The Role of Exchange Rates in the Intertemporal Risk-Return Relation in International Economies*

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Abstract

This paper investigates the role of currency denomination in the the intertemporal risk-return relation among G7 countries. Similar to the findings of previous studies, our estimation also shows that the financial markets of the G7 countries are integrated. We obtain significant pricing coefficient estimates on the global index, but insignificant estimates on country-specific risks. Different from the literature, however, we find that the intertemporal risk-return relation differ significantly under different currency denominations. The slope coefficient estimate is the largest at around seven when the returns are denominated in Japanese yen, smallest at around three to four when the returns are denominated in pound or euro and its predecessors. The slope estimates are in the middle at about five to six when the returns are denominated in the U.S. or Canadian dollar. The estimates stay in the same range and the rankings remain unchanged when we consider different specifications for the conditional covariance estimations and when we replace country-specific portfolios with industry-specific forms in estimating the intertemporal relation.

JEL Classifications: G12, C32, F03, F31.

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*This is a preliminary draft. Please do not quote. We welcome comments, including references that we have inadvertently missed.

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1. Introduction

Even with global financial market integration, deviation from purchasing power parity remains persistent, with convergence to the parity condition extremely slow. The short-run deviations are often large in magnitude, and volatile in their fluctuations (Rogoff (1996)). These deviations suggest that international goods market remain quite segmented, even though the financial market is much more integrated. The segmentations, together with specifications on consumer preferences, determine the dynamics of real exchange rates (Dumas (1992)). Reversely, the real exchange rate dynamics reveal important information about goods market segmentations and consumer preference differences across different countries.

The dynamics of nominal exchange rates further reflect the countries' differences in monetary and fiscal policies. Yet, one of the most puzzling features of exchange-rate behavior since the advent of floating exchange rates in the early 1970s is the tendency for countries with high interest rates to see their currencies appreciate rather than depreciate as the uncovered interest rate parity would suggest.[1] Fama (1984) attributes this gross violation of the uncovered interest rate parity to a time-varying risk premium. Fama shows that the implied risk premium on a currency must (1) be negatively correlated with its expected rate of depreciation and (2) have greater variance.

In an international economy with financial market integration but goods market segmentation, consumers can invest globally, but they consume in their local currency. Therefore, consumers in different countries face different risk profiles and can charge different risk premiums, due to their differences in consumption numeraires, i.e., currency denominations. More importantly, the persistent and volatile divergence of purchasing power party suggests that the exchange rate risk between two economies can be persistent and substantial. Furthermore, the gross violation of the uncovered interest rate parity suggests that not only the currency risks are priced, but also the risk premiums are likely to be time varying.

Despite the above evidence and theory on the importance of the exchange rate risk and risk premium, most empirical works on international capital asset pricing ignore the exchange rate risk and its implications for potentially different risk-return relations under different currency denominations. Empirical studies either estimate the risk-return relation under some real measures (Cavaglia, Hodrick, Vadim, and Zhang (2002)) or in most cases assume that the same risk-return relation holds across all currency de-

[1] Many studies document this phenomenon and many more attempt account for its existence. See, for instance, Bakshi and Naka (1997), Bekaert (1995), Bekaert, Wei, and Xing (2003), Canova and Marrinan (1995), Engel (1996), Flood and Rose (1996), Hodrick (1987), Phillips and McFarland (1997), and Wu and Zhang (1997).

nominations (e.g., De Santis, , and Gerard (1997), Carrieri, Errunza, and Sarkissian (2004), and Zhang (2004)). The motivation of these studies can go back to Solnik (1974), who does not give an explicit role for the exchange rate by assuming that exchange rate risk can be fully hedged, or to Grauer, Litzenberger, and Stehle (1976), who assume purchasing power parity holds. In reality, however, consumers in different countries still face the real exchange rate risk, as evidenced by the persistent and volatile deviation of the purchasing power parity. When investment returns are denominated in nominal terms, it is even harder to imagine how the same risk-return relation should hold, given the strong violation of the uncovered interest rate parity.

In this paper, we estimate and compare the intertemporal risk-return relations in the international economy under different currency denominations. The difference in the coefficient estimates reflects the difference in consumer preferences in different economies, the degree of goods market segmentation, and the pricing of exchange rate risk.

We use monthly data from January 1970 to December 2004 on the MSCI global index returns and returns on seven country indices for the G7 countries. First, we compute monthly returns on the eight indices based on one currency denomination. Then, we calculate the excess returns as the index return minus the short-term interest rate in that currency-denominated economy. For example, when we use U.S. dollar as the common currency, we use the short-term U.S. interest rate to compute the excess returns on the eight indices. When we use Japanese yen as the common currency, we use the short-term yen interest rate to compute the excess returns on the eight indices; and so on. Third, we estimate the conditional covariances between the global return and returns on the seven country indices using a bivariate GARCH specification. Finally, we estimate the common relation of the seven simultaneous equations implied by the international version of the intertemporal capital asset pricing model between the expected excess returns on the seven country indices and the corresponding conditional covariances. We repeat the above procedure and estimate the intertemporal risk-return relation under each of the seven currency denominations. Under each currency denomination, one common slope applies to all seven risk-return equations, as suggested by a global version of the intertemporal capital asset pricing model. However, the slope coefficient estimates can be different when we estimate the relation under different currency denominations.

Under all seven currency denominations, the risk-return coefficient estimates are all positive and statistically significant, suggesting that covariance with the global index is priced under all currency denominations. Nevertheless, the coefficient estimates are quite different under different currency denominations, showing the significant pricing of the exchange rate risk. In particular, the risk-return coefficient estimate is the highest at about seven when returns are denominated in Japanese yen, suggesting that Japanese investors are the most risk averse on average. On the other end of the spectrum, the coefficient estimates are the lowest at between three and four when the returns are denominated in pounds or the euros and their predecessors. The coefficient estimate are in the middle at between five and six when the returns are denominated in either U.S. or Canadian dollars.

To investigate whether country-specific risks are priced in additional to the global portfolio, we regress the country index returns on the global index returns and use the residual to denote country-specific risks. We estimate the conditional covariance between the country index returns and the residual risks in that country. We then add this covariance term to the simultaneous regressions and study whether the coefficient estimates on this term are statistically significant. Estimation shows that the coefficient estimates on this covariance term are all insignificant, suggesting that the G7 countries indeed are integrated financially. Nevertheless, with or without this residual term, the slope coefficient estimates on the covariance with the global index returns are always statistically significant, and different under different currency denominations. The ranking of the coefficient estimates does not change whether we add the residual term or not.

For robustness check, we also consider different GARCH specifications and distributional assumptions when estimating the conditional covariances. In each case, the coefficient estimates on the covariance with the global index returns remain similar and the rankings remain the same across different currency denominations.

Therefore, similar to the findings of previous studies, we also conclude that the financial markets in the seven industrialized countries are integrated. One global index captures the major risk that is priced, and country-specific risks are not priced. Different from them, however, we explicitly address the pricing of currency risk and show that indeed the risk-relation relations are different under different currency denominations. On the surface, the estimation shows that currency risks are priced, a result that is obtained from studies on uncovered interest rate parities, but somehow does not carry over to the international capital asset pricing literature. In a deeper level, our results reveal the different risk attitudes and pricing among investors that consume in different currency-denominated economies. Taken together, our study highlights the fact that the exchange rate dynamics are neither exogenously given, nor tangential to asset pricing, but rather very informative about the pricing difference across different economies.

Obviously, we are not the first to discuss the pricing of exchange rate risk. It has been recognized as early as in Alder and Dumas (1983). Nevertheless, empirical studies that recognize exchange rate risk, e.g., Dumas and Solnik (1995) and Cavaglia, Hodrick, Vadim, and Zhang (2002), often treat the exchange rate as a separate risk factor and investigate whether this risk generates a separate risk premium. Similar to these studies, our work explicitly recognizes the pricing of exchange rate risk. Different from them, however, we exploit this pricing for a different purpose. We use the exchange rate dynamics to identify the differences in risk preferences among consumers at different geographic areas who are limited to consume in different local currencies.

The paper is structured as follows. The next section lays out the theoretical foundations of an intertemporal capital asset pricing model in financially integrated international economies and shows the important role of currency denomination in revealing the different pricing attitudes across different economies. Section 3 describes the data set and the estimation methodology. Section 4 presents the results based on the country index portfolios. Section 5 discusses the results based on the industry-specific portfolios formed globally. Section 6 performs robustness analysis by considering alternative ways of estimating the covariances. Section 7 concludes.

2. Currency denomination and the intertemporal risk-return relation

In his seminal paper, Merton (1973) derives an intertemporal capital asset pricing model for a single economy. According to this model and with the assumption of constant investment opportunity, the conditional expected excess return on a stock or stock portfolio is proportional to its covariance with the market portfolio return in that economy,

$$\mu - r = A\Sigma, \tag{1}$$

where $\mu \in \mathbb{R}^n$ denotes the expected excess return on a vector of n risky assets, r denotes the riskfree rate, A reflects the average relative risk aversion of the market investors, and $\Sigma \in \mathbb{R}^{n+}$ denotes the covariance of the vector of the excess return with the market portfolio.

In a financially integrated global economy, the benefit of international diversification suggests that investors should hold the global market portfolio instead of a country-specific market portfolio. Thus, by analogy, the expected excess return on a stock or stock portfolio should be proportional to its covariance with the global market portfolio return, as a natural extension of the single-economy result in (1).

However, a key difference resides between a local economy and a global economy in that investing in foreign stocks incurs the exposure to exchange rate risk. Even if investors can trade foreign stocks electronically and thus without incurring any extra cost than trading domestic stocks in an integrated global financial market, converting investments in the foreign stocks to domestic consumptions faces the exchange rate risk between the two economies. In an ideal world where goods transportation and labor mobility are frictionless, purchasing power parity should always hold, the real exchange rate should always stay at par, and nominal exchange rate fluctuations should just reflect the nominal interest rate difference in the two economies as predicted by the uncovered interest rate parity. In such an ideal world, the exchange rate becomes a trivial numerarie switch that is immaterial to pricing. Indeed, many empirical studies on international capital asset pricing implicitly or explicitly trivialize the role of exchange rates in the pricing relation. These studies compute country-portfolio returns either using U.S. dollar as the common denomination or their respective local currency as the denomination. Then, they proceed to estimate one common risk-return relation as suggested by the single-economy relation in (1), and thus reducing the role of the exchange rate as a pure conversion of scale.

Yet in reality, although the financial market is becoming increasingly integrated globally, goods transportation is still costly and labor mobility remains limited. As a result, real exchange rate persistently deviates from the parity condition, with the deviations both large and volatile. Furthermore, the nominal exchange rate movement often defies the prediction of the uncovered interest rate parity, showing that the exchange rate risk is nothing but orthogonal to pricing. The exchange rate is not only priced, but also time-varying, potentially due to variations in the average risk attitudes of the investors in the two economies.

Thus, we argue that the exchange rate dynamics are not orthogonal to pricing, but instead are informative about how the risk premiums differ across two economies. Formally, we consider N economies where investors are free to invest in both domestic and foreign stocks, but where their consumptions are denominated to their respective local currency. As an example, a U.S. resident is free to invest in Japanese stocks, but is not free to buy commodities in Japanese yen. The cost of investing in a Nikkei-indexed portfolio is small and can be readily accomplished via many brokerage houses in the U.S.; but the cost of travelling to Japan to spend the yen is large.

Let $k = 1, 2, \cdots, N$ denotes investors that can only consume in the kth currency denomination. The expected excess returns to their global investments, when converted into their local currency, are proportional to the covariance of the investments with the global market portfolio, all denominated in the local currency:

$$\mu^k - r^k = A^k \Sigma^k, \tag{2}$$

where μ^k denotes the expected excess return in the local currency k on a vector of n risky assets, r^k denotes the kth local-economy riskfree rate, A^k reflects the average relative risk aversion of the investors whose consume in the kth currency, and Σ^k is the covariance matrix between the n risky asset returns and the global market portfolio, all measured in the kth currency.

Let j denote another currency and let S^{jk} denote the exchange rate between the two, specifically the jth currency price of the kth currency. As an example, if k denotes dollar and j denotes yen, S^{jk} denotes the yen price of one dollar. Thus, if P^k is the price of an asset in dollars, $P^k S^{jk}$ becomes the price of the asset in yen. Furthermore, if $R_t^k \equiv \ln(P_t^k / P_{t-1}^k)$ denotes the time-t one-period return on the asset in dollar, then the asset's return in yen is $R_t^j = R_t^k + s_t^{jk}$, where $s_t^{jk} \equiv \ln(S_t^{jk}/S_{t-1}^{jk})$ is the depreciation rate of the dollar against the yen.

Now let us write the intertemporal relation first in $j$the currency,

$$\mu^j - r^j = A^j \Sigma^j, \tag{3}$$

and then convert the relevant terms into the k currency denomination,

$$\begin{aligned} \mu^k + \mu^{jk} - r^j &= A^j Cov(R^j, R_g^j) = A^j Cov(R^k + s^{jk}, R_g^k + s^{jk}) \\ &= A^j \Sigma^k + A^j \left(Cov(s^{jk}, R_g^k) + Cov(s^{jk}, R^k) + Var(s^{jk}) \right), \end{aligned} \tag{4}$$

where $Cov(\cdot,\cdot)$ and $Var(\cdot)$ denote the covariance and variance of what follows, and $Rand R_g$ denote the return vector on the n risky assets and the global market portfolio return, respectively. The empirical literature has been assuming $A^k = A^j = A$ when estimating the intertemporal relation, which amounts to the strong assumption of,

$$\mu^{jk} - (r^j - r^k) = A \left(Cov(s^{jk}, R_g^k) + Cov(s^{jk}, R^k) + Var(s^{jk}) \right), \tag{5}$$

where the left hand side captures the deviations from the uncovered interest rate party, and the right hand side measures the variance and covariances of the currency returns. There is no rationale why this equality should hold in general. One extreme example that makes this equality hold to assume the validity of the uncovered interest rate parity on the left hand side and assume a deterministically moving exchange rate on the right hand side. Yet, the strong violation of the uncovered interest rate parity suggests that more likely than not $A^k \neq A^j$ for $k \neq j$. Thus, the exchange rate dynamics become informative about the risk preference difference across different economies.

3. Data and estimation

We estimate the intertemporal risk-return relation among the industrialized countries under different currency denominations according to equations (2) and (3). By comparing the slope coefficient estimates under different currency denominations, we learn about the differences in risk preferences across different economies.

3.1. Data

We estimate the intertemporal risk-return relation using both country-index portfolio returns and industry portfolio returns.

The country index portfolio returns are from MSCI on the G7 countries: the United States, Canada, Japan, the United Kingdom, Italy, Germany, and France. MSCI also provides a global market index portfolio, which we use as the global market portfolio risk factor. Data are monthly from December 1969 to December 2004. We download the gross index data in both in dollar denominations and in their respective local currencies, and compute the monthly log returns. The difference between the dollar return and the local currency return on each index defines the return on the dollar price of the local currency. We use these currency return series to convert dollar returns on all seven country portfolios into returns on each of the other six currency denominations.

The industry portfolio data are from Datastream. There are 25 industry portfolios on Datastream for each country. From these, we form equally weighed industry portfolios for the G7 countries. We again download the portfolios both in dollar denomination and their respective local currencies and convert the log returns on all 25 G7-industry portfolios into seven different currency denominations. Data for the industry portfolio are monthly from January 1973 to December 2004.

To compute excess returns, we use the short-term riskfree rate from each currency, collected from various sources to match our monthly sample on the country and industry portfolios.

Table 1 presents the means, standard deviations, skewness, excess kurtosis, Jarque-Bera (JB), and the first-order autocorrelation statistics for the country and global index portfolio returns. [2] As shown in the table, there is substantial variation in the summary statistics across different currency denominations, implying a significant impact of exchange rates on the country and global portfolio returns.[3] The skewness statistics are generally small, but negative and significant for most of countries. The excess kurtosis statistics are generally high and significantly different from zero in most cases. Tail-thickness and leptokurtosis seem to be more dominant than skewness in our sample. The Jarque-Bera statistics indicate that the excess returns on country and global market portfolio are not normally distributed.

Overall, the results show that the market risk of portfolios including domestic and foreign assets cannot be accurately described by the mean and standard deviation alone. International investors are required to measure the distribution's skewness and kurtosis to assess the risk of their investments more accurately. The summary statistics also indicate that when international portfolios are involved, estimating a risk-return relation with the assumption of normality may result in misleading estimates of the relative risk aversion coefficients.

3.2. Estimating conditional covariances

One key issue in studying the intertemporal risk-return relation is to estimate the conditional conditional covariances as they are not directly observable. To estimate monthly conditional covariances from monthly return data, we take the GARCH modeling approach of Engle (1982) and Bollerslev (1986).

Formally, under each currency denomination, we estimate the conditional covariance between excess returns on asset *i* (country-specific or industry-specific portfolios) and the global market portfolio *m* based

[2]Standard errors of the skewness and excess kurtosis estimates, computed under the null hypothesis that returns are normally distributed, are $\sqrt{6/n}$ and $\sqrt{24/n}$, respectively. Jarque-Bera, $JB = n[(S^2/6) + K^2/24]$, is a formal test statistic for testing whether the returns are normally distributed, where n denotes the number of observations, S is skewness and K is excess kurtosis. The JB statistic distributed as the Chi-square with two degrees of freedom measures the difference of the skewness and kurtosis of the series with those from the normal distribution. The critical values with two degrees of freedom at the 10%, 5%, and 1% level of significance are 4.61, 5.99, and 9.21, respectively.

[3]For example, for the U.S. equity portfolio, the mean of excess returns is in the range of 2.36% to 4.98% per annum. The standard deviations range from 15.13% to 19.28% per annum. Other statistics are also different for returns measured under different current denominations.

on the following set of specifications,

$$
\begin{aligned}
R_{i,t+1} &= \alpha_0^i + \alpha_1^i R_{i,t} + \varepsilon_{i,t+1}, && (6)\\
R_{m,t+1} &= \alpha_0^m + \alpha_1^m R_{m,t} + \varepsilon_{m,t+1}, && (7)\\
\mathbb{E}_t\left[\varepsilon_{i,t+1}^2\right] \equiv \sigma_{i,t+1}^2 &= \gamma_0^i + \gamma_1^i \varepsilon_{i,t}^2 + \gamma_2^i \sigma_{i,t}^2, && (8)\\
\mathbb{E}_t\left[\varepsilon_{m,t+1}^2\right] \equiv \sigma_{m,t+1}^2 &= \gamma_0^m + \gamma_1^m \varepsilon_{m,t}^2 + \gamma_2^m \sigma_{m,t}^2, && (9)\\
\mathbb{E}_t\left[\varepsilon_{i,t+1}\varepsilon_{m,t+1}\right] \equiv \sigma_{im,t+1} &= \gamma_0^{im} + \gamma_1^{im} \varepsilon_{i,t}\varepsilon_{m,t} + \gamma_2^{im} \sigma_{im,t}, && (10)
\end{aligned}
$$

where $R_{i,t+1}$ and $R_{m,t+1}$ denote the time $(t+1)$ excess returns on assets i and the market portfolio m over a riskfree rate, respectively. First, we use an AR(1) specification to demean the excess returns. Then we define each element of the conditional covariance matrix as a GARCH(1,1) process. We use $E_t[\cdot]$ to denote the expectation operator conditional on time t information. Hence, $\sigma_{i,t+1}^2$, $\sigma_{m,t+1}^2$, and $\sigma_{im,t+1}$ denote the time-t conditional forecasts of the return variance and covariance from time t to time $t+1$. The GARCH specifications in equations (6) to (10) do not arise directly from the ICAPM model, but they provide a close and parsimonious approximation of the form of heteroscedasticity typically encountered with economic time-series data (Bollerslev, Chou, and Kroner (1992) and Bollerslev, Engle, and Nelson (1994)). The specifications in equations (6) to (10) are direct multivariate generalizations of univariate GARCH models.[4]

We estimate the multivariate GARCH specification by maximizing the likelihood function of the return innovations. Traditionally, the likelihood functions are often specified assuming normal return innovations. Given the well-documented evidence on stock return non-normalities, our estimation consider a more general distribution specification for the return innovation that can accommodate fat tails. Specifically, we assume that the conditional distribution of the return innovation for a bivariate student-t distribution. Using ε_t and Σ_t to denote the bivariate demeaned excess return vector and the conditional covariance matrix forecasts,

$$
\varepsilon_t = \begin{bmatrix} R_{i,t} - \alpha_0^i - \alpha_1^i R_{i,t-1} \\ R_{m,t} - \alpha_0^m - \alpha_1^m R_{m,t-1} \end{bmatrix}, \quad \Sigma_t = \begin{bmatrix} \sigma_{i,t}^2 & \sigma_{im,t} \\ \sigma_{im,t} & \sigma_{m,t}^2 \end{bmatrix}, \tag{11}
$$

[4] Similar conditional covariance specifications are used in Ballie and Bollerslev (1992), Bollerslev (1990), Bollerslev, Engle, and Wooldridge (1988), Bollerslev and Wooldridge (1992), Ding and Engle (2001), Engle and Kroner (1995), Engle and Mezrich (1996), Engle, Ng, and Rothschild (1990), and Kroner and Ng (1998).

the bivariate student-t distribution is

$$f(\varepsilon_t|\mathcal{F}_{t-1}) = \frac{\Gamma((v+2)/2)}{\pi(v-2)\Gamma(v/2)}|\Sigma_t|^{-1/2}\left(1+\frac{\varepsilon_t^\top\Sigma_t^{-1}\varepsilon_t}{v-2}\right)^{-(v+2)/2}, \tag{12}$$

where $\mathcal{F}_{t-1}$ denotes the time $(t-1)$ filtration, Γ denotes the gamma function and v denotes the degree of freedom for the t-distribution. Thus, we can write the log-likelihood function as

$$\mathcal{L}(\Theta) = -\frac{1}{2}\sum_{t=1}^{T}\left[2\ln\frac{\pi(v-2)\Gamma(v/2)}{\Gamma((v+2)/2)} + \ln|\Sigma_t| + (v+2)\ln\left(1+\frac{\varepsilon_t^\top\Sigma_t^{-1}\varepsilon_t}{v-2}\right)\right], \tag{13}$$

where Θ denotes the vector of parameters in the specifications (6) to (10) and N denotes the number of monthly observations for each series. We repeat the estimation procedure for each of the seven country-index portfolios and the 25 industry portfolios, and for each portfolio under each of the seven currency denominations, altogether 224 estimations.

The summary statistics presented in Table 1 indicate the significance of non-normality in the return distribution of country-index and global market portfolios. To provide an alternative test of normality, we estimate the degrees of freedom (or tail-thickness) parameter v for the Student-t distribution for each series under each of the seven currency denominations. Table 2 reports the maximum likelihood estimates of v, the corresponding t-statistics, and the Wald test results. The Wald test is used to determine whether $1/v = 0$, which is a formal test of normality against the fat-tailedness of the return distribution. As can be seen from the p-values of the Wald statistics, the null hypothesis of $1/v = 0$ (i.e., the assumption of normality) is generally rejected in favor of the Student-t distribution. Overall, the results in Table 1 and Table 2 justify our use of the bivariate Student-t distribution in estimating the conditional covariances.

We also investigate whether country-specific risks are priced after controlling for the global market portfolio. For this purpose, first we regress returns on each country-specific portfolio i on the global market portfolio,

$$R_{i,t} = a + bR_{m,t} + E_{i,t}, \tag{14}$$

where the regression residual $E_{i,t}$ captures the country-specific risk that is orthogonal to the global portfolio risk factor. Second, we estimate conditional covariances between excess returns on each country portfolio $R_{i,t}$ and that country's residual risk $E_{i,t}$ using an analogous multivariate GARCH specification as delineated above. We label the covariance as $\omega_{i,t}$.

Table 3 presents the parameter estimates for the conditional covariances between the country-index portfolio returns and the global market portfolio returns under each of the seven currency denominations. The persistence of the conditional covariance dynamics on each series is measured by the sum of γ_1^{im} and γ_2^{im}. The results indicate that the conditional covariances are highly persistent for Japan, UK, USA, and Canada. For example, the sum of γ_1^{im} and γ_2^{im} is in the range of 0.93 to 0.99 for Japan, and 0.94 to 0.98 for UK. The persistence of covariances are much lower for Germany, France, and Italy.

Since the parameter estimates on the conditional covariances between the 25 industry portfolios and the global market portfolio returns are similar to those reported in Table 3, we do not present them in the paper. They are available upon request.

3.3. *Estimating the intertemporal coefficients*

Given the conditional covariances under each currency denomination, we estimate the intertemporal risk-return relations using the following simultaneous equations,

$$R_{i,t+1}^k = C_i^k + A^k \sigma_{im,t+1}^k + B^k \omega_{i,t+1}^k + e_{t+1}^k, \quad i = 1, 2, \cdots, n, \tag{15}$$

where k denotes the currency denomination and n denotes the number of portfolios and also the number of simultaneous equations in the estimation. In our estimation, $n = 7$ when using country-index portfolios and $n = 25$ when using industry portfolios. In equation (15), we constrain the slope coefficients (A^k, B^k) under each currency denomination to be the same across all n portfolios for internal consistency. Nevertheless, we allow the intercept C_i^k to be different across different portfolios. Under the null hypothesis of ICAPM, the intercepts should all be zero. We use deviations of the intercept estimates from zero as a test against the validity and sufficiency of our ICAPM specification. Furthermore, we allow all coefficients to differ under different currency denominations. From the different parameter estimates, we infer the pricing of the exchange rate risk and the different risk preferences for investors in different economies, whose consumption is limited to be mainly in their respective local currencies.

We estimate the system of simultaneous equations using a weighted least square method that allows us to place constraints on coefficients across equations. We compute the t-statistics of the parameter estimates accounting for heteroskedasticity and autocorrelation as well as contemporaneous cross-correlations in the errors from different equations. The estimation methodology can be regarded as an extension of the seemingly unrelated regression (SUR) method, the details of which are left in Appendix A.

Our estimation methodology with many portfolios is similar to Bollerslev, Engle, and Wooldridge (1988). They estimate the risk-return relation among the market portfolio, Treasury bonds, and Treasury bills using the time-varying conditional covariances. Their approach is based on a multivariate generalization of the GARCH-in-mean model originally proposed by Engle, Lilien, and Robins (1987).

4. Evidence from G7 country index portfolios

In the simultaneous equations in (15), under each currency denomination k, the common coefficient A^k measures the average relative risk aversion of the investors in the kth economy who can invest globally but nevertheless consumes mainly in the local currency. The common coefficient B^k measures the potential pricing of the residual risk in each country in additional to the global risk factor. Furthermore, we allow the intercept to be different for different portfolios, capturing the part of the excess return on each portfolio that cannot be explained by its conditional covariances with the global market portfolio and its country-specific residual risks. Thus, the intercepts measure the abnormal excess returns on different portfolios.

Under the assumption of financial integration, we first estimate the simple intertemporal relation between global risk factor and the country-index returns under each currency denomination k,

$$R^k_{i,t+1} = C^k_i + A^k \sigma^k_{im,t+1} + e^k_{t+1}, \quad i = 1, 2, \cdots, 7, \tag{16}$$

where the common slope estimate, A^k, of the simultaneous equations measures the average relative risk aversion of investors in the kth economy and C^k_i measures the abnormal returns on each portfolio i under each currency denomination k. Table 4 reports the parameter estimates and the t-statistics under each of the seven currency denominations.

The common slope estimates for A^k are all positive and statistically significant, suggesting that the global risk factor is priced across the G7 countries. Nevertheless, the magnitudes of the estimates differ dramatically under different currency denominations. The slope estimate is the highest at 7.2122 under yen denomination, lowest at 3.2333 under lira denomination. Broadly speaking, we can divide the estimates into three geographic groups. Japan as the representative of the Asian market has the most risk averse investors. The United States and Canada represent North American market, that generates risk aversion estimates between five and six. Finally, the four European economies all generate slope estimates less than four. The different magnitudes suggest that the exchange rate is priced and that investors in different economies have different risk preferences. Of course, with the introduction of a common currency (euro) among Germany, France, and Italy, we will lose the capability of identifying the risk preferences among investors in these countries using data after 1999. The ranking of our estimates is in line with estimates obtained from other approaches.

On the other hand, the intercept estimates are all small and statistically insignificant regardless of the currency denominations. The insignificant estimates of the intercepts suggest that one global risk factor explains well the cross-sectional differences in country portfolio returns. The abnormal excess returns from each country portfolios are not statistically different from zero.

To investigate whether the country-specific risks are priced in addition to the global risk factor, we also estimate the following simultaneous equations:

$$R^k_{i,t+1} = C^k_i + A^k\sigma^k_{im,t+1} + B^k\omega^k_{i,t+1} + e^k_{t+1}, \quad i = 1, 2, \cdots, 7, \tag{17}$$

where $\omega^k_{i,t}$ measures the conditional covariance between the country-i portfolio excess return and that country's residual risk, defined as the residual from regressing the country portfolio return on the global market portfolio return. If the common slope estimates of B^k are significant under each currency denomination, it would suggest that the country-specific risks are also priced. Table 5 reports the results on the new estimation under each of the seven currency denominations.

The estimates on A^k are similar to those obtained in Table 4, showing the robustness of the coefficient estimates. In particular, the ranking across different currency denominations remain about the same, with Japanese consumers being the most risk averse and Italian consumers the least risk averse. Also similar to the results in Table 4, the abnormal returns are not significantly different from zero.

Here, we are more interested in the slope estimates on the country-specific risk factors. The estimates on the country-specific risks are not statistically significant under any of the seven currency denominations. The insignificant estimates for B^k suggest that only the global risk factor is priced but the country-specific risks are not priced. This also provides evidence that the financial markets among the G7 countries are integrated. Investors can invest globally with little cost. Nevertheless, the significantly different slope estimates under different currency denominations indicate that consumers in different geographic areas are still limited to consume mainly in their respectively local currencies, suggesting limitations in labor movements. A Japanese investor's flying from Japan to Italy to consume in euro is still a very costly practice.

For our purpose, this limitation is actually useful in that the nominal exchange rates between these currencies are not trivial numeraire exchanges, but are informative about the differences in relative risk aversions of consumers at different locations. Had consumers had the capacity of moving freely and costlessly from Japan to Italy, purchasing power parity shall always hold and the nominal exchange rate should just reflect the difference in their respective monetary policies. We then would not be able to extract risk preference information for the exchange rate dynamics.

5. Evidence from global industry portfolios

Similar to the findings of earlier studies, we find that the G7 countries are financially integrated. Different from them, we also find that the risk-return tradeoff differs dramatically under different currency denominations. The differences in coefficient estimates provide evidence about the relative risk aversions of consumers that live at different geographic areas.

Given the financial integration, we do not need to form country-specific portfolios in our investment, nor do we need to form country-specific portfolios in estimating the risk-return relation. It is the currency denomination that generates the different coefficient estimates, not how the portfolios are formed. In principle, any ways of forming portfolios should generate similar results for the risk-return relation. To test this principle and to check the robustness of our findings, in this section, we estimate the risk-return relation using 25 industry portfolios. Datastream provides industry portfolio returns with a 25 industry partition for each of the seven countries. We aggregate the portfolio across country to obtain the average G7 return for each industry. The estimation proceeds much the same way except that now the system of simultaneous equations has 25 equations instead of seven in the country-specific portfolio case.

(to come ...)

6. Robustness analysis

To check the robustness of our findings, we consider variations in the conditional covariance estimation. Since the conditional variance and covariance of stock market returns are not observable, different approaches and specifications used in estimating the conditional variance and covariance could lead to different conclusions. In this paper, we use the bivariate GARCH(1,1) specification in equations (6) to

(10) to obtain conditional variance and covariance estimates. In this section, we study how varying the specification influences our results on the intertemporal risk-return relation.

The literature has considered two major variations on the conditional variance estimation. One incorporates exogenous instrumental variables such as short-term interest rates in the conditional variance forecasting equation (Campbell (1987)). The other is to allow downward return movements and upward return movements to have different impacts on the conditional volatility forecasts (Glosten, Jagannathan, and Runkle (1993) and Ding, Granger, and Engle (1993)). To investigate whether such variations in the variance forecasting specification alter our conclusion, we re-estimate the conditional variance and covariance using the following alternative specification,

$$R_{i,t+1} = \alpha_0^i + \alpha_1^i R_{i,t} + \varepsilon_{i,t+1}, \tag{18}$$

$$R_{m,t+1} = \alpha_0^m + \alpha_1^m R_{m,t} + \varepsilon_{m,t+1}, \tag{19}$$

$$\sigma_{i,t+1}^2 = \gamma_0^i + \gamma_1^i \varepsilon_{i,t}^2 + \gamma_2^i \sigma_{i,t}^2 + \gamma_3^i \varepsilon_{i,t}^2 I_{i,t}^- \tag{20}$$

$$\sigma_{m,t+1}^2 = \gamma_0^m + \gamma_1^m \varepsilon_{m,t}^2 + \gamma_2^m \sigma_{j,t}^2 + \gamma_3^j \varepsilon_{m,t}^2 I_{m,t}^-, \tag{21}$$

$$\sigma_{im,t+1} = \rho_{im} \sigma_{i,t+1} \sigma_{m,t+1}, \tag{22}$$

where $I_{i,t}^-$ is an indicator function that equals one when $\varepsilon_{i,t}$ is negative and zero otherwise. The indicator function generates an asymmetric GARCH effect between positive and negative shocks. In this specification, we assume constant correlation following Bollerslev (1990). The parameters in equations (18) to (22) are estimated by maximizing the log-likelihood function given in equation (13).

Table 6 presents the maximum likelihood parameter estimates and the t-statistics under each of the seven currency denominations. Similar to our earlier findings from the general covariance specification, the common slope estimates for A^k in Table 6 are all positive and statistically significant, suggesting that the global risk factor is priced across the G7 countries. However, the magnitudes of A^k differ considerably under different currency denominations. The slope estimate is again the highest at 7.8077 under yen denomination, and again the lowest at 3.2018 under lira denomination. The different magnitudes of relative risk aversion coefficients imply that the exchange rate is priced and that investors in different economies have different risk preferences. Another notable point in Table 6 is that the intercepts are all small and statistically insignificant regardless of the currency denominations. The insignificant estimates of the intercepts suggest that one global risk factor explains well the cross-sectional differences in country portfolio returns.

To examine whether the country-specific risks are priced in addition to the global risk factor, we estimate the simultaneous equations including the country's residual risk. Table 7 reports the maximum likelihood estimates of the common slope coefficients A^k and B^k under each currency denomination. The estimates of A^k are positive and significant without any exception. Their magnitudes are slightly greater than those reported in Table 6. Overall, the ranking across different currency denominations remain the same, with Japanese consumers being the most risk averse and Italian consumers the least risk averse. Also similar to the results in Table 6, the abnormal returns (or intercepts) are not significantly different from zero.

Similar to our earlier findings from the general covariance specification, the estimates on the country-specific risks are not statistically significant under any of the seven currency denominations. The insignificant estimates for B^k reported in Table 7 suggest that only the global risk factor is priced but the country-specific risks are not priced. This also provides evidence that the financial markets among the G7 countries are integrated. Investors can invest globally with little cost, but they are still limited to consume mainly in their respectively local currencies.

7. Conclusion

This paper investigates the role of currency denomination in the the intertemporal risk-return relation among G7 countries. Similar to the literature findings, our estimation also shows that the financial markets of the G7 countries are integrated. We obtain significant pricing coefficient estimates on the global index, but insignificant estimates on country-specific risks. Different from the literature, however, we find that the intertemporal risk-return relation differ significantly under different currency denominations. The slope coefficient estimate is the largest at around seven when the returns are denominated in Japanese yen, smallest at around three to four when the returns are denominated in pound or euro and its predecessors. The slope estimates are in the middle at about five to six when the returns are denominated in the U.S. or Canadian dollar. The estimates stay in the same range and the rankings remain unchanged when we consider different specifications for the conditional covariance estimations and when we replace country-specific portfolios with industry-specific forms in estimating the intertemporal relation.

Appendix A. Estimation of a system of simultaneous equations

Consider a system of n simultaneous equations, of which the typical ith equation is

$$y_i + X_i\beta_i + u_i, \tag{A1}$$

where y_i is a $N \times 1$ vector of time-series observations on the ith dependent variable, X_i is a $N \times k_i$ matrix of observations of k_i independent variables, β_i is a $k_i \times 1$ vector of unknown regression coefficients to be estimated, and u_i is a $N \times 1$ vector of random disturbance terms with mean zero. Parks (1967) proposes an estimation procedure that allows the error term to be both serially and cross-sectionally correlated.

Parks (1967) assumes that the elements of the disturbance vector u follow an AR(1) process:

$$u_{it} = \rho_i u_{it-1} + \varepsilon_{it}, \quad |\rho_i| < 1, \tag{A2}$$

where ε_{it} are random variables satisfying the conditions:

$$\begin{array}{lcllll} \mathbb{E}(\varepsilon_{it}) & = & 0, & i = 1, \cdots, n; & t = 1, \cdots, N, & \\ \mathbb{E}(\varepsilon_{it}\varepsilon_{jt}^\top) & = & \sigma_{ij}, & i, j = 1, \cdots, n; & t = 1, \cdots, N, & \\ \mathbb{E}(\varepsilon_{it}\varepsilon_{js}) & = & 0, & i, j = 1, \cdots, n; & s, t = 1, \cdots, N, & \texttt{and}\ s \neq t. \end{array} \tag{A3}$$

Equation (A1) can then be written as

$$y_i = X_i\beta_i + P_i\varepsilon_i, \tag{A4}$$

with $u_i = P_i\varepsilon_i$, where ε_i is a $N \times 1$ random vector with $\mathbb{E}(\varepsilon_i) = 0$ and $\mathbb{E}(\varepsilon_i\varepsilon_i^\top) = \sigma_{ii}I$ and

$$P_i = \begin{bmatrix} (1-\rho_i^2)-1/2 & 0 & 0 & \cdots & 0 \\ \rho_i(1-\rho_i^2)-1/2 & 1 & 0 & \cdots & 0 \\ \rho_i^2(1-\rho_i^2)-1/2 & \rho_i & 1 & \cdots & 0 \\ \vdots & & & & \\ \rho_i^{N-1}(1-\rho_i^2)-1/2 & \rho^{N-2} & \rho^{N-3} & \cdots & 1 \end{bmatrix}. \tag{A5}$$

Under this setup, Parks presents a consistent and asymptotically efficient three-step estimation technique for the regression coefficients. The first step uses single equation regressions to estimate the parameters of autoregressive model. The second step uses single equation regressions on transformed equations to estimate the contemporaneous covariances. Finally, the Aitken estimator is formed using the estimated covariance,

$$\widehat{b} = \left(X^\top\Omega^{-1}X\right)^{-1} X^\top\Omega^{-1}y, \tag{A6}$$

where $\Omega \equiv \mathbb{E}[uu^\top]$ denotes the general covariance matrix of the innovation.

In our application, we use the aforementioned methodology with the slope coefficients restricted to be the same for all portfolios. In particular, we use the same three-step procedure and the same covariance assumptions as in (A2) to (A5) to estimate the covariances and to generate the t-statistics for the parameter estimates.

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Table 1

Summary statistics on country and global portfolio returns under different currency denominations

Entries report the mean (in annual percentage), standard deviation (Std, in annual percentage), skewness, excess kurtosis, the Jarque-Bera (JB) test statistics for return distribution normality, and the first-order monthly autocorrelation (Auto) of the country and global portfolio returns under each of the seven currency denominations. Data are monthly from January 1970 to December 2004.

Country/Currency	Yen	C. Dollar	Dollar	Mark	Pound	Franc	Lira
Japan:							
Mean (%)	3.32	3.05	4.61	3.20	2.70	5.32	3.29
Std (%)	18.38	22.10	22.29	22.09	22.03	22.02	22.23
Skewness	-0.28	0.05	0.01	-0.07	-0.08	-0.11	-0.07
Kurtosis	1.30	0.72	0.42	0.62	0.40	0.67	0.47
JB	35.07	9.29	3.13	7.05	3.29	8.88	4.13
Auto	0.07	0.10	0.10	0.12	0.15	0.12	0.12
Canada:							
Mean (%)	2.42	2.15	3.71	2.30	1.80	4.43	2.39
Std (%)	21.63	17.34	19.41	21.87	20.94	21.39	21.45
Skewness	-0.61	-0.79	-0.82	-0.54	-0.76	-0.50	-0.44
Kurtosis	2.41	2.94	2.86	1.82	2.57	1.78	1.74
JB	127.93	195.78	190.42	78.88	155.76	73.25	66.96
Auto	0.01	0.07	0.05	0.09	0.05	0.08	0.08
US:							
Mean (%)	2.98	2.71	4.27	2.86	2.36	4.98	2.95
Std (%)	19.11	15.13	15.59	19.28	18.71	18.93	19.01
Skewness	-0.57	-0.51	-0.58	-0.70	-0.65	-0.62	-0.46
Kurtosis	2.14	2.29	2.40	2.28	2.41	1.89	2.06
JB	102.91	110.08	124.98	125.30	131.45	89.04	88.99
Auto	0.00	0.02	0.02	0.08	0.05	0.06	0.04
Germany:							
Mean (%)	2.83	2.56	4.11	2.70	2.20	4.83	2.79
Std (%)	22.28	21.20	21.61	20.08	21.24	20.47	20.83
Skewness	-0.42	-0.44	-0.51	-0.76	-0.65	-0.69	-0.56
Kurtosis	1.63	1.39	1.82	3.04	2.69	2.63	2.31
JB	58.53	47.45	76.53	202.76	156.47	154.53	115.34
Auto	0.01	0.01	-0.01	0.06	0.03	0.02	0.04

Table 1 (continued)

Country/Currency	Yen	C. Dollar	Dollar	Mark	Pound	Franc	Lira
UK:							
Mean (%)	3.89	3.62	5.18	3.77	3.27	5.89	3.86
Std (%)	22.71	21.76	22.19	22.60	20.20	21.84	22.69
Skewness	0.17	0.54	0.46	0.01	0.25	0.10	0.04
Kurtosis	5.10	6.63	5.82	5.42	8.60	5.44	5.98
JB	458.83	791.85	608.28	516.08	1302.73	519.41	626.41
Auto	0.06	0.07	0.07	0.13	0.08	0.11	0.07
France:							
Mean (%)	3.73	3.46	5.01	3.60	3.11	5.73	3.69
Std (%)	22.55	22.23	22.66	21.59	22.09	20.75	21.96
Skewness	-0.33	-0.27	-0.36	-0.32	-0.31	-0.35	-0.28
Kurtosis	0.98	1.62	1.55	1.25	1.92	1.22	1.21
JB	24.60	51.01	51.36	34.78	71.41	34.87	31.35
Auto	0.05	0.09	0.08	0.11	0.10	0.08	0.09
Italy:							
Mean (%)	-0.02	-0.29	1.27	-0.15	-0.64	1.98	-0.05
Std (%)	26.04	25.06	25.40	25.98	25.32	25.60	23.84
Skewness	-0.09	0.10	-0.03	0.01	0.20	0.13	0.17
Kurtosis	0.52	0.64	0.62	1.05	0.48	0.72	0.75
JB	5.29	7.94	6.80	19.33	6.90	10.40	11.82
Auto	-0.01	0.06	0.05	0.02	0.01	0.01	0.06
World:							
Mean (%)	2.72	2.45	4.01	2.60	2.10	4.73	2.69
Std (%)	15.69	13.99	14.62	16.44	15.79	16.10	16.41
Skewness	-0.79	-0.50	-0.62	-0.75	-0.84	-0.72	-0.52
Kurtosis	2.39	1.42	1.65	1.97	2.38	1.65	1.93
JB	143.61	52.69	75.30	107.64	148.86	83.70	84.32
Auto	0.03	0.09	0.08	0.14	0.12	0.12	0.10

Table 2

Maximum likelihood estimates on the degrees of freedom parameter for the Student-t distribution

For each country, the first row reports the maximum likelihood estimates of the degrees of freedom (or tail-thickness) parameter v for the Student-t distribution for each country and global market portfolio returns under each of the seven currency denominations. The second row presents the corresponding t statistics of the estimated v values in parentheses. The third row shows the Wald statistics from testing whether $1/v = 0$, which is a test of normality against the fat-tailedness of the return distribution. The fourth row gives the p-values of the Wald statistics in parentheses. The estimation is based on monthly returns (in percentages) from January 1970 to December 2004.

Country	Statistics	Yen	C. Dollar	Dollar	Mark	Pound	Franc	Lira
Japan	Estimates	6.3888	15.4366	26.9978	11.1130	14.9202	12.0571	14.4569
	t-values	(2.96)	(1.30)	(0.67)	(1.90)	(1.20)	(1.88)	(1.54)
	Wald test	8.7766	1.6852	0.4518	3.6254	1.4339	3.5181	2.3726
	p-values	(0.00)	(0.19)	(0.50)	(0.06)	(0.23)	(0.06)	(0.12)
Canada	Estimates	5.9192	6.1275	5.4675	8.1430	6.5392	7.8944	7.4976
	t-values	(3.65)	(4.09)	(3.92)	(3.33)	(3.81)	(3.27)	(3.32)
	Wald test	13.2930	16.7253	15.3841	11.0724	14.5356	10.6826	11.0522
	p-values	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
US	Estimates	6.5241	8.2923	6.9545	6.3356	6.6194	6.9381	5.2314
	t-values	(3.74)	(3.66)	(3.72)	(4.02)	(4.02)	(3.70)	(3.90)
	Wald test	13.9666	13.4164	13.8448	16.1837	16.1766	13.6635	15.1796
	p-values	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Germany	Estimates	8.3728	6.4173	6.4108	6.4647	6.1177	6.0007	6.5277
	t-values	(2.48)	(2.95)	(2.88)	(3.19)	(3.34)	(3.36)	(3.07)
	Wald test	6.1606	8.7209	8.3120	10.1480	11.1825	11.2731	9.4371
	p-values	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
UK	Estimates	6.9682	6.5172	6.0979	5.3958	5.0589	5.3416	5.1642
	t-values	(3.64)	(3.64)	(3.50)	(4.45)	(4.72)	(4.45)	(4.47)
	Wald test	13.2805	13.2316	12.2156	19.8045	22.3110	19.8233	19.9458
	p-values	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
France	Estimates	10.2321	7.1106	6.8773	9.6968	6.7369	11.5376	9.0810
	t-values	(2.41)	(2.80)	(2.63)	(2.86)	(3.28)	(2.41)	(2.58)
	Wald test	5.7997	7.8618	6.9138	8.1932	10.7906	5.8215	6.6793
	p-values	(0.02)	(0.01)	(0.01)	(0.00)	(0.00)	(0.02)	(0.01)
Italy	Estimates	10.6516	12.2473	11.0756	6.6932	15.2156	9.5763	10.8801
	t-values	(1.69)	(1.70)	(1.85)	(2.60)	(1.25)	(1.67)	(1.50)
	Wald test	2.8588	2.8785	3.4080	6.7478	1.5692	2.7749	2.2489
	p-values	(0.09)	(0.09)	(0.06)	(0.01)	(0.21)	(0.10)	(0.13)
World	Estimates	5.8358	8.8898	5.9330	6.8865	6.6998	6.9954	5.4591
	t-values	(3.91)	(2.96)	(3.58)	(3.62)	(3.94)	(3.55)	(4.15)
	Wald test	15.2906	8.7578	12.8184	13.1096	15.5205	12.6098	17.2610
	p-values	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

Table 3

Maximum likelihood estimates of the conditional covariances

Entries report the maximum likelihood estimates of the parameters that govern the dynamics of the conditional covariance between the excess returns on country-index portfolios and global market portfolios,

$$\sigma_{im,t+1} = \gamma_0^{im} + \gamma_1^{im}\varepsilon_{i,t}\varepsilon_{m,t} + \gamma_2^{im}\sigma_{im,t},$$

for the G7 countries under each of the seven currency denominations. The t-statistics are in parentheses. The estimation is based on monthly returns (in percentages) from January 1970 to December 2004.

Country	Estimates	Yen	C. Dollar	Dollar	Mark	Pound	Franc	Lira
Japan	γ_0^{im}	0.0002 (2.03)	0.0003 (2.11)	0.0001 (1.85)	0.0001 (1.97)	0.0003 (2.31)	0.0001 (1.88)	0.0001 (1.93)
	γ_1^{im}	0.0383 (2.92)	0.0395 (3.40)	0.0165 (2.17)	0.0151 (2.83)	0.0154 (3.24)	0.0113 (2.79)	0.0143 (2.62)
	γ_2^{im}	0.9529 (48.54)	0.9432 (39.87)	0.9563 (49.79)	0.9634 (41.44)	0.9311 (46.91)	0.9456 (50.12)	0.9196 (47.91)
Canada	γ_0^{im}	0.0006 (3.09)	0.0002 (2.20)	0.0002 (2.42)	0.0002 (2.02)	0.0003 (2.32)	0.0003 (2.58)	0.0004 (2.06)
	γ_1^{im}	0.0467 (2.15)	0.0552 (2.34)	0.0599 (2.74)	0.0505 (3.09)	0.0494 (3.01)	0.0608 (3.94)	0.0491 (3.83)
	γ_2^{im}	0.6931 (6.93)	0.8165 (11.01)	0.8150 (12.56)	0.8456 (10.71)	0.8223 (13.99)	0.8104 (16.70)	0.8305 (14.57)
US	γ_0^{im}	0.0001 (2.46)	0.0001 (4.36)	0.0001 (4.57)	0.0001 (3.05)	0.0001 (2.85)	0.0001 (2.84)	0.0001 (2.47)
	γ_1^{im}	0.0123 (3.37)	0.0791 (5.31)	0.0807 (4.61)	0.0480 (5.33)	0.0612 (4.13)	0.0653 (5.23)	0.0568 (5.38)
	γ_2^{im}	0.9625 (70.22)	0.8504 (35.86)	0.8237 (32.26)	0.8813 (30.99)	0.8640 (29.87)	0.8901 (31.25)	0.8754 (27.09)
Germany	γ_0^{im}	0.0007 (3.93)	0.0005 (1.85)	0.0003 (1.87)	0.0005 (2.04)	0.0004 (1.92)	0.0004 (2.21)	0.0005 (1.94)
	γ_1^{im}	0.0367 (4.86)	0.0774 (2.11)	0.0940 (2.79)	0.1003 (4.97)	0.0810 (3.00)	0.0876 (3.46)	0.0930 (3.12)
	γ_2^{im}	0.6392 (6.39)	0.6122 (6.34)	0.6435 (6.93)	0.6943 (8.34)	0.6782 (7.31)	0.6889 (7.90)	0.6533 (7.91)
UK	γ_0^{im}	0.0001 (2.68)	0.0001 (2.75)	0.0001 (2.50)	0.0001 (2.94)	0.0001 (2.84)	0.0001 (2.30)	0.0001 (2.78)
	γ_1^{im}	0.0196 (3.11)	0.0276 (4.35)	0.0264 (4.06)	0.0182 (3.07)	0.0276 (3.94)	0.0151 (3.31)	0.0248 (3.26)
	γ_2^{im}	0.9307 (39.21)	0.9502 (46.31)	0.9536 (80.54)	0.9557 (68.21)	0.9463 (65.82)	0.9656 (87.71)	0.9497 (69.05)
France	γ_0^{im}	0.0001 (2.57)	0.0006 (2.30)	0.0004 (2.00)	0.0003 (1.81)	0.0005 (1.98)	0.0005 (1.79)	0.0002 (2.09)
	γ_1^{im}	0.0478 (3.04)	0.0911 (3.21)	0.0853 (3.12)	0.0698 (3.12)	0.0744 (2.95)	0.0703 (2.60)	0.0682 (3.22)
	γ_2^{im}	0.6201 (6.10)	0.6312 (6.66)	0.6501 (7.10)	0.6389 (7.54)	0.6579 (6.23)	0.6411 (6.48)	0.6627 (6.43)
Italy	γ_0^{im}	0.0002 (2.48)	0.0007 (2.41)	0.0002 (1.91)	0.0003 (1.90)	0.0003 (2.31)	0.0004 (2.00)	0.0005 (1.82)
	γ_1^{im}	0.0392 (3.14)	0.1012 (4.22)	0.0787 (2.80)	0.0831 (3.05)	0.0992 (3.56)	0.0735 (2.97)	0.0603 (3.11)
	γ_2^{im}	0.6007 (6.00)	0.6100 (6.81)	0.6780 (7.04)	0.6739 (8.16)	0.6111 (7.77)	0.6671 (7.03)	0.6890 (8.06)

Table 4

The risk-return relation under different currency denominations with general covariance

Entries report the estimates of the following simultaneous equations under each currency denominations,

$$R^k_{i,t+1} = C^k_i + A^k \sigma^k_{im,t+1} + e^k_{i,t+1}, \quad i = 1, 2, \cdots, 7$$

for $k = 1, \cdots, 7$, where $R^k_{i,t+1}$ denotes the excess return in the country-i index portfolio denominated in the kth currency. $\sigma^k_{im,t+1}$ is estimated using the general covariance specification. The t-statistics are in parentheses. The estimation is based on monthly returns (in percentages) from January 1970 to December 2004.

Country/Currency	Yen	C. Dollar	Dollar	Mark	Pound	Franc	Lira
Common Slope:							
A^k	7.2122	5.7460	5.0346	3.7301	3.7279	3.6884	3.2333
	(2.53)	(2.51)	(2.65)	(2.22)	(2.49)	(2.40)	(2.31)
Intercepts: C^k_i							
Japan	-0.0030	-0.0092	-0.0068	-0.0119	-0.0075	0.0015	-0.0039
	(-0.69)	(-1.41)	(-1.28)	(-1.58)	(-1.31)	(0.23)	(-0.70)
Canada	-0.0086	-0.0065	-0.0074	-0.0150	-0.0098	-0.0010	-0.0058
	(-1.20)	(-1.50)	(-1.47)	(-1.77)	(-1.56)	(-0.14)	(-0.95)
US	-0.0079	-0.0069	-0.0060	-0.0143	-0.0094	0.0004	-0.0051
	(-1.13)	(-1.53)	(-1.34)	(-1.71)	(-1.50)	(0.05)	(-0.86)
Germany	-0.0065	-0.0081	-0.0065	-0.0012	-0.0073	-0.0026	-0.0035
	(-1.03)	(-1.46)	(-1.29)	(-0.22)	(-1.33)	(-0.49)	(-0.69)
UK	-0.0064	-0.0079	-0.0067	-0.0112	-0.0058	-0.0008	-0.0040
	(-0.95)	(-1.40)	(-1.23)	(-1.50)	(-1.14)	(-0.12)	(-0.68)
France	-0.0060	-0.0082	-0.0069	-0.0068	-0.0074	-0.0078	-0.0035
	(-0.95)	(-1.44)	(-1.26)	(-1.17)	(-1.27)	(-1.26)	(-0.65)
Italy	-0.0082	-0.0091	-0.0078	-0.0107	-0.0091	-0.0048	-0.0048
	(-1.33)	(-1.74)	(-1.56)	(-1.70)	(-1.65)	(-0.82)	(-1.03)

Table 5
The pricing of country-specific risks with general covariance
Entries report the estimates of the following simultaneous equations under each currency denominations,

$$R^k_{i,t+1} = C^k_i + A^k \sigma^k_{im,t+1} + B^k \omega^k_{i,t+1} e^k_{i,t+1}, \quad i = 1, 2, \cdots, 7$$

for $k = 1, \cdots, 7$, where $R^k_{i,t+1}$ denotes the excess return in the country-i index portfolio denominated in the kth currency, $\sigma^k_{im,t+1}$ measures the conditional covariance between the excess return and the global market portfolio, and $\omega^k_{i,t+1}$ measures the conditional covariance between the country portfolio excess return and the country-specific risk $E^k_{i,t+1}$, which is defined as the residual by regressing the country portfolio return on the global portfolio return. $\sigma^k_{im,t+1}$ and $\omega^k_{i,t+1}$ are estimated using the general covariance specification. The t-statistics are in parentheses. The estimation is based on monthly returns (in percentages) from January 1970 to December 2004.

Country/Currency	Yen	C. Dollar	Dollar	Mark	Pound	Franc	Lira
Common Slopes:							
A^k	8.6432	6.0107	5.8178	3.8002	3.6433	3.7115	3.2901
	(2.89)	(2.52)	(2.61)	(2.30)	(2.33)	(2.43)	(2.32)
B^k	1.9840	0.8172	0.9489	0.0674	1.2012	0.7858	0.0013
	(1.18)	(0.99)	(1.00)	(0.01)	(1.20)	(0.89)	(0.00)
Intercepts: C^k_i							
Japan	-0.0050	-0.0076	-0.0026	-0.0054	-0.0022	-0.0034	-0.0039
	(-0.56)	(-1.40)	(-0.39)	(-0.91)	(-0.30)	(-0.56)	(-0.70)
Canada	-0.0079	-0.0050	-0.0087	-0.0079	-0.0044	-0.0056	-0.0058
	(-1.04)	(-0.90)	(-1.69)	(-1.16)	(-0.57)	(-0.82)	(-0.94)
US	-0.0070	-0.0066	-0.0090	-0.0072	-0.0036	-0.0050	-0.0052
	(-0.91)	(-1.47)	(-1.72)	(-1.08)	(-0.45)	(-0.74)	(-0.84)
Germany	-0.0061	-0.0071	-0.0053	-0.0043	-0.0029	-0.0026	-0.0035
	(-0.93)	(-1.40)	(-1.03)	(-0.86)	(-0.43)	(-0.50)	(-0.69)
UK	-0.0059	-0.0072	-0.0057	-0.0056	-0.0049	-0.0031	-0.0040
	(-0.85)	(-1.33)	(-1.05)	(-0.89)	(-0.94)	(-0.50)	(-0.68)
France	-0.0057	-0.0075	-0.0057	-0.0047	-0.0032	-0.0022	-0.0035
	(-0.86)	(-1.37)	(-1.02)	(-0.84)	(-0.47)	(-0.40)	(-0.64)
Italy	-0.0081	-0.0084	-0.0058	-0.0077	-0.0051	-0.0054	-0.0044
	(-1.31)	(-1.69)	(-1.07)	(-1.30)	(-0.80)	(-0.92)	(-0.47)

Table 6

The risk-return relation under different currency denominations with constant correlation

Entries report the estimates of the following simultaneous equations under each currency denominations,

$$R^k_{i,t+1} = C^k_i + A^k \sigma^k_{im,t+1} + e^k_{i,t+1}, \quad i = 1, 2, \cdots, 7$$

for $k = 1, \cdots, 7$, where $R^k_{i,t+1}$ denotes the excess return in the country-i index portfolio denominated in the kth currency. $\sigma^k_{im,t+1}$ is estimated using the constant correlation specification. The t-statistics are in parentheses. The estimation is based on monthly returns (in percentages) from January 1970 to December 2004.

Country/Currency	Yen	C. Dollar	Dollar	Mark	Pound	Franc	Lira
Common Slope:							
A^k	7.8077	5.8114	5.1254	3.6535	3.7033	3.3876	3.2018
	(2.51)	(2.49)	(2.63)	(2.20)	(2.32)	(2.32)	(2.21)
Intercepts: C^k_i							
Japan	-0.0062	-0.0076	-0.0062	-0.0037	-0.0082	-0.0030	-0.0031
	(-0.99)	(-1.26)	(-1.42)	(-0.59)	(-1.41)	(-0.51)	(-0.59)
Canada	-0.0047	-0.0082	-0.0068	-0.0024	-0.0048	-0.0012	-0.0019
	(-0.78)	(-1.17)	(-1.29)	(-0.40)	(-1.02)	(-0.23)	(-0.37)
US	-0.0068	-0.0073	-0.0075	-0.0044	-0.0085	-0.0035	-0.0037
	(-1.06)	(-1.37)	(-1.45)	(-0.69)	(-1.17)	(-0.59)	(-0.68)
Germany	-0.0069	-0.0093	-0.0080	-0.0050	-0.0081	-0.0039	-0.0032
	(-1.21)	(-1.34)	(-1.31)	(-0.89)	(-1.10)	(-0.73)	(-0.81)
UK	-0.0022	-0.0087	-0.0070	-0.0025	-0.0065	-0.0016	-0.0020
	(-0.53)	(-1.29)	(-1.34)	(-0.45)	(-1.20)	(-0.31)	(-0.41)
France	-0.0050	-0.0081	-0.0067	-0.0019	-0.0062	-0.0011	-0.0018
	(-0.88)	(-1.40)	(-1.27)	(-0.39)	(-1.23)	(-0.24)	(-0.40)
Italy	-0.0046	-0.0086	-0.0071	-0.0020	-0.0063	-0.0007	-0.0013
	(-0.79)	(-1.21)	(-1.33)	(-0.37)	(-1.15)	(-0.14)	(-0.35)

Table 7
The pricing of country-specific risks with constant correlation
Entries report the estimates of the following simultaneous equations under each currency denominations,

$$R^k_{i,t+1} = C^k_i + A^k \sigma^k_{im,t+1} + B^k \omega^k_{i,t+1} e^k_{i,t+1}, \quad i = 1, 2, \cdots, 7$$

for $k = 1, \cdots, 7$, where $R^k_{i,t+1}$ denotes the excess return in the country-i index portfolio denominated in the kth currency, $\sigma^k_{im,t+1}$ measures the conditional covariance between the excess return and the global market portfolio, and $\omega^k_{i,t+1}$ measures the conditional covariance between the country portfolio excess return and the country-specific risk $E^k_{i,t+1}$, which is defined as the residual by regressing the country portfolio return on the global portfolio return. $\sigma^k_{im,t+1}$ and $\omega^k_{i,t+1}$ are estimated using the constant correlation specification. The t-statistics are in parentheses. The estimation is based on monthly returns (in percentages) from January 1970 to December 2004.

Country/Currency	Yen	C. Dollar	Dollar	Mark	Pound	Franc	Lira
Common Slopes:							
A^k	9.1420	6.1548	6.0425	3.7178	3.6785	3.3645	3.3880
	(3.00)	(2.49)	(2.60)	(2.37)	(2.29)	(2.30)	(2.29)
B^k	2.1129	1.4282	2.0784	-0.7502	1.8081	1.2597	0.0193
	(1.15)	(0.39)	(1.16)	(-0.80)	(1.02)	(0.98)	(0.01)
Intercepts: C^k_i							
Japan	-0.0057	-0.0078	-0.0090	-0.0033	-0.0069	-0.0026	-0.0030
	(-0.81)	(-1.10)	(-1.09)	(-0.49)	(-0.93)	(-0.47)	(-0.56)
Canada	-0.0044	-0.0087	-0.0057	-0.0050	-0.0047	-0.0010	-0.0019
	(-0.70)	(-1.05)	(-1.06)	(-1.05)	(-0.71)	(-0.18)	(-0.37)
US	-0.0064	-0.0058	-0.0087	-0.0041	-0.0075	-0.0032	-0.0035
	(-0.93)	(-1.21)	(-1.24)	(-0.62)	(-1.00)	(-0.56)	(-0.68)
Germany	-0.0068	-0.0098	-0.0058	-0.0049	-0.0061	-0.0037	-0.0033
	(-1.19)	(-1.35)	(-1.08)	(-0.87)	(-1.04)	(-0.70)	(-0.29)
UK	-0.0034	-0.0101	-0.0025	-0.0019	-0.0054	-0.0013	-0.0020
	(-0.43)	(-1.04)	(-0.39)	(-0.29)	(-0.81)	(-0.25)	(-0.41)
France	-0.0047	-0.0089	-0.0053	-0.0026	-0.0029	-0.0009	-0.0018
	(-0.79)	(-1.29)	(-1.03)	(-0.46)	(-0.16)	(-0.21)	(-0.40)
Italy	-0.0043	-0.0092	-0.0055	-0.0030	-0.0025	-0.0004	-0.0017
	(-0.72)	(-1.19)	(-1.05)	(-0.52)	(-0.46)	(-0.08)	(-0.33)