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ABSTRACT

This paper empirically examines multifactor asset pricing models for the returns and expected returns on eighteen national equity markets. The factors are chosen to measure global economic risks. Although previous studies do not reject the unconditional mean-variance efficiency of a world market portfolio, our evidence indicates that the tests are low in power, and the world market betas do not provide a good explanation of cross-sectional differences in average returns. Multiple beta models provide an improved explanation of the equity returns.

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1. Introduction

This paper studies the sources of risk and average returns in international equity markets. We examine several measures of global economic risks and ask to what extent these risk factors can explain the fluctuations in the stock markets of eighteen countries. The monthly risk measures include the returns on a world equity market portfolio, a measure of exchange risks, a Eurodollar - U.S. Treasury bill yield spread, and measures of global inflation, real interest rates, and industrial production growth. We find that the global risk factors can explain, ex-post, between 15% and 86% percent of the variance of the monthly returns over the 1970-1989 period. The world market portfolio is by far the most important factor. It alone can explain between 16% and 71% of the variance, depending on the country.

We examine the average return premiums associated with these global risks and we find significant premiums associated with the world equity index and a measure of exchange rate fluctuations, but no significant average premium associated with the other variables. Previous studies [e.g. Cumby and Glen (1990) and Harvey (1991a)] do not reject the unconditional mean-variance efficiency of the world market index. We confirm this finding in our sample using a general test. However, we find that the power of the world market betas to explain the average return differences across the countries is low. The average pricing errors of the model are reduced by introducing the additional global risk factors. The average performance of the Japanese equity market, for example, is much better explained by a model which incorporates multiple sources of risk.

The paper is organized as follows. Section 2 outlines the methodology. Section 3 describes the country returns and the global risk factor data. The empirical results are presented in section 4. Section 5 summarizes our conclusions.

2. Methodology

We first examine factor model regressions for the equity market index returns for eighteen countries:

$$r_{it} = \alpha_i + \Sigma_j{}^K{}_{=1}\, \beta_{ij}\, F_{jt} + u_{it}. \tag{1}$$

The *excess return* is $r_{it} = R_{it} - R_{ft}$, where R_{it} is the U.S. dollar return for country *i* at time *t*. R_{ft} is the dollar return of a one-month U.S. Treasury bill. (We also report results for excess returns in local currency units.) The β_{ij} are the betas of the r_{it} on the K risk factors F_j, $j=1,\ldots, K$. The error terms, u_{it} represent the "nonsystematic" excess returns, relative to the global risk factors.

We estimate most of our empirical models using Hansen's (1982) generalized method of moments (GMM), which is valid under mild statistical assumptions. In equation (1), we assume that the data vector, which is $\{r_{it}, i=1,\ldots,N, F_{jt}, j=1,\ldots,K\}$, $t=1,\ldots,T$, is generated by a strictly stationary and ergodic stochastic process. The α_i and the β_{ij} are fixed regression coefficients, which implies $E(u_{it})=E(u_{it}F_{jt})=0$. The error terms u_{it} are not assumed to be normally distributed, and the conditional variances of the u_{it}, given the F_{jt}, may depend on the values of the F_{jt}. We do not specify the functional form of this possible dependence. Instead, we report the asymptotic standard errors for the coefficients described by White (1980) and Hansen (1982). Hansen shows that the GMM coefficient estimators are consistent and asymptotically normal, derives their standard errors, and discusses the statistical assumptions more formally.

The factor model regressions provide information about the usefulness of global factors in controlling the risks of international equity investments. We are therefore interested in the factor models for their own sake. We are also interested in the relation between the risk sensitivities, β_{ij}, and the expected returns in the various national markets.

Asset pricing models attribute cross-sectional differences in expected returns to differences in betas. A general beta pricing model asserts the existence of expected premiums λ_j, j = 0,..., K; such that expected returns can be written as:

$$E(R_i) = \lambda_o + \Sigma_j{}^K{}_{=1}\, b_{ij}\, \lambda_j. \tag{2}$$

The b_{ij} are the betas (multiple regression coefficients) of the R_{it} on the K global risk factors j=1,..., K. Equation (2) implies an expression for the expected *excess* returns:

$$E(r_i) = \Sigma_j{}^K{}_{=1}\, \beta_{ij}\, \lambda_j. \tag{3}$$

Where $\beta_{ij} = b_{ij} - b_{fj}$ are the betas of the excess returns and the b_{fj} are the betas of the Treasury bill.

Beta pricing models for expected returns like equations (2) and (3) are familiar in a domestic context and are developed for an international setting by a number of authors. In order to apply a beta pricing model in a global setting, strong assumptions are needed. The national equity markets are assumed to be perfectly integrated in a global economy, with no barriers to extranational equity investments, no taxes, no transactions or information costs. Such extreme assumptions are unlikely to provide a good approximation to the actual complexity of international investments. Therefore, we interpret our results as a baseline case. Further refinements of the models, to incorporate additional considerations should produce even better explanatory power. Such refinements remain an important topic for future research.

The number and identity of the global risk factors takes on special significance in an international setting. We study models with a single factor and with multiple factors. The single factor model is a global version of the Capital Asset Pricing Model (CAPM) of Sharpe (1964) and Lintner (1965), which states that the world market portfolio is mean-variance efficient. Stulz (1981b, 1984) and Adler and Dumas (1983) provide conditions under which a single-beta CAPM based on the world market portfolio holds globally. The sufficient conditions are strong, including no exchange risk and a constant investment opportunity set, in addition to the assumptions described above.

When strict purchasing power parity fails to hold, then consumers face exchange risks for investing internationally, and exchange risks may be priced in a global asset pricing model. Adler and Dumas [1983, equation (14)] present a model in which a combination of the world market and measures of exchange risk is mean variance efficient. The exchange risk can be broken down into a separate factor for each currency, as in Dumas and Solnik (1992), or can be approximated by a single variable. We take the latter approach and study a two-beta model, using the world market portfolio and an aggregate of exchange risks as the two factors.

International equilibrium and APT models with several factors are described by Stulz (1981a), Hodrick (1981), Solnik (1983) and Ross and Walsh (1983), among others. A central intuition of such models is that the common sources of risk may command an expected return premium, while risks that can be diversified internationally should not. Korajczyk and Viallet (1989) and Heston, Rouwenhorst and Wessels (1991) find evidence of several common sources of variation in individual U.S. and European stocks. Given evidence for several common sources of variation, a number of world-wide risk factors may be important determinants of national equity market returns. We therefore study models with a number of global economic risk variables.

Equations (2) and (3) are stylized representation for a class of beta pricing models, and the content of the model is the discipline imposed in selecting the factors. Our approach is to choose the variables a priori and to investigate their importance using the factor model regressions (1). Then, we study the pricing of the most important risk factors. Our focus in this paper is on the relation between risk and long-run expected returns. That is, we investigate *unconditional* versions of the beta pricing models.[1]

We estimate and test the pricing equation (3) as a restricted seemingly unrelated regression model (SURM):

$$r_{it} = \Sigma_{j=1}^{K} \beta_{ij} (f_{jt} + \lambda_j) + u_{it}, \quad i=1,...,N, \tag{4}$$

where the f_j are the de-meaned values of the risk factors ($f_{jt} = F_{jt} - \bar{F}_j$ and $\bar{F}_j$ is the sample mean). The regression is restricted by assuming that the intercept is equal to zero. The theoretical model, equation (3), implies that $E(u_{it})=0$ in equation (4). The parameters to be estimated are the unconditional betas, β_{ij} and the expected risk premiums, λ_j. Since $E(f_{jt})=0$, we do not assume in equation (4) that the means of the factors F_j are related in any way to the expected risk premiums λ_j. This allows us to use economic variables as factors and to estimate and test the model without the need for mimicking portfolios for the factors.[2] This is an advantage over the approaches of Gibbons (1982), Gibbons, Breeden and Litzenberger (1989) and others, since the estimation of mimicking portfolios in a separate step can complicate the statistical inferences [see Wheatley (1989)].

We implement the SURM via the GMM. We therefore assume that the data vector $\{r_{it}, i=1,...,N, f_{jt}, j=1,...,K\}, t=1,...,T$, is generated by a strictly stationary and ergodic stochastic process. As before, we avoid the usual assumptions of homoskedasticity and normality, which are unlikely to hold in these data. We use a vector of ones and the contemporaneous values of the factors, F_{jt} as the instruments in the GMM. The orthogonality conditions therefore state that $E(u_{it}\ F_{jt}) = 0$ and $E(u_{it}) = 0$, for all $i=1,...,N$ and $j=1,...,K$.[3]

3. The Data

A. The Asset Returns

We study equity returns in eighteen national markets using monthly data provided by Morgan Stanley Capital International. The countries include sixteen OECD countries plus Singapore/Malaysia and Hong Kong. The country returns are value-weighted indices formed from a list of 1476 (as of December, 1989) companies. The firms represent about 65% of the market capitalization of the countries' stock markets, with some attempt to stratify the sample by industry groups, so that each industry is represented in proportion to its national weight [see Schmidt (1990)]. The stocks are generally those for which the total market value outstanding is large. Total monthly returns are measured for 1970-1989 as the capital change component of a country index plus the dividend yield, as provided by MSCI.[4] When measured in U.S. dollars the returns are in excess of the U.S. Treasury bill that is the closest to 30 days to maturity, provided by the Center for Research in Security Prices (CRSP) at the University of Chicago. When measured in local currency units the returns are in excess of a local short term interest rate from Citibase or the International Monetary Fund (IMF) [see the appendix for details]. To convert from local currency values to U.S. dollar values, the closing European interbank currency rates from MSCI are used on the last trading day of the month. The world equity market index is a value-weighted combination of the country returns.[5]

B. The Global Economic Risk Variables

Summary statistics for the variables are presented in Table 1. We include a brief discussion of each global risk variable here; details are provided in the appendix.

WDRET is the U.S. dollar return of the MSCI world equity market in excess of a short term interest rate. Asset pricing models usually include a role for a "market portfolio" as a measure of risk. Cumby and Glen (1990) test and do not reject the unconditional mean variance efficiency of the world market index. Harvey (1991a) does not reject the unconditional efficiency of the MSCI index in the set of MSCI country returns, but he does reject the conditional efficiency of the index. This raises the likelihood that previous tests are low in power. Fama and French (1992) find that unconditional betas on market indexes in the U.S. do not provide a good cross-sectional explanation of expected returns. It is therefore interesting to further examine the usefulness of a world beta to explain the country returns.

dG10FX is the log first difference in the trade-weighted U.S. dollar price of the currencies of 10 industrialized countries. The G-10 countries are defined as the G-7 (not including the U.S.) plus the Netherlands, Belgium, Sweden and Switzerland. (The G-7 countries are Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States.) This series is taken from the IMF as reported by Citibase. A positive change (dG10FX > 0) indicates a depreciation of the dollar. In Adler and Dumas (1983, equation 14), an exchange risk factor appears, which depends on exchange rates, consumer price index changes, and risk tolerance in each country. This theoretical measure is difficult to replicate empirically, so Dumas and Solnik (1992) break it down into separate factors for each country. We use a single aggregate measure as a parsimonious alternative to the approach of Dumas and Solnik (1992). Previous studies examine the pricing of exchange risks in national equity markets. They find little evidence that exchange risks are priced on average domestically, but we address a different question. We examine the pricing of a global measure of exchange risk in a multi-country asset pricing model.

G7UI is the unexpected component of a monthly global inflation measure. The G-7 inflation rate is a weighted average of the percentage changes in the consumer price indices (CPI) in the G-7 countries, using the relative shares of the total real, gross domestic product (GDP) as the weights. Inflation risk can be priced in a multi-beta model if inflation has real effects, in the general sense that global inflation is correlated with a representative investor's marginal utility. For example, higher inflation may signal higher levels of economic uncertainty which make consumers worse off. If national equity market returns differ in their exposure to changes in the global inflation outlook, there may be an inflation risk premium on global equity markets.

dG7ELT is the monthly change in a measure of long-term inflationary expectations. Chen, Roll and Ross (1986) include a measure of U.S. unexpected inflation and a measure of changes in expected U.S. inflation in their study for the U.S. We examine the pricing of global inflation on world markets. dG7ELT is formed by regressing a 48-month moving average of the G-7 inflation rate on a set of predetermined instruments and taking the first difference of the fitted values.

dTED is the change in the spread between the 90-day Eurodollar deposit rate and the 90 day U.S. Treasury bill yield. The "TED spread" is a measure of the premium on Eurodollar deposit rates in London, relative to the U.S. Treasury. Fluctuations in the spread may capture fluctuations in global credit risks.

G7RTB is a weighted average of short-term real interest rates in the G7 countries, using the shares of G-7 GDP as the weights, minus the G7 inflation rate. Real interest rates are often used in economic models to capture the state of investment opportunities. For example, Merton (1973) and Cox, Ingersoll and Ross

(1985) develop models in which interest rates are state variables. Chen, Roll and Ross (1986) and Ferson and Harvey (1991) include a real interest rate state variable in models for the U.S. market.[6]

dOIL is the change in the monthly average U.S. dollar price per barrel of crude oil.[7] A number of studies have examined premiums for oil price risk exposure using a cross-section of assets with a country. For example, Chen, Roll and Ross (1986) study oil prices as a measure of economic risk in the U.S. market and Hamao (1988) and Brown and Otsuki (1990b) study oil prices in the Japanese equity market. We study oil prices as a potential source of global market risk, to which different national markets may have differing exposures.

dG7IP is a weighted average of industrial production growth rates in the G7 countries, using a measure of relative production shares as the weights. Chen, Roll and Ross (1986) and Shanken and Weinstein (1990) examine the average pricing of U.S. industrial production in the U.S. market. Hamao (1988) examines domestic industrial production risk in the Japanese equity market and Bodurtha, Cho and Senbet (1989) study average risk premiums for domestic industrial production risks in several national equity markets. No previous study has examined the average pricing of global industrial output risks.

4. Empirical Results:

A. Factor Models

Table 2 presents factor model regressions (1) for each country using the eight global risk factors. The right hand columns of the table show the adjusted R-squares for three cases. In the far right column only a constant and the world market index are in the regressions. In the other two columns additional factors are added. The increment to the R-squares from including the other economic variables is small for most of the countries.[8] The world market portfolio is by far the most important factor from the perspective of explaining variance.

The panel at the bottom of Table 2 presents heteroskedasticity consistent Wald tests of the hypothesis that a risk factor has unconditional betas equal to zero for all of the countries. We do not reject this hypothesis for the variables dTED, G7UI, dG7IP and G7RTB; the hypothesis is rejected for the other four variables. If the beta coefficients for a particular variable are not significantly different from zero in the factor model regressions, this suggests dropping the variable from the analysis. Obviously, if the beta is zero the factor is not useful for controlling variance. Similarly, their can be no expected return premium for a factor whose betas are all zero. However, a small nonzero beta could be associated with an important average risk premium. For example a risk factor could have low correlation with asset returns but high correlation with the aggregate marginal utility of wealth.

If a risk variable does not have different coefficients across the countries then it cannot be priced even if the betas differ from zero. This is because unconditional pricing implies that expected returns differ across countries depending on differences in their sensitivities to the variable. The point estimates suggest that there are significant differences in the countries' sensitivities to several of the variables. For example, the Japanese stock market is significantly positively-related to changes in long term inflationary expectations, while the Australian market is negatively-related to the same factor. The equity markets of Hong Kong and Singapore/Malaysia are positively-related to oil price changes, while in Spain the relation is negative.

The lower panel of Table 2 presents tests of the hypothesis that a given factor has betas which are equal to a common value in all of the countries. We do not reject this hypothesis for the same set of variables for which we do not reject that the betas are jointly zero (and therefore equal). On the basis of these tests, we exclude the variables dTED, G7UI, dG7IP and G7RTB from our empirical model and we retain the other variables (WRDRET, dG10FX, dG7ELT, and dOIL).[9]

B. Expected Returns and Global Economic Risks

Table 3 presents the results of the SURM, equation (4), for two cases. The three columns on the left present a model where the world market index is the only factor. The five columns on the right present a four-factor model. The first column reports the excess return betas on the world market index, which vary from 0.38 (Austria) to 1.3 (Hong Kong). The beta in the SURM's are very close to the betas in the unrestricted regression models. The two world market beta estimates never differ by more than one standard error for a given country.

In the single-factor model with the world market factor, the factor is the excess return of a portfolio. In such a case, Shanken (1992) shows that the best estimate of the risk premium is the expected excess return itself. For the one-factor model we therefore impose the restriction that the expected risk premium λ_1 is equal to the expected excess return of the world market index. The expression $(f_{1t} + \lambda_1)$ in (4) is replaced by r_{mt}, and the number of free parameters in the SURM is reduced by one.

The goodness of fit statistic for the restricted SURM implies a right-tailed p-value of 0.295, thereby not rejecting the hypothesis that the world market portfolio is unconditionally mean-variance efficient. This is consistent with the tests in Cumby and Glen (1990) and Harvey (1991a). However, further investigation suggests that such joint tests are low in power.

The point estimate of the world market risk premium is 0.545 and its standard error is 0.271, which seems reasonable. But the pricing errors of the one-factor model for many of the countries are large. The average pricing errors are defined as the difference between the average country returns and the expected returns predicted by the model, evaluated at the sample estimates. The model leaves economically large pricing errors of 0.8% (standard error = 0.3%) per month for Japan and 0.9% (standard error = 0.7%) per month for Hong Kong. The standard errors for the average pricing errors are calculated as in Hansen (1982, lemma 4.1), which accounts for the fact that the pricing errors are evaluated at the GMM point estimates of the parameters.[10] Of course, it is hazardous to focus on the pricing errors which are the largest in a group of estimates, and to apply these standard errors to judge their significance. This is because the multiple comparisons implied by selecting the largest values are not accounted for in the individual standard errors. Still, the pricing errors suggest that the Japanese and Hong Kong stock markets out performed the world market on a beta adjusted basis over this period, while Canada, Italy and the United States have been poor performers on the same basis.

The panel below the individual country results in table 3 reports the mean absolute pricing error and a value-weighted average absolute pricing error as alternative summary measures. The value weights are the weights of the countries in the MSCI world index as of the first quarter of 1989. The value-weighted average absolute pricing error is larger (0.44% per month) than the equally weighted average (0.33% per month) primarily because Japan receives a large weight, as we believe it should.

These economically large deviations from the model, as represented by the pricing errors, are not sufficient to reject the efficiency of the world market index using the standard goodness-of-fit statistic. The joint test statistic combines the squared pricing errors together with the other orthogonality conditions, weighted by the precision with which they are estimated. The pricing errors may be large, but their precision is low, and the joint test statistic does not have enough power to reject the model.

Some additional exercises provide further evidence on the ability of the world market betas to explain the average returns. We use the cross-sectional methods of Fama and MacBeth (1973), estimating monthly regressions of the country returns on their unconditional world market betas.[11] The results differ depending on whether the cross-sectional intercept is suppressed, as it is in the SURM, or included in the regression. With the intercept suppressed the slope coefficient, an estimate of the world market premium, is 0.819% per month (standard error= 0.085%). However, when the intercept is included in the regression the slope coefficient is 0.469% (standard error = 0.349%). The average of the adjusted R-squares in the monthly cross-sectional regressions is only 4%. Although efficiency of the world market index is not rejected, the relation between the country returns and the world market betas is weak.[12] This is also evident in figure 1, which plots the average returns against the world market betas.

The right-hand columns of Table 3 present the pricing results using the four global risk variables. The goodness-of-fit test does not reject the model, producing a large right-tail p-value of 0.892. The average pricing errors for many of the countries are reduced in the multiple-beta model, relative to the single beta model. For example, the pricing errors are only 0.2% per month for Japan and 0.1% for Hong Kong. For a number of the other countries the pricing errors are an order of magnitude smaller than they were using only the world equity portfolio to measure risk. Furthermore, the mean absolute error and the value-weighted mean absolute errors are substantially reduced. This suggests that when the measures of risk are expanded to include the other variables, then much of the seemingly abnormal average performance of the Japanese and Hong Kong markets may be explained as compensation for global economic risk. However, Canada, Italy and the U.S. markets still seem to have performed poorly on a risk adjusted basis.

The average risk premium in Table 3 for the exchange risk variable dG10FX is 0.602% and is 1.8 standard errors from zero using the SURM. The average premiums for dG7ELT and dOIL are not individually statistically significant. Recall that dG10FX is measured in dollars per local currency unit. All of the betas on this variable are positive (except for the US and Canada), indicating that when the dollar depreciates the dollar excess return of foreign stocks tends to rise.[13]

The estimated risk premium for the world market index appears strongly significant in the multiple beta SURM. Introducing the additional risk factors results in a slightly larger point estimate of the premium for the world market index. The standard error of that estimate is much smaller than in the single-factor model. The results for the multiple-factor model in Table 3 appear to be different from the results of Chen, Roll and Ross (1986) for the U.S., where the introduction of multiple risk factors reduced the average premium for the U.S. stock market index to a small, statistically insignificant number. We find that part of the explanation is a difference in the methodologies.

Chen Roll and Ross (1986) used a cross-sectional regression approach which allowed for a nonzero intercept. In the SURM, the intercept is suppressed, as it should be zero under the null hypothesis. When we use cross-sectional regressions, similar to Fama and MacBeth (1973), which include an intercept and thereby estimate the multiple-beta model allowing an ad-hoc alternative hypothesis, the premium for the world market index is not significantly different from zero.[14]

C. Sensitivity Analysis

Our sample covers the decades of the 1970's and 1980's, a period in which the international investment climate saw much change. Barriers to international investment which had been in place were removed or weakened in the latter parts of the 1970's and early 1980's in several countries. Some of the national market average returns are remarkably different in the two subperiods. For example, the average excess dollar return for Japan is 1.02% per month in the first half and 1.61% in the second half. For Italy, the average is -0.67% in the first half and 1.27% in the second. It is interesting, therefore, to see if the last half of our sample produces qualitatively different results than does the first half. We estimate the SURM models for the first and second halves of our sample. Similar to the full period results, the test statistics do not reject the models. The average pricing errors are similar in the two subperiods. The value-weighted pricing error is 0.209% in the first half and 0.210% in the second half. The average pricing error for Japan is 0.346% in the first half and 0.229% in the second. The world market premium is significant in both subperiods; in fact the point estimate using the SURM is slightly larger in both subperiods than over the full sample. The premium for dOIL is 0.31% per month in the second subperiod and is significant, but the estimate is negative and insignificant in the first subperiod. A similar result is found for the foreign exchange risk variable. The premium estimate is positive in the first period and negative in the second. These results may indicate time-variation in the expected premiums.

To assess the sensitivity of our results to the use of US dollar excess returns, we change the definition of the returns to local currency units, measured in excess of a local short term interest rate. Such an excess return can be interpreted as a long position in the local stock market financed by local currency borrowing, and is therefore hedged against currency fluctuations to some extent. We find that the results on the average pricing of the world market index are virtually unaffected.[15] However, the SURM estimates of the average premium associated with dG10FX are closer to zero and are not statistically significant. We examine univariate models with dG10FX and bivariate models with dG10FX and the world market index, using Fama-MacBeth (1973) methods. The average cross-sectional relation between returns and the betas on dG10FX are depicted in figure 2 (U.S. dollar returns) and figure 3 (local currency returns). These experiments generally confirm the results of the SURMs.[16]

Measurement errors in the economic data may reduce the correlation of the global risk measures with the country returns. We therefore conduct an additional set of tests using maximum correlation portfolios for the economic risk factors, similar to Breeden, Gibbons and Litzenberger (1989) and McCurdy and Morgan (1992). The portfolio weights are the slope coefficients of the economic variable regressed on the asset returns and are fixed over time. If there is measurement error which is unrelated to returns, then the measurement error is captured in the residual when the maximum correlation portfolios are formed. We do not form the portfolios from the same set of national equity market returns that we are trying to "explain," which should reduce the impact of the overfitting problem with mimicking portfolios [Wheatley (1989)]. We use the MSCI international industry indices for this purpose. These are a set of 38 equity indices, formed by industry groups and using the common stocks of firms from many countries in each industry group. (These data are described in more detail in the appendix.)

We examine factor model regressions for the country returns using the maximum correlation portfolios as the factors. Compared with table 2, the explanatory power of the regressions are slightly higher for most, but not all of the countries. The smallest R-square is 11.8% and the largest is 85.9%. The higher R-squares are consistent with the existence of measurement error in the economic data which is unrelated to stock returns.

Given mimicking portfolios for the risk factors, the model implies that the λ_j are their expected excess returns. However, the MSCI industry indices do not include dividends, so the average returns of portfolios formed from these indices are not good estimates of the risk premiums. This is a problem similar to what Stambaugh (1983a) calls "mean deficiency." We handle the mean deficiency by treating the proxy portfolios the same way as we do the economic risk variables in the SURM; namely, we use their de-meaned values and we estimate the risk premiums as separate parameters. As the dividend component of the return is relatively smooth, its absence from the industry indices should not much affect estimates of the covariances. Comparisons with the previous tables provides further evidence on the robustness of our results.

For pricing purposes the proxy portfolios should be maximally correlated with the state variables in the universe of test assets [Breeden (1979)]. However, since we do not include the country returns in the proxy portfolios, it is possible that higher correlation with the state variables could be obtained. This can be interpreted in terms of the familiar mean-variance diagram, as in figure 4. If a set of factor portfolios determine the expected returns in a multi-beta model, a combination of them lies on the minimum variance boundary of all asset returns [Chamberlain (1983), Grinblatt and Titman (1987)]. Introducing additional assets will in general expand the minimum variance boundary. If the efficient combination of our proxy portfolios lies inside the minimum variance boundary of the test assets, then a combination of the portfolios will not price the test assets. Figure 4 shows that the unconditional minimum variance boundary formed from the industry indices contains the boundary formed from the country returns.[17]

The results of the SURM using the maximum correlation portfolios are qualitatively similar to those of Table 3. We do not reject the efficiency of a maximum correlation portfolio for the world index using the standard goodness-of-fit test, but the average pricing errors for some countries are large. The mean absolute and weighted mean pricing errors are somewhat smaller than in table 3, and they are reduced dramatically when we examine the multiple beta model. Introducing the maximum correlation portfolios for the other factors does not diminish the significance of the world market portfolio risk premium in the SURM. As in Table 3, the point estimate of its premium is higher in the multiple beta model.[18]

One interesting difference between the results using the global economic risk variables and using the maximum correlation portfolios involves the industrial output variable dG7IP. Its proxy portfolio produces the largest of the average risk premiums in the SURM, which has a t-ratio of 1.7. This suggests that the output variable may contain measurement errors that are important and are cleaned up to some extent by a maximum correlation portfolio. Shah (1989), Fama (1990), Schwert (1990) and Kothari and Shanken (1992) find that stock returns in the US are sensitive to changes in expected future output. Harris and Opler (1990) and Beckers (1991) extend such results to international data. We therefore conduct experiments in which we replace dG7IP by a one-year leading growth rate. Univariate and multivariate factor model regressions using this variable are jointly significant and the betas on this variable are significantly different across the countries.

We investigate the unconditional pricing results using the 12-month leading output growth rate as an additional risk factor. In the multivariate SURM the betas on the leading production variable do not seem to be marginally important. The premium estimate is 1.25% per month but the standard error is 1.16%. The world market premium estimate is not changed much by the introduction of the leading production variable. Using the maximum correlation portfolios, including one for the leading industrial production variable, we find generally similar results. In this case, however, the premium on the leading output factor is significant, at 2.97% per month (standard error=1.20%).

5. Concluding Remarks

We empirically examine multiple beta models for the returns and expected returns on eighteen national equity markets using a set of factors chosen to measure global economic risks. Although previous studies do not reject the unconditional mean-variance efficiency of a world equity market portfolio, we find that the world market betas provide a poor explanation of the average returns across countries. Our tests do not reject the hypothesis that the returns are consistent with a four-factor model. The average pricing errors of the multiple-beta model are only 0.2% per month for Japan and 0.1% for Hong Kong, which are much smaller than the errors of a model based on only the world market portfolio. This suggests that when the measures of risk are expanded to include such variables as exchange rates, oil prices and long-term inflationary expectations, then much of the seemingly abnormal average performance of the Japanese and Hong Kong markets may be explained as compensation for global economic risk.

Appendix

A. Equity returns

We use the Morgan Stanley Capital International (MSCI) equity indices. MSCI tracks 21 national indices of which we use 18 (Finland, Mexico and New Zealand are excluded because their data histories are shorter). Stocks from non-domiciled companies and investment funds are excluded from individual country indices. Companies with restricted float due to dominant shareholders or cross-ownership are avoided. However, there are 51 stocks from Finland, Norway, Sweden and Switzerland that are restricted. The overall weight of these restricted stocks is small, and Schmidt (1990) finds that there is little difference in the national indices when these restricted stocks are excluded. All indices have a common base of 100 in December 1969. The indices are constructed using the Laspeyres method which approximates value weighting. U.S. dollar returns are calculated by using the exchange rate available at 4:00 p.m. Central European Time. We use the MSCI world industry portfolios to construct maximum correlation portfolios for the economic variables. MSCI tracks 38 industry groups. These are: Aerospace and Military Technology, Appliances and Household Durables, Automobiles, Banking, Beverages and Tobacco, Broadcasting and Publishing, Building Materials and Components, Business and Public Services, Chemicals, Construction and Housing, Data Processing and Reproduction, Electrical and Electronics, Electronic Components and Instruments, Energy Equipment and Services, Energy Sources, Financial Services, Food and Household Products, Forest Products and Paper, Gold Mines, Health and Personal Care, Industrial Components, Insurance, Leisure and Tourism, Machinery and Engineering, Merchandising, Metals (Non-Ferrous), Metals (Steel), Miscellaneous Materials and Commodities, Multi-Industry, Recreation, Other Consumer Goods, Real Estate, Telecommunication, Textiles and Apparel, Transportation-Airlines, Transportation-Road and Rail, Transportation-Shipping, Utilities-Electrical and Gas, and Wholesale and International Trade. All of the world industry indices have a base value of 100 in December 1969. The indices are calculated in U.S. dollars but do not include dividends. Both the industry and national indices are calculated in excess of the 30-day U.S. Treasury bill which is available from Ibbotson Associates. The correlations of the maximum correlation portfolios and the MSCI world index (wrdret) are shown below:

CORRELATIONS OF THE MAXIMUM CORRELATION PORTFOLIOS

wrdret	1.000	0.997	0.318	0.557	0.000	0.478	0.165	-0.103	0.251
mwrdret		1.000	0.319	0.559	0.000	0.479	0.166	-0.103	0.252
mdted			1.000	0.095	-0.143	0.272	-0.120	-0.118	0.146
mdG10fx				1.000	-0.044	0.147	0.118	-0.080	0.189
mG7ui					1.000	0.065	-0.018	0.210	-0.560
mdG7elt						1.000	-0.282	0.004	-0.155
mdoil							1.000	-0.053	0.304
mdG7ip								1.000	-0.173
mG7rtb									1.000

B. The World Risk Factors

WDRET = The world return is the arithmetic return on the Morgan Stanley Capital International world equity index (including dividends) less the Ibbotson Associates one month bill rate.

dTED = The change in the Eurodollar-Treasury yield spread is the difference between the 90-day Eurodollar yield (Citibase FYUR3M) and the 90 day Treasury bill yield (Citibase FYGM3 secondary market, converted from discount to true yield to maturity).

dG10FX = The change in the G-10 foreign exchange rate is the difference in the trade weighted dollar price of foreign exchange for 10 industrialized countries (G-7 plus the Netherlands, Belgium, Sweden, and Switzerland) (Citibase FXG10).

G7UI = The unexpected inflation for the G-7 countries is derived from a time-series model applied to an aggregate G-7 inflation rate. The G-7 inflation rate is constructed by weighing the individual countries' inflation rates (Citibase: PC6CA, PC6FR, PC6IT, PC6JA, PC6UK, PC6WG and ZUNEW) by their shares in the previous quarter's real U.S. dollar G-7 gross domestic product. These weights change through time. The time series model is ARIMA(0,1,2)(0,1,2) and the parameter estimates are:

	Parameter	Std. Error	T-ratio
Intercept	0.00000	0.000057	0.10
MA1,1	0.432613	0.061754	7.01
MA1,2	0.271394	0.061544	4.41
MA2,1	-0.305806	0.065162	-4.69
MA2,2	-0.180382	0.065377	-2.76

The parameters are estimated with 250 monthly observations. The chi-square test for significance of the first six residual autocorrelations has a p-value of .111 and the corresponding statistic for the first 12 autocorrelations has a p-value of .275.

dG7ELT = Change in long term expected G-7 inflation is a result of projecting the four year moving average of G-7 inflation on a set of predetermined instrumental variables. The predetermined instruments are (1) the level of one-month short-term U.S. Treasury bill yield, (2) the dividend yield of the MSCI value-weighted world stock market index, (3) a spread between the yields to maturity of ten-year U.S. Treasury bonds and 90-day U.S. Treasury bills, (4) the lagged value of the Eurodollar (TED) - U.S. Treasury spread, (5) the lagged return on the MSCI world market index, and (6) a dummy variable for the month of January. The regression models the expected long term inflation and dG7ELT is the first difference of the fitted values of the regression.

dOIL = The change in the natural log of the average U.S. dollar price of per barrel at the wellhead from 1974-1989 and the posted West Texas Intermediate price from 1969-1973. Since the West Texas price is consistently higher than the average wellhead price, the 1969-1973 data is grossed down by 65%. This represents the average premium of West Texas over the average during 1974-1976.

dG7IP = The change in G-7 industrial production is calculated by weighing local industrial production

indices by the following (fixed) factors: Canada .04314, France .09833, Germany .05794, Italy .13093, Japan .07485, U.K. .11137, U.S. .48343 which are the weights in G-7 gross domestic product in the third quarter of 1969. The logarithmic difference in this aggregate index is the growth in G-7 industrial production.
G7RTB = The G-7 real interest rate is calculated by aggregating individual countries' short term interest rates. The following interest rates are used (Citibase FYCA3M-Canada 90 day Treasury bill, FYFR3M-France 90 day bill, FYGE3M-Germany 90 day bill, FYIT6M-Italy 180 day bill, FYCMJP-Japan commercial paper 1969-1976 and FYJP3M-Japan Gensaki rate 1977-1989, FYUK3M-United Kingdom 90 day bill, FYUS3M-United States 90 day bill.) The aggregate G-7 interest rate is calculated by using the countries's previous quarter's shares in G-7 gross domestic product. The real G-7 interest rate is calculated by subtracting the G-7 inflation rate.

C. Short term interest rates

These are used to calculate excess returns in local currency units. The data are as follows: Australia-13 week bill (IFS 61C), Austria-Money market rate (IFS 60B), Belgium-3 month bill (Citibase FYBE3M), Canada-3 month bill (IFS 60C), Denmark-Discount rate 1969-1971 (IFS 60A), Call money rate 1972-1989 (IFS 60B), France-3 month interbank (Citibase FYFR3M), Germany-Frankfurt 90-day rate (Citibase FYWG3M), Hong Kong-No data, U.S. 3-month bill used, Italy-6 month bill (Citibase FYIT6M), Japan-Call money rate 1969-1976 (Citibase FYCMJP), Gensaki rate, 1977-1989 (Citibase FYJP3M), Netherlands-Call money rate 1969-1978:11 (IFS 60B), 3 month bill 1979:12-1989, Norway-Prime rate 1969-1971:1, Call money rate 1971:12-1989 (IFS60B), Singapore/Malaysia-no data, U.S. bill, Spain-Prime rate 1969-1973:12, Call money rate 1974-1976 (IFS 60B), 3 month bill 1977-1989 (IFS 60C), Sweden-3 month bill (IFS 60C), Switzerland-3 month deposit rate (Citibase FYSW3M), United Kingdom-3 month bill (Citibase FYUK3M), United States-3 month bill (Citibase FYUS3M)

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Footnotes:

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1. See Ferson and Harvey (1993) for an examination of *conditional* asset pricing models using similar variables.

2. Mimicking portfolios are defined as portfolios that may be substituted for the factors in a factor model regression and whose expected excess returns are the risk premiums [e.g. Lehmann and Modest (1988), Huberman, Kandel and Stambaugh (1987)].

3. We use iterated GMM following Ferson and Foerster (1993), who found that such an approach has superior finite sample properties when compared with a two step procedure.

4. The dividend yield is 1/12 of the previous year's dividend divided by the level of the index at the end of a month. See the appendix for details.

5. MSCI attempts to avoid the double counting of firms whose equity is traded on the stock markets of more than one country. There are, however, other problems with the index. French and Poterba (1991) show that the MSCI world index gives too much weight to Japan because the amount of cross corporate ownership of shares in Japan has been unusually high. Alternative indices, such as the FT-Actuaries world index, suffer from the same problem. Harvey (1991a) reports that in March of 1989 Japan accounted for 43% of the MSCI world index and 41% of the FT-Actuaries index. We choose the MSCI data over the FT-Actuaries data, because the latter are only available from 1981.

6. Although the correlation between G7RTB and G7UI is relatively high (at -0.56), it is not perfect because the G7 nominal interest rates are not part of the conditioning information used to form G7UI and because G7RTB is not prewhitened.

7. We use a spliced series of the posted west Texas intermediate crude and the average U.S. wellhead price, as described in the appendix. These are not the best indicators of market prices, but they are the best available to us for this period. Futures markets for crude oil did not develop until 1983 (heating oil futures began trading in 1978). Chen, Roll and Ross (1986) used the energy component of the producer price index. Given the prevalence of long

term oil price contracts over much of the sample, this measure is not likely to better reflect current oil market conditions.

8. This is consistent with the evidence of Wasserfallen (1989) who finds little sensitivity of international stock returns to macroeconomic news. However, Wasserfallen uses the residuals from a vector autoregression as his factors, while in table 2 we define innovations relative to the unconditional means.

9. We replicated the tests in table 2 using only G7UI or only dG7ELT as the inflation variable. The test results for the two inflation variables, as well as the other six risk factors, were not sensitive to which of the inflation variables was included, or if both were included. Thus, dG7ELT does not appear to proxy for unexpected inflation in these data. We also found that the coefficients on the world market index for most of the countries, as presented in table 2, are similar to their simple regression betas on only the world market index.

10. The GMM parameter estimates are found by minimizing a quadratic form, g'Wg, where W is the fixed weighting matrix, the inverse of a consistent estimate of the covariance matrix of the orthogonality conditions, g, at the true parameter values. Our pricing error is the sample mean of one of the orthogonality conditions, evaluated at the point estimates. Let gd be the sample mean of the gradient of the orthogonality conditions with respect to the parameter vector. The formula for the covariance matrix of the sample

mean of the orthogonality conditions is: $[W^{-1} - gd(gd'W\ gd)^{-1}\ gd']/T$. The covariance matrix is evaluated at the consistent sample estimates.

11. The cross-sectional regression procedure of Fama and MacBeth (1973) assumes that the security returns are correlated cross-sectionally and are heteroskedastic, which implies that the usual regression standard errors can be misleading. The time-series average of the monthly cross-sectional regression slopes is used as the estimate of the expected premium. The standard errors are calculated as the standard error of the mean using the time series of the monthly estimates. This assumes that the series of the monthly coefficient estimates (which are themselves portfolio returns) are uncorrelated over time. Shanken (1992) reviews the statistical assumptions required for the Fama MacBeth approach and shows how it is related to Maximum likelihood and other approaches.

12. A single cross-sectional regression of the average excess returns on the world market betas, using the table 3 estimates, produces an adjusted R-square of only 3.2%.

13. This is consistent with previous studies [e.g., Jorion (1991)] who measure exchange rates as dollars per currency unit and find positive (but insignificant) average premiums.

14. In a single cross-sectional regression (with an intercept) of the average returns on the four betas from table 3, the adjusted R-square is 71.2% and three of the four premiums appear significant. The world market premium has the smallest t-statistic, equal to 1.74.

15. Korajczyk and Viallet (1989) report similar results in their study of four countries using a world CAPM and APT factors.

16. The main exception is that when returns are measured in local currency units the average premium on dG10FX is marginally significant in the univariate and bivariate Fama-MacBeth models, while in the SURM and in the four-factor models the premium is not significant.

17. If the industry returns were shifted up by an amount, approximately reflecting the missing dividend yield, it is clear from figure 5 that the country index boundary would still be contained within the adjusted boundary. Of course, the figure does not account for any estimation error in the boundary, which may be large.

18. The joint tests for zero betas and for betas that are equal across the national markets, like in table 2, are conducted using the proxy portfolios in place of the economic variables. Only the portfolios for the variables dOIL and dG7RTB fail to produce significant regression betas. We therefore use

six factor portfolios in most of these experiments. We have also replicated the SURM's using four maximum correlation portfolios for the variables that were examined in table 3. We find similar results for the average pricing errors and the test statistics. The estimates of the world market premium and the other premiums are also similar to those of the six maximum correlation portfolio model. The main difference is that the premium for dG10FX is 1.38% per month in the six factor model and only 0.244% in the four factor model. Neither of these is more than two standard errors from zero.

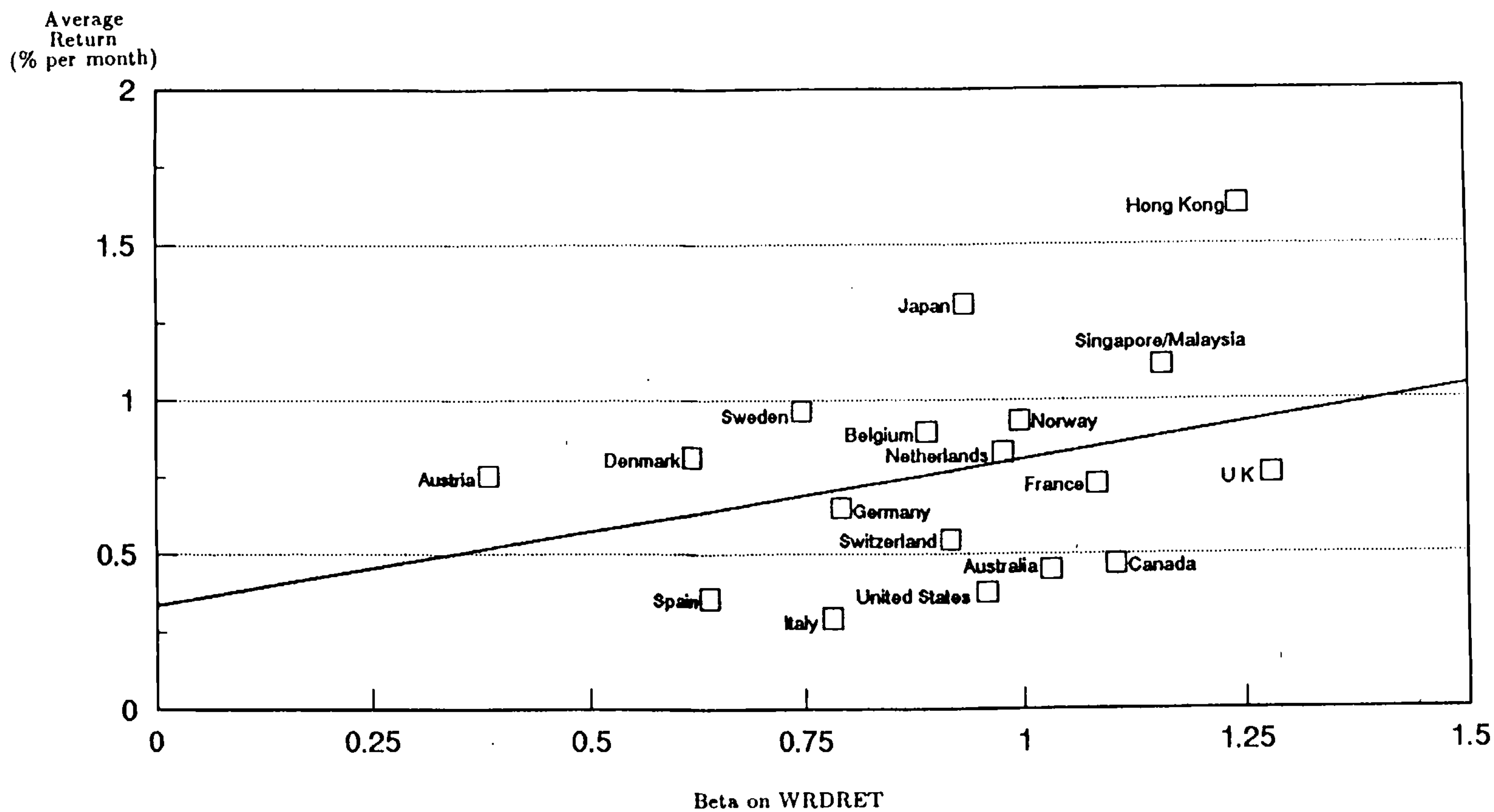

Fig. 1. Unconditional beta coefficients of country index returns calculated in U.S. dollars regressed on the world market return (WRDRET). The sample is February 1970–December 1989.

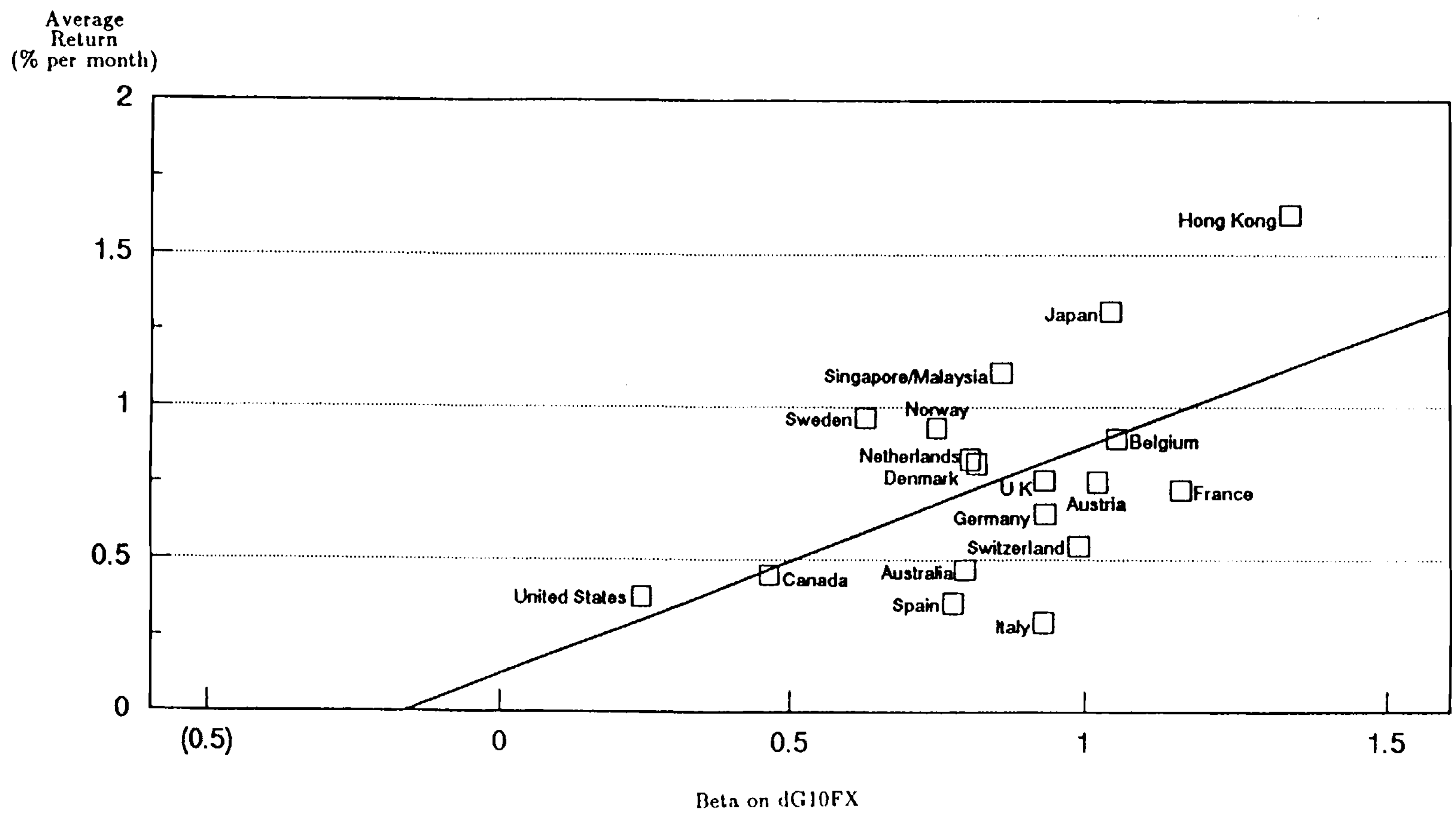

Fig. 2. Unconditional beta coefficients of country index returns calculated in U.S. dollars regressed on the log change in the trade-weighted U.S. dollar price of the currencies of 10 industrialized countries (dG10FX). The sample is February 1970–December 1989.

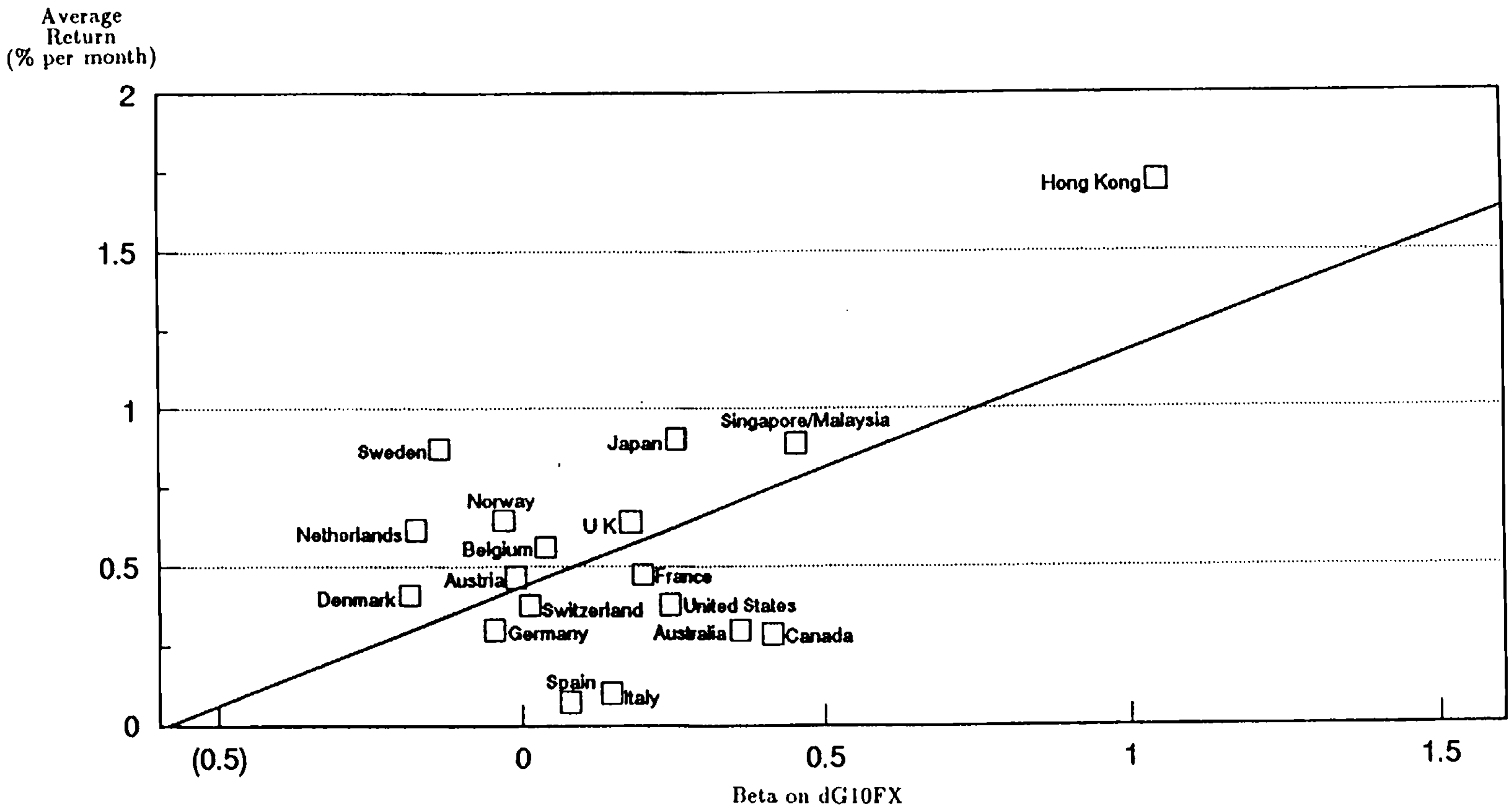

Fig. 3. Unconditional beta coefficients of country index returns calculated in local currency regressed on the log change in the trade-weighted U.S. dollar price of the currencies of 10 industrialized countries (dG10FX). The sample is February 1970–December 1989.

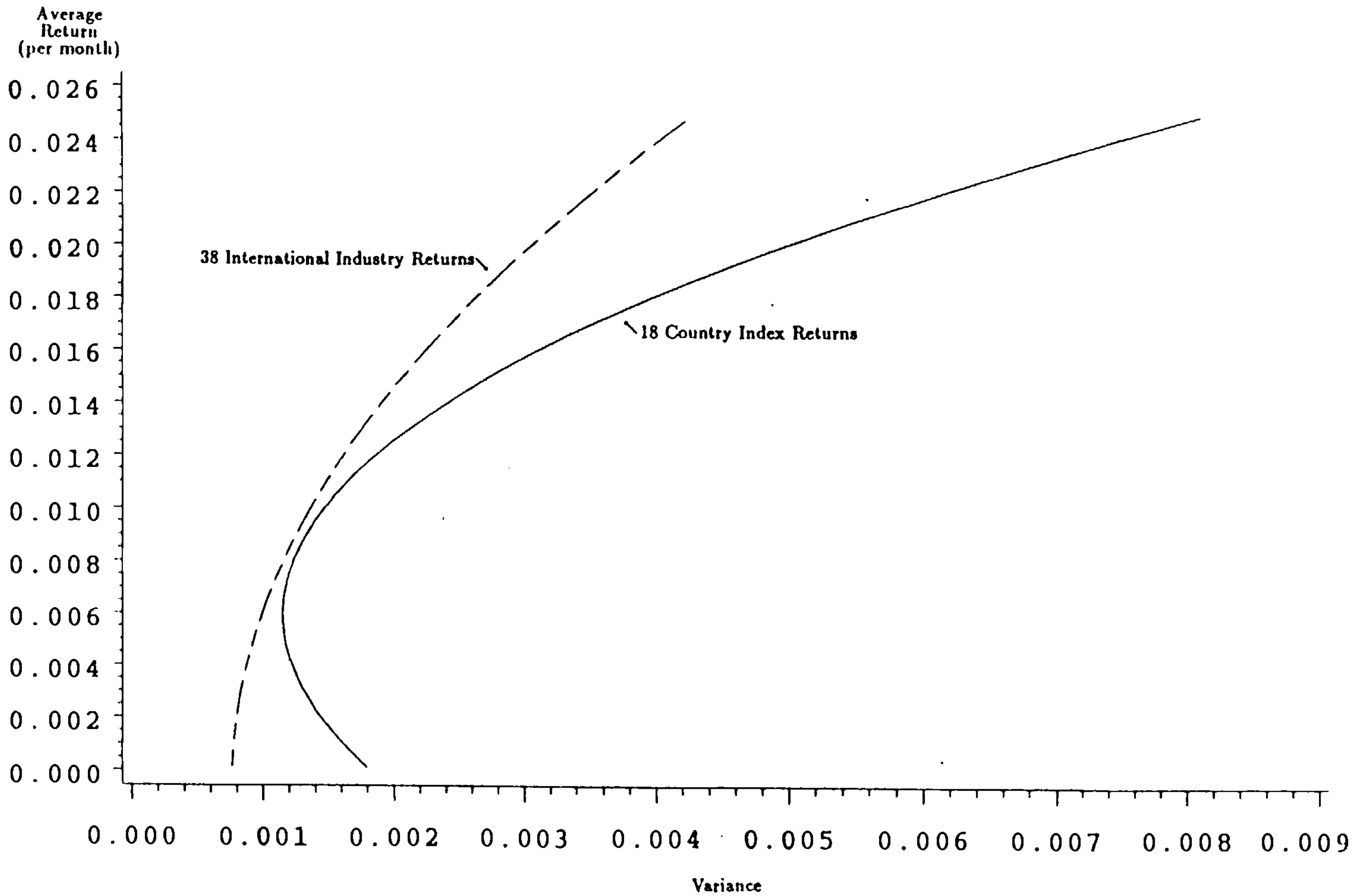

Fig. 4. Two unconditional minimum variance frontiers. The solid curve is calculated using the returns on 18 country portfolios. The dash curve represents the unconditional minimum variance frontier calculated using 38 international industry returns. The industry returns do not include dividends. The sample is February 1970–December 1989.

Table 1

Summary statistics for the global risk factors: 1970:2–1989:12 (239 observations)

Variable	Symbol	Mean	Std. dev.	ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	ρ_{24}
World risk factors									
World excess return	wdret	0.545	4.189	0.15	-0.03	0.05	0.03	0.09	0.00
Change in Eurodollar-Treasury yield	dted	-0.046	3.988	-0.08	-0.09	-0.07	-0.13	0.02	0.09
Log change in G-10 foreign exchange rate	dG10fx	0.104	2.099	0.31	0.06	0.09	0.08	0.06	-0.07
Unexpected G-7 inflation	dG7ui	-0.005	0.204	0.00	0.04	-0.02	-0.09	0.01	0.08
Change in long-term G-7 expected inflation	dg7elt	-0.039	1.275	-0.34	-0.12	0.09	-0.16	-0.08	0.01
Change in price of oil	doil	0.062	0.861	0.56	0.22	0.08	0.03	-0.02	-0.02
Change in G-7 industrial production	dg7ip	0.215	0.817	0.11	0.27	0.24	0.15	-0.08	-0.17
G-7 real interest rate	g7rtb	0.132	0.316	0.67	0.52	0.53	0.52	0.67	0.55

Correlations of the world risk factors

Variable	wdret	dted	dG10fx	G7ui	dG7elt	doil	dG7ip	G7rtb
wdret	1.000							
dted	-0.154	1.000						
dG10fx	0.314	0.010	1.000					
G7ui	-0.005	0.166	-0.015	1.000				
dG7elt	-0.253	0.238	-0.110	0.005	1.000			
doil	-0.066	-0.107	-0.143	0.207	-0.098	1.000		
dG7ip	-0.053	0.128	-0.075	0.128	0.053	-0.055	1.000	
G7rtb	0.101	-0.058	-0.059	-0.564	-0.007	-0.289	-0.031	1.000

The world excess return is the arithmetic return on the Morgan Stanley Capital International world equity index (including dividends) less the Ibbotson Associates one-month bill rate. The change in the Eurodollar-Treasury yield spread is the difference between the 90-day Eurodollar yield and the 90 day Treasury bill yield (from Federal Reserve Bulletin). The log change in the G-10 foreign exchange rate is the difference in the trade weighted dollar per foreign exchange rate of 10 industrialized countries (G-7 plus the Netherlands, Belgium, Sweden, and Switzerland) from International Monetary Fund. The unexpected inflation for the G-7 countries is derived from a time-series model applied to an aggregate G-7 inflation rate where the (varying) weights in the aggregation are determined by country weights in total G-7 gross domestic product. Change in long term expected G-7 inflation is a result of projecting the four year moving average of G-7 inflation on the set of instrumental variables specified below. The change in the price of oil is the change in the average U.S. dollar price of per barrel at the wellhead from 1974-1989 and the posted West Texas Intermediate price from 1969-1973. This variable is divided by 100. The change in G-7 industrial production is calculated by weighting local industrial production index levels by the following weights: Canada .04314, France .09833, Germany .05794, Italy .13093, Japan .07485, U.K. .11137 U.S. .48343 which are the weights in G-7 gross domestic product in the third quarter of 1969. The growth rate is the logarithmic difference in the aggregate industrial production index. The G-7 real interest rate is calculated by aggregating individual countries' short term interest rates minus inflation rates using (varying) weights based on quarterly shares in G-7 gross domestic product.

Table 2

Regressions of the asset returns on the world risk factors: 1970:2–1989:12 (239 observations)

Source of risk / Country	intercept	wdret	dted	dG10fx	G7ui	dG7elt	doil	dG7ip	G7rtb	$\overline{R}^2$ 8 risk factors	$\overline{R}^2$ 4 risk factors	$\overline{R}^2$ only wdret
Australia	0.001	0.999*	-0.187	0.117	3.262	-0.944*	0.269	-0.203	-1.334	0.351	0.344	0.319
	(0.004)	(0.146)	(0.102)	(0.219)	(2.349)	(0.392)	(0.648)	(0.505)	(1.641)			
Austria	0.003	0.217*	0.045	0.881*	2.380	-0.281	-0.133	0.128	1.984	0.155	0.159	0.075
	(0.003)	(0.093)	(0.081)	(0.232)	(2.107)	(0.509)	(0.556)	(0.409)	(1.379)			
Belgium	0.003	0.814*	-0.050	0.549*	1.663	0.451	-0.517	-0.239	1.156	0.422	0.427	0.386
	(0.003)	(0.089)	(0.087)	(0.186)	(1.581)	(0.298)	(0.575)	(0.378)	(1.061)			
Canada	-0.000	1.053*	0.003	-0.200	-0.339	-0.294	0.445	-0.066	-0.733	0.552	0.559	0.549
	(0.003)	(0.068)	(0.064)	(0.130)	(1.481)	(0.227)	(0.312)	(0.319)	(0.992)			
Denmark	0.003	0.557*	0.042	0.509*	-0.372	0.177	0.325	0.098	0.830	0.220	0.230	0.209
	(0.003)	(0.074)	(0.065)	(0.163)	(1.982)	(0.193)	(0.411)	(0.483)	(1.242)			
France	0.002	1.007*	0.018	0.511*	2.058	0.172	-0.301	-0.519	0.366	0.390	0.395	0.380
	(0.004)	(0.092)	(0.105)	(0.198)	(2.192)	(0.352)	(0.449)	(0.492)	(1.375)			
Germany	0.003	0.698*	0.022	0.461*	-0.243	-0.155	-0.215	-0.548	0.651	0.306	0.311	0.292
	(0.004)	(0.097)	(0.095)	(0.204)	(2.142)	(0.448)	(0.478)	(0.463)	(1.508)			
Hong Kong	0.006	1.255*	0.420	0.669	-2.437	0.130	1.356*	1.741*	-1.049	0.187	0.170	0.162
	(0.008)	(0.215)	(0.272)	(0.554)	(4.379)	(0.413)	(0.777)	(0.783)	(2.808)			
Italy	-0.004	0.670*	-0.289*	0.512*	-1.019	0.367	-1.020	0.705	1.437	0.217	0.199	0.175
	(0.005)	(0.106)	(0.127)	(0.239)	(2.800)	(0.377)	(0.794)	(0.518)	(1.657)			
Japan	0.006	0.893*	-0.052	0.535*	1.230	0.776*	-0.350	0.292	1.092	0.452	0.457	0.408
	(0.004)	(0.084)	(0.089)	(0.147)	(1.730)	(0.267)	(0.357)	(0.388)	(1.217)			
Netherlands	0.003	0.936*	0.069	0.208	0.148	-0.176	0.087	-0.322	0.466	0.542	0.545	0.544
	(0.003)	(0.072)	(0.075)	(0.121)	(1.461)	(0.225)	(0.292)	(0.309)	(1.056)			
Norway	0.002	0.973*	-0.026	0.224	6.107*	-0.074	0.992	0.657	0.046	0.278	0.262	0.253
	(0.006)	(0.133)	(0.125)	(0.245)	(2.385)	(0.403)	(0.560)	(0.511)	(1.815)			
Singapore/Malaysia	0.005	1.142*	-0.098	0.230	-4.099	-0.355	1.907*	1.103	-3.368	0.304	0.297	0.265
	(0.006)	(0.187)	(0.129)	(0.239)	(2.867)	(0.323)	(0.456)	(0.746)	(2.192)			
Spain	-0.002	0.550*	-0.105	0.416*	4.370*	0.237	-1.048	0.094	1.803	0.198	0.196	0.170
	(0.004)	(0.105)	(0.088)	(0.197)	(2.017)	(0.240)	(0.653)	(0.459)	(1.429)			
Sweden	0.003	0.718*	-0.059	0.214	0.756	0.199	-0.160	0.690	1.052	0.243	0.246	0.249
	(0.004)	(0.098)	(0.091)	(0.204)	(2.129)	(0.329)	(0.624)	(0.387)	(1.445)			
Switzerland	0.001	0.873*	0.019	0.469*	-2.456	0.360	0.227	-0.261	0.112	0.472	0.471	0.447
	(0.003)	(0.078)	(0.081)	(0.153)	(1.672)	(0.207)	(0.362)	(0.353)	(1.148)			
United Kingdom	0.003	1.273*	0.117	0.093	-2.830	-0.320	0.104	-0.067	-1.947	0.451	0.453	0.457
	(0.005)	(0.153)	(0.091)	(0.164)	(2.750)	(0.265)	(0.291)	(0.597)	(1.956)			
United States	-0.001	1.004*	-0.005	-0.415*	-1.269	-0.322	0.110	-0.054	-0.651	0.776	0.777	0.741
	(0.002)	(0.041)	(0.038)	(0.091)	(0.880)	(0.131)	(0.181)	(0.191)	(0.603)			
$\chi^2_{18}\ \beta_{ij} = 0$		1930.48	17.429	107.471	24.965	34.485	34.289	19.402	13.704			
p-value		[0.000]	[0.494]	[0.000]	[0.126]	[0.011]	[0.012]	[0.367]	[0.748]			
$\chi^2_{17}\ \beta_{ij} = \beta_j$		120.346	17.219	87.107	24.758	34.409	31.755	19.316	13.402			
p-value		[0.000]	[0.440]	[0.000]	[0.100]	[0.007]	[0.016]	[0.311]	[0.709]			

Table 2 (continued)

The R-squares are adjusted for degrees of freedom. * denotes more than two standard errors from zero. When 4 risk factors are used the variables are *wdret*, *dG10fx*, *dG7elt* and *doil*. Standard errors and chi-square statistics are heteroskedasticity consistent. The world excess return, *wdret* is the arithmetic return on the Morgan Stanley Capital International world equity index (including dividends) less the Ibbotson Associates one month bill rate. The change in the Eurodollar-Treasury yield spread, *dted*, is the difference between the 90-day Eurodollar yield and the 90 day Treasury bill yield (from Federal Reserve Bulletin). The log change in the G-10 foreign exchange rate, *dG10fx*, is the difference in the trade weighted dollar per foreign exchange rate of 10 industrialized countries (G-7 plus the Netherlands, Belgium, Sweden, and Switzerland) from International Monetary Fund. The unexpected inflation for the G-7 countries, *G7ui*, is derived from a time-series model applied to an aggregate G-7 inflation rate where the (varying) weights in the aggregation are determined by country weights in total G-7 gross domestic product. Change in long term expected G-7 inflation, *dG7elt*, is a result of projecting the four year moving average of G-7 inflation on a set of instrumental variables. The change in the price of oil, *doil*, is the change in the average U.S. dollar price of per barrel at the wellhead from 1974-1989 and the posted West Texas Intermediate price from 1969-1973. This variable is divided by 100. *dG7ip* is the change in G-7 industrial production. The growth rate is the logarithmic difference in this aggregate index. The G-7 real interest rate, *G7rtb*, is calculated by aggregating individual countries' short term interest rates minus inflation rates using weights based on quarterly shares in G-7 gross domestic product.

Table 3

Unconditional asset pricing tests using generalized method of moments:

$$u_t = r_t - (f + \lambda \otimes \iota)\beta$$

where r represents the asset excess returns, f are the demeaned world risk factors, λ are the risk premiums associated with the factors and β are the asset sensitivites to the world risk factors. In the first test in columns 2-4, there is a single risk factor, the world market return, and $f + \lambda \otimes \iota$ is replaced with the excess market return. In this case, the reported risk premium is the average market excess return over the sample and the standard error of the mean is in parentheses. In the second test in columns 5-9, f is specified to be four demeaned world risk factors. The instruments in the GMM estimations are a constant and the risk measure(s). The χ^2 test is the minimized values of the GMM criterion function for the system. The sample is 1970:2–1989:12 (239 observations)

	One factor model			Four factor model				
Source of risk Country	wdret β	Average return(%)	Average pricing error(%)	wdret β	dG10fx β	dG7elt β	doil β	Average pricing error(%)
Australia	1.115 (0.140)	0.468	-0.140 (0.427)	0.980 (0.145)	0.130 (0.210)	-1.273 (0.395)	0.693 (0.620)	-0.151 (0.263)
Austria	0.386 (0.072)	0.756	0.546 (0.355)	0.251 (0.088)	0.823 (0.216)	0.124 (0.489)	-0.256 (0.501)	0.094 (0.205)
Belgium	0.899 (0.077)	0.897	0.407 (0.300)	0.831 (0.086)	0.510 (0.171)	0.511 (0.288)	-0.591 (0.457)	-0.022 (0.221)
Canada	1.031 (0.066)	0.451	-0.111 (0.250)	1.025 (0.066)	-0.261 (0.113)	-0.381 (0.209)	0.486 (0.289)	-0.122 (0.209)
Denmark	0.651 (0.076)	0.816	0.461 (0.321)	0.578 (0.071)	0.481 (0.154)	0.272 (0.191)	0.149 (0.366)	0.031 (0.254)
France	1.066 (0.086)	0.729	0.148 (0.369)	1.016 (0.089)	0.434 (0.178)	0.403 (0.321)	-0.330 (0.364)	-0.293 (0.311)
Germany	0.782 (0.083)	0.651	0.224 (0.328)	0.717 (0.093)	0.424 (0.186)	0.206 (0.404)	-0.273 (0.429)	-0.119 (0.255)
Hong Kong	1.310 (0.166)	1.630	0.915 (0.745)	1.182 (0.207)	0.706 (0.585)	0.299 (0.436)	1.125 (0.674)	0.119 (0.558)
Italy	0.783 (0.095)	0.296	-0.130 (0.450)	0.706 (0.097)	0.392 (0.235)	0.267 (0.309)	-1.201 (0.760)	-0.315 (0.368)
Japan	0.918 (0.089)	1.313	0.813 (0.303)	0.915 (0.079)	0.519 (0.133)	0.777 (0.261)	-0.299 (0.314)	0.227 (0.199)
Netherlands	0.984 (0.064)	0.830	0.293 (0.239)	0.931 (0.070)	0.202 (0.111)	-0.014 (0.220)	0.029 (0.245)	0.039 (0.197)
Norway	1.028 (0.113)	0.932	0.371 (0.456)	0.960 (0.124)	0.185 (0.233)	0.081 (0.385)	1.233 (0.523)	-0.077 (0.367)
Singapore/Malaysia	1.183 (0.169)	1.114	0.469 (0.510)	1.091 (0.181)	0.263 (0.234)	-0.489 (0.294)	1.955 (0.424)	-0.027 (0.325)
Spain	0.641 (0.095)	0.361	0.011 (0.374)	0.580 (0.106)	0.312 (0.190)	0.249 (0.222)	-0.991 (0.580)	-0.141 (0.306)
Sweden	0.774 (0.081)	0.964	0.542 (0.347)	0.737 (0.095)	0.179 (0.188)	0.318 (0.292)	-0.242 (0.577)	0.298 (0.293)
Switzerland	0.915 (0.070)	0.548	0.049 (0.272)	0.866 (0.077)	0.410 (0.147)	0.397 (0.204)	0.008 (0.341)	-0.405 (0.196)
United Kingdom	1.274 (0.134)	0.761	0.066 (0.369)	1.254 (0.149)	0.100 (0.161)	-0.180 (0.224)	0.056 (0.294)	-0.169 (0.317)
United States	0.968 (0.039)	0.380	-0.147 (0.151)	0.994 (0.037)	-0.395 (0.082)	-0.378 (0.126)	0.071 (0.175)	-0.026 (0.084)
Mean absolute pricing error			0.325					0.149
Value-weighted[a] mean absolute pricing error			0.444					0.154
	wdret λ	χ^2_{18} [p-value]		wdret λ	dG10fx λ	dG7elt λ	doil λ	χ^2_{14} [p-value]
Risk premiums	0.545 (0.271)	20.697 [0.295]		0.717 (0.049)	0.602 (0.340)	0.211 (0.252)	0.155 (0.167)	7.950 [0.892]

[a] Value weights are the proportion of the Morgan Stanley Capital International world index represented by each of the 18 country indices. Standard errors and chi-square statistics are heteroskedasticity consistent. The world excess return, *wdret* is the arithmetic return on the Morgan Stanley Capital International world equity index (including dividends) less the Ibbotson Associates one month bill rate. The log change in the G-10 foreign exchange rate, *dG10fx*, is the difference in the trade weighted dollar per foreign exchange rate of 10 industrialized countries (G-7 plus the Netherlands, Belgium, Sweden, and Switzerland). Change in long term expected G-7 inflation, *dG7elt*, is a result of projecting the four year moving average of G-7 inflation on a set of instrumental variables. The change in the price of oil, *doil*, is the change in the average U.S. dollar price of per barrel at the wellhead from 1974-1989 and the posted West Texas Intermediate price from 1969-1973. This variable is divided by 100.