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Abstract

We study the properties of foreign exchange risk premiums that can explain the forward bias puzzle, defined as the tendency of high-interest rate currencies to appreciate rather than depreciate. These risk premiums 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 the empirical analysis, we estimate risk premiums using an affine multi-currency term structure model and find that model-implied risk premiums yield unbiased predictions for exchange rate excess returns. While interest rate risk affects the level of risk premiums, the time variation in excess returns is almost entirely driven by currency risk. Furthermore, risk premiums are closely related to global risk aversion, countercyclical to the state of the economy, and 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 premiums 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 premiums must be at the heart of this observation.

In this paper, we re-examine the relation between the term structure of interest rates and exchange rates by expressing the link 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. We start from noting that forward rates are generally biased predictors of future spot exchange rates, and expected spot rate changes comprise a time-varying risk premium in addition to the forward 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

[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

pioneering work of Backus et al. (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 exact functional form of the term structure of foreign exchange risk premiums in closed form. This allows us to jointly match the term structures of interest rates and the term structure of foreign exchange risk premiums in the estimation procedure. Using daily data for six major US dollar exchange rates over the last 20 years, we generate model-implied exchange rate expectations and risk premiums for horizons ranging from 1 day to 4 years.

The empirical results reveal that the RA-UIP model is capable of identifying time-varying risk premiums that match observed exchange rate behavior. In particular, they 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 the model generates unbiased predictions for exchange rate excess returns. This implies that accounting for risk premiums 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 exchange rate 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 excess returns can almost entirely be explained by the premium for currency risk.

We also provide empirical evidence that risk premiums 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 premiums in dollar exchange rates are countercyclical to the US economy. Overall, a

model estimation is hence based on daily data. However, as discussed below, we relate the model-implied risk premiums to observable economic variables later in the paper to refine our understanding of the drivers of the latent factors.

large part of expected excess returns can be explained by fundamentals deemed relevant in traditional exchange rate models.

Related Literature in More Detail 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 et al. (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 under 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, e.g. Dewachter and Maes (2001), Ahn (2004), Inci and Lu (2004), Mosburger and Schneider (2005), and Anderson et al. (2009), elaborate on the effects of local versus global factors in an international economy.[2] Brandt and Santa-Clara (2002) and Anderson et al. (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 premiums are significantly related to both the estimated volatility of the pricing kernels and the volatility of exchange rates. The estimated risk premiums mostly satisfy the Fama (1984) necessary conditions for explaining the forward

[2] Another recent related article is Leippold and Wu (2007). Instead of using an affine model, they propose a class of multi-currency quadratic models.

bias puzzle, although the puzzle remains in several cases.[3]

The choice of variables and the results from our analysis of the economic drivers of foreign exchange risk premiums is consistent with recent research. Our evidence that expected excess returns are (i) related to global risk aversion is consistent with the flight-to-quality and flight-to-liquidity arguments in Lustig et al. (2010b) and Brunnermeier et al. (2008), (ii) countercyclical to the state of the US economy is in line with e.g. Lustig and Verdelhan (2007), De Santis and Fornari (2008), and Lustig et al. (2010b), and (iii) driven by traditional exchange rate fundamentals is supported by 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 relation 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. We present the results in Section 4 and discuss extensions and robustness checks in Section 5. Section 6 presents empirical evidence that financial and macroeconomic variables are important drivers of the foreign exchange risk premium. Section 7 concludes. The Appendix provides technical details on derivations and some estimation procedures. A separate Internet Appendix reports the parameter estimates in detail and provides additional empirical results related to extensions and robustness checks.

2 Exchange Rates, Interest Rates and No-Arbitrage

This section defines the fundamental relations linking exchange rates and interest rates, and shows the implications of imposing the no-arbitrage condition in this context. This results in the risk-adjusted variants of UIP and FUH, which are shown to imply intuitive

[3] 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 et al. (1993), Bekaert and Hodrick (1993), Bansal et al. (1995), Bekaert (1996), Bekaert et al. (1997), Lustig and Verdelhan (2007), Brunnermeier et al. (2008), Farhi and Gabaix (2008), Jurek (2009), Lustig et al. (2010a), Verdelhan (2010), Bansal and Shaliastovich (2009), and Farhi et al. (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 (2010), and the profitability and economic value of currency speculation, e.g. Burnside et al. (2010), and Della Corte et al. (2009).

properties for the foreign exchange risk premium.

2.1 Uncovered Interest Parity and Forward Unbiasedness

We express exchange rates as domestic currency prices per unity of foreign currency. S_t denotes the spot exchange rate, $F_{t,T}$ is the forward exchange rate for an exchange of currencies at time $T > t$, s_t and $f_{t,T}$ are the corresponding log exchange rates. The domestic and foreign T-period yields of the respective zero bonds are $y_{t,T} \equiv -\log p_{t,T}$ and $y^\star_{t,T} \equiv -\log p^\star_{t,T}$. Assuming risk-neutrality and rational expectations, UIP postulates that the expected exchange rate change must equal the yield differential or equivalently, because Covered Interest Parity (CIP) holds, the forward premium:

$$\mathbb{E}^{\mathbb{P}}_t[\Delta s_{t,T}] = f_{t,T} - s_t = y_{t,T} - y^\star_{t,T},$$

where $\Delta s_{t,T} = s_T - s_t$ and $\mathbb{E}^{\mathbb{P}}_t$ denotes the conditional expectation under the physical probability measure. UIP further implies that excess returns, $rx_{t,T} \equiv s_T - f_{t,T}$, should be unpredictable and it also forms the economic foundation for the FUH that the forward rate should be an unbiased predictor of the future spot exchange rate, $f_{t,T} = \mathbb{E}^{\mathbb{P}}_t[s_T]$. Empirical tests are usually performed by estimating the 'Fama regressions' (Fama, 1984)

$$\Delta s_{t,T} = \alpha + \beta(y_{t,T} - y^\star_{t,T}) + \eta_{t,T}, \tag{1}$$

$$rx_{t,T} = \alpha + \gamma(y_{t,T} - y^\star_{t,T}) + \eta_{t,T}, \tag{2}$$

where $\gamma = \beta - 1$. The null hypotheses that UIP is valid holds if $\alpha = 0$, $\beta = 1$, and $\eta_{t,T}$ is 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'.

Fama (1984) argues that the forward bias may be caused by a time-varying risk premium $\lambda_{t,T}$ that is priced in forward rates, $f_{t,T} = \mathbb{E}^{\mathbb{P}}_t[s_T] + \lambda_{t,T}$. The omission of $\lambda_{t,T}$ in the Fama regressions results in a negative β estimate if

$$\begin{aligned} &\mathbb{C}ov^{\mathbb{P}}\left[\lambda_{t,T}, \mathbb{E}_t^{\mathbb{P}}[\Delta s_{t,T}]\right] < 0 \\ &\left|\mathbb{C}ov^{\mathbb{P}}\left[\lambda_{t,T}, \mathbb{E}_t^{\mathbb{P}}[\Delta s_{t,T}]\right]\right| > \mathbb{V}^{\mathbb{P}}\left[\mathbb{E}_t^{\mathbb{P}}[\Delta s_{t,T}]\right]. \end{aligned} \tag{3}$$

The first condition is that the risk premium's covariance with expected exchange rate changes is negative, the second is that the absolute value of this covariance is greater than the variance of expected changes. However, attempts to explain the forward bias puzzle using risk premiums have only had limited success so far.

2.2 Risk-Adjusted UIP and FUH under No-Arbitrage

We relax the assumption of risk-neutrality and derive risk-adjusted counterparts to the conventional UIP and FUH that endogenize time-varying risk premiums in the spirit of Fama (1984). Since the price of a forward contract changes over time due to both spot rate and interest rate fluctuations, we investigate the relation between spot and forward exchange rates in a no-arbitrage setting with stochastic interest rates. We choose $p_{t,T}$ as the numeraire where the associated probability measure is the T-forward measure $\mathbb{Q}_{\mathbb{T}}$.[4] Combining the no-arbitrage pricing equation with CIP gives

$$F_{t,T} = \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right] = \mathbb{E}_t^{\mathbb{Q}}\left[\frac{d\mathbb{Q}_{\mathbb{T}}}{d\mathbb{Q}} S_T\right]. \tag{4}$$

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 $B_t = e^{\int_0^t r_s\,ds}$, where r is the short rate of interest. Only in the case of deterministic interest rates, the Radon-Nikodym derivative $\frac{d\mathbb{Q}_{\mathbb{T}}}{d\mathbb{Q}} = 1$ and hence $\mathbb{Q}$ equals $\mathbb{Q}_{\mathbb{T}}$. 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).

Under the assumption of rational expectations, taking conditional expectation yields

[4] See for example Björk (2004, p. 355), or Mele (2009, p. 242).

the natural right-hand sides of predictive relations for log exchange rate returns

$$\begin{aligned}\Delta s_{t,T} &= \mathbb{E}_t^{\mathbb{P}}\left[s_T - s_t\right] + \varepsilon_{t,T} \\ &= \mathbb{E}_t^{\mathbb{P}}\left[s_T\right] - \big(\log F_{t,T} - (y_{t,T} - y_{t,T}^{\star})\big) + \varepsilon_{t,T} \\ &= \nu_{t,T} + (y_{t,T} - y_{t,T}^{\star}) + \varepsilon_{t,T}\end{aligned} \tag{5}$$

and excess returns

$$rx_{t,T} = \nu_{t,T} + \varepsilon_{t,T} \tag{6}$$

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 (5), 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 $\nu_{t,T}$. This component $\nu_{t,T}$ drives excess returns and since it 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) as $\lambda_{t,T} = -\nu_{t,T}$. 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

$$\mathbb{E}_t^{\mathbb{P}}\left[s_T\right] = \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[s_T\right] - \Big(\mathbb{E}_t^{\mathbb{Q}}\left[s_T\right] - \mathbb{E}_t^{\mathbb{P}}\left[s_T\right]\Big) - \Big(\mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[s_T\right] - \mathbb{E}_t^{\mathbb{Q}}\left[s_T\right]\Big) \tag{7}$$

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

$$\begin{aligned}\lambda_{t,T} &= \log \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right] - \mathbb{E}_t^{\mathbb{P}}\left[s_T\right] \\ &= \underbrace{\Big(\mathbb{E}_t^{\mathbb{Q}}\left[s_T\right] - \mathbb{E}_t^{\mathbb{P}}\left[s_T\right]\Big)}_{\text{pure currency risk}} + \underbrace{\Big(\log \mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right] - \mathbb{E}_t^{\mathbb{Q}}\left[s_T\right]\Big)}_{\text{impact of stochastic rates}}.\end{aligned} \tag{8}$$

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.[6]

[5] 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. (5). These Jensen terms are considered to be very small in currency markets, though; see e.g. the survey of Engel (1996). For completeness and comparison, we provide analogue derivations without logs in Appendix A.

[6] We provide a formal derivation of Eq. (8) in Appendix A.2.

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

3.1 Affine Multi-Country Term Structure Model

The RA-FUH and RA-UIP expressions derived in the previous section are model-free relations that extend the conventional FUH and UIP in that they are free of risk preferences and consistent with no-arbitrage. To make these relations amenable for empirical work, we employ a parametric framework that allows to evaluate expressions (5) and (8) 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 et al. (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 et al. (2001) for their *completely 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 premiums). We describe the details of the model in the next subsection. However, two extensions deserve to be mentioned here. First, in contrast to Backus et al. (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. (8). 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 premiums. As a result, we can derive the term structure of foreign exchange risk premiums in closed form.

[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 and 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.

3.1.1 A Continuous-Time Model for an International Economy

For the econometric analysis, to put structure on the coefficients and error terms appearing in the predictive equation (5), 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, comprised of latent state variables X_t and the observed log exchange rate s_t: $Z \equiv (Z_t)_{t\geq 0, Z_0 = z_0 \in \mathcal{D}} \equiv (X_{1t}, X_{2t}, X_{3t}, X_{4t}, s_t) \equiv (X_t, s_t)$, 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 four-factor setting for the international economy. To ensure arbitrage-free markets, we start with a relation between the two countries' pricing kernels that ensures consistent pricing

$$\frac{M_t^\star}{M_0^\star} \equiv \frac{S_t}{S_0}\frac{M_t}{M_0}. \tag{9}$$

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 et al. (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 (9), 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,

$$dX_t = (a^{\mathbb{P}} + b^{\mathbb{P}} X_t)dt + \sigma(X_t)dW_t^{\mathbb{P}}, \tag{10}$$

where

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

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. Factors X_1 and X_2 are square-root processes that drive conditional variance. Factors X_3 and X_4 are conditionally Gaussian to accommodate negative correlation between the state variables, which the yield data usually require; see e.g. Dai and Singleton (2000). With a setting comprised only of square-root processes, correlation would be constrained to be positive, both instantaneously and for a fixed time $\tau > 0$. The dynamics of the domestic pricing kernel are

$$\frac{dM_t}{M_t} = -r_t dt - \Lambda(X_t)^{\top} dW_t^{\mathbb{P}}, \tag{12}$$

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

$$\Lambda(x) = \sigma(x)^{-1} \left(a^{\mathbb{P}} + b^{\mathbb{P}} x - (a^{\mathbb{Q}} + b^{\mathbb{Q}} x) \right), \text{ where} \tag{13}$$

$$a^{\mathbb{Q}} \equiv \begin{pmatrix} a_1^{\mathbb{Q}} \\ a_2^{\mathbb{Q}} \\ 0 \\ 0 \end{pmatrix}, \; b^{\mathbb{Q}} \equiv \begin{pmatrix} b_{11}^{\mathbb{Q}} & 0 & 0 & 0 \\ b_{21}^{\mathbb{Q}} & b_{22}^{\mathbb{Q}} & 0 & 0 \\ b_{31}^{\mathbb{Q}} & b_{32}^{\mathbb{Q}} & b_{33}^{\mathbb{Q}} & 0 \\ b_{41}^{\mathbb{Q}} & b_{42}^{\mathbb{Q}} & b_{43}^{\mathbb{Q}} & b_{44}^{\mathbb{Q}} \end{pmatrix}. \tag{14}$$

To unambiguously determine the unconditional mean of the short rate, which is affected by the constant factor loading δ_0 and the unconditional means of X_3 and X_4 in a very similar way, we impose $a_3^{\mathbb{Q}} = a_4^{\mathbb{Q}} = 0$. The parameters $a_1^{\mathbb{Q}}$ and $a_2^{\mathbb{Q}}$ are identified through the behavior of the square-root factors X_1 and X_2, in particular near the boundary of the state space. The market price of risk specification Λ follows Cheridito et al. (2007); it is admissible if $2a_1^{\mathbb{P}} > 1, 2a_2^{\mathbb{P}} > 1$ in addition to the admissibility conditions from Duffie et al.

(2003). For stationarity we impose $b^{\mathbb{P}}_{11} < 0, b^{\mathbb{P}}_{22} < 0$. 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

$$\frac{dM^\star_t}{M^\star_t} = -r^\star_t dt - (\Lambda(X_t)^\top - \Sigma\,\sigma(X_t))\,dW^{\mathbb{P}}_t, \tag{15}$$

where the drift of X_t under $\mathbb{Q}_\star$ (the foreign $\mathbb{Q}$ measure) solves[8]

$$a^{\mathbb{Q}_\star} + b^{\mathbb{Q}_\star} x = a^{\mathbb{P}} + b^{\mathbb{P}} x - \sigma(x)(\Lambda(x)^\top - \Sigma\sigma(x))^\top. \tag{16}$$

Computing the solution to Eqs. (12) and (15) and using Eq. (9) we find that the foreign exchange rate S_t evolves according to

$$\frac{dS_t}{S_t} = (r_t - r^\star_t + \Sigma\,\sigma(X_t)\,\Lambda(X_t))dt + \Sigma\sigma(X_t)dW^{\mathbb{P}}_t, \tag{17}$$

where $\Sigma \equiv (\Sigma_1, \Sigma_2, \Sigma_3, \Sigma_4)$, and $r^\star_t \equiv \delta^\star_0 + \delta^\star_1 X_t$ with $\delta^\star_1 \equiv (\delta^\star_{11}, \delta^\star_{12}, \delta^\star_{13}, \delta^\star_{14})$. The corresponding log dynamics of s_t are then

$$ds_t = \left(r_t - r^\star_t + \Sigma\,\sigma(X_t)\,\Lambda(X_t) - \frac{1}{2}\,\Sigma\,\sigma(X_t)\,\sigma(X_t)^\top\,\Sigma^\top\right)dt + \Sigma\sigma(X_t)dW^{\mathbb{P}}_t, \tag{18}$$

which turn out to be affine in X_t.[9]

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 with only 4 driving Brownian motions. Nevertheless, Z_t constitutes an affine Markov process under probability measures $\mathbb{P}, \mathbb{Q}$, and $\mathbb{Q}_\star$.[10] For a fixed time horizon $T > t$ it turns out that the *conditional* covariance matrix of

[8] In addition to the admissibility conditions, drifts (14) and (16) also satisfy $2a^{\mathbb{Q}}_1 > 1, 2a^{\mathbb{Q}}_2 > 1$, and $2a^{\mathbb{Q}_\star}_1 > 1, 2a^{\mathbb{Q}_\star}_2 > 1$ to ensure existence of the change of measure from $\mathbb{P}$ to $\mathbb{Q}$ as well as $\mathbb{P}$ to $\mathbb{Q}_\star$, respectively. For stationarity $b^{\mathbb{Q}}_{11} < 0, b^{\mathbb{Q}}_{22} < 0$ and $b^{\mathbb{Q}_\star}_{11} < 0, b^{\mathbb{Q}_\star}_{22} < 0$.

[9] 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 (9), 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 (18), is that $\nabla s(X_t) = \Sigma$. The assumed evolution of the foreign pricing kernel in (15) is not the only choice for an admissible pricing kernel, but it is the only way of maintaining affine dynamics of the exchange rate, which greatly improves tractability.

[10] 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 et al. (2008) introduce

$Z_T|Z_t$ is non-singular, in contrast to the instantaneous one. As a consequence of the affine formulation we have that yields and spot predictions based on RA-UIP in Eq. (5) are all affine in the state variables Z_t

$$\bar{y}_{t,T} = -\left(A(T-t) + B(T-t)\, Z_t\right), \tag{19}$$

$$\bar{y}^{\star}_{t,T} = -\left(A^{\star}(T-t) + B^{\star}(T-t)\, Z_t\right), \tag{20}$$

$$\mathbb{E}^{\mathbb{P}}_t \left[s_T\right] = AQ(T-t) + BQ(T-t)\, Z_t, \tag{21}$$

$$\begin{aligned} \log \mathbb{E}^{\mathbb{Q}_{\mathbb{T}}}_t \left[S_T\right] &= \log \frac{\mathbb{E}^{\mathbb{Q}}_t \left[e^{-\int_t^T \delta_0 + \delta_1 X_s\, ds}\, e^{S_T}\right]}{p_{t,T}} \\ &= \phi(T-t,u) - A(T-t) + \left(\psi(T-t,u) - B(T-t)\right) Z_t \\ &\equiv \mathcal{A}(T-t) + \mathcal{B}(T-t)\, Z_t, \end{aligned} \tag{22}$$

where a bar indicates 'model-implied'. $A(T-t)$, $B(T-t)$ (and $A^{\star}(T-t), B^{\star}(T-t)$) in Eqs. (19) and (20) are the solutions $\psi(T-t,0)$ and $\phi(T-t,0)$ from the ODE in (B.6) with domestic (foreign) $\mathbb{Q}$ parameters respectively; see Appendix B.1 for details.[11] Eq. (21) can be computed using formula (B.3) with a selection vector F with non-zero entry only for s, and ϕ and ψ in (22) solve the ODE in Eq. (B.6) with initial condition $u = (0,0,0,0,1)$.

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 set of observables that we need to fit is large. One can therefore think of these driving state variables as a low-dimensional 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 as well as in terms of the conceptual setup. First, our

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.1.

[11] Writing the yield equations (19) and (20) in terms of the enlarged state vector Z instead of X is just a matter of notational convenience as $\frac{\partial B_s(\tau)}{\tau} = 0$ together with the initial condition $B_s(0) = 0$ imply a zero factor loading for any maturity τ. 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.6).

methodological framework is Bayesian, which yields a posterior distribution of both latent state variables and the parameters of the model. Employing 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.[12] Second, we consider the *joint* dynamics of the latent state variables with the exchange rate. The evolution of the exchange rate therefore affects the distribution of the parameters. Third, in the estimation procedure we do not only fit bond yields in the US and the foreign country but simultaneously also match the predictive relation implied by RA-UIP derived in Eq. (5). 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 premiums. Details of the estimation procedure can be found in Appendix D.

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. (5) and (6) using the estimation procedure outlined in the previous section.[13]

As a first step, we check whether the model risk premium fulfills the conditions formulated by Fama (1984), given in Eq. (3): first, the covariance between the model-implied 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.

[12] 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).

[13] 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.

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 the RA-UIP model predicted excess returns $\widehat{rx}_{t,T}$

$$rx_{t,T} = \alpha' + \beta' \widehat{rx}_{t,T} + \eta'_{t,T} \tag{23}$$

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 the 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.[14] In other words, we have a twofold motivation for applying these criteria: first, we gain additional insight on the model's goodness of fit as compared to only considering the R^2 of regression (23). Second, we complement the evidence on the predictability of excess returns by assessing the predictive ability of the model per se as well as relative to the benchmark predictions based on UIP and a random walk (RW) without drift. These results will show whether empirical exchange rate dynamics are more adequately characterized by RA-UIP, UIP or the RW.

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

[14]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).

correctly predicted. The remaining criteria are defined as functions of squared prediction errors of the 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 the model as compared to the benchmark is given by

$$R2 = 1 - \frac{MSE^M}{MSE^B}. \tag{24}$$

Positive values indicate that the model performs better than the benchmark.

The CW test statistic is defined as

$$CW = MSE^B - MSE^M + N^{-1} \sum_{n=1}^{N} \left(\Delta\widehat{rx}^B_{t,T} - \Delta\widehat{rx}_{t,T} \right)^2, \tag{25}$$

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. (25) 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 the empirical analysis, we apply the block bootstrap procedure described in Appendix E 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.[15] The predictions are based on the full time-t information set $\mathcal{F}_t$. Using an $\mathcal{F}_t$-measurable test function h_t, we test the null hypothesis that predictions based on the 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, $\Delta L_T = SE^B_T - SE^M_T$. The test function we use is $h_t = (1, \Delta L_t)^\top$. The GW statistic is given by

$$GW = N \left(N^{-1} \sum_{n=1}^{N} h_t \Delta L_T \right)^\top \widehat{\Omega}_N^{-1} \left(N^{-1} \sum_{n=1}^{N} h_t \Delta L_T \right) \tag{26}$$

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

where $\widehat{\Omega}_N^{-1}$ is a consistent estimate of the variance of $h_t \Delta L_T$.[16] 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 the model risk premiums 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 Indicators at the 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.

[16] 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).

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 (1). 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.[17] As outlined in Section 2.1, the two Fama regressions in Eqs. (1) and (2) contain 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 Estimation Results

In this section, we discuss results related to how well our model fits the US and foreign term structures of interest rates and we give economic interpretations to the latent factors that drive the international economy. Further estimation results (parameter estimates, confidence intervals, and properties of market prices of risk) are reported in detail in the Internet Appendix.

4.3.1 Yield Pricing Errors

In Table 4 we present results that show that the model fits the data reasonably well in that the pricing errors for the term structures are satisfactory. The errors reported in Table 4

[17] 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.

are in the range of other recent studies, e.g. Brennan and Xia (2006) and Anderson et al. (2009), even though the sample periods used in estimation are different. 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 et al. (2001) and Brandt and Santa-Clara (2002). Moreover we discuss how one can conceptually add more countries and present an example of a three-country estimation in Section 5.2. 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 premiums, 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 difference 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 the bilateral estimation effectively delivers a unique US pricing kernel, although the uniqueness is not imposed in the model.

4.3.2 Interpretation of Latent Factors

While we examine the drivers of foreign exchange risk premiums later in Section 6, we now perform a factor rotation to gain insights on the forces behind the state variables governing the international economy. Collin-Dufresne et al. (2008) show that the latent factors underlying single-country affine term structure models can be rotated into variables with unambiguous economic interpretations. Building on the results of Litterman and Scheinkman

(1991), they further show how to obtain model-independent estimates of the state variables, which allows to estimate their globally identifiable representation and facilitates the interpretation of multi-factor models. We perform three rotations and compare the model-implied processes to their corresponding model-free estimates. The results reported below show that the factor dynamics are strongly related to the information in the US yield curve and to the carry factor (i.e. the interest rate differential) between the US and the foreign country. Technical details and resulting factor loadings are given in Appendix C.

With the first rotation, we investigate how the estimated factor dynamics are related to the US term structure expressed in terms of the level of the instantaneous short rate, the slope, and the quadratic variations of both. We start by rotating the third state variable (the first Gaussian) into the level of the US short rate r_t and subsequently define the slope μ_t as the instantaneous drift of r_t. The remaining two state variables are rotated into the quadratic variations of the short rate and of the slope. As a result, we obtain an observable representation of the model in terms of the instantaneous US short rate level (r_t), slope (μ_t), short rate variance (V_t), and slope variance (U_t),

$$\begin{pmatrix} dV_t \\ dU_t \\ dr_t \\ d\mu_t \end{pmatrix} = \left(\begin{pmatrix} \varphi_V \\ \varphi_U \\ \varphi_r \\ \varphi_\mu \end{pmatrix} + \begin{pmatrix} \vartheta_V & \vartheta_{VU} & 0 & 0 \\ \vartheta_{UV} & \vartheta_U & 0 & 0 \\ 0 & 0 & 0 & 1 \\ \vartheta_{\mu V} & \vartheta_{\mu U} & \vartheta_{\mu r} & \vartheta_\mu \end{pmatrix} \begin{pmatrix} V_t \\ U_t \\ r_t \\ \mu_t \end{pmatrix} \right) dt$$
$$+ \begin{pmatrix} c_1 & c_2 & 0 & 0 \\ d_1 & d_2 & 0 & 0 \\ \delta_{11} & \delta_{12} & \delta_{13} & \delta_{14} \\ \varrho_1 & \varrho_2 & \varrho_3 & \varrho_4 \end{pmatrix} \operatorname{diag} \begin{pmatrix} f_0 + f_1 V_t + f_2 U_t \\ g_0 + g_1 V_t + g_2 U_t \\ y_0 + y_1 V_t + y_2 U_t \\ z_0 + z_1 V_t + z_2 U_t \end{pmatrix} \begin{pmatrix} dW_{1t} \\ dW_{2t} \\ dW_{3t} \\ dW_{4t} \end{pmatrix}.$$

We also follow Collin-Dufresne et al. (2008) in estimating the model-free state variables. We perform a principal components analysis (PCA) to obtain the first three principal components of yield levels and express yield curve derivatives (i.e. level and slope) as sums of derivatives of the PCA loading functions. Using maturities of up to one year, we use lower-order polynomials to extrapolate the loading functions down to zero. We then calculate the model-free estimates of the short rate level (L_t) and slope (Sl_t) based on the fitted polynomials.

Table 5 presents the correlations of the model-implied processes and their model-free counterparts in columns labeled Rotation 1. The level correlations range from 98.5% to 99.9% across all countries. The slope correlations are between 82% and 86% for CHF, DEM-EUR, and GBP, between 55% and 61% for AUD and CAD, and 26% for JPY. For the first three, we also find high correlations for the variance processes, 66% to 68% for V_t and 45% to 54% for U_t.[18] Overall, the results show that the information in the US yield curve plays a fundamental role in the international economy, as one would expect for a model of USD exchange rates.

In the second rotation, we again rotate the third state variable into the US short rate, r_t, and then rotate the fourth into the differential of the US and the foreign short rate, $r_t - r_t^*$, to obtain a carry factor. Our motivation to do so is twofold. First, the short rate differential represents the expected instantaneous depreciation under the risk-neutral measure. Second, research on the cross-section of foreign exchange excess returns suggests that the riskiness of different currencies can be understood in terms of a dollar risk factor and a carry risk factor; see e.g. Lustig et al. (2010a). V_t and U_t now represent the quadratic variations of the US short rate level and the level differential. We obtain the model-free estimates for the level differential analogously to those of the US level, and Table 5 reports results in columns labeled Rotation 2. For all countries we consistently find again that we match the US level. For CHF, DEM-EUR, GBP, and JPY we also find high correlations of $r_t - r_t^*$ with the model-free level differential (56% to 74%) and also the related variance processes U_t and $QV_t[L - L^*]$ (45% to 75%). For these countries, the foreign yield curve contains valuable information not contained in the US curve and thus carry risk is an important factor. For AUD and CAD, the information in the foreign term structure seems

[18] A lower correlation of μ_t with Sl_t and the mixed variance results simply reflect that the estimated model also has to match the term structures of foreign yields and foreign exchange risk premiums while the model-free estimates are only based on the US yield curve. As a result, model-implied and model-free estimates exhibit very different correlations to expected depreciation ($\mathbb{E}_t^{\mathbb{Q}}[ds]$) and exchange rate variance ($QV_t[ds]$) and these covariations with exchange rate variables have an impact on and may overlay other correlations. For instance, V_t and U_t also pick up exchange rate variance and are thus highly correlated with $QV_t[ds]$ while there is no common pattern for their model-free counterparts. The covariations with $\mathbb{E}_t^{\mathbb{Q}}[ds]$ play a similar role, the extent of μ_t correlations can be explained analogously, and similar arguments apply to the other rotations below.

less relevant as compared to that in the US curve.

While the second rotation indicates that the short rate differential adds information beyond the US curve, we now check whether the US term structure adds information when the carry factor has been already accounted for. We do so because Lustig et al. (2010a) find that the US risk factor essentially captures average excess returns across currencies. In a setting like ours, the carry for each country potentially already incorporates this information because of its bilateral nature. We thus modify the rotation in that we first rotate the third state variable into $r_t - r_t^*$ and subsequently the fourth into r_t. Table 5 reports results in columns labeled Rotation 3. We find that the level differential has a correlation of 84% for CAD and more than 96% for all other countries. The correlations of the US level range from 1% to 17%, which suggests that the carry factor already comprises most, but not all of the information contained in the US curve.

Fig. 1 plots the US risk factors and carry risk factors implied by the model and their model-independent counterparts. Overall, the results show that both the US term structure as well as the carry between the US and the foreign country are driving forces behind the latent factor international economy. These results are consistent with recent studies on the cross-section of currency returns; however, in our bilateral setting, the carry factor conveys most of the information.

4.4 Model Evaluation

To evaluate the RA-UIP model we employ the criteria described in Section 3.3. The empirical results reveal that the model predictions are unbiased and have higher accuracy than the UIP and RW benchmarks.

4.4.1 Fama Conditions and Unbiasedness of Model Predictions

We first verify whether the model risk premium fulfills the conditions formulated by Fama (1984), as described in Eq. (3), 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 6. The results show that both conditions are fulfilled for all currencies except the GBP. Specifically, for the

GBP the first condition (negative covariance) is satisfied 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 the model.

Table 7 presents results for regression (23) by reporting parameter estimates along with block-bootstrapped standard errors in parentheses as well as t-statistics for the null hypothesis of unbiasedness $\beta' = 1$.[19] 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 described 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 horizon increases. 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 estimates of β' are not statistically different from unity 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.

[19] 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 heteroskedasticity. 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.

4.4.2 Predictability of Excess Returns

In Table 8, we present results for the predictive ability criteria discussed in Section 3.3. The HR, $R2$, CW, and GW measures allow us to gain insight on the model's goodness of fit as compared to only considering the R^2 of the predictive regression. Furthermore, we complement previous evidence on the predictability of excess returns based on the model per se and as compared to the benchmark predictions based on UIP and the RW.

The HR indicates that the 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 is achieved for the 1-year and 4-year horizons with the largest values across currencies ranging from 63% to 97%.[20] There is evidence that the 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. (24), indicate that the 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 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.[21] In other words, the mean-squared prediction errors of the model are much smaller than those of the benchmarks providing another piece of evidence that the RA-UIP model fits the empirical behavior of exchange rates better than UIP and 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 E. The CW p-values generally decrease with the prediction horizon and indicate that the model predictions

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

[21] 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. 389) or Boudoukh et al. (2006). Note that we have a different predictor and different dependent variable for each horizon.

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 the 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 premiums that drive (excess) returns.[22]

4.5 Decomposing Foreign Exchange Risk Premiums

Following the derivations of the RA-FUH and RA-UIP in Section 2.2, we show in Eq. (8) 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 9 displays descriptive statistics for estimated risk premiums 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 premiums.

These results suggest that the variation in foreign exchange risk premiums - and hence deviations from UIP constituting the forward bias puzzle - are largely driven by the pure

[22]The finding that no-arbitrage improves predictions has similarly been documented in the term structure literature, see e.g. Ang and Piazzesi (2003), Christensen et al. (2010), Diez de los Rios (2009) and Almeida and Vicente (2008).

currency risk component. We redo the empirical model evaluation analysis in Section 4.4 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 premiums may be relevant in other contexts, for example assessing the profitability of currency speculation, which we do not investigate in this paper.

5 Extensions and Robustness Checks

We perform various extensions and robustness checks to further validate the model and to support the empirical findings. We first show that models with a smaller number of latent factors are not capable of jointly matching the term structures of interest rates and foreign exchange risk premiums. Second, we provide evidence from models for more than two countries. We also show that extending the information set by currency options does not qualitatively change the results, and finally that our conclusions are not affected by the recent financial crisis. Detailed empirical results are given in the Internet Appendix.

5.1 Smaller Models with Two or Three Factors

In our setting, the international economy is driven by four latent factors. In this Section we investigate smaller models and report pricing errors, predictive regression estimates, and predictive ability statistics for models with three factors in Tables A.8 to A.10 and for models with two factors in Tables A.11 to A.13.

We find that models with three factors also produce risk premiums that have predictive ability but at the expense of substantially larger yield pricing errors. Furthermore, the differences of the RMSEs of US yields across models (even across non-JPY models) are as high as 36 basis points (this is 18 times the maximum RMSE difference in the four factor model). This raises concerns about the ability to estimate a unique US pricing kernel. The two factor model produces, as expected, less predictability, higher pricing errors, and the RMSEs of US yields deviate by more than 170 basis points across models.

The results illustrate that there is a substantial tradeoff between jointly fitting the term structures of domestic and foreign interest rates and that of foreign exchange risk premiums. As compared to the standard model with four factors, smaller models appear to be overstrained in accomplishing this task.

5.2 Multiple Exchange Rates

As mentioned above, we estimate bilateral models for country pairs because it is computationally infeasible to estimate one model for the US and all foreign countries at once. As a consequence, the latent process that implicitly prices the US term structure can change across estimations. While the results in Section 4.3.1 suggest that differences are small, we discuss here how to extend the framework to multiple exchange rates such that a unique US pricing kernel is ensured ex-ante. As an example, we present a three-country model estimation.

Indexing each foreign country with $n = 1, ..., N$ the bilateral relations and dynamics established versus the US remain unchanged, but additionally consistent pricing between all foreign economies has to be ensured. This means that RA-UIP does not only imply the predictive relation in Eq. (5) for each country n against the US but also for changes in the log cross-rates between all foreign countries. For the cross rate $s^{j,n}$, expressed as currency j price per unit of currency n, the respective relation is given by

$$\Delta s^{j,n}_{t,T} = \Delta s^{n}_{t,T} - \Delta s^{j}_{t,T} = (\nu^{n}_{t,T} - \nu^{j}_{t,T}) + (y^{j}_{t,T} - y^{n}_{t,T}) + \varepsilon^{j,n}_{t,T} \tag{27}$$

which shows that cross-rate returns depend on the foreign countries' yields and the differential of their USD risk premiums.

No-arbitrage requires that the relation defined in Eq. (9) holds for all foreign countries, which allows to specify the dynamics of each foreign country's pricing kernel and its USD exchange rate as in Section 3.1.1. Consistent pricing across foreign economies additionally requires to rule out triangular arbitrage between all countries

$$\frac{M^j_t}{M^j_0} = \left[\frac{S^j_t}{S^j_0}\frac{S^n_0}{S^n_t}\right]\frac{M^n_t}{M^n_0}$$

where the term in $[\cdot]$ represents the cross-rate for exchanging currencies of countries n and j. Taking logs and limits implies the arbitrage-free dynamics of the log cross-rate $ds_t^{j,n} = ds_t^n - ds_t^j$. Since ds_t^n and ds_t^j are specified as ds_t in Section 3.1.1 we can compute the corresponding cross-rate predictions.

In estimation, one has to jointly match the term structure of interest rates in all countries, the term structure of USD foreign exchange risk premiums for each foreign country, and the (implicit) term structure of risk premiums in cross-rates. That is, first, the pricing equation has to be extended such that it comprises yields of all N foreign countries. Second, the predictive relation for USD exchange rate returns in Eq. (5) has to be implemented for all countries. Third, the predictive relation for the cross-rates in Eq. (27) has to be matched for all combinations of foreign countries. With this extended setup, the estimation procedure follows the routine described in Appendix D.

While the concept is straightforward, an empirical implementation of an N-country model involves serious numerical and computational difficulties. The specification of our two-country model is five-dimensional (four latent factors and the observed exchange rate) and has 45 parameters. With each additional foreign economy the model grows at least by one additional observed exchange rate. Since four factors are already a parsimonious choice in the two-country setting, increasing the number of factors would be desirable when adding countries. However, adding exchange rates and factors exacerbates the curse of dimensionality, thereby impeding computational feasibility and eventually making an empirical implementation virtually impossible.[23]

We implement a six-dimensional three-country model (four latent factors and two exchange rates) and discuss two estimations involving the US and Switzerland. In the first estimation we use the DEM-EUR (which behaves as judged by our estimation very similar to the CHF) as the third currency and report results in Table A.14. The predictive regression and predictive accuracy results resemble the patterns of the bilateral setting but they are less pronounced in terms of significance: we find that model-implied risk premiums are

[23] Our Markov-Chain Monte Carlo estimation requires billions of evaluations of matrix exponentials (see formula (B.3)) and matrix inverses (for the transition density approximations and the predictive density). As the size of the coefficients of these quantities grows faster than exponentially with the number of state variables, even switching from six to seven would render the estimation procedure prohibitively slow.

unbiased predictors and significantly different from zero for horizons of 3 months or longer and that the model's ability to predict excess returns is higher than that of the UIP and RW benchmarks. The downside, however, is a sizeable increase in yield pricing errors. For the US, where we have now imposed a unique pricing kernel ex-ante, the yield RMSEs range from 7 basis points to 111 basis points, i.e. they are more than double the RMSEs of the bilateral models at the short end and more than six times larger at the long end.

In the second estimation, we use the JPY (which appears to be more different from the CHF than the other currencies) as the third currency and report results in Table A.15. Again, the predictability results exhibit similar but less pronounced patterns and the yield pricing errors are higher as compared to the bilateral models, in particular for the CHF.

Overall, the three-country results suggest that four factors are not enough, in this multiple exchange rates setting, to jointly model three term structures of interest rates as well as the two corresponding term structures of USD risk premiums. Given that larger models are beyond computational feasibility, it appears reasonable to use bilateral models that effectively deliver a unique pricing kernel (rather than impose it ex-ante) but fit yields and exchange rate dynamics more accurately. However, for the purposes of this paper, it is important to emphasize that the unbiasedness of model-implied risk premiums reported in our core results is robust to using a larger model that imposes a unique US pricing kernel.

5.3 Information in Currency Options

One issue that has arisen in the literature on affine term structure models is that bonds may be insufficient to span fixed income markets and that derivatives may be needed to fully identify pricing kernels.[24] In our model, exchange rate dynamics are driven by the difference in the innovations of two pricing kernels. In the international economy there is no source of risk that exclusively affects exchange rates and hence currency derivatives combine the information embedded in domestic and foreign fixed income derivatives.[25] To analyze whether currency options convey additional information about foreign exchange

[24]See the work of Collin-Dufresne and Goldstein (2002) on unspanned stochastic volatility and the subsequent literature building on their work.

[25]That is, since all factors affect exchange rate as well as domestic and foreign interest rate dynamics, currency derivatives can be hedged/replicated using domestic and foreign fixed income derivatives.

risk premiums we rely on the concept of model-free implied variance (MFIV).

Britten-Jones and Neuberger (2000) show that MFIV equals the expected realized variance under the risk neutral measure. MFIV is fully determined by current option prices and defined as

$$MFIV_{t,T} = \frac{2}{T-t}\left[\int_0^{F_{t,T}} \frac{P_{t,T}(K)}{p_{t,T}K^2}dK + \int_{F_{t,T}}^{\infty} \frac{C_{t,T}(K)}{p_{t,T}K^2}dK\right]$$

where $P_{t,T}(K)$ and $C_{t,T}(K)$ are the respective time-t prices of T-period European put and call options with strike K.[26] To calculate MFIV we use daily currency option data obtained from JP Morgan comprising 1-month implied volatilities for five points, which is standard in currency markets (Carr and Wu, 2007): at-the-money forward (ATMF), 10-delta call, 10-delta put, 25-delta call, and 25-delta put.[27] To calculate implied volatilities and option prices for other strikes, we follow the suggestions of Jiang and Tian (2005).

To incorporate the information conveyed by MFIV, we augment the estimation procedure to require that the model-implied expectation of realized variance matches MFIV. We assess whether MFIV has additional information content for foreign exchange risk premiums by comparing estimation results with and without currency options. For all currencies, the sample period is January 24, 1996 to October 10, 2008, except for the DEM-EUR series, for which the sample starts on January 1, 1998. Our empirical analysis suggests that conditioning on the information in currency options does not have a material effect on how well the model matches foreign exchange risk premiums. In general, when we regress realized excess returns on model predictions from both estimations, the slope coefficients in Eq. (2) and the R^2s are very similar; see Tables A.16 and A.18. The predictive accuracy of both models as compared to the UIP and RW benchmarks is very similar as well; see Tables A.17 and A.19. These results suggest that the specification of our model is flexible enough to capture the variance dynamics of exchange rates and hence, for the purpose

[26] Jiang and Tian (2005) discuss how to inter- and extrapolate when only a finite range of strike prices is available and show that resulting approximation errors are small. They also demonstrate that the MFIV concept is still valid if the underlying asset price process has jumps and they provide evidence that MFIV contains more information than other volatility predictors. For a recent application of the MFIV concept to foreign exchange markets see Della Corte et al. (2010).

[27] Since the data provides implied volatilities and deltas, but not prices directly, we infer strike prices from deltas and implied volatilities and calculate option prices using Garman and Kohlhagen (1983). Note that in FX markets the convention is to multiply put deltas by -100 and call deltas by 100.

of this paper, it is not necessary to additionally condition on the information in currency options.

5.4 Sample excluding the Financial Crisis

As mentioned above, we bootstrap zero yields from money market and swap rates based on the argument put forward by Feldhütter and Lando (2008) that these are the best parsimonious proxy for riskless rates. Due to the recent financial crisis this choice may not be innocuous because the rates may be confounded with credit risk. We therefore repeat the empirical analysis for a sample that excludes the financial crisis by only using data until the end of 2006. We present yield errors, predictive regression estimates, and predictive ability statistics in Tables A.20 to A.22. The results are quantitatively very similar and qualitatively identical to those reported for the full sample.

6 Drivers of the Risk Premium

The above results provide strong empirical support for the existence of time-varying risk premiums 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 at the CBOE, which is highly correlated with similar volatility indexes in other countries; see e.g. Lustig et al. (2010b). Furthermore, the VIX can also be viewed as a proxy for funding liquidity constraints, noted in Brunnermeier et al. (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 premiums on high yield currencies while they accept lower (or more negative) risk premiums on low yield currencies, consistent with 'flight-to-quality' and 'flight-to-liquidity' arguments.[28]

[28] 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 et al. (2008).

Recent research suggests that risk premiums on US exchange rates are countercyclical to the US economy, similar to risk premiums in other markets; see e.g. Lustig and Verdelhan (2007), De Santis and Fornari (2008), and Lustig et al. (2010b). 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 the model risk premium is countercyclical, the relation between expected excess returns and output growth should be negative, whereas the relation 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 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 relation between these fundamentals and expected excess returns should be positive.

Table 10 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 premiums 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 11. 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 it ranges 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 the 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 in which 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 the GBP data set. Also, finding that the CHF's link to observable fundamentals is weak but that its link to global risk is strong seems consistent with Switzerland being viewed as a 'safe haven' and primarily as a destination for flight-to-quality.

7 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 premiums 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 a multi-currency context to explicitly account for these properties of forward rates and embedded risk premiums. We take the model to US exchange rate data and find that estimated model expectations and risk premiums 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 premiums can be sufficient to resolve the forward bias puzzle without additionally requiring departures from rational expectations.

Furthermore, we provide empirical evidence that risk premiums 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, and that they also depend on macroeconomic variables (output growth, money supply growth, consumption growth) such that risk premiums in dollar exchange rates are countercyclical to the US economy.

We disentangle the risk premiums into compensation for currency risk and interest rate risk. We find that the time 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 premiums 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. (5) and (6) 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.

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

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

with $\nu_{t,T} = \left(\mathbb{E}_t^{\mathbb{P}}\left[S_T\right]/\mathbb{E}_t^{\mathbb{Q}_{\mathbb{T}}}\left[S_T\right] - 1\right)e^{(y_{t,T}-y^{\star}_{t,T})}$. 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 relation in Eq. (7) is formally established from

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

B Technical Details Related to the Model

B.1 Conditional Moments of Polynomial Processes

It is shown in Cuchiero et al. (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 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

$$
P_t f(x) \equiv \mathbb{E}\left[f(X_t)|X_0 = x\right] \in \mathrm{Pol}_{\leq l}(\mathcal{D}) \quad \text{for} \quad f \in \mathrm{Pol}_{\leq l}(\mathcal{D}), \tag{B.1}
$$

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

$$
\mu(x) \equiv a + b\,x, \quad V(x) \equiv G + H\,x = G + H_1 x_1 + \cdots + H_i x_i, \tag{B.2}
$$

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

$$P_t f = F\, e^{t\,A}\, E^\top, \tag{B.3}$$

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

$$\mathcal{A}e_i = \sum_{j=1}^{q} a_{ij} e_j. \tag{B.4}$$

For discounted exponential moments we have that

$$\mathbb{E}_t\left[e^{-\int_t^T \delta_0 + \delta_X\, X_s\, ds}\, e^{uX_T}\right] = e^{\phi(\tau,u)+\psi(\tau,u)X_t}, \tag{B.5}$$

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

$$\begin{aligned}\frac{d\psi(\tau,u)}{d\tau} &= -\delta_X + b\,\psi(\tau,u) + \frac{1}{2}\psi(\tau,u)^\top\, H\, \psi(\tau,u), \quad \psi(0,u) = u \\ \frac{d\phi(\tau,u)}{d\tau} &= -\delta_0 + a\,\psi(\tau,u) + \frac{1}{2}\psi(\tau,u)^\top\, G\, \psi(\tau,u), \quad \phi(0,u) = 0.\end{aligned} \tag{B.6}$$

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

B.2 Second Moment of Forecast Errors

Assuming $L \leq T$ we are interested in model-implied covariance structure of the error terms from Eq. (5)

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

$II.$ can be computed according to Eq. (21). 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.3).

C Details Related to Factor Rotations

The dynamics of the latent factors (X_{1t}, X_{2t}, X_{3t}, X_{4t}) are governed by two square root and two Gaussian processes. In all rotations, the first step is to rotate X_{3t} into

$$
\pi_t = \kappa_0 + \sum_{j=1}^{4} \kappa_j X_{jt}
$$

where π_t is the US short rate in Rotations 1 and 2, i.e. $\kappa = \delta$, and the short rate differential in Rotation 3, i.e. $\kappa = \delta - \delta^*$. The π_t dynamics are

$$
d\pi_t = \left(\omega_0 + \omega_1 X_{1t} + \omega_2 X_{2t} + \omega_3 \pi_t + \omega_4 X_{4t}\right) dt + \sum_{j=1}^{4} \kappa_j \sigma_j dW_{jt}
$$

where σ_j denotes the j-th element of $\sigma(X_t)$ and

$$
\omega_3 = \frac{\kappa_4 b_{43}}{\kappa_3} + b_{33}, \qquad \omega_0 = \sum_{i=1}^{4} \kappa_i a_i - \kappa_0 \omega_3, \qquad \omega_j = \sum_{i=j}^{4} \kappa_i b_{ij} - \kappa_j \omega_3 \text{ for } j = \{1,2,4\}.
$$

Given these π_t dynamics, we then rotate X_{4t} into process Π_t which either represents the instantaneous slope of the US term structure, the level differential, or the US level:

$$
\Pi_t = \Omega_0 + \Omega_1 X_{1t} + \Omega_2 X_{2t} + \Omega_3 \pi_t + \Omega_4 X_{4t}, \qquad \text{where in}
$$

- R1 (slope): $\Omega = \omega$, based on slope $\mu_t \equiv \omega_0 + \omega_1 X_{1t} + \omega_2 X_{2t} + \omega_3 \pi_t + \omega_4 X_{4t}$.
- R2 $(r_t - r_t^*)$: $\Omega_3 = \omega_3 - \delta_3^*/\delta_3$ and $\Omega_j = \omega_j - \delta_j^* + (\delta_3^*/\delta_3)\delta_j$ for $j = \{0, 1, 2, 4\}$.
- R3 (r_t): $\Omega_3 = \omega_3 + \delta_3^*/\delta 3$ and $\Omega_j = \omega_j + \delta_j^* - (\delta_3^*/\delta_3)\delta_j$ for $j = \{0, 1, 2, 4\}$.

The dynamics of Π_t are

$$d\Pi_t = (\lambda_0 + \lambda_1 X_{1t} + \lambda_2 X_{2t} + \lambda_3 \pi_t + \lambda_4 \Pi_t)dt + \sum_{j=1}^{4} \varrho_j \sigma_j dW_{jt}$$

where

$$\begin{aligned}
\lambda_0 &= \Omega_1 a_1 + \Omega_2 a_2 + \Omega_3 \omega_0 + \Omega_4 a_4 - \Omega_0(\Omega_3(\omega_4/\Omega_4) + b_{44}) + (b_{43}/\kappa_3)(\Omega_0\kappa_4 - \Omega_4\kappa_0), \\
\lambda_1 &= \Omega_1 b_{11} + \Omega_2 b_{21} + \Omega_3 \omega_1 + \Omega_4 b_{41} - \Omega_1(\Omega_3(\omega_4/\Omega_4) + b_{44}) + (b_{43}/\kappa_3)(\Omega_1\kappa_4 - \Omega_4\kappa_1), \\
\lambda_2 &= \Omega_2 b_{22} + \Omega_3 \omega_2 + \Omega_4 b_{42} - \Omega_2(\Omega_3(\omega_4/\Omega_4) + b_{44}) + (b_{43}/\kappa_3)(\Omega_2\kappa_4 - \Omega_4\kappa_2), \\
\lambda_3 &= \Omega_3 \omega_3 - \Omega_3(\Omega_3(\omega_4/\Omega_4) + b_{44}) + (b_{43}/\kappa_3)(\Omega_3\kappa_4 + \Omega_4), \\
\lambda_4 &= \Omega_3(\omega_4/\Omega_4) + b_{44} - (b_{43}/\kappa_3)\kappa_4, \\
\varrho_3 &= \kappa_3 \Omega_3, \\
\varrho_j &= (\Omega_3 \kappa_j + \Omega_j) \qquad \text{for } j = \{1, 2, 4\}.
\end{aligned}$$

Next, we compute the quadratic variation of π_t and Π_t and define

$$V_t \equiv c_0 + c_1 X_{1t} + c_2 X_{2t} \qquad U_t \equiv d_0 + d_1 X_{1t} + d_2 X_{2t}$$

where

$$\begin{aligned}
c_0 &= \kappa_3^2 + \kappa_4^2 & d_0 &= \varrho_3^2 + \varrho_4^2 \\
c_j &= \kappa_j^2 + \kappa_3^2 \beta_j + \kappa_4^2 \gamma_j & d_j &= \varrho_j^2 + \varrho_3^2 \beta_j + \varrho_4^2 \gamma_j \qquad \text{for } j = \{1, 2\}.
\end{aligned}$$

Note that lower bounds for the variances of π_t and Π_t are given by $\kappa_3^2 + \kappa_4^2$ and $\rho_3^2 + \rho_4^2$ respectively. Solving for X_1 and X_2 we get

$$\begin{aligned}
X_1 &= \frac{c_2 (d_0 - U) + d_2 (V - c_0)}{c_1 d_2 - c_2 d_1} \equiv f_0 + f_1 V + f_2 U, \\
X_2 &= \frac{c_1 (U - d_0) + d_1 (c_0 - V)}{c_1 d_2 - c_2 d_1} \equiv g_0 + g_1 V + g_2 U.
\end{aligned}$$

From this, we compute the joint dynamics of (V, U), rewrite π and Π dynamics in terms of V and U and finally obtain the dynamics of the observable system

$$\begin{pmatrix} dV_t \\ dU_t \\ d\pi_t \\ d\Pi_t \end{pmatrix} = \left(\begin{pmatrix} \varphi_1 \\ \varphi_2 \\ \varphi_3 \\ \varphi_4 \end{pmatrix} + \begin{pmatrix} \vartheta_{11} & \vartheta_{12} & 0 & 0 \\ \vartheta_{21} & \vartheta_{22} & 0 & 0 \\ \vartheta_{31} & \vartheta_{32} & \vartheta_{33} & \vartheta_{34} \\ \vartheta_{41} & \vartheta_{42} & \vartheta_{43} & \vartheta_{44} \end{pmatrix} \begin{pmatrix} V_t \\ U_t \\ \pi_t \\ \Pi_t \end{pmatrix} \right) dt$$

$$+ \begin{pmatrix} c_1 & c_2 & 0 & 0 \\ d_1 & d_2 & 0 & 0 \\ \kappa_1 & \kappa_2 & \kappa_3 & \kappa_4 \\ \varrho_1 & \varrho_2 & \varrho_3 & \varrho_4 \end{pmatrix} \operatorname{diag} \begin{pmatrix} f_0 + f_1 V_t + f_2 U_t \\ g_0 + g_1 V_t + g_2 U_t \\ y_0 + y_1 V_t + y_2 U_t \\ z_0 + z_1 V_t + z_2 U_t \end{pmatrix} \begin{pmatrix} dW_{1t} \\ dW_{2t} \\ dW_{3t} \\ dW_{4t} \end{pmatrix}$$

where

$$\begin{aligned}
\varphi_1 &= c_1(a_1 + b_{11} f_0) + c_2(a_2 + b_{21} f_0 + b_{22} g_0), \\
\varphi_2 &= d_1(a_1 + b_{11} f_0) + d_2(a_2 + b_{21} f_0 + b_{22} g_0), \\
\varphi_3 &= \omega_0 + \omega_1 f_0 + \omega_2 g_0 - (\omega_4/\Omega_4)(\Omega_0 + f_0 \Omega_1 + g_0 \Omega_2), \\
\varphi_4 &= \lambda_0 + \lambda_1 f_0 + \lambda_2 g_0, \\
\vartheta_{1j} &= c_1 b_{11} f_j + c_2(b_{21} f_j + b_{22} g_j) && \text{for } j = \{1, 2\}, \\
\vartheta_{2j} &= d_1 b_{11} f_j + d_2(b_{21} f_j + b_{22} g_j) && \text{for } j = \{1, 2\}, \\
\vartheta_{3j} &= \omega_1 f_j + \omega_2 g_j - (\omega_4/\Omega_4)(f_j \Omega_1 + g_j \Omega_2) && \text{for } j = \{1, 2\}, \\
\vartheta_{33} &= \omega_3 - (\omega_4/\Omega_4)\Omega_3, \\
\vartheta_{34} &= \omega_4/\Omega_4, \\
\vartheta_{4j} &= \lambda_1 f_j + \lambda_2 g_j && \text{for } j = \{1, 2\}, \\
\vartheta_{4j} &= \lambda_j && \text{for } j = \{3, 4\}, \\
y_0 &= 1 + \beta_1 f_0 + \beta_2 g_0, \\
y_j &= \beta_1 f_j + \beta_2 g_j && \text{for } j = \{1, 2\}, \\
z_0 &= 1 + \gamma_1 f_0 + \gamma_2 g_0, \\
z_j &= \gamma_1 f_j + \gamma_2 g_j && \text{for } j = \{1, 2\}.
\end{aligned}$$

D 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 Section 3.1.1; in total we have 45 parameters. The model ought to fit zero-coupon yields of the respective currencies, represent the joint evolution of the latent state variables with the foreign exchange rate, 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 $D \equiv \mathrm{diag}(12, \cdots, 1/4)$, seven foreign zero-yields $y^{\star}$ with the same maturities, and the log exchange rate s_t. We assume that the exchange rate is observed without error and 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. (19)–(20). We assume that the pricing errors enter additively into the pricing equations

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

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

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 (5) 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 the model. Appendix B.2 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

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

We sample from the posterior distribution

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

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.[29] Denote with $\phi(z; \upsilon, \Omega)$ 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.[30] With this approximation we obtain $p(Z \mid \theta)$ in density (D.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 the 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 gener-

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

[30] 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.3) in Appendix B.1. 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 (2010) 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 et al. (2010).

ate 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. We report parameter estimates of the models in the separate Internet Appendix in Section AA.

E Block Bootstrap Procedure

We use the tests proposed by Clark and West (2007) and Giacomini and White (2006) to assess the predictive ability of the 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), Hall et al. (1995), Politis and White (2004), Patton et al. (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 (1), $\Delta s_{t,T} = \alpha + \beta(y_{t,T} - y^{\star}_{t,T}) + \eta_{t,T}$, 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: Yield Pricing Errors

The table reports annualized root mean squared errors in basis points for the domestic US T-period yields (Panel A) and the respective foreign yields (Panel B). The rows indicate the model estimated, the column headers indicate the yield maturities T. 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: Interpretation of Latent State Variables: US Risk Factors and Carry Risk Factors

The table reports results related to the three factor rotations discussed in Section 4.3.2. For each rotation, we report the correlation (in percentage points) of the model-implied variables to the respective model-independent estimates in blocks of four columns each: the first two columns report results for the US short rate level (r_t), the slope (μ_t), and the level differential ($r_t - r_t^*$) implied from the model. V_t and U_t are the corresponding quadratic variations. In the rows, L_t denotes the model-free estimate of the US short rate level, Sl_t the estimate for the slope, and $L_t - L_t^*$ for the short rate differential. $QV_t[\cdot]$ denotes the respective quadratic variation. In the last two rows and columns we report correlations to $\mathbb{Q}$-expected depreciation ($\mathbb{E}_t^{\mathbb{Q}}[ds]$) and to the model-implied variance of the exchange rate ($QV_t[ds]$). Coefficients of 100% represent correlations greater than 99.5% for space reasons indicate The results are based on parameter and states variable estimates of the model using daily data from 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.

	Rotation 1				Rotation 2				Rotation 3					
	r_t	μ_t	V_t	U_t	r_t	$r_t - r_t^*$	V_t	U_t	$r_t - r_t^*$	r_t	V_t	U_t	$\mathbb{E}_t^{\mathbb{Q}}[ds]$	$QV_t[ds]$
AUD														
L_t	99.7	-5.0	49.4	58.5	99.7	-38.7	49.4	42.2	82.7	15.8	30.1	29.9	83.0	33.6
Sl_t	-11.5	55.8	32.6	38.2									-8.7	22.6
$L_t - L_t^*$					84.6	-9.2	15.8	1.6	97.0	13.0	-17.3	-17.7	97.4	-12.3
$QV_t[L]$	28.5	28.1	20.6	42.4	28.5	32.6	20.6	6.5	43.1	2.8	-14.0	-14.4	43.2	-8.5
$QV_t[Sl]$	44.5	31.3	52.0	57.4									40.5	33.3
$QV_t[L - L^*]$					45.5	-3.8	-6.8	-19.9	60.0	1.6	-36.1	-36.4	60.1	-32.1
$\mathbb{E}_t^{\mathbb{Q}}[ds]$	83.8	-11.0	13.9	42.3	83.8	-6.7	13.9	-0.1	100.0	-2.4	-18.8	-19.1	100.0	-13.9
$QV_t[ds]$	32.2	57.1	92.7	68.8	32.2	-7.3	92.7	98.2	-13.4	-2.8	99.8	99.7	-13.9	100.0
CAD														
L_t	99.7	-3.5	-66.3	-72.8	99.7	-55.0	-66.3	-73.3	54.0	11.8	-73.3	-73.3	58.7	-72.9
Sl_t	-12.7	61.3	-16.5	-31.7									1.3	-26.5
$L_t - L_t^*$					64.0	0.6	-93.3	-74.7	84.1	9.3	-75.0	-74.8	90.2	-80.7
$QV_t[L]$	31.6	19.3	-21.3	-34.2	31.6	-8.4	-21.3	-32.2	12.7	2.5	-32.0	-32.1	13.5	-30.0
$QV_t[Sl]$	50.1	27.1	-28.5	-50.8									16.0	-43.3
$QV_t[L - L^*]$					34.6	-43.7	4.7	-5.1	-4.8	2.6	-5.0	-5.1	-4.7	-2.8
$\mathbb{E}_t^{\mathbb{Q}}[ds]$	59.0	8.4	-89.2	-63.0	59.0	13.8	-89.2	-70.6	98.9	-27.5	-71.0	-70.7	100.0	-76.6
$QV_t[ds]$	-72.8	-53.5	94.5	97.4	-72.8	-12.7	94.5	99.4	-71.4	-10.4	99.5	99.4	-76.6	100.0
CHF														
L_t	99.9	-15.1	25.6	10.8	99.9	-49.6	25.6	22.8	20.9	8.7	32.8	33.1	20.6	29.1
Sl_t	-22.6	86.2	-32.7	-38.8									20.8	-31.0
$L_t - L_t^*$					22.8	64.5	-87.6	-88.9	99.0	-4.6	-83.8	-83.6	99.1	-85.9
$QV_t[L]$	27.2	-6.3	65.8	60.5	27.2	-60.2	65.8	64.9	-51.0	7.8	67.7	67.7	-51.1	66.8
$QV_t[Sl]$	26.1	-5.5	59.4	54.3									-45.6	60.3
$QV_t[L - L^*]$					41.8	-74.9	74.2	73.1	-53.0	12.6	76.6	76.7	-53.2	75.5
$\mathbb{E}_t^{\mathbb{Q}}[ds]$	20.8	17.2	-89.0	-93.8	20.8	66.7	-89.0	-90.2	100.0	-11.1	-85.2	-85.1	100.0	-87.2
$QV_t[ds]$	29.0	-23.5	99.9	97.3	29.0	-89.1	99.9	99.7	-86.9	7.3	99.9	99.9	-87.2	100.0
DEM-EUR														
L_t	99.8	-9.2	33.5	-5.4	99.8	-36.0	33.5	15.1	40.2	6.6	32.5	36.3	40.0	21.9
Sl_t	-22.3	82.3	-24.2	-36.8									19.3	-28.8
$L_t - L_t^*$					43.6	57.7	-67.9	-81.1	98.5	-8.5	-68.7	-65.4	98.6	-76.8
$QV_t[L]$	27.3	-4.5	68.7	56.7	27.3	-55.3	68.7	64.4	-44.4	11.8	68.5	69.1	-44.6	66.3
$QV_t[Sl]$	26.3	-4.4	62.7	51.3									-39.9	60.4
$QV_t[L - L^*]$					12.8	-63.4	75.1	74.4	-60.7	14.8	75.2	74.8	-60.8	75.1
$\mathbb{E}_t^{\mathbb{Q}}[ds]$	40.0	19.4	-71.2	-92.8	40.0	62.1	-71.2	-84.0	100.0	-15.3	-72.1	-68.8	100.0	-79.9
$QV_t[ds]$	22.0	-18.6	98.9	94.6	22.0	-85.2	98.9	99.7	-79.5	11.3	99.1	98.3	-79.9	100.0
GBP														
L_t	99.8	-26.5	42.2	17.2	99.8	-66.7	42.2	36.6	-9.1	1.1	37.9	38.1	-10.2	32.5
Sl_t	-21.0	82.5	-32.9	-51.2									45.8	-41.2
$L_t - L_t^*$					-11.3	73.8	-91.9	-94.8	96.3	-4.3	-94.2	-94.1	97.4	-96.3
$QV_t[L]$	26.0	2.8	67.0	52.8	26.0	-49.4	67.0	64.8	-50.9	5.1	65.4	65.4	-51.9	62.7
$QV_t[Sl]$	25.0	3.8	58.4	44.7									-43.9	54.1
$QV_t[L - L^*]$					4.5	-24.8	46.3	45.3	-38.1	1.7	45.6	45.6	-38.9	44.2
$\mathbb{E}_t^{\mathbb{Q}}[ds]$	-8.5	38.2	-90.5	-98.3	-8.5	75.1	-90.5	-93.7	99.8	-19.2	-93.1	-93.0	100.0	-95.5
$QV_t[ds]$	31.0	-32.7	98.7	97.3	31.0	-82.0	98.7	99.8	-94.1	3.2	99.6	99.6	-95.5	100.0
JPY														
L_t	98.5	-65.1	72.8	63.5	98.5	-60.8	72.8	81.8	30.2	17.5	93.6	88.6	30.3	86.1
Sl_t	-10.7	26.0	13.4	-10.2									27.8	9.9
$L_t - L_t^*$					31.6	56.1	84.3	-15.2	98.6	-64.7	26.2	1.9	98.6	67.9
$QV_t[L]$	28.5	-6.8	-19.4	56.0	28.5	-80.7	-19.4	45.7	-50.8	64.6	23.5	37.3	-50.7	-5.1
$QV_t[Sl]$	27.6	-9.9	-16.6	50.3									-45.6	-3.7
$QV_t[L - L^*]$					31.5	-83.4	-18.5	53.1	-52.1	70.1	29.3	44.3	-51.9	-2.3
$\mathbb{E}_t^{\mathbb{Q}}[ds]$	34.4	-55.4	86.4	-41.2	34.4	55.3	86.4	-11.7	100.0	-66.2	29.8	5.6	100.0	70.8
$QV_t[ds]$	89.2	-64.8	96.3	32.4	89.2	-16.0	96.3	60.0	70.8	-13.2	87.8	73.2	70.8	100.0

Table 6: Fama Conditions

The table shows the relevant covariances ($\mathbb{C}ov^{\mathbb{P}}$) and variances ($\mathbb{V}^{\mathbb{P}}$) for the Fama-conditions, Eq. (3). 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 7: Regressions of Excess Returns on Expected Excess Returns

The table shows the results from estimating, by ordinary least squares, the regression (23), $ER_{t,T} = \alpha' + \beta' \widehat{ER}_{t,T} + \eta'_{t,T}$, 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 8: Ability to Predict Excess Returns

The table reports results related to the predictive ability of the 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 E 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 9: Decomposing Foreign Exchange Risk Premiums

This table reports means and standard deviations (in parentheses) of annualized foreign exchange risk premiums 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 2.2, in particular Eq. (8). The descriptives are calculated from daily model estimates of the risk premiums. 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 10: 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***

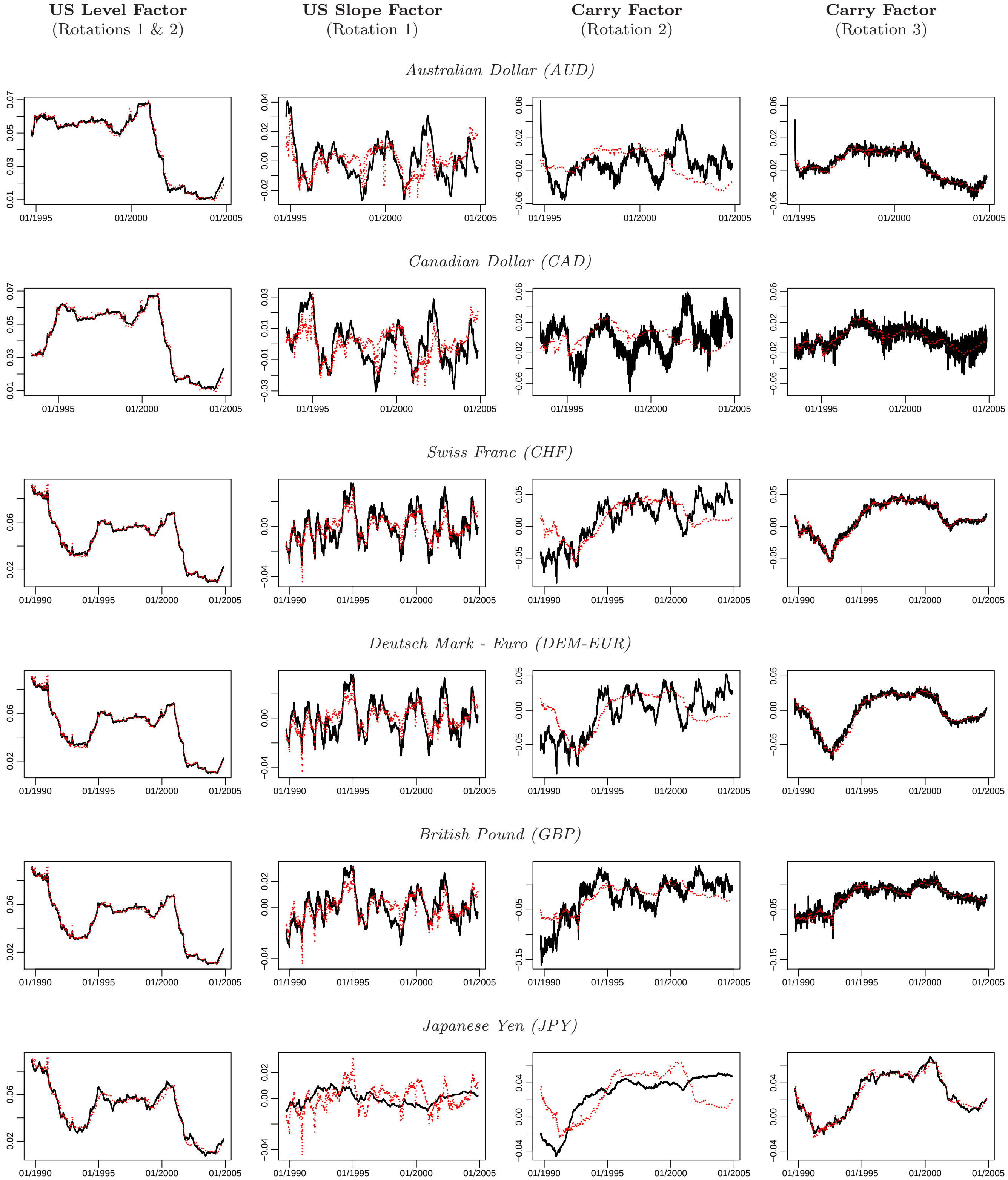
Table 11: Regressions of Expected Excess Returns on Financial and Fundamental Variables

The table presents results of regressing expected excess returns on the 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

Figure 1: Interpretation of Latent State Variables: US Risk Factors and Carry Risk Factors

The figure plots the US risk factors and Carry risk factors as described in Section 4.3.2. The solid (black) lines represent model-implied estimates obtained through factor rotations. The dashed lines (red) are the corresponding model-independent estimates. The first column plots the US short rate level from Rotations 1 and 2, the second the US slope from Rotation 1, the third the carry factor from Rotation 2, and the fourth the carry factor from Rotation 3. Estimations are based on daily data from 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.



Internet Appendix for

"Properties of Foreign Exchange Risk Premiums" (not for publication)

This separate Internet Appendix first reports and discusses detailed empirical results related to parameter estimations. We then present a number of Tables which are discussed and referenced in the main text but are not included in the paper.

AA Details Related to Model Estimation Results

We present the parameter estimates for the two-country models of the US and the six foreign countries estimated using the zero yields of the two countries and the respective spot exchange rate applying the procedure described in Section 3.2. Tables A.1 to A.6 report 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. All confidence intervals are fairly tight, only for 9 of the 264 parameters we report the confidence interval includes zero and most of these are significant at the 10 percent level.

We conduct further checks to validate the accuracy of the estimation results. At first sight, some of the stochastic volatility parameters (β and γ) appear large and as an additional plausibility check we compare model-implied quadratic variations of the short rate and the instantaneous exchange rate return to their empirical counterparts. The results in Panel A of Table A.7 reveal that the levels of implied and observed variations are very similar.

We also show that the properties of model-implied US bond risk premiums are consistent with those reported in other studies. Duffee (2002) demonstrates that affine term structure models can replicate observed term structure characteristics only if the specification of the market price of risk is flexible enough. A first check reveals that the risk premiums implied by the model change signs and are highly variable, a necessary condition to match the observed data. Following Duffee (2002), we assess the specification of the market price of risk by analyzing whether the model is capable to replicate the empirical relation between expected returns and the slope of the yield curve. We generate yield predictions for maturities of 6 months, 2 years, and 4 years (the longest maturity in our data set) at prediction horizons of 3 months, 6 months, and 1 year, and regress the prediction errors on the slope defined as the 4-year minus the 3-month yield. Panel B of Table A.7 shows that the t-statistics are small with only a few exceptions which implies that the model captures the information contained in the slope. Overall, the results suggest that the market price of risk specification is indeed consistent with the prevailing literature on US term structure risk premiums.

Table A.1: AUD Model Parameters

The table shows parameter estimates for the AUD 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	2.5% quantile	97.5% quantile
ζ	0.0009	0.0008	0.0009
$\zeta^\star$	0.0015	0.0015	0.0015
β_1	3945.8700	1064.9200	5327.0700
β_2	7218.8000	6402.5400	8345.5400
γ_1	124.9850	16.1548	1976.2000
γ_2	94173.0000	91804.1000	95765.3000
Σ_1	-0.0054	-0.0061	-0.0050
Σ_2	-0.0000	-0.0007	0.0008
Σ_3	0.0000	0.0000	0.0000
Σ_4	-0.0000	-0.0000	-0.0000
$a_1^{\mathbb{P}}$	0.6289	0.5062	1.1658
$a_2^{\mathbb{P}}$	0.6526	0.5073	1.5410
$a_3^{\mathbb{P}}$	-193.6010	-215.1030	-174.1270
$a_4^{\mathbb{P}}$	-101.8440	-131.1690	-67.1335
$b_{11}^{\mathbb{P}}$	-0.0049	-0.0581	-0.0005
$b_{21}^{\mathbb{P}}$	0.7347	0.5389	1.0467
$b_{22}^{\mathbb{P}}$	-0.2674	-0.3727	-0.2017
$b_{31}^{\mathbb{P}}$	-257.3190	-288.8670	-245.2310
$b_{32}^{\mathbb{P}}$	93.0770	88.7177	105.7550
$b_{33}^{\mathbb{P}}$	-0.3096	-0.3570	-0.2367
$b_{41}^{\mathbb{P}}$	-168.6240	-212.0130	-148.0130
$b_{42}^{\mathbb{P}}$	22.5160	14.2609	34.0064
$b_{43}^{\mathbb{P}}$	12.7164	12.2719	13.4160
$b_{44}^{\mathbb{P}}$	-76.8625	-79.7563	-74.9951
$a_1^{\mathbb{Q}}$	17.8304	17.4561	18.8995
$a_2^{\mathbb{Q}}$	0.5334	0.5008	0.5918
$b_{11}^{\mathbb{Q}}$	-0.7539	-0.7913	-0.7338
$b_{21}^{\mathbb{Q}}$	0.1902	0.1718	0.2058
$b_{22}^{\mathbb{Q}}$	-0.0138	-0.0188	-0.0094
$b_{31}^{\mathbb{Q}}$	219.0930	214.0410	220.3490
$b_{32}^{\mathbb{Q}}$	-43.1930	-43.7912	-42.1947
$b_{33}^{\mathbb{Q}}$	-0.5065	-0.5188	-0.4870
$b_{41}^{\mathbb{Q}}$	14.2100	9.2637	14.1644
$b_{42}^{\mathbb{Q}}$	5.1591	2.9440	6.3413
$b_{43}^{\mathbb{Q}}$	10.0722	9.7209	10.5316
$b_{44}^{\mathbb{Q}}$	-66.5951	-66.8218	-66.5717
δ_0	$1.07E{-04}$	$6.74E{-06}$	$6.28E{-04}$
δ_1	$-1.33E{-03}$	$-1.40E{-03}$	$-1.28E{-03}$
δ_2	$8.15E{-04}$	$7.96E{-04}$	$8.55E{-04}$
δ_3	$8.67E{-06}$	$8.55E{-06}$	$8.71E{-06}$
$\delta_0^\star$	$2.21E{-04}$	$5.82E{-06}$	$7.77E{-04}$
$\delta_1^\star$	$-1.27E{-03}$	$-1.30E{-03}$	$-1.23E{-03}$
$\delta_2^\star$	$1.18E{-03}$	$1.17E{-03}$	$1.22E{-03}$
$\delta_4^\star$	$1.55E{-05}$	$1.52E{-05}$	$1.56E{-05}$

Table A.2: CAD Model Parameters

The table shows parameter estimates for the CAD 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	2.5% quantile	97.5% quantile
ζ	0.0008	0.0008	0.0008
$\zeta^\star$	0.0019	0.0019	0.0020
β_1	8.8785	7.3745	10.9926
β_2	9.1753	6.6249	11.3678
γ_1	56999.7000	56461.6000	57182.7000
γ_2	15979.8000	15840.0000	17406.1000
Σ_1	0.0008	-0.0008	0.0012
Σ_2	-0.0066	-0.0071	-0.0052
Σ_3	0.0017	0.0016	0.0020
Σ_4	-0.0001	-0.0001	-0.0001
$a_1^{\mathbb{P}}$	4.1751	3.6036	5.2474
$a_2^{\mathbb{P}}$	5.8534	4.4874	6.9351
$a_3^{\mathbb{P}}$	-108.8080	-114.4710	-99.2117
$a_4^{\mathbb{P}}$	65.7446	54.3099	72.2453
$b_{11}^{\mathbb{P}}$	-0.5467	-0.7063	-0.4303
$b_{21}^{\mathbb{P}}$	0.0213	0.0008	0.0767
$b_{22}^{\mathbb{P}}$	-1.3170	-1.4719	-1.0494
$b_{31}^{\mathbb{P}}$	-1.0493	-1.9556	0.1114
$b_{32}^{\mathbb{P}}$	16.1108	14.5503	17.5721
$b_{33}^{\mathbb{P}}$	-0.1661	-0.2292	-0.1007
$b_{41}^{\mathbb{P}}$	147.9280	146.2700	200.2120
$b_{42}^{\mathbb{P}}$	163.0730	143.6460	171.8130
$b_{43}^{\mathbb{P}}$	447.1020	435.8450	464.9060
$b_{44}^{\mathbb{P}}$	-220.3550	-227.3720	-212.8180
$a_1^{\mathbb{Q}}$	3.1116	3.0477	3.2396
$a_2^{\mathbb{Q}}$	1.3518	1.2708	1.3920
$b_{11}^{\mathbb{Q}}$	-0.5245	-0.5438	-0.5161
$b_{21}^{\mathbb{Q}}$	0.0014	0.0000	0.0035
$b_{22}^{\mathbb{Q}}$	-0.1073	-0.1142	-0.1016
$b_{31}^{\mathbb{Q}}$	-12.1488	-12.1899	-11.7413
$b_{32}^{\mathbb{Q}}$	2.8612	2.6061	2.9571
$b_{33}^{\mathbb{Q}}$	-0.7100	-0.7265	-0.6883
$b_{41}^{\mathbb{Q}}$	-29.4461	-30.3831	-17.0176
$b_{42}^{\mathbb{Q}}$	-45.9201	-49.5772	-44.1277
$b_{43}^{\mathbb{Q}}$	367.3170	363.1390	370.3030
$b_{44}^{\mathbb{Q}}$	-191.1730	-191.7990	-187.4800
δ_0	$9.76E{-}02$	$9.71E{-}02$	$9.92E{-}02$
δ_1	$-3.03E{-}05$	$-1.26E{-}04$	$6.62E{-}05$
δ_2	$9.45E{-}05$	$-2.13E{-}05$	$1.76E{-}04$
δ_3	$4.07E{-}04$	$4.04E{-}04$	$4.22E{-}04$
$\delta_0^\star$	$5.39E{-}02$	$5.33E{-}02$	$5.55E{-}02$
$\delta_1^\star$	$2.16E{-}03$	$2.05E{-}03$	$2.24E{-}03$
$\delta_2^\star$	$4.15E{-}03$	$4.05E{-}03$	$4.45E{-}03$
$\delta_4^\star$	$1.86E{-}04$	$1.84E{-}04$	$1.93E{-}04$

Table A.3: CHF Model Parameters

The table shows parameter estimates for the CHF 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	2.5% quantile	97.5% quantile
ζ	0.0008	0.0008	0.0008
$\zeta^\star$	0.0019	0.0019	0.0019
β_1	1089.9000	1019.6000	1595.0700
β_2	774.9630	492.9500	880.3900
γ_1	84422.8000	83847.2000	85020.9000
γ_2	113087.0000	112107.0000	113172.0000
Σ_1	0.0104	0.0087	0.0115
Σ_2	0.0010	-0.0008	0.0021
Σ_3	0.0003	0.0003	0.0003
Σ_4	-0.0001	-0.0001	-0.0001
$a_1^{\mathbb{P}}$	9.4311	7.8729	11.1957
$a_2^{\mathbb{P}}$	2.7025	1.5711	4.3386
$a_3^{\mathbb{P}}$	-881.0210	-917.5050	-856.4220
$a_4^{\mathbb{P}}$	402.4230	392.9370	419.1870
$b_{11}^{\mathbb{P}}$	-0.9812	-1.1133	-0.8642
$b_{21}^{\mathbb{P}}$	0.0068	0.0002	0.0259
$b_{22}^{\mathbb{P}}$	-0.4022	-0.5408	-0.2992
$b_{31}^{\mathbb{P}}$	-140.0270	-148.4500	-130.9710
$b_{32}^{\mathbb{P}}$	84.9015	77.2622	88.2037
$b_{33}^{\mathbb{P}}$	-1.4743	-1.5142	-1.3515
$b_{41}^{\mathbb{P}}$	188.7960	181.1880	198.8390
$b_{42}^{\mathbb{P}}$	-222.9010	-237.4920	-214.3050
$b_{43}^{\mathbb{P}}$	66.8164	65.5973	69.7816
$b_{44}^{\mathbb{P}}$	-106.5780	-111.6860	-105.0680
$a_1^{\mathbb{Q}}$	5.8042	5.7219	5.9142
$a_2^{\mathbb{Q}}$	4.8991	4.6643	5.0358
$b_{11}^{\mathbb{Q}}$	-0.6038	-0.6107	-0.5987
$b_{21}^{\mathbb{Q}}$	0.0010	0.0000	0.0016
$b_{22}^{\mathbb{Q}}$	-0.2122	-0.2154	-0.2103
$b_{31}^{\mathbb{Q}}$	-140.5590	-142.6140	-139.4780
$b_{32}^{\mathbb{Q}}$	75.0871	74.5818	79.1018
$b_{33}^{\mathbb{Q}}$	-0.9140	-0.9249	-0.9064
$b_{41}^{\mathbb{Q}}$	17.7682	16.4450	19.0031
$b_{42}^{\mathbb{Q}}$	-12.6246	-14.5794	-10.5460
$b_{43}^{\mathbb{Q}}$	60.2569	59.2545	60.8944
$b_{44}^{\mathbb{Q}}$	-97.8089	-98.6446	-97.5368
δ_0	$7.54E-02$	$7.48E-02$	$7.64E-02$
δ_1	$7.76E-06$	$-4.38E-05$	$3.97E-05$
δ_2	$-2.83E-04$	$-3.02E-04$	$-2.68E-04$
δ_3	$3.35E-05$	$3.34E-05$	$3.36E-05$
$\delta_0^\star$	$2.37E-05$	$5.55E-07$	$8.38E-05$
$\delta_1^\star$	$1.63E-03$	$1.60E-03$	$1.65E-03$
$\delta_2^\star$	$1.45E-03$	$1.42E-03$	$1.55E-03$
$\delta_4^\star$	$2.07E-05$	$2.04E-05$	$2.09E-05$

Table A.4: DEM-EUR Model Parameters

The table shows parameter estimates for the DEM-EUR 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	2.5% quantile	97.5% quantile
ζ	0.0009	0.0008	0.0009
$\zeta^{\star}$	0.0024	0.0023	0.0024
β_1	882.5680	459.1370	934.3260
β_2	1074.9200	991.9040	1337.9900
γ_1	84726.4000	83971.7000	84942.0000
γ_2	114691.0000	114280.0000	115001.0000
Σ_1	0.0154	0.0145	0.0169
Σ_2	0.0050	0.0043	0.0070
Σ_3	0.0002	0.0002	0.0002
Σ_4	-0.0001	-0.0001	-0.0000
$a_1^{\mathbb{P}}$	3.5550	2.3926	3.9369
$a_2^{\mathbb{P}}$	0.9073	0.5428	1.8332
$a_3^{\mathbb{P}}$	-670.3850	-679.2770	-644.4510
$a_4^{\mathbb{P}}$	318.5340	311.0800	324.4080
$b_{11}^{\mathbb{P}}$	-1.0044	-1.0937	-0.8944
$b_{21}^{\mathbb{P}}$	0.0021	0.0001	0.0156
$b_{22}^{\mathbb{P}}$	-0.0018	-0.0164	-0.0001
$b_{31}^{\mathbb{P}}$	-56.5448	-62.3560	-44.9247
$b_{32}^{\mathbb{P}}$	44.1639	41.0341	48.0625
$b_{33}^{\mathbb{P}}$	-0.9645	-1.0206	-0.8544
$b_{41}^{\mathbb{P}}$	179.8400	162.2240	200.5020
$b_{42}^{\mathbb{P}}$	-193.4750	-201.6950	-180.5390
$b_{43}^{\mathbb{P}}$	59.7758	55.1904	60.9749
$b_{44}^{\mathbb{P}}$	-88.6530	-91.1689	-83.5696
$a_1^{\mathbb{Q}}$	4.4062	4.3022	4.4802
$a_2^{\mathbb{Q}}$	4.1133	3.9891	4.2511
$b_{11}^{\mathbb{Q}}$	-0.6916	-0.6966	-0.6831
$b_{21}^{\mathbb{Q}}$	0.0003	0.0000	0.0011
$b_{22}^{\mathbb{Q}}$	-0.0976	-0.0996	-0.0940
$b_{31}^{\mathbb{Q}}$	-120.9930	-121.8150	-120.0630
$b_{32}^{\mathbb{Q}}$	57.8475	56.1198	58.7773
$b_{33}^{\mathbb{Q}}$	-1.0489	-1.0551	-1.0290
$b_{41}^{\mathbb{Q}}$	17.3027	16.4805	19.3163
$b_{42}^{\mathbb{Q}}$	0.2710	-2.1334	1.6969
$b_{43}^{\mathbb{Q}}$	56.5762	54.4855	57.1519
$b_{44}^{\mathbb{Q}}$	-89.6759	-89.8414	-89.2115
δ_0	$5.25E{-}02$	$5.25E{-}02$	$5.33E{-}02$
δ_1	$-1.60E{-}04$	$-2.18E{-}04$	$-1.36E{-}04$
δ_2	$1.25E{-}05$	$-1.75E{-}05$	$1.79E{-}05$
δ_3	$3.27E{-}05$	$3.24E{-}05$	$3.30E{-}05$
$\delta_0^{\star}$	$1.49E{-}05$	$2.70E{-}07$	$7.78E{-}05$
$\delta_1^{\star}$	$2.69E{-}03$	$2.64E{-}03$	$2.75E{-}03$
$\delta_2^{\star}$	$1.41E{-}03$	$1.38E{-}03$	$1.44E{-}03$
$\delta_4^{\star}$	$1.97E{-}05$	$1.95E{-}05$	$2.03E{-}05$

Table A.5: GBP Model Parameters

The table shows parameter estimates for the GBP 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	2.5% quantile	97.5% quantile
ζ	0.0009	0.0008	0.0009
$\zeta^{\star}$	0.0026	0.0026	0.0026
β_1	183.4170	20.9248	650.3200
β_2	633.9690	515.4800	751.1730
γ_1	57912.7000	57391.9000	58294.4000
γ_2	74981.2000	74400.8000	75048.4000
Σ_1	0.0120	0.0106	0.0129
Σ_2	0.0044	0.0040	0.0051
Σ_3	0.0001	0.0001	0.0001
Σ_4	-0.0000	-0.0000	-0.0000
$a_1^{\mathbb{P}}$	5.4736	4.8151	5.9215
$a_2^{\mathbb{P}}$	0.5409	0.5068	1.3256
$a_3^{\mathbb{P}}$	-683.4750	-710.9810	-672.9920
$a_4^{\mathbb{P}}$	-291.8470	-315.4370	-271.6110
$b_{11}^{\mathbb{P}}$	-1.4459	-1.5097	-1.3147
$b_{21}^{\mathbb{P}}$	3.8638	3.5038	4.4227
$b_{22}^{\mathbb{P}}$	-1.3794	-1.5159	-1.2520
$b_{31}^{\mathbb{P}}$	-161.5000	-169.3630	-148.2510
$b_{32}^{\mathbb{P}}$	77.2220	70.1285	82.8690
$b_{33}^{\mathbb{P}}$	-0.7799	-0.8441	-0.6852
$b_{41}^{\mathbb{P}}$	291.5330	275.6010	301.0580
$b_{42}^{\mathbb{P}}$	-185.7750	-191.5280	-174.5420
$b_{43}^{\mathbb{P}}$	94.4833	89.8578	97.1711
$b_{44}^{\mathbb{P}}$	-183.5080	-190.4420	-178.7140
$a_1^{\mathbb{Q}}$	3.3095	3.1992	3.4611
$a_2^{\mathbb{Q}}$	6.0921	5.7417	6.3318
$b_{11}^{\mathbb{Q}}$	-0.6272	-0.6433	-0.6156
$b_{21}^{\mathbb{Q}}$	0.0096	0.0006	0.0412
$b_{22}^{\mathbb{Q}}$	-0.2516	-0.2572	-0.2441
$b_{31}^{\mathbb{Q}}$	-130.3660	-131.7770	-128.7370
$b_{32}^{\mathbb{Q}}$	47.8954	46.0986	48.9368
$b_{33}^{\mathbb{Q}}$	-0.6362	-0.6462	-0.6163
$b_{41}^{\mathbb{Q}}$	22.3227	17.2650	24.0451
$b_{42}^{\mathbb{Q}}$	-23.7638	-28.5961	-20.5912
$b_{43}^{\mathbb{Q}}$	91.5406	86.4298	92.3040
$b_{44}^{\mathbb{Q}}$	-181.2760	-181.4790	-180.9580
δ_0	$5.10E{-02}$	$4.94E{-02}$	$5.20E{-02}$
δ_1	$-2.62E{-04}$	$-2.80E{-04}$	$-1.59E{-04}$
δ_2	$-2.12E{-04}$	$-2.44E{-04}$	$-1.89E{-04}$
δ_3	$3.44E{-05}$	$3.42E{-05}$	$3.45E{-05}$
$\delta_0^{\star}$	$2.69E{-02}$	$2.58E{-02}$	$2.74E{-02}$
$\delta_1^{\star}$	$2.30E{-03}$	$2.22E{-03}$	$2.41E{-03}$
$\delta_2^{\star}$	$5.47E{-04}$	$4.96E{-04}$	$5.86E{-04}$
$\delta_4^{\star}$	$5.80E{-05}$	$5.72E{-05}$	$6.10E{-05}$

Table A.6: 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	2.5% quantile	97.5% quantile
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	$1.57E{-}04$	$1.55E{-}04$	$1.58E{-}04$
$\delta_0^\star$	$1.00E{-}03$	$9.83E{-}04$	$1.03E{-}03$
$\delta_1^\star$	$1.80E{-}03$	$1.79E{-}03$	$1.80E{-}03$
$\delta_2^\star$	$-7.58E{-}05$	$-7.61E{-}05$	$-7.41E{-}05$
$\delta_4^\star$	$1.10E{-}04$	$1.10E{-}04$	$1.11E{-}04$

Table A.7: Summary of Additional Checks Related to Estimation Results

Panel A compares the means of the model-implied quadratic variation of the short rate and the instantaneous exchange rate return to their empirical counterparts.
Panel B reports the results of regressing errors of model-generated yield predictions on the slope of the term structure. The yield maturities are given in the first column, the prediction horizon in the second. The slope is defined as the 4-year yield (the longest maturity in our data set) minus the 3-month yield. The t-statistics in brackets are calculated using Newey and West (1987) standard errors with the optimal truncation lag chosen as suggested by Andrews (1991).
For both Panels, 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.

Panel A: Model Implied vs. Observed Quadratic Variations

	QV of Short Rate		QV of FX Returns	
	Model Implied	Observed	Model Implied	Observed
AUD	0.0110	0.0048	0.1102	0.1048
CAD	0.0044	0.0050	0.0607	0.0592
CHF	0.0056	0.0062	0.1268	0.1134
DEM-EUR	0.0055	0.0062	0.1087	0.1043
GBP	0.0053	0.0062	0.0914	0.0897
JPY	0.0058	0.0062	0.1265	0.1103

Panel B: Relation Between Yield Prediction Errors and Slope

Maturity	Horizon	AUD	CAD	CHF	DEM-EUR	GBP	JPY
6 months	3	-0.0459	-0.0256	-0.2670	-0.1943	-0.2298	-0.1153
		[-0.23]	[-0.15]	[-2.24]	[-1.54]	[-1.65]	[-0.79]
2 years	3	-0.1336	-0.0378	-0.4823	-0.3722	-0.4279	-0.1618
		[-0.32]	[-0.08]	[-1.61]	[-1.13]	[-1.16]	[-0.44]
4 years	3	-0.2138	-0.0148	-0.6695	-0.5488	-0.6032	-0.1189
		[-0.17]	[-0.01]	[-0.97]	[-0.78]	[-0.98]	[-0.18]
6 months	6	-0.1417	-0.0426	-0.2921	-0.2540	-0.2852	-0.0268
		[-0.64]	[-0.21]	[-2.02]	[-1.67]	[-1.34]	[-0.16]
2 years	6	-0.3553	-0.1246	-0.5450	-0.4914	-0.5187	-0.1785
		[-0.95]	[-0.21]	[-1.77]	[-1.35]	[-1.24]	[-0.51]
4 years	6	-0.5811	-0.2652	-0.7799	-0.7363	-0.7176	-0.2890
		[-0.70]	[-0.28]	[-1.28]	[-1.15]	[-1.27]	[-0.52]
6 months	12	-0.1418	-0.0299	-0.2638	-0.2442	-0.2416	0.0434
		[-0.66]	[-0.14]	[-1.87]	[-1.63]	[-1.00]	[0.31]
2 years	12	-0.3771	-0.1341	-0.4989	-0.4774	-0.4637	-0.1452
		[-1.19]	[-0.24]	[-1.55]	[-1.41]	[-1.25]	[-0.57]
4 years	12	-0.6450	-0.3229	-0.7408	-0.7419	-0.6698	-0.3417
		[-1.50]	[-0.49]	[-1.77]	[-1.60]	[-1.52]	[-1.04]

Table A.8: Yield Pricing Errors: Model with Three Factors

The table reports annualized root mean squared errors in basis points for the domestic US T-period yields (Panel A) and the respective foreign yields (Panel B). The rows indicate the model estimated, the column headers indicate the yield maturities T. 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 monthy	6 months	1 year	2 years	3 years	4 years
AUD	10	13	13	17	37	57	75
CAD	10	13	12	18	45	67	87
CHF	12	15	15	20	40	64	88
DEM-EUR	11	15	18	28	54	83	111
GBP	13	16	15	23	49	77	102
JPY	14	20	24	28	50	88	128

Panel B: Foreign Yields

	1 month	3 monthy	6 months	1 year	2 years	3 years	4 years
AUD	13	16	15	17	38	58	83
CAD	16	21	22	21	37	61	89
CHF	8	11	12	15	31	50	66
DEM-EUR	11	14	13	19	41	59	77
GBP	17	22	22	28	47	72	104
JPY	8	10	12	19	32	56	88

Table A.9: Regressions of Excess Returns on Expected Excess Returns: Model with Three Factors

The table shows the results from estimating, by ordinary least squares, the regression (23), $ER_{t,T} = \alpha' + \beta' \widehat{ER}_{t,T} + \eta'_{t,T}$, 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	0.0007	0.0003	−0.0021	−0.0055
se(α')	(0.0001)	(0.0006)	(0.0023)	(0.0074)	(0.0233)	(0.0821)
β'	0.3371***	0.4496***	0.6816***	0.9922***	1.1123***	0.9888**
se(β')	(0.0787)	(0.1222)	(0.2349)	(0.3124)	(0.2991)	(0.3904)
$t[\beta' = 1]$	[-8.42]	[-4.51]	[-1.36]	[-0.02]	[0.38]	[-0.03]
R^2	0.0056	0.0211	0.0804	0.2909	0.5409	0.4621
CAD						
α'	0.0000	−0.0000	−0.0003	−0.0007	0.0010	0.0008
se(α')	(0.0001)	(0.0003)	(0.0012)	(0.0033)	(0.0080)	(0.0329)
β'	0.3067	0.7001***	0.8483***	0.8837***	1.0394***	0.9991***
se(β')	(0.1893)	(0.2193)	(0.2565)	(0.2560)	(0.1806)	(0.2184)
$t[\beta' = 1]$	[-3.66]	[-1.37]	[-0.59]	[-0.45]	[0.22]	[-0.00]
R^2	0.0010	0.0168	0.0745	0.1895	0.6030	0.6388
CHF						
α'	0.0002**	0.0008	0.0030	0.0075	0.0085	−0.0026
se(α')	(0.0001)	(0.0006)	(0.0025)	(0.0071)	(0.0235)	(0.0465)
β'	1.2685***	1.0701***	0.9294**	1.0224**	1.4033**	1.0769***
se(β')	(0.3299)	(0.3521)	(0.3807)	(0.4121)	(0.6147)	(0.4045)
$t[\beta' = 1]$	[0.81]	[0.20]	[-0.19]	[0.05]	[0.66]	[0.19]
R^2	0.0023	0.0078	0.0225	0.0570	0.2141	0.2508
DEM-EUR						
α'	0.0000	0.0001	0.0008	0.0023	0.0006	0.0025
se(α')	(0.0001)	(0.0005)	(0.0023)	(0.0064)	(0.0201)	(0.0520)
β'	0.8955***	0.7215**	0.5804*	0.5556*	0.7896**	0.7821***
se(β')	(0.3181)	(0.2893)	(0.2978)	(0.3108)	(0.3585)	(0.2939)
$t[\beta' = 1]$	[-0.33]	[-0.96]	[-1.41]	[-1.43]	[-0.59]	[-0.74]
R^2	0.0036	0.0111	0.0284	0.0576	0.2490	0.3003
GBP						
α'	0.0001	0.0005	0.0023	0.0081	0.0133	0.0013
se(α')	(0.0001)	(0.0004)	(0.0019)	(0.0072)	(0.0206)	(0.0392)
β'	0.1106	0.1760	0.3369	0.1876	0.7897	0.9221***
se(β')	(0.1327)	(0.2344)	(0.5209)	(1.1003)	(0.7584)	(0.2728)
$t[\beta' = 1]$	[-6.70]	[-3.52]	[-1.27]	[-0.74]	[-0.28]	[-0.29]
R^2	0.0008	0.0015	0.0019	0.0007	0.0499	0.3342
JPY						
α'	−0.0000	−0.0001	−0.0002	0.0017	0.0087	−0.0271
se(α')	(0.0001)	(0.0007)	(0.0026)	(0.0054)	(0.0169)	(0.0476)
β'	0.1143**	−0.0104	0.1630	0.8783***	1.0931***	0.7604***
se(β')	(0.0500)	(0.1690)	(0.3405)	(0.2791)	(0.2377)	(0.1983)
$t[\beta' = 1]$	[-17.73]	[-5.98]	[-2.46]	[-0.44]	[0.39]	[-1.21]
R^2	0.0012	0.0000	0.0015	0.0819	0.4373	0.4723

Table A.10: Ability to Predict Excess Returns: Model with Three Factors

The table reports results related to the predictive ability of the 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 E 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	0.5547	0.5882	0.6583	0.6917	0.7750	0.7750	0.5547	0.5882	0.6583	0.6917	0.7750	0.7750
$R2$	0.0057	0.0218	0.0836	0.2947	0.5461	0.5014	0.0044	0.0150	0.0526	0.2321	0.4467	0.4366
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]	[0.119]	[0.044]	[0.031]	[<0.01]	[<0.01]	[<0.01]	[0.180]	[0.084]	[0.068]	[<0.01]	[0.016]	[<0.01]
CAD												
HR	0.5279	0.5669	0.6029	0.5515	0.7353	0.6985	0.5279	0.5669	0.6029	0.5515	0.7353	0.6985
$R2$	0.0011	0.0170	0.0752	0.1909	0.6104	0.6997	-0.0002	0.0110	0.0529	0.1350	0.5171	0.6689
p-value[CW]	[0.078]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.188]	[0.017]	[0.012]	[0.011]	[<0.01]	[<0.01]
p-value[GW]	[0.321]	[0.234]	[0.199]	[0.053]	[0.014]	[<0.01]	[0.304]	[0.323]	[0.265]	[0.096]	[0.039]	[<0.01]
CHF												
HR	0.5306	0.5272	0.5667	0.5889	0.7778	0.4944	0.5306	0.5272	0.5667	0.5889	0.7778	0.4944
$R2$	0.0024	0.0080	0.0235	0.0590	0.2141	0.2615	0.0017	0.0049	0.0101	0.0262	0.1152	0.1417
p-value[CW]	[<0.01]	[<0.01]	[0.030]	[0.014]	[<0.01]	[<0.01]	[<0.01]	[0.040]	[0.166]	[0.095]	[<0.01]	[<0.01]
p-value[GW]	[0.192]	[0.075]	[0.153]	[0.044]	[0.033]	[<0.01]	[0.197]	[0.099]	[0.198]	[0.070]	[0.078]	[<0.01]
DEM-EUR												
HR	0.5425	0.5740	0.5556	0.6111	0.7556	0.7778	0.5425	0.5740	0.5556	0.6111	0.7556	0.7778
$R2$	0.0038	0.0118	0.0315	0.0638	0.2530	0.3034	0.0032	0.0089	0.0179	0.0254	0.1293	0.2110
p-value[CW]	[<0.01]	[<0.01]	[0.011]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.058]	[0.064]	[<0.01]	[<0.01]
p-value[GW]	[0.107]	[0.100]	[0.079]	[0.024]	[0.010]	[<0.01]	[0.066]	[0.056]	[0.113]	[0.053]	[0.038]	[<0.01]
GBP												
HR	0.5230	0.5474	0.5222	0.5333	0.6167	0.6944	0.5230	0.5474	0.5222	0.5333	0.6167	0.6944
$R2$	0.0014	0.0040	0.0131	0.0322	0.1372	0.5384	0.0008	0.0012	0.0002	-0.0035	0.0390	0.5087
p-value[CW]	[0.046]	[0.121]	[0.125]	[0.131]	[0.023]	[<0.01]	[0.071]	[0.178]	[0.241]	[0.193]	[<0.01]	[<0.01]
p-value[GW]	[<0.01]	[0.256]	[0.395]	[0.242]	[0.099]	[<0.01]	[0.011]	[0.381]	[0.469]	[0.254]	[0.045]	[<0.01]
JPY												
HR	0.5144	0.5032	0.5556	0.6722	0.7944	0.7889	0.5144	0.5032	0.5556	0.6722	0.7944	0.7889
$R2$	0.0012	0.0000	0.0016	0.0824	0.4402	0.5574	0.0004	-0.0035	-0.0157	0.0406	0.3157	0.3042
p-value[CW]	[0.022]	[0.464]	[0.245]	[<0.01]	[<0.01]	[<0.01]	[0.025]	[0.629]	[0.702]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.130]	[0.163]	[0.334]	[0.063]	[<0.01]	[<0.01]	[0.170]	[0.190]	[0.215]	[0.090]	[<0.01]	[0.016]

Table A.11: Yield Pricing Errors: Model with Two Factors

The table reports annualized root mean squared errors in basis points for the domestic US T-period yields (Panel A) and the respective foreign yields (Panel B). The rows indicate the model estimated, the column headers indicate the yield maturities T. 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 monthy	6 months	1 year	2 years	3 years	4 years
AUD	11	14	15	17	35	56	73
CAD	12	16	16	19	39	60	80
CHF	51	88	122	163	204	227	250
DEM-EUR	12	16	17	22	35	60	87
GBP	41	71	99	133	166	192	223
JPY	53	91	127	170	215	242	268

Panel B: Foreign Yields

	1 month	3 monthy	6 months	1 year	2 years	3 years	4 years
AUD	25	45	66	99	142	172	206
CAD	27	44	61	80	119	149	175
CHF	7	11	12	14	27	41	54
DEM-EUR	64	109	148	186	242	266	280
GBP	15	19	18	25	48	76	108
JPY	8	12	14	20	28	53	89

Table A.12: Regressions of Excess Returns on Expected Excess Returns: Model with Two Factors

The table shows the results from estimating, by ordinary least squares, the regression (23), $ER_{t,T} = \alpha' + \beta' \widehat{ER}_{t,T} + \eta'_{t,T}$, 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	0.0002	0.0006	0.0002	−0.0018	0.0096
se(α')	(0.0001)	(0.0006)	(0.0025)	(0.0075)	(0.0230)	(0.0800)
β'	0.3742*	0.6022***	0.7650***	0.9456***	0.9972***	0.9994**
se(β')	(0.2118)	(0.1776)	(0.2256)	(0.2552)	(0.2414)	(0.3902)
$t[\beta' = 1]$	[-2.95]	[-2.24]	[-1.04]	[-0.21]	[-0.01]	[-0.00]
R^2	0.0013	0.0157	0.0907	0.2922	0.5287	0.4564
CAD						
α'	0.0000	0.0002	0.0006	0.0018	0.0062	−0.0067
se(α')	(0.0001)	(0.0003)	(0.0015)	(0.0043)	(0.0114)	(0.0346)
β'	0.1413	0.4709**	0.5806**	0.6425**	0.8001***	1.1422***
se(β')	(0.1983)	(0.1993)	(0.2559)	(0.2835)	(0.2209)	(0.2628)
$t[\beta' = 1]$	[-4.33]	[-2.66]	[-1.64]	[-1.26]	[-0.90]	[0.54]
R^2	0.0002	0.0091	0.0489	0.1438	0.4941	0.6238
CHF						
α'	0.0002	0.0007	0.0013	0.0040	0.0018	−0.0245
se(α')	(0.0002)	(0.0008)	(0.0028)	(0.0079)	(0.0251)	(0.0548)
β'	0.0933***	0.2136**	0.1061	0.6615	0.6311	0.7357
se(β')	(0.0360)	(0.0951)	(0.2591)	(0.5645)	(0.9380)	(0.8102)
$t[\beta' = 1]$	[-25.18]	[-8.27]	[-3.45]	[-0.60]	[-0.39]	[-0.33]
R^2	0.0105	0.0142	0.0008	0.0106	0.0132	0.0383
DEM-EUR						
α'	0.0001	0.0004	0.0016	0.0038	0.0005	−0.0172
se(α')	(0.0001)	(0.0006)	(0.0026)	(0.0075)	(0.0248)	(0.0683)
β'	0.0145	0.0417	0.3150	0.7166	1.4695*	1.5042*
se(β')	(0.0317)	(0.0718)	(0.3176)	(0.7526)	(0.8492)	(0.8787)
$t[\beta' = 1]$	[-31.05]	[-13.34]	[-2.16]	[-0.38]	[0.55]	[0.57]
R^2	0.0000	0.0003	0.0048	0.0161	0.1113	0.1459
GBP						
α'	0.0002**	0.0009**	0.0036	0.0111	0.0153	−1.0082
se(α')	(0.0001)	(0.0004)	(0.0036)	(0.0202)	(0.0923)	(0.7068)
β'	−2.3075***	−1.6457	−0.7420	−0.6750	0.7247	8.6550
se(β')	(0.8252)	(1.0596)	(2.5434)	(4.9694)	(4.5353)	(5.5053)
$t[\beta' = 1]$	[-4.01]	[-2.50]	[-0.68]	[-0.34]	[-0.06]	[1.39]
R^2	0.0046	0.0047	0.0008	0.0012	0.0037	0.2887
JPY						
α'	−0.0000	−0.0002	−0.0011	−0.0031	−0.0004	0.0424
se(α')	(0.0001)	(0.0006)	(0.0027)	(0.0079)	(0.0246)	(0.0877)
β'	−0.0852	0.7226**	0.5436	0.6216	1.1548**	1.9730**
se(β')	(0.1576)	(0.3058)	(0.3631)	(0.4858)	(0.5606)	(0.7784)
$t[\beta' = 1]$	[-6.89]	[-0.91]	[-1.26]	[-0.78]	[0.28]	[1.25]
R^2	0.0001	0.0084	0.0131	0.0314	0.1490	0.5432

Table A.13: Ability to Predict Excess Returns: Model with Two Factors

The table reports results related to the predictive ability of the 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 E 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	0.5148	0.5598	0.6333	0.7000	0.7833	0.7833	0.5148	0.5598	0.6333	0.7000	0.7833	0.7833
$R2$	0.0014	0.0164	0.0939	0.2961	0.5340	0.4960	0.0002	0.0096	0.0633	0.2336	0.4320	0.4306
p-value[CW]	[0.044]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.108]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.377]	[0.177]	[0.034]	[<0.01]	[<0.01]	[<0.01]	[0.549]	[0.281]	[0.073]	[<0.01]	[0.017]	[<0.01]
CAD												
HR	0.5243	0.5619	0.5441	0.5221	0.6838	0.6176	0.5243	0.5619	0.5441	0.5221	0.6838	0.6176
$R2$	0.0002	0.0094	0.0495	0.1452	0.5036	0.6872	-0.0010	0.0033	0.0266	0.0861	0.3847	0.6551
p-value[CW]	[0.313]	[0.025]	[0.014]	[<0.01]	[<0.01]	[<0.01]	[0.522]	[0.073]	[0.043]	[0.030]	[<0.01]	[<0.01]
p-value[GW]	[0.346]	[0.266]	[0.274]	[0.140]	[0.050]	[<0.01]	[0.154]	[0.206]	[0.338]	[0.231]	[0.119]	[<0.01]
CHF												
HR	0.5453	0.5411	0.4944	0.5222	0.5056	0.5056	0.5453	0.5411	0.4944	0.5222	0.5056	0.5056
$R2$	0.0105	0.0144	0.0018	0.0127	0.0132	0.0521	0.0099	0.0114	-0.0118	-0.0216	-0.1109	-0.1017
p-value[CW]	[<0.01]	[<0.01]	[0.353]	[0.170]	[0.397]	[0.035]	[<0.01]	[<0.01]	[0.355]	[0.510]	[0.975]	[0.351]
p-value[GW]	[<0.01]	[0.060]	[0.196]	[0.258]	[0.272]	[0.029]	[<0.01]	[0.121]	[0.195]	[0.168]	[0.096]	[0.012]
DEM-EUR												
HR	0.5025	0.5006	0.4889	0.5667	0.6167	0.5722	0.5025	0.5006	0.4889	0.5667	0.6167	0.5722
$R2$	0.0002	0.0009	0.0080	0.0225	0.1161	0.1496	-0.0004	-0.0019	-0.0060	-0.0176	-0.0303	0.0368
p-value[CW]	[0.438]	[0.381]	[0.232]	[0.058]	[<0.01]	[0.012]	[0.474]	[0.454]	[0.433]	[0.426]	[0.266]	[0.220]
p-value[GW]	[0.362]	[0.269]	[0.257]	[0.176]	[0.057]	[<0.01]	[0.437]	[0.322]	[0.176]	[0.241]	[0.168]	[0.013]
GBP												
HR	0.4954	0.4905	0.5111	0.5167	0.5667	0.6833	0.4954	0.4905	0.5111	0.5167	0.5667	0.6833
$R2$	0.0051	0.0072	0.0120	0.0326	0.0953	0.5068	0.0045	0.0044	-0.0009	-0.0030	-0.0077	0.4751
p-value[CW]	[0.999]	[0.862]	[0.221]	[0.062]	[0.025]	[<0.01]	[0.997]	[0.802]	[0.431]	[0.379]	[0.130]	[<0.01]
p-value[GW]	[0.069]	[0.200]	[0.460]	[0.273]	[0.118]	[<0.01]	[0.090]	[0.351]	[0.537]	[0.358]	[0.186]	[<0.01]
JPY												
HR	0.5192	0.5259	0.5944	0.6278	0.6833	0.9556	0.5192	0.5259	0.5944	0.6278	0.6833	0.9556
$R2$	0.0001	0.0084	0.0132	0.0319	0.1533	0.6169	-0.0007	0.0049	-0.0039	-0.0122	-0.0349	0.3976
p-value[CW]	[0.622]	[<0.01]	[0.028]	[0.010]	[<0.01]	[<0.01]	[0.700]	[0.023]	[0.142]	[0.150]	[0.343]	[0.086]
p-value[GW]	[0.315]	[0.139]	[0.209]	[0.184]	[0.078]	[<0.01]	[0.147]	[0.328]	[0.439]	[0.397]	[0.189]	[0.021]

Table A.14: Results of Three-Country Estimation: USD, CHF, and DEM-EUR

The sample period is September 18, 1989 to October 10, 2008. The results in Panel A are based on daily data, those in Panels B and C on non-overlapping observations for horizons up to 1 month and on monthly frequency for horizons of 3 months and beyond.
Panel A reports annualized root mean squared errors in basis points for the domestic US T-period yields and the respective foreign yields of the CHF and DEM-EUR. The results are based on daily observations.
Panel B shows the results from estimating, by ordinary least squares, the regression (23), $ER_{t,T} = \alpha' + \beta'\widehat{ER}_{t,T} + \eta'_{t,T}$, 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.
Panel C reports results related to the predictive ability of the 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 E which accounts for autocorrelation and heteroscedasticity.

Panel A: Yield Pricing Errors

	1 month	3 monthy	6 months	1 year	2 years	3 years	4 years
USD	7	7	11	31	71	94	113
CHF	16	24	34	51	82	100	115
DEM-EUR	20	29	37	49	81	106	131

Panel B: Regressions of Excess Returns on Expected Excess Returns

	1 day	1 week	1 month	3 months	1 year	4 years
CHF						
α'	0.0000	0.0002	0.0009	0.0019	0.0045	0.0112
se(α')	(0.0001)	(0.0007)	(0.0026)	(0.0068)	(0.0204)	(0.0394)
β'	0.1020	−0.0380	0.3269	0.7236**	0.8974***	1.0545***
se(β')	(0.2873)	(0.3406)	(0.3328)	(0.2933)	(0.3342)	(0.3824)
$t[\beta' = 1]$	[-3.13]	[-3.05]	[-2.02]	[-0.94]	[-0.31]	[0.14]
R^2	0.0000	0.0000	0.0058	0.0656	0.2155	0.2613
DEM-EUR						
α'	0.0001	0.0003	0.0010	0.0018	0.0025	0.0123
se(α')	(0.0001)	(0.0006)	(0.0024)	(0.0063)	(0.0195)	(0.0495)
β'	0.2147	0.1266	0.4632	0.7469**	1.0139***	1.2367***
se(β')	(0.2586)	(0.2987)	(0.3235)	(0.3131)	(0.3735)	(0.4215)
$t[\beta' = 1]$	[-3.04]	[-2.92]	[-1.66]	[-0.81]	[0.04]	[0.56]
R^2	0.0002	0.0003	0.0136	0.0767	0.2694	0.3018

Panel C: Ability to Predict Excess Returns

	Model vs. UIP						Model vs. RW					
	1d	1w	1m	3m	1y	4y	1d	1w	1m	3m	1y	4y
CHF												
HR	0.5172	0.5424	0.6056	0.6667	0.7611	0.7778	0.5172	0.5424	0.6056	0.6667	0.7611	0.7778
R2	0.0001	0.0002	0.0069	0.0676	0.2155	0.2720	-0.0006	-0.0028	-0.0067	0.0352	0.1168	0.1538
p-value[CW]	[0.429]	[0.600]	[0.198]	[<0.01]	[<0.01]	[<0.01]	[0.796]	[0.920]	[0.536]	[0.049]	[<0.01]	[<0.01]
p-value[GW]	[0.535]	[0.443]	[0.163]	[0.053]	[0.011]	[<0.01]	[0.275]	[0.221]	[0.228]	[0.105]	[0.032]	[<0.01]
DEM-EUR												
HR	0.5346	0.5638	0.5833	0.6722	0.7667	0.7722	0.5346	0.5638	0.5833	0.6722	0.7667	0.7722
R2	0.0003	0.0009	0.0167	0.0827	0.2734	0.3049	-0.0003	-0.0019	0.0029	0.0451	0.1530	0.2127
p-value[CW]	[0.197]	[0.307]	[0.073]	[<0.01]	[<0.01]	[<0.01]	[0.573]	[0.749]	[0.292]	[0.029]	[<0.01]	[<0.01]
p-value[GW]	[0.472]	[0.434]	[0.075]	[0.017]	[<0.01]	[<0.01]	[0.345]	[0.308]	[0.087]	[0.034]	[0.034]	[<0.01]

Table A.15: Results of Three-Country Estimation: USD, CHF, and JPY

The sample period is September 18, 1989 to October 10, 2008. The results in Panel A are based on daily data, those in Panels B and C on non-overlapping observations for horizons up to 1 month and on monthly frequency for horizons of 3 months and beyond.
Panel A reports annualized root mean squared errors in basis points for the domestic US T-period yields and the respective foreign yields of the CHF and JPY. The results are based on daily observations.
Panel B shows the results from estimating, by ordinary least squares, the regression (23), $ER_{t,T} = \alpha' + \beta'\widehat{ER}_{t,T} + \eta'_{t,T}$, 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.
Panel C reports results related to the predictive ability of the 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 E which accounts for autocorrelation and heteroscedasticity.

Panel A: Yield Pricing Errors

	1 month	3 monthy	6 months	1 year	2 years	3 years	4 years
USD	9	13	16	27	45	62	77
CHF	30	50	66	86	114	139	166
JPY	11	16	21	26	29	47	80

Panel B: Regressions of Excess Returns on Expected Excess Returns

	1 day	1 week	1 month	3 months	1 year	4 years
CHF						
α'	0.0000	0.0003	0.0015	0.0046	0.0094	0.0009
se(α')	(0.0001)	(0.0006)	(0.0025)	(0.0066)	(0.0203)	(0.0411)
β'	−0.0907	0.0484	0.3397	0.6150**	1.0054***	0.9061*
se(β')	(0.1365)	(0.1850)	(0.2864)	(0.3113)	(0.3025)	(0.5072)
$t[\beta' = 1]$	[-7.99]	[-5.14]	[-2.31]	[-1.24]	[0.02]	[-0.19]
R^2	0.0002	0.0001	0.0050	0.0311	0.1884	0.1834
JPY						
α'	−0.0001	−0.0002	−0.0004	−0.0011	0.0044	−0.0125
se(α')	(0.0001)	(0.0007)	(0.0028)	(0.0069)	(0.0174)	(0.0482)
β'	0.0578**	0.0648	0.0177	0.3868**	1.0183***	0.8909***
se(β')	(0.0279)	(0.0594)	(0.1169)	(0.1856)	(0.2252)	(0.2137)
$t[\beta' = 1]$	[-33.71]	[-15.75]	[-8.40]	[-3.30]	[0.08]	[-0.51]
R^2	0.0025	0.0038	0.0001	0.0316	0.4014	0.5641

Panel C: Ability to Predict Excess Returns

	Model vs. UIP						Model vs. RW					
	1d	1w	1m	3m	1y	4y	1d	1w	1m	3m	1y	4y
CHF												
HR	0.5149	0.5272	0.5833	0.6000	0.7444	0.7722	0.5149	0.5272	0.5833	0.6000	0.7444	0.7722
$R2$	0.0002	0.0003	0.0061	0.0331	0.1884	0.1951	-0.0004	-0.0028	-0.0075	-0.0005	0.0863	0.0645
p-value[CW]	[0.666]	[0.218]	[0.108]	[0.016]	[<0.01]	[<0.01]	[0.784]	[0.402]	[0.409]	[0.153]	[<0.01]	[0.028]
p-value[GW]	[0.293]	[0.481]	[0.254]	[0.025]	[<0.01]	[<0.01]	[0.297]	[0.376]	[0.319]	[0.102]	[0.045]	[<0.01]
JPY												
HR	0.5086	0.5411	0.5111	0.6167	0.7889	0.9389	0.5086	0.5411	0.5111	0.6167	0.7889	0.9389
$R2$	0.0025	0.0038	0.0002	0.0321	0.4045	0.6344	0.0017	0.0003	-0.0171	-0.0120	0.2721	0.4251
p-value[CW]	[<0.01]	[0.125]	[0.573]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.136]	[0.751]	[0.150]	[<0.01]	[0.024]
p-value[GW]	[0.221]	[0.186]	[0.330]	[0.154]	[<0.01]	[<0.01]	[0.276]	[0.163]	[0.202]	[0.302]	[<0.01]	[0.018]

Table A.16: Regressions of Excess Returns on Expected Excess Returns: Sample 01/1996 to 10/2008

The table shows the results from estimating, by ordinary least squares, the regression (23), $ER_{t,T} = \alpha' + \beta'\widehat{ER}_{t,T} + \eta'_{t,T}$, 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 January 24, 1996 to October 10, 2008 for AUD, CAD, CHF, GBP, and JPY. For DEM-EUR the sample period is January 1, 1998 to October 10, 2008.

	1 day	1 week	1 month	3 months	1 year	4 years
AUD						
α'	0.0001	0.0003	0.0012	0.0026	0.0011	0.0064
se(α')	(0.0001)	(0.0005)	(0.0029)	(0.0077)	(0.0233)	(0.0963)
β'	0.7261***	0.8357***	1.0245***	1.0313***	1.0901***	0.9526***
se(β')	(0.2349)	(0.2428)	(0.3043)	(0.3118)	(0.2652)	(0.3439)
$t[\beta' = 1]$	[-1.17]	[-0.68]	[0.08]	[0.10]	[0.34]	[-0.14]
R^2	0.0042	0.0240	0.1301	0.3152	0.6360	0.4240
CAD						
α'	0.0001	0.0002	0.0008	0.0015	0.0007	0.0003
se(α')	(0.0001)	(0.0003)	(0.0013)	(0.0035)	(0.0079)	(0.0334)
β'	0.1503	0.8019***	1.0462***	1.0350***	1.0555***	1.0222***
se(β')	(0.2145)	(0.3022)	(0.2827)	(0.2516)	(0.1503)	(0.1313)
$t[\beta' = 1]$	[-3.96]	[-0.66]	[0.16]	[0.14]	[0.37]	[0.17]
R^2	0.0002	0.0162	0.1101	0.2750	0.7415	0.7596
CHF						
α'	−0.0000	−0.0001	0.0002	0.0016	0.0007	−0.0030
se(α')	(0.0002)	(0.0007)	(0.0027)	(0.0079)	(0.0174)	(0.0423)
β'	0.3789*	0.4834***	0.6959**	0.9280***	1.0777***	0.9830***
se(β')	(0.1947)	(0.1652)	(0.2815)	(0.2868)	(0.2201)	(0.1940)
$t[\beta' = 1]$	[-3.19]	[-3.13]	[-1.08]	[-0.25]	[0.35]	[-0.09]
R^2	0.0029	0.0106	0.0516	0.1910	0.5782	0.5321
DEM-EUR						
α'	0.0001	0.0004	0.0022	0.0058	−0.0065	0.0073
se(α')	(0.0002)	(0.0007)	(0.0027)	(0.0074)	(0.0086)	(0.0648)
β'	0.2149	0.5708***	0.6283***	0.7359***	1.0743***	0.9155***
se(β')	(0.2185)	(0.2212)	(0.2188)	(0.1886)	(0.0742)	(0.2359)
$t[\beta' = 1]$	[-3.59]	[-1.94]	[-1.70]	[-1.40]	[1.00]	[-0.36]
R^2	0.0008	0.0198	0.0811	0.2421	0.8373	0.5997
GBP						
α'	0.0001	0.0005	0.0011	0.0023	0.0044	0.0056
se(α')	(0.0001)	(0.0004)	(0.0014)	(0.0045)	(0.0163)	(0.0402)
β'	0.1824*	0.3815**	0.8514***	0.9699***	0.9266***	0.9392***
se(β')	(0.1098)	(0.1715)	(0.2178)	(0.1989)	(0.2302)	(0.1534)
$t[\beta' = 1]$	[-7.45]	[-3.61]	[-0.68]	[-0.15]	[-0.32]	[-0.40]
R^2	0.0007	0.0043	0.0612	0.2285	0.4901	0.5703
JPY						
α'	0.0002	0.0008	0.0014	0.0004	−0.0153	−0.0711*
se(α')	(0.0002)	(0.0010)	(0.0032)	(0.0065)	(0.0255)	(0.0376)
β'	1.4600***	1.5081***	1.1789**	1.1704**	1.2898**	0.5786**
se(β')	(0.5018)	(0.5177)	(0.4824)	(0.5731)	(0.6333)	(0.2443)
$t[\beta' = 1]$	[0.92]	[0.98]	[0.37]	[0.30]	[0.46]	[-1.72]
R^2	0.0030	0.0168	0.0382	0.0754	0.1846	0.2989

Table A.17: Ability to Predict Excess Returns: Sample 01/1996 to 10/2008

The table reports results related to the predictive ability of the 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 E 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 January 24, 1996 to October 10, 2008 for AUD, CAD, CHF, GBP, and JPY. For DEM-EUR the sample period is January 1, 1998 to October 10, 2008.

	Model vs. UIP						Model vs. RW					
	1d	1w	1m	3m	1y	4y	1d	1w	1m	3m	1y	4y
AUD												
HR	0.5525	0.6022	0.6857	0.7238	0.8762	0.7905	0.5525	0.6022	0.6857	0.7238	0.8762	0.7905
$R2$	0.0043	0.0247	0.1326	0.3187	0.6370	0.4835	0.0029	0.0173	0.1043	0.2574	0.5602	0.3965
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]	[0.169]	[0.063]	[0.039]	[<0.01]	[<0.01]	[<0.01]	[0.255]	[0.089]	[0.054]	[<0.01]	[<0.01]	[<0.01]
CAD												
HR	0.5381	0.5739	0.6476	0.5905	0.8857	0.7333	0.5381	0.5739	0.6476	0.5905	0.8857	0.7333
$R2$	0.0004	0.0172	0.1148	0.2823	0.7475	0.8228	-0.0009	0.0102	0.0873	0.2211	0.6794	0.7999
p-value[CW]	[0.406]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.645]	[0.040]	[0.012]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.454]	[0.273]	[0.172]	[0.042]	[<0.01]	[<0.01]	[0.206]	[0.407]	[0.235]	[0.073]	[<0.01]	[<0.01]
CHF												
HR	0.5403	0.5630	0.6476	0.7524	0.8667	0.9143	0.5403	0.5630	0.6476	0.7524	0.8667	0.9143
$R2$	0.0031	0.0116	0.0550	0.1996	0.5980	0.5398	0.0019	0.0054	0.0298	0.1407	0.4876	0.5037
p-value[CW]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.018]	[0.035]	[0.016]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.212]	[0.284]	[0.228]	[0.016]	[<0.01]	[<0.01]	[0.326]	[0.429]	[0.379]	[0.031]	[<0.01]	[<0.01]
DEM-EUR												
HR	0.5355	0.5766	0.6463	0.6463	0.8780	0.8902	0.5355	0.5766	0.6463	0.6463	0.8780	0.8902
$R2$	0.0009	0.0206	0.0854	0.2522	0.8390	0.7251	-0.0001	0.0157	0.0642	0.2039	0.8025	0.7374
p-value[CW]	[0.157]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.288]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.297]	[0.074]	[0.130]	[<0.01]	[<0.01]	[<0.01]	[0.306]	[0.105]	[0.175]	[0.012]	[<0.01]	[<0.01]
GBP												
HR	0.5350	0.5522	0.6000	0.7048	0.7238	0.8190	0.5350	0.5522	0.6000	0.7048	0.7238	0.8190
$R2$	0.0015	0.0092	0.0867	0.2878	0.5742	0.7129	0.0004	0.0031	0.0567	0.2088	0.4335	0.6554
p-value[CW]	[0.145]	[0.028]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.214]	[0.079]	[<0.01]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.262]	[0.022]	[0.173]	[0.022]	[0.016]	[<0.01]	[0.341]	[0.097]	[0.218]	[0.034]	[0.026]	[<0.01]
JPY												
HR	0.5407	0.6000	0.5905	0.6952	0.6952	1.0000	0.5407	0.6000	0.5905	0.6952	0.6952	1.0000
$R2$	0.0034	0.0188	0.0465	0.0979	0.2751	0.8639	0.0025	0.0143	0.0256	0.0395	0.0357	0.1591
p-value[CW]	[<0.01]	[<0.01]	[0.025]	[0.018]	[0.015]	[<0.01]	[<0.01]	[<0.01]	[0.073]	[0.067]	[0.051]	[<0.01]
p-value[GW]	[0.153]	[0.080]	[0.293]	[0.145]	[0.071]	[<0.01]	[0.150]	[0.078]	[0.243]	[0.146]	[0.068]	[<0.01]

Table A.18: Regressions of Excess Returns on Expected Excess Returns: Sample 01/1996 to 10/2008 including Currency Options

The table shows the results from estimating, by ordinary least squares, the regression (23), $ER_{t,T} = \alpha' + \beta'\widehat{ER}_{t,T} + \eta'_{t,T}$, 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 January 24, 1996 to October 10, 2008 for AUD, CAD, CHF, GBP, and JPY. For DEM-EUR the sample period isJanuary 1, 1998 to October 10, 2008.

	1 day	1 week	1 month	3 months	1 year	4 years
AUD						
α'	0.0000	0.0001	0.0008	0.0016	−0.0012	0.0213
se(α')	(0.0001)	(0.0005)	(0.0025)	(0.0062)	(0.0196)	(0.0907)
β'	0.8295***	0.8741***	0.9148***	0.9433***	1.0499***	0.9226***
se(β')	(0.1979)	(0.2266)	(0.2468)	(0.2277)	(0.2275)	(0.3378)
$t[\beta' = 1]$	[-0.86]	[-0.56]	[-0.35]	[-0.25]	[0.22]	[-0.23]
R^2	0.0053	0.0292	0.1267	0.3194	0.6843	0.4217
CAD						
α'	0.0001	0.0003	0.0013	0.0028	0.0017	−0.0061
se(α')	(0.0001)	(0.0004)	(0.0014)	(0.0034)	(0.0078)	(0.0349)
β'	0.2050**	0.4329*	0.8983***	0.9465***	1.0325***	1.0217***
se(β')	(0.1020)	(0.2545)	(0.2489)	(0.2213)	(0.1471)	(0.1395)
$t[\beta' = 1]$	[-7.79]	[-2.23]	[-0.41]	[-0.24]	[0.22]	[0.16]
R^2	0.0007	0.0071	0.1013	0.2716	0.7402	0.7454
CHF						
α'	−0.0000	0.0000	0.0004	0.0006	−0.0062	−0.0044
se(α')	(0.0001)	(0.0006)	(0.0026)	(0.0079)	(0.0179)	(0.0423)
β'	0.5639***	0.7340***	0.7205***	0.8088***	0.9364***	1.0130***
se(β')	(0.1391)	(0.1687)	(0.2331)	(0.2357)	(0.1962)	(0.1995)
$t[\beta' = 1]$	[-3.13]	[-1.58]	[-1.20]	[-0.81]	[-0.32]	[0.06]
R^2	0.0054	0.0239	0.0574	0.1638	0.5470	0.5294
DEM-EUR						
α'	0.0001	0.0004	0.0022	0.0059	−0.0061	−0.0002
se(α')	(0.0002)	(0.0007)	(0.0027)	(0.0074)	(0.0088)	(0.0687)
β'	0.1598	0.4812***	0.6352***	0.7666***	1.1301***	0.9498***
se(β')	(0.1688)	(0.1722)	(0.2219)	(0.1930)	(0.0791)	(0.2484)
$t[\beta' = 1]$	[-4.98]	[-3.01]	[-1.64]	[-1.21]	[1.64]	[-0.20]
R^2	0.0005	0.0136	0.0740	0.2344	0.8314	0.5927
GBP						
α'	0.0001	0.0004	−0.0109**	−0.0297**	−0.0433*	−0.0777
se(α')	(0.0001)	(0.0005)	(0.0045)	(0.0123)	(0.0243)	(0.0527)
β'	−0.0026	0.2746	2.4790***	2.5660***	2.2722***	1.7400***
se(β')	(0.1448)	(0.3704)	(0.7249)	(0.7243)	(0.4975)	(0.3669)
$t[\beta' = 1]$	[-6.92]	[-1.96]	[2.04]	[2.16]	[2.56]	[2.02]
R^2	0.0000	0.0006	0.0643	0.2027	0.5105	0.5737
JPY						
α'	−0.0003	−0.0012	−0.0030	−0.0055	−0.0145	−0.0513**
se(α')	(0.0002)	(0.0008)	(0.0034)	(0.0093)	(0.0248)	(0.0216)
β'	1.4620***	1.6955***	0.9862***	0.6151**	0.7675***	0.6841***
se(β')	(0.3360)	(0.3787)	(0.3556)	(0.2820)	(0.2921)	(0.1309)
$t[\beta' = 1]$	[1.38]	[1.84]	[-0.04]	[-1.37]	[-0.80]	[-2.41]
R^2	0.0058	0.0419	0.0574	0.0486	0.1947	0.4152

Table A.19: Ability to Predict Excess Returns: Sample 01/1996 to 10/2008 including Currency Options

The table reports results related to the predictive ability of the 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 E 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 January 24, 1996 to October 10, 2008 for AUD, CAD, CHF, GBP, and JPY. For DEM-EUR the sample period isJanuary 1, 1998 to October 10, 2008.

	Model vs. UIP						Model vs. RW					
	1d	1w	1m	3m	1y	4y	1d	1w	1m	3m	1y	4y
AUD												
HR	0.5385	0.5891	0.6857	0.6857	0.8762	0.7905	0.5385	0.5891	0.6857	0.6857	0.8762	0.7905
$R2$	0.0054	0.0299	0.1291	0.3229	0.6852	0.4815	0.0041	0.0225	0.1007	0.2619	0.6186	0.3941
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]	[0.147]	[0.084]	[0.044]	[<0.01]	[<0.01]	[<0.01]	[0.211]	[0.120]	[0.064]	[<0.01]	[<0.01]	[<0.01]
CAD												
HR	0.5355	0.5522	0.6476	0.6095	0.8857	0.7048	0.5355	0.5522	0.6476	0.6095	0.8857	0.7048
$R2$	0.0009	0.0082	0.1061	0.2790	0.7462	0.8123	-0.0004	0.0011	0.0783	0.2175	0.6778	0.7880
p-value[CW]	[0.173]	[0.056]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.311]	[0.190]	[0.012]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.423]	[0.384]	[0.185]	[0.040]	[<0.01]	[<0.01]	[0.397]	[0.472]	[0.252]	[0.071]	[<0.01]	[<0.01]
CHF												
HR	0.5416	0.5630	0.5905	0.7238	0.8571	0.9143	0.5416	0.5630	0.5905	0.7238	0.8571	0.9143
$R2$	0.0056	0.0249	0.0608	0.1728	0.5683	0.5371	0.0045	0.0188	0.0357	0.1119	0.4497	0.5007
p-value[CW]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.021]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.115]	[0.055]	[0.222]	[0.011]	[<0.01]	[<0.01]	[0.149]	[0.097]	[0.342]	[0.014]	[<0.01]	[<0.01]
DEM-EUR												
HR	0.5355	0.5905	0.6341	0.6463	0.8902	0.8780	0.5355	0.5905	0.6341	0.6463	0.8902	0.8780
$R2$	0.0006	0.0143	0.0783	0.2446	0.8331	0.7203	-0.0004	0.0094	0.0570	0.1959	0.7953	0.7329
p-value[CW]	[0.264]	[0.014]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.422]	[0.040]	[<0.01]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.428]	[0.176]	[0.130]	[<0.01]	[<0.01]	[<0.01]	[0.388]	[0.296]	[0.176]	[0.011]	[<0.01]	[<0.01]
GBP												
HR	0.5255	0.5391	0.5714	0.5619	0.6667	0.6476	0.5255	0.5391	0.5714	0.5619	0.6667	0.6476
$R2$	0.0009	0.0055	0.0897	0.2639	0.5912	0.7152	-0.0002	-0.0006	0.0598	0.1823	0.4562	0.6581
p-value[CW]	[0.511]	[0.073]	[0.020]	[<0.01]	[<0.01]	[<0.01]	[0.658]	[0.227]	[0.085]	[0.040]	[0.019]	[<0.01]
p-value[GW]	[0.318]	[0.043]	[0.149]	[0.034]	[<0.01]	[<0.01]	[0.497]	[0.171]	[0.176]	[0.051]	[0.016]	[<0.01]
JPY												
HR	0.5146	0.5587	0.5714	0.6095	0.6952	1.0000	0.5146	0.5587	0.5714	0.6095	0.6952	1.0000
$R2$	0.0062	0.0439	0.0655	0.0718	0.2840	0.8865	0.0054	0.0395	0.0450	0.0116	0.0475	0.2986
p-value[CW]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.010]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.030]	[0.041]	[<0.01]
p-value[GW]	[0.059]	[<0.01]	[0.064]	[0.074]	[0.082]	[<0.01]	[0.088]	[0.016]	[0.114]	[0.120]	[0.085]	[<0.01]

Table A.20: Yield Pricing Errors: Sample until December 2006

The table reports annualized root mean squared errors in basis points for the domestic US T-period yields (Panel A) and the respective foreign yields (Panel B). The rows indicate the model estimated, the column headers indicate the yield maturities T. The results are based on daily observations for the sample periods are October 12, 1994 to December 29, 2006 for AUD; June 1, 1993 to December 29, 2006 for CAD; and September 18, 1989 to December 29, 2006 for CHF, DEM-EUR, GBP, and JPY.

Panel A: US Yields

	1 month	3 monthy	6 months	1 year	2 years	3 years	4 years
AUD	3	4	6	10	9	13	20
CAD	3	3	6	9	9	12	16
CHF	3	3	5	10	9	10	17
DEM-EUR	3	3	6	12	10	11	18
GBP	3	3	5	10	10	11	17
JPY	9	11	11	16	35	52	68

Panel B: Foreign Yields

	1 month	3 monthy	6 months	1 year	2 years	3 years	4 years
AUD	6	7	8	15	17	24	37
CAD	7	9	10	16	23	37	57
CHF	7	8	8	13	27	39	51
DEM-EUR	8	10	10	16	33	46	65
GBP	9	10	11	24	35	52	77
JPY	3	3	5	9	11	10	17

Table A.21: Regressions of Excess Returns on Expected Excess Returns: Sample until December 2006

The table shows the results from estimating, by ordinary least squares, the regression (23), $ER_{t,T} = \alpha' + \beta'\widehat{ER}_{t,T} + \eta'_{t,T}$, 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 December 29, 2006 for AUD; June 1, 1993 to December 29, 2006 for CAD; and September 18, 1989 to December 29, 2006 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	−0.0003	−0.0011	0.0005	0.0072
se(α')	(0.0001)	(0.0006)	(0.0023)	(0.0060)	(0.0166)	(0.0927)
β'	0.4105**	0.5598**	0.5944**	0.8118***	1.0773***	0.9135
se(β')	(0.1805)	(0.2612)	(0.2480)	(0.1944)	(0.1117)	(0.8124)
$t[\beta' = 1]$	[-3.27]	[-1.69]	[-1.64]	[-0.97]	[0.69]	[-0.11]
R^2	0.0026	0.0181	0.0702	0.3328	0.7487	0.3540
CAD						
α'	−0.0001***	−0.0002	−0.0013	−0.0028	−0.0001	−0.0062
se(α')	(0.0000)	(0.0003)	(0.0012)	(0.0032)	(0.0096)	(0.0330)
β'	0.0551	0.8163***	0.6201**	0.6877***	1.1416***	1.1312**
se(β')	(0.1839)	(0.3114)	(0.2599)	(0.2391)	(0.3003)	(0.4722)
$t[\beta' = 1]$	[-5.14]	[-0.59]	[-1.46]	[-1.31]	[0.47]	[0.28]
R^2	0.0000	0.0109	0.0279	0.0989	0.5029	0.4660
CHF						
α'	−0.0000	0.0000	0.0001	−0.0004	−0.0029	−0.0019
se(α')	(0.0001)	(0.0006)	(0.0026)	(0.0071)	(0.0213)	(0.0433)
β'	0.4882*	0.5187**	0.6354**	0.8708***	0.9152***	0.9834***
se(β')	(0.2580)	(0.2625)	(0.2710)	(0.2659)	(0.3307)	(0.3263)
$t[\beta' = 1]$	[-1.98]	[-1.83]	[-1.35]	[-0.49]	[-0.26]	[-0.05]
R^2	0.0008	0.0042	0.0246	0.1039	0.2516	0.3266
DEM-EUR						
α'	0.0000	0.0002	0.0010	0.0019	0.0007	0.0053
se(α')	(0.0001)	(0.0005)	(0.0024)	(0.0063)	(0.0208)	(0.0522)
β'	0.5307*	0.5015*	0.6511**	0.8167***	0.9338**	0.8856**
se(β')	(0.2720)	(0.2825)	(0.3060)	(0.3161)	(0.3939)	(0.3886)
$t[\beta' = 1]$	[-1.73]	[-1.76]	[-1.14]	[-0.58]	[-0.17]	[-0.29]
R^2	0.0009	0.0037	0.0251	0.0903	0.2680	0.2490
GBP						
α'	0.0001	0.0004	0.0014	0.0018	0.0148	0.0206
se(α')	(0.0001)	(0.0005)	(0.0018)	(0.0048)	(0.0152)	(0.0352)
β'	0.1791	0.3065	0.5734	0.8253	0.2882	0.7391**
se(β')	(0.4893)	(0.5289)	(0.5965)	(0.7488)	(0.7941)	(0.3450)
$t[\beta' = 1]$	[-1.68]	[-1.31]	[-0.72]	[-0.23]	[-0.90]	[-0.76]
R^2	0.0000	0.0003	0.0036	0.0191	0.0097	0.2438
JPY						
α'	0.0001	0.0004	0.0014	0.0023	0.0066	−0.0109
se(α')	(0.0001)	(0.0007)	(0.0027)	(0.0061)	(0.0189)	(0.0284)
β'	1.1401***	1.0564***	0.8779***	0.7680***	0.9763***	0.9308***
se(β')	(0.2325)	(0.2364)	(0.2510)	(0.2025)	(0.2113)	(0.1433)
$t[\beta' = 1]$	[0.60]	[0.24]	[-0.49]	[-1.15]	[-0.11]	[-0.48]
R^2	0.0063	0.0232	0.0645	0.1042	0.4662	0.7611

Table A.22: Ability to Predict Excess Returns: Sample until December 2006

The table reports results related to the predictive ability of the 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 E 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 December 29, 2006 for AUD; June 1, 1993 to December 29, 2006 for CAD; and September 18, 1989 to December 29, 2006 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	0.5275	0.5668	0.6364	0.7273	0.8384	0.7374	0.5275	0.5668	0.6364	0.7273	0.8384	0.7374
$R2$	0.0027	0.0188	0.0733	0.3407	0.7496	0.3592	0.0020	0.0150	0.0564	0.2996	0.7097	0.3541
p-value[CW]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[0.018]	[0.014]	[<0.01]	[<0.01]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.258]	[0.050]	[0.144]	[<0.01]	[<0.01]	[<0.01]	[0.276]	[0.059]	[0.179]	[<0.01]	[<0.01]	[<0.01]
CAD												
HR	0.5319	0.5406	0.5304	0.5478	0.7391	0.5217	0.5319	0.5406	0.5304	0.5478	0.7391	0.5217
$R2$	0.0006	0.0135	0.0419	0.1247	0.5055	0.4962	-0.0004	0.0086	0.0249	0.0649	0.3952	0.4778
p-value[CW]	[0.243]	[<0.01]	[0.012]	[<0.01]	[<0.01]	[0.016]	[0.496]	[0.020]	[0.049]	[0.011]	[<0.01]	[0.020]
p-value[GW]	[0.241]	[0.184]	[0.153]	[0.015]	[0.024]	[<0.01]	[0.274]	[0.280]	[0.262]	[0.029]	[0.054]	[<0.01]
CHF												
HR	0.5383	0.5516	0.6101	0.6604	0.8302	0.7547	0.5383	0.5516	0.6101	0.6604	0.8302	0.7547
$R2$	0.0008	0.0042	0.0247	0.1039	0.2527	0.3475	-0.0000	0.0005	0.0084	0.0662	0.1484	0.2050
p-value[CW]	[0.044]	[0.037]	[0.017]	[<0.01]	[<0.01]	[<0.01]	[0.188]	[0.157]	[0.089]	[0.012]	[<0.01]	[<0.01]
p-value[GW]	[0.350]	[0.231]	[0.152]	[0.026]	[0.016]	[<0.01]	[0.497]	[0.364]	[0.282]	[0.054]	[0.061]	[<0.01]
DEM-EUR												
HR	0.5354	0.5530	0.5597	0.6101	0.7736	0.7610	0.5354	0.5530	0.5597	0.6101	0.7736	0.7610
$R2$	0.0009	0.0038	0.0256	0.0907	0.2680	0.2501	0.0003	0.0010	0.0120	0.0533	0.1482	0.1395
p-value[CW]	[0.022]	[0.031]	[0.012]	[<0.01]	[<0.01]	[<0.01]	[0.081]	[0.110]	[0.045]	[<0.01]	[<0.01]	[<0.01]
p-value[GW]	[0.368]	[0.192]	[0.074]	[0.017]	[0.021]	[<0.01]	[0.390]	[0.226]	[0.115]	[0.028]	[0.058]	[<0.01]
GBP												
HR	0.5308	0.5487	0.5723	0.6478	0.6604	0.6352	0.5308	0.5487	0.5723	0.6478	0.6604	0.6352
$R2$	0.0004	0.0017	0.0099	0.0334	0.0568	0.4197	-0.0001	-0.0002	0.0008	0.0124	-0.0007	0.4543
p-value[CW]	[0.271]	[0.214]	[0.150]	[0.080]	[0.181]	[<0.01]	[0.262]	[0.198]	[0.115]	[0.023]	[0.107]	[<0.01]
p-value[GW]	[0.066]	[0.532]	[0.339]	[0.214]	[0.178]	[<0.01]	[<0.01]	[0.390]	[0.336]	[0.130]	[0.127]	[<0.01]
JPY												
HR	0.5277	0.5530	0.5786	0.6541	0.7925	0.9560	0.5277	0.5530	0.5786	0.6541	0.7925	0.9560
$R2$	0.0064	0.0235	0.0657	0.1076	0.4712	0.7942	0.0054	0.0194	0.0466	0.0603	0.3464	0.6843
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]	[0.014]	[0.016]	[0.146]	[0.064]	[<0.01]	[<0.01]	[<0.01]	[0.023]	[0.212]	[0.091]	[<0.01]	[<0.01]