"EQUITY RISK PREMIA AND THE PRICING OF FOREIGN EXCHANGE RISK"

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Abstract

We investigate the relation between the risk premia observed in forward foreign exchange markets and international equity markets. If these markets share common sources of risk then the time variation in forward risk premia should be related to the forward contract's sensitivity to well-diversified equity benchmark portfolios and the time variation in the risk premia of those benchmark portfolios. We find that the forward contracts have a component of their conditional mean returns that is not reflected in their relation to the equity markets.

There is an large body of empirical work which indicates that forward prices for foreign exchange are not unbiased predictors of future spot exchange rates, [e.g., Hansen and Hodrick (1980, 1983), Bilson (1981), Korajczyk (1985), Mark (1985), Cumby (1988), and the extensive review by Hodrick (1987)]. That is, the evidence indicates that $E_{t-1}[S_t] \neq G_{t-1}$, where G_{t-1} is the forward exchange rate set at time t-1 for delivery at time t, S_t is the spot exchange rate at time t, and $E_{t-1}[\cdot]$ denotes expectations conditional on information available at time t-1.

This evidence has been variously interpreted as evidence of (a) inefficiencies in the forward market; (b) forward risk premia which vary through time; and (c) "peso" problems in which the anticipation of rare, but important, events influence the pricing of assets in ways that induce ex post bias in the forward rates when the observation period is short.

There is also a growing body of evidence that the risk premia on common stocks vary through time. This can be seen through the evidence of seasonality in stock returns [e.g., Gultekin and Gultekin (1983)] as well as evidence on the relation between equity risk premia and observable instruments for time varying expected returns [e.g., Keim and Stambaugh (1986) or Fama and French (1988)].

The purpose of this paper is to investigate the relation between the forecastable components of returns in the forward exchange market and international equity markets. In particular we wish to determine whether the observed risk premia in the forward market can be explained by the premia observed in the equity markets.

I. The Implications of an Intertemporal Asset Pricing Model

We begin by utilizing the first order conditions from a standard representative agent's discrete-time utility maximization problem, [see Lucas (1982) or Hodrick (1987, Chp.2)] which states that

$$E_{t-1}[Q_{mt}R_{jt}] = 1 \tag{1}$$

where R_{jt} is the gross nominal (currency m) return on asset j from t-1 to t and Q_{mt} is the marginal rate

of substitution of currency m between t-1 and t. Now let R_{ft} denote the gross return on a nominally (in currency m) riskless asset. Relation (1) implies that $R_{ft}E_{t-1}[Q_{mt}] = 1$. This, plus the definition of conditional covariance, leads to

$$E_{t-1}(R_{jt}) - R_{ft} = -R_{ft}Cov_{t-1}[Q_{mt},R_{jt}] \tag{2}$$

where $Cov_{t-1}[\cdot,\cdot]$ denotes the covariance conditional on time t-1 information.

From this we obtain a conditional asset pricing relation in terms of a benchmark portfolio which is on the conditional mean/variance frontier. Assume that there is a traded asset with returns given by

$$R_{mt} = Q_{mt}/E_{t-1}(Q_{mt})^2. \tag{3}$$

Hansen and Richard (1987) show that the conditional mean/variance frontier can be formed by linear combinations of R_{mt} and R_{ft}. That is, if $R_{bt} = \omega_{t-1}R_{mt} + (1-\omega_{t-1})R_{ft}$ then the benchmark portfolio return, R_{bt}, is conditionally mean/variance efficient. This allows us to write (2) as

$$E_{t-1}(R_{jt}) - R_{ft} = \beta_{jt-1}[E(R_{bt}) - R_{ft}] \tag{4}$$

where $\beta_{jt-1} = Cov_{t-1}[R_{jt},R_{bt}]/Var_{t-1}[R_{bt}]$.

This is true for any asset or portfolio, j. Consider the following portfolio. Invest S_{t-1} dollars in R_f, where S_{t-1} is the current spot exchange rate for the foreign currency, and enter into a forward contract for one unit of the currency. The excess rate of return on this portfolio is

$$r_{Gt} = (S_t - G_{t-1} + R_{ft}S_{t-1})/S_{t-1} - R_{ft} = (S_t - G_{t-1})/S_{t-1}.$$

Thus, we get an expression for the equilibrium difference between the forward exchange rate and the expected future spot rate

$$E(r_{Gt}) = \beta_{Gt-1}[E(R_{bt}) - R_{ft}]. \tag{5}$$

The single-beta relation in (5) has been the point of departure for a number of studies which treat the excess return on the benchmark portfolio as a latent variable [e.g., Hansen and Hodrick (1983), Hodrick and Srivastava (1984), and Giovannini and Jorion (1987)]. The latent variable approach utilizes the fact that movements in asset expected returns should be proportional to movements in the expected return on the benchmark portfolio, where the constant of proportionality is the conditional beta. That is, if[4]

$$E(r_{bt}) = [E(R_{bt}) - R_{ft}] = \alpha'\underline{z}_{t-1},$$

then $$E(r_{jt}) = \beta_{jt-1}\alpha'\underline{z}_{t-1} \tag{6}$$

for all j. If we assume that betas are constant through time ($\beta_{jt} = \beta_j$) or explicitly model the time variation in betas, then (6) implies testable restrictions on a pooled times series/cross-section of asset returns. These types of restrictions are tested in the studies cited above. We take a different approach in which we assume that the benchmark portfolio can be constructed from a linear combination of portfolios constructed to mirror the common movements across a set of international common stocks.

II. The Role of a Factor Model

We assume that asset returns follow a factor structure with the factors spanning the state variables that describe the evolution of the investment opportunity set. This implies that our

benchmark portfolio return, R_{bt}, is a linear combination of the returns on factor mimicking portfolios. This, in turn, implies that the risk premia in the forward exchange market should be determined by the forward contracts' conditional covariances with the factor mimicking portfolios. This allows us to use security return data to estimate the return on our benchmark portfolio.

Our assumed factor structure is given by

$$r_{jt} = E(r_{jt}) + b_{j1}\delta_{1t} + \ldots + b_{jk}\delta_{kt} + \varepsilon_{jt} \tag{7}$$

where b_{ji} is the sensitivity of asset j to factor i, δ_{it} is the realization of factor i in period t, and ε_{jt} is the diversifiable component of asset j's return. We assume that $E(\varepsilon_{jt}) = 0$; $E(\delta_{it}) = 0$ and $E(\delta_{it}\varepsilon_{j\tau}) = 0$ for all i, j, t, and τ. Let n denote the number of assets in the economy and V^n denote the covariance matrix of $\underline{\varepsilon}^n = (\varepsilon_{1t}, \ldots, \varepsilon_{nt})$. The diversifiability of the ε's implies that the eigenvalues of V^n are bounded as n approaches infinity, [see Chamberlain and Rothschild (1983)].

Constantinides (1989) discusses the relation between the pricing implications of (1)-(4) and the intertemporal arbitrage pricing theory. In general the marginal rate of substitution, Q_{mt}, is a function of all information available at t, denoted Φ_t. Let $\underline{s}_t$ denote the p-vector of state variables which, given information available at t-1, represent a sufficient statistic for Q_{mt} [that is $Q_{mt}(\Phi_t) = \hat{Q}_{mt}(\Phi_{t-1}, \underline{s}_t)$]. The expected returns on asset j will be determined by the factor sensitivities b_{ij} (i = 1, ..., k) as well as the covariance of ε_j with the state variables, $\underline{s}_t$. If we assume that the factors span the state variable, that is $\underline{s}_t = \Omega\underline{\delta}_t$,[2] then the covariance between ε_{jt} and $\underline{s}_t$ is zero since $E(\varepsilon_{jt}\delta_{it}) = 0$. Thus, under this spanning assumption, the expected returns on assets are determined by their conditional covariances with the factors. If we can construct portfolios which are perfectly correlated with the factors then the asset pricing model implies that our benchmark portfolio is a linear combination of these factor mimicking portfolios. Thus we can price assets, including forward contracts for foreign exchange, relative to these factor mimicking portfolios.

We use the asymptotic principal components technique of Connor and Korajczyk (1986, 1988)

to construct factor mimicking portfolios from the returns on common stocks. If exchange rates and common equities are influenced by the same factors or state variables, then we should be able to use the time variation in the expected returns on the factor mimicking portfolios to explain the apparent time variation in the risk premia in the forward market.

III. Data Description and Construction of Factor Mimicking Portfolios

We examine eight exchange rates in relative to the U.S. dollar: the British pound, the Canadian dollar, the Dutch guilder, the French franc, the Italian lira, the Japanese yen, the Swiss franc and the German mark. We use end-of-month exchange rates from January 1974 to December 1988. The starting date was chosen to coincide with the beginning of the current floating exchange rate regime (1973). One-month spot and forward exchange rates and one-month eurocurrency rates were taken from the Data Resources Incorporated (DRI) data files. The forward and spot exchange rates (bids and asks) represent opening prices in New York. The eurocurrency rates (bids and asks) are mid-morning rates in London (mid-morning in Paris for the euro-pound). Euro currency rates for Canada, Italy, and Japan are not available until August 1979, October 1980, and September 1979, respectively. For months before these dates we construct implied eurocurrency rates using the interest rate parity theorem.[3] In our empirical work below, we use averages of bid and ask quotes.[4]

Our sample of equity returns includes stocks from Australia, France, Japan, the United Kingdom, and the United States and spans the period January 1974 to December 1988. A summary of our data sources is presented in Table 1. The equity sample includes all assets traded on the Australian Stock Exchange, the New York and American Stock Exchanges as well as the NASDAQ system, the first section of the Tokyo Stock Exchange, the London Stock Exchange and the U.K. unlisted securities market. It includes approximately 20% of the Paris Bourse listed stocks.

Monthly returns adjusted for dividends and stock splits, are transformed into US dollar returns using end-of-month exchange rates from the DRI data files. To compute excess returns we use the US Treasury Bill returns from Ibbotson Associates (1988).

Given that the spot and forward exchange rates are quoted at the open in New York, eurocurrency rates are quoted mid-morning in Europe, and equity prices are quoted at the close of the respective exchanges, we have non-synchronous observations. The extent of the non-synchroneity is not large relative to the one-month observation period. Below we report various diagnostics which indicate that the results are not likely to be driven by non-synchroneity.

To estimate the excess returns on the factor mimicking portfolios we use the asymptotic principal components technique of Connor and Korajczyk (1986, 1988). The asymptotic principal components procedure can easily accommodate the large number of stocks in our sample (between 8010 and 11659 with return observations in a given month).

The procedure assumes the factor structure in (7) and that an exact multifactor pricing relationship holds, i.e.

$$E(r_j) = b_{j1}\gamma_1 + b_{j2}\gamma_2 + \dots + b_{jk}\gamma_k \tag{8}$$

Let T be the number of time periods; n the number of securities; R^n the n x T matrix of excess returns; F the k x T matrix of realized factors plus risk premia ($F_{it} = \delta_{it} + \gamma_{it}$) and B^n the n x k matrix of factor loadings. Note that the estimation procedure allows the risk premia, γ_{it}, to vary through time. Equation (8) implies that:

$$R^n = B^nF + \varepsilon^n \tag{9}$$

with: $E(F\varepsilon^{n\prime}) = 0$, $E(\varepsilon^n) = 0$, and $E(\varepsilon^n\varepsilon^{n\prime}) = V^n$.

Let Ω^n be the T x T matrix defined by $\Omega^n = R^{n\prime}R^n/n$ and $\hat{F}^n$ the k x T matrix of the first k eigenvectors of Ω^n. Under the assumption that asset returns follow an approximate k-factor model [in the sense of Chamberlain and Rothschild (1983)] Connor and Korajczyk (1986) show that $\hat{F}^n$ converges towards a non-singular linear transformation of F as n goes to infinity. We assume that our equity

sample size is large enough to consider $\hat{F}^n$ estimated from the sample as a transform of F. Consequently, the first k eigenvectors of Ω^n are estimates of the excess returns on factor mimicking portfolios. In order to use all available data in our sample we employ an extension of the principal components technique from Connor and Korajczyk (1988) which does not require that asset returns must exhibit continuous time series of returns. As a consequence we avoid any survivorship bias. While these types of factor portfolios do not fully explain the pricing of international equities, they perform well relative to some common alternative models, [Korajczyk and Viallet (1989)].

A major difficulty in any application of factor analysis is the determination of the appropriate number of factors. We choose to present our results using five factors. We have performed our tests using different numbers of factors and found that the main results of the paper are robust to changes in the number of factors.

IV. Empirical Results

We present evidence on the time series properties of exchange rate changes $[(S_t - S_{t-1})/S_{t-1}]$, forward contract returns $[(S_t - G_{t-1})/S_{t-1}]$, and forward premia $[(G_{t-1} - S_{t-1})/S_{t-1}]$ in Table 2. Consistent with previous evidence, the percentage change in exchange rates and the forward return show little evidence of autocorrelation while the forward premium exhibits significant autocorrelation. Also, the forward premium exhibits much lower variability than the other series whose standard deviations are between six and nineteen times the size of the standard deviation in the forward premium.

We begin our formal tests by documenting the fact that returns on forward contracts have forecastable components. Hence, forward prices cannot be conditionally unbiased predictors of future spot prices. In Table 3 we present results in which we regress $r_{Gt} = (S_t - G_{t-1})/S_{t-1}$ on a constant and the forward premium observed at time t-1, $(G_{t-1} - S_{t-1})/S_{t-1}$. If G_{t-1} is a conditionally unbiased predictor, then the intercepts and the slope coefficients should be equal to zero. We use data on eight exchange rates relative to the US dollar. The time period is January 1974 to December 1988. We reject the hypothesis that the intercepts are jointly zero and the hypothesis that the slope coefficients are jointly

zero. Thus, the results in Table 3 confirm the findings of others in that there is reliable evidence that forward exchange rates are not unbiased predictors of future spot rates. We also estimate the regression over several subperiods. The results are not reported here, but we reject the unbiasedness hypothesis over each subperiod.

We now wish to determine whether the time variation in expected returns on the forward contracts are explained by the risk premia in the factor mimicking portfolios. A necessary condition for this is that the factor risk premia are related, *ex ante*, to the instruments useful in predicting returns on the forward contracts. We test this by regressing the period t excess returns of the first five factor portfolios on the time t - 1 forward premia for all eight currencies. This is a standard 5-equation multivariate regression. If the coefficients on the forward premia are all zero, then we will not be able to explain the time variation in forward contract risk premia exhibited in table 3. A joint test that the coefficients are zero across all five equations produces a test statistic of 70.15 which should have a χ^2 distribution with 40 degrees of freedom (eight forward premia for each of the five factor portfolios). The p-value of this test statistic is 0.002 so we reject the hypothesis that the factor risk premia are unrelated to the forward premia at the beginning of the month.[5]

Part of the above explanatory power might be due to the problem of non-synchronous trading. For example, the prices determining the time t - 1 forward premia (quoted at the New York open) are observed after the time t - 1 stock prices in Japan which are used to calculate time t returns. To determine whether this non-synchroneity problem is causing the results cited above we regress the period t excess returns of the first five factor portfolios on the time t - 2 forward premia for all eight currencies. A joint test that the coefficients are zero across all five equations produces a χ^2_{40} test statistic of 65.65 (p-value = 0.007) so we still reject the hypothesis that the factor risk premia are unrelated to the forward premia at the beginning of the month at, for example, the 0.05 significance level.

We initially assume that the conditional factor betas of the forward returns are constant through time. We regress each currency's excess forward return, r_{Ct}, on a constant, the excess returns

on five factor mimicking portfolios, and the observable forward premium at the beginning of the period. If the factor mimicking portfolios represent the benchmark portfolio, r_{bt}, and the conditional betas are truly constant, then the intercept *and* the slope coefficient on the forward premium should be zero. The results are shown in Table 4. The factors, in general, have significant explanatory power as can be seen from the increase in R^2 from Table 3 to Table 4. We do not report the factor betas because their interpretation is made difficult by the standard rotational indeterminacy problem of factor analytic or principal components based methods. Inclusion of the factor mimicking portfolios does not change the joint significance of the intercepts or of the coefficients on the forward premium. In particular, the coefficients seem to change very little with the inclusion of the factor mimicking return. Thus, while the forward returns are significantly correlated with the factor returns, there remains a time-varying component of the forward returns which is unrelated to the time-varying component of equity returns.

The asset pricing model (5) is stated in terms of conditional betas. The results above use the assumption of constant conditional betas as a means of identifying the model. Rejection of the restrictions could be due to the fact that conditional betas are not constant [Hansen and Richard (1987)]. Giovannini and Jorion (1987, 1989) and Mark (1988) find that the performance of the latent variable model, described in section I, as well as a CAPM-based model is significantly improved when they allow for time variation in the conditional betas. We allow the forward contracts to have factor betas which are functions of the beginning-of-period forward premium and a dummy which splits the period into two subperiods 8-1973 to 9-1979 and 10-1979 to 12-1988.[6] While we can reject the hypothesis that the factor betas are constant, we still reject the hypothesis that the intercept and forward premium coefficients are zero. The results are reported in Table 5.

The *ex post* return on the forward contract and be decomposed into a real interest rate differential (across countries) plus the change in the real exchange rate [see, for example, Korajczyk (1985, p.349)][7]

$$r_{Gt} = (R_t^* - I_t^*) - (R_t - I_t) + \Delta\rho_t \tag{10}$$

where R_t^* (R_t) is the foreign (domestic) interest rate, I_t^* (I_t) is the foreign (domestic) inflation rate, and $\Delta\rho_t$ is the change in the real exchange rate. We now wish to determine whether the time varying component of $E[r_{Gt}]$ is due primarily to predictable components of interest rate differentials or changes in the real exchange rate, or both. Also, we are interested in the relation between the factor portfolio risk premia and the separate components of the forward return.

In Table 6 we present the results from regressing each component of $E[r_{Gt}]$ on five factor portfolio returns and the forward premium at time t - 1. The regressions in Table 6 impose the constant conditional covariance assumption used in Tables 3 and 4. The evidence indicates that the ability of the beginning-of-period forward premium to predict r_{Gt} is due to both components of r_{Gt}, the real interest differential *and* the change in the real exchange rate. The factor portfolio returns do not seem to capture all of the time variation in $E[r_{Gt}]$ in either of the components.

V. Conclusion

We investigate the relation between the risk premia observed in forward foreign exchange markets and international equity markets. If these markets share common sources of risk then the time variation in forward risk premia should be related to the forward contract's sensitivity to well-diversified equity benchmark portfolios and the time variation in the risk premia of those benchmark portfolios. We use a large cross-section (8010-11659) of international equity returns from five countries to estimate the factor benchmark portfolios.

We find that the equity risk premia and the forward risk premia are related but that forward contracts have a component of their conditional mean returns that is not reflected in their relation to the equity markets. This additional component in mean returns is found in both the real interest rate differential and real exchange rate change which make up the forward return. Potential explanations of this phenomenon are:

(a) There are sources of risk peculiar to the forward markets that are not reflected in equity markets or the factor portfolios fail to span the state variables relevant for determining the intertemporal marginal rate of substitution, Q_{mt}. For example we may need to construct factor portfolios from equity returns representing a larger sample of countries. A related argument is that the exchange rate risk is reflected in higher order factors. We have performed the tests with ten factor portfolios without changing the inferences reached above.

(b) The methods of constructing benchmark portfolios or modelling the time-variation in conditional betas have failed to reflect important influences in the true benchmark returns. For example, the factor extraction procedure puts some restrictions on how betas can vary through time. Also, our choice of modelling the time variation in forward contracts' conditional betas may be a cause of rejection. Other authors [e.g., Mark (1988) and Giovannini and Jorion (1989)] have utilized an autoregressive conditional heteroscedasticity (ARCH) approach, with mixed results.

(c) The equity and exchange markets are not fully integrated.

(d) Pricing related to rare events (the "peso" problem) leads to measured "risk premia" which are related to perceived changes in the probability of these events.

Because of the importance of the real interest differential in the roward risk premium, constructing benchmark portfolios more closely linked to the term structure may lead to additional insights. However, the predictability of the change in the real exchange rate, in addition to the real interest differential, indicates that the forward risk premium is not solely a term structure phenomenon. In addition to econometric innovations and additional data, an analysis of the effects of changes in capital controls or of the intervention process followed by monetary authorities may help us understand the predictable components of forward foreign exchange contracts.

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Table 1. Equity Data: sources and sample sizes.

A. Equity return sources

Country (Exchange)	Source of Data	Number of Firms in Sample[a]
Australia (Australian Stock Exchange)	Centre for Research in Finance Australian Graduate School of Management	3352
France (Paris Bourse)	Compagnie des Agents de Change	112
Japan (Tokyo Stock Exchange - First Section)	Japan Securities Research Institute	1130
United Kingdom (London Stock Exchange and Unlisted Securities Market)	London Share Price Data Base London Business School	5379
United States (NYSE, ASE, NASDAQ)	Center for Research in Security Prices University of Chicago	13614

B. Number of firms with observed returns per month.

Maximum	Average	Minimum
11659	9442	8010

[a] Number of firms represents the number of firms with any monthly returns data available in the period January 1974 to December 1988.

Table 2. Sample statistics: autocorrelations, means and standard deviations. Monthly data from January 1974 to December 1988.

	Autocorrelations												Mean*	σ*
Country	ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_7	ρ_8	ρ_9	ρ_{10}	ρ_{11}	ρ_{12}		
	$(S_t - S_{t-1})/S_{t-1}$													
United Kingdom	0.06	0.08	-0.03	0.07	0.07	-0.04	0.05	0.04	0.01	0.03	0.16	-0.02	-0.09	3.28
Canada	-0.07	-0.08	0.11	0.01	0.09	-0.04	-0.08	0.15	0.01	0.07	0.18	-0.15	-0.09	1.34
Netherlands	0.01	0.18	-0.01	0.05	0.06	-0.04	0.08	0.00	0.02	0.05	0.17	-0.06	0.24	3.28
France	-0.06	0.14	0.07	0.10	0.09	-0.01	0.11	-0.01	0.02	0.04	0.13	-0.04	-0.09	3.30
Italy	0.03	0.15	0.06	0.05	0.15	-0.02	0.13	0.04	0.06	0.05	0.10	-0.01	-0.38	3.05
Japan	0.05	-0.01	0.09	0.07	0.07	-0.09	0.09	0.02	0.01	-0.05	0.10	0.05	0.51	3.51
Switzerland	0.01	0.13	-0.03	0.02	0.05	-0.08	0.10	-0.03	-0.03	-0.01	0.14	-0.04	0.51	3.96
West Germany	-0.02	0.16	-0.02	0.03	0.05	-0.05	0.12	-0.01	0.04	0.04	0.17	-0.07	0.29	3.38
	$(S_t - G_{t-1})/S_{t-1}$													
United Kingdom	0.10	0.09	-0.02	0.07	0.07	-0.04	0.05	0.04	0.01	0.03	0.16	-0.01	0.13	3.35
Canada	-0.03	-0.07	0.10	0.01	0.08	-0.04	-0.07	0.15	0.02	0.08	0.16	-0.16	0.01	1.37
Netherlands	0.04	0.19	0.02	0.07	0.07	-0.03	0.08	0.00	0.02	0.05	0.18	-0.04	0.00	3.33
France	-0.03	0.15	0.07	0.10	0.08	-0.02	0.10	-0.04	0.00	0.04	0.11	-0.05	0.15	3.35
Italy	0.05	0.16	0.05	0.03	0.13	-0.06	0.09	0.00	0.03	0.01	0.07	-0.03	0.23	3.11
Japan	0.09	0.03	0.10	0.09	0.07	-0.09	0.09	0.02	0.01	-0.04	0.10	0.06	0.28	3.53
Switzerland	0.04	0.15	-0.01	0.04	0.06	-0.06	0.11	-0.02	-0.01	0.01	0.15	-0.03	0.05	4.03
West Germany	0.00	0.17	0.00	0.04	0.06	-0.04	0.12	-0.01	0.05	0.05	0.18	-0.06	-0.01	3.41

Table 2 (continued)

Country	ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_7	ρ_8	ρ_9	ρ_{10}	ρ_{11}	ρ_{12}	Mean	σ
	$(G_{t-1} - S_{t-1})/S_{t-1}$													
United Kingdom	0.84	0.71	0.56	0.44	0.38	0.32	0.27	0.25	0.22	0.18	0.18	0.13	-0.21	0.30
Canada	0.80	0.62	0.48	0.37	0.35	0.38	0.35	0.36	0.30	0.19	0.13	0.05	-0.10	0.14
Netherlands	0.11	0.05	0.02	0.06	0.05	0.05	0.03	0.06	0.03	0.00	-0.01	0.00	0.24	0.52
France	0.62	0.50	0.40	0.28	0.21	0.25	0.24	0.21	0.24	0.13	0.20	0.13	-0.23	0.37
Italy	0.64	0.51	0.38	0.29	0.23	0.21	0.15	0.16	0.10	0.02	0.06	0.07	-0.61	0.49
Japan	0.30	0.26	0.53	0.26	0.27	0.20	0.18	0.24	0.21	0.19	0.21	0.18	0.23	0.51
Switzerland	0.87	0.75	0.63	0.56	0.54	0.53	0.54	0.55	0.54	0.50	0.45	0.42	0.46	0.27
West Germany	0.80	0.56	0.39	0.25	0.23	0.27	0.34	0.44	0.47	0.40	0.31	0.26	0.30	0.18

* Mean and standard deviation (σ) are expressed as % per month.
Exchange rates are expressed in U.S. dollars per foreign currency. S_t is the spot exchange rate at the end of month t and G_{t-1} is the 30-day forward rate at the end of month t - 1.
Asymptotic standard errors for the autocorrelations range from 0.075 to 0.077 under the null hypothesis that $\rho = 0$.

Table 3. Seemingly unrelated regression of forward returns on forward premia. Monthly returns from January 1974 to December 1988. N = 180.

$$(S_t - G_{t-1})/S_{t-1} = \alpha_0 + \alpha_1(G_{t-1} - S_{t-1})/S_{t-1} + v_t$$

Country	α_0 x 100	α_1	R^2
United Kingdom	-0.30	-2.02	0.071
	(-1.13)	(-3.50)	
Canada	-0.26	-2.76	0.072
	(-2.22)	(-4.09)	
Netherlands	0.21	-0.86	0.033
	(0.85)	(-7.20)	
France	0.04	-0.46	0.016
	(0.15)	(-1.69)	
Italy	-0.36	-0.97	0.038
	(-1.35)	(-4.15)	
Japan	0.44	-0.72	0.011
	(1.61)	(-1.87)	
Switzerland	0.90	-1.86	0.056
	(2.39)	(-3.48)	
West Germany	0.37	-1.26	0.026
	(1.37)	(-3.60)	

T- statistics in parentheses.

Wald tests: (a) $\alpha_0 = 0$ for all countries: $\chi^2_8 = 25.59$, p-value = 0.001 (b) $\alpha_1 = 0$ for all countries: $\chi^2_8 = 105.24$, p-value < 0.001 and (c) $\alpha_0 = 0$ and $\alpha_1 = 0$ for all countries: $\chi^2_{16} = 109.93$, p-value < 0.001.

Table 4. Seemingly unrelated regression of forward returns on five factor mimicking portfolios and forward premia. Monthly returns from January 1974 to December 1988. N = 180.

$$(S_t - G_{t-1})/S_{t-1} = \alpha_0 + \beta_1\hat{F}_{1t} + \ldots + \beta_5\hat{F}_{5t} + \alpha_1(G_{t-1} - S_{t-1})/S_{t-1} + v_t$$

Country	α_0 x 100	α_1	R^2
United Kingdom	-0.63	-1.15	0.420
	(-2.79)	(-2.20)	
Canada	-0.36	-2.48	0.158
	(-3.09)	(-3.74)	
Netherlands	-0.26	-0.90	0.272
	(-1.15)	(-8.49)	
France	-0.45	-0.44	0.294
	(-1.98)	(-1.60)	
Italy	-0.75	-0.99	0.261
	(-2.93)	(-4.24)	
Japan	-0.37	-0.43	0.528
	(-1.78)	(-1.42)	
Switzerland	0.27	-1.68	0.277
	(0.73)	(-3.24)	
West Germany	-0.13	-1.20	0.245
	(-0.52)	(-3.88)	

T- statistics in parentheses.

Wald tests: (a) $\alpha_0 = 0$ for all countries: $\chi^2_8 = 29.58$, p-value < 0.001 (b) $\alpha_1 = 0$ for all countries: $\chi^2_8 = 116.86$, p-value < 0.001 and (c) $\alpha_0 = 0$ and $\alpha_1 = 0$ for all countries: $\chi^2_{16} = 129.98$, p-value < 0.001.

Table 5. Seemingly unrelated regression of forward returns on five factor mimicking portfolios and forward premia. Time varying betas. Monthly returns from January 1974 to December 1988. N = 180.

$$(S_t - G_{t-1})/S_{t-1} = \alpha_0 + \beta_{1t}\hat{F}_{1t} + \ldots + \beta_{5t}\hat{F}_{5t} + \alpha_1(G_{t-1} - S_{t-1})/S_{t-1} + v_t$$

$$\beta_{it} = b_{i0} + b_{i1}[(G_{t-1} - S_{t-1})/S_{t-1}] + b_{i2}D_t$$

Country	α_0 x 100	α_1	R^2
United Kingdom	-0.79	-1.53	0.482
	(-3.49)	(-2.78)	
Canada	-0.38	-1.79	0.292
	(-3.46)	(-2.78)	
Netherlands	-0.21	-1.71	0.329
	(-0.94)	(-6.53)	
France	-0.52	-0.64	0.313
	(-2.27)	(-1.95)	
Italy	-0.96	-1.15	0.318
	(-3.67)	(-4.33)	
Japan	-0.16	-0.97	0.581
	(-0.76)	(-2.69)	
Switzerland	0.55	-2.60	0.345
	(1.45)	(-4.55)	
West Germany	-0.03	-1.86	0.285
	(-0.10)	(-4.08)	

T - statistics in parentheses. D_t equals unity after September 1979 and zero otherwise.

Wald tests: (a) $\alpha_0 = 0$ for all countries: $\chi^2_5 = 41.51$, p-value < 0.001 (b) $\alpha_1 = 0$ for all countries: $\chi^2_5 = 80.77$, p-value < 0.001 (c) $\alpha_0 = 0$ and $\alpha_1 = 0$ for all countries: $\chi^2_{16} = 97.49$, p-value < 0.001 (d) $b_{i1} = 0$ for all countries, i = 1, ... 5: $\chi^2_{40} = 78.07$, p-value < 0.001 (e) $b_{i2} = 0$ for all countries, i = 1, ... 5: $\chi^2_{40} = 76.74$, p-value < 0.001 (f) $b_{i1} = 0$ and $b_{i2} = 0$ for all countries, i = 1, ... 5: $\chi^2_{80} = 159.32$, p-value < 0.001.

Table 6. Seemingly unrelated regression of real interest differentials and real exchange rate changes on five factor mimicking portfolios and forward premia. Monthly returns from January 1974 to December 1988. N = 180.

$$(R_t^* - I_t^*) - (R_t - I_t) = \alpha_0 + \beta_1\hat{F}_{1t} + \ldots + \beta_5\hat{F}_{5t} + \alpha_1(G_{t-1} - S_{t-1})/S_{t-1} + v_t$$

$$\Delta\rho_t = a_0 + b_1\hat{F}_{1t} + \ldots + b_5\hat{F}_{5t} + a_1(G_{t-1} - S_{t-1})/S_{t-1} + u_t$$

A. Real Interest Rate Differentials

Country	α_0 x 100	α_1	R^2
United Kingdom	-0.05	0.12	0.080
	(-0.86)	(0.76)	
Canada	-0.03	-0.32	0.055
	(-0.72)	(-1.73)	
Netherlands	-0.04	-0.06	0.077
	(-0.93)	(-1.00)	
France	-0.02	-0.34	0.106
	(-0.70)	(-5.01)	
Italy	-0.03	-0.22	0.067
	(-0.47)	(-2.75)	
Japan	-0.21	0.38	0.105
	(-3.13)	(3.93)	
Switzerland	0.21	-0.85	0.237
	(3.49)	(-7.98)	
West Germany	0.11	-0.44	0.199
	(2.49)	(-4.05)	

Table 6. (continued)

B. Changes in Real Exchange Rates			
Country	a_0 x 100	a_1	R^2
United Kingdom	-0.57	-1.27	0.420
	(-2.48)	(-2.38)	
Canada	-0.33	-2.10	0.126
	(-2.72)	(-3.05)	
Netherlands	-0.22	-0.87	0.257
	(-0.94)	(-7.17)	
France	-0.44	-0.17	0.293
	(-1.99)	(-0.67)	
Italy	-0.81	-0.93	0.228
	(-3.16)	(-3.96)	
Japan	-0.16	-0.83	0.519
	(-0.76)	(-2.72)	
Switzerland	0.17	-1.08	0.252
	(0.46)	(-2.09)	
West Germany	-0.16	-1.02	0.246
	(-0.62)	(-3.19)	

T - statistics in parentheses.

Wald tests: (a) $\alpha_0 = 0$ for all countries: $\chi^2_8 = 26.74$, p-value < 0.001 (b) $\alpha_1 = 0$ for all countries: $\chi^2_8 = 105.51$, p-value < 0.001 (c) $\alpha_0 = 0$ and $\alpha_1 = 0$ for all countries: $\chi^2_{16} = 171.64$, p-value < 0.001 (d) $a_0 = 0$ for all countries, i = 1, ... 5: $\chi^2_8 = 25.37$, p-value = 0.001 (e) $a_1 = 0$ for all countries, i = 1, ... 5: $\chi^2_8 = 91.98$ p-value < 0.001 (f) $a_0 = 0$ and $a_1 = 0$ for all countries: $\chi^2_{16} = 100.53$, p-value < 0.001.

Endnotes

1. We will use lower case r's to denote excess returns, while upper case R's denote gross returns.

2. Note that this requires $p \leq k$. Here $\underline{\delta}$ denotes the k x 1 vector of factor realizations and Ω is a p x k matrix with rank p.

3. Let R and R^* denote the domestic (U.S. dollar) and foreign eurocurrency rates, respectively. As before S and G denote spot and forward exchange rates. Interest rate parity implies that $R^* = (1 + R)(S/G) - 1$. When there are bid/ask spreads in the spot, forward, and euro-dollar market we can put conservative bounds on the bid and ask foreign eurocurrency rates. Let subscript a (b) denote ask (bid). A lower bound on R_b^* is $(1 + R_b)(S_b/G_a) - 1$ while an upper bound on R_a^* is $(1 + R_a)(S_a/G_b) - 1$.

4. Bossaerts and Hillion (1989) argue that use of bid/ask averages can lead to inconsistent estimates of the relation between forward prices and expected future spot prices. We have estimated (but do not report here) various models using an instrumental variables procedure similar to the one they suggest. We do not find that our inferences change much. It may be that this issue is more important when using daily data (as they do) as opposed to monthly data.

5. The evidence in Dominguez (1988) is also consistent with a significant relation between equity risk premia and foreign exchange risk.

6. The split at October 1979 was chosen because it corresponds with a change in monetary policy followed by the Federal Reserve Board.

7. The relation is exact with continuous compounding and approximate with discrete compounding.

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