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Can country-specific interest rate factors explain the forward premium anomaly?

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Abstract

The forward premium anomaly refers to the fact that changes in spot exchange rates are negatively related to interest rate differentials between home and foreign countries, which is contrary to the predictions of the uncovered interest rate parity (UIRP). We propose a regression model of the interest rate differentials across countries (known as carry trade) adjusted for a time-varying exchange rate risk premium which can explain the anomaly and provide forecasts of exchange rate changes in accordance to the theory. The proposed model is based on estimates of the exchange rate risk premium implied by a simple and empirically attractive two-country affine term structure model with global and local factors. We also show that the forecasting power of the model compares favorably to the random walk model of exchange rates, considered as benchmark in the literature.

Keywords UIRP · Two-country affine term structure model · Forward premium anomaly · Exchange rate forecasting · Expectations hypothesis

JEL Classification G14 · E43 · F31

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1 Introduction

One of the well-known puzzles in the exchange rate literature is the failure of interest rate differentials across countries to forecast the correct direction of future exchange rate movements, predicted by the uncovered interest rate parity (UIRP). In fact, the slope coefficient of a regression of the one-month ahead exchange rate change between two countries on their one-month maturity interest rate differential (known as carry trade) predicts appreciation of the home currency per unit of the foreign currency, instead of depreciation. This means that countries with higher interest rates will face appreciation of their currency, compared to those with lower interest rates. This phenomenon is known in the literature as the forward premium anomaly (or puzzle). For a survey, see Lothian and Wu (2011).

Many theories have been suggested in the literature to explain the anomaly. Among these theories, the existence of a time-varying exchange rate risk premium constitutes a natural one. According to this theory, investors in the FOREX market are not risk neutral, but they are covered for adverse exchange rate movements (see, e.g., Mark 2001).[1] This can obscure investors' expectations embodied in interest rate differentials about future changes of exchange rates, since these differentials are also determined by exchange rate risk premium movements in international bond and money markets. To model the time-variation of this risk premium, recent research relies on the factors and their associated risk-premium sources driving the domestic and foreign term structure of interest rates (see, e.g., Inci and Lu (2004), Ahn (2004), and Ang and Chen (2010), and Diez de los Rios (2011), among others). Based on affine international term structure models, this strand of research is mainly interested in examining whether there exist global and/or country-specific (local) factors that can explain the slope coefficient bias of the UIRP regressions and, hence, the forward premium anomaly. The global factors are often associated with macroeconomic changes in inflation and real activity, linked to world business cycle effects (see Diebold et al. 2006, or more recently Argyropoulos and Tzavalis 2016), while the country-specific ones are assumed to reflect asymmetries in the home (or foreign) economy, which are not common across countries. These asymmetries influence, separately, the asset and bond prices across countries and, as noted by Backus et al. (2001), they are necessary conditions for the explanation of the anomaly.

In this paper, we add to the above literature in two ways: Firstly, we suggest an empirically attractive two-country affine term structure model which provides estimates of the exchange rate risk premium by modelling simultaneously interest rates and exchange rate dynamics. These estimates are then used to examine if time-variation in the risk premium can explain the forward premium anomaly. To this end, we suggest an extension of the UIRP regression model adjusted for time-varying risk

[1] Alternative theories include the following: peso problems due to missed regime shifts by investors (see Evans and Karen 1995), exchange rate stabilization monetary policy rules smoothing out exchange rate changes (see McCallum 1994), sticky prices yielding time-varying real exchange rate deviations from the PPP (purchasing power parity) (see, e.g., see Meese and Rogoff 1988, and more recently Boudoukh et al. 2016) and measurement errors combined with small magnitudes of interest rate differentials, see, e.g., Lothian and Wu (2011).

premium effects. Secondly, we evaluate the performance of the model to forecast future exchange rate movements net of the risk premium effects. Using data from five developed economies (Germany, Japan, Canada, the US and UK), we show that the suggested regression model provides forecasts of the future exchange rate movements which are in accordance with the expectations hypothesis. We also find that, in terms of forecasting power, the suggested model compares favorably, in terms of the root mean square and absolute forecasting error metrics, to the random walk model of exchange rates which is considered as the benchmark model in the literature (see Rossi 2013 for a survey).

The rest of the paper is organized as follows. In Section 2, we present the model. In Section 3, we provide estimation results of the model, as well as the UIRP regression model adjusted for the risk premium effects and we evaluate the forecasting performance of the model. Finally, Section 4 concludes the paper.

2 A two-country affine term structure model

We consider a two-country essentially affine term structure model where assets are denominated in either home (h) or foreign (f) currencies. The prices of bonds and interest rates of both countries are driven by two common factors (state variables), denoted as x_{1t} and x_{2t}, reflecting common macroeconomic conditions across countries, and one local factor, denoted as $x_{3t}^{(j)}$, reflecting differences (or asymmetries) across countries.[2] The above variables, collected in vector $X_t^{(j)} = (x_{1t}, x_{2t}, x_{3t}^{(j)})'$, $j = \{h, f\}$, follow the Gaussian processes:

$$\begin{pmatrix} dx_{1t} \\ dx_{2t} \\ dx_{3t}^{(j)} \end{pmatrix} = \begin{pmatrix} k_{11} & 0 & 0 \\ 0 & k_{22} & 0 \\ 0 & 0 & k_{33}^{(j)} \end{pmatrix} \left[\begin{pmatrix} \theta_1 \\ \theta_2 \\ \theta_3^{(j)} \end{pmatrix} - \begin{pmatrix} x_{1t} \\ x_{2t} \\ x_{3t}^{(j)} \end{pmatrix} \right] dt + \begin{pmatrix} \sigma_{11} & 0 & 0 \\ 0 & \sigma_{22} & 0 \\ 0 & 0 & \sigma_{33}^{(j)} \end{pmatrix} \begin{pmatrix} dW_{1t} \\ dW_{2t} \\ dW_{3t}^{(j)} \end{pmatrix}, \tag{1}$$

where $W_{it}^{(j)}$, $i = \{1, 2, 3\}$, denote uncorrelated Wiener processes, $k_{ii}^{(j)}$ and $\sigma_{ii}^{(j)}$ denote the mean reversion and variance parameters of $x_{it}^{(j)}$, while $\theta_i^{(j)}$ is the long-run mean of $x_{it}^{(j)}$.

The current price of a zero-coupon nominal bond of country j with maturity interval τ, denoted as $B_t^{(j)}(\tau)$, paying one unit of currency j in period $t + \tau$, can be derived based on the following Euler equation:

$$B_t^{(j)}(\tau) = E_t \left(\frac{M_{t+\tau}^{(j)}}{M_t^{(j)}} \right), \tag{2}$$

[2] This assumption is also consistent with affine term structure models assuming that three factors can explain the total term structure variation (see, e.g., Dai and Singleton 2002, or more recently Argyropoulos and Tzavalis 2015).

where $E_t(.)$ is the conditional expectations operator and $M_t^{(j)}$ is the pricing kernel. The latter follows the stochastic process:

$$\frac{dM_t^{(j)}}{M_t^{(j)}} = -r_t^{(j)}dt - \Lambda_t^{(j)\prime} dW_t^{(j)}, \tag{3}$$

where $W_t^{(j)} = (W_{1t}, W_{2t}, W_{3t}^{(j)})'$ and $r_t^{(j)}$ is the instantaneous nominal interest rate specified as

$$r_t^{(j)} = x_{1t} + x_{2t} + x_{3t}^{(j)} = \delta_1^{(j)\prime} X_t^{(j)}, \tag{4}$$

where $\delta_1^{(j)} = (1, 1, 1)'$ is a standard normalization condition often made in the literature (e.g. Kim and Orphanides 2012), and $\Lambda_t^{(j)}$ is a vector of risk pricing functions depending on $X_t^{(j)}$. We specify $\Lambda_t^{(j)}$ to be in accordance to Eq. 1 as follows:

$$\Lambda_t^{(j)} = \Sigma^{(j)-1}\left(\lambda_0^{(j)} + \lambda_1^{(j)} X_t^{(j)}\right), \tag{5}$$

where $\Sigma^{(j)} = \begin{pmatrix} \sigma_{11} & 0 & 0 \\ 0 & \sigma_{22} & 0 \\ 0 & 0 & \sigma_{33}^{(j)} \end{pmatrix}$, $\lambda_0^{(j)} = \begin{pmatrix} \lambda_{0,1} \\ \lambda_{0,2} \\ \lambda_{0,3}^{(j)} \end{pmatrix}$ and $\lambda_1^{(j)} = \begin{pmatrix} \lambda_{1,11} & 0 & 0 \\ 0 & \lambda_{1,22} & 0 \\ 0 & 0 & \lambda_{1,33}^{(j)} \end{pmatrix}$.

Based on Eqs. 3 and 5, we can derive from Eq. 2 the following relationship for the nominal term structure of interest rates of country j and maturity interval τ, denoted as $R_t^{(j)}(\tau)$:

$$R_t^{(j)}(\tau) = A^{(j)}(\tau) + D_1(\tau)x_{1t} + D_2(\tau)x_{2t} + D_3^{(j)}(\tau)x_{3t}^{(j)}, \tag{6}$$

where $R_t^{(j)}(\tau) = -\frac{1}{\tau}\log B_t^{(j)}(\tau)$ and

$$D_i^{(j)}(\tau) = (1/\tau)\left(1 - e^{-(k_{ii}^{(j)}+\lambda_{1,ii}^{(j)})\tau}\right)\left(k_{ii}^{(j)} + \lambda_{1,ii}^{(j)}\right)^{-1}\delta_1^{(j)},$$

for $i = \{1, 2, 3\}$, and $A^{(j)}(\tau)$ is a scalar function.[3] This relationship can be used to obtain an analytic formula of the forward risk premium, depending on state variables $x_{3t}^{(j)}$, $j = \{h, f\}$. To this end, define the exchange rate S_t as the number of units of home currency (e.g., US dollars) per unit of foreign currency. Then, by the law of one price, the bond price $B_t^{(j)}(\tau)$ must be correctly priced in both the home and foreign markets, otherwise profitable arbitrage opportunities will arise. In terms of pricing relationship Eq. 2, this law means

$$E_t\left(\frac{M_{t+\tau}^{(f)}}{M_t^{(f)}}\right) = E_t\left(\frac{S_{t+\tau}}{S_t}\frac{M_{t+\tau}^{(h)}}{M_t^{(h)}}\right),$$

[3] See Dai and Singleton (2002).

which implies

$$M_t^{(f)} = S_t M_t^{(h)}, \text{ or } m_t^{(f)} = s_t + m_t^{(h)}, \tag{7}$$

where $m_t^{(j)} \equiv \log M_t^{(j)}$ and $s_t \equiv \log S_t$. The above equation shows that the exchange rate S_t (or its logarithm s_t) is uniquely determined by pricing kernels $M_t^{(j)}$, $j = \{h, f\}$ (or $m_t^{(j)}$ respectively). The stochastic process of s_t can be derived by applying Ito's lemma to Eq. 3. This yields

$$dm_t^{(j)} = -\left(r_t^{(j)} - \frac{1}{2}\Lambda_t^{(j)\prime}\Lambda_t^{(j)}\right)dt - \Lambda_t^{(j)} dW_t^{(j)}.$$

Using $m_t^{(f)} = s_t + m_t^{(h)}$, the last relationship implies the following stochastic process for s_t:

$$ds_t = \left((r_t^{(h)} - r_t^{(f)}) + \frac{1}{2}(\Lambda_t^{(h)\prime}\Lambda_t^{(h)} - \Lambda_t^{(f)\prime}\Lambda_t^{(f)})\right)dt + \left(\Lambda_t^{(h)} - \Lambda_t^{(f)}\right)' dW_t^{(j)}. \tag{8}$$

Taking the conditional expectation of this process, at time t, gives

$$E_t(ds_t) = \left((r_t^{(h)} - r_t^{(f)}) + \frac{1}{2}(\Lambda_t^{(h)\prime}\Lambda_t^{(h)} - \Lambda_t^{(f)\prime}\Lambda_t^{(f)})\right)dt. \tag{9}$$

The last relationship clearly shows that the expected exchange rate change $E_t(ds_t)$ equals the interest rate differential $(r_t^{(h)} - r_t^{(f)})$ plus a component, given as $\wp_t(\tau) = \frac{1}{2}(\Lambda_t^{(h)\prime}\Lambda_t^{(h)} - \Lambda_t^{(f)\prime}\Lambda_t^{(f)})dt$, which represents the exchange rate risk premium. The sources of this premium are those of the nominal term structures of the home and foreign countries, captured by risk price functions $\Lambda_t^{(j)}$, $j = \{h, f\}$. Substituting Eqs. 5 into 9 yields the following relationship of $\wp_t(\tau)$:

$$\begin{aligned}\wp_t(\tau) = \frac{1}{2}\Bigg[&\left(\lambda_{0,3}^{(h)}\right)^2 - \left(\lambda_{0,3}^{(f)}\right)^2 + 2\left(\lambda_{0,3}^{(h)}\lambda_{1,33}^{(h)}x_{3t}^{(h)} - \lambda_{0,3}^{(f)}\lambda_{1,33}^{(f)}x_{3t}^{(f)}\right)\\ &+ \left(\lambda_{1,33}^{(h)}\right)^2\left(x_{3t}^{(h)}\right)^2 - \left(\lambda_{1,33}^{(f)}\right)^2\left(x_{3t}^{(f)}\right)^2\Bigg]dt\end{aligned} \tag{10}$$

which shows that time-variation of $\wp_t(\tau)$ is due to local factors $x_{3t}^{(j)}$, $j = \{h, f\}$, and their squared terms $(x_{3t}^{(j)})^2$, determining $R_t^{(j)}(\tau)$. From this relationship, we can easily see that time-variation of $\wp_t(\tau)$ obscures the true information of $(r_t^{(h)} - r_t^{(f)})$ about future exchange rate changes $E_t(ds_t)$, and thus violates the predictions of the UIRP about $E_t(ds_t)$.[4] This is one of the major explanations for the forward premium anomaly, offered in the literature (see Fama 1980, Baillie and Bollerslev 2000, or more recently Coakley et al. 2004).

[4]Similar relationships to Eq. 10 have been derived by Backus et al. (2001) and Ang and Chen (2010) based on stochastic discount factor approach in the discrete-time framework.

3 Estimation of the model

The two-country term structure model, presented in the previous section, can be used to retrieve estimates of forward risk premium $\wp_t(\tau)$ and then to investigate if it can explain the forward premium anomaly. To this end, we estimate the following system of equations, simultaneously:

$$R_{t+1}^{(j)}(\tau) = A^{(j)}(\tau) + D^{(j)}(\tau)' X_{t+1}^{(j)} + e_{t+1}^{(j)}(\tau),\ j = \{h, f\}, \tag{11}$$

$$\Delta x_{t+1}^{(j)} = \theta_i^{(j)}(1 - e^{-k_{ii}^{(j)}\Delta t}) + (e^{-k_{ii}^{(j)}\Delta t} - 1)x_{it}^{(j)} + \omega_{i,t+1}^{(j)},\ i = \{1, 2, 3\}, \tag{12}$$

$$\begin{aligned}\Delta s_{t+1} - (r_t^{(h)} - r_t^{(f)}) = &\frac{1}{2}\Bigg[\left(\lambda_{0,3}^{(h)}\right)^2 - \left(\lambda_{0,3}^{(f)}\right)^2 + 2\left(\lambda_{0,3}^{(h)}\lambda_{1,33}^{(h)}x_{3t}^{(h)} - \lambda_{0,3}^{(f)}\lambda_{1,33}^{(f)}x_{3t}^{(f)}\right)\\ &+ \left(\lambda_{1,33}^{(h)}\right)^2\left(x_{3t}^{(h)}\right)^2 - \left(\lambda_{1,33}^{(f)}\right)^2\left(x_{3t}^{(f)}\right)^2\Bigg] + \zeta_{t+1}.\end{aligned} \tag{13}$$

where $D^{(j)}(\tau) = \left(D_1(\tau), D_2(\tau), D_3^{(j)}(\tau)\right)'$, $r_t^{(h)} - r_t^{(f)} = x_{3t}^{(h)} - x_{3t}^{(f)}$ (see Eq. 4) and $e_{t+1}^{(j)}(\tau)$, $\omega_{i,t+1}^{(j)}$ and ζ_{t+1} constitute zero-mean error terms. Equation 11 corresponds to the affine term structure relationship (6) and (13) is a discretization of Eq. 10 for the process of exchange rate one period ahead based on Euler's method and Eq. 12 is a discretization of the processes of Eq. 1, for $i = \{1, 2, 3\}$. We estimate Eq. 11 for $\tau = \{3, 6\}$ months, given that the term structure of interest rates relationships (6) for $\tau = \{1, 12\}$ months are used in the estimation of the above system as identities to retrieve values of x_{1t}, x_{2t} and $x_{3t}^{(j)}$ from term structure of interest rates data, by inverting (6).[5]

We estimate (11)–(13) as a system of seemingly unrelated regressions (SUR) based on the non-linear least squares (NLLS) method. This method provides robust results to density misspecification of the error terms of the above system and accounts for cross-equation correlation. To allow for possible trivial bond market frictions, which may affect the term structure data, we leave unrestricted the intercept of regression (11) of the system. Finally, note that, corresponding to enormous evidence in the literature that one of the state variables of term structure models capture level effects (see, e.g., Chen and Tsang 2013), we also impose the restriction $k_{11} + \lambda_{1,11} = 0$ in the estimation of the system, implying $D_1(\tau) = 1$, for all τ. This restriction means that x_{1t} plays the role of the level term structure factor.

Table 1 presents estimates of the system of Eqs. 11–13 for the US dollar (USD) against the currencies of four other countries, namely Germany (DM/Euro), Canada

[5]Note that the common factor structure of x_{1t} and x_{2t} across countries allows us to retrieve local factors $x_{3t}^{(j)}$ based on the spreads of interest rates $R_t^{(h)}(\tau) - R_t^{(f)}(\tau)$, for $\tau = \{\tau_1, \tau_2\}$, by inverting the following relationships:

$$R_t^{(h)}(\tau_1) - R_t^{(f)}(\tau_1) = A^{(h)}(\tau_1) - A^{(f)}(\tau_1) + D_3^{(h)}(\tau_1)x_{3t}^{(h)} - D_3^{(f)}(\tau_1)x_{3t}^{(f)}$$
$$R_t^{(h)}(\tau_2) - R_t^{(f)}(\tau_2) = A^{(h)}(\tau_2) - A^{(f)}(\tau_2) + D_3^{(h)}(\tau_2)x_{3t}^{(h)} - D_3^{(f)}(\tau_2)x_{3t}^{(f)}$$

Then, given $x_{3t}^{(h)}$ and $x_{3t}^{(f)}$, we can retrieve x_{1t} and x_{2t} by also inverting (6), for τ_1 and τ_2.

Table 1 Estimates of system (11)–(13)

$x_{it}^{(j)}$	x_{1t}	x_{2t}	$x_{3t}^{(h)}$	$x_{3t}^{(f)}$	x_{1t}	x_{2t}	$x_{3t}^{(h)}$	$x_{3t}^{(f)}$
	USD/DM(Euro)				USD/CAD			
$k_{ii}^{(j)}$	9.41***	7.87***	0.94***	17.47	4.08***	0.62***	0.54	8.99***
	(0.19)	(1.26)	(0.26)	(15.50)	(0.53)	(0.15)	(0.40)	(2.05)
$\lambda_{0,i}^{(j)}$	2.33***	−5.88***	−1.41	2.89***	5.64***	−3.62***	1.69**	0.14
	(0.00)	(0.00)	(1.29)	(0.96)	(0.00)	(0.00)	(0.84)	(1.07)
$\lambda_{1,ii}^{(j)}$	−9.41***	−1.37	−0.86***	−0.33	−4.08***	5.35***	0.40	1.16***
	(0.19)	(2.47)	(0.25)	(0.65)	(0.53)	(3.18)	(0.41)	(0.31)
	USD/GBP				USD/JPY			
$k_{ii}^{(j)}$	8.98***	3.02***	0.11	18.62***	9.52***	8.44***	1.09***	17.54
	(0.16)	(0.55)	(0.40)	(2.86)	(0.22)	(1.93)	(0.43)	(12.35)
$\lambda_{0,i}^{(j)}$	−9.47***	14.04***	3.38***	0.17	3.04***	−6.68***	−1.16	2.91
	(0.00)	(0.00)	(0.64)	(0.98)	(0.00)	(0.00)	(3.43)	(2.05)
$\lambda_{1,ii}^{(j)}$	−8.98***	4.09**	0.93***	1.35***	−9.52***	−2.04	−0.72*	0.74*
	(0.16)	(1.70)	(0.24)	(0.33)	(0.22)	(2.57)	(0.38)	(0.42)

Standard errors are in parentheses. (*,**,***) indicate significance at the 10%, 5% and 1% level respectively

(CAD), UK (GBP) and Japan (JPY).[6] Note that we present estimates of the key coefficients of the model $k_{ii}^{(j)}$, $\lambda_{0,i}^{(j)}$ and $\lambda_{1,ii}^{(j)}$, for reasons of space. Asymptotic standard errors are in parentheses. For reasons of space, the table presents estimates of the key parameters of the model which imply significant time-variation in the forward risk premium $\wp_t(\tau)$ and affect its dynamics, captured by state variables $x_{3t}^{(j)}$ (i.e., $k_{ii}^{(j)}$ and, $\lambda_{0,i}^{(j)}$ and $\lambda_{1,ii}^{(j)}$). Our estimates are based on monthly data covering the period 1979:01 – 2018:01. All data series are from Datastream. For $R_t^{(j)}(\tau)$, we use Eurocurrency deposit interest rates of $\tau = \{1, 3, 6, 12\}$ months.[7] Both $R_t^{(j)}(\tau)$ and exchange rate changes Δs_t are given at annual basis and in percentage terms. In the Appendix, we present a table of correlation coefficients among interest rates $R_t^{(j)}(\tau)$, for all countries j and maturity intervals τ (see Table 6). The high (close to unity) values of these coefficients indicate that there is strong correlation among $R_t^{(j)}(\tau)$ across all countries. This result is consistent with the common factor assumptions of our suggested two-country term structure model.

[6]For Germany, we have used the euro conversion rate of the Deutsche Mark (DM) (i.e., 1.95583) to participate in the EMS to calculate the series of the currency rate of this country with respect to USD after the introduction of Euro as the single currency in 1999. See also Diez de los Rios and Sentana (2011).

[7]Note that Eurocurrency deposits are essentially zero-coupon bonds whose payoffs at maturity are the principal plus the interest payment. Eurocurrency deposit rates are used in many studies testing the predictions of the UIRP (see, e.g., Olmo and Pilbeam 2011).

The results of Table 1 show that the estimates of $k_{ii}^{(j)}$ and $\lambda_{0,i}^{(j)}$ or $\lambda_{1,ii}^{(j)}$ tend to be statistically significant at the 5%, or less, level. This is true for all four pairs of currencies examined. For all these pairs, either the estimates of $\lambda_{0,i}^{(j)}$ or $\lambda_{1,ii}^{(j)}$ are statistically significant, implying a significant time-variation of the exchange rate risk premium $\wp_t(\tau)$, as predicted by relationship (10). Note that, under the assumptions of the affine term structure models, the sign of the estimates of coefficients $\lambda_{0,i}^{(j)}$ and $\lambda_{1,ii}^{(j)}$ is not necessary to be positive, or negative (see, e.g., Dai and Singleton 2002). In the Appendix, based on the results of Table 1, we graphically present estimates of series $\wp_t(\tau)$ (see Fig. 1) and descriptive statistics of them, namely, their mean, standard deviation and max-min values (see Table 6) as well as their correlation coefficients between all pairs of $\wp_t(\tau)$, for all currencies considered (see Table 5). These results indicate that there are substantial movements of $\wp_t(\tau)$ over time, for all j. The low values of the correlation coefficients reported in Table 5 can be attributed to the influence of the local factors $x_{3t}^{(j)}$, $j = \{h, f\}$, driving $\wp_t(\tau)$, reflecting asymmetries in the bond and money markets across the home and foreign countries.

3.1 UIRP regressions adjusted for time-varying risk premium effects

Having estimated risk premium $\wp_t(\tau)$, at each point t, based on the system of Eqs. 11–13, next we examine if the empirical failures of the UIRP to predict the correct direction of exchange rate changes Δs_{t+1} can be attributed to the time-variation of this premium. To this end, we estimate the following regression model based on the least squares (LS) method:

$$\Delta s_{t+1} = a + \beta \left(R_t^{(h)}(1) - R_t^{(f)}(1) - \wp_t(\tau) \right) + \varepsilon_{t+1}, \tag{14}$$

where $\wp_t(\tau)$ is obtained by Eq. 10, based on the values of $\lambda_{0,i}^{(j)}$ and $\lambda_{1,ii}^{(j)}$, as well as $x_{3t}^{(j)}$, $j = \{h, f\}$, obtained by the estimation of system (11)–(13). Note that, in regression (14), we subtract $\wp_t(\tau)$ from interest rate differentials $R_t^{(h)}(1) - R_t^{(f)}(1)$ to adjust for the effects of $\wp_t(\tau)$ on spread $R_t^{(h)}(1) - R_t^{(f)}(1)$, directly. This can be justified by the covered interest rate parity (CIRP), predicting that $R_t^{(h)}(1) - R_t^{(f)}(1)$ contains joint information about $\wp_t(\tau)$ and $E_t(\Delta s_{t+1})$ (see Eq. 9).[8]

Estimation of regression model (14) without adjusting for the effects of $\wp_t(\tau)$ on $R_t^{(h)}(1) - R_t^{(f)}(1)$ will lead to biased estimates of slope coefficient β. To see this more clearly, consider regression model (14) ignoring $\wp_t(\tau)$, i.e.,

$$\Delta s_{t+1} = a + \beta \left(R_t^{(h)}(1) - R_t^{(f)}(1) \right) + u_{t+1} \tag{15}$$

[8]Note that $\wp_t(\tau)$ constitutes a good proxy of the exchange rate risk premium implied by interest rate differential $R_t^{(h)}(1) - R_t^{(f)}(1)$, for a small maturity τ and interval of time. This regression model corresponds to Eq. 9 - see Ahn (2004), and it is in the spirit of the regression models of Tzavalis (2003), and Argyropoulos and Tzavalis (2019) adjusting the term spread of interest rates for risk premium effects in order to forecast future changes in short-term interest rates or inflation rate changes, respectively.

where $u_{t+1} = \varepsilon_{t+1} - \beta \wp_t(\tau)$. This regression model is often employed in the literature to test the prediction of the UIRP under the rational expectations hypothesis, which implies $\beta = 1$. The structural form of the error term u_{t+1}, i.e., $u_{t+1} = \varepsilon_{t+1} - \beta \wp_t(\tau)$, implies that the LS estimates of β based on regression (15) will be biased due to contemporaneous correlation between the regressor term $R_t^{(h)}(1) - R_t^{(f)}(1)$ and risk premium term $\wp_t(\tau)$, entered into u_{t+1}. As Eqs. 6 and 10 indicate, both $R_t^{(h)}(1) - R_t^{(f)}(1)$ and $\wp_t(\tau)$ are affine functions of common state variables x_{1t}, x_{2t} and $x_{3t}^{(j)}$, $j = \{h, f\}$, and thus will be contemporaneously correlated. In the literature, one stream of research (see, e.g., Ahn 2004) is focused on investigating if the bias of the slope coefficient β of model (15) due to $\wp_t(\tau)$ can be explained by a two-currency affine term structure model. Instead, our approach, based on regression model (14), enables us to test if the UIRP allowing for a time-varying risk premium is satisfied by the data, based on sample estimates of slope coefficient β.

Table 2 presents estimates of the slope coefficients a and β of regression model (14). Since $\wp_t(\tau)$ is estimated by the data and, thus, $R_t^{(h)}(1) - R_t^{(f)}(1) - \wp_t(\tau)$ constitutes a generated regressor, the table also presents estimates of a and β based on the instrumental variables (IV) estimator, using as instruments lagged values of interest rate differentials, for $\tau = \{3, 6, 12\}$. We test the validity of these instruments based on the overidentifying restrictions J-test, whose p-values are reported in the table. Note that, for comparison reasons, the table also presents estimation results for the standard UIRP model (15), which does not adjust for risk premium effects. The IV estimates for this regression model may also deal with the problem of the contemporaneous correlation between regression term $R_t^{(h)}(1) - R_t^{(f)}(1)$ and $\wp_t(\tau)$, omitted from this regression.[9] This obviously requires the use of valid instrumental variables, which are uncorrelated with $\wp_t(\tau)$.

The results of the table indicate that the time-variation of $\wp_t(\tau)$ can indeed explain the empirical failure of the UIRP. The regression model (14), adjusted for $\wp_t(\tau)$, provides estimates of the slope coefficient β which are in line with the expectations hypothesis. Indeed, the estimates of β are close to unity and significant at the 5%, or less, level, for all countries considered, with the exception of Canada where they are not significant. But, even for this country the estimates of β have the correct sign and they are close to unity. These results hold for both the LS and IV estimates reported in the table. The very low values of the coefficient of determination R^2, reported in the table for both the unadjusted and adjusted for forward risk premium UIRP regression models estimated, can be attributed to the high volatility of exchange rates. This is consistent with the empirical literature of exchange rate models (see, e.g., Engel 2014). Finally, note that the IV estimates of the standard UIRP regression cannot remove the bias in β which can be attributed to the problem of invalid instruments, as argued before. This is one of the merits of using model (14) to test the predictions of the UIRP. Since $R_t^{(h)}(1) - R_t^{(f)}(1)$ is adjusted by $\wp_t(\tau)$, there is no need to rely on instrumental variables to remove the risk premium bias of the slope coefficient β.

[9] This method has been suggested in the term structure of interest rates literature to capture the effects of the risk premium on the term spread in forecasting future short-term interest rates by Driffill et al. (1997).

Table 2 UIRP regressions adjusted for risk premium effects

	USD/DM(Euro)		USD/CAD		USD/GBP		USD/JPY	
	Estimates of model: $\Delta s_{t+1} = \alpha + \beta(R_t^{(h)}(1) - R_t^{(f)}(1)) + \varepsilon_{t+1}$							
	LS	IV	LS	IV	LS	IV	LS	IV
a	0.43	0.28	−0.53	0.21	−3.38*	−4.44*	6.35***	6.07***
	(1.90)	(2.11)	(1.27)	(1.12)	(1.91)	(2.72)	(2.41)	(2.43)
β	−0.79	−0.77	−0.63	−0.84	−1.56*	−2.05	−1.85***	−1.74***
	(0.78)	(0.90)	(0.57)	(0.62)	(0.95)	(1.39)	(0.68)	(0.70)
R^2	0.003		0.001		0.010		0.015	
	Estimates of model: $\Delta s_{t+1} = \alpha + \beta(R_t^{(h)}(1) - R_t^{(f)}(1) - \wp_t) + u_{t+1}$							
	LS	IV	LS	IV	LS	IV	LS	IV
a	−0.03	−0.42	0.20	0.16	2.74	3.17	−0.22	−0.20
	(1.70)	(1.74)	(1.14)	(1.19)	(2.17)	(2.94)	(1.95)	(2.11)
β	0.98***	1.02**	0.64	0.66	1.06***	1.22**	0.94***	0.98***
	(0.40)	(0.52)	(0.45)	(1.02)	(0.01)	(0.61)	(0.33)	(0.46)
$pval_{J-test}$		0.84		0.20		0.14		0.16
R^2	0.017		0.003		0.023		0.019	

Standard errors are in parentheses. (*,**, ***) indicate significance at the 10%, 5% and 1% level respectively. $pval_{J-test}$ are the p-values of the J-test

To see whether the estimation results for model (14) remain robust to proxy variables for risk premium $\wp_t(\tau)$ suggested in the literature (see, e.g., Chen and Tsang 2013), in Table 3 we present LS and IV estimates of both the adjusted and unadjusted UIRP regression models, estimated before, which also include in their right hand side (RHS) such proxies. In particular, these variables include changes over time of the following variables: $R_t^{(h)}(1)-R_t^{(f)}(1)$, denoted $\Delta(Short)$, $\left(R_t^{(h)}(120) - R_t^{(h)}(1)\right) - \left(R_t^{(f)}(120) - R_t^{(f)}(1)\right)$, denoted $\Delta(Term)$, and $L_t^{(h)} - L_t^{(f)}$, denoted $\Delta(Level)$, where $L_t^{(j)}$ is the average of interest rates with maturities $\tau = \{1, 120\}$ months. The results of the table clearly indicate that the estimates of the slope coefficient β for both the adjusted and unadjusted UIRP regression models remain unaffected by the presence of the above proxies. These results mean that the adjusted model (14) is adequate to capture the exchange rate risk premium bias of spread $R_t^{(h)}(1) - R_t^{(f)}(1)$ in forecasting future exchange rate changes Δs_{t+1}. Another interesting conclusion that can be drawn from the results of Table 3 is that the proxy variables for $\wp_t(\tau)$ cannot capture the bias of $R_t^{(h)}(1) - R_t^{(f)}(1)$ alone.

3.2 Out-of-sample forecasting performance

In this section we examine if the regression model (14) also improves upon the performance of the standard UIRP regression model, unadjusted for forward risk premium

Table 3 UIRP regressions with additional regressors

	USD/DM(Euro)		USD/CAD		USD/GBP		USD/JPY	
Regressors	Estimates of model (14) with additional regressors							
	LS	IV	LS	IV	LS	IV	LS	IV
$R_t^{(h)}(1) - R_t^{(f)}(1) - \wp_t(\tau)$	1.08***	1.10**	0.44	0.31	1.06***	1.55**	1.10*	1.23*
	(0.41)	(0.56)	(0.45)	(1.22)	(0.39)	(0.78)	(0.67)	(0.76)
$\Delta(Short)$	−2.13	−6.15*	−2.52	−4.08**	−0.18	−0.40	−7.45	−14.81
	(2.02)	(3.38)	(1.56)	(1.97)	(1.72)	(2.39)	(7.16)	(9.53)
$\Delta(Term)$	−11.60*	−17.54	19.89***	24.41*	−0.99	5.69	−23.39*	−28.41
	(6.77)	(24.65)	(7.09)	(14.30)	(6.21)	(26.03)	(12.48)	(26.88)
$\Delta(Level)$	0.34	−5.40	10.82***	14.44**	6.17*	4.20	−9.01	−21.16*
	(3.87)	(13.27)	(3.95)	(7.41)	(3.42)	(10.69)	(5.86)	(12.08)
$pval_{J-test}$		0.78		0.13		0.09		0.09
R^2	0.038		0.026		0.054		0.028	
	Estimates of the standard UIRP regression with additional regressors							
	LS	IV	LS	IV	LS	IV	LS	IV
$R_t^{(h)}(1) - R_t^{(f)}(1)$	−1.06	−1.10	−0.48	−0.74	−2.14**	−1.68*	−1.20	−1.12
	(0.78)	(0.92)	(0.60)	(0.74)	(0.93)	(1.00)	(0.93)	(1.00)
$\Delta(Short)$	−1.90	−6.31*	−2.38	−3.07	0.78**	0.16	−7.17	−16.07
	(2.10)	(3.54)	(1.63)	(2.09)	(2.03)	(2.68)	(7.37)	(12.47)
$\Delta(Level)$	−9.25	−10.72	19.70***	18.40	−1.32	4.87	−21.58*	−13.37
	(6.79)	(23.11)	(6.99)	(14.93)	(6.19)	(17.21)	(12.47)	(37.56)
$\Delta(Term)$	2.10	−0.35	11.30***	12.81***	6.80**	10.19	−9.66*	−37.64**
	(3.98)	(11.58)	(3.84)	(7.42)	(3.36)	(8.27)	(5.93)	(18.98)
$pval_{J-test}$		0.73		0.15		0.10		0.28
R^2	0.025		0.026		0.049		0.024	

Standard errors are in parentheses. (*,**, ***) indicate significance at the 10%, 5% and 1% level respectively. $pval_{J-test}$ are the p-values of the J-test

effects, to forecast the future exchange rate change Δs_{t+1}. To this end, we carry out an out-of-sample forecasting exercise comparing the performance of these models to the random walk with drift. The latter constitutes a benchmark metric for evaluating exchange rate models (see, e.g., Hodrick 1987). In our exercise, we also consider the regression model (14) which is augmented with variables $\Delta(Short)$, $\Delta(Term)$ and $\Delta(Level)$, used as proxies of $\wp_t(\tau)$.

To derive forecasts of the above models, we rely on recursive estimates of them. That is, we estimate the models based on an initial in-sample window of observations and then re-estimate them until the end of the sample by adding one observation at a time. This is repeated for an out-of-sample interval covering the period from 2016:01 to 2018:12. Note that the above estimation procedure of model (14) also requires

Table 4 Out-of-sample performance

	USD/DM(Euro)	USD/CAD	USD/GBP	USD/JPY
	$M1$: Regression (14)			
MAE	18.83	21.34	18.31	27.44
$RMSE$	22.40	24.40	28.40	37.74
$DM_{(M1 \text{ vs } M2)}$	−0.71	−1.50**	−1.51**	2.11**
$DM_{(M1 \text{ vs } M3)}$	−0.82	−3.44***	−1.90**	1.18
$DM_{(M1 \text{ vs } M4)}$	−0.50	−0.14	−0.37	0.60
	$M2$: Standard UIRP Regression			
MAE	19.08	21.49	19.05	26.81
$RMSE$	22.78	24.52	28.50	37.06
$DM_{(M2 \text{ vs } M3)}$	−0.76	−3.48***	−1.96**	0.53
$DM_{(M2 \text{ vs } M4)}$	1.48*	1.40*	1.64**	−0.30
	$M3$:UIRP Regression augmented with proxies for $\wp_t(\tau)$			
MAE	19.25	22.96	19.64	26.31
$RMSE$	23.01	26.27	28.93	35.66
$DM_{(M3 \text{ vs } M4)}$	1.02	3.66**	1.96**	−1.40***
	$M4$: RW model			
MAE	18.99	21.36	21.76	27.01
$RMSE$	22.65	24.36	34.16	36.93

$RMSE$ and MAE are the root mean square and mean absolute error metrics, respectively

DM denotes the Diebold-Mariano test statistic. (*, **, ***) indicate significance at the 10%, 5% and 1% level respectively

recursive estimation of the system of Eqs. 11–13 to obtain recursive estimates of the forward premium $\wp_t(\tau)$, too.

The results of our forecasting exercise are reported in Table 4. To evaluate the forecasting performance of the alternative models considered, the table reports values of the root mean square and absolute errors metrics, denoted as $RMSE$ and MAE, respectively. It also reports Diebold and Mariano (1995) test statistic (denoted DM) of each model against the other three models considered, separately. This statistic can test whether there exist significant differences between the forecasts of two non-nested models (say $M1$ and $M2$). A negative sign and statistically significant value of DM means that model $M1$ has significant forecasting power compared to model $M2$.[10]

[10]Note that, for models $M1$ and $M2$, the DM test statistic is based on the forecast loss function $d_{t+j} = \left(u_{t+j}^{(M1)}\right)^2 - \left(u_{t+j}^{(M2)}\right)^2$. Given d_{t+j}, DM is defined as $DM = \left(\frac{1}{T-T_0}\sum_{j=T_0+1}^{T} d_{t+j}\right)\hat{\sigma_d}^{1/2}$, where T_0 is the initial, in-sample window of the sample and $\hat{\sigma_d}$ is the long-run variance of d_{t+j}, which can be consistently estimated based on Newey and West (1987) estimator.

The results of the table lead to the following conclusions: First, in terms of metrics $RMSE$ and MAE, the forecasting performance of regression model (14) (denoted as $M1$) is better than the standard UIRP regression model (denoted as $M2$), which doesn't allow for forward risk premium effects. This is true for three out of four foreign currencies considered, namely GBP, DM(Euro) and CAD, with the exception of JPY. Secondly, for these three currencies, model (14) has also better forecasting power than the UIRP model augmented with the forward risk premium proxy variables (denoted as $M3$) based on $RMSE$ and MAE. Note that, for CAN and GBP, the forecasting superiority of model $M1$ compared to models $M2$ and $M3$ can be also more formally justified by the values of the test statistic DM, reported in the table, which are significant at 5%, or less, level. For JPY, the results of the table indicate that the best model in terms of metrics $RMSE$ and MAE is model $M3$. Third, regarding comparisons to the random walk model (denoted as $M4$), the results of the table indicate that, for currencies GBP, DM(Euro) and CAD, model $M1$ compares favorably to model $M4$ in terms of metrics $RMSE$ and MAE, while, for JPY, this holds for model $M3$. The values of DM do not indicate any significant superiority either for model $M1$ or $M4$, at the 5% level.

4 Conclusions

This paper contributes into the exchange rate literature by suggesting a regression model of interest rate differentials across countries to forecast future exchange rate changes which is adjusted for the exchange rate risk premium effects. To capture these effects, the paper is based on a simple and empirically attractive model of exchange rate risk premium implied by a two-country affine term structure model. Estimation of the model shows that it can predict the correct direction of exchange rate changes one-month ahead, predicted by the UIRP, which supports the view that the risk premium can explain the forward premium anomaly. The forecasting performance of the model compares favorably to the random walk model of exchange rates, which is considered as benchmark in the literature.

The estimation of the two-country term structure model considered indicates that the risk premium effects can be indeed attributed to the local (non-common) factors between the home and foreign economies. The paper shows that these factors capture substantial and time-varying asymmetries across countries, which can obscure the expectations in the interest rates differentials about future exchange rate movements. These asymmetries can be attributed to real economy and monetary policy differences across countries. An interesting extension of this work would be to investigate separately the effects of these differences on the exchange rate risk premium and their relationship with the forward premium anomaly.

Appendix

In this appendix, we present descriptive statistics, including correlation coefficients among interest rates $R_t^{(j)}(\tau)$, for maturity intervals $\tau = \{1, 3, 6, 12\}$ months, and risk premium $\wp_t(\tau)$ across all countries j (see Tables 5, 6 and 7). Figure 1 graphically presents the estimates of $\wp_t(\tau)$.

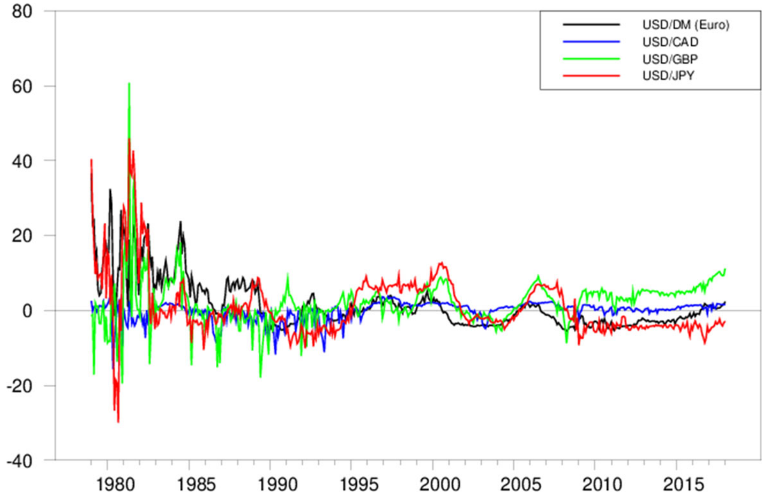

Fig. 1 Exchange rate risk premium $\wp_t(\tau)$

Table 5 Correlation coefficients among interest rates $R_t^{(j)}(\tau)$, for all countries j and maturity intervals τ

	$R_{t(1)}^{(US)}$	$R_{t(3)}^{(US)}$	$R_{t(6)}^{(US)}$	$R_{t(12)}^{(US)}$	$R_{t(1)}^{(DM)}$	$R_{t(3)}^{(DM)}$	$R_{t(6)}^{(DM)}$	$R_{t(12)}^{(DM)}$	$R_{t(1)}^{(CA)}$	$R_{t(3)}^{(CA)}$
$R_{t(1)}^{(US)}$	1.00									
$R_{t(3)}^{(US)}$	1.00	1.00								
$R_{t(6)}^{(US)}$	1.00	1.00	1.00							
$R_{t(12)}^{(US)}$	0.99	1.00	1.00	1.00						
$R_{t(1)}^{(DM)}$	0.78	0.78	0.78	0.77	1.00					
$R_{t(3)}^{(DM)}$	0.79	0.79	0.78	0.78	1.00	1.00				
$R_{t(6)}^{(DM)}$	0.80	0.80	0.79	0.79	1.00	1.00	1.00			
$R_{t(12)}^{(DM)}$	0.81	0.80	0.80	0.80	0.99	1.00	1.00	1.00		
$R_{t(1)}^{(CA)}$	0.94	0.94	0.94	0.94	0.86	0.86	0.87	0.87	1.00	
$R_{t(3)}^{(CA)}$	0.95	0.95	0.95	0.95	0.85	0.86	0.87	0.87	1.00	1.00
$R_{t(6)}^{(CA)}$	0.95	0.95	0.95	0.95	0.85	0.85	0.86	0.87	0.99	1.00
$R_{t(12)}^{(CA)}$	0.94	0.94	0.95	0.95	0.84	0.85	0.86	0.87	0.99	0.99
$R_{t(1)}^{(UK)}$	0.88	0.88	0.87	0.87	0.86	0.86	0.87	0.87	0.92	0.92
$R_{t(3)}^{(UK)}$	0.88	0.88	0.88	0.88	0.86	0.86	0.87	0.88	0.92	0.92
$R_{t(6)}^{(UK)}$	0.89	0.89	0.89	0.89	0.86	0.86	0.87	0.88	0.93	0.93
$R_{t(12)}^{(UK)}$	0.89	0.89	0.89	0.89	0.86	0.86	0.87	0.88	0.92	0.93
$R_{t(1)}^{(JP)}$	0.79	0.79	0.79	0.79	0.81	0.81	0.82	0.82	0.86	0.86
$R_{t(3)}^{(JP)}$	0.80	0.80	0.81	0.81	0.81	0.82	0.82	0.83	0.87	0.87
$R_{t(6)}^{(JP)}$	0.81	0.81	0.82	0.82	0.81	0.82	0.83	0.83	0.88	0.88
$R_{t(12)}^{(JP)}$	0.82	0.82	0.83	0.83	0.81	0.82	0.83	0.83	0.89	0.89

Table 5 (continued)

	$R_{t(6)}^{(CA)}$	$R_{t(12)}^{(CA)}$	$R_{t(1)}^{(UK)}$	$R_{t(3)}^{(UK)}$	$R_{t(6)}^{(UK)}$	$R_{t(12)}^{(UK)}$	$R_{t(1)}^{(JP)}$	$R_{t(3)}^{(JP)}$	$R_{t(6)}^{(JP)}$	$R_{t(12)}^{(JP)}$
$R_{t(1)}^{(US)}$										
$R_{t(3)}^{(US)}$										
$R_{t(6)}^{(US)}$										
$R_{t(12)}^{(US)}$										
$R_{t(1)}^{(DM)}$										
$R_{t(3)}^{(DM)}$										
$R_{t(6)}^{(DM)}$										
$R_{t(12)}^{(DM)}$										
$R_{t(1)}^{(CA)}$										
$R_{t(3)}^{(CA)}$										
$R_{t(6)}^{(CA)}$	1.00									
$R_{t(12)}^{(CA)}$	1.00	1.00								
$R_{t(1)}^{(UK)}$	0.92	0.92	1.00							
$R_{t(3)}^{(UK)}$	0.92	0.92	1.00	1.00						
$R_{t(6)}^{(UK)}$	0.93	0.93	1.00	1.00	1.00					
$R_{t(12)}^{(UK)}$	0.93	0.93	0.99	1.00	1.00	1.00				
$R_{t(1)}^{(JP)}$	0.86	0.87	0.88	0.88	0.87	0.87	1.00			
$R_{t(3)}^{(JP)}$	0.88	0.88	0.88	0.88	0.88	0.87	1.00	1.00		
$R_{t(6)}^{(JP)}$	0.89	0.89	0.89	0.88	0.88	0.88	0.99	1.00	1.00	
$R_{t(12)}^{(JP)}$	0.90	0.90	0.88	0.88	0.88	0.88	0.99	0.99	1.00	1.00

Table 6 Descriptive statistics of risk premium $\wp_t(\tau)$

	USD/DM(Euro)	USD/CAD	USD/GBP	USD/JPY
Mean	1.52	−0.16	2.06	0.92
Max	36.64	4.08	60.78	45.95
Min	−5.66	−15.69	−19.48	−30.01
Std. Dev.	6.67	2.50	6.11	8.26

Max. stands for maximum value, Min. for minimum value and Std. Dev. for standard deviation

Table 7 Correlation coefficients of risk premium $\wp_t(\tau)$ between all pairs of currencies

	USD/DM(Euro)	USD/CAD	USD/GBP	USD/JPY
USD/DM(Euro)	1.00			
USD/CAD	0.04	1.00		
USD/GBP	0.16	0.23	1.00	
USD/JPY	0.54	0.12	0.31	1.00

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