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Exchange rate and foreign inflation risk premiums in global equity returns

Maria Vassalou *

Graduate School of Business, Columbia University, 416 Uris Hall, 3022 Broadway, New York, NY 10027-6902, USA

Abstract

We test for the pricing of exchange rate and foreign inflation risk in equities. Our tests are motivated by the empirical implications of the models of Solnik (1974b) [Solnik, B., 1974b. The international pricing of risk: an empirical investigation of the world capital market structure. Journal of Finance 365–377] as revised by Sercu (1980) [Sercu, P., 1980. A generalization of the international asset pricing model. Revue de l'Association Française de Finance **1**, 91–135], Grauer et al. (1976) [Grauer, F., Litzenberger, R., Stehle, R., 1976. Sharing rules and equilibrium in an international capital market under uncertainty. Journal of Financial Economics **3**, 233–256], and Adler and Dumas (1983) [Adler, M., Dumas, B., 1983. International portfolio choice and corporation finance: a synthesis. Journal of Finance **38**, 925–984]. Both exchange rate and foreign inflation risk factors can explain part of the within-country cross-sectional variation in returns. Our results have important implications for hedging exchange rate risk. They also demonstrate that home bias, at least in US equity portfolios, cannot be the result of US investors' efforts to hedge their domestic inflation.

Keywords: International asset pricing; Foreign inflation risk premiums; Exchange rate risk premiums

1. Introduction

The benefits of international diversification have been known for many decades, but it is only recently that investors have started allocating a significant portion of their portfolio holdings in foreign equities. To manage the risk of international port-

* Tel.: +1-212-854-4104; fax: +1-212-316-9180.
E-mail address: mv91@columbia.edu (M. Vassalou).

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folios, investors need to know the factors that explain the cross-sectional and cross-country variation in global equity returns.

Several studies have documented that the world market factor is an important determinant of asset returns (see, e.g., Solnik, 1974a; Stehle, 1977; Jorion and Schwartz, 1986; Korajczyk and Viallet, 1989; Harvey, 1991). There is also evidence that exchange rate and world inflation risk factors can explain part of the cross-country differences in the returns of equities (see, for instance, Dumas and Solnik (1995), and Ferson and Harvey (1994) among others). However, there is still no sufficient evidence to suggest that exchange rate and world inflation factors can also explain the within-country cross-sectional differences in returns. Our study contributes to this literature by testing for the presence of exchange rate and foreign inflation risk premiums in the cross-section of equity returns using individual security data from 10 developed markets.[1]

The hypotheses we test are motivated by three international asset pricing models, namely those of Adler and Dumas (1983), Solnik (1974b), Sercu (1980), and Grauer et al. (1976). We find that both exchange rate and foreign inflation risk factors can explain part of the within-country variation in average returns.

To test for the pricing of exchange rate risk, we combine information for a cross-section of exchange rates into two indexes. The first one, the common component index, combines information that is common to all exchange rates, whereas the second, the residual component index, captures fluctuations that are specific to the individual exchange rates. Our procedure has the advantage of reducing the dimensionality of exchange rates whereas at the same time it results in the inclusion of more information about changes in exchange rates in our tests than the single index approach adopted in previous studies. Our results reveal that at least part of the exchange rate risk premium in equities is attached to the residual components of exchange rates which were overlooked in previous studies. These components have important implications for the pricing and hedging of exchange rate risk.

Both the Grauer, Litzenberger, and Stehle model and the Adler and Dumas model suggest that inflation risk is priced. However, the testable implications of the two models are different. The Adler and Dumas model suggests the presence of as many inflation risk premiums in equities as there are countries. In contrast, the Grauer, Litzenberger, and Stehle model suggests that equities carry a single risk premium with respect to inflation. This is a result of their assumption that the purchasing power parity holds, and therefore that all stochastic inflation rates collapse to a single rate when they are expressed in terms of the same reference currency. To test the implications of the two models and to discriminate between them empirically, we test for the number of inflation risk premiums present in equity returns. In particular, we test whether US inflation is priced in all 10 countries and whether, in its presence, additional inflation rates earn a risk premium. We find that US inflation risk is priced

[1] Note that Jorion (1991) tested for the pricing of exchange rate risk in US equities and found that it is not priced. Furthermore, Chen et al. (1986) and Ferson and Harvey (1991) tested for the pricing of US inflation risk in US equities and found mixed results.

in all countries of our sample. This is interesting since it implies that home bias in US portfolios cannot be the result of US investors' efforts to hedge domestic inflation. Finally, world inflation risk, orthogonal to US inflation risk, does not appear to carry a statistically significant risk premium when US inflation uncertainty and exchange rate uncertainty is taken into account.

The rest of the study is organized as follows. Section 2 outlines the three international asset pricing models which motivate our tests and lays out our econometric approach. Section 3 describes the data and the portfolio construction methodology. Section 4 discusses the empirical results. We conclude in Section 5 with a summary.

2. Methodology

2.1. The three international CAPMs

In the international CAPM models of Adler and Dumas (1983), Solnik (1974b), Sercu (1980), and Grauer et al. (1976), expected excess returns of risky assets are linear functions not only of their betas with respect to the world market portfolio, but also with exchange rate or inflation risk factors.

The Adler and Dumas (AD) model assumes that investors of the L+1 countries have potentially different consumption preferences, and hence they measure inflation by different price indexes. Assume there are N risky assets of which the first $n=N-L$ are stock securities, and the remaining L are nominal bank deposits denominated in the L currencies. These L deposits are nominally risky when they are expressed in terms of the reference currency. Apart from the fact that they accumulate interest, their prices are essentially the exchange rates vis-à-vis the reference currency. The N+1st security is a bank deposit denominated in units of the reference currency and is instantaneously nominally riskless. In equilibrium, investors hold a combination of the world market portfolio, and an inflation hedge portfolio which hedges against the inflation risk of their country. We define all returns to be excess returns. Then the pricing relation of the Adler and Dumas model can be stated as follows:

$$E(R_k)=\gamma_0+\sum_{l=1}^{L+1}\gamma_l^{\pi}\beta_{kl}^{\pi}+\gamma^{w}\beta_k^{w} \tag{1}$$

where; $E(R_k)$ is the expected excess (over the risk-free interest rate) log return (per period) of asset k; γ_l^{π} is the expected excess return (risk premium) of a portfolio which is as highly correlated as possible with the inflation rate in country l; β_{kl}^{π} is the regression beta of asset k with the inflation rate of country l; γ^{w} is the expected excess return (risk premium) of the world market portfolio; and β_k^{w} is the regression beta of the return of asset k with the return on the world market portfolio.

The AD model implies that $\gamma_0=0$, $\gamma_l^{\pi}=E(i_l)-\gamma_0$, and $\gamma^{w}=E(R^{w})-\gamma_0$, where i_l denotes the inflation rate of country l, and R^{w} the return on the world market portfolio. If $\gamma_0\neq 0$ then a Black (1972)-type of version of the AD model should be correct.[2] In

[2] See Jorion and Schwartz (1986) for a similar argument.

that case, the return on the zero-beta portfolio should be equal to the risk-free rate plus γ_0

Solnik (1974a), and the revised version of his model as it appears in Sercu (1980), the S-S model as we will call it, assumes that the inflation rate of country l expressed in its home currency, is zero or nonstochastic. As a result, the L+1 inflation hedge funds of the Adler and Dumas model collapse to L exchange rate hedge funds, and the pricing relation of the S-S model becomes:

$$E(R_k)=\gamma_0+\sum_{l=1}^{L}\gamma_l^f\beta_{kl}^f+\gamma^w\beta_k^w \tag{2}$$

where; $\gamma_l^f=E(r_l^f)-\gamma_0$ is the expected excess return (risk premium) of a portfolio which is perfectly correlated with the return of bond l expressed in the reference currency, r_l^f, (i.e. the exchange rate between currency l and the reference currency L+1); and β_{kl}^f is the regression beta of the return on asset l with the return on the exchange rate between currencies l and L+1.

Finally, the Grauer, Litzenberger, and Stehle (GLS) model, which is the most restrictive of the three specifications, assumes that inflation is stochastic, but purchasing power parity (PPP) holds. Under these assumptions, it is easy to show that the AD model collapses to:

$$E(R_k)=\gamma_0+\gamma_l^\pi\beta_{kl}^\pi+\gamma^w\beta_k^w,\ \ \forall l=n+1,...N+1 \tag{3}$$

The AD, S-S, and GLS models assume that the first and second moments of security returns are constant. Under this assumption, conditional and unconditional moments are identical, and hence, the investment opportunity set is constant. We will test for the pricing of exchange rate and foreign inflation risk, as implied by these models, using unconditional moments.

2.2. *Econometric approach*

The models described in (1), (2), and (3) allow us to test for the pricing of exchange rate and inflation risk, but not for their relative importance. To test the latter hypothesis, we “nest” the three models into one specification. To do that, we overparameterize the AD model in the following manner:

$$E(R_k)=\gamma_0+\sum_{l=1}^{L+1}\gamma_l^\pi\beta_{kl}^\pi+\sum_{l=1}^{L}\gamma_l^f\beta_{kl}^f+\gamma^w\beta_k^w \tag{4}$$

In what follows, we use as our base the S-S and GLS models to estimate exchange rate and foreign inflation risk premiums, and model (4) to evaluate their *relative* importance in the pricing of equities. We will continue to refer to relation (4) as the AD model. Note, however, that this is strictly correct only if the inflation terms in (4) are stated in local currency rather than the reference currency L+1.

2.2.1. The reduction of dimensionality in the L exchange rates variables

Because exchange rates tend to move together to a large extent, the inclusion of changes of several exchange rates in the same regression model creates severe multicollinearity problems. In addition, to test effectively for the pricing of exchange rate risk, one needs to include changes in a large number of exchange rates which can result in the estimation of a large number of risk premiums at the expense of efficiency. The purpose of this section is to propose and lay out a methodology which will address both issues by simultaneously resolving the multicollinearity problem and minimizing the number of exchange rate risk premiums that need to be estimated.

Previous studies have chosen to include either an index of changes in exchange rates (see, e.g., Jorion, 1991; Ferson and Harvey, 1994), or changes in a small number of exchange rates (e.g., Dumas and Solnik, 1995).

In this paper, we include information on nine exchanges rates (those implied by the 10 countries in our data sample) combined in two indexes. One, the *common component index*, measures movements which tend to be common across all exchange rates. The second index, called the *residual exchange rate index*, aggregates the fluctuations which are specific to the individual exchange rates. This procedure has three advantages: first, it resolves the multicollinearity problem, second it reduces the dimensionality of exchange rates, and third, as it is shown below, it results in the inclusion of more information about changes in exchange rates than the single index method.

Similarly to previous studies, we work with changes rather than levels in exchange rates, and we measure exchange rates in logs. Under the assumption that exchange rates follow a random walk, changes in exchange rates represent innovations.[3]

Our procedure involves the following steps: We project the changes in each of the L exchange rates on the changes of the remaining $L-1$ exchange rates through the following regression:

$$r^f_{jt}=\delta_{0j}+\sum_{1\leq l\leq L,\, l\neq j}\delta_{lj}r^f_{lt}+\epsilon_{jt} \tag{5}$$

where $E(\epsilon_{jt})=0$; $cov(r^f_{lt},\epsilon_{jt})=0$, $\forall 1\leq l\leq L$. The residuals ϵ_j of the changes in the j exchange rate represent the component of r^f_j that is not explained by the changes in the remaining exchange rates, i.e. *the residual component* of r^f_j. The *common (or systematic) component* of the L exchange rates κ_j, is defined as follows:

$$\kappa_{jt}=r^f_{jt}-\delta_{0j}-\epsilon_{jt} \tag{6}$$

We furthermore define the deviation of the common component of the L exchange rates from its mean as $\eta_{jt}=\kappa_{jt}-\bar{\kappa}_j$ where $\bar{\kappa}_j$ denotes the sample mean of κ_j. By con-

[3] There is some evidence of predictability in changes in exchange rates—see for instance, Bekaert and Hodrick (1992). However, this predictability is small, and therefore, the random walk remains a reasonable approximation of the process followed by exchange rates.

struction, $E(\eta_{jt})$=0; $cov(\epsilon_{jt},\eta_{jt})$=0. Up to this point, there is no loss of information from the decomposition presented. Each exchange rate is simply a linear combination of its common and idiosyncratic component. We then construct two equally weighted indexes corresponding to the sets of residuals obtained from (7) and (8):

$$e_t=\frac{1}{L}\sum_{j=1}^{L}\epsilon_{jt} \tag{7}$$

and

$$\lambda_t=\frac{1}{L}\sum_{j=1}^{L}\eta_{jt} \tag{8}$$

The variable e_t, is the average residual component of changes in all L exchange rates, whereas λ_t describes the average common component shared by changes in the same exchange rates. The creation of the two indexes is necessary in order to minimize the number of exchange rate betas and risk premiums that need to be estimated. This gives rise to some loss of information. We evaluate below the information contained in the two indexes, as well as the information lost.

Table 1 compares the common and residual component indexes with an equally weighted index.[4] The comparison is performed in US dollars which is the reference currency of this study. Panel A presents correlation coefficients. The correlation of λ_t with the equally weighted index is 0.991 which means that the common component index is virtually identical to the equally weighted index of all exchange rates. Furthermore, the residual component index has a correlation of 0.228 with the common component index and a correlation of 0.355 with the equally weighted index of changes in all exchange rates. In Panel B of Table 1 we report the adjusted R-squares from OLS regressions of changes in exchange rates on the constructed indexes. It is interesting to note that the common and residual component indexes can jointly explain a larger proportion of the variation in exchange rates than the equally weighted index. Although there is always loss of information by grouping exchange rates into indexes, this loss is smaller when exchange rates are grouped into two indexes with the method proposed here rather than in a single index. The additional information about changes in exchange rates contained in the residual component index should increase the power of our tests regarding the pricing of exchange rate risk relative to those of previous studies which used the single index approach. This is important since Jorion (1991), for instance, who used the single index approach to test for the pricing of exchange rate risk, failed to reject the hypothesis that changes in exchange rates receive a zero risk premium.

Finally, it is worthwhile to mention that in a principal component analysis, not presented here, the first factor was effectively an equally weighted average of the changes in all exchange rates, whereas several of the other factors could be inter-

[4] For a description of the data, see Section 4.1.

Table 1
Statistical properties of the common and idiosyncratic exchange rate indexes[a]

Panel A: Correlation coefficients

		corr(λ,EW)	*corr(λ,e)*	*corr(e,EW)*
		0.991*	0.228*	0.355*

Panel B: Coefficients of determination

Country	$R^2(\lambda)$	$R^2(e)$	*adj.* $R^2(\lambda,e)$	$R^2(EW)$
Australia	0.20	0.49	0.58	0.28
Canada	0.08	0.09	0.15	0.11
France	0.85	0.02	0.86	0.83
Italy	0.71	0.06	0.72	0.71
Germany	0.93	0.02	0.96	0.86
Japan	0.57	0.21	0.66	0.62
Switzerland	0.80	0.07	0.81	0.80
Netherlands	0.95	0.02	0.98	0.89
UK	0.60	0.17	0.66	0.64

[a] The correlation coefficient between the common component index and the equally weighted index is denoted by *corr(λ,EW)* and the correlation coefficient between the common component index and the residual index by *corr(λ,e)*. Furthermore, *corr(e,EW)* denotes the correlation coefficient between the residual component index and the equally weighted index. The coefficients of determination from regressions of the changes in exchange rates on the common component index, and the residual component index, are denoted by $R^2(\lambda)$ and $R^2(e)$, respectively. The corrected coefficient of determination from bivariate regressions on both the common and residual component indexes is denoted with *adj.* $R^2(\lambda,e)$ whereas $R^2(EW)$ denotes the coefficient of determination from regressions of the changes in exchange rates on the equally weighted index of all exchange rates. The reference currency is the US dollar. Exchange rates are quoted as foreign currency per US dollar. The statistics are calculated using monthly observations from January 1973 to December 1990, i.e., 216 observations in total.
*significant at 1% level.
**significant at 5% level.

preted as the residual components from the above decomposition. One may therefore understand the common component index as representing the first factor from the principal component analysis, and the residual component index as a combination of the rest.

2.2.2. Testing for the pricing of foreign inflation risk

Recall that the GLS model suggests the presence of a single inflation risk premium in equity returns whereas the AD model specifies that the inflation uncertainty of all countries should be priced. To empirically discriminate the two models we need to test for the number of inflation risk premiums contained in international equities. The GLS model provides no guidance as of which inflation rate should be priced. This is because in the GLS model PPP holds, and all country inflation rates collapse to a single inflation rate when expressed in terms of the same reference currency.

Given the prominence of the United States in the world economy and the international capital markets, and the fact that the US dollar is the reference currency of this study, we choose to test whether US inflation uncertainty is priced in the equity returns of all countries in our sample. This hypothesis corresponds to testing the GLS model. In order to empirically distinguish the AD model from the GLS model, we also test whether inflation uncertainty which is unrelated to US inflation is also priced in international equity returns. If both null hypotheses of zero inflation risk premiums are rejected, we can reject the GLS model in favor of the AD model. If, however, we only reject the first hypothesis, we can conclude that the empirical predictions of the AD model do not dominate those of the GLS model.[5] Finally, if we only reject the second hypothesis then we can again reject the GLS model in favor the AD model, since its prediction that a single inflation rate is priced would not be empirically supported.

To test the above hypotheses, we construct an index of world inflation that contains the inflation rates of all countries in our sample other than the US. This index uses GDP weightings and is expressed in US dollars:[6]

$$i_t^g = \sum_{l=1}^{L} \phi_l i_{lt} \tag{9}$$

where: i_{lt} denotes the inflation rate of country l, where l=1,...L. The L+1[st] inflation rate is the US (reference currency) inflation rate; φ_l is a GDP weight that aims to proxy national wealth weights and is updated on a yearly basis; and, i_t^g denotes world inflation.

To the extent that inflation processes across countries are not independent, the index of world inflation can be correlated with the US inflation. Since we are interested in the pricing of world inflation risk which is residual to the US inflation risk, we render the two series orthogonal to each other through the following projection:

$$i_t^g = v_0 + v_1 i_{L+1t} + v_t \tag{10}$$

where: v_t denotes the residual world inflation orthogonal to the US inflation. It follows again that $E(i_{L+1t})=E(v_t)=cov(i_{L+1t},v_t)=0$

Given that the models are tested using unconditional moments, the variables in (9) and (10) should represent levels of inflation rates rather than their innovations

[5] Note that we cannot formally reject the AD model since it is the most general specification examined in this study.

[6] The weighting scheme employed is motivated by the Adler and Dumas model. Note that even under the simplifying assumption of equal average risk tolerances across countries, equation (14) in Adler and Dumas (1983) can be written using the notation of this study as:

$$E(R_k) = \gamma^w \beta_k^w + \gamma^\pi \sum_{l=1}^{L+1} W_l \beta_{kl}^\pi$$

where W_l denotes the wealth of country l and γ^π is the world inflation risk premium. Note that the inflation risk premium of country l in Eqs. (1, 3) and (4) of this study is related to γ^π in the following manner: $\gamma_l^\pi = \gamma^\pi W_l$. This implies that the world inflation index needs to be weighted in such a way so as to reflect the relative wealth of each country. I am thankful to Piet Sercu for pointing out this need. A proxy for a country's wealth is its gross domestic product (GDP).

(unexpected inflation). Note, however, that the the inflation series are nonstationary[7], and therefore, regressions of equity returns on inflation levels would be unbalanced. To avoid this problem, we filter the inflation series using an ARIMA(0,1,1) model and use in our tests the innovations which represent unexpected inflation.[8] Therefore, in the rest of the paper, the variables in relations (9) and (10) should be understood as denoting innovations rather than levels of inflation rates.[9]

2.3. The econometric specification of the three competing models

Based on the data transformations performed in Section 2.2.1, the changes in the *L* exchange rates that appear in the S-S model can be substituted by the common and residual component indexes of exchange rates. In other words, we will assume that

$$\sum_{l=1}^{L}\gamma_l^f\beta_{kl}^f=\gamma^e\beta_k^e+\gamma^\lambda\beta_k^\lambda \tag{11}$$

where: γ^e is the excess residual component index of changes in exchange rates. It represents the exchange rate risk premium with respect to the residual component index; β_k^e is the regression beta of the return on asset *k* with the residual component index; γ^λ is the excess common component index of changes in exchange rates, and it is the risk premium with respect to the common component index of changes in the *L* exchange rates; and β_k^λ is the regression beta of the return on asset *k* with the common component index.

If we substitute the above relation in (2) the S-S model becomes:

$$E(R_k)=\gamma_0+\gamma^e\beta_k^e+\gamma^\lambda\beta_k^\lambda+\gamma^w\beta_k^w \tag{12}$$

In a similar manner, we assume that the *L*+1 inflation rates in the Adler and Dumas model can be well approximated by the innovations in the US inflation and the index of residual world inflation. In particular, we assume that:

$$\sum_{l=1}^{L+1}\gamma_l^\pi\beta_{kl}^\pi=\gamma^i\beta_k^i+\gamma^v\beta_k^v \tag{13}$$

where: γ^i is the US unanticipated inflation rate in excess of the risk free rate, and denotes the US unanticipated inflation risk premium; β_k^i is the regression beta of the return of asset *k* with the US unanticipated inflation rate. γ^v is the residual world

[7] For tests of unit roots in inflation series see Crowder (1996), Siklos and Wohar (1997) and Vassalou (1994, 1996).

[8] The ARIMA(0,1,1) model is widely used for inflation forecasting. For evidence on the comparative performance of the model see Fama and Gibbons (1984).

[9] Recall that the use of inflation innovations in our tests is not inconsistent with the way we treat returns on the world market portfolio and changes in exchange rates. As noted earlier, under the random walk hypothesis, both the returns on the world market portfolio and changes in exchange rates represent innovations.

inflation rate in excess of the risk-free rate. It denotes the residual world inflation risk premium; and β_k^v is the regression beta of the return on asset k with the world unanticipated inflation rate.

Under the above assumptions on the L exchange rates and the L+1 inflation rates, the AD model can be written as:

$$E(R_k)=\gamma_0+\gamma^i\beta_k^i+\gamma^v\beta_k^v+\gamma^e\beta_k^v+\gamma^\lambda\beta_k^\lambda+\gamma^w\beta_k^w \quad (14)$$

Finally, in line with the assumptions of the GLS model, we assume that all inflation rates collapse to the US inflation rate when they are expressed in terms of US dollars. We therefore state the GLS model as follows:

$$E(R_k)=\gamma_0+\gamma^i\beta_k^i+\gamma^w\beta_k^w \quad (15)$$

To formulate empirical tests for the three models, we decompose the rate of return on asset k into an expected component $E(R_{kt})$ and a set of innovations. In the case of the S-S model this yields:

$$R_{kt}=E(R_{kt})+\beta_k^w(R_t^w-E(R_t^w))+\beta_k^e e_t+\beta_k^\lambda\lambda_t+\zeta_{kt} \quad (16)$$

Substituting (11) into (14) gives

$$R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^e\beta_k^e+\beta_k^e e_t+\gamma^\lambda\beta_k^\lambda+\beta_k^\lambda\lambda_t+\zeta_{kt} \quad (17)$$

We repeat the same procedure for the AD model which now becomes:

$$R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^e\beta_k^e+\beta_k^e e_t+\gamma^\lambda\beta_k^\lambda+\beta_k^\lambda\lambda_t+\gamma^i\beta_k^i+\beta_k^i i_t+\gamma^v\beta_k^v+\beta_k^v v_t +\xi_{kt} \quad (18)$$

Similarly, the GLS model can be written as:

$$R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^i\beta_k^i+\beta_k^i i_i+\psi_{kt} \quad (19)$$

Under the null hypotheses that exchange rate and inflation risks are not priced, $\gamma^e=\gamma^\lambda=\gamma^i=\gamma^v=0$. Also, note that $E(\zeta_{kt})=E(\psi_{kt})=E(\xi_{kt})=0$.

To estimate the models (17), (18), and (19) we construct K portfolios of security returns for each of the 10 countries in our sample, following the methodology described in Section 3.2. Each model is estimated in a system of $K\times10$ equations, allowing for contemporaneous correlations in error terms as in seemingly unrelated regressions (SUR). The β and γ coefficients in each model are estimated simultaneously which avoids the errors-in-variables biases of the coefficients imbedded in two-stage procedures, such as the classic Fama–MacBeth (1993) methodology.

Notice that despite the fact that our tests include the world market risk factor, the models (17), (18), and (19) do not include a term for the world market risk premium.

However, an estimate for the world market risk premium can be easily calculated through the equation $\gamma^w = E(R^w) - \gamma_0$. When β and γ coefficients are estimated simultaneously, one γ coefficient is rendered redundant. We chose that coefficient to be the world market risk premium since the pricing of the world market factor is well known and well documented in the literature and does not warrant further investigation. The purpose of this paper is to test for the pricing of exchange rate and inflation risk, as well as their relative importance, *over and above* that of the world market factor. The methodology adopted here was initially proposed in Gibbons (1982), and subsequently used in Jorion and Schwartz (1986), among others. It is particularly suitable for estimations using the generalized methods of moments (GMM).

We map the models into an iterated GMM procedure, and employ the Newey–West (1987) estimator. Tests of unconditional mean-variance efficiency using GMM were first performed in MacKinlay and Richardson (1991). Evidence in Ferson and Foerster (1994) suggests that the iterated GMM procedure has better small sample performance than the one-step GMM estimation, while it maintains the same asymptotic distribution theory. The truncation parameter q in the Newey–West estimator was set equal to six, and corresponds to the number of residual autocorrelations that were found statistically significant at the 10% level.[10]

In the Newey–West estimator, the weighting matrix employed is "optimal" in Hansen's (1982) sense, and the minimized sample analog of the quadratic function follows a χ^2 distribution. This means that we can directly perform Hansen's (1982) *J*-test on the overidentifying restrictions of the models. Furthermore, we can use Newey–West's (1987) *D*-test to compare the minimized objective function of the restricted models (i.e. models (17), and (19)) with that of the unrestricted model (i.e. model (18)).

As additional diagnostic tests, we compute mean pricing errors (ME) and root mean square pricing errors (RMSE) from the models estimated. Finally, we compute adjusted R-squares.

3. Data and portfolio construction

3.1. Data

Our study uses monthly stock returns from 10 countries namely, Australia, Canada, France, Italy, Switzerland, the Netherlands, Japan, Germany, UK, and USA.[11]

[10] Because in several cases, the significant autocorrelations were not equal to six, we repeated the main tests of this paper setting $q=3$ which corresponds to the number of the *majority* of autocorrelations significant at the 10% level. The t-values changed only trivially with this modification, and our results remained qualitatively the same.

[11] Other international asset pricing tests that use individual security returns include those of Korajczyk and Viallet (1989), and Jorion and Schwartz (1986). The first study uses security returns from the USA, UK, Japan, and France, while the second one from USA and Canada.

Our sample runs from January 1973 to December 1990. Data for the USA are from the files of the Center for Research in Security Prices (CRISP), whereas data for the UK are from the London Share Price Database (LSPD) compiled at London Business School. The remaining data are extracted from Datastream.

Note that Datastream allows the downloading of monthly prices and dividend yields for individual securities, but not of total returns. Therefore, total returns for eight of the 10 countries in our sample have been calculated by spreading evenly the monthly dividends throughout each year. This method, which represents the only option we had available, may smooth the series to a certain extent but it is not expected to affect the means in any meaningful way. Furthermore, the parameter estimates should also be unaffected. Sharpe and Cooper (1972) have shown that beta estimates remain the same, independently of whether we use for our regressions total returns, or simply, capital gains. We replicated their tests using a small number of securities from each of the 10 countries. Our results are consistent with theirs. For that reason, we do not report them.

All country samples are comprised of security returns with continuous record. This, in principle, can impart a survivorship bias in our results. In Section 4.4.4 we test for this possibility and conclude that our results on the pricing of exchange rate and inflation risk cannot be affected in a significant manner by the survivorship bias in our database.

Table 2 reports the number of securities in each country sample and summarizes

Table 2
Distributional properties of stock returns[a]

Country	Number of tests	Filter Ho: 0.27%	Reversals Ho: 1%	mean	variance	skewness	kurtosis
Australia	95	1.30%	0.70%	0.008	0.021	−0.655	7.990
Canada	32	1.07%	0.72%	0.007	0.007	−0.197	3.270
France	104	1.05%	0.60%	0.011	0.011	0.109	2.225
Germany	143	1.02%	1.07%	0.006	0.006	0.179	3.910
Italy	65	1.18%	0.71%	0.013	0.013	0.439	5.800
Japan	112	1.34%	1.37%	0.011	0.011	0.589	3.740
Netherlands	101	1.32%	0.53%	0.011	0.011	−0.010	7.230
Switzerland	67	1.16%	1.00%	0.005	0.005	−0.243	4.664
UK	600	1.45%	1.02%	0.013	0.013	0.113	4.297
USA	400	1.22%	1.03%	0.098	0.098	0.450	6.985

[a] All tests were performed in the local currency. The second column gives the number of securities available in each country sample. Only securities with a continuous record for the entire time period were included (i.e. Jan. 1973 to Dec. 1990). The third column describes the results of the filter rule test which aims to detect the percentage of observations outside a ± three standard deviations interval. The fourth column reports the percentage of observations that exhibit reversals outside the 1.47 standard deviations interval. Under the null hypothesis that the data follow a normal distribution, only 1% of the observations should exhibit reversals. The remaining four columns report the average values of the first four moments across securities of the same country. Under the null hypothesis of a normal distribution, both the skewness and kurtosis should be equal to zero.

the distributional properties of the individual security returns. The purpose of the tests presented in this table is to verify the quality of our individual security returns.

A filtering rule was imposed to check for the number of observations that lie outside an interval of plus or minus three standard deviations, and the results were compared with those expected under the hypothesis that the data are sampled from a normal distribution. Our evidence coincides with results from previous studies (see, e.g., Fama, 1965) which indicate the presence of fat tails in the distributions of asset returns.

We also check all samples for the presence of reversals relative to the mean. To render the test powerful, we choose a rule that allows 1% of the observations to exhibit reversals. This rule permits us to make meaningful inferences without picking reversals of insignificant value. Under the normal distribution hypothesis, the 1% rule prescribes us to check for reversals outside the 1.47 standard deviations interval. The results show that in all samples the percentage of observations that passed the rule is very close to the theoretical one. This indicates that there are no “flags” or other significant punching errors in our data.

Finally, we compute the first four moments for each security in our sample, and report the average values of these moments for each country sample. The third moment denotes skewness, and under the null that the data are generated by a normal distribution, it should be equal to zero. We find that the data are slightly skewed to the right, with the exception of Australia, Canada, and Switzerland where small negative values were obtained. The fourth moment denotes kurtosis. The estimated values confirm the results of the filtering rule and show that all samples are to some degree leptokurtic with that phenomenon being more apparent in Australia, the Netherlands, and the USA.

Overall, our results on the distributions of security returns are consistent with those in the literature. They suggest that our data are of comparable quality to those in other studies, and they confirm the need to employ a distribution-free estimator for our tests, since security returns do not appear to be normally distributed.

In the rest of the paper, we proxy the world stock market portfolio with the Morgan Stanley Capital International (MSCI) world index. Country indexes for the 10 countries in our sample are also from MSCI. Spot exchange rate data are extracted from the OECD files. The series of 30-day Treasury Bill (T-bill) interest rates was obtained from the “Encorr” database maintained by Ibbotson Associates. Inflation rates are from the IMF Series and are calculated from each country’s Consumer Price Index (CPI). Data for Australian inflation are available only on a quarterly basis. The monthly inflation series is therefore computed by spreading evenly the quarterly inflation over the three month’s period. Finally, GDP data were obtained from the International Financial Statistics (IFS), 1991 Yearbook.

Table 3 provides summary statistics for all data used in the study. Our reference currency is the US dollar. The statistics are means, standard deviations, and autocorrelations up to order twelve. All variables are calculated in excess of the holding period return on the US 30-day T-bill. Portfolio 1 for a given country represents the portfolio with the lowest betas against the world market portfolio, the equally weighted index of all changes in exchange rates, and the GDP weighted index of

Table 3
Summary statistics

	Mean	S.D.	Autocorrelations ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_7	ρ_8	ρ_9	ρ_{10}	ρ_{11}	ρ_{12}
PANEL A. Equity returns														
Australian portfolios														
Portfolio #1	−0.0012	0.0746	−0.049	−0.055	−0.070	0.074	−0.019	−0.037	−0.144	0.244*	−0.154	0.075	−0.232*	0.004
Portfolio #2	−0.0015	0.0742	−0.195	0.034	−0.079	0.100	−0.079	−0.037	−0.098	0.190	−0.078	0.039	−0.159	0.011
Portfolio #3	−0.007	0.0794	−0.009	0.083	−0.117	0.026	0.048	−0.072	−0.039	0.058	−0.019	−0.003	−0.100	0.073
Portfolio #4	−0.0068	0.0827	−0.088	0.062	−0.003	0.052	−0.092	0.015	−0.136	0.154	−0.151	−0.021	−0.089	−0.002
Portfolio #5	−0.0062	0.0892	−0.085	0.109	−0.114	0.127	−0.128	−0.038	−0.173	0.235*	−0.208*	0.05	−0.117	0.008
Portfolio #6	−0.0154	0.1046	−0.051	0.198*	−0.053	0.169	−0.105	0.123	−0.134	0.152	−0.218*	0.025	−0.127	−0.027
Portfolio #7	0.0008	0.0831	−0.115	0.092	−0.013	0.034	−0.041	−0.025	−0.079	0.206*	−0.186	0.076	−0.08	0.011
Portfolio #8	−0.0094	0.1019	−0.068	0.215*	−0.198	0.112	−0.144	0.031	−0.15	0.244*	−0.083	0.144	−0.153	0.068
Canadian portfolios														
Portfolio #1	−0.0012	0.0555	0.071	0.055	−0.019	−0.079	0.002	−0.075	0.021	−0.003	−0.074	−0.07	−0.187	−0.141
Portfolio #2	0.0022	0.0473	−0.081	0.107	−0.079	0.030	−0.102	−0.019	−0.082	0.119	−0.21	−0.067	−0.04	0.048
Portfolio #3	−0.0024	0.0408	0.037	0.01	0.143	−0.121	−0.142	−0.02	0.028	0.022	0.052	−0.113	−0.049	0.033
Portfolio #4	0.0001	0.0626	−0.078	0.08	−0.046	−0.157	−0.146	0.019	−0.195	0.215*	−0.100	0.033	−0.005	0.113
Portfolio #5	−0.0037	0.0703	−0.101	0.076	−0.051	−0.050	−0.080	0.049	−0.04	0.047	−0.064	−0.149	−0.120	0.049
Portfolio #6	0.0072	0.0708	0.004	0.057	0.071	−0.006	−0.118	0.012	−0.017	0.086	−0.153	−0.003	−0.063	−0.106
Portfolio #7	0.0002	0.0710	−0.168	0.09	−0.194	−0.163	−0.073	−0.017	−0.04	0.107	0.017	−0.073	−0.129	0.053
Portfolio #8	−0.0044	0.0638	0.028	0.006	0.023	0.046	−0.082	0.082	−0.089	0.084	−0.243*	−0.107	−0.193	0.048
French portfolios														
Portfolio #1	0.0087	0.0688	0	0.013	−0.01	−0.013	−0.008	−0.103	−0.038	−0.155	−0.106	−0.023	0.033	0.141
Portfolio #2	0.0045	0.0705	0.039	−0.072	0.097	0.089	−0.071	0.088	0.041	0.031	−0.029	0.017	−0.04	0.040

Portfolio #3	0.0068	0.0594	0.008	0.123	0.002	0.059	−0.168	0.021	−0.174	−0.078	−0.181	0.003	−0.022	0.197*
Portfolio #4	0.0038	0.0614	0.112	0.038	0.006	0.013	−0.084	0.042	−0.019	0.013	−0.078	−0.035	−0.01	0.128
Portfolio #5	0.0009	0.0777	0.019	−0.027	0	0.034	0.024	−0.037	−0.059	−0.186	−0.096	−0.058	−0.026	0.113
Portfolio #6	0.0134	0.0814	0.001	−0.028	0.063	−0.019	−0.001	0.061	0.039	−0.122	−0.031	−0.019	0.011	0.180
Portfolio #7	0.0021	0.0752	−0.006	−0.022	0.001	−0.031	−0.081	−0.1	−0.056	−0.196	−0.097	0.047	0.024	0.114
Portfolio #8	0.005	0.0793	−0.069	0.009	0.061	−0.005	0.057	0.005	0.044	−0.107	−0.015	−0.011	−0.032	0.124
German portfolios														
Portfolio #1	0.0120	0.0505	0.065	0.213*	0.030	0.095	0.057	−0.128	−0.198*	−0.125	−0.080	−0.047	−0.091	0.023
Portfolio #2	0.0098	0.0564	0.031	0.069	−0.058	0.094	0.029	−0.097	−0.141	−0.081	−0.092	−0.018	−0.044	−0.037
Portfolio #3	0.0075	0.0501	0.154	0.050	−0.112	0.161	0.096	−0.061	−0.118	−0.084	−0.079	0.017	−0.011	−0.004
Portfolio #4	0.0097	0.0537	0.065	0.044	−0.209*	0.108	0.035	−0.016	−0.238*	0	−0.068	0.108	−0.071	−0.023
Portfolio #5	0.0114	0.0594	−0.078	0.086	−0.103	0.115	0.135	−0.158	−0.042	−0.113	0.019	−0.062	−0.034	0.001
Portfolio #6	0.0077	0.063	0.038	−0.059	−0.031	0.146	0.074	−0.135	−0.16	−0.036	−0.115	−0.088	0.016	0.030
Portfolio #7	0.0096	0.0596	−0.003	0.046	−0.076	0.140	0.148	−0.086	−0.03	−0.116	0.038	−0.014	0.064	−0.03
Portfolio #8	0.0118	0.0589	0.015	0.008	−0.007	0.114	0.086	−0.143	−0.064	−0.072	0.014	−0.051	−0.054	−0.087
Italian portfolios														
Portfolio #1	−0.0004	0.0746	−0.062	−0.044	0.227*	−0.092	0.014	0.075	−0.157	−0.182	0.077	−0.124	−0.160	0.087
Portfolio #2	−0.0014	0.0846	−0.046	−0.001	0.266*	−0.047	−0.073	0.135	−0.079	−0.149	0.065	−0.086	−0.164	0.105
Portfolio #3	0.0035	0.0635	0.098	0.037	0.122	−0.035	−0.005	0.097	−0.040	0.040	0.069	−0.164	−0.193	−0.024
Portfolio #4	0.0034	0.0808	−0.09	0.028	0.222*	−0.059	0.008	−0.003	0.063	−0.015	0.060	−0.143	−0.028	0.019
Portfolio #5	0.0037	0.0919	−0.011	0.065	0.168	−0.060	0.063	−0.038	−0.062	−0.013	0.020	−0.145	−0.068	0.025
Portfolio #6	0.0026	0.0892	−0.024	0.071	0.203*	−0.141	−0.014	0.060	−0.140	0.009	0.069	−0.184	−0.005	−0.018
Portfolio #7	0.0047	0.0824	−0.046	0.069	0.266*	−0.162	0.067	−0.064	−0.139	−0.064	−0.098	−0.052	−0.054	−0.059
Portfolio #8	−0.0009	0.0759	0.011	0.047	0.234*	−0.004	−0.002	0.106	−0.036	−0.12	0.078	−0.069	−0.052	0.046
Japanese portfolios														
Portfolio #1	0.0055	0.0676	0.001	−0.003	−0.112	0.050	0.067	0.020	0.074	0.014	0.064	0.047	0.002	0.047
Portfolio #2	0.0073	0.0674	0.005	0.033	−0.121	−0.051	−0.015	0.020	−0.028	−0.102	0.066	−0.008	−0.024	0.093
Portfolio #3	0.0086	0.0669	−0.100	0.021	−0.089	0.097	−0.056	0.044	0.070	0.045	0.009	0.100	−0.055	−0.027

(continued on next page)

Table 3 (*continued*)

	Mean	S.D.	Autocorrelations ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_7	ρ_8	ρ_9	ρ_{10}	ρ_{11}	ρ_{12}
Portfolio #4	0.0040	0.0600	−0.108	0.033	0.003	0.051	−0.002	0.003	0.003	−0.008	0.013	0.031	0.009	0.010
Portfolio #5	0.0149	0.0647	−0.137	0.044	−0.084	−0.111	−0.02	0.053	0.004	−0.192	0.064	−0.040	0.001	0.092
Portfolio #6	0.0087	0.0710	−0.058	−0.001	−0.138	0.043	−0.038	0.078	0.005	−0.052	0.078	−0.069	−0.029	0.029
Portfolio #7	0.0150	0.0715	−0.191	0.023	−0.144	0.046	−0.002	−0.015	−0.029	−0.077	−0.064	0.060	0.059	0.036
Portfolio #8	0.0095	0.0673	−0.146	0.038	−0.185	−0.028	0.013	0.094	0.045	−0.150	0.010	−0.012	−0.066	0.103
Dutch portfolios														
Portfolio #1	0.0140	0.0641	−0.145	0.005	0.045	−0.040	0.159	−0.055	−0.001	0.049	−0.187	0.101	0.064	0.004
Portfolio #2	0.0060	0.0524	−0.008	−0.055	0.02	−0.025	0.181	0.073	−0.143	−0.094	−0.108	−0.002	0.169	−0.044
Portfolio #3	0.0050	0.0605	−0.009	−0.028	−0.084	0.085	0.240*	−0.003	−0.081	−0.162	−0.127	0.040	−0.071	0.033
Portfolio #4	0.0088	0.0592	−0.015	0.039	−0.044	0.066	0.176	0.099	−0.069	0.016	−0.081	0.105	0.111	0.017
Portfolio #5	0.0052	0.0626	−0.061	−0.009	−0.019	−0.047	0.077	−0.027	−0.116	−0.074	−0.123	−0.069	0.009	−0.041
Portfolio #6	0.0042	0.0585	0.068	−0.146	−0.069	−0.086	0.148	0.029	−0.134	−0.106	−0.218*	−0.027	0.109	0.004
Portfolio #7	0.0028	0.0668	0.067	0.001	0.01	0.047	0.106	−0.018	−0.068	−0.142	−0.143	−0.073	−0.101	−0.124
Portfolio #8	0.0081	0.0557	−0.093	−0.119	−0.02	−0.015	0.185	−0.034	−0.002	−0.183	−0.113	−0.021	0.004	0.101
Swiss portfolios														
Portfolio #1	0.0107	0.0484	−0.015	0.0115	−0.132	0.072	−0.014	0.084	−0.011	−0.037	0.000	−0.044	−0.033	−0.017
Portfolio #2	0.0084	0.0488	0.043	0.076	−0.067	0.158	0.041	0.061	−0.072	−0.090	−0.060	−0.065	0.020	−0.044
Portfolio #3	0.0050	0.0473	0.113	0.083	−0.084	0.102	0.034	0.137	0.094	0.017	0.034	−0.070	−0.046	−0.082
Portfolio #4	0.0085	0.0447	0.108	0.142	−0.139	0.111	0.072	0.016	0.048	−0.124	0.002	−0.037	0.071	−0.073
Portfolio #5	0.0070	0.0573	−0.011	0.061	−0.023	0.100	0.054	0.014	0.028	−0.018	−0.063	0.019	−0.033	0.027
Portfolio #6	0.0095	0.0643	−0.017	0.023	−0.043	0.064	0.081	0.001	−0.069	−0.041	−0.119	−0.110	0.077	−0.081
Portfolio #7	0.0022	0.0565	0.045	−0.042	−0.16	0.067	0.018	0.133	−0.117	0.007	−0.176	−0.096	0.03	0.031
Portfolio #8	0.0099	0.0591	0.003	0.157	−0.112	0.199*	0.071	0.090	−0.014	−0.061	−0.133	−0.052	0.08	−0.081

UK portfolios[a]														
Portfolio #1	0.0041	0.0666	−0.008	−0.032	0.086	−0.047	0.123	0.012	−0.056	0.047	−0.059	0.015	0.042	0.009
Portfolio #2	0.0038	0.0675	0.018	−0.069	0.045	−0.039	0.109	−0.002	−0.089	0.042	−0.079	0.001	0.038	0.041
Portfolio #3	0.0036	0.0536	0.123	−0.002	0.068	−0.028	0.124	−0.026	−0.062	0.068	−0.056	0.039	0.021	0.077
Portfolio #4	0.0008	0.0582	0.062	0.006	0.058	−0.042	0.142	0.002	−0.057	0.032	−0.100	0.028	0.049	0.062
Portfolio #5	0.0045	0.0779	−0.043	0.045	0.001	−0.079	0.124	0.009	−0.106	0.086	−0.104	0.020	0.026	0.028
Portfolio #6	0.0027	0.0806	−0.094	−0.014	0.036	−0.087	0.114	0.027	−0.129	0.096	−0.074	0.010	0.045	−0.001
Portfolio #7	0.0018	0.0706	−0.015	−0.041	0.059	−0.078	0.089	0.048	−0.084	0.036	−0.111	−0.001	0.039	0.030
Portfolio #8	0.0038	0.0713	−0.053	−0.046	0.035	−0.092	0.059	−0.006	−0.067	0.070	−0.144	0.053	0.002	−0.021
US portfolios														
Portfolio #1	−0.0026	0.0493	0.021	0.084	−0.008	−0.059	0.043	0.132	0.025	0.067	−0.130	0.007	−0.031	0.113
Portfolio #2	−0.0018	0.0461	−0.058	−0.007	−0.016	−0.019	−0.025	0.000	−0.038	−0.028	−0.026	−0.032	0.113	−0.009
Portfolio #3	0.0018	0.0439	−0.030	0.104	−0.015	0.024	−0.056	0.094	−0.042	0.068	−0.093	0.019	0.025	0.073
Portfolio #4	0.0008	0.0471	−0.085	0.087	−0.091	0.031	−0.043	0.223*	−0.189	0.126	−0.082	−0.014	−0.048	0.202*
Portfolio #5	0.0002	0.0694	−0.14	0.118	−0.021	−0.031	−0.086	0.131	−0.121	0.094	−0.062	−0.038	−0.137	0.022
Portfolio #6	0.0023	0.063	−0.059	0.132	0.018	−0.025	−0.072	0.212*	−0.034	0.093	−0.086	−0.036	−0.112	0.045
Portfolio #7	0.0018	0.063	−0.117	0.033	−0.036	−0.054	−0.087	0.234*	−0.024	0.101	−0.134	−0.028	−0.078	0.053
Portfolio #8	0.001	0.0704	−0.005	0.188	0.085	0.042	0.099	0.219*	−0.065	0.060	−0.071	−0.064	−0.021	−0.123
MSCI indexes														
Australia	0.0018	0.0804	−0.176	0.067	−0.073	0.043	−0.093	−0.009	−0.098	0.251*	−0.207*	0.105	−0.187	0.000
Canada	0.0011	0.0574	−0.027	0.023	0.009	−0.146	−0.146	0.011	−0.106	0.125	−0.084	−0.072	−0.152	−0.002
France	0.0099	0.0685	−0.005	−0.03	0.045	0.002	−0.025	−0.072	−0.043	−0.059	−0.038	−0.046	0.034	0.173
Germany	0.0137	0.0638	−0.026	0.036	−0.048	0.127	0.173	−0.162	−0.039	−0.053	0.057	−0.058	−0.018	−0.013
Italy	0.0036	0.0773	−0.019	0.005	0.284*	−0.011	−0.062	0.012	−0.071	−0.039	0.003	−0.138	−0.114	0.047
Japan	0.0097	0.0613	−0.120	0.148	−0.08	0.1444	0.052	0.116	0.044	−0.136	0.085	0.005	0.018	0.033
Netherlands	0.0098	0.0519	−0.022	−0.115	−0.009	−0.078	0.027	−0.087	0.023	−0.061	0.009	0.005	−0.009	0.000

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Table 3 (*continued*)

	Mean	S.D.	Autocorrelations											
			ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_7	ρ_8	ρ_9	ρ_{10}	ρ_{11}	ρ_{12}
Switzerland	0.0103	0.0547	−0.044	0.032	−0.032	0.097	0.016	−0.018	−0.085	−0.024	−0.012	−0.129	−0.096	0.016
UK	0.0106	0.0717	−0.058	−0.038	−0.028	−0.106	0.045	−0.004	−0.071	0.077	−0.091	−0.004	0.053	−0.008
USA	0.0033	0.0487	−0.145	0.075	−0.006	−0.033	0.041	0.027	−0.124	0.054	−0.064	−0.049	−0.051	0.024
World index	0.0053	0.0428	−0.134	0.126	−0.004	0.013	−0.01	0.057	−0.045	−0.039	−0.032	−0.029	−0.028	0.051
PANEL B. Other variables														
US T-bill	0.0065	0.0021	0.805*	0.707*	0.687*	0.689*	0.606*	0.532*	0.474*	0.451*	0.401*	0.294*	0.242*	0.270*
Unexpected inflation														
USA	−0.0005	0.0029	0.159	−0.065	0.014	−0.086	0.229*	0.226*	0.071	0.074	−0.229*	−0.051	0.027	−0.004
World	0.0009	0.0427	−0.011	−0.044	−0.161	0.096	−0.008	−0.063	0.043	0	0.17	0.064	−0.031	−0.081
Foreign exchange common and residual component indexes in US dollars														
Common	−0.0065	0.0233	0.151	0.098	−0.099	0.119	0.066	−0.007	0.077	0.008	0.149	0.097	0.007	−0.05
Idiosyncratic	−0.0065	0.0038	0.283*	0.193	0.235*	0.172	0.283*	0.104	−0.044	0.101	0.021	0.048	−0.116	−0.022

[a] The statistics are based on the even monthly observations from 1973:2–1990:12 (108 observations). All variables are in US dollars in excess of the holding period return on the US 30-day Treasury Bill. Portfolio 1 for a given country represents the portfolio with the lowest betas against the world market portfolio, the equally weighted index of all changes in exchange rates, and the GDP weighted index of world unexpected inflation. Portfolio 8 for a given country represents the portfolio with the highest betas against all three variables. Since only even observations are used in the tests, the autocorrelation ρ_2, for example, refers to the autocovariance of the current even month return (at time *t*) with the lagged by two even months return, divided by the variance computed from all even observations.

*Significant at the 5% level based on an approximate standard error of $1/\sqrt{108}=0.0962$.

world unexpected inflation. Portfolio 8 for a given country represents the portfolio with the highest betas against all three variables. The portfolio construction methodology is described in detail in the following section.

3.2. Portfolio construction

The AD, S-S, and GLS models imply that the independent variables against which we need to gain dispersion are the return to the world market portfolio, the inflation series, and the changes in exchange rates. Given the data transformation performed in Sections 2.2.1 and 2.2.2, and the structure imposed on the models as stated in relations (17), (18), and (19), we choose as instrumental variables for the classification of stocks into portfolios the world market portfolio beta, the beta of the world inflation index (not orthogonal to the US inflation), and the beta with respect to an equally weighted index of changes in all exchange rates.

The two inflation variables in (18) are orthogonal to each other, and therefore, classifying securities according to the beta coefficient of the world inflation index (not orthogonal to US inflation) offers dispersion against both variables. Furthermore, recall that the correlation between the equally weighted and residual indexes is 0.355, while that of the equally weighted and common component indexes is equal to 0.991. Therefore, a classification of securities using the equally weighted index will offer dispersion against both the common component and the residual component indexes of exchange rates. Evidently, the dispersion gained against the common component index will be higher than that against the residual index, but again the variation of changes in exchange rates explained by the common component index is also considerably higher.

To avoid problems related to selection bias, the estimation of these beta coefficients should be independent of the beta estimates obtained in our tests. Since our sample spans the period from January 1973 to December 1990, each security has a total of 216 monthly observations. To construct the portfolios we use Chen's (1983) methodology, and therefore, we separate the observations into two groups of odd and even months. We use odd observations to estimate betas and even ones to calculate the returns of the portfolios, and estimate the models.

The advantage of this approach relative to the classic Fama–MacBeth (1973) procedure is that it allows us to use observations from the whole time period covered by our sample. In the Fama–MacBeth procedure, the first five years of data are used only to classify stocks into portfolios. A further advantage of the Chen procedure is that it assumes stability only between betas estimated using odd and even observations, a rather weak assumption. In contrast, Fama–MacBeth assume stability of betas across time which is a stronger assumption. However, a disadvantage of Chen's portfolio construction approach is that the total number of observations used is smaller than the number of observations we would have used with the Fama–MacBeth approach.

There are six possible ways in which we can classify securities into portfolios given that we want to obtain dispersion against three variables. However, our aim is to maximize dispersion against the exchange rate and inflation betas since the focus of our paper is the pricing of exchange rate and inflation risk premiums, and exchange rate and inflation betas tend to be more noisy than world market betas. Given these considerations, we choose to first classify securities according to their world betas into two portfolios, then we subdivide each portfolio into two portfolios according to the exchange rate betas, and finally, all portfolios are split into two according to the world inflation betas. This classification gives maximum dispersion against inflation betas and least dispersion against world market betas. Since we do not estimate world market risk premiums in this study, the limited dispersion gained against world market betas should not be problematic. In addition, we will show in Section 4.4.3 that, although the dispersion of world market betas across portfolios is smaller than that of the exchange rate and inflation betas, the standard errors of the world market betas are substantially smaller than the standard errors of exchange rate and inflation betas. This justifies our choice to aim for larger dispersion of exchange rate and inflation betas across portfolios. Finally, it will be shown that the dispersion of exchange rate betas and their standard errors are comparable to those of inflation betas, although our portfolio construction approach does not maximize dispersion against exchange rate betas. Again, this piece of evidence renders support to our choice to aim for maximum dispersion against inflation betas.

To avoid mis-specification biases, the estimation of betas was carried out according to the implications of the three theoretical models. In particular, for each set of betas, we chose to estimate them according to the most general specification in which they appear. World and inflation betas were estimated jointly, as specified in the AD model. Exchange rate betas were estimated together with the world betas, as it is implied by the S-S model. However, from these estimations we retained only the exchange rate betas. Eight portfolios were formed for each country, i.e. a total of 80 portfolios for all 10 countries.

The above procedure was repeated for the two subperiods of 108 total observations. The portfolio returns for the entire period are obtained by appending the portfolio returns of the first subperiod to those of the second subperiod. This is done in order to account for possible non-stationarities in betas, and it is equivalent to updating the membership of securities in the eight portfolios twice during the entire period.

4. Empirical results

4.1. Country-specific exchange rate risk premiums

Table 4 presents the results from the estimation of the S-S model. They are obtained by estimating betas, intercepts, and risk premiums for all countries simultaneously. We observe that at least one of the exchange rate indexes is priced in six out of the 10 countries in our sample. The common component index is priced in Germany, Japan, Switzerland, and Netherlands. Note that in Germany, Japan, and Switzerland, the risk premium is negative and varies between −0.15% per month for Germany and −0.7% for Japan. This means that hedging the common component

Table 4
Estimation of exchange rate risk premiums[a]

Country/coefficient	γ_0	γ^λ	γ^e	ME*	RMSE†	R^2‡
Australia	0.0337 (0.36)	0.0101 (0.49)	−0.0011 (−0.05)	0.0008	0.0691	34.20
Canada	0.0020 (0.61)	−0.0092 (−0.53)	0.0037 (2.36)	−0.0020	0.0493	31.70
France	0.0289 (0.99)	−0.0317 (−0.84)	−0.0020 (−0.16)	0.0003	0.0545	40.45
Germany	0.0123 (3.35)	−0.0015 (−2.26)	0.0051 (2.67)	0.0003	0.0423	45.92
Japan	0.0118 (2.11)	−0.0072 (−2.67)	0.0092 (4.63)	0.0012	0.0548	34.24
Switzerland	0.0647 (0.04)	−0.0051 (−2.46)	−0.0204 (−0.40)	−0.0023	0.0363	54.00
Italy	0.0393 (0.14)	−0.0834 (−0.12)	−0.0288 (−0.10)	0.0002	0.0745	11.04
Netherlands	0.0025 (1.03)	0.0118 (2.96)	0.0079 (6.72)	0.0023	0.0476	36.10
UK	0.0450 (0.23)	−0.0277 (−0.15)	−0.0353 (−0.20)	0.0018	0.0512	42.73
USA	−0.0013 (−0.60)	0.0043 (0.62)	0.0042 (2.53)	−0.0008	0.0379	68.26
Hansen's J_T test§: $\chi^2(50)=54.51$; P-value=0.30						

[a] The following model is estimated using an iterated generalized methods of moments (GMM) procedure, and employing the Newey–West estimator (the truncation parameter q was set equal to 6):

$$R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^e\beta_k^e+\beta_k^e e_t+\gamma^\lambda\beta_k^\lambda+\beta_k^\lambda\lambda_t+\zeta_{kt}$$

where R_k is the return of equity portfolio k. Eight portfolios are constructed for each of the 10 countries, i.e. 80 portfolios in total. The variables e and λ denote the return on the residual and common component exchange rate indexes. All returns are calculated in US dollars in excess of the holding period return on the Treasury Bill which is closest to 30 days to maturity. The data span the period from 1973:01–1990:12. The coefficients γ^e and γ^λ denote the risk premiums with respect to the residual and common components of changes in the nine exchange rates against the US dollar. The coefficients β_k^w, β_k^e and β_k^λ denote the betas of portfolio k with the world market portfolio, the residual, and common component indexes of changes in exchange rates, respectively. Their estimates are not reported. The coefficient γ_0 should be equal to zero if the model is true. As a proxy for the return on the world market portfolio, R^w, we use the return on the MSCI world index. Country-specific risk premiums in the 80 equations system are estimated simultaneously for all countries. We use only even observations for the estimation of the model, i.e. 108 observations in total, since odd observations were previously utilized for the construction of portfolios. Beta and gamma coefficients are estimated simultaneously. Corrected t-values for heteroskedasticity and serial correlation appear in parentheses beside the coefficient estimates.

*Mean error from model (17) across portfolios.

†Root mean square error from model (17) across portfolios.

‡Adjusted coefficient of determination, calculated in each case using the sum of square residuals from the relevant portfolio returns.

§Hansen's GMM test.

exchange rate exposure during the period studied would have resulted into an increase in the return of equity portfolios in these countries, since hedging the exposure to the common component index would decrease or eliminate its exchange rate risk premium. This would not have been the case in the Netherlands where the common component risk premium is positive. Similarly to the results of Jorion (1991), the common component index which is virtually identical to a simple index of all exchange rates, is not priced in the USA. Note, however, that the residual component index carries a risk premium of 0.42% per month in the US which is statistically significant at the 5% level. In addition, the residual component is also priced in Canada, Germany, Japan, and the Netherlands. The magnitude of its risk premium in these countries varies from 0.365% in Canada to 0.921% in Japan.

The results with respect to the pricing of the residual component index are important because they affect significantly our conclusions regarding the pricing of exchange rate risk, and therefore, the effect that exchange rate hedging may have on the return of equity portfolios. In the absence of the residual component index from our tests we would have incorrectly concluded that exchange rate risk is not priced in Canada and the US, and consequently, hedging exchange rate risk would only reduce the volatility of equity portfolios in these countries, leaving their return unchanged. By the same token, we would have concluded that hedging exchange rate risk in Germany and Japan would result in a bigger improvement of the performance of their equity portfolios than it appears to be the case in the presence of the residual exchange rate component.

If the S-S model is the correct model, the intercept γ_0 must be equal to zero. We see that γ_0 is statistically different from zero at the 5% level only for Germany and Japan which means that the model performs relatively well in the other countries.

To further evaluate the ability of the S-S model to explain equity returns in the 10 countries we use a series of criteria. We first compute the mean pricing error (ME) across the eight portfolios of each country. We find that the ME is positive in seven countries but negative in the remaining three. A positive ME indicates that the model tends to overestimate expected returns of equity portfolios. The absolute magnitude of the ME varies from 2.7% ($0.023\times12\times100$) per annum (pa) for Switzerland and the Netherlands, to 0.9% pa for Australia and the US. Therefore, although the model tends to misprice equities, the absolute magnitude of the mispricing is relatively small.

We also compute the standard deviation of the pricing error which is given by the root mean square error (RMSE). To interpret these numbers, we need to know the average standard deviation of the portfolios. From Table 3 we know that, for instance, the average standard deviation across the Italian portfolios is 0.08 or 27.7% pa. The RMSE across Italian portfolios in Table 4 is 0.074 or 25.63% pa. This means that the model reduces the average standard deviation of the Italian portfolios by 2.13%. This order of standard deviation reduction is among the smallest observed across countries. One of the largest reductions is found in the US portfolios where the standard deviation is reduced by 6.4%. Next in line are the UK with a reduction of 5.9%, Germany with 4.9%, and Japan with 3.84%. The above figures reveal that a substantial part of the variation in returns remains unexplained. This is not necessarily worrisome, since the ME are small.

The adjusted *R*-square provides an additional intuitive measure for the performance of the model. It is calculated by adding up the sum of square residuals from the eight equations of each country's portfolios. It appears that the S-S model performs best in the US (68.26%) and worst in Italy (11.04%). This is of course consistent with the results of the RMSE.

We compute Hansen's *J*-statistic which follows a $\chi^2(50)$=54.51. The *P*-value of the test is equal to 0.30 which suggests that the model cannot be rejected. This is not surprising, given that the model performs well in most countries in our sample.

4.2. Country-specific US inflation risk premiums

Table 5 reports the results from the estimation of the GLS model. Similar to the case of the S-S model, intercepts, betas, and risk premiums were estimated for all countries simultaneously. Interestingly, US unanticipated inflation is priced at the 5% level not only in the US portfolios but also in all others, apart from the UK and Swiss portfolios where it is priced at the 10% level. This is contrary to the general belief that only domestic inflation is priced in a country's equities, and suggests that at least US inflation can be hedged using not only US equities but also equities from other countries. In that sense, home bias in US portfolios cannot be due to investors' demands for hedging domestic inflation.

Note that the US inflation risk premium attached to Canadian and US equities is of almost identical magnitude. This is an interesting but also intuitive result since, in general, Canada closely follows the US monetary policy and maintains a similar level of inflation to that of the US. Notice also that the US inflation risk premium attached to foreign equities is larger than that found in US (and Canadian) equities. This may be the effect of expectations of inflation transmissions from the US to the rest of the countries, although our tests are not geared towards providing a definite explanation for this result. The US inflation risk premium is always positive and its magnitude is of a similar order to that of the exchange rate risk premiums.

The country intercepts indicate that the GLS model performs well in all countries except in Germany and the Netherlands where positive and statistically significant intercepts were found.

The ME are again positive in seven countries, but their absolute magnitudes are larger than in the case of the S-S model. This suggests that the S-S model may price equities more accurately than the GLS model. The largest mispricings are found in Australia and the Netherlands where the GLS model underprices Australian equity by 4.7% on average, and overprices Dutch equity by 3.69%. Note, however, that the US securities are underpriced by only 0.1% pa which is a quite smaller mispricing than the one produced by the S-S model. This result does not mean that the GLS model is better in pricing US securities. The RMSE of the GLS model are larger than those of the S-S model for all countries. In other words, although the GLS model produces a smaller pricing error in the case of US securities, it also leaves a larger proportion of the standard deviation of returns unexplained.

Table 5
Estimation of US inflation risk premium[a]

Country	γ_0	γ^i	ME*	RMSE†	R^2‡
Australia	0.0033 (−0.96)	0.0113 (5.06)	−0.0039	0.0717	30.36
Canada	−0.0002 (−0.09)	0.0040 (3.49)	−0.0026	0.0502	30.76
France	0.0035 (1.31)	0.0094 (8.85)	0.0002	0.0569	36.34
Germany	0.0053 (2.84)	0.0059 (5.47)	0.0016	0.0492	29.03
Japan	0.032 (1.50)	0.0065 (5.88)	0.0011	0.0589	25.49
Switzerland	0.0090 (1.11)	0.0201 (1.68)	0.0003	0.0435	37.00
Italy	0.0017 (0.58)	0.0061 (8.85)	0.0009	0.0753	11.10
Netherlands	0.0067 (3.33)	0.0082 (8.00)	0.0031	0.0523	25.93
UK	−0.0066 (−0.90)	0.0229 (1.71)	0.0015	0.0533	38.96
USA	−0.0022 (−1.69)	0.0048 (4.77)	−0.0001	0.0418	64.76
Hansen's J_T-test§: $\chi^2(60)$=43.9; P-value=0.94					

[a] The following model is estimated using an iterated generalized methods of moments (GMM) procedure, and employing the Newey–West estimator (the truncation parameter q was set equal to 6):

$$R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^i\beta_k^i+\beta_k^i i_t+\psi_{kt}$$

where R_k is the return of equity portfolio k. Eight portfolios are constructed for each of the 10 countries, i.e. 80 portfolios in total. The variable i denotes the innovations in US inflation rate. All variables are in US dollars in excess of the holding period return on the Treasury Bill which is closest to 30 days to maturity. The data span the period from 1973:01–1990:12. The coefficient γ^i denotes the US unanticipated inflation risk premium. The coefficients β_k^w and β_k^i denote the betas of portfolio k with the world market portfolio, and the US unanticipated inflation. Beta estimates are not reported in this Table. The coefficient γ_0 should be equal to zero if the model is true. As a proxy for the return on the world market portfolio, R^w, we use the return on the MSCI world index. Country-specific risk premiums in the 80 equations system are estimated for all countries simultaneously. We used only even observations for the estimation of the model, since the odd observations were previously used for the construction of portfolios. Beta and gamma coefficients are estimated simultaneously. Corrected t-values for heteroskedasticity and serial correlation appear in parentheses beside the coefficient estimates.
*Mean error from model (19) across portfolios.
†Root mean square error from model (19) across portfolios.
‡Adjusted coefficient of determination, calculated in each case using the sum of square residuals from the relevant portfolio returns.
§ Hansen's GMM test.

As would be expected, the adjusted R-squares are also smaller in all cases relative to those computed for the S-S model. The reduction is not dramatic, however. The GLS model can explain 65% of the variation of returns in the US, and 11% in Italy. The fact that the model does not perform poorly in absolute terms is confirmed by Hansen's *J*-test. The *P*-value implies that the model cannot be rejected.

4.3. *Joint estimation of country-specific exchange rate and inflation risk premiums*

Since the tests of the S-S and GLS models suggest that both exchange rate and US inflation uncertainty is priced, it is worthwhile to examine to what extent the estimated exchange rate risk premiums proxy for inflation risk premiums and vice versa. The relative importance of exchange rate and foreign inflation risk premiums in equities is tested using model (18) which is our empirical version of the AD model.

Table 6 presents the results. They are obtained once again by estimating all coefficients for all countries simultaneously. We observe that the exchange rate and US inflation risk premiums are markedly different in terms of value from those reported in Tables 4 and 5. In addition, in most cases, they are not statistically significant at the 5% level. This is due to multicollinearity problems arising from the simultaneous inclusion of exchange rate and inflation variables in the model. In particular, the residual world inflation index has correlations of 0.34 and 0.96 with the residual and common component exchange rate indexes, respectively. Note, however, that the correlations of US unanticipated inflation with the residual and common component factors are only 0.05 and −0.10, respectively.

A comparison of the ME and RMSE with those reported in Tables 4 and 5 is more revealing. The ME in five countries are now negative. In terms of absolute magnitude, they are lower in Canada, Germany, Italy, the Netherlands, and the UK, relative to those from the previous models, but higher in the other countries. The lowest ME are found in Japan and the UK and have an absolute magnitude of approximately 0.06% pa. The highest ME are in France and the Netherlands. French assets are underpriced by 1.98% pa, and Dutch assets are overpriced by 1.02% pa. The RMSE are somewhat smaller and the reduction is of the order of 0.17% pa.

The adjusted R-squares are larger than in the previous models in eight countries with the increase being more pronounced in Canada, Switzerland, and to a lesser extent in Germany and the United States. A slight decrease in the adjusted R-squares is observed in the Netherlands, and the UK. As expected, Hansen's test cannot reject the model.

Overall, it appears that a model which includes both exchange rate and inflation risk factors can price international equities better than the restricted models examined. Explicit hypotheses tests on the relative performance of the three models are provided in the Section 4.4.2.

4.4. Diagnostics

4.4.1. The equality of risk premiums across countries

Recall that the S-S and AD models assume that PPP does not hold. When PPP does not hold, capital markets may be partially segmented, and therefore, investors in different countries may require different risk premiums for bearing the same risk. Capital markets can exhibit partial segmentation, that is, segmentation which does not arise from the presence of explicit frictions in the markets, such as restrictions to ownership, when certain sources of risk are not perfectly hedgeable. Note that in the case of the S-S model, capital markets *are not* partially segmented because exchange rate risk can be perfectly hedged.[12] In that sense, exchange rate risk pre-

[12] For a discussion of this point, see Section VII in the Adler and Dumas (1983) paper, and their footnote 86, in particular.

Table 6
Joint estimation of exchange rate and inflation risk premiums[a]

Country	γ_0	γ^{λ}	γ^{e}	γ^{i}	γ^{v}	ME*	RMSE†	R^2‡
Australia	−0.0169 (−0.14)	0.0032 (0.70)	0.0112 (0.25)	0.0124 (0.66)	0.0026 (0.12)	0.0007	0.0679	34.64
Canada	0.0096 (0.21)	0.0270 (1.26)	0.0021 (1.11)	0.0023 (1.12)	−0.0070 (−0.20)	−0.0007	0.0490	47.94
France	0.0027 (0.25)	0.0038 (1.47)	0.0033 (0.39)	0.0004 (0.40)	0.0076 (0.35)	−0.0017	0.0538	42.94
Germany	0.0162 (1.12)	−0.0013 (−0.13)	0.0091 (2.54)	0.0056 (1.09)	0.0131 (0.39)	−0.0003	0.0414	51.55
Japan	0.0110 (0.18)	−0.0350 (−0.80)	0.0220 (0.23)	−0.0011 (−0.10)	0.0030 (1.04)	−0.0001	0.0544	36.70
Switzerland	0.0332 (0.24)	0.0015 (0.89)	0.0189 (2.26)	0.0493 (1.16)	0.0014 (1.09)	0.0007	0.0356	61.87
Italy	−0.0027 (−0.42)	0.0055 (1.15)	0.0063 (1.99)	0.0022 (0.28)	0.0076 (1.29)	−0.0002	0.0740	14.06
Netherlands	0.0511 (0.15)	−0.0099 (−0.70)	0.0119 (0.34)	0.0419 (1.61)	−0.0025 (−0.55)	0.0009	0.0471	35.76
UK	0.0191 (0.19)	−0.0046 (−0.19)	−0.0063 (−0.60)	0.0010 (0.32)	−0.0080 (−0.71)	0.0000	0.0510	41.91
USA	0.0036 (0.09)	0.0118 (0.14)	0.0088 (0.11)	0.0002 (1.96)	0.0070 (0.60)	0.0001	0.0374	74.29
Hansen's J_T-test§: $\chi^2(30)=35.65$; $P-$value=0.22								

[a] The following model is estimated using an iterated generalized methods of moments (GMM) procedure, and employing the Newey–West estimator (the truncation parameter q was set equal to 6):

$$R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^e\beta_k^e+\beta_k^e e_t+\gamma^\lambda\beta_k^\lambda+\beta_k^\lambda\lambda_t+\gamma^i\beta_k^i+\beta_k^i i_t+\gamma^v\beta_k^v+\beta_k^v v_t+\xi_{kt}$$

where R_k is the return of equity portfolio k. Eight portfolios are constructed for each of the 10 countries, i.e. 80 portfolios in total. The variables e and λ denote the return on the residual and common component exchange rate indexes. The variable i denotes the innovations in US inflation rate, whereas the variable v the innovations in residual world inflation orthogonal to US unanticipated inflation. All variables are in US dollars in excess of the holding period return on the Treasury Bill which is closest to 30 days to maturity. The data span the period from 1973:01–1990:12. The coefficients γ^e, γ^λ denote risk premiums with respect to the residual and common components of changes in the exchange rates, and γ^i, γ^v denote risk premiums with respect to the US unanticipated inflation rate, and the residual world unanticipated inflation rate. The coefficients $\beta_k^w,\beta_k^e,\beta_k^\lambda,\beta_k^i$ and β_k^v denote betas of portfolio k with the world market portfolio, the idiosyncratic and common component exchange rate indexes, the US unanticipated inflation, and the residual world unanticipated inflation, respectively. The beta estimates are not reported here. If the model is true, then γ_0=0. As a proxy for the return on the world market portfolio, R^w, we use the return on the MSCI world index. Country-specific risk premiums in the 80 equations system are estimated simultaneously. Only even observations for the period 1973:01–1990:12 are used, i.e. 108 observations in total, since odd observations were previously used for the construction of portfolios. Corrected t-values for heteroskedasticity and serial correlation appear in parentheses beside the coefficient estimates.

*Mean error from model (18) across portfolios.

†Root mean square error from model (18) across portfolios.

‡Adjusted coefficient of determination, calculated in each case using the sum of square residuals from the relevant portfolio returns.

§Hansen's GMM test.

miums in the S-S model do not need to vary across countries. In contrast to the S-S and AD models, the GLS model assumes that PPP holds, and therefore, it implies that capital markets are perfectly integrated and that risk premiums are equal across countries.

The hypothesis of equality of risk premiums across countries is formally tested in this section by computing the Newey–West *D*-statistic. This involves two steps. We first estimate each model allowing the risk premiums to vary across countries, and save the final weighting matrix. We then use this weighting matrix to re-estimate the model under the restriction of equality of risk premiums across countries. The difference of the minimized objective functions from the two estimations is χ^2 distributed with degrees of freedom equal to the number of restrictions that the restricted model imposes on the unrestricted one.

The results from the above test are reported in Table 7, Panel A. For all three models, the *P*-value of the Newey–West test is small enough to lead us to reject the hypothesis of equality of risk premiums across countries at any conventional level of significance. Testing for the hypothesis of equality of risk premiums across coun-

Table 7
Diagnostics[a]

	GLS model	S-S model	AD model
Panel A: Equality of risk premiums across countries			
Unrestricted model	$\chi^2(50)=54.51$ *P*-value: 0.30	$\chi^2(60)=43.9$ *P*-value: 0.94	$\chi^2(30)=35.65$ *P*-value: 0.22
Restricted model	$\chi^2(77)=105.68$ *P*-value: 0.0167	$\chi^2(78)=92.81$ *P*-value: 0.12	$\chi^2(75)=88.67$ *P*-value: 0.13
Newey–West *D*-test	$\chi^2(27)=51.17$ *P*-value: 0.003	$\chi^2(18)=48.91$ *P*-value: 0.0001	$\chi^2(25)=53.02$ *P*-value: 0.0009
Panel B: Relative performance of the three models			
Unrestricted model			$\chi^2(30)=35.65$ *P*-value: 0.22
Restricted model	$\chi^2(60)=79.75$ *P*-value: 0.045	$\chi^2(50)=71.43$ *P*-value: 0.02	
Newey–West *D*-test	$\chi^2(30)=44.10$ *P*-value: 0.047	$\chi^2(3)=35.78$ *P*-value: 0.016	

[a] Panel A tests the hypothesis of equality of risk premiums across countries using the three alternative models. For each model, the row labeled “unrestricted model” reports the results from Hansen’s *J*-test from estimation of the model with country-specific risk premiums. The row labeled “restricted model” reports the results from estimating the model under the restriction of equality of risk premiums across countries. This estimation uses the final weighting matrix from the estimation of the unrestricted model. The row labeled “Newey–West *D*-test” reports the chi-square test for the difference of the minimized objective function from the two estimations. Panel B tests the relative performance of the three models. The unrestricted model in this case is the AD model with country-specific risk premiums. The unrestricted models are the GLS and S-S models with country-specific risk premiums. All tests are performed using the 80 equations systems.

tries is one of the two possible ways to discriminate empirically among the three models. Our result suggests that world capital markets are less than perfectly integrated, and leads to the rejection of the S-S and GLS models. The second way to discriminate empirically among the three models is presented below.

4.4.2. The relative performance of the three models

Panel B of Table 7 reports the results from Newey–West *D*-tests on the relative performance of the three models when they are used to estimate country-specific risk premiums. It is important to compare the performance of the three models allowing the risk premiums to vary across countries since the evidence in Section 4.4.1 suggests that capital markets may be partially segmented. To compute the statistic in this case, we re-estimate the S-S and GLS models using the weighting matrix from the estimation of the AD model. This is possible, since models S-S and GLS models are nested with the AD model. The computed *D*-statistic rejects at the 5% level the GLS and the S-S models in favor of the AD model. This implies that although exchange rate and inflation risk factors are correlated, their simultaneous presence in the AD model improves in statistical terms the performance of the restricted models. In other words, the inflation risk factors contain, on average across countries, useful information for the pricing of equities beyond the information contained in the exchange rate risk factors.

4.4.3. The beta coefficients

Figs. 1–5 provide a graphical representation of the statistical properties of the beta coefficients from the estimation of the AD model. On the upper part of the graph, we depict the individual beta coefficient estimates for the eight portfolios of each country, ± one standard error. This part of the graph illustrates the amount of within-country variation in betas, in relation to their standard errors. On the lower part, we graph the average beta value in each country. This shows the cross-country variation in betas. Summary statistics on all beta coefficients estimated from the three models are reported in Table 8.

Both the Figures and the statistics in Table 8 suggest that the cross-country dispersion of beta coefficients is larger than their average cross-sectional dispersion. This is especially the case for the exchange rate and inflation betas. World market betas

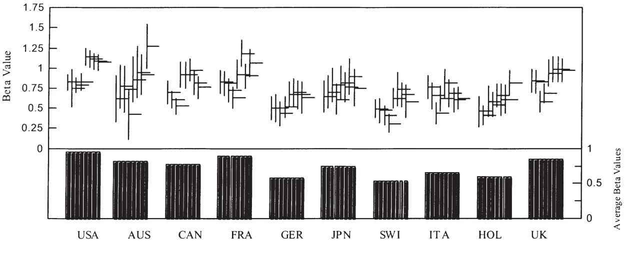

Fig. 1. World market betas. The coefficients are from the estimation of model (18). The upper part of the graph depicts the individual beta coefficient for each of the 80 portfolios, ± one standard error. The lower part of the graph reports the average beta value in each country.

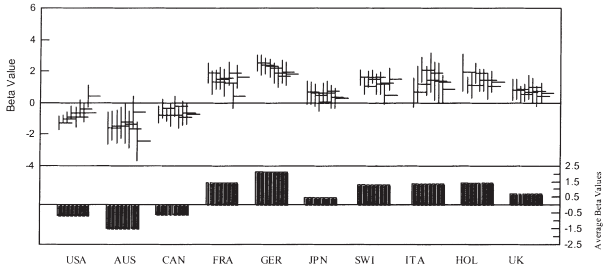

Fig. 2. Common component exchange rate betas. The coefficients are from the estimation of model (18). The upper part of the graph depicts the individual beta coefficient for each of the 80 portfolios, ± one standard error. The lower part of the graph reports the average beta value in each country.

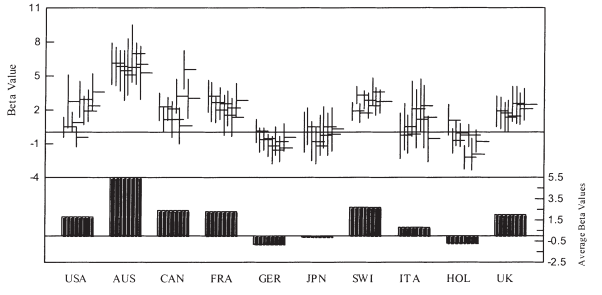

Fig. 3. Residual component exchange rate betas. The coefficients are from the estimation of model (18). The upper part of the graph depicts the individual beta coefficient for each of the 80 portfolios, ± one standard error. The lower part of the graph reports the average beta value in each country.

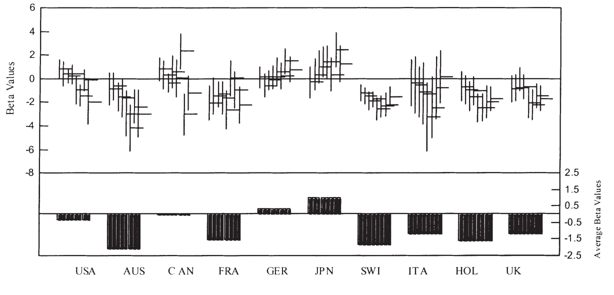

Fig. 4. US unanticipated inflation betas. The coefficients are from the estimation of model (18). The upper part of the graph depicts the individual beta coefficient for each of the 80 portfolios, ± one standard error. The lower part of the graph reports the average beta value in each country.

tend to possess substantial cross-sectional dispersion, and they are estimated with smaller standard errors. Residual component exchange rate betas are more noisy than common component exchange rate betas, but they exhibit larger within-country and cross-country variation. Similarly, US inflation betas are more noisy than residual world inflation betas, but they also possess larger within-country and cross-country

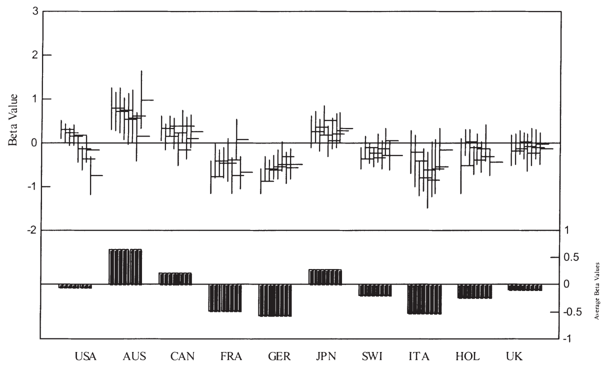

Fig. 5. World unanticipated inflation betas. The coefficients are from the estimation of model (18). The upper part of the graph depicts the individual beta coefficient for each of the 80 portfolios, ± one standard error. The lower part of the graph reports the average beta value in each country.

Table 8
Summary statistics on beta coefficient estimates[a]

Beta coefficient	Average coefficient value	Average standard error	Average t-value	Average cross-sectional dispersion	Cross-country dispersion
The Solnik-Sercu model: $R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^e\beta_k^e+\beta_k^e e_t+\gamma^\lambda\beta_k^\lambda+\beta_k^\lambda\lambda_t+\zeta_{kt}$					
β^w	0.751	0.162	5.030	0.139	0.199
β^λ	0.414	0.247	1.977	0.152	0.659
β^e	1.197	1.401	0.804	0.680	1.904
The Grauer, Litzenberger, and Stehle model: $R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^i\beta_k^i+\beta_k^i i_t+\psi_{kt}$					
β^w	0.847	0.147	6.176	0.125	0.166
β^i	0.112	1.359	0.125	0.765	1.08
The Adler and Dumas model: $R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^e\beta_k^e+\beta_k^e e_t+\gamma^\lambda\beta_k^\lambda+\beta_k^\lambda\lambda_t+\gamma^i\beta_k^i+\beta_k^i i_t+\gamma^v\beta_k^v+\beta_k^v v_t+\xi_{kt}$					
β^w	0.739	0.164	4.869	0.136	0.197
β^λ	0.572	0.744	0.914	0.382	1.214
β^e	1.587	1.502	1.015	0.809	2.060
β^i	−0.869	1.321	−0.689	0.847	1.322
β^v	−0.091	0.386	−0.314	0.217	0.452

[a] The summary statistics for the beta coefficients are from the 80 equations systems. World betas are denoted by β^w, common component exchange rate betas by β^λ, residual component exchange rate betas by β^e, US unanticipated inflation betas by β^i, and finally, residual (to the US unanticipated inflation) world inflation betas by β^v. The "average coefficient value" is the average value of the beta coefficient across the 80 portfolios. Average standard errors and t-values across the 80 portfolios are reported in the columns labeled "average standard error", and "average t-value". To calculate the "average cross-sectional dispersion", we calculate the standard deviation for the beta coefficient across the eight portfolios of each country, and we report the average standard deviation across the 10 countries. The "cross-country dispersion" refers to the standard deviation of the beta coefficients across the 80 portfolios.

variation. It appears that the estimation of country-specific exchange rate and foreign inflation risk premiums in this study was possible due to the substantial within-country variation in the residual component exchange rate betas and the US inflation betas. The construction of the residual component exchange rate index for testing the presence of exchange rate risk premiums in international equities and the use of US inflation as a proxy of foreign inflation risk premiums constitute contributions of this study.

4.4.4. Tests of survivorship bias

In Section 3.1 we noted that all individual security returns used in the construction of our portfolios have a continuous record for the whole time-period covered by our study. It is therefore possible that our results regarding the pricing of exchange rate and foreign inflation risk in equities suffer from survivorship bias.

We test for this possibility in the following way. We perform cross-country tests for the pricing of exchange rate and foreign inflation risk using two alternative databases. Our first set of tests use the 80 portfolios constructed in this study. The second set of tests uses MSCI country indexes.[13] Cross-country tests estimate the average magnitude and significance of exchange rate and foreign inflation risk premiums across countries. The pricing of exchange rate and foreign inflation risk across countries has been documented in previous studies. Our aim here is to use these tests in order to evaluate the possible effect that the survivorship bias in our data may have on our country-specific estimates of exchange rate and foreign inflation risk premiums.

Table 9 presents the results. It is interesting to note that independently of whether exchange rate and foreign inflation risk premiums are estimated using as left-hand-side variables the 80 portfolios or the 10 MSCI indexes, the magnitude of the risk premiums are remarkably similar in the two estimations. This reveals that the cross-country estimates of exchange rate and inflation risk premiums using the 80 portfolios do not suffer in any significant manner from the survivorship bias present in our individual security returns. It is therefore unlikely that the estimates of within-country exchange rate and foreign inflation risk premiums are subject to survivorship biases.

Table 9 also reports the ME and RMSE for all models and for both estimations. The difference in their magnitudes between estimations is very small. In all cases, the ME and RMSE from estimations that use the MSCI indexes are smaller than those from estimations that use the 80 portfolios. This is not surprising since in tests that use the 10 MSCI indexes, the model needs to explain only the cross-country differences in returns and not also their within-country differences. Finally, the adjusted R-squares and Hansen's tests reveal similar information about the models

[13] It is worthwhile to mention that the constituents of the MSCI indexes are not necessarily the same securities used for the construction of the eighty portfolios. Furthermore, membership of securities in the MSCI indexes is updated semi-annually, and therefore, the MSCI indexes are free from survivorship bias.

Table 9
Estimation of average cross-country exchange rate and foreign inflation risk premiums[a]

	γ_0	γ^λ	γ^e	γ^i	γ^u	ME*	RMSE[†]	$R^{2‡}$	J_T[§]
The Solnik-Sercu model: $R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^e\beta_k^e+\beta_k^e e_t+\gamma^\lambda\beta_k^\lambda+\beta_k^\lambda\lambda_t+\zeta_{kt}$									
80 portfolios	−0.0085	0.0105	0.0071			0.00017	0.05307	39.01	$\chi^2(50)=54.5$
	(−1.75)	(19.16)	(30.99)						P-value: 0.30
MSCI indexes	0.0005	0.0105	0.0065			0.00015	0.05099	49.59	$\chi^2(7)=3.65$
	(2.63)	(5.69)	(15.58)						P-value: 0.82
The Grauer, Litzenberger, and Stehle model: $R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^i\beta_k^i+\beta_k^i i_t+\psi_{kt}$									
80 portfolios	−0.0003			0.0071		0.00012	0.05627	36.26	$\chi^2(78)=74.8$
	(−0.72)			(44.02)					P-value: 0.58
MSCI indexes	0.0008			0.0078		0.0001	0.05244	43.29	$\chi^2(8)=3.51$
	(4.88)			(6.76)					P-value: 0.90
The Adler and Dumas model: $R_{kt}=\gamma_0(1-\beta_k^w)+\beta_k^w R_t^w+\gamma^e\beta_k^e+\beta_k^e e_t+\gamma^\lambda\beta_k^\lambda+\beta_k^\lambda\lambda_t+\gamma^i\beta_k^i+\beta_k^i i_t+\gamma^v\beta_k^v+\beta_k^v v_t+\xi_{kt}$									
80 portfolios	−0.0006	0.0106	0.0075	0.0081	0.0051	0.00005	0.05249	44.44	$\chi^2(75)=75.63$
	(−1.09)	(13.74)	(38.33)	(34.73)	(2.46)				P-value: 0.52
MSCI indexes	0.0003	0.0134	0.0076	0.0089	0.0014	0.00003	0.04898	50.51	$\chi^2(5)=1.19$
	(0.83)	(3.86)	(6.08)	(4.5)	(0.12)				P-value: 0.94

[a] Average cross-country exchange rate and foreign inflation risk premiums are estimated using an iterated generalized methods of moments (GMM) procedure, and employing the Newey–West estimator (the truncation parameter q was set equal to 6). Each of the three models, namely the Solnik–Sercu model, the Grauer, Litzenberger, and Stehle model and the Adler and Dumas model are estimated using two alternative sets of data. The first estimation uses as left-hand-side (LHS) variables the 80 portfolios constructed according to the methodology described in Section 3.2. The second estimation uses as LHS variables the MSCI country indexes of the 10 countries in our sample. All variables are expressed in terms of US dollars. The data span the period from 1973:01–1990:12. The coefficients γ^e, γ^λ, denote risk premiums with respect to the residual and common components of changes in the exchange rates, and γ^i, γ^v, denote risk premiums with respect to the US unanticipated inflation rate, and the residual world unanticipated inflation rate, orthogonal to the US unanticipated inflation rate. The coefficients $\beta_k^w,\beta_k^e,\beta_k^\lambda,\beta_k^i$ and β_k^v denote betas of portfolio k with the world market portfolio, the residual and common component exchange rate indexes, the US unanticipated inflation, and the residual world unanticipated inflation, respectively. The estimates of beta coefficients are not reported in this Table. If the model is true, then $\gamma_0=0$. As a proxy for the return on the world market portfolio, R^w, we use the return on the MSCI world index. When the models are estimated using as LHS variables the 80 portfolios, we use only the even observations for the period 1973:01–1990:12, i.e. 108 observations in total. The odd observations were utilized for the construction of portfolios. The estimations of the models that use as LHS variables the 10 MSCI indexes employ all data points, i.e. 216 observations in total. In all cases, beta and gamma coefficients are estimated simultaneously. T-values corrected for heteroskedasticity and serial correlation appear in parentheses below the coefficient estimates.

independently of whether they are estimated using the MSCI indexes or the 80 portfolios.

4.4.5. The robustness of the risk premiums estimates

Recall that each of the three models is estimated in a system of eighty equations which may raise concerns regarding the small sample properties of our estimator. For instance, the results reported in Table 4 involve the estimation of (80×79)/2=3,160 variance-covariance terms plus 270 parameters, a total of 3,430 moments. With 8,640 total observations in the system, the ratio of observations per moment is 2.5 which may be considered low.[14]

We provide additional estimations of country-specific exchange rate and inflation risk premiums in Table 10. In these tests, we estimate the risk premiums for each country separately. The restriction imposed by these tests compared to those reported in Tables 4–6 is that the covariance matrix of residuals across countries is zero. This restriction holds under the null hypothesis that each model is true. By imposing this restriction, we can estimate each model country by country in systems of eight equations. The only shortcoming of this restriction is that we can no longer test the hypothesis of equality of risk premiums across countries. For this test, we continue to rely on the estimations from the 80 equation systems.

The ratios of observations per moment increase now to 18.8 for the GLS model, 15.7 for the S-S model, and 11.8 for the AD model. Although these significant increases in the ratios, compared to those for Tables 4–6, may not eliminate completely any small-sample biases that can potentially exist in our results, they relieve them greatly. In addition, it is possible that the efficiency of the Newey–West estimator in the tests of Tables 4–6 is low as result of the size of the systems. This may be the case despite the fact that the Newey–West estimator is always positive definite. To account for this possibility, we also report, in square brackets, *t*-values from nonlinear least squares standard errors.

Table 10 reveals that our results remain qualitatively the same with those of Tables 4–6, when the risk premiums are estimated for each country separately. Exchange rate and inflation risk is again priced in the same countries as before. Exceptions are found in the case of Japan, where the common component exchange rate risk premium is not priced anymore, as is the case for the inflation risk premium in the GLS model for Japan. Furthermore, the common component exchange rate risk premium in Switzerland is now positive. Finally, in Italy the residual component exchange rate risk premium is now priced, whereas the common component exchange rate risk premium is priced only marginally. Given that we estimate 70 risk premiums in total, such few differences between alternative estimations are to be expected and do not affect the economic interpretation of the results of this study. It is also interesting to note that our inference regarding the pricing of exchange rate

[14] Similarly, the ratio of observations per moment for Tables 5 and 6 are 8,640/3,340=2.6 and 8,640/3,610=2.4, respectively.

Table 10
Estimation of the S-S, GLS and AD models: individual-country tests[a]

Country/model	γ_0	γ^{λ}	γ^{e}	γ^{i}	γ^{ν}	J_T*
Australia						
S-S	0.0214	0.0727	0.0016			$\chi^2(5)$=0.81
	(0.56)	(1.43)	(0.19)			*P*-value: 0.97
	[0.52]	[0.83]	[0.16]			
GLS	−0.0106			0.0141		$\chi^2(6)$=6.29
	(−2.08)			(3.31)		*P*-value: 0.39
	[−2.25]			[2.98]		
AD	0.0140	0.0707	−0.0011	0.0271	0.0113	$\chi^2(12)$=0.43
	(0.29)	(0.43)	(−0.06)	(0.28)	(0.37)	*P*-value: 0.99
	[0.28]	[0.41]	[−0.06]	[0.25]	[0.35]	
Canada						
S-S	0.0022	−0.0040	0.0024			$\chi^2(5)$=2.70
	(0.50)	(−0.20)	(1.06)			*P*-value: 0.74
	[0.53]	[−0.24]	[1.03]			
GLS	−0.0014			0.0044		$\chi^2(6)$=7.43
	(−0.73)			(3.37)		*P*-value: 0.28
	[−0.75]			[3.33]		
AD	−0.0307	−0.0278	0.0233	0.0243	−0.0018	$\chi^2(12)$=1.43
	(−0.25)	(−0.21)	(0.31)	(0.33)	(−0.02)	*P*-value: 0.99
	[−0.19]	[−0.16]	[0.22]	[0.23]	[−0.01]	
France						
S-S	0.0011	−0.0151	−0.038			$\chi^2(8)$=2.99
	(0.16)	(−0.31)	(−0.33)			*P*-value: 0.70
	[0.15]	[−0.32]	[−0.30]			
GLS	0.0011			0.0127		$\chi^2(6)$=2.97
	(0.12)			(5.57)		*P*-value: 0.81
	[0.13]			[3.56]		
AD	0.0031	0.0051	0.0063	0.0112	−0.0094	$\chi^2(12)$=1.72
	(0.38)	(0.49)	(1.5)	(6.33)	(−0.43)	*P*-value: 0.99
	[0.25]	[0.37]	[0.54]	[2.43]	[−0.40]	
Germany						
S-S	0.0095	0.0014	0.0063			$\chi^2(5)$=5.16
	(3.43)	(0.3)	(4.31)			*P*-value: 0.39
	[2.15]	[0.33]	[3.65]			
GLS	0.0055			0.0056		$\chi^2(6)$=6.31
	(2.25)			(4.73)		*P*-value: 0.39
	[2.10]			[4.39]		
AD	0.0146	−0.0131	0.0029	0.0004	−0.0403	$\chi^2(12)$=0.82
	(1.48)	(−0.73)	(0.55)	(0.08)	(−0.98)	*P*-value: 0.99
	[1.20]	[−0.67]	[0.56]	[0.07]	[−0.90]	
Japan						
S-S	0.0144	−0.0118	0.0111			$\chi^2(5)$=0.818
	(1.76)	(−0.79)	(2.73)			*P*-value: 0.97
	[1.78]	[−0.97]	[2.52]			
GLS	0.0087			−0.0047		$\chi^2(6)$=6.18
	(1.07)			(−0.41)		*P*-value: 0.40
	[0.98]			[−0.43]		

(continued on next page)

Table 10 (*continued*)

Country/model	γ_0	γ^{λ}	γ^{e}	γ^{i}	γ^{ν}	J_T*
AD	0.0146	−0.0115	0.0107	0.003	−0.0105	$\chi^2(12)$=0.46
	(0.83)	(−0.59)	(1.59)	(0.3)	(−0.26)	*P*-value: 0.99
	[0.81]	[−0.73]	[1.76]	[0.28]	(−0.22)	
Switzerland						
S-S	0.0023	0.0148	0.0025			$\chi^2(5)$=8.62
	(0.31)	(2.19)	(0.77)			*P*-value: 0.13
	[0.12]	[3.22]	[1.60]			
GLS	−0.0063			0.0305		$\chi^2(6)$=8.05
	(−1.55)			(2.05)		*P*-value: 0.23
	[−1.45]			[1.97]		
AD	−0.0004	−0.0062	0.0093	−0.0634	−0.0181	$\chi^2(12)$=4.41
	(−0.03)	(−0.35)	(2.08)	(−0.77)	(−1.48)	*P*-value: 0.97
	[−0.25]	[−0.33]	[1.99]	[−0.75]	[−1.37]	
Italy						
S-S	−0.0024	0.0157	0.0086			$\chi^2(5)$=2.77
	(−0.48)	(1.95)	(6.14)			*P*-value: 0.73
	[−0.54]	[1.69]	[4.37]			
GLS	−0.0007			0.0053		$\chi^2(6)$=2.94
	(−0.16)			(5.42)		*P*-value: 0.81
	[−0.22]			[4.62]		
AD	−0.0131	−0.0293	0.007	0.0122	0.0178	$\chi^2(12)$=1.20
	(−0.38)	(−0.34)	(1.26)	(0.84)	(0.61)	*P*-value: 0.99
	[−0.34]	[−0.33]	[1.15]	[0.79]	[0.59]	
Netherlands						
S-S	0.0045	0.0093	0.0092			$\chi^2(5)$=3.23
	(2.53)	(2.18)	(10.86)			*P*-value: 0.66
	[1.40]	[2.00]	[6.57]			
GLS	0.0059			0.006		$\chi^2(6)$=2.48
	(2.65)			(8.38)		*P*-value: 0.87
	[1.84]			[5.31]		
AD	0.0088	0.0106	0.0021	0.0166	0.0023	$\chi^2(12)$=0.94
	(0.61)	(0.87)	(0.17)	(1.06)	(0.07)	*P*-value: 0.99
	[0.48]	[0.15]	[0.62]	[0.86]	[0.07]	
UK						
S-S	−0.025	0.0143	0.0314			$\chi^2(5)$=0.34
	(−0.46)	(0.44)	(0.6)			*P*-value: 0.99
	[−0.46]	[0.55]	[0.57]			
GLS	−0.0024			0.015		$\chi^2(6)$=8.72
	(−1.04)			(1.7)		*P*-value: 0.19
	[−0.95]			[1.65]		
AD	−0.0251	0.0157	0.0317	0.0751	−0.0059	$\chi^2(12)$=0.33
	(−0.31)	(0.1)	(0.39)	(0.41)	(−0.02)	*P*-value: 0.99
	[−0.29]	[0.09]	[0.33]	[0.31]	[−0.02]	
USA						
S-S	−0.0018	0.0045	0.0051			$\chi^2(5)$=3.99
	(−0.65)	(0.57)	(3.96)			*P*-value: 0.55
	[−0.68]	[0.61]	[3.88]			
GLS	−0.0015			0.005		$\chi^2(6)$=3.19
	(−0.84)			(4.74)		*P*-value: 0.78
	[−0.90]			[5.10]		

(continued on next page)

Table 10 (*continued*)

Country/model	γ_0	γ^λ	γ^e	γ^i	γ^v	J_T*
AD	0.0009	0.0147	0.0068	0.0045	0.0206	$\chi^2(12)=2.67$
	(0.28)	(1.71)	(4.59)	(3.73)	(1.33)	*P*-value: 0.99
	[0.27]	[1.46]	[4.03]	[3.55]	[1.11]	

[a] This table presents single-country tests of the S-S, GLS and AD models. Exchange rate and inflation risk premiums are estimated for each country separately in systems of eight equations. The tests are conducted in US$ terms. The data span the period from 1973:01–1990:12. The coefficients γ^e, γ^λ, denote risk premiums with respect to the residual and common components of changes in the exchange rates, and γ^i, γ^v, denote risk premiums with respect to the US unanticipated inflation rate, and the residual world unanticipated inflation rate. If the model is true, then $\gamma_0=0$. Only even observations for the period 1973:01–1990:12 are used, i.e. 108 observations in total, since odd observations were previously used for the construction of portfolios. Corrected *t*-values for heteroskedasticity and serial correlation appear in parentheses below the coefficient estimates. The numbers in square brackets denote *t*-values for the same coefficients calculated, however, using nonlinear least squares standard errors.
*Hansen's GMM test.

and inflation risk is the same whether it is carried out on the basis of nonlinear least squares or Newey–West standard errors.

5. Conclusions

This paper examined the ability of exchange rate and foreign inflation risk factors to explain the within-country differences in average returns. The hypotheses tested were motivated by three international asset pricing models, namely the Solnik (1974b) model as revised by Sercu (1980), the Grauer et al. (1976) model, and the Adler and Dumas (1983) model. Our results suggest that exchange rate and foreign inflation risks are generally priced in the equity returns of the 10 countries in our sample.

We decomposed changes in a cross-section of exchange rates into a component which is common across exchange rates, and a residual component. Based on this decomposition, we estimated exchange rate risk premiums and found that in several countries the exchange rate risk premium is at least partly attached to the residual component of changes in exchange rates. The pricing of the residual component of exchange rates has not been previously examined in the literature. Our results suggest that it has important implications for the pricing and hedging of exchange rate risk.

Following the implications of the Grauer, Litzenberger, and Stehle model, we tested whether US inflation risk is priced in the cross-section of equity returns of all 10 countries. We found that US inflation risk is indeed priced in all countries. This result constitutes the second contribution of the paper and it is contrary to the popular belief that only domestic inflation may be priced in the equities of a given country. It implies that home bias in US portfolios cannot be the result of investors' efforts to hedge US inflation.

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