Asset price based estimates of sterling exchange rate risk premia

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Abstract

In this paper we report estimates of the effective sterling, sterling/Deutsche mark and sterling/US dollar risk premia over a monthly 1987-2001 sample, generated using a conditional factor model for the stochastic discount factor of a representative 'worldwide' investor. The model relates this stochastic discount factor to the real return on a 'worldwide' stock portfolio, with the model parameters varying with variations in the slope of the 'world' term structure of interest rates. Econometric tests indicate that this model is accepted by the data. The corresponding parameter estimates are used to compute the risk premium for the three aforementioned sterling exchange rates. A graphical analysis indicates that, in terms of magnitude, our measure of the exchange rate risk premium is mainly of importance for the sterling/Deutsche mark exchange rate. Risk-adjusted test regressions for uncovered interest rate parity *vis-à-vis* the major European currencies provide some confirmation for this.

Key words: Uncovered interest rate parity, exchange rate risk premia, conditional linear factor models, habit persistence in consumption.

JEL classification: F31, G12, G15.

Summary

Many structural exchange rate models, as well as open-economy policy models, use uncovered interest rate parity (UIP) as a building block, despite the fact that UIP is strongly violated for floating currencies. Several explanations for this phenomenon have been put forward, including the presence of time-varying risk premia. Existing empirical models of the foreign exchange rate risk premium, however, are not able to generate risk premium estimates that are sufficiently variable to explain the variability in deviations from UIP.

In this paper we attempt to estimate the risk premium for several bilateral sterling exchange rates as well as the sterling effective exchange rate index (ERI). Within intertemporal utility optimisation models, the foreign exchange rate risk premium equals the conditional covariance between the future exchange rate change and the future marginal rate of substitution of the representative investor. In conventional models the marginal rate of substitution equals a linear function of future consumption growth, which is often proxied by the future real return on a stock market portfolio. This motivates the use of an unconditional (or otherwise known as static) linear factor model for this marginal rate of substitution, with either consumption growth or the real stock return as a factor. In this paper we allow for habit persistence in the consumption behaviour of a representative international investor when we derive our measure of the marginal rate of substitution. This derivation can be used to motivate the use of a *conditional* linear factor model for the marginal rate of substitution in which it still is related to the future real return on the agent's stock portfolio, but the model parameters are time-varying and this time variation is related to movements in the slope of the term structure of interest rates. The slope of the term structure is used, as this variable has predictive power for future turning points in the real return on the stock market portfolio. Another novel feature relative to the existing literature is that our risk premium measures are related to a representative investor who operates on a global level instead of a representative investor from a particular country.

Our estimates of unconditional and conditional factor models for the global representative investor show that, in contrast to the unconditional factor model, the conditional factor model is accepted on a monthly 1987-2001 sample of nine major sterling exchange rates. We combine the resulting conditional estimates of the marginal rate of substitution for the global investor with the covariance between the relative change in a particular sterling rate and the real return on a 'world' stock portfolio to proxy the risk premium in the effective sterling exchange rate, the sterling/DM rate and the sterling/dollar rate. The resulting sterling risk premia exhibit large swings and seem especially important for the

sterling/DM rate. A graphical analysis of the estimated sterling risk premia shows, however, that the impact of the risk premium movements on sterling exchange rates seems to be limited to the short to medium run.

The foreign exchange risk premium is unobservable, and it therefore is difficult to assess whether our estimates of the foreign exchange risk premium are accurate. However, our estimates of both the marginal rate of substitution and exchange rate risk premia indicate that our approach has some empirical validity. Risk-adjusted UIP test regressions indeed indicate that relative to the major European currencies the usage of our estimated sterling exchange rate risk premia improves the parameter estimates slightly in favour of UIP, albeit not significantly so.

1 Introduction

Many structural exchange rate models as well as open-economy policy models use uncovered interest rate parity (UIP) as a building block. However, Fama (1984) has shown that conditional UIP is strongly violated for floating currencies, ie a regression of subsequent relative nominal exchange rate change on the forward discount typically produces a negative estimated coefficient.[1] Several explanations for this phenomenon has been put forward including the presence of time-varying risk premia.

Existing empirical models of the foreign exchange rate risk premium, however, are not able to generate risk premium estimates which are sufficiently variable enough to explain the variability in UIP deviations. In this paper we use a new approach to try and get more believable estimates of the time variation in the exchange rate risk premium. We do this by including in the representative consumers utility function a degree of habit persistence in consumption. This provides the rationale for using a conditional linear factor model to proxy the stochastic discount factor (SDF) as suggested by Cochrane (1996) in which the parameters vary with movements in the 'world' term structure of interest rates. Based on the parameter estimates from the SDF model for a representative international investor we construct estimates of the foreign exchange risk premium for various bilateral sterling exchange rates. Although our estimates of the exchange risk premia are in nominal terms the estimates are related to the real returns of the representative international investor. This is of importance as 'The real return on a financial asset will depend on the environment and preferences of the risk-neutral agent' (Engel (1996, page 132)), and as such one would expect that our estimates are 'true' exchange rate risk premia in the sense that they are related to the marginal rate of substitution of the representative agent.

The paper is structured as follows. In Section 2 we outline the theoretical model on which our risk premium estimates are based. The estimation method, the corresponding estimates of the risk premium and their interpretation can be found in Section 3. Concluding remarks make up Section 4.

2 A linear factor model with time-varying parameters

If there are no arbitrage opportunities across asset prices, then there exists a so-called stochastic discount factor that prices these assets. In this section we motivate what kind of

(1) Engel (1996) cites several studies conducted since Fama (1984), such as McCallum (1994), which all find that conditional UIP fails. Flood and Rose (1996), however, report results on fixed bilateral exchange rates which seem to yield more favorable results *pro* UIP. This could indicate that the UIP puzzle is mainly a characteristic of floating exchange rates.

empirical model we are going to use to proxy this stochastic discount factor, which in turn is of importance for us to be able to back out empirical estimates of the sterling risk premia. Section 2.1 provides the general consumption-based framework in which we define the stochastic discount factor and link it to individual sterling risk premia, while in Section 2.2 we show how different types of consumer preferences result in different stochastic discount factor models and how we derive our particular conditional linear factor model.

2.1 Factor price models and consumption

Suppose we have a representative international investor who can trade freely in an international financial asset k which is denominated in its own currency. The intertemporal choice problem faced by this investor can be summarised as maximising a time-separable utility function

$$\max \mathrm{E}_t \left[\sum_{j=0}^{\infty} \delta^j U(C_{t+j}) \right] \tag{1}$$

where E_t is the mathematical expectations operator, C_{t+j} is the level of consumption at time $t+j$ for our representative international investor, $U(C_{t+j})$ is the corresponding level of utility, and δ is the time discount factor (the degree of impatience of the investor). The optimal investment decision of this investor is described by the first-order conditions across N (foreign) assets which corresponds with **(1)**, ie

$$U'(C_t) = \delta \mathrm{E}_t \left[R_{k,t+1} U'(C_{t+1}) \right], \; k = 1, \ldots, N \tag{2}$$

in which $R_{k,t+1}$ is the real return on the k^{th} asset and $U'(c_t)$ is the marginal utility of consumption (which equals the first-order derivative of the utility relative to consumption). We can rewrite **(2)** in terms of the stochastic discount factor (SDF) or pricing kernel

$$M_{t+1} = \delta U'(C_{t+1}) / U'(C_t) \tag{3}$$

which in this context equals the discounted ratio of the expected future and present marginal utilities of consumption. This SDF therefore measures at which rate the representative investor is willing to substitute future consumption in period $t+1$ for present consumption in period t. Through **(3)** we rewrite **(2)** as

$$1 = \mathrm{E}_t \left[R_{k,t+1} M_{t+1} \right], \; k = 1, \ldots, N \tag{4}$$

The expected real return in terms of the international investor's price level from taking an uncovered, one-period investment in the k^{th} nominal bond denominated in a base currency can be described through **(4)** as

$$1 = E_t \left[M_{t+1} (1 + i_{k,t+1}) \frac{P_t}{P_{t+1}} \frac{S_{k,t+1}}{S_{k,t}} \right], \; k = 1, \ldots, N \tag{5}$$

In **(5)** $i_{k,t+1}$ is the one-period nominal return on a bond maturing at $t+1$ denominated in the currency of country k, P_t is the relevant price level for the representative international investor at t, and $S_{k,t}$ is the price of a unit of currency k in terms of the base currency.[(2)] Now, consider taking a covered position in the investment of the k^{th} bond. In this case, the asset pricing equation **(5)** becomes,

$$1 = \mathrm{E}_t\left[M_{t+1}(1+i_{k,t+1})\frac{P_t}{P_{t+1}}\frac{F_{k,t}}{S_{k,t}}\right],\ k=1,\ldots,N \qquad \textbf{(6)}$$

where $F_{k,t}$ is the one-period forward price of a unit of currency k in terms of the base currency.

Taking the difference between equations **(5)** and **(6)**, and making use of the fact that $i_{k,t+1}$ is known at time t, we have:

$$\mathrm{E}_t\left[M_{t+1}\frac{P_t}{P_{t+1}}\frac{S_{k,t+1}-F_{k,t}}{S_{k,t}}\right] = 0,\ k=1,\ldots,N \qquad \textbf{(7)}$$

Asset price equation **(7)** states that the conditionally expected value of risk-adjusted profits in the forward currency market at time t should be zero. Note that **(7)** also holds when the utility function has a more general function than the time-separable, constant discount form. If we assume that the variables in **(7)** are log-normally distributed then we can rewrite it as

$$\begin{aligned}\mathrm{E}_t(\Delta s_{k,t+1}) - (f_{k,t}-s_{k,t}) = -\frac{1}{2}Var_t(\Delta s_{k,t+1})&\\ + Cov_t(\Delta s_{k,t+1},\Delta p_{t+1,t}) - Cov_t(\Delta s_{k,t+1}, m_{t+1})&\end{aligned} \qquad \textbf{(8)}$$

where lower-case letters denote logarithms of the previously defined upper-case letters. In terms of Engel (1996) the 'true' exchange rate risk premium, ie corrected for the Jensen's inequality term and the inflation risk premium $Cov_t(\Delta s_{k,t+1},\Delta p_{t+1,t})$, equals

$$rp_t = -Cov_t(\Delta s_{k,t+1}, m_{t+1}) \qquad \textbf{(9)}$$

The risk premium definition **(9)** is 'true' in the sense that it is a function of the marginal utility of consumption of the representative investor and thus results from the optimising behaviour of this investor. The representative investor has an urge to smooth his consumption through time in order to minimise his consumption volatility. He does not care about the volatility of the returns of his investments, which is in **(8)** proxied by the sum of the Jensen's inequality term and the inflation risk premium, as long as his consumption is stabilised. The representative investor has therefore a preference to buy bonds of those countries that move *anti-cyclically* with its consumption as it helps to smooth the expected consumption path. Bonds of countries that move *pro-cyclically* with the representative investor's consumption will only be attractive to the investor when its return contains a risk premium as holding the asset would cause a more volatile

(2) A rise in $S_{k,t}$ represents a depreciation of the domestic currency.

consumption path. Hence, the 'true' risk premium **(9)** depends negatively on the covariance between the exchange rate return and the SDF as the latter essentially measures the expectation of the intertemporal consumption growth path, see **(3)**.

To get an empirical estimate of the risk premium through **(9)** one has to specify a model for the log SDF m_{t+1}. One way of modelling m_{t+1} is to use a linear factor model, ie

$$m_{t+1} = a + b' \boldsymbol{f}_{t+1} \quad \textbf{(10)}$$

where $\boldsymbol{f}_{t+1}$ is a $k \times 1$ vector of factors that contain information about the marginal utility of consumption, as the SDF is related to the present and future levels of this marginal utility (see **(3)**). This specification in turn poses the question of what should one use for factors $\boldsymbol{f}_{t+1}$? In order to be able to answer that question, one has to take a stance on how the preferences of our representative international investor are specified, and we deal with this in the next subsection.

2.2 Deriving the appropriate factors

A standard assumption in the literature is to assume that the utility function used in **(1)** is of the constant rate of risk aversion (CRRA) form, ie

$$U(C_t) = \frac{1}{1-\gamma} C_t^{1-\gamma} \quad \textbf{(11)}$$

and therefore the corresponding SDF equals

$$M_{t+1} = \delta \left(\frac{C_t}{C_{t+1}} \right)^{\gamma} \quad \textbf{(12)}$$

where γ is the relative risk aversion parameter. Hence, higher future consumption growth decreases the preparedness of the investor to postpone a fraction of its present consumption as the future consumption prospects have become more positive (obviously the *vice versa* applies when future consumption growth decreases).

Equation **(12)** implies a constant linear factor model in which m_{t+1} depends on relative real consumption growth

$$m_{t+1} = a + b \Delta c_{t+1} \quad \textbf{(13)}$$

where $a = \ln(\delta)$ and $b = -\gamma$. We can now express the exchange rate risk premium as a function of relative real consumption growth, ie

$$rp_t = \gamma Cov_t(\Delta s_{k,t+1}, \Delta c_{t+1}) \quad \textbf{(14)}$$

Thus, if the future covariance in **(14)** is positive then the relative return on the domestic bond moves procyclically with consumption growth and investing in domestic bonds is not an appropriate tool to hedge against consumption volatility. The investor therefore has to receive a higher risk premium in order to induce him to hold the investment. Many empirical studies have analysed risk premium model **(14)**. However, for plausible values of the risk-averse parameter γ estimates of the risk premium **(14)** are not variable enough in order to be able to explain the observed deviations from UIP.(3) And when the risk-aversion parameter γ is estimated on an unrestricted basis it attains unrealistically high values. Mark (1985), for example, finds estimates of $\gamma > 40$ whereas Engel (1996) reports several estimates for γ in excess of 100. Based on these results it thus becomes clear that risk premia estimates utilising standard utility function specifications will not go far in explaining the observed deviations from UIP.

One way of introducing more variability and time variation in our exchange rate risk premium is to introduce time-varying parameters in **(13)** in order to have a more volatile SDF

$$m_{t+1} = a_t + b_t \Delta c_{t+1} \tag{15}$$

This implicates a move from an *unconditional* factor model to a *conditional* factor model, ie the SDF is now fully conditional on the investor's information set in period t. It is worthwhile to point out that **(13)** always implies **(15)** but the reverse need not be the case, see Cochrane (2001, Section 8.3). A theoretical motivation for using such a conditional factor model for the price kernel would be based on introducing habit persistence in consumer behaviour that would result in non-time separable utility. In this vein we follow Campbell and Cochrane (2000) and assume that **(1)** is maximised based on

$$U(C_t - X_t) = \frac{(C_t - X_t)^{1-\gamma} - 1}{1-\gamma} \tag{16}$$

where X_t is the level of habit persistence. As in Campbell and Cochrane (2000) we assume that habit responds slowly to consumption. The surplus consumption ratio, ie $V_t = \frac{C_t - X_t}{C_t}$, is therefore assumed to comply in its logarithmic form with the following AR(1) model,

$$v_{t+1} = (1-\phi)\bar{v} + \phi v_t + \lambda(v_t)(\Delta c_{t+1} - g), \quad 0 < \phi < 1 \tag{17}$$

where $\bar{v}$ and g are the steady-state levels of the surplus ratio and consumption growth respectively whereas $\lambda(v_t)$ is an unobserved sensitivity function in the spirit of Campbell and Cochrane (2000).

The corresponding SDF can now be written as

$$M_{t+1} = \delta \left(\frac{V_{t+1}}{V_t} \frac{C_{t+1}}{C_t} \right)^{-\gamma} \tag{18}$$

and in combination with **(17)** this in turn implies the following conditional factor model

$$m_{t+1} = a + b(v_t) + d(v_t)(\Delta c_{t+1}) \tag{19}$$

(3) Plausible values of γ are in the range of 2-5, see Engel (1996).

The log surplus ratio v_t is in practice unobservable and one has to proxy its behaviour in **(19)** by a scaling variable z_t, see also Lettau and Ludvigson (2001), ie

$$m_{t+1} = (a_0 + a_1 z_t) + (b_0 + b_1 z_t)\Delta c_{t+1} \quad \textbf{(20)}$$

which is equal to a scaled conditional linear factor model in the spirit of Cochrane (1996). In order to show the difference in risk premium behaviour *vis-à-vis* the case without habit persistence we can write the exchange rate risk premium as the following function of real consumption growth:

$$rp_t = \gamma Cov_t(\Delta s_{k,t+1}, \Delta c_{t+1}) + \gamma Cov_t(\Delta s_{k,t+1}, \Delta v_{t+1}) \quad \textbf{(21)}$$

Provided that the log surplus ratio v_t is variable enough the extra covariance term in **(21)** could generate an appropriately volatile exchange rate risk premium.

However, the use of a conditional factor model based on **(19)** still requires the use of real consumption growth as a factor, and consumption data might well be unavailable at higher data frequencies such as on a monthly basis. Moreover, Campbell (1993) argues that consumption of asset market participants may be poorly approximated by aggregate consumption, and using an Epstein-Zin utility function as well as log-linearising the investor's budget restriction he substitutes in his asset pricing model real consumption growth with an aggregate real stock market return. In more general terms, one can propose, as in the intertemporal capital asset pricing model (ICAPM) of Merton (1973), that optimal consumption is a function of one or more 'state variables', which are variables that reflect the degree in which the investor can maximise its consumption stream, ie $C_t = g(Z_t^C)$. The SDF will therefore become a function of these 'state variables' Z_t^C,

$$M_{t+} = \delta U'(g(Z_{t+1}^C))/U'(g(Z_t^C)) \quad \textbf{(22)}$$

One obvious candidate 'state variable' for consumption is wealth, in particular if one follows through the line of reasoning underlining the permanent income hypothesis of consumer behaviour. Hence, one could replace in **(13)** and **(15)** Δc_{t+1} as a factor by the real return on the investor's wealth portfolio r_{t+1}^W.

Using r_{t+1}^W as a factor instead of real consumption growth, however, does not solve the issue of data availability at higher frequencies. As pointed out by, for example, Lettau and Ludvigson (2001), the real return on wealth is a composite of the real return on the representative market portfolio as well as the real return on human capital. Data on the latter are for most countries not readily available at monthly or higher frequencies. We therefore assume that r_{t+1}^W can be reasonably proxied by the real return on a representative market portfolio r_{t+1}^m, and therefore our conditional linear factor model equals:

$$m_{t+1} = (a_0 + a_1 z_t) + (b_0 + b_1 z_t) r_{t+1}^m \quad \textbf{(23)}$$

As z_t should have explanatory power for future real returns on the market portfolio we shall use in the estimates in Section 3 the slope of the term structure of interest rates as the z_t variable.[(4)] Obviously, the strength of this relationship depends on how accurate the relationship is between real consumption growth and r_t^W as well as that of the relationship between r_t^W and r_t^m. Depending on the strengths of these relationships our factor $f_{t+1} = r_t^m$ could potentially be contaminated by measurement error, and this in turn could affect the power of **(23)** to proxy fluctuations in the log SDF m_{t+1} and thus the exchange rate risk premium; see Wickens and Smith (2001) for more details. Also, as pointed out in Cochrane (2001, page 170), measures of r_t^m are usually much more volatile than real consumption growth, and it is not likely that even an accurately measured consumption stream for investors is as volatile as r_t^m. Despite these caveats, we think that specification **(23)** could be a flexible and useful tool to proxy fluctuations in the SDF and thus, through **(9)**, the exchange rate risk premia, which can be done without having to properly identify structural parameters such as γ in **(14)** or **(21)**.

3 Estimation

In this section we use specification **(23)** from Section 2 to estimate risk premia for several major sterling exchange rates. Section 3.1 provides a description of both how we have specified **(23)** and the utilised estimation methodology. Next, Section 3.2 reports and interprets the estimation results.

3.1 Method and data

We derived in Section 2 that the intertemporal choice faced by the representative international investor yielded a fundamental asset pricing equation as represented by **(7)**. Traditionally, studies aimed at producing consumption-based estimates of the exchange rate risk premium utilise the SDF for the investor of a particular country. However, in this paper we follow Ayuso and Restoy (1996) and focus in the estimation on the SDF from **(29)** for a representative 'world' investor. According to Ayuso and Restoy (1996) '...it seems preferable to analyze international asset prices on the basis of their contribution to the overall risk and return of an international well diversified portfolio.' (Ayuso and Restoy (1996, page 373)). Also, focusing on a 'world' investor with preferences related to a 'world' consumption level and endowed with 'world' wealth could at least partly

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(4) Conditioning variable z_t should have predictive power for future turning points in r_t^m. Both the literature as well as a brief analysis in Section 3.1 shows that this is the case for the slope of the term structure. Other variables could potentially have this property too, such as credit spreads in bond future markets and the return on human wealth, but data availability issues limit us to using the term structure slope.

circumvent the aggregation problem due to long-lasting PPP deviations when we would focus on investors of a particular nationality. Hence, we use a world aggregate price index (composed of a weighted average of domestic aggregate consumer price indices), P_t^w, and the real return on a world stock portfolio, r_t^w, to estimate the SDF for this 'world investor'.

Scaling variable z_t should reflect the time variation in the future excess returns on the representative investor's portfolio and thus z_t should have explanatory power for the real returns utilised in the SDF equation. Several studies have indicated that the slope of the term structure of interest rates has predictive power for real stock returns in the medium to long run, see eg Hodrick (1992) and Fama and French (1993). Hence, we have chosen the term structure as our conditioning variable, ie

$$z_t = i_t^{lr} - i_t^{sr} \tag{24}$$

where i_t^{lr} (i_t^{sr}) is the long-term (short-term) nominal interest rate for the representative investor at time t.

In order to assess whether indeed the world term structure of interest rates z_t has explanatory power for the real returns on the world portfolio at the h-month horizon $r_{t+h,t}^w$, we estimate the regression

$$r_{t+h,t}^w = \alpha + \beta z_t + \epsilon_{t+h,t}, \quad h = 1, 3, 6, 12, 24, 36 \tag{25}$$

where $\hat{\epsilon}_{t+h,t}$ is the regression residual. For $h > 1$ the residuals $\hat{\epsilon}_{t+h,t}$ become serially correlated as the left-hand side variable in **(25)** contains an overlap (the horizon becomes larger than a month and the data are sampled on a monthly frequency). We compute the appropriate standard errors by pre-whitening the residuals with an AR(1) model and applying the procedures from Newey and West (1987, 1994). The estimation results are reported in Table A. These results show, in compliance with previous studies, that indeed the term structure has explanatory power for the real stock returns as the coefficient corresponding with z_t becomes significant from the two-year horizon onwards and the R^2 becomes substantially larger.

One caveat of the overlapping data regression model **(25)** is that for large h finite sample inference on parameter β becomes biased due to a summation of a large number of autocovariances, as shown by Hodrick (1992) through Monte Carlo experiments. An alternative way of doing inference on β in **(25)** is based on the observation that the covariance term corresponding with β in **(25)**

$$Cov\left[\sum_{j=1}^{h} r_{t+j,t+j-1}^w;\ z_t\right] \tag{26}$$

Table A: Explanatory power of the 'world' term structure for the real returns on the 'world' portfolio, 1987:12-2001:4

h	β	$t_{\bar{\beta}}$	R^2
1	0.01 (0.01)	–	0.9%
3	0.02 (0.02)	1.34	3.0%
6	0.03 (0.04)	1.53	7.3%
12	0.07 (0.08)	1.65	16.9%
24	0.15* (0.06)	1.70	41.1%
36	0.21** (0.07)	2.14*	50.2%

Note: Columns β and R^2 result from regression **(25)** at different horizons of h months. The values within parentheses are the standard errors for β (see **(25)**) corrected for serial correlation through the procedures from Newey and West (1987, 1994) combined with pre-whitening. The column $t_{\bar{\beta}}$ reports the t-statistic for the null hypothesis that $\bar{\beta} = 0$ in **(28)** based on White (1980) heteroskedasticity-robust standard errors. An * (**) indicates that β in **(25)** or $\bar{\beta}$ in **(28)** is significantly different from 0 at the 5% (1%) significance level.

is numerically identical to

$$Cov\left[r^{w}_{t+1,t};\ \sum_{j=0}^{h-1} z_{t-j}\right] \quad \textbf{(27)}$$

where $r^{w}_{t+j,t+j-1}$ is the real return on the 'world' stock portfolio between month $t+j$ and $t+j-1$. Covariance **(27)** is the numerator of the OLS estimator of $\bar{\beta}$ in

$$r^{w}_{t+1,t} = \bar{\alpha} + \bar{\beta}\left[\sum_{j=0}^{h-1} z_{t-j}\right] + \bar{\epsilon}_{t+1,t}, \quad h = 1,3,6,12,24,36 \quad \textbf{(28)}$$

Due to the numerical equality between **(26)** and **(27)** specifications **(25)** and **(28)** are asymptotically equivalent. This approach is advocated by Jegadeesh (1991) and Cochrane (1991) to investigate the long-run link between stock returns and predictive variables, and by Groen (1999) to test the long-run link between exchange rates and monetary fundamentals. Hodrick (1992) shows that in finite samples inference based on **(28)** is better behaved than inference based on **(25)**.

We report in Table A for $h > 1$ the t-statistics for $H_0: \bar{\beta} = 0$ from **(28)** as a robustness check for our previous inference on β from **(25)**.(5) At each value of h regression **(28)** is well behaved, ie Lagrange-Multiplier tests for residual serial correlation (at lag orders 1, 3, 6 and 12) and ARCH-effects (lag order 1) could not detect the significant presence of these phenomena. Heteroskedasticity tests based on White (1980), however, were able to detect significant heteroskedasticity at all horizons. We therefore have based the reported t-values on White (1980) heteroskedasticity-robust standard errors. Based on these t-values we have to conclude that up to the 24-month horizon there is no statistically significant link between the real 'world' stock return and the 'world' term structure slope, but at the 36-month horizon the t-statistic based on **(28)** confirms the significance of the link. Hence, in the medium to long run our 'world' term structure slope has explanatory power for turning points in the real 'world' stock return.

As **(7)** represents a set of moment conditions, the generalised method of moments (GMM) approach seems to be a natural way of estimating the SDF. The moment conditions set out under **(7)** are expectations conditional on all available information at time t, ie I_t. As we assume rational expectations we have for any set of instruments $Q_t \subset I_t$ that

$$E\left[Q_t M_{t+1} \rho_{t+1} \pi_{k,t+1}\right] = 0,\ k = 1, \ldots, N \quad \textbf{(29)}$$

holds, where

$$M_{t+1} = \exp\left[(a_0 + a_1 z_t) + (b_0 + b_1 z_t) r^{w}_{t+1}\right]$$

whereas

$$\rho_{t+1} = \frac{P^{w}_{t+1}}{P^{w}_{t}} \quad \text{and} \quad \pi_{k,t+1} = \frac{S_{k,t+1} - F_{k,t}}{S_{k,t}}$$

(5) At $h = 1$ **(25)** and **(28)** involve exactly the same regression.

for $k = 1, \ldots, N$. The moment conditions in **(29)** imply that on average profits in the forward currency markets are zero which guarantees the absence of arbitrage opportunities. We therefore use GMM based on the moment conditions set out under **(29)** to estimate the parameters in the SDF equation and use the corresponding GMM test for overidentification restrictions in order to assess whether the estimated SDF is rejected by the data or not. In our GMM estimations we select lags of the inflation rate, the forward market profits, the return of wealth and the conditioning variable as our set of instruments, ie

$$Q_t' = \left(r_t^w, \ldots, r_{t-l}^w, z_t, \ldots, z_{t-l}, \rho_t, \ldots, \rho_{t-l}, \pi_{1,t}, \ldots, \pi_{1,t-l}, \pi_{N,t}, \ldots, \pi_{N,t-l}\right) \qquad \textbf{(30)}$$

The instruments in **(30)** are those variables which are likely to be of importance to market participants in solving their forecasting problem with regard to the return on their investments. As can be asserted from **(29)** and **(30)** we combine in our estimation of the parameters in the SDF equation information across N currencies as currency changes are highly contemporaneously correlated across countries, see also Mark (1985).

After we have estimated the parameters in the SDF equation through **(29)**, we use in a final step these parameter estimates to construct our measures of the exchange rate risk premium. The general form of the exchange rate risk premium **(9)**, however, measures the conditional covariance between an observed variable $\Delta s_{k,t+1}$ and an unobserved variable m_{t+1}. In **(29)**, on the other hand, the SDF is a function of the real return on the world stock portfolio and therefore we proxy the exchange rate risk premium **(9)** through a function which depends on the covariance between $\Delta s_{k,t+1}$ and r_{t+1}^w, ie

$$rp_{k,t} = -(\hat{b}_0 + \hat{b}_1 z_t) Cov(\Delta s_{k,t+1}, r_{t+1}^w), \quad k = 1, \ldots, N \qquad \textbf{(31)}$$

where $\hat{b}_0$ and $\hat{b}_1$ are the parameter estimates resulting from applying GMM on **(29)**. Intuitively **(31)** states that the time variation in the future returns r_{t+1}^w is proxied by the slope of the term structure z_t, and thus this determines the time variation in the exchange rate risk premium.

Based on the line of reasoning in Section 2 and in this subsection we can summarise our approach of estimating exchange rate risk premia, which in general is of the form **(9)**, as follows. First, based on an intertemporal utility maximising framework we can motivate that the SDF of the representative agent conditionally depends on the real return on its wealth portfolio and the time variation of the parameters in this function depends on the slope of the term structure of interest rates. Next, by imposing the absence of arbitrage opportunities in forward currency markets we estimate this SDF function for a 'world' investor utilising the real return on a 'world' portfolio, a 'world' aggregate price level and the slope of the 'world' term structure through applying GMM on **(29)**. Finally, the resulting parameter estimates and the observed covariance between the exchange rate return and the real return on the 'world' stock portfolio are used to proxy the unobserved covariance between the exchange rate return and the SDF through **(31)**.

3.2 Estimation results

We use monthly data from December 1987 up to April 2001. The forward and spot exchange rates used are the bilateral sterling rates *vis-à-vis* the Australian dollar, Belgian franc, the Canadian dollar, the Dutch guilder, the French franc, the German mark, the Italian lira, the Japanese yen and the US dollar.[6] Thus to estimate the SDF there are nine primary moment conditions, ie one for each currency bilateral, and we therefore have $N = 9$ in **(29)**. In **(29)** we also use proxies for the real return on the world portfolio, the world aggregate price level and the world term structure based on weighted averages of these variables across the major industrialised countries. We have chosen to focus on the industrialised countries as the capital flows among these countries are more or less unhindered.[7]

The most optimal GMM estimates of the conditional factor model for the log SDF, estimated through **(29)**, are reported in the second column of Table B. These are the most optimal in the sense that parameters a_0 and b_0 were insignificant in the preliminary estimates, and thus we estimated **(29)** under the restriction $a_0 = b_0 = 0$. These estimates are based on the instrument variables set as defined in **(30)** with 7 lags. The number of lags in the instruments were selected through sequentially estimating **(29)** with an increasing number of lags (starting off with lags=1). The optimal number of lags is set equal to the number of lags at which the test for overidentification restrictions accepted the specification. From the second column of Table B one can observe that all the parameters in the SDF equation are time varying and the corresponding test of overidentification restrictions is not able to reject this particular SDF model. The corresponding parameter estimates are significantly negative which is in compliance with the negative relationship between the SDF and its determinants under habit persistence, see **(18)**.

To provide a benchmark for our parameter estimates from the conditional factor model we also estimate an unconditional or static version of the linear factor model for the SDF, ie

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(6) The Belgian franc, the Dutch guilder, the French franc, the German mark and the Italian lira effectively disappeared as independent currencies with the start of the European Monetary Union. However, we maintain them as separate currencies in our estimated system. Already before 1999 there was a high degree of comovement across these currencies as member countries of the European Exchange Rate Mechanism (ERM) essentially linked their monetary policy to that of Germany. In the estimation the GMM weighting matrix (ie the long-run covariance matrix across currencies) would pick up this comovement due to the ERM and the EMU.

(7) A more elaborate description of the data can be found in the appendix.

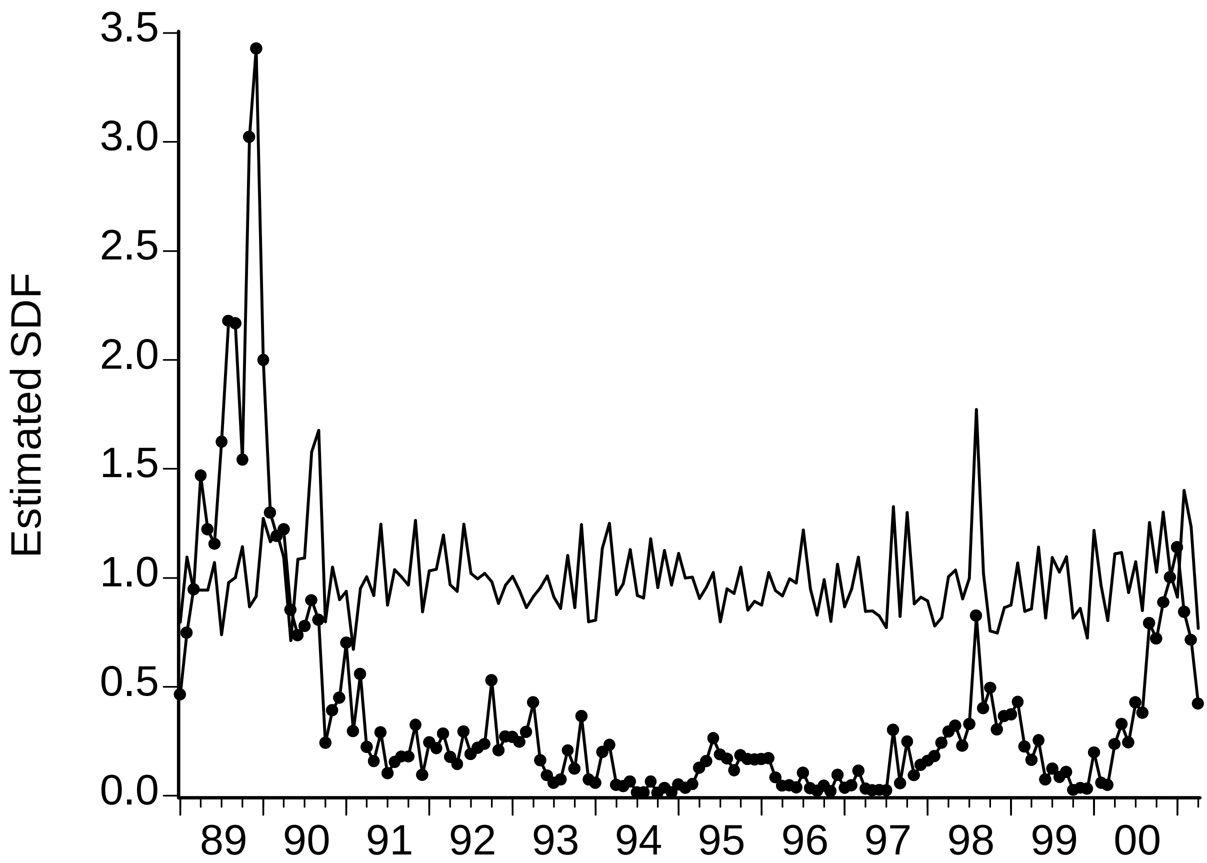
Table B: GMM estimates of the log SDF using both an unconditional and a conditional factor model, 1987:12-2001:4

	No habit persistence	Habit persistence	Habit persistence (1987:12-1995:06)
a_0	—	—	—
a_1	N.A.	-1.53 (0.06)	-1.32 (0.07)
b_0	-0.05 (0.01)	—	—
b_1	N.A.	-0.10 (0.01)	-0.09 (0.02)
OR	531 [0.00]	788.45 [0.18]	569.32 [0.16]
Lags	4	7	5

Note: The table contains estimation results for **(29)**. A '—' indicates that the corresponding variable has been dropped due to the insignificance of the parameter. The 'N.A.' stands for 'not applicable'. Values within parentheses are the standard errors corrected for serial correlation through the procedures from Newey and West (1987, 1994), whereas the values in squared brackets are the p-values for the overidentifying restrictions test 'OR' under the null hypothesis that the specification is valid. 'Lags' are the number of lags for the instrumental variables.

we estimate **(29)** with GMM under the restriction $a_1 = b_1 = 0$. The corresponding parameter estimates are reported in the first column of Table B, but for none of the attempted lag specifications for the instrumental variables, $l = 1, \ldots, 7$, was the overidentifying restrictions test able to accept this SDF model without habit persistence on the data. The results reported in the first column of Table B are those for the estimated unconditional factor model with the correct (negative) sign of the parameters, see **(12)**. Hence, the results from the first 2 columns of Table B indicate that a conditional factor model of the SDF, with the real return on the 'world' stock index as a factor and the slope of the 'world' term structure as a scaling variable, seems to provide a better description of investment behaviour on the forward currency markets than the unconditional model without a scaling variable.

Chart 1: Estimates of the stochastic discount factor for the representative international investor based on an unconditional and a conditional factor model



The solid line is the estimate of the stochastic discount factor based on the unconditional factor model and the dotted line is the one based on the conditional factor model.

To illustrate the difference between the two SDF models we have plotted in Chart 1 fitted values of the *ex-ante* SDF using the parameter estimates from the first two columns of Table B. As can be discerned from this chart the SDF exhibits a much more clustered pattern when it is based on the conditional factor representation **(23)** of the log SDF when an unconditional (ie time invariant) representation is used. In other words, with the conditional representation the high values are clustered together and the same applies to the low values, whereas in the case of the unconditional representation the SDF almost behaves like a white noise process around the value 1. Thus, an exchange rate risk premium based on the conditional factor representation **(23)** for the SDF could potentially be more variable than one using an unconditional factor SDF model.

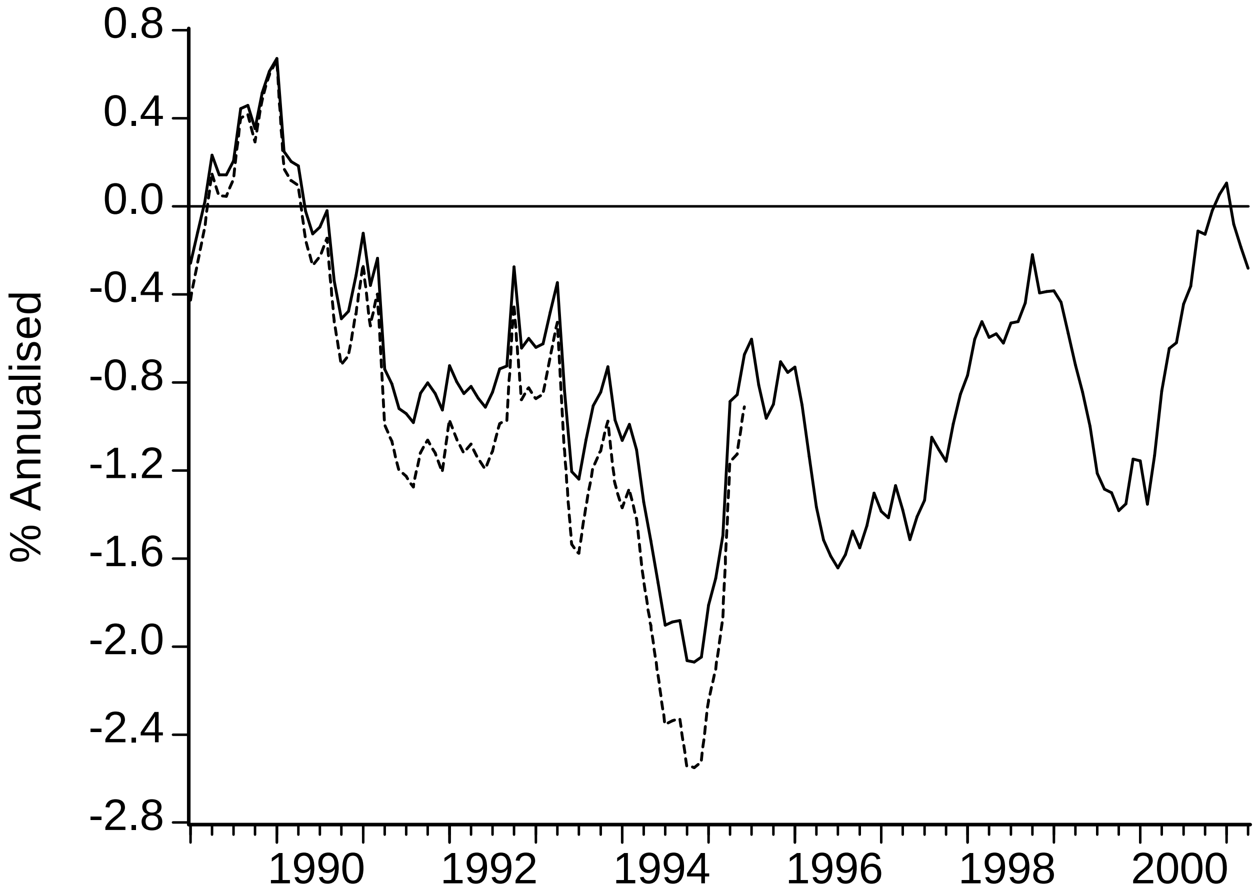
The parameter estimates from the second column of Table B can now be used through **(31)** to construct measures of the sterling risk premia. In the case of the United Kingdom the effective sterling exchange rate is mainly driven by the bilateral sterling rates *vis-à-vis* the euro area and the United States. Hence, we also focus on the risk premia estimates for the German mark (as a proxy for the euro) and the US dollar relative to the United Kingdom (ie we express the exchange rates as the number of foreign currency per unit of pound sterling so that an increase indicates a depreciation of the foreign currency and *vice versa*).

We have in Chart 2 extracted a proxy of the effective sterling risk premium from the nine bilateral sterling rates for the 1989:01-2001:04 sample as a benchmark for our bilateral estimates.[8] As movements in the effective sterling exchange rate are mainly due to movements in the sterling/euro rate, we have plotted the sterling/DM risk premium in Chart 3 as a proxy for the sterling/euro risk premium. From the chart it becomes apparent that in general the effective sterling risk premium mimics the dynamic behaviour of the sterling/DM risk premium, albeit that the effective sterling risk premium is in absolute value smaller than the sterling/DM one. A graphical description of the sterling/dollar risk premium is reported in Chart 4. Note that the dynamics in our measure of the exchange rate risk premium are driven by the slope of the 'world' term structure of interest rates. The value and the sign of the estimated risk premia for different sterling rates are therefore determined by the level of the covariance of the exchange rate return and the real return on the 'world' stock portfolio, see **(31)**.

As the covariance of the exchange rate return and the real return on the 'world' stock portfolio has been positive for the sterling/dollar rate, the sterling/dollar risk premium in

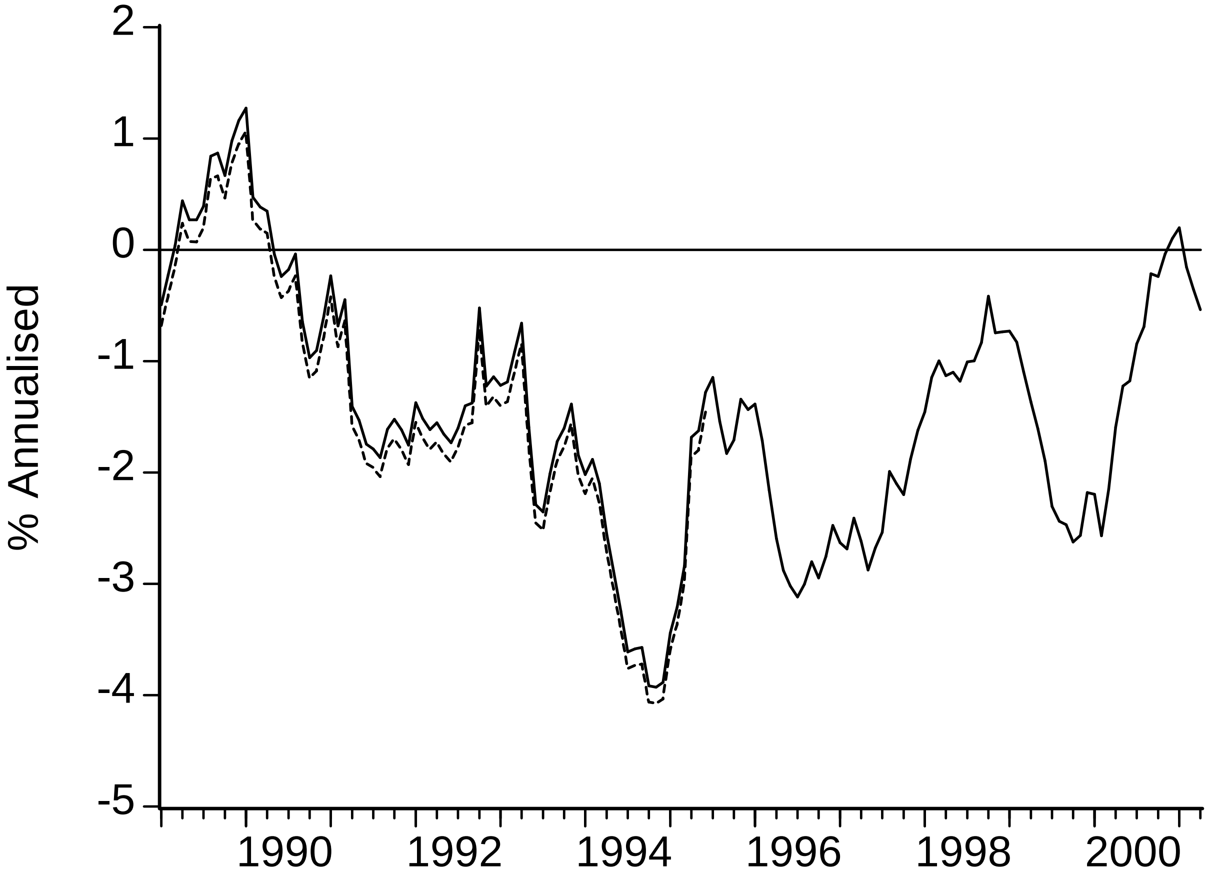
[8] We focus in the figures on the 1989:01-2001:04 sample as we have a lag order of 7 in the set of instrumental variables for the GMM estimation of **(29)** and the risk premium estimates for the pre-1989 period are therefore unreliable. The appendix describes how we have constructed the measure of the effective sterling risk premium.

Chart 2: Estimates of the effective sterling risk premium for the periods 1989:01-2001:04 and 1989:01-1995:06



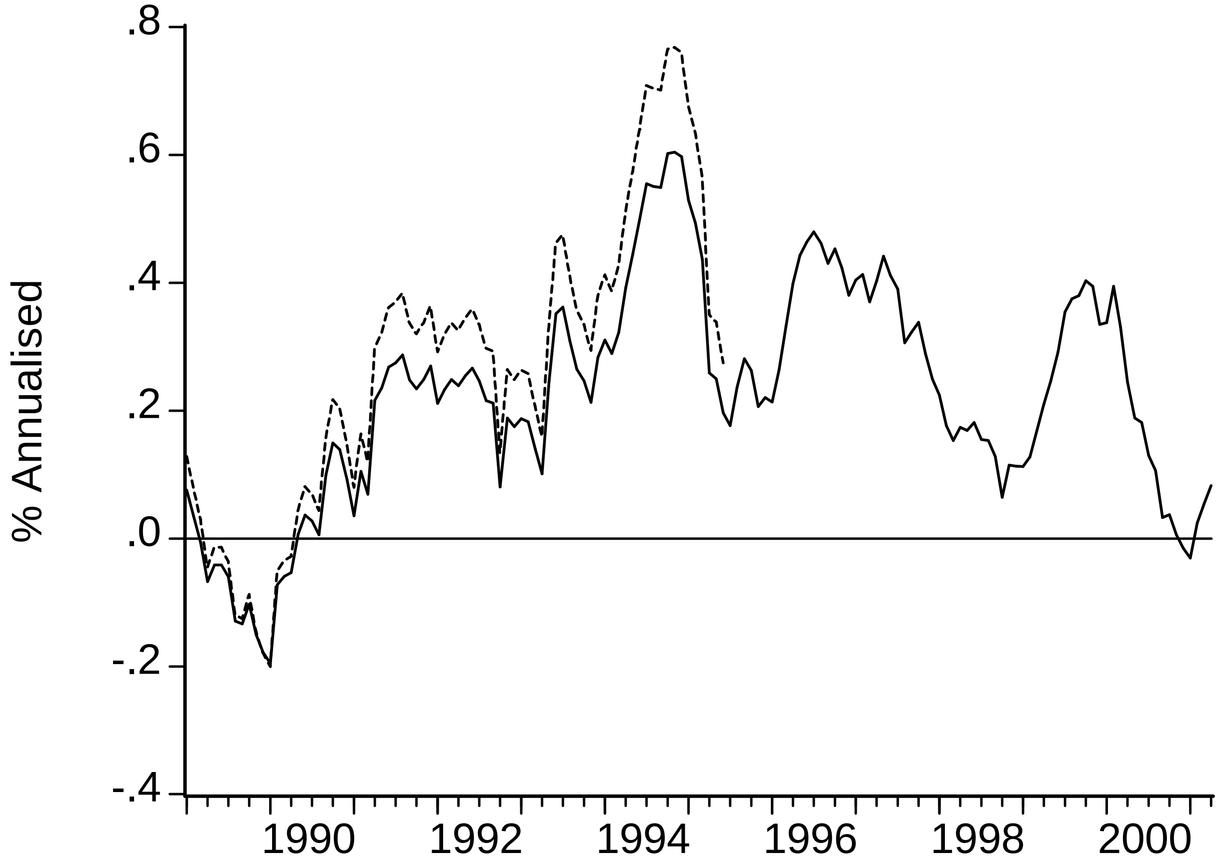
The solid line is the estimated risk premium over the whole sample, whereas the dashed line is the one over the subsample. A **minus** sign indicates that sterling is **preferred**.

Chart 3: Estimates of the sterling/Deutsche mark risk premium for the periods 1989:01-2001:04 and 1989:01-1995:06



The solid line is the estimated risk premium over the whole sample, whereas the dashed line is the one over the subsample. A **minus** sign indicates that sterling is **preferred**.

Chart 4: Estimates of the sterling/US dollar risk premium for the periods 1989:01-2001:04 and 1989:01-1995:06



The solid line is the estimated risk premium over the whole sample, whereas the dashed line is the one over the subsample. A **minus** sign indicates that sterling is **preferred**.

Chart 4 moves in the opposite direction *vis-à-vis* the sterling/DM risk premium. This most likely reflects the view of investors that sterling moves in line with the euro (and DM before 1999) albeit in a less than proportional way given the pattern of the sterling/euro risk premium in Chart 3. However, the sterling/dollar risk premium is very small in magnitude and as a consequence any large-scale movements in the sterling/dollar rate are most likely due to other factors than the risk premium. One can also observe from Charts 2-4 that the estimated risk premia have the same sign for a prolonged period of time, and in fact only change sign at near the start and near the end of the sample period respectively. There are several possible reasons for this phenomenon. First, as stated above, the dynamics of the SDF, and hence the exchange rate risk premia, are dominated by the scale variable z_t. This scale variable is the slope of the term structure of 'world' interest rates, so one can expect z_t to behave fairly persistently. Also, given the short data span, ie 1989-2001, it might well be that our estimates of the different covariances of the bilateral exchange rate returns with the real return on the 'world' stock portfolio are contaminated by finite sample bias. As these estimates determine the level of our estimated exchange rate risk premia, it might well be the case that this finite sample bias results in an estimated level that is too far away from the zero line, and hence the number of switches in sign is simply too low. Finally, there is the possibility that the unconditional covariance between the bilateral exchange rate return and the real return on the 'world' stock portfolio in **(31)** has not been constant throughout the period, and we shall explore this possibility in a bit more detail.

Our empirical proxy of the sterling risk premium **(31)** is a function of the unconditional covariance between the return on the sterling exchange rate of currency *k* and the real return on the 'world' stock portfolio. Potentially this covariance could have changed over the sample and thus affect the robustness of our sterling risk premium estimates. To check the robustness of our risk premium we have repeated our two-step procedure on the first half of the sample, ie 1987:12-1995:06.[9] That is, we first estimate with GMM the parameters in the SDF model under habit persistence through **(29)** on the 1987:12-1995:06 subsample. Next, we combine these parameter estimates with the estimated $Cov(\Delta s_{k,t+1}, r^w_{t+1})$ over the 1987:12-1995:06 period to proxy the subsample exchange rate risk premium with **(31)**. The subsample SDF parameter estimates can be found in the last column of Table B. Like in the full sample case, see the second column of the table, the overidentifying restrictions test accepts the model on the data although the lag order in the instrumental variables set is slightly lower than for the full sample. The parameter estimates for the 1987:12-1995:06 subsample are comparable to the full sample estimates in the second column, and any differences between the subsample risk premium estimates and the full sample-based estimates would be entirely due to differences in the estimated $Cov(\Delta s_{k,t+1}, r^w_{t+1})$.

The dashed lines in Charts 2, 3 and 4 are the empirical proxies of the subsample sterling risk premia computed through **(31)** in which we have combined the parameter estimates from the last column in Table B with the 1989:01-1995:06 estimates of $Cov(\Delta s_{k,t+1}, r^w_{t+1})$ for the effective sterling exchange rate, sterling/DM and sterling/dollar rates.[10] In case of the sterling/DM risk premium the subsample-based estimated risk premium can hardly be discerned from the full sample-based estimated risk premium. In fact, only in the case of

[9] As the 1987:12-2001:04 sample is already fairly short, the 1987:12-1995:06 sample is the only usable subsample in terms of the number of observations.
[10] Again, we focus on the sample from 1989 onwards, as we have a lag order of 5 in the GMM estimation of the subsample SDF.

the sterling/dollar risk premium is one able to observe a visible difference between the subsample-based and full sample-based estimations. This is especially the case during 1994, which is most likely caused by the worldwide turbulence in bond markets that occurred throughout that year. However, the difference in magnitude between the subsample-based and full sample-based estimates of the sterling/dollar risk premium is not very large. Hence, we can conclude that our estimated sterling risk premia have been reasonably stable over the sample.

In order to show any possible comovements between the exchange rate and the risk premium, we have constructed in Charts 5, 6 and 7 charts of the accumulated changes in the nominal exchange rate and the inverse of the corresponding risk premium for the effective sterling, sterling/Deutsche mark and sterling/US rates.[11] Based on this specification we would expect a positive relationship between the two series: a decrease in the sterling risk premium results in an increased preference for sterling which should result in a sterling appreciation. Roughly speaking this positive relationship seems to be present in the data, as both series exhibit a corresponding V-shaped pattern over the sample, especially for the effective sterling and the sterling/Deutsche mark relationships. There are, however, episodes that the exchange rate moved in the opposite direction than implied by the corresponding risk premium movements, in particular at the start, during 1995 and the end of the sample, and these diverging movements are particularly stark for the sterling/US rate. Next to that, we also observe that the magnitude of the changes in the exchange rate is larger than that of the risk premium, in particular in the case of the US dollar/sterling exchange rate.

The swings in the effective sterling risk premium in Chart 2 can be linked to several events using Chart 5. The 1992 ERM crisis induced a sharp decrease in the degree of preference for sterling which died out quite quickly in the aftermath of the crisis. The corresponding depreciation of sterling, however, surpasses the increase in riskiness. The Barings crisis during 1995 resulted in sterling becoming less preferred. This is not reflected in the exchange rate movement and in fact it moved in the opposite direction. The sharp sterling appreciation in 1996-97 is accompanied by a sharp increase in the preference of sterling. During 1998, however, this increased preference for sterling was reversed, possibly due to a positive impact of the approaching start of the European Monetary Union (EMU) on the currencies of EMU member states, although sterling continued appreciating. In 2000 the risk premium seem to have caught up with the exchange rate movements.

[11] The nominal exchange rate is here defined as the number of foreign currencies per pound sterling and therefore a rise indicates an appreciation of sterling.

Chart 5: Accumulated changes in the effective sterling exchange rate and the inverse of the corresponding risk premium, 1989:01-2001:04

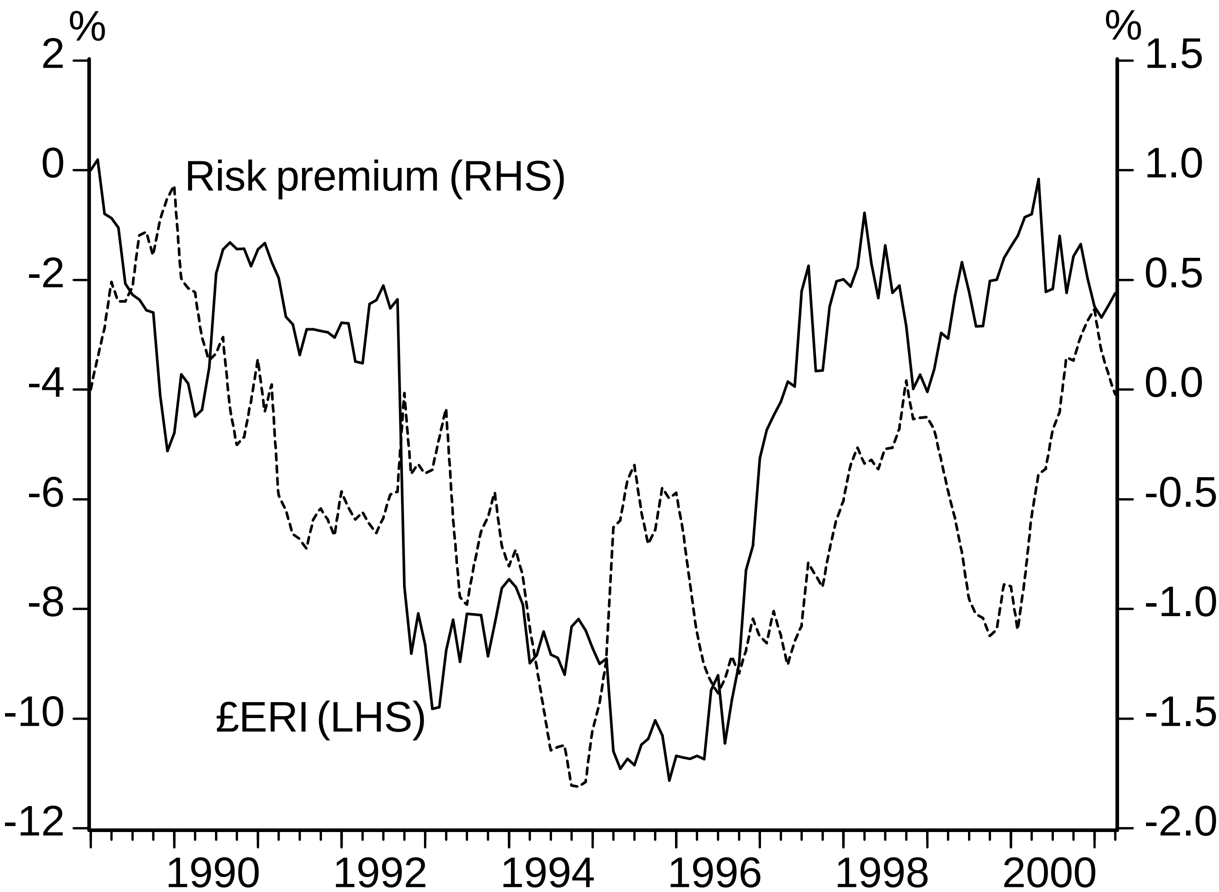


Chart 6: Accumulated changes in the sterling/Deutsche mark exchange rate and the inverse of the corresponding risk premium, 1989:01-2001:04

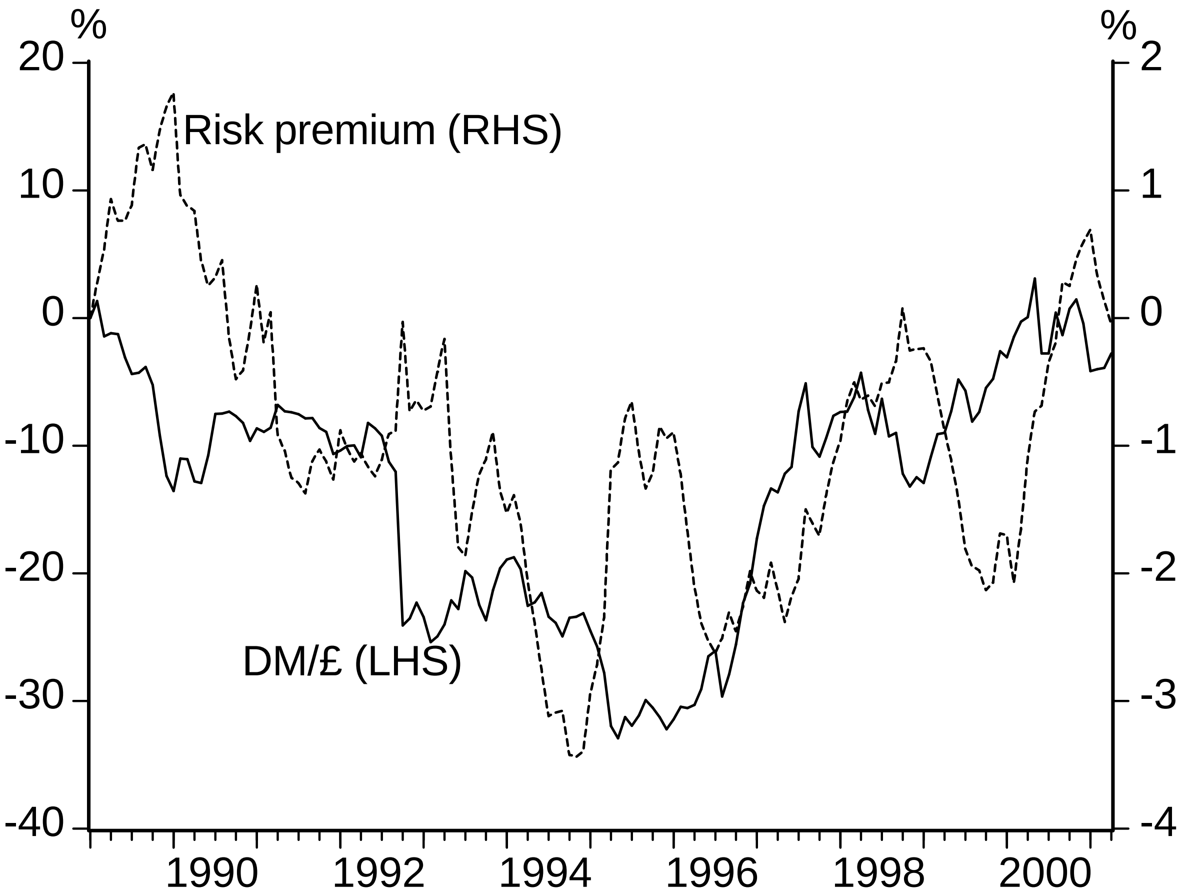
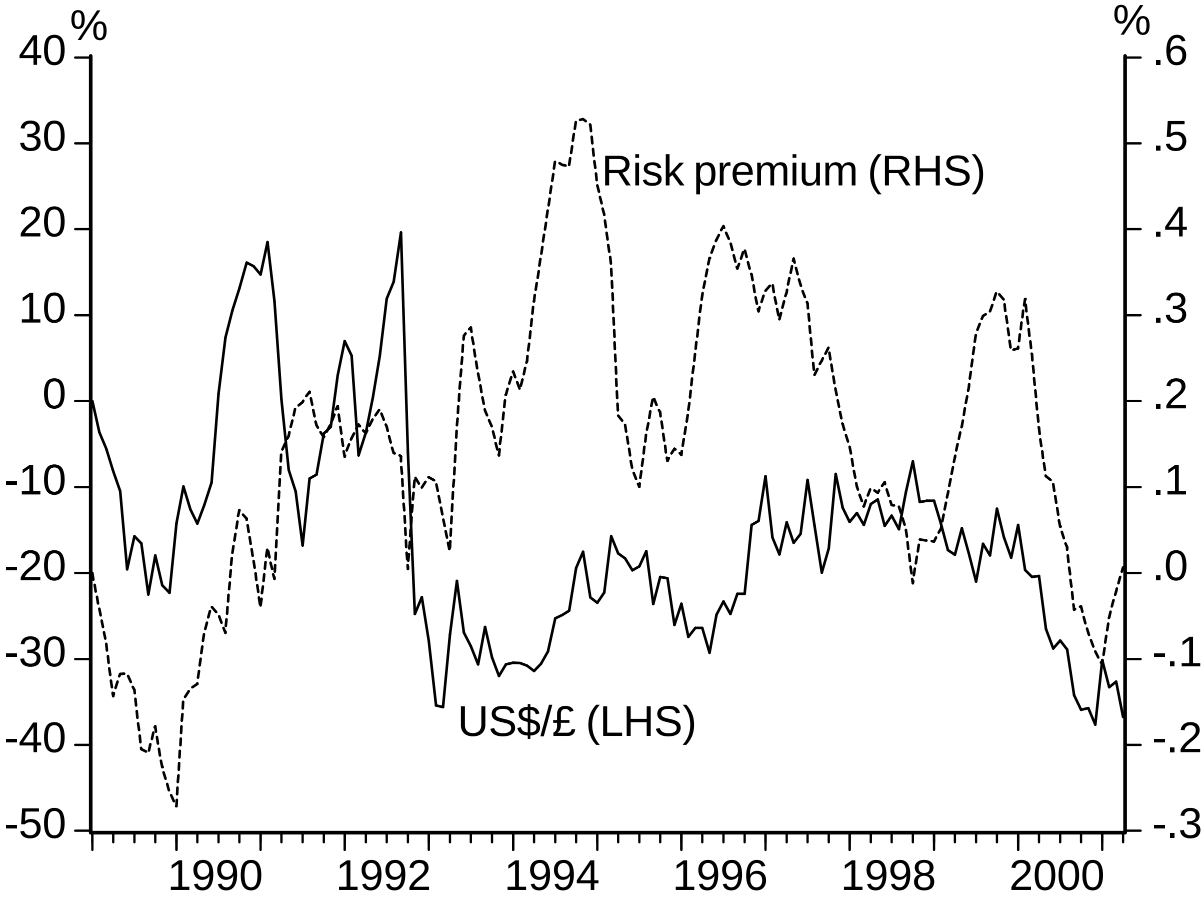


Chart 7: Accumulated changes in the sterling/US dollar exchange rate and the inverse of the corresponding risk premium, 1989:01-2001:04



When we focus on the 1996-97 sterling appreciation and recognise that the sterling/DM bilateral rate is the main driver of our measure of the effective sterling exchange rate, we observe in Chart 6 that during the second half of 1996 that the sharp sterling appreciation *vis-à-vis* the DM was accompanied by a sharp increase in the degree to which sterling is preferred to the Deutsche mark. During the second half of 1996, however, the risk premium stabilises whereas the sterling/DM exchange rate continues to appreciate. In fact while the sterling/DM rate appreciates over the 1996-99 period, the risk premium movement in favour of sterling during the first half of the period is reversed during the second half. This indicates that any role played by the sterling/DM risk premium during the 1996-99 sterling appreciation was only temporary in nature. If we assume that the sterling/DM risk premium is a proper proxy for the sterling/euro risk premium, one can observe that the sterling behaviour relative to the euro from mid-1999 onwards seems to be in line with the corresponding sterling/euro risk premium movements.

The previous graphical analysis indicates that there is some comovement between our estimates of the sterling risk premia and the observed exchange rate movements. Therefore it could be possible that the inclusion of our risk premium estimates in the standard UIP regression would improve its fit. According to UIP the expected future relative change in the exchange rate equals the current interest rate differential at proper maturities. Combining this with covered interest rate parity, risk neutrality and rational expectations we get for the k^{th} currency[12]

$$\Delta s_{k,t+1} = f_{k,t} - s_{k,t} \quad \textbf{(32)}$$

which implies that excess exchange rate returns are zero-mean iid distributed series. Under risk aversion, however, the forward-spot differential is augmented by the risk premium (ie $rp_{k,t}$), the inflation risk premium and the Jensen's inequality term

$$\Delta s_{k,t+1} = (f_{k,t} - s_{k,t}) + rp_{k,t} + Cov_t(\Delta s_{k,t+1}, \Delta p_{t+1,t}) + (-\frac{1}{2}Var_t(\Delta s_{k,t+1})) \quad \textbf{(33)}$$

and thus excess exchange rate returns are equal to the sum of risk premium, the inflation risk term and the Jensen's inequality term, see **(8)** and **(9)**.

In practice we test the appropriateness of **(32)** and **(33)** through regression

$$\Delta s_{k,t+1} = \delta_0 + \delta_1 \text{DUM-ERM}_t + \delta_2(f_{k,t} - s_{k,t}) + \epsilon_{k,t+1,t} \quad \textbf{(34)}$$

or its risk-adjusted version

$$\Delta s_{k,t+1} = \delta_0 + \delta_1 \text{DUM-ERM}_t + \delta_2(f_{k,t} - s_{k,t} + rp_{k,1}) + \epsilon_{k,t+1,t} \quad \textbf{(35)}$$

where $rp_{k,t}$ equals our estimated risk premium **(31)** and in **(35)** $\delta_0 + \epsilon_{k,t+1}$ contains the impact of the inflation risk premium and the Jensen's inequality terms. At a minimum we

[12] Covered interest rate parity states that interest rate differentials and the difference between the log of the forward exchange rate and log spot exchange rate are equal, ie $i_t^* - i_t = f_{k,t} - s_{k,t}$.

have to find that $\delta_2 = 1$ in order to accept (risk-adjusted) UIP. In **(34)** and **(35)** DUM-ERM$_t$ is a dummy variable which attains a value of 1 in September 1992 when pound sterling was forced out of the ERM due to speculative attacks. Note that in case of **(35)** applying OLS will yield inconsistent estimates of the parameter standard errors as the risk-adjusted forward premium $(f_{k,t} - s_{k,t} + rp_{k,1})$ is a generated regressor in the sense of Pagan (1984). To cope with this problem we estimate **(35)** with two-stage least squares (TSLS) where we use an intercept, DUM-ERM$_t$ and lags of $(f_{k,t} - s_{k,t} + rp_{k,1})$ as instrument variables. The estimation results for both **(34)** and **(35)** applied on our nine bilateral sterling rates can be found in Table C.

The first three columns of Table C contain the results for the plain UIP regression **(34)**. For most currencies we have the result that the forward premium has no significant impact on the relative future exchange rate change. There are, however, exceptions in the form of the Australia/UK and US/UK exchange rates. For these two bilateral sterling rates we not only find that the forward premium has a significant impact on the future rate of depreciation, the corresponding coefficient (δ_2) is insignificantly different from 1. Hence, for this period the UIP condition seems to hold for the Australia/UK and US/UK rates, and in particular for the US/UK rate this could explain the low magnitude for the corresponding risk premium estimates in Chart 4.

The results for the risk-adjusted UIP model **(35)** can be found in the last three columns of Table C. For the Australia/UK and US/UK rates the estimation results deteriorate relative to the unadjusted UIP case, as risk-neutral UIP seems to hold for these rates. However, the results are a bit more positive for the major European sterling bilateral exchange rates. That is the parameters of the corresponding risk-adjusted forward premia increase substantially in value and become closer to the optimal value of 1. This result could be the consequence of the observation that *vis-à-vis* European currencies risk-averse behaviour is much more important than relative to non-European currencies, eg compare the risk premium estimates in Charts 3 and 4. Note, however, that none of the estimation results in the last three columns of Chart C indicate that δ_2 in **(35)** becomes significantly different from 0.[13]

[13] Interestingly, for all regressions ARCH-LM tests on the residuals did not detect any conditional residual heteroskedasticity, which implies that the inflation risk and Jensen's inequality terms are in **(33)** and **(35)** constant. This corresponds with a general finding in the literature that at monthly frequencies GARCH effects in exchange rate returns, which are present at higher data frequencies, are absent, see eg Hsieh (1989).

Table C: UIP regressions with and without risk adjustment, 1989:01-2001:4

	UIP			*Risk-adjusted UIP*		
	δ_0	δ_1	δ_2	δ_0	δ_1	δ_2
Aus/UK	0.00	-0.15^{***}	1.16^{*}	0.01	-0.15^{***}	0.88
	(0.00)	(0.03)	(0.66)	(0.00)	(0.03)	(1.25)
			$[0.81]$			
Bel/UK	0.00	-0.14^{***}	-0.06	0.00	-0.14^{***}	0.05
	(0.00)	(0.02)	(0.05)	(0.00)	(0.02)	(0.42)
Can/UK	0.00	-0.10^{***}	0.61	0.00	-0.10^{***}	0.60
	(0.00)	(0.02)	(0.54)	(0.01)	(0.03)	(1.82)
Fra/UK	0.00	-0.15^{***}	0.23	0.00	-0.15^{***}	0.60
	(0.00)	(0.02)	(0.43)	(0.00)	(0.02)	(1.20)
Ger/UK	0.00	-0.14^{***}	0.26	0.00	-0.14^{***}	0.86
	(0.00)	(0.02)	(0.42)	(0.00)	(0.02)	(1.39)
Ita/UK	0.00	-0.00	0.02	0.00	-0.00	0.17
	(0.00)	(0.02)	(0.41)	(0.00)	(0.02)	(0.84)
Jap/UK	-0.00	-0.18^{***}	-0.09	0.02	-1.38	2.45
	(0.00)	(0.04)	(0.77)	(0.06)	(2.09)	(15.51)
Net/UK	0.00	-0.14^{***}	0.13	0.00	-0.14^{***}	0.58
	(0.00)	(0.02)	(0.42)	(0.00)	(0.02)	(1.20)
US/UK	0.00	-0.14^{***}	0.98^{*}	0.00	-0.15^{***}	0.45
	(0.00)	(0.03)	(0.50)	(0.00)	(0.03)	(1.34)
			$[0.97]$			

Note: The columns labelled δ_0, δ_1 and δ_2 contain estimates of the corresponding parameters of regression model **(34)** or **(35)**, which are based on OLS in the standard UIP case **(34)** and on TSLS in the case of risk-adjusted UIP **(35)**. An *(**)[***] indicates that the coefficient is significantly different from 0 at the 10% (5%) [1%] significance level based on a standard t-test (parameter standard errors are in parentheses). Conditional on a significant δ_2, the value in squared brackets equals a p-value of a t-test for $\mathrm{H}_0: \ \delta_2 = 1$.

4 Concluding remarks

In this paper we attempted to estimate the risk premium for a number of bilateral sterling exchange rates as UIP in its risk-neutral form is known to fail empirically. In intertemporal utility optimisation models the foreign exchange rate risk premium equals the conditional covariance between the future exchange rate change and the future marginal rate of substitution (or SDF) for the representative investor. In conventional models this SDF is a linear function of future consumption growth, which in turn is often substituted by the future real return on a stock portfolio. Empirical studies show that in most cases these SDF models are rejected by the data, see Engel (1996), and therefore foreign exchange rate risk premium measures that are based on this approach are not valid either.

We have, however, followed a more novel route in the sense that, like Lettau and Ludvigson (2001), we allow for habit persistence in the consumption behaviour of the representative international investor when we derive our measure of the SDF. This provides a motivation for the usage of the Cochrane (1996) conditional factor model to estimate the SDF of the representative agent. In this conditional factor model the SDF is related to the future real return on the agent's stock portfolio, but the model parameters are time varying and this time variation is related to movements in the slope of the term structure of interest rates. We have chosen the slope of the term structure as this variable has predictive power for future turning points in the real return on the stock portfolio. Another novelty relative to the existing literature is that our risk premium measures are related to an representative investor who operates on a global level instead of a representative investor from a particular country.

Our empirical results can be summarised as follows. Estimates of 'worldwide' factor models of the SDF show that, in contrast to an unconditional SDF model, the conditional factor model is accepted on a monthly 1987-2001 sample of nine major sterling exchange rates. We combine the parameter estimates of the 'worldwide' SDF model under habit persistence with the covariance between the relative change in a particular sterling rate and the real return on a 'world' stock portfolio to proxy the risk premium in the effective sterling exchange rate, the sterling/DM rate and the sterling/dollar rate. Resulting sterling risk premia exhibit large swings and seem especially of importance for the sterling/DM rate. A graphical analysis of the estimated sterling risk premia shows, however, that the impact of the risk premium movements on sterling exchange rates seems to be limited to the short to medium run. A check of both our SDF as well as risk premium estimates on a subsample indicated that our results are fairly stable.

As in reality the foreign exchange risk premium is unobservable, it is quite difficult to assess whether our estimates of the foreign exchange risk premium is close to the actual one. However, both the SDF parameter estimates as well as the risk premium estimates, which are based on the SDF parameter estimates, indicate that our approach has some empirical validity. Risk-adjusted UIP test regressions indeed indicate that relative to the major European currencies the usage of our estimated sterling exchange rate risk premia improve the parameter estimates slightly in favour of UIP, albeit not significantly so.

Appendix: Data

Our sample is on a monthly frequency and starts in December 1987 and ends in April 2001. The spot and forward sterling exchange rates are obtained from Datastream and these are for the last Friday of the month. Thus the profits on forward exchange rate speculation are based on the last Friday of the current month relative to the last Friday of the previous month. We use the sterling forward and spot rates for Australia, Belgium, Canada, France, Germany, Italy, Japan, the Netherlands and the United States.

The monthly return on the 'world' portfolio of wealth is measured through the monthly real return on equity. To measure the monthly *nominal* stock market return we use the relative change of the Morgan Stanley Capital International (MSCI) stock price index in local currency across all industrialised countries. This index is only available from December 1987 onwards, which explains the starting date of our analysis. The 'world' aggregate price level is proxied through the IMF's weighted average CPI across the industrialised countries, which is obtained from the IMF's *International Financial Statistics* (IFS). Based on this 'world' aggregate price level one can construct a 'world' inflation measure which in turn can be used to make the return on the 'world' MSCI index *real* (this is r_t^w). The slope of the 'world' term structure is proxied by a weighted average of the difference between the yield on the ten-year Treasury bond and the yield on the three-month Treasury bill across the G7 countries, where the weights are set to the relative GDP size in 1991. The ten-year T-bond yields and the three-month T-bill yields are retrieved from the IFS database.

Our measure of the effective sterling risk premium is constructed as a weighted average of the risk premium estimates across the nine sterling bilateral exchange rates. The weights reflect the relative size of the aforementioned nine countries in the broad effective sterling measure.

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