information as eq. (5), the intercepts and the residuals are the same and $\gamma = \beta - 1$. Accordingly, the null hypothesis of unpredictable excess returns is given by $\alpha = 0$, $\gamma = 0$, and $\eta_{t,T}$ being serially uncorrelated.

Empirical research has consistently rejected UIP; for surveys see Hodrick (1987), Froot and Thaler (1990), Engel (1996). It is now considered a stylized fact that estimates of β are closer to minus unity than plus unity, implying that higher interest rate currencies tend to appreciate when UIP predicts them to depreciate. This finding is commonly referred to as the 'forward bias puzzle'. Empirical research also provides evidence that excess returns are predictable on the basis of the lagged interest rate differential or the forward premium.

Fama (1984) argues that the forward bias may be caused by a time-varying risk premium that eqs. (5) and (6) do not account for. Denoting by $\lambda_{t,T}$ the risk premium eq. (4) then reads

$$\mathbb{E}^{\mathbb{P}}_t[s_T] + \lambda_{t,T} = f_{t,T}. \tag{7}$$

Subtracting s_t from both sides, the expected change in the exchange rate is

$$\mathbb{E}^{\mathbb{P}}_t[s_T] - s_t = (f_{t,T} - s_t) - \lambda_{t,T}. \tag{8}$$

From the theory of linear regressions we know that, if a time varying risk premium exists, its omission in the Fama regression (5) results in a negative β estimate if the following conditions are satisfied

$$\begin{aligned} &\mathbb{C}ov^{\mathbb{P}}\left[\lambda_{t,T}, \mathbb{E}_t^{\mathbb{P}}[s_T - s_t]\right] < 0 \\ &\left|\mathbb{C}ov^{\mathbb{P}}\left[\lambda_{t,T}, \mathbb{E}_t^{\mathbb{P}}[s_T - s_t]\right]\right| > \mathbb{V}^{\mathbb{P}}\left[\mathbb{E}_t^{\mathbb{P}}[s_T - s_t]\right]. \end{aligned} \tag{9}$$

In other words, eq. (9) states that if a time-varying risk premium exists which exhibits negative covariance with expected exchange rate changes (first condition) and the absolute value of this covariance is greater than the variance of expected changes (second condition), the omission of the risk premium leads to a negative β estimate in the Fama regression. However, attempts to explain the forward bias puzzle using risk premia have only met with limited success so far.

2.3 Risk-Adjusted UIP and FUH under No-Arbitrage

The FUH states that the forward exchange rate constitutes the expectation of the underlying spot exchange rate it is written on. As in eq. (4), the expectation is usually taken under the assumption of risk neutrality. To relax this assumption, we derive the relationship between spot and forward exchange rates from the principles of no-arbitrage. As we show below, this allows us to formulate risk-adjusted counterparts to the conventional UIP and FUH that endogenize time-varying risk premia in the spirit of Fama (1984).

The CIP eq. (1) reveals that the connection between spot and forward exchange rates is determined by the zero-coupon bonds in the respective currencies. With this obvious exposure to interest rate risk, we investigate the relationship between spot and forward exchange rates in a no-arbitrage setting in which we allow interest rates to vary stochastically. This feature merely reflects the observation that the price of the forward contract changes over time due to both spot rate and interest rate fluctuations.

The natural starting point for our derivation is the Fundamental Theorem of Asset Pricing (FTAP), which states that claims deflated by a traded asset as a numeraire are martingales under specific probability measures. For a claim Π payable at time $T \geq t$ with

N a traded asset serving as numeraire, the FTAP establishes that $\Pi_t/N_t = \mathbb{E}_t^{\mathbb{Q}_N}[\Pi_T/N_T]$ where $\mathbb{Q}_N$ is the probability measure induced by the numeraire N. In our application, it is useful to choose $p_{t,T}$ as the numeraire; see for example Björk (2004, p. 355ff) , or Mele (2009, p. 242ff). The associated probability measure is the T-forward measure $\mathbb{Q}_{\mathbb{T}}$ and the pricing equation is given by

$$\Pi_t = p_{t,T}\mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}[\Pi_T] . \tag{10}$$

Consider a long position in a standard forward contract that pays off $S_T - F_{t,T}$ at time T. Since the initial value is zero by convention, applying eq. (10) leads to

$$\begin{aligned} 0 &= p_{t,T}\mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}[S_T - F_{t,T}] \text{, and therefore} \\ F_{t,T} &= \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}[S_T] . \end{aligned} \tag{11}$$

Hence, under no-arbitrage the forward rate is the expected spot rate under the T-forward measure $\mathbb{Q}_{\mathbb{T}}$ and in general not under the risk neutral measure $\mathbb{Q}$ associated with the bank account B_t.[4] To see this in detail, consider a long forward position again. With B_t solving $dB_t = r_t dt$, where r is the corresponding short rate, the price of a domestic T-period zero bond is given by $p_{t,T} = \mathbb{E}_t^{\mathbb{Q}}\left[e^{-\int_t^T r(u)du}\right]$. Applying the FTAP yields

$$\begin{aligned} 0 &= \mathbb{E}_t^{\mathbb{Q}}\left[e^{-\int_t^T r(u)du}(S_T - F_{t,T})\right] \\ F_{t,T} &= \mathbb{E}_t^{\mathbb{Q}}\left[\frac{e^{-\int_t^T r(u)du}}{p_{t,T}} S_T\right] = \mathbb{E}_t^{\mathbb{Q}}[\Omega_T S_T] , \end{aligned} \tag{12}$$

where $\Omega_T = \frac{e^{-\int_t^T r(u)du}}{p_{t,T}}$. If the short rate process is deterministic, $\Omega_T = 1$ and $\mathbb{Q}_{\mathbb{T}}$ and $\mathbb{Q}$ are the same. If the short-rate process is stochastic, Ω_T represents the change of numeraire from B_t to $p_{t,T}$ and corresponding measures $\mathbb{Q}$ to $\mathbb{Q}_{\mathbb{T}}$, i.e. $\Omega_T \equiv \frac{d\mathbb{Q}_{\mathbb{T}}}{d\mathbb{Q}}$ is the

[4]Equivalently, note that a foreign bond position is equal to a claim paying S_T at T and hence $p^{\star}_{t,T}S_t = p_{t,T}\mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}[S_T]$, which also implies that $F_{t,T} = \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}[S_T]$.

Radon-Nikodym derivative, and thus

$$F_{t,T} = \mathbb{E}_t^{\mathbb{Q}}\left[\Omega_T S_T\right] = \mathbb{E}_t^{\mathbb{Q}}\left[\frac{d\mathbb{Q}_{\mathbb{T}}}{d\mathbb{Q}} S_T\right] = \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right]. \tag{13}$$

We term the unbiasedness of the forward rate as a predictor for the expected spot rate under the T-forward measure the risk-adjusted FUH (RA-FUH).

Building on the RA-FUH, we derive a predictive regression for exchange rate changes and excess returns that resemble the Fama regressions in eqs. (5) and (6) but additionally account for a time-varying risk premium and are consistent with no-arbitrage. We follow the related literature and derive these relations for log returns; for completeness, we provide the analogue derivation without logs along with some technical details spared from the main text in Appendix A.

Under the assumption of rational expectations, taking conditional expectation yields the natural right-hand side of a predictive relation for the log exchange rate return:

$$\begin{aligned}
\Delta s_{t,T} &= \mathbb{E}_t^{\mathbb{P}}\left[s_T - s_t\right] + \varepsilon_{t,T} \\
&= \mathbb{E}_t^{\mathbb{P}}\left[s_T\right] - \left(\log F_{t,T} - (y_{t,T} - y^{\star}_{t,T})\right) + \varepsilon_{t,T} \\
&= \mathbb{E}_t^{\mathbb{P}}\left[s_T\right] - \log \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right] + (y_{t,T} - y^{\star}_{t,T}) + \varepsilon_{t,T} \\
&= \nu_{t,T} + (y_{t,T} - y^{\star}_{t,T}) + \varepsilon_{t,T},
\end{aligned} \tag{14}$$

with $\nu_{t,T} = \mathbb{E}_t^{\mathbb{P}}\left[\log S_T\right] - \log \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right]$. Expression (14), which we term risk-adjusted UIP (RA-UIP), shows that, in the absence of arbitrage exchange rate returns are governed by the yield differential - as postulated by UIP - but additionally comprise a time-varying component $\nu_{t,T}$. We also rewrite eq. (14) as a predictive relation for excess returns, analogously to eq. (6),

$$rx_{t,T} = \nu_{t,T} + \varepsilon_{t,T}, \tag{15}$$

showing that excess returns are driven by the time-varying component $\nu_{t,T}$. As $\nu_{t,T}$ is determined by the difference in expectations of the (log) spot exchange rate under the physical and the T-forward measure, it reflects risk adjustments. Hence RA-UIP explicitly identifies the risk premium postulated by Fama (1984) in equation (8) as $\lambda_{t,T} = -\nu_{t,T}$. As a consequence, the estimates of the slope coefficients in the conventional Fama regressions (5) and (6) under the assumption of no-arbitrage are subject to an omitted variable bias. More precisely, a time-varying risk premium exists and forward exchange rates in general deviate from future spot exchange rates unless interest rates are deterministic (i.e. $\mathbb{Q}_{\mathbb{T}} = \mathbb{Q}$) and agents are risk-neutral (i.e. $\mathbb{P} = \mathbb{Q}$).[5] To see this in more detail, note that

$$\mathbb{E}_t^{\mathbb{P}}\left[s_T\right] = \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[s_T\right] - \left(\mathbb{E}_t^{\mathbb{Q}}\left[s_T\right] - \mathbb{E}_t^{\mathbb{P}}\left[s_T\right]\right) - \left(\mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[s_T\right] - \mathbb{E}_t^{\mathbb{Q}}\left[s_T\right]\right) \tag{16}$$

which allows us to decompose the risk premium $\lambda_{t,T} = -\nu_{t,T}$ as[6]

$$\begin{aligned}\lambda_{t,T} &= \log \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right] - \mathbb{E}_t^{\mathbb{P}}\left[s_T\right] \\ &= \underbrace{\left(\mathbb{E}_t^{\mathbb{Q}}\left[s_T\right] - \mathbb{E}_t^{\mathbb{P}}\left[s_T\right]\right)}_{\text{pure currency risk}} + \underbrace{\left(\log \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right] - \mathbb{E}_t^{\mathbb{Q}}\left[s_T\right]\right)}_{\text{impact of stochastic rates}}\end{aligned} \tag{17}$$

The first term is a pure currency risk component which reflects corrections for risk aversion, the second term takes into account the impact of interest rates' stochastic nature on the risk premium.

3 The Empirical Model, Estimation and Evaluation of RA-UIP

3.1 Model

The RA-FUH and RA-UIP derived in the previous section are general, model-free relations that extend the conventional FUH and UIP in that they are free of risk preferences

[5] Note that even in this extreme case, the risk premium takes into account some mechanical Jensen's type terms, as then $\nu_{t,T} = \mathbb{E}_t^{\mathbb{P}}\left[\log S_T\right] - \log \mathbb{E}_t^{\mathbb{P}}\left[S_T\right]$ in eq. (14). These Jensen terms are considered to be very small in foreign exchange markets, though; see e.g. the survey of Engel (1996). To compare this to the derivation without logarithms see Appendix A.

[6] We provide a formal derivation of this relationship in Appendix A.2.

and consistent with no-arbitrage. In order to make these relations amenable for empirical work, we employ a parametric framework that allows to evaluate expressions (14) and (17) in closed form. We use a continuous-time, arbitrage-free dynamic multi-country affine term structure model with four latent factors to model the international financial market.[7] The design of the model is guided by the pioneering work of Backus, Foresi, and Telmer (2001) as well as the insights of Brennan and Xia (2006). Our *extended affine* model is flexible enough to meet the conditions formulated by Backus, Foresi, and Telmer (2001) for their *complete affine* model (asymmetric effects of state variables on state prices in different countries or negative nominal interest rates with positive probability) as well as the relations emphasized by Brennan and Xia (2006) in their *essentially affine* model (association between volatilities of pricing kernels, exchange rates, and risk premia). Since the model mainly serves as a workhorse to empirically assess the RA-FUH and RA-UIP rather than constituting a major contribution of our paper, we save the formal description and technical details of the model for Appendix B.1. Yet, two extensions deserve to be mentioned here as they are important for our work: first, in contrast to Backus, Foresi, and Telmer (2001), we use a multi-period setting to account for fluctuations in the bond market; this allows us to disentangle pure currency risk from interest rate risk as in the decomposition in eq. (17). Second, while Brennan and Xia (2006) use a linear first order approximation in time around the infinitesimal moments of the risk premium, our model produces exact, horizon-dependent risk premia. As a result, we can derive the term structure of foreign exchange risk premia in closed form.

3.2 Model Estimation

The model described above is formulated in terms of latent state variables. Relative to the small number of these driving state variables, the panel of asset prices that we need to fit is large. One can therefore think of these driving state variables as a low-dimensional

[7] It is well-established practice in the term structure literature to employ 3 factors (Litterman and Scheinkman, 1991). For international markets Leippold and Wu (2007) recommend using up to 7 factors. To keep the model as small as possible to focus on the economic ideas of this paper, we do not estimate such a large model. We choose 4 factors to reflect the co-movement between yields in different countries and to capture common factors in a parsimonious way.

representation of observed asset prices, very similar to factor reduction. Our estimation procedure differs from those used in previous research on multi-country affine term structure models in both the methodology applied as well as in terms of the conceptual setup. First, our methodological framework is Bayesian, which yields a posterior distribution of both latent state variables and the parameters behind our model. To obtain draws from this high-dimensional and complex distribution we employ Markov Chain Monte Carlo (MCMC) methods. The Bayesian methodology allows us to perform parameter inference without resorting to asymptotics, and it provides a very natural way to cope with latent state variables by treating them as parameters.[8] Second and more importantly, in the estimation procedure we do not only match bond yields in the US and the foreign country but simultaneously also match the predictive relationship implied by RA-UIP derived in eq. (14). In other words, we jointly fit the domestic and foreign term structures of interest rates as well as the term structure of foreign exchange risk premia. This innovative approach turns out to be crucial for matching the empirical properties of exchange rate (excess) returns. Details of the estimation procedure can be found in Appendix C.

3.3 Model Evaluation

In contrast to the standard formulation of UIP, the RA-UIP introduced in this paper explicitly accounts for a time-varying risk premium that arises from the assumption of no-arbitrage. This section describes how we assess whether the model is capable of identifying the risk premium. The RA-UIP model predictions for exchange rate changes $\Delta\widehat{s}_{t,T}$ and excess returns $\widehat{rx}_{t,T}$ are obtained from eqs. (14) and (15) using the estimation procedure outlined in the previous section.[9]

As a first step, we check whether our model risk premium fulfills the conditions formulated by Fama (1984), given in eq. (9): first, the covariance between the model-implied

[8]This is a non-negligable advantage over Maximum Likelihood estimation, where the state variables are either integrated out, some prices are assumed to be observed without error to back out the state variables, or filters are employed which are either expensive to evaluate, or approximations. For GMM estimation similar constraints apply, see for instance the implied-state GMM approach in Pan (2002).

[9]To be precise, the expressions are evaluated at the multivariate median of the parameter posterior distribution along with a smoothed estimate of the trajectory of the latent state variables.

risk premium, $\widehat{\lambda}_{t,T} = -\widehat{\nu}_{t,T}$, and expected exchange rate changes, $\Delta\widehat{s}_{t,T}$, is negative; second, the absolute value of this covariance is greater than the variance of expected exchange rate changes. If the model risk premium satisfies these conditions, its omission in the Fama regression causes a negative β estimate.

The next step is to analyze whether the risk premium allows for unbiased predictions of excess returns and hence spot rate changes (or whether the risk premium just accounts for part of the forward bias). We therefore regress observed excess returns on our RA-UIP model predicted excess returns $\widehat{rx}_{t,T}$

$$rx_{t,T} = \alpha' + \beta' \widehat{rx}_{t,T} + \eta'_{t,T} \tag{18}$$

and test whether $\alpha' = 0$ and whether the slope coefficients are statistically significant and if $\beta' = 1$. If we cannot reject that $\alpha' = 0$ and $\beta' = 1$, this indicates that accounting for the risk premium *can* be sufficient to resolve the forward bias puzzle without additionally requiring departures from rational expectations.

Finally, we assess the predictive accuracy of our model by using four additional evaluation criteria: the hit-ratio (HR), an $R2$-measure, the test proposed by Clark and West (2007) based on mean squared prediction errors (CW), and the Giacomini and White (2006) test for conditional predictive ability (GW). The predictions are all in-sample predictions, because our focus is not to provide forecasting models but to evaluate departures from UIP.[10] In other words, we have a twofold motivation for applying these criteria: first, we gain additional insight on our model's goodness of fit as compared to only considering the R^2 of the predictive regression. Second, we complement the evidence on the predictability of excess returns by assessing the predictive ability of our model per se as well as relative to the benchmark predictions based on UIP and a random-walk (RW) without drift. These results will also show whether empirical exchange rate dynamics are more adequately characterized by RA-UIP, UIP or the RW.

[10] Moreover, some recent research argues that it is not clear whether out-of-sample tests of predictability are powerful enough to discriminate among competing predictive variables or models, showing that in-sample tests can be more reliable under certain conditions (e.g. Campbell and Thompson (2008) and the references therein).

We apply the four evaluation criteria to compare the accuracy of the RA-UIP model predictions for excess returns, $\widehat{rx}_{t,T}$, to predictions based on the benchmarks. The UIP predicted exchange rate change is given by $\Delta\widehat{s}^{UIP}_{t,T} = (y_{t,T} - y^{\star}_{t,T})$ and the corresponding excess return prediction is $\widehat{rx}^{UIP}_{t,T} = 0$. The RW predictions are $\Delta\widehat{s}^{RW}_{t,T} = 0$ and $\widehat{rx}^{RW}_{t,T} = -(y_{t,T} - y^{\star}_{t,T})$. HR is calculated as the proportion of times the sign of the excess return is correctly predicted. The remaining criteria are defined as functions of squared prediction errors of our model, SE^M, and of the respective benchmark B, SE^B (where B is either UIP or RW); the respective means are denoted by MSE^M and MSE^B. The $R2$ measure of our model as compared to the benchmark is given by

$$R2 = 1 - \frac{MSE^M}{MSE^B}. \tag{19}$$

Positive values indicate that our model performs better than the benchmark.[11]

The CW test statistic is defined as

$$CW = MSE^B - MSE^M + N^{-1} \sum_{n=1}^{N} \left(\Delta\widehat{rx}^B_{t,T} - \Delta\widehat{rx}_{t,T} \right)^2, \tag{20}$$

where N is the number of observations in the sample. The CW test allows to compare the predictive ability of the RA-UIP model as compared to that of the nested alternatives. In contrast to other tests which are only based on the difference in MSEs, e.g. Diebold and Mariano (1995), the last term in eq. (20) adjusts for the upward bias in MSE^M caused by parameter estimates in the larger model whose population values are zero and just introduce noise. In our empirical analysis, we apply the block bootstrap procedure described in Appendix D to obtain p-values for the CW test statistics.

To assess the conditional predictive ability of the RA-UIP model, we implement the GW test for the full sample as follows.[12] The predictions are based on the full time-t information set $\mathcal{F}_t$. Using a $\mathcal{F}_t$-measurable test function h_t, we test the null hypothesis that

[11] Note that the $R2$ measure is based on the same information as the test by Diebold and Mariano (1995).

[12] Although the main focus of Giacomini and White (2006) is on rolling window methods, their results also hold for a fixed estimation sample (cf. p. 1548).

predictions based on our model and the benchmark predictions have equal conditional predictive ability, $H_{0,h} : \mathbb{E}[h_t \Delta L_T] = 0$. ΔL_T denotes the differential in loss functions of the two competing predictions at t for time T; for the case of the squared prediction error loss function, $\Delta L_T = SE_T^B - SE_T^M$. The test function we use is $h_t = (1, \Delta L_t)'$. The GW statistic is given by

$$GW = N\left(N^{-1}\sum_{n=1}^{N} h_t \Delta L_T\right)' \widehat{\Omega}_N^{-1}\left(N^{-1}\sum_{n=1}^{N} h_t \Delta L_T\right) \tag{21}$$

where $\widehat{\Omega}_N^{-1}$ is a consistent estimate of the variance of $h_t \Delta L_T$.[13] The empirical results will be based on block-bootstrapped p-values for the GW test statistic.

4 Empirical Analysis

4.1 Data

Daily interest rate and spot exchange rate data are obtained from Datastream. Riskless zero-coupon yields are bootstrapped from money market (Libor) rates with maturities of 1, 3, and 6 months and swap rates with maturities of 1, 2, 3 and 4 years. Feldhütter and Lando (2008) show that swap rates are the best parsimonious proxy for riskless rates. The model estimation is performed on daily zero-yields and spot exchange rates for the US dollar against the Australian dollar (AUD), Canadian dollar (CAD), Swiss franc (CHF), the merged Deutsch mark and euro series (DEM-EUR), the British pound (GBP) and Japanese yen (JPY). The sample periods are October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY.

To relate our model risk premia to financial market and macroeconomic variables, we also obtain daily data for the VIX S&P 500 implied volatility index. Data for industrial production and narrow money supply are obtained from the OECD Main Economic

[13] To obtain a HAC consistent estimate for $(T - t) > 1$ we use the weight function as in Newey and West (1987) with the truncation lag being equal to $(T - t) - 1$, as suggested by Giacomini and White (2006).

Indicators at a monthly frequency for all countries except industrial production in Australia and Switzerland, which is only available quarterly. The sample periods match those mentioned above with the exception of the VIX series which starts in January 1990. To measure US consumption growth, we use consumption data (available quarterly), the consumer price index, and population figures from the International Monetary Fund's *International Financial Statistics* database.

4.2 Descriptive Statistics and Fama Regressions

The empirical analysis presented here is based on non-overlapping observations for prediction horizons of 1 day, 1 week, and 1 month. For the longer horizons of 3 months, 1 year, and 4 years we choose a monthly frequency to maintain a reasonable number of data points. Tables 1 and 2 report descriptive statistics for annualized exchange rate returns and yield differentials.

As a preliminary exercise, we estimate the conventional Fama regression (5). The results reported in Table 3 are consistent with the 'forward bias' documented in previous research. While the estimates of the intercept α are in most cases small and statistically insignificantly different from zero, the β estimates are generally negative and different from the UIP theoretical value of unity for all currencies. For the GBP, estimates across all six horizons are positive but only the 4-year β estimate is statistically significant at conventional significance levels.[14] As outlined in Section 2.2, the second Fama regression in eq. (6) contains the same information because $\gamma = \beta - 1$. Since $t[\gamma = 0] = t[\beta = 1]$ the results are in line with previous evidence that excess returns are predictable on the basis of the lagged interest differential (forward premium).

4.3 Model Evaluation

Next, we present results that show that our model fits the data reasonably well in that the pricing errors for the term structures are satisfactory. Then, we provide empirical

[14]These values are likely to reflect two major UIP reversions the GBP experienced in our sample: the ERM crisis in 1992 and for the 4-year horizon also the impact of the current financial crisis on the UK and its currency.

results for the model evaluation criteria described in Section 3.3.

Table 4 describes the model's ability to capture the term structure of interest rates. Since it is computationally infeasible to estimate the US and all foreign term structures of interest rates jointly with the corresponding exchange rates all at once, we estimate bilateral models for country pairs following Backus, Foresi, and Telmer (2001) and Brandt and Santa-Clara (2002). We report the root mean squared pricing errors of the domestic US yields (Panel A) and the respective foreign yields (Panel B) measured in basis points for each of the six bilateral models. As an alternative, one could estimate single currency term structure models (as in Brennan and Xia (2006)) and perform an ex-post analysis of the currency implications. The advantage of this alternative is that one ensures ex-ante that the US pricing kernel is unique, the disadvantage being that one disregards all information available from currency forwards and the dynamics of the exchange rate. In our context of foreign exchange risk premia, we choose to estimate bilateral models and then compare the US yields (and their pricing errors) implied by these models. Inspection of the RMSE of US yields in Table 4 reveals a deviation of maximally 2 basis points for the longest maturity, while for shorter maturities the RMSEs are identical across estimations. The exception is the JPY model, which exhibits larger RMSEs for US yields but smaller for foreign yields as compared to the other models. We perform various additional tests (e.g. pairwise regression of US yield pricing errors from the bilateral model estimations, not reported) and cannot reject the null hypothesis that the implied US term structure is the same across models. This means that our bilateral estimation effectively delivers a unique US pricing kernel, although the uniqueness is not imposed in the model.

The errors reported in Table 4 are in the range of Brennan and Xia (2006) and Anderson, Hammond, and Ramezani (2009). While Brennan and Xia (2006) only report an estimate for the standard deviation of the pricing error, a comparison with the RMSE from Anderson, Hammond, and Ramezani (2009) reveals an interesting phenomenon. While pricing errors in Anderson, Hammond, and Ramezani (2009) exhibit a tub-shaped pattern as a function of the maturity of yields, the pricing errors in our study monotonically increase with maturity.[15]

4.3.1 Fama Conditions and Unbiasedness of Model Predictions

Next, we verify whether our model risk premium fulfills the conditions formulated by Fama (1984), as described in eq. (9), such that the omission of the risk premium causes a negative β estimate. We report the covariances between the risk premium and expected exchange rate changes and the variance of expected changes in Table 5. The results show that both conditions are fulfilled for all currencies except the GBP. Specifically , for the GBP the first condition (negative covariance) is fulfilled across all six horizons but the second condition is not. However, the violation of the second condition is not surprising as it is consistent with the positive β estimates for the GBP in Table 3. We rather view this as a corroboration of the flexibility of our model.

Table 6 presents results for the predictive regression (18) by reporting parameter estimates along with block-bootstrapped standard errors in parentheses as well as t-statistics for the null hypothesis of unbiasedness $\beta' = 1$.[16] The table also reports the R^2 of the regressions but we defer a detailed discussion of the model fit to the next subsection where we evaluate the predictive ability criteria motivated in Section 3.3. In brief, we find strong evidence that excess return predictions based on the model risk premium are unbiased. All estimates of the intercept α' are very small and not significantly different from zero. All estimates of the slope coefficient β' are positive (except GBP at the 1-day horizon) and become closer to unity and more significant as the prediction horizons increase. Parameter estimates are significantly positive across all horizons for AUD, CAD, CHF, and DEM-EUR, for horizons longer than 1 month for the JPY, and at the 4-year horizon for the GBP. At the same time the β's are not statistically different from unity

[15] When comparing our pricing errors to those reported in other studies, one has to keep in mind that we use daily data from September 1989 to October 2008, while Brennan and Xia (2006) use monthly data from January 1985 to May 2002 and Anderson, Hammond, and Ramezani (2009) uses weekly data only from May 1998 to August 2005.

[16] We calculate block-bootstrapped standard errors for all subsequent regressions. The block-bootstrap procedure avoids the necessity to rely on asymptotic theory but still allows to handle serial correlation and heteroscedasticity. We also calculate, but do not report, Newey and West (1987) standard errors with the optimal truncation lag chosen as suggested by Andrews (1991). These standard errors are very similar or slightly smaller than those obtained from the block-bootstrap procedure.

except at the 1-day horizon for the CHF and horizons up to one month for the JPY. The less pronounced evidence for the GBP is again consistent with the comparably smaller forward bias as judged by the Fama regression results in Table 3.

To reiterate, the findings related to the Fama conditions and the unbiasedness of model predictions are consistent with the notion that the time-varying risk premium accounts for the forward bias puzzle. While results from the Fama conditions show that the risk premium has the general properties to cause a downward bias in the β estimate of the Fama regression across horizons, the unbiasedness results strengthen this evidence as they indicate that accounting for the risk premium can be sufficient to resolve the puzzle without requiring departures from rational expectations.

4.3.2 Predictability of Excess Returns

In Table 7, we present results for the predictive ability criteria discussed in Section 3.3. As motivated above, we do not pursue the goal of providing out-of-sample exchange rate forecasting models but to understand the dynamics of deviations from UIP. The HR, $R2$, CW, and GW measures allow us to gain insight on our model's goodness of fit as compared to only considering the R^2 of the predictive regression. We furthermore complement previous evidence on the predictability of excess returns based on our model per se and as compared to the benchmark predictions based on UIP and the RW.

The HR indicates that our model predictions have high directional accuracy: while the HR is slightly above 50% for the 1-day horizon, it dramatically increases across horizons for all currencies. The highest HR are achieved for the 1-year and 4-year horizons with the largest values across currencies ranging from 63% to 97%.[17] Thus, we have first evidence that our model fits the data very well in that it replicates the sign of excess returns, i.e. UIP deviations.

The values reported for the $R2$-measure, as defined in eq. (19), indicate that our model outperforms both benchmarks. The $R2$s are positive for all currencies across all horizons against the UIP benchmark. The $R2$s are also positive across currencies and horizons

[17]The Pesaran and Timmermann (1992) test statistics for directional accuracy also suggest that most of the HR are highly significant. Results are omitted to save space but available on request.

against the RW benchmark with the exception of negative values at the short horizons for the GBP (up to 1 week) and the JPY (up to 1 month). A common feature across all currencies is that the highest $R2$ is reached for the longest horizons, ranging from 30% to 79% against UIP and from 21% to 67% against the RW.[18] In other words, the mean-squared prediction errors of our model are much smaller than those of the benchmarks providing another piece of evidence that our RA-UIP model fits the empirical behavior of exchange rates better than UIP or the RW.

The results for the Clark and West (2007) test and the Giacomini and White (2006) test for conditional predictive ability further support that the model predictions are more accurate than those of the benchmarks. We report p-values for the test statistics which are obtained from the block-bootstrap procedure described in Appendix D. The CW p-values generally decrease with the prediction horizon and indicate that our model predictions significantly outperform UIP predictions for 4 currencies at the 1-day and 1-week horizon and for all 6 currencies at horizons of 1 month or longer. The results for the RW benchmark generally follow the same pattern but exhibit more variability in terms of significance at the shorter horizons. The GW results indicate that the model dominates UIP and RW also in terms of conditional predictive ability. Again, the p-values exhibit some cross-currency variability for shorter horizons, but they indicate significantly stronger predictive ability of the model as compared to UIP at horizons beyond 1 month for AUD, CAD, CHF, and DEM-EUR; for the GBP and JPY results are significant at the 1-year and 4-year horizons. The results for the RW benchmark are very similar.

Overall, the predictions from our model dominate those based on the benchmarks, thereby providing evidence that the empirical behavior of exchange rates is more accurately characterized by RA-UIP as compared to UIP or the RW. The superior predictive ability arises from the fact that the model-implied no-arbitrage conditions allow to identify the risk premia that drive (excess) returns.[19]

[18] The increasing predictability with longer horizons does not result from a mechanical link between short- and long-horizon predictions similar to the arguments of e.g. Cochrane (2001), p. 389f, or Boudoukh, Richardson, and Whitelaw (2006). Note that we have a different predictor and different dependent variable for each horizon.

[19] The finding that no-arbitrage improves predictions has similarly been documented in the term structure literature, see e.g. Ang and Piazzesi (2003), Christensen, Diebold, and Rudebusch (2007), Diez de

4.4 Decomposing Foreign Exchange Risk Premia

Following the derivations of the RA-FUH and RA-UIP in Section 2.3, we show in eq. (17) that the foreign exchange risk premium can be decomposed into a pure currency risk component and a second component that accounts for the fact that interest rates are stochastic. Table 8 displays descriptive statistics for estimated risk premia and their components on an annualized basis.

The average premium for pure currency risk can be positive or negative. Consistent with intuition we find that compensation for bearing interest rate risk is strictly positive. The average interest rate risk premium contributes, depending on the currency, a sizable level to the overall risk premium. However, the standard deviations are very small compared to those of the overall risk premia.

These results suggest that the variation in foreign exchange risk premia - and hence deviations from UIP constituting the forward bias puzzle - are largely driven by the pure currency risk component. We redo the empirical model evaluation analysis in Section 4.3 based on model expectations comprising only the pure currency risk component. We find that the results (not reported) are qualitatively identical to those above and that quantitative differences are very small. Nevertheless, although the interest rate risk component does not vary much, its sizable contribution to the average level of foreign exchange risk premia may be relevant in numerous other contexts, for example assessing the profitability and economic value of currency speculation, which we do not investigate in this paper.

5 Drivers of the Risk Premium

The above results provide strong empirical support for the existence of time-varying risk premia as stated by RA-UIP. In this section we show that the time-variation in expected excess returns is closely related to global risk measures and to macroeconomic variables.

Our proxy for global risk is based on the VIX S&P 500 implied volatility index traded

los Rios (2009) and Almeida and Vicente (2008).

at the CBOE, which is highly correlated with similar volatility indices in other countries; see e.g. Lustig, Roussanov, and Verdelhan (2008). Furthermore, the VIX can also be viewed as a proxy for funding liquidity constraints, see e.g. Brunnermeier, Nagel, and Pedersen (2008). If the VIX captures global risk appetite and funding liquidity constraints, expected currency excess returns should be negatively related to the VIX multiplied by the sign of the yield differential, $sVIX_t \equiv VIX_t \times \text{sign}[y_t - y_t^\star]$: in times of global market uncertainty and higher funding liquidity constraints, investors demand higher risk premia on high yield currencies while they accept lower (or more negative) risk premia on low yield currencies, consistent with 'flight-to-quality' and 'flight-to-liquidity' arguments.[20]

Recent research suggests that risk premia on US exchange rates are countercyclical to the US economy, similarly to risk premia in other markets; see e.g. Lustig and Verdelhan (2007), De Santis and Fornari (2008), and Lustig, Roussanov, and Verdelhan (2008). As proxies for the state of the US economy, we use industrial production (IP_t) as a measure of output, and M1 as a measure for narrow money supply (NM_t). Using monthly data, the growth rates ΔIP_t and ΔNM_t are defined as 1-year log changes. If our model risk premium is countercyclical, the relationship between expected excess returns and output growth should be negative whereas the relationship with money growth should be positive.

Lustig and Verdelhan (2007) show that high interest rate currencies depreciate on average when domestic consumption growth is low while low interest rate currencies appreciate under the same conditions. They argue that low interest rate currencies hence provide domestic investors with a hedge against aggregate domestic consumption growth risk. We construct a quarterly series of US consumption based on total private consumption deflated by the consumer price index and divided by population figures to obtain per capita consumption. Consumption growth is defined as the 1-year log change. To account for the asymmetric effect of low versus high interest rate currencies, we multiply consumption growth by the sign of the yield differential. The findings of Lustig and

[20] We also use the TED spread (difference between the 3-month Eurodollar rate and the 3-month Treasury rate) as an alternative proxy. The results are similar to those based on the VIX reported in the paper; this is in line with Brunnermeier, Nagel, and Pedersen (2008).

Verdelhan (2007) suggest that expected excess returns should be negatively related to signed consumption growth $s\Delta CO_t$.

Finally, we relate the risk premium to macroeconomic variables deemed relevant in traditional monetary models of the exchange rate. As a proxy for exchange rate fundamentals we use the "observable fundamentals" as in Engel and West (2005), defined as the country differential in money supply minus the country differential in output. We measure output and money supply in the foreign countries analogously to the US variables and define the change in observable fundamentals as $\Delta OF_t = (\Delta NM_t - \Delta NM_t^\star) - (\Delta IP_t - \Delta IP_t^\star)$. Traditional exchange rate models suggest that the relationship between these fundamentals and expected excess returns should be positive.

Table 9 presents contemporaneous correlations of expected excess returns with the variables described above; the significance indicated by the asterisks is judged by block bootstrapped standard errors which are not reported to save space. The correlations strongly support our priors as all coefficients are signed correctly across currencies and horizons, in most cases with a high level of significance. These results thus suggest that foreign exchange risk premia are driven by global risk perception and macroeconomic variables in a way that is consistent with economic intuition.

We also run univariate regressions of expected excess returns on the signed VIX, signed consumption growth, and the observable fundamentals, as well as multivariate regressions on combinations of these variables. We report OLS estimates in Table 10. The univariate results confirm the correlation analysis for the three proxies in terms of sign and statistical significance of coefficients, in most cases accompanied with large explanatory power (as judged by the R^2). The signed VIX has lowest explanatory power for the GBP, but for all other currencies it is substantial: at the 1-day horizon the R^2 ranges from 0.13 to 0.58, at the 1-year horizon they range from 0.31 to 0.62. The observable fundamentals have similar explanatory power across currencies (except CHF) with the R^2 ranging between 0.32 and 0.51. The results for signed consumption growth exhibit the largest cross-currency variability in terms of explanatory power, with R^2s ranging from 0.08 to 0.14 for the GBP and JPY, from 0.18 and to 0.29 for CHF, and from 0.54 to 0.61 for AUD, CAD, and DEM-EUR.

In our multivariate regression analysis we combine the observable fundamentals with either the signed VIX or signed consumption growth. Signs and significance of coefficients are similar to the univariate regressions but the explanatory power can be substantially larger. The R^2s are lowest for the CHF with values between 0.23 and 0.34. For CAD and JPY the specification with signed VIX fits the data somewhat better; e.g. for the CAD the R^2s are 0.83 (3 months) and 0.75 (1 year). In case of the AUD, the specification with signed consumption growth fits better with an R^2 of around 0.72 for both horizons. The results for DEM-EUR (R^2s of 0.64 and 0.67) and GBP (0.51 and 0.43) are very similar for both specifications.

Overall, we find that the model risk premium is related to global risk aversion, countercyclical to the US economy, and associated with traditional exchange rate fundamentals. The few cases where significance is less pronounced or explanatory power is lower may even corroborate our results. For example, the absence of a strong relation between the GBP and the global risk proxy is consistent with the comparably smaller forward bias in our GBP data set. Also, finding that the CHF's link to observable fundamentals is weak but that its link to global risk is strong is consistent with Switzerland being viewed as a 'safe haven' and primarily as a destination for flight-to-quality.

6 Conclusion

There is a large literature documenting the empirical failure of uncovered interest rate parity and of the forward unbiasedness hypothesis: the forward premium is a biased predictor for subsequent exchange rate changes, and the forward rate is a biased predictor for the future spot exchange rate. In this paper we show from the principle of no-arbitrage that currency forwards are in general biased predictors for spot exchange rates, because they not only reflect expected spot rates but additionally comprise time-varying risk premia that compensate for both currency risk and interest rate risk. We develop an expression for the risk premium and employ it in a prediction model resembling the Fama

(1984) regression. Expected exchange rate returns are driven by the yield differential but additionally comprise a time-varying risk premium (Fama's omitted variable), which we estimate from a multi-currency term structure model.

For the empirical analysis, we extend affine term structure models applied in multi-currency contexts to explicitly account for these properties of forward rates and embedded risk premia. We take the model to US exchange rate data and find that estimated model expectations and risk premia satisfy the necessary conditions for explaining the forward bias puzzle. Moreover, the model is capable of producing unbiased predictions for excess returns and hence we conclude that accounting for risk premia can be sufficient to resolve the forward bias puzzle without additionally requiring departures from rational expectations.

Furthermore, we provide strong empirical evidence that risk premia are closely linked to economic variables that proxy for global risk, the US business cycle, and traditional exchange rate fundamentals. Our results suggest that expected excess returns reflect flight-to-quality and flight-to-liquidity considerations. Expected excess returns also depend on macroeconomic variables (output growth, money supply growth, consumption growth) such that risk premia in dollar exchange rates are countercyclical to the US economy.

We disentangle the risk premia into compensation for currency risk and interest rate risk. We find that the variation in expected excess returns is almost entirely driven by currency risk. The premium for interest rate risk exhibits very little variation but contributes substantially to the level of risk premia for some currencies. Given its sizable contribution to the overall level of compensation for risk in foreign exchange markets, interest rate risk should be explicitly accounted for in future research, for instance, when assessing the profitability and economic value of currency speculation.

A Additional Derivations for RA-UIP and RA-FUH

A.1 Predictive relations without logarithms

Analogously to eqs. (14) and (15) we derive the predictive relations for changes of the spot exchange rate and excess returns without taking logarithms. For the sake of easier readability, we use the same notation for $\varepsilon_{t,T}$, $\nu_{t,T}$, and $\lambda_{t,T}$ here for the case of no logarithms as in the main text where we use logarithms.

Defining $\Delta S_{t,T} \equiv (S_T - S_t)/S_t$. Under the assumption of rational expectations, taking conditional expectation yields the natural right-hand side of a predictive relation for the exchange rate return

$$\begin{aligned}
\Delta S_{t,T} &= \mathbb{E}_t^{\mathbb{P}}\left[S_T\right]/S_t - 1 + \varepsilon_{t,T} \\
&= \Big(\mathbb{E}_t^{\mathbb{P}}\left[S_T\right]/\mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right]\Big)e^{(y_{t,T}-y^{\star}_{t,T})} - 1 + \varepsilon_{t,T} \\
&= \nu_{t,T} + e^{(y_{t,T}-y^{\star}_{t,T})} - 1 + \varepsilon_{t,T},
\end{aligned} \tag{A.1}$$

with $\nu_{t,T} = \Big(\mathbb{E}_t^{\mathbb{P}}\left[S_T\right]/\mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right] - 1\Big)e^{(y_{t,T}-y^{\star}_{t,T})}$. Hence, unless $\mathbb{Q}_{\mathbb{T}} = \mathbb{P}$, i.e. under risk-neutrality and deterministic short rates, there is a time-varying risk premium, $\lambda_{t,T} = -\nu_{t,T}$. Analogously, we find that excess returns defined as $RX_{t,T} = (S_T - F_{t,T})/S_t$ comprise the time-varying risk premium

$$\begin{aligned}
RX_{t,T} &= \frac{\mathbb{E}_t^{\mathbb{P}}\left[S_T\right] - \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right]}{S_t} + \varepsilon_{t,T}, \\
&= \frac{\mathbb{E}_t^{\mathbb{P}}\left[S_T\right] - \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right]}{\mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right]}e^{(y_{t,T}-y^{\star}_{t,T})} + \varepsilon_{t,T}, \\
&= \nu_{t,T} + \varepsilon_{t,T}.
\end{aligned} \tag{A.2}$$

A.2 Decomposition of the risk premium

The relationship in eq. (16) is formally established from the following:

$$\begin{aligned}
\mathbb{E}_t^{\mathbb{Q}_\mathbb{T}}[s_T] &= \mathbb{E}_t^{\mathbb{Q}}\left[\frac{d\mathbb{Q}^T}{d\mathbb{Q}} s_T\right] \\
&= \mathbb{E}_t^{\mathbb{Q}}[s_T] + \mathbb{C}ov_t^{\mathbb{Q}}\left[\frac{d\mathbb{Q}^T}{d\mathbb{Q}}, s_T\right] \\
&= \mathbb{E}_t^{\mathbb{P}}\left[\frac{d\mathbb{Q}}{d\mathbb{P}} s_T\right] + \mathbb{C}ov_t^{\mathbb{Q}}\left[\frac{d\mathbb{Q}^T}{d\mathbb{Q}}, s_T\right] \\
&= \mathbb{E}_t^{\mathbb{P}}[s_T] + \mathbb{C}ov_t^{\mathbb{P}}\left[\frac{d\mathbb{Q}}{d\mathbb{P}}, s_T\right] + \mathbb{C}ov_t^{\mathbb{Q}}\left[\frac{d\mathbb{Q}^T}{d\mathbb{Q}}, s_T\right] \\
&= \mathbb{E}_t^{\mathbb{P}}[s_T] + \left(\mathbb{E}_t^{\mathbb{Q}}[s_T] - \mathbb{E}_t^{\mathbb{P}}[s_T]\right) + \left(\mathbb{E}_t^{\mathbb{Q}_\mathbb{T}}[s_T] - \mathbb{E}_t^{\mathbb{Q}}[s_T]\right).
\end{aligned}$$

B Affine Multi-Country Term Structure Model

B.1 A Continuous-Time Model for an International Economy

For our econometric analysis, to put structure on the coefficients and error terms appearing in the predictive equation (14), we endow the international financial market with a model for interest rate risk and currency risk. This section therefore engineers a continuous-time, arbitrage-free dynamic term structure model for two economies, along with the exchange rate. The workhorse for this exercise is the framework of affine diffusion processes.

We assume that the international economy is driven by a time-homogeneous, partially observed Markov diffusion process $Z \equiv (Z_t)_{t\geq 0, Z_0=z_0\in\mathcal{D}} =: (X_{1t}, X_{2t}, X_{3t}, X_{4t}, s_t) = (X_t, s_t)$, living on state space $\mathcal{D} = \mathbb{R}^2_{++} \times \mathbb{R}^3$, where $\mathbb{R}_{++} \equiv \{x \in \mathbb{R} : x > 0\}$. To reflect the co-movement between yields in different countries and to capture common factors in a parsimonious way we choose a latent 4-factor setting for our international economy.To ensure arbitrage-free markets we start with a relation between the two countries' pricing kernels that ensures consistent pricing

$$\frac{M_t^\star}{M_0^\star} \equiv \frac{S_t}{S_0}\frac{M_t}{M_0}. \tag{B.1}$$

Here, M is the global pricing kernel in domestic currency, and $M^{\star}$ is the global pricing kernel in foreign currency. This relation has been established by Backus, Foresi, and Telmer (2001). Graveline (2006) notes that it ensures that the foreign pricing kernel is the minimum-variance (MV) kernel, provided the domestic kernel is the MV kernel. This condition puts restrictions on the dynamic behavior of the pricing kernels and the spot exchange rate. It will only be possible to specify the dynamics of two of the three constituents of (B.1), while the third will be determined endogenously. Our dynamic specification builds on these ideas. The general guideline is to maintain a tractable model with maximum flexibility. We start with affine dynamics of the latent factors X_t,

$$dX_t = (a^{\mathbb{P}} + b^{\mathbb{P}} X_t)dt + \sigma(X_t)dW_t^{\mathbb{P}}, \text{ where} \tag{B.2}$$

$$a^{\mathbb{P}} \equiv \begin{pmatrix} a_1^{\mathbb{P}} \\ a_2^{\mathbb{P}} \\ a_3^{\mathbb{P}} \\ a_3^{\mathbb{P}} \end{pmatrix}, \; b^{\mathbb{P}} \equiv \begin{pmatrix} b_{11}^{\mathbb{P}} & 0 & 0 & 0 \\ b_{21}^{\mathbb{P}} & b_{22}^{\mathbb{P}} & 0 & 0 \\ b_{31}^{\mathbb{P}} & b_{32}^{\mathbb{P}} & b_{33}^{\mathbb{P}} & 0 \\ b_{41}^{\mathbb{P}} & b_{42}^{\mathbb{P}} & b_{43}^{\mathbb{P}} & b_{44}^{\mathbb{P}} \end{pmatrix}, \; \sigma(X_t) \equiv \operatorname{diag} \begin{pmatrix} \sqrt{X_{1t}} \\ \sqrt{X_{2t}} \\ \sqrt{1 + \beta_1 X_{1t} + \beta_2 X_{2t}} \\ \sqrt{1 + \gamma_1 X_{1t} + \gamma_2 X_{2t}} \end{pmatrix}, \tag{B.3}$$

and $dW^{\mathbb{P}} = d(W_{1t}^{\mathbb{P}}, \ldots, W_{4t}^{\mathbb{P}})^{\top}$. The constant coefficients in $\sigma(X_t)$ are restricted to unity for identification purposes. The dynamics of the domestic pricing kernel are

$$\frac{dM_t}{M_t} = -r_t dt - \Lambda(X_t)^{\top} dW_t^{\mathbb{P}}, \tag{B.4}$$

where $\Lambda : \mathbb{R}_{++}^2 \times \mathbb{R}^2 \mapsto \mathbb{R}^4$ is the solution to

$$\Lambda(x) = \sigma(x)^{-1} \left(a^{\mathbb{P}} + b^{\mathbb{P}} x - (a^{\mathbb{Q}} + b^{\mathbb{Q}} x) \right), \text{ where} \tag{B.5}$$

$$a^{\mathbb{Q}} \equiv \begin{pmatrix} a_1^{\mathbb{Q}} \\ a_2^{\mathbb{Q}} \\ 0 \\ 0 \end{pmatrix}, \; b^{\mathbb{Q}} \equiv \begin{pmatrix} b_{11}^{\mathbb{Q}} & 0 & 0 & 0 \\ b_{21}^{\mathbb{Q}} & b_{22}^{\mathbb{Q}} & 0 & 0 \\ b_{31}^{\mathbb{Q}} & b_{32}^{\mathbb{Q}} & b_{33}^{\mathbb{Q}} & 0 \\ b_{41}^{\mathbb{Q}} & b_{42}^{\mathbb{Q}} & b_{43}^{\mathbb{Q}} & b_{44}^{\mathbb{Q}} \end{pmatrix}. \tag{B.6}$$

The market price of risk specification Λ follows Cheridito, Filipović, and Kimmel (2007);

it is admissible if $2a_1^{\mathbb{P}} > 1, 2a_2^{\mathbb{P}} > 1$, and $b_{11}^{\mathbb{P}} < 0, b_{22}^{\mathbb{P}} < 0$ in addition to the admissibility conditions from Duffie, Filipović, and Schachermayer (2003). We define $r_t \equiv \delta_0 + \delta_1 X_t$ with $\delta_1 = (\delta_{11}, \delta_{12}, \delta_{13}, \delta_{14})$. We also define the dynamics of the foreign pricing kernel as:

$$\frac{dM_t^\star}{M_t^\star} = -r_t^\star dt - (\Lambda(X_t)^\top - \Sigma\,\sigma(X_t))\,dW_t^{\mathbb{P}}, \tag{B.7}$$

where the drift of X_t under $\mathbb{Q}_\star$ (the foreign $\mathbb{Q}$ measure) solves[21]

$$a^{\mathbb{Q}_\star} + b^{\mathbb{Q}_\star} x = a^{\mathbb{P}} + b^{\mathbb{P}} x - \sigma(x)(\Lambda(x)^\top - \Sigma\sigma(x))^\top. \tag{B.8}$$

Computing the solution to eqs. (B.4) and (B.7) we get by eq. (B.1) that the foreign exchange rate S_t evolves according to

$$\frac{dS_t}{S_t} = (r_t - r_t^\star + \Sigma\,\sigma(X_t)\,\Lambda(X_t))dt + \Sigma\sigma(X_t)dW_t^{\mathbb{P}}, \tag{B.9}$$

where $\Sigma \equiv (\Sigma_1, \Sigma_2, \Sigma_3, \Sigma_4)$, and $r_t^\star \equiv \delta_0^\star + \delta_1^\star X_t$ with $\delta_1^\star \equiv (\delta_{11}^\star, \delta_{12}^\star, \delta_{13}^\star, \delta_{14}^\star)$. The corresponding log dynamics of s_t are then

$$ds_t = \left(r_t - r_t^\star + \Sigma\,\sigma(X_t)\,\Lambda(X_t) - \frac{1}{2}\,\Sigma\,\sigma(X_t)\,\sigma(X_t)^\top\,\Sigma^\top\right)dt + \Sigma\sigma(X_t)dW_t^{\mathbb{P}}, \tag{B.10}$$

which turn out to be affine in X_t.[22]

The instantaneous covariance matrix of $Z_t = (X_t, s_t)$ is singular (while $\sigma(x)\sigma(x)^\top$ is non-singular), since we have a 5-dimensional process for only 4 driving Brownian motions. Nevertheless, Z_t constitutes an affine Markov process under probability measures $\mathbb{P}, \mathbb{Q}$,

[21]In addition to the admissibility conditions, drifts (B.6) and (B.8) also satisfy $2a_1^{\mathbb{Q}} > 1, 2a_2^{\mathbb{Q}} > 1$, $b_{11}^{\mathbb{Q}} < 0, b_{22}^{\mathbb{Q}} < 0$, and $2a_1^{\mathbb{Q}_\star} > 1, 2a_2^{\mathbb{Q}_\star} > 1$, $b_{11}^{\mathbb{Q}_\star} > 0, b_{22}^{\mathbb{Q}_\star} > 0$, to ensure existence of the change of measure from $\mathbb{P}$ to $\mathbb{Q}$ as well as $\mathbb{P}$ to $\mathbb{Q}_\star$, respectively.

[22]A natural way to look at the dynamics of the exchange rate would start from the assumption that s_t is some twice differentiable function $s(X_t)$ of the state vector X_t with diffusion matrix $\sigma(X_t)$. One could then apply Ito's rule and conclude that the instantaneous volatility of s_t is given by $\nabla s(X_t)\sigma(X_t)$. Unfortunately we do not know the function $s(X_t)$. No-arbitrage gives us relation (B.1), which is revealing about the dynamics, but not the state of the exchange rate. What we can infer from this relation, but only together with our specification (B.10), is that $\nabla s(X_t) = \Sigma$. Footnote 24 also emphasizes how absence of arbitrage provides information about the evolution, but not the state of the foreign exchange rate, from the context of the arbitrage-free CIP.

and $\mathbb{Q}_\star$.[23] For a fixed time horizon $T > t$ it turns out that the *conditional* covariance matrix of $Z_T|Z_t$ is non-singular, in contrast to the instantaneous one. As a consequence of our affine formulation we have that yields and spot predictions based on RA-UIP in eq. (14) are all affine in the state variables Z_t

$$\bar{y}_{t,T} = -\left(A(T-t) + B(T-t)\, Z_t\right) \tag{B.11}$$

$$\bar{y}^\star_{t,T} = -\left(A^\star(T-t) + B^\star(T-t)\, Z_t\right) \tag{B.12}$$

$$\mathbb{E}^{\mathbb{P}}_t\left[s_T\right] = AQ(T-t) + BQ(T-t)\, Z_t \tag{B.13}$$

$$\begin{aligned} \log \mathbb{E}^{\mathbb{Q}_T}_t\left[S_T\right] &= \log \frac{\mathbb{E}^{\mathbb{Q}}_t\left[e^{-\int_t^T \delta_0 + \delta_1 X_s\, ds}\, e^{S_T}\right]}{p_{t,T}} \\ &= \phi(T-t, u) - A(T-t) + \left(\psi(T-t, u) - B(T-t)\right) Z_t \\ &=: \mathcal{A}(T-t) + \mathcal{B}(T-t)\, Z_t, \end{aligned} \tag{B.14}$$

where a bar indicates 'model-implied'. $A(T-t)$, $B(T-t)$ (and $A^\star(T-t), B^\star(T-t)$) in eqs. (B.11) and (B.12) are the solutions $\psi(T-t, 0)$, resp. $\phi(T-t, 0)$ from the ODE in (B.20) with domestic (foreign) $\mathbb{Q}$ parameters; see Appendix B.2 for details.[24] Eq. (B.13) can be computed using formula (B.17) with a selection vector F with non-zero entry only for s, and ϕ and ψ in (B.14) solve the ODE in eq. (B.20) with initial condition $u = (0, 0, 0, 0, 1)$.

B.2 Conditional Moments of Polynomial Processes

It is shown in Cuchiero, Teichmann, and Keller-Ressel (2008) that affine processes such as the one used in the present paper are a subclass of polynomial processes. Polynomial processes are particularly attractive because their conditional moments are polynomials in

[23] For an introduction to affine models and a rigorous treatment of the existence of exponential and polynomial moments, see Filipović and Mayerhofer (2009). In a recent paper Cuchiero, Teichmann, and Keller-Ressel (2008) introduce the class of *polynomial* processes, of which affine diffusion processes are a subclass. For polynomial processes, conditional polynomial moments map to polynomials in the state variables. They can be computed in closed-form according to a formula which is reviewed in Appendix B.2.

[24] It is a tedious, yet rewarding exercise to check that $\mathcal{A}(T-t) + A(T-t) - A^\star(T-t) + (\mathcal{B}(T-t) + B(T-t) - B^\star(T-t)\, Z_t) = S_t$ holds for any Z_t (i.e. whether CIP holds) by investigating the ODE (B.20).

the state variables. The coefficients of the polynomial are determined by the parameters of the process and the time horizon. To be more precise, consider a time-homogeneous (affine) Markov process $X \equiv (X_t)_{t\geq 0, X_0=x_0\in\mathcal{D}}$ living on state space $\mathcal{D} \subset \mathbb{R}^N$. Denote the finite dimensional vector space of all polynomials of degree less than or equal to l by $\mathrm{Pol}_{\leq l}(\mathcal{D})$. An affine process X induces the semigroup

$$P_t f(x) \equiv \mathbb{E}\left[f(X_t)|X_0 = x\right] \in \mathrm{Pol}_{\leq l}(\mathcal{D}) \quad \text{for} \quad f \in \mathrm{Pol}_{\leq l}(\mathcal{D}), \tag{B.15}$$

which maps polynomial moments to polynomials. For affine $(X_t) \in \mathbb{R}^i_+ \times \mathbb{R}^{N-i}$ define

$$\mu(x) \equiv a + b\,x, \quad V(x) \equiv G + H\,x = G + H_1 x_1 + \cdots + H_i x_i, \tag{B.16}$$

where G is a $N \times N$ matrix and H is a $N \times N \times N$ cube. Polynomial moments can be computed using the semigroup's infinitesimal generator

$$\mathcal{A}f(x) = \frac{1}{2}\sum_{j,l=1}^{N} V_{jl}(x)\frac{\partial^2 f(x)}{\partial x_j \partial x_l} + \sum_{j=1}^{N} \mu_j(x)\frac{\partial f(x)}{\partial x_j}.$$

Choose a basis $E \equiv< e_1, \ldots, e_q >$ of $\mathrm{Pol}_{\leq k}(\mathcal{D})$, where $q = \dim \mathrm{Pol}_{\leq k}(\mathcal{D}) = \sum_{j=0}^{k}\binom{N-1+j}{j}$, and a selection vector $F \equiv< f_1, \ldots, f_q >$. Conditional polynomial moments are then computed according to

$$P_t f = F\, e^{t\,A}\, E^\top, \tag{B.17}$$

where $A = (a_{ij})_{i,j=1,\ldots,q}$ is defined implicitly through

$$\mathcal{A}e_i = \sum_{j=1}^{q} a_{ij} e_j. \tag{B.18}$$

For discounted exponential moments we have that

$$\mathbb{E}_t\left[e^{-\int_t^T \delta_0 + \delta_X\, X_s\, ds}\, e^{uX_T}\right] = e^{\phi(\tau,u)+\psi(\tau,u)X_t}, \tag{B.19}$$

where $\phi(\tau, u)$ and $\phi(\tau, u)$ solve a system of Riccati equations (cf. Filipović and Mayerhofer, 2009) with $\tau \equiv T - t$

$$\begin{aligned} \frac{d\psi(\tau,u)}{d\tau} &= -\delta_X - b\,\psi(\tau,u) + \frac{1}{2}\psi(\tau,u)^\top H\,\psi(\tau,u), \quad \psi(0,u) = u \\ \frac{d\phi(\tau,u)}{d\tau} &= -\delta_0 + a\,\psi(\tau,u) + \frac{1}{2}\psi(\tau,u)^\top G\,\psi(\tau,u), \quad \phi(0,u) = 0. \end{aligned} \tag{B.20}$$

For $u = (0, 0, \ldots, 0)$ we recognize the bond price equation, for which we will suppress the second argument in the coefficients.

B.3 Second Moment of Forecast Errors

Assuming $L \leq T$ we are interested in model-implied covariance structure of the error terms from eq. (14)

$$\begin{aligned} \mathbb{C}ov_t\left[\varepsilon_{t,T}, \varepsilon_{t,L}\right] &= \mathbb{C}ov_t\left[s_T, s_L\right] \\ &= \mathbb{E}_t^{\mathbb{P}}\left[s_T\, s_L\right] - \mathbb{E}_t^{\mathbb{P}}\left[s_T\right]\mathbb{E}_t^{\mathbb{P}}\left[s_L\right] \\ &= \underbrace{\mathbb{E}_t^{\mathbb{P}}\left[\mathbb{E}_L^{\mathbb{P}}\left[s_T\right] s_L\right]}_{I.} - \underbrace{\mathbb{E}_t^{\mathbb{P}}\left[s_T\right]\mathbb{E}_t^{\mathbb{P}}\left[s_L\right]}_{II.} \end{aligned}$$

$II.$ can be computed according to eq. (B.13). For $I.$ we get

$$\begin{aligned} \mathbb{E}_t^{\mathbb{P}}\left[\mathbb{E}_L^{\mathbb{P}}\left[s_T\right]\, s_L\right] &= \mathbb{E}_t^{\mathbb{P}}\left[\left(AQ(T-L) + BQ(T-L)Z_L\right) s_L\right] \\ &= AQ(T-L)\left(AQ(L-t) + BQ(L-t)Z_t\right) + BQ(T-L)\,\mathbb{E}_t^{\mathbb{P}}\left[Z_L s_L\right] \end{aligned}$$

The vector of cross-sectional moments $\mathbb{E}_t^{\mathbb{P}}\left[Z_L s_L\right]$ is a quadratic form in the state variables and can be computed using formula (B.17).

C Model Estimation

Let $\theta = \left\{a_1^{\mathbb{P}}, a_2^{\mathbb{P}}, \ldots, \delta_{13}^{\star}, \delta_{14}^{\star}\right\}$ be the set of parameters governing the dynamics of the processes driving the economy described in Appendix B.1; in total we have 45 parameters. The model ought to fit zero-coupon yields of the respective currencies as well as predict changes in the log spot rate. The observed data are seven US zero-yields $y = \{y_t\}$, where $y_t = (y_{t,t+1m}, y_{t,t+3m}, y_{t,t+6m}, y_{t,t+1y}, y_{t,t+2y}, y_{t,t+3y}, y_{t,t+4y})^{\top} D$ and $D \equiv \mathrm{diag}(12, \cdots, 1/4)$, and seven foreign zero-yields $y^{\star}$ with the same maturities. We assume that the yields are observed with cross-sectionally and intertemporally i.i.d. errors $\varrho_t \sim \mathrm{MVN}(0, \Sigma_{\varrho^{\star}})$, and $\varrho_t^{\star} \sim \mathrm{MVN}(0, \Sigma_{\varrho^{\star}})$, respectively. Let $\bar{y} = \{\bar{y}_t\}$, where $\bar{y}_t = (\bar{y}_{t,t+1m}, \ldots \bar{y}_{t,t+4y})^{\top} D$, and $\bar{y}^{\star} = \{\bar{y}_t^{\star}\}$, where $\bar{y}_t^{\star} = (\bar{y}_{t,t+1m}^{\star}, \ldots \bar{y}_{t,t+4y}^{\star})^{\top} D$ denote the corresponding model-implied quantities from eqs. (B.11)–(B.12). We assume that the pricing errors enter additively into the pricing equations

$$y_t = \bar{y}_t + \varrho_t \tag{C.1}$$

$$y_t^{\star} = \bar{y}_t^{\star} + \varrho_t^{\star}. \tag{C.2}$$

For parsimony we assume that the covariance matrices of the errors are diagonal with parameters ζ, and $\zeta^{\star}$, where $\Sigma_{\varrho} = \mathrm{diag}(\zeta, \cdots, \zeta)$, and $\Sigma_{\varrho^{\star}} = \mathrm{diag}(\zeta^{\star}, \cdots, \zeta^{\star})$, respectively. The predictive equation (14) is implemented for horizons of 1 day, 1 week, 1 month, 3 months, 1 year and 4 years. With $\varepsilon_t \equiv (\varepsilon_{t,t+1d}, \ldots \varepsilon_{t,t+4y})$ we specify the covariance matrix of the forecast errors $\Sigma_{\varepsilon_t} \equiv \mathbb{V}_t^{\mathbb{P}}[\varepsilon_t]$ in the predictive regression such that it reflects the cross-sectional covariance structure of our model. Appendix B.3 derives how it can be computed as a function of state variables and the model parameters. We specify the errors to be normally distributed with mean zero and these model-implied covariances.

Estimation is performed using Bayesian methodology where we employ the usual uninformed prior

$$\pi(\theta_i) \propto \begin{cases} \mathbb{1}_{\{\theta_i \text{ admissible}\}} & \theta_i \in \mathbb{R} \\ \frac{\mathbb{1}_{\{\theta_i \text{ admissible}\}}}{\theta_i} & \theta_i \in \mathbb{R}_+ \end{cases}. \tag{C.3}$$

We sample from the posterior distribution

$$p(X, \theta \mid y, y^\star, s) \propto p(y, y^\star \mid Z, \theta)\, p(Z \mid \theta) \pi(\theta) \tag{C.4}$$

by in turn drawing from (Hammersley and Clifford, 1970)

$$p(X \mid y, y^\star, s, \theta) \propto p(y, y^\star \mid Z, \theta)\, p(Z \mid \theta)$$

and

$$p(\theta \mid y, y^\star, s, X) \propto p(y, y^\star \mid Z, \theta)\, p(Z \mid \theta) \pi(\theta)$$

using MCMC methods.[25] Denote with $\phi(x; \upsilon, \Omega)$ the density of the multivariate normal distribution with mean υ and covariance Ω. We approximate transition densities $p(Z_t \mid Z_{t-1}, \theta)$ with a normal distribution, which has been shown previously to perform well in likelihood-based inference.[26] With this approximation we obtain $p(Z \mid \theta)$ in density (C.4)

$$p(Z \mid \theta) = \prod_{n=2}^{N} p(Z_n \mid Z_{n-1}, \theta) \approx \prod_{n=2}^{N} \phi\left(Z_t; \mathbb{E}^{\mathbb{P}}\left[Z_n \mid Z_{n-1}\right], \mathbb{V}_t^{\mathbb{P}}\left[Z_n \mid Z_{n-1}\right]\right)$$

and also

$$p(y, y^\star \mid Z, \theta) = \prod_{n=1}^{N} \phi\left(y_n; \bar{y}_n, \Sigma_\varrho\right) \phi\left(y_n^\star; \bar{y}_n^\star, \Sigma_{\varrho^\star}\right) \phi\left(\varepsilon_n; 0, \Sigma_{\varepsilon_n}\right).$$

Due to the high-dimensional and nonlinear nature of our problem we sample the parameters and the latent states using Metropolis-Hastings steps with random-walk proposal densities. By construction this proposal yields autocorrelated draws. We therefore generate 10,000,000 samples of which we discard the first 5,000,000. From the remaining draws we take every 1,000th draw to obtain (approximately) independent draws from the posterior distribution.

[25] A comprehensive reference for MCMC methods in finance is Johannes and Polson (2009).

[26] We approximate $p(Z_t \mid Z_{t-1}, \theta) \approx \phi(Z_t; \mathbb{E}^{\mathbb{P}}\left[Z_t \mid Z_{t-1}\right], \mathbb{V}_t^{\mathbb{P}}\left[Z_t \mid Z_{t-1}\right])$, where mean $\mathbb{E}^{\mathbb{P}}\left[Z_t \mid Z_{t-1}\right]$ and covariance $\mathbb{V}_t^{\mathbb{P}}\left[Z_t \mid Z_{t-1}\right]$ are the first two (true) conditional moments, which are again computed using formula (B.17) in Appendix B.2. An alternative likelihood approximation is developed in Aït-Sahalia (2008). It has been used successfully in connection with affine term structure models in Aït-Sahalia and Kimmel (2009) and with affine equity models in Aït-Sahalia and Kimmel (2007) within a maximum likelihood context. An adaption of MCMC algorithms to use closed-form likelihood approximations within Bayesian methodology is presented in Stramer, Bognar, and Schneider (2009).

As a representative, we present the parameter estimates for the two-country model of the US and Japan estimated using the zero yields of the two countries and the JPY spot exchange rate applying the procedure described in Section 3.2. Table C.1 at the end of this document reports point estimates and corresponding 95 percent confidence intervals. Point estimates are computed as the draw from the posterior distribution with minimal L1 distance to the other draws. Confidence intervals are computed from the empirical posterior distribution. Parameter estimates for the other currencies are available upon request.

D Block Bootstrap Procedure

We use the tests proposed by Clark and West (2007) and Giacomini and White (2006) to assess the predictive ability of our model. The null hypothesis of the *CW* test is that the nested models have equal (adjusted) mean squared errors; under the alternative hypothesis the larger model exploits (additional) predictive information and has a lower mean squared error. The null hypothesis of the *GW* test is that the models have equal conditional predictive ability; the test statistic is based on the series of squared prediction error differentials. The bootstrap procedure described below computes how often an economy in which there is no predictability would produce as much predictability as found in actual data.

Specifically, we impose a data generating process of no predictability. We consider an overlapping block resampling scheme which can handle serial correlation and also heteroscedasticity; see e.g. Künsch (1989), Politis and Romano (1992), Hall, Horowitz, and Jing (1995), Politis and White (2004), Patton, Politis, and White (2009). Let y_t be the dependent variable and $\widehat{y}_t$ the prediction of that variable, and proceed as follows:

1. Run the regression of form $y_t = \alpha + \beta \widehat{y}_t + \varepsilon_t$, compute the *CW* and *GW* test-statistics, and set $\tilde{y}_t = \hat{\varepsilon}_t$.

2. Form an artificial sample $S_t^* = (y_t^*, \widehat{y}_t^*)$ by randomly sampling, with replacement, b overlapping blocks of length l from the sample $(\tilde{y}, \widehat{y}_t)$.

3. Run the regression $y_t^* = \alpha^* + \beta^* \widehat{y_t^*} + \varepsilon_t^*$, and compute the CW^* and GW^* test-statistics.

4. Repeat steps 2 and 3 5,000 times.

5. Determine the one-sided p-values of the two test-statistics by computing the proportional number of times that $CW^* > CW$ and $GW^* > GW$.

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Table 1: Descriptive Statistics of Exchange Rate Changes

Log exchange rate returns are based on non-overlapping observations for horizons up to 1 month and on monthly frequency for horizons of 3 months and beyond. All figures are annualized. N denotes the number of observations. AC$(T-t)$ denotes the autocorrelation for the lag being equal to the horizon. The sample periods are October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY.

	1 day	1 week	1 month	3 months	1 year	4 years
AUD						
N	2632	527	120	120	120	120
Mean	0.0042	0.0061	0.0065	0.0025	0.0020	0.0089
Std Dev	0.1048	0.1012	0.0962	0.1009	0.1193	0.1311
Skewness	−0.1745	−0.3243	−0.1458	−0.0555	0.0146	−0.1600
Kurtosis	6.3153	3.6681	2.9266	2.9369	2.5032	1.6528
AC$(T-t)$	0.0050	−0.0063	0.1390	0.0776	0.1909	−0.2202
CAD						
N	2989	598	136	136	136	136
Mean	0.0049	0.0055	0.0045	0.0041	0.0077	0.0168
Std Dev	0.0592	0.0601	0.0586	0.0600	0.0607	0.0817
Skewness	0.1058	0.0807	0.2504	0.6931	0.7804	0.3879
Kurtosis	5.2707	3.7735	3.1555	3.9702	3.2926	1.5467
AC$(T-t)$	−0.0065	−0.0902	0.0951	0.0312	0.2476	0.3284
CHF						
N	3954	791	180	180	180	180
Mean	0.0234	0.0230	0.0239	0.0222	0.0138	0.0122
Std Dev	0.1134	0.1151	0.1131	0.1174	0.1100	0.0929
Skewness	0.1323	−0.0520	−0.0506	−0.1887	0.0220	−0.3004
Kurtosis	4.8408	3.9049	3.4349	2.8253	2.2132	2.2479
AC$(T-t)$	0.0098	−0.0370	0.0899	−0.0864	−0.0380	−0.5532
DEM-EUR						
N	3954	791	180	180	180	180
Mean	0.0167	0.0165	0.0170	0.0151	0.0077	0.0072
Std Dev	0.1043	0.1061	0.1044	0.1109	0.1080	0.1042
Skewness	0.0218	−0.1681	−0.1188	−0.1078	0.1037	−0.1305
Kurtosis	4.6383	3.7138	3.6990	2.6264	2.0779	1.9378
AC$(T-t)$	0.0149	−0.0175	0.1361	−0.0764	0.0383	−0.4480
GBP						
N	3954	791	180	180	180	180
Mean	0.0109	0.0105	0.0109	0.0114	0.0071	0.0067
Std Dev	0.0897	0.0960	0.0960	0.0983	0.0876	0.0693
Skewness	−0.1615	−0.8473	−1.0329	−1.1814	−0.3579	−0.0093
Kurtosis	5.6681	8.8557	6.5192	8.1755	3.5891	1.9332
AC$(T-t)$	0.0587	0.0211	0.0772	−0.0528	−0.0481	−0.4144
JPY						
N	3954	791	180	180	180	180
Mean	0.0209	0.0208	0.0222	0.0212	0.0207	0.0106
Std Dev	0.1103	0.1178	0.1118	0.1206	0.1054	0.0879
Skewness	0.5513	0.9126	0.4784	0.3244	−0.4827	0.2869
Kurtosis	7.5747	8.6013	4.0976	3.5989	2.5784	3.3482
AC$(T-t)$	0.0282	−0.0728	0.0927	−0.0405	0.0882	−0.6362

Table 2: Descriptive Statistics of Yield Differentials

The results are based on non-overlapping observations for horizons up to 1 month and on monthly frequency for horizons of 3 months and beyond. All figures are annualized. N denotes the number of observations. AC($T-t$) denotes the autocorrelation for the lag being equal to the horizon. The sample periods are October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY.

	1 day	1 week	1 month	3 month	1 year	4 years
AUD						
N	2632	527	120	120	120	120
Mean	−0.0131	−0.0131	−0.0131	−0.0128	−0.0119	−0.0100
Std Dev	0.0010	0.0023	0.0048	0.0084	0.0162	0.0214
Skewness	−0.3051	−0.3061	−0.3349	−0.3190	−0.2261	−0.0673
Kurtosis	1.7540	1.7549	1.7769	1.7445	1.6728	1.4826
AC($T-t$)	0.9994	0.9969	0.9852	0.9630	0.7311	−0.7606
CAD						
N	2989	598	136	136	136	136
Mean	−0.0007	−0.0007	−0.0007	−0.0009	−0.0016	−0.0022
Std Dev	0.0007	0.0017	0.0035	0.0060	0.0110	0.0163
Skewness	0.3745	0.3753	0.3558	0.3259	0.2664	−0.2217
Kurtosis	2.4859	2.4823	2.5052	2.5196	2.5426	2.1107
AC($T-t$)	0.9981	0.9929	0.9639	0.8690	0.4487	−0.5120
CHF						
N	3954	791	180	180	180	180
Mean	0.0112	0.0112	0.0112	0.0113	0.0130	0.0184
Std Dev	0.0016	0.0035	0.0074	0.0125	0.0214	0.0247
Skewness	−0.5354	−0.5367	−0.5466	−0.5492	−0.4674	−0.4514
Kurtosis	2.4617	2.4654	2.493	2.5214	2.5549	3.0721
AC($T-t$)	0.9995	0.9978	0.9900	0.9650	0.7859	−0.4463
DEM-EUR						
N	3954	791	180	180	180	180
Mean	−0.0033	−0.0033	−0.0032	−0.0028	−0.0008	0.0034
Std Dev	0.0016	0.0035	0.0074	0.0125	0.0213	0.0235
Skewness	−0.7088	−0.7087	−0.7178	−0.6905	−0.5951	−0.4391
Kurtosis	2.5272	2.5248	2.5444	2.5393	2.5838	2.9784
AC($T-t$)	0.9998	0.9988	0.9936	0.9730	0.7332	−0.4389
GBP						
N	3954	791	180	180	180	180
Mean	−0.0239	−0.0239	−0.0238	−0.0235	−0.0209	−0.0134
Std Dev	0.0014	0.0031	0.0065	0.0109	0.0181	0.0228
Skewness	−0.7826	−0.7731	−0.7769	−0.7799	−0.7458	−0.5988
Kurtosis	2.4733	2.4506	2.4422	2.4927	2.6521	2.8806
AC($T-t$)	0.9991	0.9964	0.9859	0.9549	0.6958	−0.0064
JPY						
N	3954	791	180	180	180	180
Mean	0.0262	0.0262	0.0263	0.0269	0.0292	0.0333
Std Dev	0.0015	0.0034	0.0071	0.0121	0.0221	0.0319
Skewness	−0.1771	−0.1774	−0.1777	−0.1353	−0.0510	−0.1614
Kurtosis	1.7206	1.7215	1.7298	1.6821	1.6267	1.8823
AC($T-t$)	0.9997	0.9981	0.9918	0.9745	0.7942	−0.1129

Table 3: Fama Regressions

The table shows the results from estimating, by ordinary least squares, the Fama regression (5), $\Delta s_{t,T} = \alpha + \beta(y_{t,T} - y^{\star}_{t,T}) + \eta_{t,T}$, for the horizons indicated in the column headers. Values in parentheses are asymptotic autocorrelation and heteroscedasticity consistent standard errors following Newey and West (1987). $t[\beta = 1]$ is the t-statistic for testing $\beta = 1$. R^2 is the in-sample coefficient of determination. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. The results are based on non-overlapping observations for horizons up to 1 month and on monthly frequency for horizons of 3 months and beyond. The sample periods are October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY.

	1 day	1 week	1 month	3 months	1 year	4 years
AUD						
α	-0.0003	-0.0013	-0.0057^{*}	-0.0176^{**}	-0.0582^{**}	0.0052
se(α)	(0.0002)	(0.0008)	(0.0032)	(0.0069)	(0.0269)	(0.1364)
β	-5.5010^{***}	-5.6732^{***}	-5.6021^{***}	-5.5060^{***}	-5.0384^{***}	-0.7535
se(β)	(1.9883)	(1.9086)	(1.7643)	(1.8159)	(1.3612)	(1.2085)
$t[\beta = 1]$	[-3.27]	[-3.50]	[-3.74]	[-3.58]	[-4.44]	[-1.45]
R^2	0.0029	0.0166	0.0787	0.2097	0.4709	0.0151
CAD						
α	0.0000	0.0001	0.0002	0.0004	0.0026	0.0635
se(α)	(0.0001)	(0.0003)	(0.0015)	(0.0035)	(0.0102)	(0.0765)
β	-3.4228^{**}	-3.4443^{**}	-2.8355^{**}	-2.9106^{***}	-3.0959^{***}	-0.4018
se(β)	(1.4524)	(1.4718)	(1.4214)	(1.0993)	(0.9108)	(1.2704)
$t[\beta = 1]$	[-3.05]	[-3.02]	[-2.70]	[-3.56]	[-4.50]	[-1.10]
R^2	0.0019	0.0091	0.0288	0.0852	0.3144	0.0065
CHF						
α	0.0002^{**}	0.0008	0.0035	0.0098	0.032	0.1296^{***}
se(α)	(0.0001)	(0.0006)	(0.0027)	(0.0086)	(0.0273)	(0.0423)
β	-1.4813	-1.419	-1.4412	-1.3672	-1.3929	-1.0922
se(β)	(1.1402)	(1.1567)	(1.1429)	(1.2871)	(1.0399)	(0.7152)
$t[\beta = 1]$	[-2.18]	[-2.09]	[-2.14]	[-1.84]	[-2.30]	[-2.93]
R^2	0.0004	0.0019	0.0089	0.0211	0.0736	0.0845
DEM-EUR						
α	0.0001	0.0003	0.0012	0.0032	0.0064	0.0419
se(α)	(0.0001)	(0.0005)	(0.0023)	(0.0059)	(0.0204)	(0.0768)
β	-0.6817	-0.6919	-0.8104	-1.0400	-1.6348	-0.9614
se(β)	(1.0521)	(1.0695)	(1.0568)	(1.131)	(1.1785)	(0.8931)
$t[\beta = 1]$	[-1.60]	[-1.58]	[-1.71]	[-1.80]	[-2.24]	[-2.20]
R^2	0.0001	0.0005	0.0033	0.0138	0.1035	0.0471
GBP						
α	0.0001	0.0003	0.0013	0.0041	0.0131	0.1118^{*}
se(α)	(0.0001)	(0.0007)	(0.0031)	(0.0068)	(0.0245)	(0.0632)
β	0.2833	0.2496	0.1932	0.1842	0.2879	1.5835^{***}
se(β)	(1.0295)	(1.1018)	(1.1073)	(1.5776)	(1.3194)	(0.4945)
$t[\beta = 1]$	[-0.70]	[-0.68]	[-0.73]	[-0.52]	[-0.54]	[1.18]
R^2	0.0000	0.0001	0.0002	0.0004	0.0036	0.2715
JPY						
α	0.0003	0.0014	0.0066^{*}	0.0205^{**}	0.0933^{***}	0.1764
se(α)	(0.0002)	(0.0009)	(0.0036)	(0.0082)	(0.0155)	(0.1174)
β	-1.9643^{*}	-1.9416	-2.0449^{*}	-2.152^{**}	-2.4908^{***}	-1.0064^{*}
se(β)	(1.1533)	(1.2303)	(1.1661)	(1.0076)	(0.7335)	(0.6056)
$t[\beta = 1]$	[-2.57]	[-2.39]	[-2.61]	[-3.13]	[-4.76]	[-3.31]
R^2	0.0007	0.0031	0.017	0.0467	0.2731	0.1331

Table 4: Pricing Errors for US and Foreign Yields

The table reports the root mean squared errors in basis points for the domestic US yields (Panel A) and the respective foreign yields (Panel B). The rows indicate the model estimated, the column headers indicate the yield maturities. The results are based on daily observations for the sample periods October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY.

Panel A: US Yields

	1 month	3 months	6 months	1 year	2 years	3 years	4 years
Model AUD	3	3	6	10	10	12	19
Model CAD	3	3	6	10	9	12	17
Model CHF	3	3	5	10	9	11	17
Model DEM-EUR	3	3	5	11	10	11	18
Model GBP	3	3	5	11	11	11	19
Model JPY	9	11	10	15	34	51	66

Panel B: Foreign Yields

	1 month	3 months	6 months	1 year	2 years	3 years	4 years
Model AUD	6	7	8	15	17	24	37
Model CAD	7	8	9	16	23	35	54
Model CHF	7	8	8	12	25	37	49
Model DEM-EUR	8	10	10	15	33	47	64
Model GBP	9	9	10	23	34	50	74
Model JPY	4	3	4	10	12	11	19

Table 5: Fama Conditions

The table shows the relevant covariances ($\mathbb{C}ov^{\mathbb{P}}$) and variances ($\mathbb{V}^{\mathbb{P}}$) for the Fama-conditions, eq. (9). The values are annualized and multiplied by 10,000. $\widehat{\lambda}_{t,T}$ is the model-implied risk premium, $\Delta\widehat{s}_{t,T}$ denotes the model predicted exchange rate return. The results are based on non-overlapping observations for horizons up to 1 month and on monthly frequency for horizons of 3 months and beyond. The sample periods are October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY.

	1 day	1 week	1 month	3 months	1 year	4 years
AUD						
$\mathbb{C}ov^{\mathbb{P}}\left[\widehat{\lambda}_{t,T}, \Delta\widehat{s}_{t,T}\right]$	-0.83	-3.32	-11.39	-29.02	-77.58	-70.87
$\mathbb{V}^{\mathbb{P}}\left[\Delta\widehat{s}_{t,T}\right]$	0.76	2.99	9.96	25.07	66.02	61.27
CAD						
$\mathbb{C}ov^{\mathbb{P}}\left[\widehat{\lambda}_{t,T}, \Delta\widehat{s}_{t,T}\right]$	-0.19	-0.55	-2.17	-5.80	-16.97	-34.68
$\mathbb{V}^{\mathbb{P}}\left[\Delta\widehat{s}_{t,T}\right]$	0.17	0.45	1.73	4.62	13.84	31.49
CHF						
$\mathbb{C}ov^{\mathbb{P}}\left[\widehat{\lambda}_{t,T}, \Delta\widehat{s}_{t,T}\right]$	-0.37	-1.37	-4.79	-12.34	-31.07	-25.40
$\mathbb{V}^{\mathbb{P}}\left[\Delta\widehat{s}_{t,T}\right]$	0.33	1.19	4.00	9.98	23.18	17.11
DEM-EUR						
$\mathbb{C}ov^{\mathbb{P}}\left[\widehat{\lambda}_{t,T}, \Delta\widehat{s}_{t,T}\right]$	-0.24	-1.17	-4.70	-12.98	-36.80	-36.85
$\mathbb{V}^{\mathbb{P}}\left[\Delta\widehat{s}_{t,T}\right]$	0.19	0.93	3.70	10.10	28.15	28.26
GBP						
$\mathbb{C}ov^{\mathbb{P}}\left[\widehat{\lambda}_{t,T}, \Delta\widehat{s}_{t,T}\right]$	-0.06	-0.30	-1.17	-3.24	-12.79	-23.47
$\mathbb{V}^{\mathbb{P}}\left[\Delta\widehat{s}_{t,T}\right]$	0.08	0.40	1.52	4.04	14.97	30.92
JPY						
$\mathbb{C}ov^{\mathbb{P}}\left[\widehat{\lambda}_{t,T}, \Delta\widehat{s}_{t,T}\right]$	-0.40	-1.92	-7.33	-18.77	-49.27	-64.57
$\mathbb{V}^{\mathbb{P}}\left[\Delta\widehat{s}_{t,T}\right]$	0.33	1.57	5.80	14.43	36.90	52.87

Table 6: Regressions of Excess Returns on Expected Excess Returns

The table shows the results from estimating, by ordinary least squares, the regression (18), $ER_{t,T} = \alpha' + \beta' \widehat{ER}_{t,T} + \eta'_{t,T}$, for the horizons indicated in the column headers. Values in parentheses are block-bootstrapped standard errors. $t[\beta' = 1]$ is the t-statistic for testing $\beta' = 1$. R^2 is the in-sample coefficient of determination. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. The results are based on non-overlapping observations for horizons up to 1 month and on monthly frequency for horizons of 3 months and beyond. The sample periods are October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY.

	1 day	1 week	1 month	3 months	1 year	4 years
AUD						
α'	0.0000	0.0001	0.0006	−0.0002	−0.0022	−0.0008
se(α')	(0.0001)	(0.0005)	(0.0020)	(0.0056)	(0.0199)	(0.0768)
β'	0.6991***	0.7899***	0.8429***	0.9949***	1.0710***	0.9674***
se(β')	(0.2205)	(0.2321)	(0.2410)	(0.2557)	(0.2696)	(0.3398)
$t[\beta' = 1]$	[-1.36]	[-0.91]	[-0.65]	[-0.02]	[0.26]	[-0.10]
R^2	0.0040	0.0225	0.0974	0.3024	0.6115	0.4348
CAD						
α'	0.0000	0.0000	−0.0002	−0.0005	0.0006	−0.0033
se(α')	(0.0001)	(0.0003)	(0.0012)	(0.0032)	(0.0075)	(0.0325)
β'	0.6147**	1.0648***	0.9415***	0.9376***	1.1202***	1.0536***
se(β')	(0.2775)	(0.2748)	(0.2897)	(0.2723)	(0.1904)	(0.2182)
$t[\beta' = 1]$	[-1.39]	[0.24]	[-0.20]	[-0.23]	[0.63]	[0.25]
R^2	0.0023	0.0211	0.0687	0.1680	0.5861	0.6246
CHF						
α'	0.0000	0.0002	0.0008	0.0015	−0.0032	0.0079
se(α')	(0.0001)	(0.0006)	(0.0022)	(0.0061)	(0.0179)	(0.0347)
β'	0.5346**	0.6430**	0.6969**	0.9428***	0.8991***	0.9614***
se(β')	(0.2601)	(0.3070)	(0.3003)	(0.2702)	(0.3120)	(0.2878)
$t[\beta' = 1]$	[-1.79]	[-1.16]	[-1.01]	[-0.21]	[-0.32]	[-0.13]
R^2	0.0010	0.0052	0.0228	0.1004	0.2545	0.3457
DEM-EUR						
α'	0.0001	0.0005	0.0021	0.0051	0.0031	0.0118
se(α')	(0.0001)	(0.0005)	(0.0021)	(0.0059)	(0.0170)	(0.0458)
β'	0.5342*	0.5285*	0.6591**	0.8313***	0.8960***	0.8595***
se(β')	(0.2996)	(0.3071)	(0.3004)	(0.2830)	(0.3048)	(0.2930)
$t[\beta' = 1]$	[-1.55]	[-1.54]	[-1.13]	[-0.60]	[-0.34]	[-0.48]
R^2	0.0008	0.0038	0.0246	0.0941	0.2949	0.3018
GBP						
α'	0.0001	0.0006	0.0022	0.0051	0.0164	0.0185
se(α')	(0.0001)	(0.0006)	(0.0023)	(0.0063)	(0.0146)	(0.0351)
β'	−0.0149	0.2083	0.4410	0.6531	0.5146	0.7822***
se(β')	(0.7139)	(0.7057)	(0.6928)	(0.6235)	(0.4889)	(0.2461)
$t[\beta' = 1]$	[-1.42]	[-1.12]	[-0.81]	[-0.56]	[-0.99]	[-0.88]
R^2	0.0000	0.0001	0.0026	0.0157	0.0470	0.3362
JPY						
α'	0.0000	0.0000	0.0008	0.0027	0.0021	−0.0055
se(α')	(0.0001)	(0.0006)	(0.0023)	(0.0054)	(0.0175)	(0.0231)
β'	0.0194	0.1862	0.3990	0.6050***	0.9421***	0.9678***
se(β')	(0.2869)	(0.2739)	(0.2708)	(0.2293)	(0.2110)	(0.1356)
$t[\beta' = 1]$	[-3.42]	[-2.97]	[-2.22]	[-1.72]	[-0.27]	[-0.24]
R^2	0.0000	0.0006	0.0117	0.0587	0.4191	0.7516

Table 7: Ability to Predict Excess Returns

The table reports results related to the predictive ability of our model as compared to the UIP and RW benchmarks. Hit-ratios (HR) are calculated as the proportion of times the sign of the excess return is correctly predicted by the model. $R2 = 1 - MSE_M/MSE_B$ where MSE_M denotes the mean squared prediction error of the model and MSE_B that of the benchmark. CW and GW denote the test-statistics of Clark and West (2007) and Giacomini and White (2006) as described in section 3.3. The one-sided p-values of the test-statistics in square brackets are obtained from the block bootstrap procedure described in Appendix D which accounts for autocorrelation and heteroscedasticity. The results are based on non-overlapping observations for horizons up to 1 month and on monthly frequency for horizons of 3 months and beyond. The sample periods are October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY.

	Model vs. UIP						Model vs. RW					
	1d	1w	1m	3m	1y	4y	1d	1w	1m	3m	1y	4y
AUD												
HR	0.5410	0.5769	0.6417	0.7250	0.8333	0.7667	0.5410	0.5769	0.6417	0.7250	0.8333	0.7667
$R2$	0.0041	0.0231	0.1005	0.3062	0.6159	0.4760	0.0029	0.0163	0.0701	0.2445	0.5318	0.4080
p-value[CW]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.152]	[0.051]	[0.027]	[<0.01]	[<0.01]	[<0.01]	[0.228]	[0.075]	[0.051]	[<0.01]	[<0.01]	[<0.01]
CAD												
HR	0.5433	0.5585	0.5662	0.5882	0.7500	0.6250	0.5433	0.5585	0.5662	0.5882	0.7500	0.6250
$R2$	0.0024	0.0213	0.0693	0.1694	0.5939	0.6879	0.0011	0.0153	0.0469	0.1120	0.4966	0.6559
p-value[CW]	[0.025]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.091]	[<0.01]	[0.015]	[0.010]	[<0.01]	[<0.01]
p-value[GW]	[0.460]	[0.173]	[0.216]	[0.043]	[0.014]	[<0.01]	[0.633]	[0.254]	[0.321]	[0.079]	[0.039]	[<0.01]
CHF												
HR	0.5311	0.5373	0.5889	0.6500	0.8167	0.7778	0.5311	0.5373	0.5889	0.6500	0.8167	0.7778
$R2$	0.0010	0.0054	0.0238	0.1023	0.2546	0.3551	0.0003	0.0024	0.0104	0.0711	0.1608	0.2505
p-value[CW]	[0.028]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.090]	[0.033]	[0.032]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.414]	[0.242]	[0.108]	[0.021]	[0.010]	[<0.01]	[0.482]	[0.229]	[0.223]	[0.038]	[0.030]	[<0.01]
DEM-EUR												
HR	0.5387	0.5626	0.5722	0.6222	0.7889	0.7667	0.5387	0.5626	0.5722	0.6222	0.7889	0.7667
$R2$	0.0010	0.0045	0.0277	0.1000	0.2987	0.3049	0.0004	0.0016	0.0141	0.0631	0.1826	0.2127
p-value[CW]	[0.011]	[0.013]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.052]	[0.056]	[0.018]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.340]	[0.211]	[0.037]	[<0.01]	[<0.01]	[<0.01]	[0.368]	[0.239]	[0.072]	[0.015]	[0.024]	[<0.01]
GBP												
HR	0.5187	0.5196	0.5444	0.6056	0.6667	0.6778	0.5187	0.5196	0.5444	0.6056	0.6667	0.6778
$R2$	0.0006	0.0027	0.0138	0.0468	0.1346	0.5397	-0.0000	-0.0001	0.0009	0.0116	0.0361	0.5102
p-value[CW]	[0.174]	[0.104]	[0.062]	[0.029]	[0.029]	[<0.01]	[0.214]	[0.133]	[0.107]	[0.023]	[<0.01]	[<0.01]
p-value[GW]	[0.103]	[0.405]	[0.261]	[0.126]	[0.104]	[<0.01]	[0.037]	[0.340]	[0.332]	[0.128]	[0.041]	[<0.01]
JPY												
HR	0.5228	0.5221	0.5556	0.6333	0.7611	0.9722	0.5228	0.5221	0.5556	0.6333	0.7611	0.9722
$R2$	0.0000	0.0006	0.0118	0.0593	0.4221	0.7917	-0.0008	-0.0029	-0.0053	0.0164	0.2936	0.6725
p-value[CW]	[0.419]	[0.206]	[0.048]	[<0.01]	[<0.01]	[<0.01]	[0.916]	[0.715]	[0.345]	[0.073]	[<0.01]	[<0.01]
p-value[GW]	[0.541]	[0.263]	[0.124]	[0.137]	[<0.01]	[<0.01]	[0.149]	[0.145]	[0.181]	[0.224]	[<0.01]	[<0.01]

Table 8: Decomposing Foreign Exchange Risk Premia

This table reports means and standard deviations (in parentheses) of annualized foreign exchange risk premia and their components, i.e. the pure currency risk component and the component that accounts for the fact that interest rates are stochastic; for the decomposition see Section 2.3, in particular eq. (17). The descriptives are calculated from daily model estimates of the risk premia. The sample periods are October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY.

	1 day	1 week	1 month	3 months	1 year	4 years
AUD						
Risk Premium	−0.0167	−0.0172	−0.0174	−0.0170	−0.0153	−0.0198
	(0.1521)	(0.1338)	(0.1185)	(0.1127)	(0.0951)	(0.0460)
- Pure currency risk	−0.0229	−0.0234	−0.0235	−0.0230	−0.0217	−0.0268
	(0.1519)	(0.1337)	(0.1184)	(0.1125)	(0.0949)	(0.0458)
- Impact of stochastic rates	0.0062	0.0062	0.0060	0.0061	0.0064	0.0070
	(0.0011)	(0.0011)	(0.0011)	(0.0011)	(0.0011)	(0.0011)
CAD						
Risk Premium	−0.0083	−0.0080	−0.0074	−0.0071	−0.0076	−0.0188
	(0.0746)	(0.0592)	(0.0550)	(0.0524)	(0.0455)	(0.0315)
- Pure currency risk	−0.0101	−0.0098	−0.0092	−0.0089	−0.0095	−0.0211
	(0.0750)	(0.0597)	(0.0556)	(0.0529)	(0.0460)	(0.0317)
- Impact of stochastic rates	0.0019	0.0019	0.0018	0.0018	0.0019	0.0022
	(0.0007)	(0.0007)	(0.0006)	(0.0006)	(0.0005)	(0.0004)
CHF						
Risk Premium	−0.0066	−0.0061	−0.0055	−0.0059	−0.0049	0.0083
	(0.1045)	(0.0921)	(0.0837)	(0.0790)	(0.0662)	(0.0316)
- Pure currency risk	−0.0149	−0.0144	−0.0135	−0.0139	−0.0130	0.0001
	(0.1062)	(0.0941)	(0.0858)	(0.0812)	(0.0685)	(0.0330)
- Impact of stochastic rates	0.0082	0.0083	0.0080	0.0079	0.0081	0.0082
	(0.0046)	(0.0046)	(0.0044)	(0.0043)	(0.0039)	(0.0026)
DEM-EUR						
Risk Premium	0.0082	0.0074	0.0050	0.0014	−0.0067	−0.0012
	(0.0886)	(0.0882)	(0.0850)	(0.0819)	(0.0705)	(0.0356)
- Pure currency risk	0.0023	0.0015	−0.0007	−0.0043	−0.0124	−0.0068
	(0.0898)	(0.0894)	(0.0861)	(0.0830)	(0.0715)	(0.0359)
- Impact of stochastic rates	0.0059	0.0059	0.0057	0.0056	0.0057	0.0056
	(0.0026)	(0.0026)	(0.0025)	(0.0024)	(0.0022)	(0.0016)
GBP						
Risk Premium	−0.0187	−0.0191	−0.0211	−0.0239	−0.0230	−0.0200
	(0.0396)	(0.0393)	(0.0374)	(0.0371)	(0.0371)	(0.0222)
- Pure currency risk	−0.0229	−0.0233	−0.0252	−0.0279	−0.0268	−0.0233
	(0.0394)	(0.0391)	(0.0374)	(0.0371)	(0.0370)	(0.0217)
- Impact of stochastic rates	0.0042	0.0042	0.0041	0.0040	0.0039	0.0033
	(0.0020)	(0.0019)	(0.0019)	(0.0018)	(0.0016)	(0.0010)
JPY						
Risk Premium	0.0397	0.0386	0.0343	0.0272	0.0112	0.0220
	(0.1116)	(0.1102)	(0.1031)	(0.0970)	(0.0814)	(0.0465)
- Pure currency risk	0.0316	0.0305	0.0264	0.0192	0.0021	0.0104
	(0.1095)	(0.1080)	(0.1010)	(0.0948)	(0.0796)	(0.0467)
- Impact of stochastic rates	0.0081	0.0081	0.0079	0.0080	0.0090	0.0115
	(0.0036)	(0.0035)	(0.0033)	(0.0033)	(0.0032)	(0.0025)

Table 9: Correlations of Expected Excess Returns with Financial and Fundamental Variables

The table presents contemporaneous correlations of expected excess returns with the VIX signed by the yield differential ($sVIX_t$), the 1-year log changes in US industrial production (ΔIP_t) and US narrow money supply (ΔNM_t), the observable fundamentals, $\Delta OF_t = (\Delta NM_t - \Delta NM_t^\star) - (\Delta IP_t - \Delta IP_t^\star)$, and the 1-year log change in CPI deflated private consumption per capita in the US ($s\Delta CO_t$). *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. The significance is judged by block-bootstrapped standard errors which are not reported. The results are based on non-overlapping observations for horizons up to 1 month and on monthly frequency for horizons of 3 months and beyond. The sample periods are October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY. Analysis involving the VIX start in January 1990.

	1 day	1 week	1 month	3 months	1 year	4 years
AUD						
$sVIX_t$	−0.5660***	−0.6286***	−0.6920***	−0.7630***	−0.7652***	−0.7294***
ΔIP_t			−0.4781***	−0.4893***	−0.5140***	−0.5685***
ΔNM_t			0.5938***	0.6333**	0.6676**	0.7798***
$s\Delta CO_t$				−0.7786***	−0.7745***	−0.7205***
ΔOF_t				0.6842***	0.6917***	0.6874***
CAD						
$sVIX_t$	−0.6124***	−0.8098***	−0.7999***	−0.8228***	−0.7873***	−0.6773***
ΔIP_t			−0.4795***	−0.5106***	−0.5637***	−0.5963***
ΔNM_t			0.7905***	0.7866***	0.7638***	0.6802***
$s\Delta CO_t$				−0.7759***	−0.7456***	−0.6507***
ΔOF_t			0.6792***	0.7169***	0.6691***	0.5499***
CHF						
$sVIX_t$	−0.3596***	−0.3803**	−0.4101**	−0.4660**	−0.5536**	−0.5375**
ΔIP_t			−0.3064**	−0.3661**	−0.4571***	−0.5114***
ΔNM_t			0.7553***	0.8150***	0.8727***	0.8781***
$s\Delta CO_t$				−0.4251**	−0.5357***	−0.5795***
ΔOF_t				0.3212	0.3740*	0.3400
DEM-EUR						
$sVIX_t$	−0.7623***	−0.7666***	−0.7344***	−0.7632***	−0.7838***	−0.7780***
ΔIP_t			−0.3703***	−0.4055***	−0.4306***	−0.4414***
ΔNM_t			0.8243***	0.8471***	0.8625***	0.8393***
$s\Delta CO_t$				−0.7359***	−0.7575***	−0.7771***
ΔOF_t			0.6314***	0.6793***	0.6948***	0.6471***
GBP						
$sVIX_t$	−0.1359	−0.1489	−0.0979	−0.1888	−0.2985	−0.1588
ΔIP_t			−0.2387	−0.3201*	−0.3439**	−0.1180
ΔNM_t			0.6558***	0.7111***	0.6389***	0.3176*
$s\Delta CO_t$				−0.2767	−0.3706***	−0.4078***
ΔOF_t			0.6656***	0.7138***	0.6161***	0.3726**
JPY						
$sVIX_t$	−0.5929***	−0.5915***	−0.5715***	−0.5963*	−0.6547**	−0.7079
ΔIP_t			−0.5746***	−0.5796***	−0.5794***	−0.5071***
ΔNM_t			0.6986***	0.7472***	0.6938***	0.3707**
$s\Delta CO_t$				−0.3126*	−0.3256**	−0.2732
ΔOF_t			0.5626***	0.6124***	0.6548***	0.6039***

Table 10: Regressions of Expected Excess Returns on Financial and Fundamental Variables

The table presents results of regressing expected excess returns on our proxies for global risk (VIX signed with the yield differential, $sVIX_t$), exchange rate fundamentals (observable fundamentals, $\Delta OF_t = (\Delta NM_t - \Delta NM_t^{\star}) - (\Delta IP_t - \Delta IP_t^{\star})$), US consumption growth ($s\Delta CO_t$), and combinations thereof. Numbers in parentheses are block bootstrapped standard errors. R^2 is the in-sample coefficient of determination. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. The results are based on non-overlapping observations for horizons up to 1 month and on quarterly frequency for horizons of 3 months and beyond. The sample periods are October 12, 1994 to October 10, 2008 for AUD; June 1, 1993 to October 10, 2008 for CAD; and September 18, 1989 to October 10, 2008 for CHF, DEM-EUR, GBP, and JPY. Analysis involving the VIX start in January 1990.

	Global Risk				FX Fundamentals			Global Risk and FX Fundamentals						Cons. Growth		Cons. Growth and FX Fundamentals			
	1 day	1 month	3 months	1 year	1 month	3 months	1 year	1 month		3 months		1 year		3 months	1 year	3 months		1 year	
	$sVIX_t$	$sVIX_t$	$sVIX_t$	$sVIX_t$	ΔOF_t	ΔOF_t	ΔOF_t	$sVIX_t$	ΔOF_t	$sVIX_t$	ΔOF_t	$sVIX_t$	ΔOF_t	$s\Delta CO_t$	$s\Delta CO_t$	$s\Delta CO_t$	ΔOF_t	$s\Delta CO_t$	ΔOF_t
AUD																			
coeff	−0.0113***	−0.0329***	−0.1018***	−0.3340***		0.2317***	0.7662***			−0.0754***	0.1398***	−0.2457***	0.4667***	−1.0633***	−3.4601***	−0.8508***	0.1638***	−2.7508***	0.5468***
se	(0.0028)	(0.0026)	(0.0240)	(0.0827)		(0.0883)	(0.2924)			(0.0178)	(0.0410)	(0.0583)	(0.1332)	(0.2749)	(0.9684)	(0.1610)	(0.0372)	(0.5943)	(0.1244)
R^2	0.3204	0.4789	0.5822	0.5855		0.4681	0.4784			0.7134	0.7134	0.7221	0.7221	0.6062	0.5998	0.8161	0.8161	0.8183	0.8183
CAD																			
coeff	−0.0060***	−0.0178***	−0.0523***	−0.1685***	0.0592***	0.1916***	0.6020***	−0.0139***	0.0364***	−0.0399***	0.1170***	−0.1307***	0.3580***	−0.4978***	−1.6104***	−0.4005***	0.1141***	−1.3173***	0.3438***
se	(0.0010)	(0.0012)	(0.0107)	(0.0394)	(0.0130)	(0.0532)	(0.2042)	(0.0025)	(0.0080)	(0.0072)	(0.0295)	(0.0305)	(0.1199)	(0.1580)	(0.5296)	(0.0977)	(0.0265)	(0.4168)	(0.1178)
R^2	0.3750	0.6399	0.6770	0.6199	0.4613	0.5139	0.4477	0.7830	0.7830	0.8308	0.8308	0.7469	0.7469	0.6020	0.5559	0.7798	0.7798	0.6983	0.6983
CHF																			
coeff	−0.0057**	−0.0164***	−0.0527**	−0.2076**		0.0683	0.2622			−0.0449*	0.0278	−0.1816**	0.0938	−0.5555**	−2.3086***	−0.4766**	0.0472	−2.0195***	0.1727
se	(0.0026)	(0.0028)	(0.0244)	(0.0957)		(0.0519)	(0.1821)			(0.0240)	(0.0457)	(0.0854)	(0.1407)	(0.2294)	(0.7005)	(0.2069)	(0.0479)	(0.6075)	(0.1596)
R^2	0.1293	0.1681	0.2172	0.3065		0.1032	0.1399			0.2298	0.2298	0.3196	0.3196	0.1807	0.2870	0.2263	0.2263	0.3431	0.3431
DEM-EUR																			
coeff	−0.0088***	−0.0256***	−0.0768***	−0.2649***	0.0807***	0.2502***	0.8567***	−0.0185***	0.0381***	−0.0539***	0.1194**	−0.1888***	0.3969**	−0.7927***	−2.7320***	−0.5827***	0.1292**	−2.0155***	0.4411**
se	(0.0014)	(0.0017)	(0.0112)	(0.0422)	(0.0129)	(0.0433)	(0.1511)	(0.0053)	(0.0147)	(0.0151)	(0.0560)	(0.0507)	(0.1781)	(0.1055)	(0.3528)	(0.1347)	(0.0529)	(0.4161)	(0.1749)
R^2	0.5810	0.5393	0.5824	0.6144	0.3986	0.4615	0.4827	0.5907	0.5907	0.6385	0.6385	0.6694	0.6694	0.5415	0.5738	0.6307	0.6307	0.6664	0.6664
GBP																			
coeff	−0.0011	−0.0023	−0.0135	−0.0837	0.0228***	0.0749***	0.2519***	−0.0012	0.0216***	−0.0114	0.0730***	−0.0765	0.2428***	−0.1557	−0.8123**	−0.1614*	0.0640***	−0.8308***	0.2093***
se	(0.0014)	(0.0016)	(0.0227)	(0.0735)	(0.0035)	(0.0176)	(0.0713)	(0.0032)	(0.0034)	(0.0139)	(0.0153)	(0.0568)	(0.0542)	(0.1262)	(0.3829)	(0.0839)	(0.0160)	(0.2613)	(0.0606)
R^2	0.0185	0.0096	0.0357	0.0891	0.4430	0.5095	0.3796	0.4105	0.4105	0.5137	0.5137	0.4337	0.4337	0.0766	0.1373	0.5023	0.5023	0.4378	0.4378
JPY																			
coeff	−0.0132***	−0.0345***	−0.1033*	−0.3734*	0.0540***	0.1664**	0.5885**	−0.0228**	0.0357***	−0.0656	0.1135**	−0.2468	0.3809**	−0.7193**	−2.4775**	−0.5591**	0.1393*	−1.8981**	0.5036*
se	(0.0043)	(0.0040)	(0.0615)	(0.2009)	(0.0151)	(0.0778)	(0.2672)	(0.0113)	(0.0133)	(0.0606)	(0.0572)	(0.1931)	(0.1903)	(0.3490)	(1.1041)	(0.2820)	(0.0778)	(0.8525)	(0.2688)
R^2	0.3515	0.3266	0.3556	0.4287	0.3165	0.3750	0.4288	0.4245	0.4245	0.4785	0.4785	0.5565	0.5565	0.0977	0.1060	0.3741	0.3741	0.4365	0.4365

Table C.1: JPY Model Parameters

The table shows parameter estimates for the JPY data set. Point estimates are computed as the draw from the posterior distribution with minimal L1 distance to the other draws. Confidence intervals are computed from the empirical posterior distribution.

Parameter	Point Estimate	95% confidence interval	
ζ	0.0025	0.0024	0.0025
$\zeta^\star$	0.0009	0.0009	0.0009
β_1	4.1424	2.7331	4.9026
β_2	11.7506	10.4242	12.4701
γ_1	8.4781	7.0951	10.2487
γ_2	28.1971	24.8945	30.9722
Σ_1	-0.0114	-0.0130	-0.0103
Σ_2	0.0083	0.0079	0.0096
Σ_3	-0.0039	-0.0043	-0.0035
Σ_4	-0.0035	-0.0035	-0.0034
$a_1^{\mathbb{P}}$	0.5231	0.5005	0.5782
$a_2^{\mathbb{P}}$	0.5192	0.5008	0.6168
$a_3^{\mathbb{P}}$	-56.3956	-58.3480	-51.0532
$a_4^{\mathbb{P}}$	29.0699	24.4765	31.9081
$b_{11}^{\mathbb{P}}$	-0.2277	-0.2972	-0.1854
$b_{21}^{\mathbb{P}}$	0.0103	0.0005	0.0748
$b_{22}^{\mathbb{P}}$	-0.6423	-0.6896	-0.6059
$b_{31}^{\mathbb{P}}$	-1.4762	-1.8189	-1.1673
$b_{32}^{\mathbb{P}}$	0.5269	0.2286	0.6306
$b_{33}^{\mathbb{P}}$	-0.0019	-0.0176	0.0000
$b_{41}^{\mathbb{P}}$	-111.1620	-113.7310	-108.6060
$b_{42}^{\mathbb{P}}$	-61.9570	-62.9497	-60.9826
$b_{43}^{\mathbb{P}}$	21.9912	21.6463	22.4145
$b_{44}^{\mathbb{P}}$	-8.6476	-8.7531	-8.5571
$a_1^{\mathbb{Q}}$	2.8375	2.7935	2.8690
$a_2^{\mathbb{Q}}$	9.8192	9.7005	9.9073
$b_{11}^{\mathbb{Q}}$	-0.0008	-0.0029	-0.0000
$b_{21}^{\mathbb{Q}}$	0.0170	0.0137	0.0222
$b_{22}^{\mathbb{Q}}$	-0.2741	-0.2772	-0.2709
$b_{31}^{\mathbb{Q}}$	4.9797	4.9220	5.0458
$b_{32}^{\mathbb{Q}}$	1.7195	1.6984	1.7344
$b_{33}^{\mathbb{Q}}$	-0.6353	-0.6424	-0.6270
$b_{41}^{\mathbb{Q}}$	-111.1400	-111.6350	-110.2580
$b_{42}^{\mathbb{Q}}$	-59.6544	-59.9086	-58.9453
$b_{43}^{\mathbb{Q}}$	21.1715	20.9889	21.2557
$b_{44}^{\mathbb{Q}}$	-7.9357	-7.9568	-7.8205
δ_0	$2.58E{-06}$	$2.67E{-07}$	$3.74E{-05}$
δ_1	$3.29E{-04}$	$3.20E{-04}$	$3.39E{-04}$
δ_2	$1.05E{-03}$	$1.05E{-03}$	$1.06E{-03}$
δ_3	$1.57E{-04}$	$1.55E{-04}$	$1.58E{-04}$
$\delta_0^\star$	$1.00E{-03}$	$9.83E{-04}$	$1.03E{-03}$
$\delta_1^\star$	$1.80E{-03}$	$1.79E{-03}$	$1.80E{-03}$
$\delta_2^\star$	$-7.58E{-05}$	$-7.61E{-05}$	$-7.41E{-05}$
$\delta_4^\star$	$1.10E{-04}$	$1.10E{-04}$	$1.11E{-04}$