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# The World Price of Covariance Risk

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**ABSTRACT**

In a financially integrated global market, the conditionally expected return on a portfolio of securities from a particular country is determined by the country's world risk exposure. This paper measures the conditional risk of 17 countries. The reward per unit of risk is the world price of covariance risk. Although the tests provide evidence on the conditional mean variance efficiency of the benchmark portfolio, the results show that countries' risk exposures help explain differences in performance. Evidence is also presented which indicates that these risk exposures change through time and that the world price of covariance risk is not constant.

In a world with increasingly integrated financial services, why do industrialized countries have much different average stock returns? Why have Japanese stocks done so well compared to all other countries through 1989 and so poorly recently? If we view countries as stock portfolios in a global market, asset pricing theory suggests that cross-sectional differences in countries' risk exposures should explain the cross-sectional variation in expected returns.

This paper tests whether conditional versions of the Sharpe (1964) and Lintner (1965) asset pricing model are consistent with behavior of returns in 17 countries. Country risk is defined as the conditional sensitivity (or covariance) of the country return to a world stock return. This risk is allowed to vary through time. The reward per unit of sensitivity is the world price of covariance risk. Conditional covariances are calculated for each country. The differences in the countries' conditional covariances should explain the differences in national performance if there is only one source of risk.

The empirical results indicate that the time-varying covariances are able to capture some, but not all, of the dynamic behavior of the country returns. This could be due to incomplete market integration, the existence of more than one source of risk, or some other misspecification. The world price of covariance risk is also calculated. This measure exhibits significant time

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variation, which indicates the investor's expected compensation per unit of country risk exposure changes through time in a partially predictable way.

My paper is organized as follows. In the first section, the econometric methodology is introduced that measures the conditional moments of the country and world stock returns. The data are documented in the second section. The empirical results are presented in the third section. Some concluding remarks are offered in the final part.

## I. Methodology

### *A. An International CAPM with Time-Varying Moments*

A number of studies have examined asset pricing relations with international data. The tests of the Sharpe (1964) and Lintner (1965) asset pricing model have been executed by Solnik (1974), Stehle (1977), and others.[1] Multifactor asset pricing models have been tested by Cho, Eun, and Senbet (1986), Hamao (1988), Gultekin, Gultekin, and Penati (1989), and Korajczyk and Viallet (1989). Finally, Wheatley (1988) tested the restrictions implied by the consumption-based capital asset pricing model. All of these studies assess unconditional moment restrictions implied by the models, i.e., do cross-sectional differences in *average* risk explain the differences in *average* returns?

This study focuses on *conditional* asset pricing restrictions. The conditional version of the Sharpe (1964) and Lintner (1965) capital asset pricing model restricts the conditionally expected return on an asset to be proportional to its covariance with the market portfolio. The proportionality factor is the price of covariance risk which is the expected compensation (expected return) that the investor receives for taking on a unit of covariance risk. The model is:

$$E[r_{jt} \mid \Omega_{t-1}] = \frac{E[r_{mt} \mid \Omega_{t-1}]}{\operatorname{Var}[r_{mt} \mid \Omega_{t-1}]} \operatorname{Cov}[r_{jt}, r_{mt} \mid \Omega_{t-1}], \tag{1}$$

where $r_{jt}$ is the return on a portfolio of country $j$ equity from time t − 1 to t in excess of a risk free return, $r_{mt}$ is the excess return on the world market portfolio, and $\Omega_{t-1}$ is the information set that investors use to set prices. The ratio of the conditionally expected return on the market index $E[r_{mt} \mid \Omega_{t-1}]$ to the conditional variance of the market index $\operatorname{Var}[r_{mt} \mid \Omega_{t-1}]$ is the world price of covariance risk.

There are a number of issues that arise when applying the model to international data. For the Sharpe-Lintner model to hold internationally, Stulz (1981) demonstrates that some auxiliary assumptions must be made. A sufficient assumption is perfect correlation between the world market portfolio and world consumption. This assumption or equivalent distributional assumptions are implicit in this study's application of the Sharpe-Lintner

[1] See the references in Solnik (1977) and Stulz (1984).

model.[2] Alternatively, one can view the model as testing the mean-variance efficiency of the world market portfolio.

The empirical implementation of the model takes the view of a global investor whose returns are calculated in U.S. dollars. In other words, the investor is unhedged in exchange rates. Consistent with this assumption, the nominal return on the U.S. Treasury bill that is 30 days to maturity is conditionally risk-free. That is, when the investor buys a 30-day bill, the nominal return is conditionally known. Furthermore, the excess return is real because the U.S. inflation component in the stock return is cancelled out by the inflation component in the bill return.[3]

## *B. Econometric Specifications*

Some additional structure must be imposed before equation (1) is testable. In particular, a model must be specified for the conditional first moments. Assume that investors process information using a linear filter:[4]

$$u_{jt} = r_{jt} - \mathbf{Z}_{t-1}\boldsymbol{\delta}_j, \tag{2}$$

where $u_{jt}$ is the investor's forecast error for the return on country $j$, $\mathbf{Z}_{t-1}$ are $l$ information variables that are available to the investor, and $\boldsymbol{\delta}_j$ is a set of time-invariant weights that the investor uses to derive the conditionally expected returns.

Given the assumption on the conditional first moments, we can rewrite (1):

$$\mathbf{Z}_{t-1}\boldsymbol{\delta}_j = \frac{\mathbf{Z}_{t-1}\boldsymbol{\delta}_m}{E[u_{mt}^2|\mathbf{Z}_{t-1}]} E[u_{jt}u_{mt}|\mathbf{Z}_{t-1}], \tag{3}$$

where $u_{mt}$ is the investor's forecast error for the return on the world market portfolio. Notice that $E[u_{mt}^2|\mathbf{Z}_{t-1}]$ is the definition of conditional variance and $E[u_{jt}u_{mt}|\mathbf{Z}_{t-1}]$ is the conditional covariance. Also, equation (3) is conditioned on $\mathbf{Z}_{t-1}$ which is the subset of the true information set.[5] Next,

[2]Stulz's (1981) international capital asset pricing model assumes that representative investor has state-independent von Neuman-Morgenstern expected utility. Some recent research has considered non-von Neuman-Morgenstern utility. In particular, the assumption that the investor is indifferent about the resolution of uncertainty is dropped. Epstein and Zin (1988) and Giovannini and Weil (1988) provide a description of the conditions where the conditional Sharpe-Lintner model obtains.

[3]However, this does not imply that real returns are independent of inflation. Evidence on the relation between stock returns and inflation is presented in Gultekin (1983) and Solnik (1983).

[4]Sufficient distributional conditions that imply linear conditional expectations involve the joint distribution of the returns and the information variables falling into the class of spherically invariant distributions. This class of distributions is described in Vershik (1964) and Blake and Thomas (1968) and applied to conditionally expected stock returns in Harvey (1990).

[5]Since $\mathbf{Z} \subset \boldsymbol{\Omega}$ (the true information set), the expectation of the true conditional covariance is not the covariance conditioned on $\mathbf{Z}$. Conditioning on the specified information, $E[\mathrm{Cov}(r_{jt}, r_{mt}|\boldsymbol{\Omega}_{t-1})|\mathbf{Z}_{t-1}] = \mathrm{Cov}(r_{jt}, r_{mt}|\mathbf{Z}_{t-1}) - \mathrm{Cov}(E[r_{jt}|\boldsymbol{\Omega}_{t-1}], E[r_{mt}|\boldsymbol{\Omega}_{t-1}]|\mathbf{Z}_{t-1})$. A similar result holds for the conditional variance (where $r_{jt} = r_{mt}$). As a result, equation (3) should be viewed as an approximation. The model conditioned upon $\boldsymbol{\Omega}$ is untestable because $\boldsymbol{\Omega}$ is not observable.

multiply both sides of (3) by the conditional variance:

$$E\left[u_{mt}^{2}\mathbf{Z}_{t-1}\boldsymbol{\delta}_{j}\,|\,\mathbf{Z}_{t-1}\right] = E\left[u_{jt}u_{mt}\mathbf{Z}_{t-1}\boldsymbol{\delta}_{m}\,|\,\mathbf{Z}_{t-1}\right]. \tag{4}$$

Notice that the conditionally expected returns on the market and the country portfolio are moved inside the expectation operators. This can be done because they are known conditional on the information $\mathbf{Z}_{t-1}$. The deviation from the expectation is:

$$h_{jt} = u_{mt}^{2}\mathbf{Z}_{t-1}\boldsymbol{\delta}_{j} - u_{jt}u_{mt}\mathbf{Z}_{t-1}\boldsymbol{\delta}_{m}, \tag{5}$$

where $h_{jt}$ is a disturbance that should be unrelated to the information under the null hypothesis that the model is true. If $h_{jt}$ is divided by the conditional variance of the world market return, it can be interpreted as the deviation of the country's return from the return predicted by the model. In other words, $h_{jt}$ is a pricing error. A negative pricing error implies the model is overpricing while a positive pricing error indicates that the model is underpricing.

The econometric model to test the asset pricing restrictions is formed by combining equations (2) and (5):

$$\varepsilon_{t} = \left(\mathbf{u}_{t} \quad u_{mt} \quad \mathbf{h}_{t}\right) = \begin{pmatrix} \left[\mathbf{r}_{t} - \mathbf{Z}_{t-1}\boldsymbol{\delta}\right]' \\ \left[r_{mt} - \mathbf{Z}_{t-1}\boldsymbol{\delta}_{m}\right]' \\ \left[u_{mt}^{2}\mathbf{Z}_{t-1}\boldsymbol{\delta} - u_{mt}\mathbf{u}_{t}\mathbf{Z}_{t-1}\boldsymbol{\delta}_{m}\right]' \end{pmatrix}, \tag{6}$$

where $\mathbf{u}$ is a $1 \times n$ (number of countries) vector of innovations in the conditional means of the country returns. The model implies that $E[\varepsilon_{t}\,|\,\mathbf{Z}_{t-1}] = 0$. With $n$ countries, there are $n + 1$ columns of innovations in the conditional means ($\mathbf{u}$ and $u_{m}$) and $n$ columns in $\mathbf{h}$. If there are $l$ information variables, there are $[l \times (2n + 1)]$ orthogonality conditions. However, there are $[l \times (n + 1)]$ parameters to estimate, which implies there are $l \times n$ overidentifying conditions.[6]

Hansen's (1982) generalized method of moments (GMM) is used to estimate the parameters in equation (6). The GMM forms a vector of the orthogonality conditions $\mathbf{g} = vec(\varepsilon'\mathbf{Z})$ where $\varepsilon$ is the matrix of forecast errors for $T$ observations and $2n + 1$ equations, and $\mathbf{Z}$ is a $T \times l$ matrix of observations on thepredetermined instrumental variables. The parameter vector $\boldsymbol{\delta}$ is chosen to make the orthogonality conditions as close to zero as possible by minimizing the quadratic form $\mathbf{g}'\mathbf{w}\mathbf{g}$ where the $\mathbf{w}$ is symmetric weighting matrix that defines the metric used to make $\mathbf{g}$ close to zero. The consistent estimate of $\mathbf{w}$ is formed by

$$\left[\sum_{t=1}^{T}(\varepsilon_{t}\otimes\mathbf{Z}_{t-1})'(\varepsilon_{t}\otimes\mathbf{Z}_{t-1})\right]^{-1}.$$

However, $\varepsilon$ depends on the parameters. As a result, the estimation proceeds in stages. An initial estimate of the parameters is obtained by using an

[6]This formulation is explored in the context of returns on the New York Stock Exchange by Harvey (1989) and Huang (1989).

identity matrix for **w**. These parameters are used to calculate $\varepsilon$ and a new weighting matrix. The estimation procedure is repeated with this new weighting matrix. Hansen provides the conditions that guarantee that the estimates are consistent and asymptotically normal.

The minimized value of this quadratic form is distributed $\chi^2$ under the null hypothesis with degrees of freedom equal to the number of orthogonality conditions minus the number of parameters. This $\chi^2$ statistic, which is known as the test of the *overidentifying* restrictions, provides a goodness of fit test for the model. A high $\chi^2$ statistic means that the disturbances are correlated with the instrumental variables. This is a symptom of model misspecification.

In a system of many equations, the test of the overidentifying restrictions does not tell us where the model is failing. One possible solution is to estimate equation (6) for individual countries. Even with one country, (6) provides a test of the model's restriction that the conditionally expected excess return on a country portfolio is proportional to its conditional covariance with the world return. However, the single country test does not impose the cross-country restriction that the proportionality factor (the world price of covariance risk) is the same for each country. The single country tests are weaker because fewer restrictions are being imposed. However, statistical rejections in the single country estimation may provide valuable insights as to where the model is failing. Another possibility is to examine subsets of the disturbances; in particular, the errors implied by the asset pricing model's restrictions **e**. An additional test is to regress the disturbances for a particular country portfolio on the set of instruments. If the model is correct, the $R^2$ should be zero.

### C. *World Price of Covariance Risk*

In the framework of equation (6), all of the conditional moments—the means, variances, and covariances—are allowed to change through time. If some of these moments are constant, then more powerful tests can be constructed by imposing this additional structure.

Traditionally, asset pricing tests have focused on whether expected returns are proportional to the expected return on a benchmark portfolio. This restriction can be imposed and tested:

$$\mathbf{k}_t = \mathbf{r}_t - r_{mt}\boldsymbol{\beta}, \tag{7}$$

where $\beta$ is a $n$-vector of coefficients. This coefficient vector can represent the ratios of conditional covariances of the country excess returns to the conditional variance of the benchmark return.[7] The model implies that

[7]There is an alternative interpretation of equation (7). Since the coefficient is not restricted in the estimation to be the ratio of the conditional covariance of the country excess return to the conditional variance of the benchmark return, (7) can be interpreted as a single factor latent variables test [see Hansen and Hodrick (1983), Gibbons and Ferson (1985), and Ferson (1990)]. In this test, the coefficient represents the ratio of the covariance of the country's return and the single factor to the covariance of the benchmark's return and the single factor.

$E[\mathbf{k}_t | \mathbf{Z}_{t-1}] = \mathbf{0}$ where $\mathbf{k}_t$ is the pricing error associated with this implementation of the asset pricing model. There are $l \times n$ orthogonality conditions and $n$ parameters to estimate leaving $l \times (n-1)$ orthogonality conditions to be tested. An advantage of equation (7) is that the models for conditional means need not be specified.

Another version of the model assumes that the reward to volatility ratio is constant. In our context, the reward to volatility is the world price of covariance risk. Imposing this restriction results in:

$$\mathbf{e}_t = \mathbf{r}_t - \lambda \mathbf{u}_t u_{mt}, \tag{8}$$

where $\lambda$ is the ratio of the conditionally expected return on the market divided by the conditional variance, and $\mathbf{e}_t$ is the pricing error associated with the assumption of a constant price of covariance risk. In contrast to equation (7), it is necessary to have a model for the conditional means in (8). The system is:

$$\boldsymbol{\varepsilon}_t = (\mathbf{u}_t \quad u_{mt} \quad \mathbf{e}_t) = \begin{pmatrix} [\mathbf{r}_t - \mathbf{Z}_{t-1}\boldsymbol{\delta}]' \\ [r_{mt} - \mathbf{Z}_{t-1}\boldsymbol{\delta}_m]' \\ [\mathbf{r}_t - \lambda(u_{mt}\mathbf{u}_t)]' \end{pmatrix}'. \tag{9}$$

With $n$ assets, there are $n + 1$ columns in $\mathbf{u}$ and $u_m$ and $n$ columns in $\mathbf{e}$. If there are $l$ instrumental variables, there are $[l \times (2n + 1)]$ orthogonality conditions and $[1 + l \times (n + 1)]$ parameters to estimate. As a result, there are $l \times n - 1$ overidentifying restrictions to be tested.[8]

It is possible to simplify the estimation in equation (9) by noting that $E[u_{mt}u_{jt} | \mathbf{Z}_{t-1}] = E[u_{mt}r_{jt} | \mathbf{Z}_{t-1}]$.[9] This allows us to drop the $n$ equations for the conditional means of the country returns. The more parsimonious system is:

$$\boldsymbol{\eta}_t = (u_{mt} \quad \mathbf{e}_t) = \begin{pmatrix} [r_{mt} - \mathbf{Z}_{t-1}\boldsymbol{\delta}_m]' \\ [\mathbf{r}_t - \lambda(u_{mt}\mathbf{r}_t)]' \end{pmatrix}'. \tag{10}$$

This system has $n + 1$ equations and $l \times (n + 1)$ orthogonality conditions. With $l + 1$ parameters, there are $l \times n - 1$ overidentifying restrictions. This

[8]Using a different data set, Giovannini and Jorion (1989) provide maximum likelihood estimation of equation (8) where the covariance is parametered to be a non-stochastic function of the current information set.

[9]This follows from

$$\begin{aligned} E[u_{mt}u_{jt} | \mathbf{Z}_{t-1}] &= E[u_{mt}(r_{jt} - \mathbf{Z}_{t-1}\boldsymbol{\delta}_j) | \mathbf{Z}_{t-1}] \\ &= E[u_{mt}r_{jt} | \mathbf{Z}_{t-1}] - E[u_{mt}\mathbf{Z}_{t-1}\boldsymbol{\delta}_j | \mathbf{Z}_{t-1}] \\ &= E[u_{mt}r_{jt} | \mathbf{Z}_{t-1}] - E[u_{mt} | \mathbf{Z}_{t-1}]\mathbf{Z}_{t-1}\boldsymbol{\delta}_j \\ &= E[u_{mt}r_{jt} | \mathbf{Z}_{t-1}], \end{aligned}$$

since $E[u_{mt} | \mathbf{Z}_{t-1}] = 0$.

is the same number of restrictions as equation (9), and hence the systems (9) and (10) are asymptotically equivalent. However, in the context of the estimation, the dimensionality of the **w** matrix is much smaller in equation (10) and much more computationally manageable.

### D. Time Variation in the Reward Per Unit of Risk

The constancy of the world price of covariance risk can be tested. The definition of a constant reward to volatility ratio is:

$$\frac{E[r_{mt} | \mathbf{Z}_{t-1}]}{E[u_{mt}^2 | \mathbf{Z}_{t-1}]} = \lambda. \tag{11}$$

Multiply both sides by the conditional variance:

$$E[r_{mt} | \mathbf{Z}_{t-1}] = \lambda E[u_{mt}^2 | \mathbf{Z}_{t-1}]. \tag{12}$$

This implies that the conditional mean of the market return is proportional to the conditional variance. To test this assumption, the following system can be estimated:

$$\boldsymbol{\xi}_t = (u_{mt} \quad e_{mt}) = \begin{pmatrix} [r_{mt} - \mathbf{Z}_{t-1}\boldsymbol{\delta}_m]' \\ [r_{mt} - \lambda u_{mt}^2]' \end{pmatrix}'. \tag{13}$$

This system has $l - 1$ overidentifying conditions.

It is also interesting to look at country-specific reward to volatility, i.e., the ratio of the country's conditionally expected return to its own conditional variance. If global markets are not financially integrated, then the ratio of country expected return to country variance is the relevant measure that transforms conditional covariance into expected return. This country-specific reward to volatility can be estimated by substituting the country returns into equation (13). The system also provides a test of whether the ratio is constant through time.

### E. Country Performance

A country's performance is determined by its return in excess of the expected return given its riskiness. The pricing error is a measure of performance. A positive pricing error implies that the country earned more than expected given its level of risk. Of course, performance is measured under the null hypothesis that the model is correct. If the model is misspecified, then we cannot say which countries earned abnormal returns.

A useful measure of performance is the mean pricing error. In the context of system (9) which assumes a constant world price of covariance risk, this measure is defined as:

$$\text{Mean Error} = \frac{1}{T}\sum_{t=1}^{T} e_{jt}. \tag{14}$$

For Japan, if this measure is small or negative, then the results would indicate that the Japanese market has not done as well as is popularly believed. The patterns of the abnormal performance measure could also be examined in particular subperiods.

Another summary measure is the mean absolute pricing error:

$$\text{Mean Absolute Error} = \frac{1}{T}\sum_{t=1}^{T} |e_{jt}|, \tag{15}$$

where $|\cdot|$ is the absolute value operator. The mean pricing error indicates whether the performance is at the level of the expectations on average. Two countries may have the same mean error but very different performance through time. The mean absolute value captures the magnitude of the deviations from the mean. However, there is no reason to believe that the mean absolute error should be zero. This measure is included as a summary measure of the difference between the expected returns conditional on the model being correct and the actual returns.[10]

## II. Data and Summary Statistics

### A. Data Sources

Most of the data in this study are drawn from Morgan Stanley Capital International (MSCI). Monthly data on equity indices for 16 OECD countries and Hong Kong are available from December 1969 to May 1989.[11] These indices are value weighted and are calculated with dividend reinvestment. Morgan Stanley also calculates a value weighted world equity index which serves as the world market portfolio.

The MSCI international indices are composed of stocks that broadly represent stock composition in the different countries. Almost all the stocks (99%) can be readily purchased by non-nationals.[12] Although the MSCI indices are weighted towards larger capitalization stocks, the returns are similar to widely quoted country index returns. For example, there is a 99.1% correlation between the MSCI U.S. return and the New York Stock Exchange value-weighted return calculated by the Center for Research in Security Prices (CRSP) at the University of Chicago. For Japan, there is a 93.8%

[10]An alternative measure of the deviation is the root mean squared error. Since similar patterns were found in the mean absolute errors and the root mean squared errors, only the mean absolute errors are reported.

[11]The 16 OECD countries are Australia, Austria, Belgium, Canada, Denmark, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, the United Kingdom, and the United States. Morgan Stanley also has data on Finland and New Zealand but only since December 1987. Data is available for Singapore/Malaysia, but dividend data is not available for the full period. As a result, these countries are omitted from the empirical analysis.

[12]Cumby and Glen (1990) note that only 1% of the securities followed by Morgan Stanley Capital International are not available to non-nationals. This group is composed of Swedish bank stocks and some Swiss registered shares.

**Appendix Table A-I**

## The Composition of the Morgan Stanley Capital International Indices On March 31, 1989[a]

| Country | Number of Companies Included | Weight in MSCI World Index | Market value of Companies billion U.S. $ |
|---|---|---|---|
| Austria | 15 | 0.1 | 5.6 |
| Belgium | 22 | 0.6 | 32.9 |
| Denmark | 27 | 0.3 | 15.1 |
| Finland | 21 | 0.3 | 13.2 |
| France | 83 | 2.6 | 132.9 |
| Germany | 57 | 2.7 | 141.7 |
| Italy | 68 | 1.4 | 73.8 |
| Netherlands | 24 | 1.3 | 70.0 |
| Norway | 18 | 0.2 | 12.7 |
| Spain | 31 | 1.0 | 50.0 |
| Sweden | 38 | 1.0 | 52.3 |
| Switzerland | 52 | 1.6 | 77.6 |
| U.K. | 136 | 8.4 | 438.6 |
| Europe | 592 | 21.4 | 1116.4 |
| Australia | 66 | 1.3 | 71.0 |
| Hong Kong | 21 | 0.9 | 44.7 |
| Japan | 265 | 42.9 | 2224.9 |
| New Zealand | 13 | 0.2 | 8.0 |
| Singapore/Malaysia | 53 | 0.6 | 29.3 |
| EAFE | 1021 | 67.3 | 3494.3 |
| Canada | 89 | 2.5 | 131.0 |
| South African Gold Mines | 21 | 0.2 | 10.3 |
| U.S. | 335 | 30.0 | 1555.5 |
| World | 1466 | 100.0 | 5191.1 |

[a]From Morgan Stanley Capital International *Perspective* First Quarter, 1989.

correlation between the MSCI return and the Nikkei 255 return.[13] An important difference between the MSCI indices and other national indices such as CRSP is the exclusion of investment companies and foreign domiciled companies. These stocks are excluded to avoid double counting. Appendix Table A-I provides a description of the number of companies included in the country indices and the market value of these companies as of March 31, 1989. The weight that each country commands in the MSCI world index is also reported.

[13]Over the 1970:2–1988:12 period, the mean return on the CRSP value weighted index is 0.40% per month with 4.83% standard deviation. Over the same period, the MSCI U.S. index has a mean return of 0.33% and standard deviation of 4.70%. The difference in mean return and standard deviation is due to the MSCI index using fewer small stocks. For Japan, the mean return on the Nikkei 225 is 1.13% with a standard deviation of 5.70%. The MSCI Japan index has an average return of 1.34% and a standard deviation of 6.09%.

McDonald (1989) and French and Poterba (1989) show that the MSCI world index gives too much weight to the Japanese stocks because of the large amount of cross-corporate ownership. I investigated an alternative index, the FT-Actuaries World Index which is compiled by *The Financial Times*, Goldman, Sachs and Co., and Country NatWest/Wood Mackenzie. Unfortunately, the FT-Actuaries index suffers from the same problem. In March 1989, Japan composed 42.9% of the MSCI index and as of June 1989, 40.7% of the FT-Actuaries index. The cross ownership problem is not restricted to Japan. Substantial intercorporate ownership is prevalent in other countries such as Germany.

All returns are calculated in excess of the U.S. Treasury bill that is closest to 30 days to maturity on the last trading day of the month. Data from 1970–1988 are drawn from the CRSP Government Bond File. The data for 1989 are from the *Wall Street Journal*. Holding period returns are calculated in the same way as Fama (1984).

The selection of conditioning information is an important step. The instrumental variables should approximate the information that investors use to set prices. Given that expected returns change through time, the instrumental variables should have the ability to predict returns.

The empirical strategy involves a prespecification of two categories of instrumental variables: common and local instruments. The common set of instruments consists of an identical set of instruments for all countries. In contrast, the local instruments include country-specific variables. According to the model, time variation in the conditionally expected country returns has three potential sources: variation in the world expected return, changes in the volatility of the world return, and time-varying conditional covariances of the country return with the world return. The common instrument set is important for the first two sources. Local information, in addition to the common instruments, may be important in detecting changes in the country's conditional covariances.

The specification of the common instrumental variables were drawn from studies of U.S. stock returns since there is little research on time-variation in international returns. The information set contains: the lagged world excess stock return, a dummy variable for the month of January, the dividend yield on the Standard and Poor's 500 stock price index, the U.S. term structure premia, and the U.S. default risk yield spread.

The first information variable is the lagged excess return on the world index. Many studies beginning with Fama (1965) have documented some degree of autocorrelation in returns. A dummy variable for the month of January is also included. Keim (1983) documents that U.S. returns in January are systematically higher. Gultekin and Gultekin (1983) find disproportionately large January returns in many industrialized countries.

The U.S. dividend price ratio is also included in the information set. Fama and French (1988, 1989) show that this is an important explanatory variable for U.S. stock returns. Cutler, Poterba and Summers (1989) show that many international stock returns are also influenced by the dividend yield. Following Harvey (1989), the dividend yield on the S&P 500 is expressed in excess of the 1-month bill rate.

A measure of default risk is the fourth information variable. Keim and Stambaugh (1986) and Fama and French (1989) show that the junk bond spread is able to predict returns. The junk bond spread is the difference in yields between Moody's Baa and Aaa rated bonds. A shorter maturity term structure variable is also included. Following Campbell (1987) and Harvey (1989), the excess return on a 3-month bill is included as the final information variable. Campbell and Hamao (1989) show that measures of the term structure are able to explain returns in Japan as well as the United States.

A number of local instruments are considered: the lagged own-country return, the country-specific dividend yields, foreign exchange rate changes, local short-term interest rates and local long-term to short-term interest rate spreads. In the model testing, three sets of instrumental variables are used. The first group is the common instruments. In the second set, *Local Instruments A*, the common instruments are augmented by the inclusion of the local dividend yields. In the third set, *Local Instruments B*, the common instruments are again augmented by the local dividend yield. In addition, the own-country lagged excess return replaces the lagged word excess return.

### *B. Summary Statistics*

Unconditional means, standard deviations, and autocorrelations of the monthly returns are provided in Panel A of Table I. The highest mean excess return over the sample is from the Hong Kong market. Hong Kong also has the highest volatility. The United States has one of the lowest average returns. However, the volatility of the U.S. stock returns is lower than any other country.

While the first-order autocorrelation of the U.S. returns is not significant, there are some country returns that exhibit significant autocorrelation. High first-order autocorrelations are found in Austria, Belgium, Italy, Japan, Norway, and Spain. Significant seasonal autocorrelations are found in the returns of Austria and Denmark.

The world market portfolio is the value-weighted average of the country returns. The world portfolio has a lower standard deviation than any individual country.[14] Comparing the country portfolios to the world portfolio, there are seven countries (including the U.S.) that are unconditionally dominated by the world market portfolio. That is, given a choice between investing in one of these countries and the world portfolio, the world portfolio is a better investment for the risk-averse investor because it delivers a lower unconditional standard deviation and a higher unconditionally expected return.

[14]The differences in standard deviations across countries do not appear to be driven by the number of stocks included in the country indices. A cross-sectional regression (not reported) of the standard deviations on the number of firms included in each country index reported in Appendix Table A-I failed to detect a statistically significant relation.

**Table I**

**Summary Statistics for the Country Returns and the Instrumental Variables**

The statistics are based on monthly data from 1970:2–1989:5 (232 observations). The country returns are calculated in U.S. dollars in excess of the holding period return on the Treasury bill that is closest to 30 days to maturity. The dividend yields are the average (over the past year) monthly dividends divided by the current month price level. The returns and dividend yields are from Morgan Stanley Capital International. The instrumental variables are: the return for holding a 90-day U.S. Treasury bill for 1 month less the return on a 30-day bill (excess U.S. 3-month bill), the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds (U.S. junk bond), and the dividend yield on the Standard and Poor's 500 stock index less the return on a 30-day bill (excess U.S. dividend yield).

| | | | Autocorrelation | | | | | |
|---|---|---|---|---|---|---|---|---|
| Variable | Mean | Std. Dev. | $\rho_1$ | $\rho_2$ | $\rho_3$ | $\rho_4$ | $\rho_{12}$ | $\rho_{24}$ |
| A. Equity returns | | | | | | | | |
| Australia | .00440 | .08240 | −.012 | −.054 | .012 | .020 | −.035 | .011 |
| Austria | .00554 | .05382 | .146* | .191* | .117 | .015 | .176* | .019 |
| Belgium | .00867 | .06018 | .143* | .067 | .044 | .043 | .078 | .082 |
| Canada | .00440 | .05898 | .017 | −.104 | .074 | −.016 | −.044 | .034 |
| Denmark | .00718 | .05542 | .065 | .209* | .100 | .121 | −.177* | .088 |
| France | .00647 | .07362 | .109 | .045 | .121 | .029 | −.008 | −.011 |
| Germany | .00502 | .05984 | .033 | .057 | .062 | .073 | .003 | .025 |
| Hong Kong | .01684 | .12778 | .062 | −.035 | −.007 | −.050 | −.001 | −.043 |
| Italy | .00221 | .07774 | .171* | −.031 | .112 | .089 | .050 | .028 |
| Japan | .01341 | .06094 | .162* | .003 | .110 | .091 | .081 | −.024 |
| Netherlands | .00767 | .05562 | .076 | −.015 | .048 | −.107 | .066 | −.027 |
| Norway | .00930 | .08293 | .186* | −.007 | .152* | −.060 | .009 | .032 |
| Spain | .00355 | .06502 | .156* | .029 | −.037 | .116 | .044 | .115 |
| Sweden | .00938 | .06217 | .057 | .015 | .092 | .056 | .010 | −.015 |
| Switzerland | .00462 | .05676 | .080 | −.023 | .079 | .044 | .035 | −.024 |
| United Kingdom | .00736 | .07925 | .118 | −.069 | .099 | .016 | .001 | .016 |
| United States | .00373 | .04695 | .051 | −.049 | .004 | .001 | .065 | −.055 |
| World | .00533 | .04175 | .171* | −.026 | .064 | .027 | .083 | −.017 |

**Table I**—*Continued*

| Variable | Mean | Std. Dev. | Autocorrelation $\rho_1$ | $\rho_2$ | $\rho_3$ | $\rho_4$ | $\rho_{12}$ | $\rho_{24}$ |
|---|---|---|---|---|---|---|---|---|
| B. Dividend yields | | | | | | | | |
| Australia | .00365 | .00234 | .973* | .945* | .916* | .887* | .644* | .265* |
| Austria | .00359 | .00093 | .940* | .892* | .856* | .803* | .391* | .024 |
| Belgium | .00261 | .00057 | .977* | .957* | .931* | .910* | .718* | .444* |
| Canada | .00788 | .00247 | .985* | .966* | .948* | .931* | .783* | .537* |
| Denmark | .00362 | .00150 | .992* | .981* | .968* | .953* | .757* | .494* |
| France | .00427 | .00139 | .976* | .946* | .918* | .889* | .693* | .405* |
| Germany | .00357 | .00084 | .975* | .946* | .916* | .886* | .665* | .349* |
| Hong Kong | .00313 | .00138 | .957* | .899* | .840* | .783* | .471* | .069 |
| Italy | .00231 | .00065 | .944* | .882* | .817* | .762* | .289* | −.107 |
| Japan | .00168 | .00092 | .993* | .985* | .976* | .966* | .888* | .682* |
| Netherlands | .00491 | .00108 | .965* | .924* | .888* | .851* | .610* | .307* |
| Norway | .00298 | .00108 | .970* | .937* | .903* | .865* | .511* | .091 |
| Spain | .00587 | .00333 | .992* | .983* | .975* | .966* | .870* | .675* |
| Sweden | .00331 | .00113 | .978* | .954* | .938* | .919* | .765* | .539* |
| Switzerland | .00225 | .00038 | .934* | .873* | .819* | .770* | .435* | .115 |
| United Kingdom | .00426 | .00107 | .932* | .848* | .786* | .708* | .347* | .164* |
| United States | .00352 | .00077 | .981* | .952* | .923* | .895* | .642* | .425* |
| C. Other instrumental variables | | | | | | | | |
| Excess U.S. 3 month bill | .00078 | .00141 | .276* | .014 | .000 | .001 | −.096 | .006 |
| U.S. junk bond | .00108 | .00037 | .947* | .875* | .825* | .787* | .408* | .063 |
| Excess U.S dividend yield | −.00243 | .00183 | .900* | .819* | .744* | .681* | .416* | .160* |

*Significant at the 5% level based on an approximate standard error of $1/\sqrt{232} = .065$.

Interestingly, the world portfolio exhibits significant first-order autocorrelation, indicating that there is some predictable variation.

The unconditional means, standard deviations, and autocorrelations of the countries' dividend yields are provided in Panel B of Table I. Japan has, by far, the lowest dividend yield. Since the yield is an equally weighted 12-month moving average of dividends divided by the current month's price level, a high degree of autocorrelation is expected. The first 12 autocorrelations are significantly different from zero in all 17 countries.

Summary statistics are also provided for some of the common instrumental variables in Panel C of Table I. The excess returns on the 3-month Treasury bill have significant first order autocorrelation. Both the U.S. junk bond spread and the U.S. dividend yield spread show slower mean reversion. The U.S. junk bond spread autocorrelation drops to zero by the 24th lag. As already noted, the dividend yield is highly autocorrelated by construction. Notice that the mean dividend yield spread is negative, indicating that on average the bill rate is higher than the dividend yield on the S&P 500.

Unconditional correlations of the equity returns and the instrumental variables are provided in Table II. The first panel reveals that all stock returns move together on average. However, they may not move as closely together as one would expect. For example, the correlation between U.S. returns and U.K. returns is 49%; the correlation between U.S. and Japanese returns is only 27%.

The second panel shows that the dividend yields are not all positively correlated. The Australian dividend yield is negatively correlated with most other dividend yields. The U.S. and Japanese dividend yields are uncorrelated (−3%). The correlation of the U.S. and Canadian dividend yields is 84% which probably reflects the high degree of integration of the two economies.

To complement the summary statistics in Tables I and II, Figure 1 provides the traditional graph of mean return against variance. The unconditional minimum variance frontier calculated from the index returns is also plotted. Note that the returns are not excess returns. Unconditionally, the bill rate is not "risk free".

There are a number of interesting features to Figure 1. First, notice that Hong Kong is much different from the other portfolios—it has by far the highest volatility. Second, the two portfolios closest to the minimum variance frontier are the U.S. and Japan. Unconditionally, Japan does not dominate the United States. Third, the world market portfolio is the closest portfolio to the frontier.[15] Unconditional asset pricing tests would assess whether the world portfolio is far enough from the frontier to reject the restrictions of the asset pricing model. For example, Gibbons, Ross, and Shanken (1989) propose an exact $F$-statistic that tests the null hypothesis that the intercepts in the

[15] I also compared the minimum variance frontier based on 17 countries 1970:2–1987:12 to the minimum variance frontier based on 12 U.S. industry portfolios over the same period. The frontier based on the industry portfolios was always inside (less efficient than) the world frontier.

multivariate regression of the asset excess returns on the market excess return are jointly zero. When this test is executed on the 17 country portfolios, the probability value is 0.304.[16] Adler and Dumas (1983) provide an alternative specification where foreign exchange portfolios are included in the multivariate regression framework. The $p$-values are not substantially altered when this model is tested. Hence, these tests do not provide evidence against the hypothesis of unconditional mean variance efficiency. However, an examination of the individual regressions in both models reveals that Japan has a statistically significant intercept. So while the standard multivariate tests of the unconditional mean-variance efficiency do not provide evidence against the null hypothesis, these tests may not be very powerful.[17]

## III. Empirical Results

### A. *The Predictability of Country Stock Returns with the Common Instruments*

The predictability of the international equity returns using a common set of instruments is studied in Table III. The results indicate that there is significant time variation in most of the country returns. Furthermore, the value-weighted portfolio of all countries is the most predictable. The $R^2$ for the world market portfolio is 13.3%.

The results contrast with some other work on predicting international stock returns. For example, using country-specific dividend yields, Cutler, Poterba and Summers (1989) are only able to explain about 1% of the variance of the monthly returns over 1960–1988. They are able to account for 0.5% of the Japanese returns and 1.0% of the U.S. returns. This compares to 6.7% and 12.5% for the two countries, respectively, using the common information variables in Table III. Using a number of combinations of variables that include the U.S. and Japanese dividend yields, short-term rates, and long-term to short-term rate spreads, Campbell and Hamao (1989) report a 6.5% (largest) $R^2$ for Japan and a 10.0% (largest) $R^2$ for the U.S. over the 1971–1987 period.[18]

There are a number of interesting observations from Table III. For Japanese stock returns, the most important explanatory variable is the lagged world return.[19] The January dummy is more than one standard error from zero in

[16]The $F$-statistic for the same test with the Group of 7 ($G$-7) countries (Canada, France, Germany, Italy, Japan, United Kingdom, and United States) has a $p$-value of 0.103. The only country that has a significant intercept is Japan.

[17]Using a sample of 13 country indices and monthly data from 1982:1 to 1988:6, Cumby and Glen (1990) also cannot reject the null hypothesis of the mean-variance efficiency of the MSCI world index.

[18]Campbell and Hamao (1989) do not use the MSCI indices for Japan and the United States. They extend the index created by Hamao (1988) which is a value weighted index drawn from stocks listed in the first and second sections of the Tokyo Stock Exchange.

[19]I thought that this might be due to Japan being across the date line. The last trading day of the month for Europe and North America is only a few hours away from the first trading day of the month for Japan. However, when this regression is re-executed with the lagged Japanese return replacing the world index, the results are similar. Hamao, Masulis, and Ng (1990) also provide evidence of significant first-order autocorrelation in the Nikkei 225 return.

**Table II**

**Unconditional Correlations of the Country Returns and the Instrumental Variables**

The correlations are based on monthly data from 1970:2-1989:5 (232 observations). The country returns are calculated in U.S. dollars in excess of the holding period return on the Treasury bill that is closest to 30 days to maturity. The dividend yields are the average (over the past year) monthly dividends divided by the current monthly price level. The returns and dividend yields are from Morgan Stanley Capital International. The instrumental variables are: the excess return on the world index (*rwd*), the one month return for holding a 90-day U.S. Treasury bill less the return on a 30-day bill (*xustb3*), the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds (*usjunk*), and the dividend yield on the Standard and Poor's 500 stock index less the return on a 30-day bill (*xusdiv*).

A. Equity returns

| Portfolio | Aa | Au | Be | Ca | De | Fr | Ge | HK | It | Ja | Ne | No | Sp | Sw | Sz | UK | US |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| World | .57 | .29 | .62 | .75 | .47 | .61 | .54 | .41 | .42 | .63 | .74 | .50 | .41 | .50 | .68 | .67 | .86 |
| Australia | - | .16 | .32 | .59 | .28 | .37 | .28 | .36 | .24 | .28 | .40 | .41 | .32 | .36 | .42 | .46 | .47 |
| Austria | | - | .46 | .18 | .30 | .43 | .55 | .21 | .24 | .25 | .43 | .28 | .29 | .27 | .49 | .23 | .12 |
| Belgium | | | - | .38 | .45 | .64 | .63 | .32 | .42 | .46 | .65 | .53 | .40 | .44 | .67 | .50 | .41 |
| Canada | | | | - | .30 | .44 | .30 | .29 | .27 | .28 | .55 | .45 | .28 | .35 | .48 | .52 | .72 |
| Denmark | | | | | - | .35 | .39 | .31 | .26 | .39 | .45 | .33 | .30 | .30 | .44 | .35 | .33 |
| France | | | | | | - | .57 | .24 | .44 | .41 | .58 | .47 | .36 | .33 | .62 | .52 | .42 |
| Germany | | | | | | | - | .28 | .34 | .42 | .66 | .37 | .34 | .39 | .74 | .38 | .33 |
| Hong Kong | | | | | | | | - | .22 | .32 | .42 | .29 | .24 | .27 | .34 | .35 | .29 |
| Italy | | | | | | | | | - | .38 | .36 | .25 | .35 | .30 | .38 | .34 | .22 |
| Japan | | | | | | | | | | - | .45 | .17 | .35 | .33 | .43 | .35 | .27 |
| Netherlands | | | | | | | | | | | - | .52 | .38 | .43 | .73 | .62 | .56 |
| Norway | | | | | | | | | | | | - | .25 | .38 | .48 | .40 | .44 |
| Spain | | | | | | | | | | | | | - | .31 | .33 | .30 | .25 |
| Sweden | | | | | | | | | | | | | | - | .47 | .39 | .38 |
| Switzerland | | | | | | | | | | | | | | | - | .55 | .49 |
| United Kingdom | | | | | | | | | | | | | | | | - | .49 |
| United States | | | | | | | | | | | | | | | | | - |

**Table II**—*Continued*

B. Dividend yields

| Country | Aa | Au | Be | Ca | De | Fr | Ge | HK | It | Ja | Ne | No | Sp | Sw | Sz | UK | US |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Australia | - | .33 | −.50 | −.24 | −.33 | −.20 | −.29 | .35 | .19 | −.50 | −.10 | .04 | −.09 | −.22 | −.13 | −.02 | .01 |
| Austria | | - | .07 | .41 | .05 | .58 | .32 | .74 | .14 | −.14 | .70 | .35 | .23 | .23 | .73 | .73 | .60 |
| Belgium | | | - | .47 | .64 | .63 | .62 | −.04 | .19 | .63 | .52 | .34 | .36 | .60 | .47 | .27 | .40 |
| Canada | | | | - | .27 | .86 | .74 | .41 | −.08 | .00 | .73 | .74 | .81 | .55 | .71 | .60 | .84 |
| Denmark | | | | | - | .41 | .49 | −.30 | .19 | .76 | .51 | .16 | .03 | .84 | .33 | .38 | .28 |
| France | | | | | | - | .69 | .38 | .06 | .20 | .85 | .68 | .56 | .64 | .81 | .67 | .74 |
| Germany | | | | | | | - | .22 | −.20 | .12 | .77 | .41 | .63 | .66 | .70 | .59 | .76 |
| Hong Kong | | | | | | | | - | .04 | −.41 | .44 | .31 | .47 | −.12 | .56 | .57 | .58 |
| Italy | | | | | | | | | - | .37 | −.06 | .14 | −.08 | .10 | .02 | −.06 | .02 |
| Japan | | | | | | | | | | - | .16 | −.00 | −.22 | .49 | .12 | .09 | −.03 |
| Netherlands | | | | | | | | | | | - | .54 | .47 | .68 | .85 | .81 | .78 |
| Norway | | | | | | | | | | | | - | .64 | .37 | .42 | .28 | .66 |
| Spain | | | | | | | | | | | | | - | .33 | .41 | .32 | .79 |
| Sweden | | | | | | | | | | | | | | - | .49 | .54 | .54 |
| Switzerland | | | | | | | | | | | | | | | - | .83 | .74 |
| United Kingdom | | | | | | | | | | | | | | | | - | .69 |
| United States | | | | | | | | | | | | | | | | | - |

C. Other instrumental variables

| Variable | *rwd* | *xustb3* | *usjunk* | *xusdiv* |
|---|---|---|---|---|
| Excess world equity return | - | .17 | .19 | .22 |
| Excess U.S. 3 month bill | | - | .36 | −.16 |
| U.S. junk bond | | | - | −.31 |
| U.S. Excess dividend yield | | | | - |

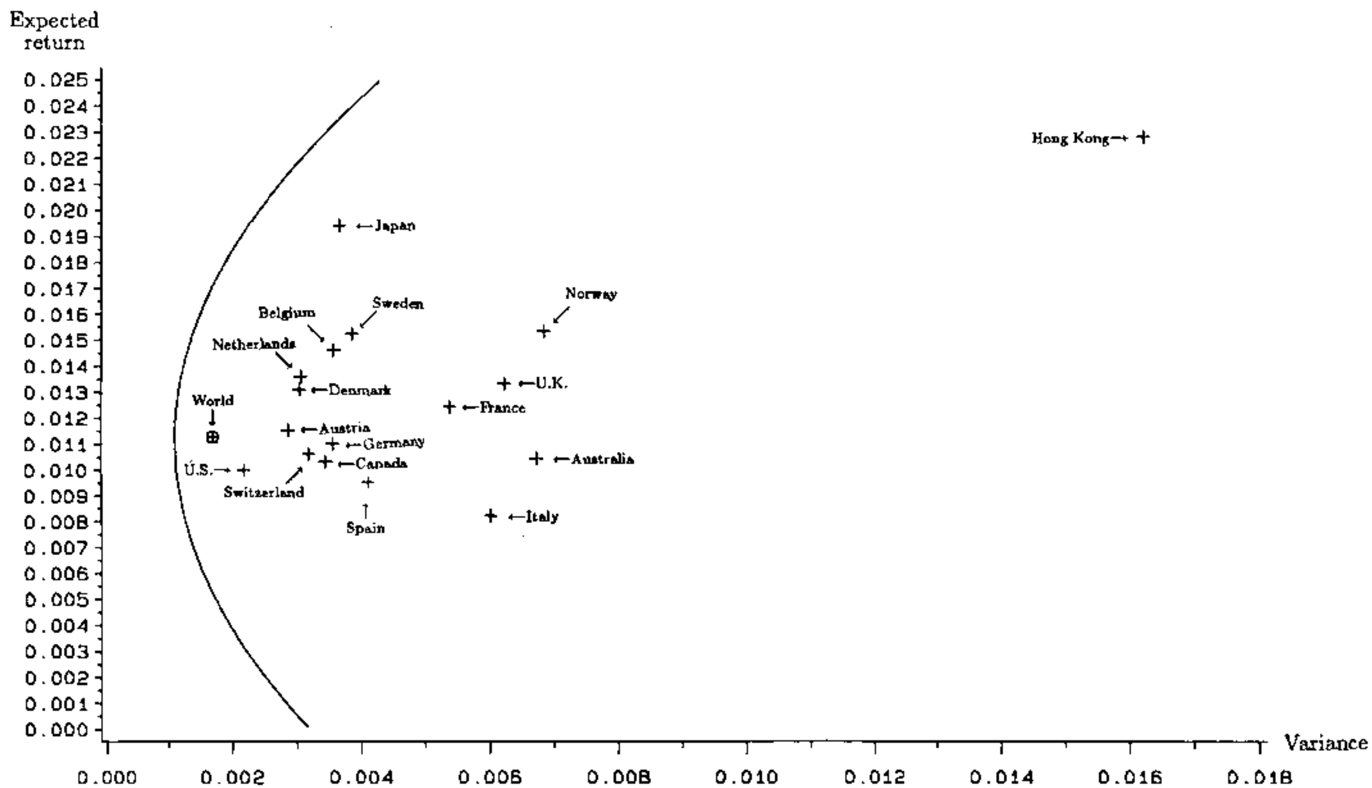


**Figure 1. The unconditional minimum variance frontier.** The minimum variance frontier is calculated from the unconditional means, variances, and covariances of 17 country returns. The returns are in U.S. dollars and are from Morgan Stanley Capital International. The data are from 1970:2–1989:5 (232 observations).

13 of the 17 countries. In Austria, there is a negative January effect. The excess return on the 3-month bill is two standard errors from zero in Australia, Belgium, Canada, the Netherlands, Switzerland, the United Kingdom, and the United States. The junk bond spread is more than two standard errors from zero in only two countries by 1.5 standard errors from zero in five additional countries. The dividend yield spread is more than two standard errors from zero in 12 of the 17 countries.[20]

The results in Table III can be interpreted as evidence against the hypothesis that the conditional mean returns in the countries are constant. Indeed, the examination of conditional asset pricing models is only well motivated if there is evidence of time-varying expected returns. An $F$-test shows that 15 of the 18 regressions are significant at the 10% level, 13 at the 5% level, and 10 at the 1% level.

*B. The Predictability of Country Stock Returns with the Local Instruments*

Table IV presents evidence on the predictability of international equity returns using both common and local information variables. The far left-hand column presents adjusted $R^2$ values using the common information variables. In some cases, they do not exactly match results reported in Table III because the sample may be slightly different.

Eight combinations of instrumental variables are used moving from the common information set (column 1) to a completely local information set (column 8). For comparison, the results of Cutler, Poterba and Summers (1989) and Campbell and Hamao (1989) are also reported. The local information variables include the country specific lagged return, the own dividend yield, the rate of change in the foreign exchange rate, the local short-term interest rate, and the long-term to short-term interest rate spread.

There are a number of interesting results in Table IV. First, the common information variables appear to capture the bulk of the predictable variation in the country returns. Comparing the common information regressions (column 1) to the completely local information regressions (column 8) only two countries show higher $R^2$ values using the local information. Explanatory power for Austria slightly increases from 5.6% to 5.9% when the local information variables are used. The $R^2$ for Norway increases from 1.9% to 4% with the local variables.

The other columns of Table IV show the effect of mixing the common information variables with the local information variables. Interestingly, the lagged U.S. dollar foreign exchange rate change has virtually no explanatory power. Only Norway and Sweden are affected by the change in the foreign exchange rate. Perhaps the most surprising result is the lack of importance of the local short-term interest rates and the long-term to short-term interest rate spread. Comparing columns 4 (common instruments with country-specific

[20]The regressions were re-estimated with own-currency returns rather than U.S. dollar returns. The explanatory power of these regressions was largely unaffected by using the own-currency returns. These results are available on request.

Table III

## Regressions of Country Returns on the Common Set of Instrumental Variables

The regressions are based on monthly data from 1970:2–1989:5 (232 observations). The country returns are calculated in U.S. dollars in excess of the holding period return on the Treasury bill that is closest to 30 days to maturity. The equity data are from Morgan Stanley Capital International. $t$-statistics in brackets are heteroskedasticity consistent. The model estimated is:

$$r_{j,t} = \delta_{j,0} + \delta_{j,1} rwd_{t-1} + \delta_{j,2} jan_t + \delta_{j,3} xustb3_{t-1} + \delta_{j,4} usjunk_{t-1} + \delta_{j,5} xusdiv_{t-1} + \varepsilon_{j,t} \quad (2)$$

The instrumental variables are: a constant, the excess return on the world index ($rwd$), a dummy variable for the month of January ($jan$), the return for holding a 90-day U.S. Treasury bill for 1 month less the return on a 30-day bill ($xustb3$), the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds ($usjunk$), and the dividend yield on the Standard and Poor's 500 stock index less the return on a 30-day bill ($xusdiv$).

| Portfolio | $\delta_0$ | $\delta_1$ | $\delta_2$ | $\delta_3$ | $\delta_4$ | $\delta_5$ | $\bar{R}^{2\,a}$ |
|---|---|---|---|---|---|---|---|
| Australia | 0.008 | 0.189 | 0.022 | 13.312 | −1.131 | 6.306 | 0.073 |
| | [0.499] | [1.267] | [1.013] | [3.160] | [−0.079] | [2.348] | |
| Austria | 0.037 | 0.139 | −0.034 | 2.352 | −22.002 | 3.074 | 0.058 |
| | [3.701] | [1.659] | [−2.873] | [1.044] | [−2.250] | [1.873] | |
| Belgium | 0.017 | −0.017 | 0.018 | 6.423 | 4.233 | 8.068 | 0.058 |
| | [1.593] | [−0.169] | [1.536] | [2.303] | [0.391] | [3.594] | |
| Canada | 0.004 | 0.035 | 0.017 | 12.598 | 2.911 | 5.871 | 0.107 |
| | [0.351] | [0.379] | [1.052] | [3.730] | [0.253] | [2.595] | |
| Denmark | 0.002 | −0.148 | 0.016 | 0.923 | 19.047 | 6.831 | 0.032 |
| | [0.180] | [−1.747] | [1.096] | [0.440] | [1.766] | [3.713] | |
| France | 0.014 | 0.071 | 0.018 | 2.289 | 3.403 | 6.283 | 0.013 |
| | [0.932] | [0.627] | [0.874] | [0.576] | [0.262] | [1.973] | |
| Germany | 0.005 | 0.098 | −0.006 | 2.490 | 11.088 | 5.579 | 0.021 |
| | [0.435] | [0.913] | [−0.380] | [0.920] | [1.034] | [2.613] | |

**Table III**—*Continued*

| Portfolio | $\delta_0$ | $\delta_1$ | $\delta_2$ | $\delta_3$ | $\delta_4$ | $\delta_5$ | $\bar{R}^{2}$ [a] |
|---|---|---|---|---|---|---|---|
| Hong Kong | 0.026 | 0.305 | 0.065 | 4.368 | −2.551 | 6.756 | 0.029 |
| | [0.858] | [1.702] | [2.529] | [0.702] | [−0.092] | [1.473] | |
| Italy | 0.006 | 0.210 | 0.027 | 1.379 | −4.783 | 1.057 | 0.005 |
| | [0.353] | [1.788] | [1.434] | [0.440] | [−0.318] | [0.320] | |
| Japan | 0.016 | 0.287 | 0.005 | −0.416 | 9.191 | 5.822 | 0.067 |
| | [1.281] | [2.749] | [0.435] | [−0.160] | [0.716] | [2.661] | |
| Netherlands | 0.001 | −0.011 | 0.026 | 6.205 | 15.940 | 7.154 | 0.076 |
| | [0.091] | [−0.108] | [1.756] | [2.351] | [1.429] | [3.380] | |
| Norway | 0.033 | 0.083 | 0.044 | 5.398 | −27.865 | 0.932 | 0.020 |
| | [2.016] | [0.491] | [2.339] | [1.089] | [−1.764] | [0.280] | |
| Spain | 0.019 | 0.172 | 0.017 | 3.757 | −18.714 | 0.329 | 0.009 |
| | [1.614] | [1.648] | [1.035] | [1.343] | [−1.736] | [0.144] | |
| Sweden | −0.013 | 0.115 | 0.023 | 1.441 | 24.151 | 3.040 | 0.032 |
| | [−0.974] | [0.974] | [1.540] | [0.486] | [1.805] | [1.160] | |
| Switzerland | 0.009 | −0.049 | 0.009 | 5.970 | 8.468 | 7.850 | 0.052 |
| | [0.853] | [−0.463] | [0.586] | [2.335] | [0.759] | [3.606] | |
| United Kingdom | −0.007 | −0.039 | 0.044 | 9.103 | 25.052 | 9.432 | 0.079 |
| | [−0.516] | [−0.259] | [1.495] | [2.837] | [1.609] | [3.109] | |
| United States | −0.014 | −0.092 | 0.020 | 8.289 | 22.980 | 6.175 | 0.125 |
| | [−1.550] | [−0.913] | [1.682] | [3.080] | [2.735] | [4.137] | |
| World | −0.005 | 0.032 | 0.018 | 6.602 | 16.848 | 6.015 | 0.133 |
| | [−0.664] | [0.455] | [1.751] | [3.495] | [2.256] | [4.448] | |

[a] Coefficient of determination adjusted for degrees of freedom.

Table IV

## International Evidence on the Predictability of Equity Returns Using Common and Country Specific Instrumental Variables

The $R^2$ values are adjusted for degrees of freedom. The country returns are calculated in U.S. dollars in excess of the holding period return on the U.S. Treasury bill that is closest to 30 days to maturity. The equity data are from Morgan Stanley Capital International. The regressions are estimated with eight different sets of conditioning information. The instrumental variables are: excess world return ($rwd_{t-1}$), a January dummy ($jan_t$), the difference in returns on U.S. 3- and 1-month Treasury bills ($xustb3_{t-1}$), the spread between Moody's Baa and Aaa rated U.S. bonds ($usjunk_{t-1}$), the U.S. dividend yield in excess of the 1-month Treasury bill ($xusdiv_{t-1}$), the equity returns in each country ($r_{j,t-1}$), the dividend yields for each country ($div_{j,t-1}$), the return on the U.S. exchange rate for each country ($fx_{j,t-1}$), the level of short term interest rates in each country ($i_{j,t-1}$), and the difference between long term government bond yields and short term yields in each country ($term_{j,t-1}$).

| Instruments → Portfolio and (number of observations) | $rwd_{t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $xusdiv_{t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $xusdiv_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $div_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $xusdiv_{j,t-1}$, $div_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $div_{j,t-1}$, $fx_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $div_{j,t-1}$, $i_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $div_{j,t-1}$, $i_{j,t-1}$, $term_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $div_{j,t-1}$, $i_{j,t-1}$, $term_{j,t-1}$ | Cutler, Poterba, and Summers [1989] | Campbell and Hamao [1989] |
|---|---|---|---|---|---|---|---|---|---|---|
| Australia 70:2–89:1 (228) | .078 | .070 | .073 | .087 | .080 | .070 | .022 | .011 | .002 | – |
| Austria 71:12–89:1 (216) | .056 | .062 | .049 | .059 | .050 | .054 | .057 | .059 | .017 | – |
| Belgium[a] 70:2–89:5 (232) | .058 | .062 | .064 | .080 | .051 | – | – | – | –.004 | – |
| Canada 70:2–89:5 (232) | .107 | .111 | .082 | .107 | .084 | .091 | .028 | .000 | .001 | – |
| Denmark 72:2–89:1 (204) | .035 | .018 | –.016 | .013 | –.004 | .008 | .013 | .008 | –.004 | – |
| France[b] 70:2–89:5 (229) | .013 | .020 | .001 | .019 | –.004 | .005 | .007 | .010 | –.004 | – |

**Table IV**—*Continued*

| Instruments → <br> Portfolio and (number of observations) | $rwd_{t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $xusdiv_{t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $xusdiv_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $div_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $xusdiv_{j,t-1}$, $div_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $div_{j,t-1}$ $fx_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $xustb3_{t-1}$, $div_{j,t-1}$, $i_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $usjunk_{t-1}$, $div_{j,t-1}$, $i_{j,t-1}$, $term_{j,t-1}$ | $r_{j,t-1}$, $jan_t$, $div_{j,t-1}$, $i_{j,t-1}$, $term_{j,t-1}$ | Cutler, Poterba, and Summers [1989] | Campbell and Hamao [1989] |
|---|---|---|---|---|---|---|---|---|---|---|
| Germany 70:2–89:5 (232) | .021 | .017 | −.007 | .013 | −.012 | .003 | .000 | −.002 | −.002 | - |
| Hong Kong[a] 70:2–89:5 (232) | .029 | .021 | .008 | .017 | .007 | - | - | - | .040 | - |
| Italy 71:2–89:5 (220) | .001 | .021 | .021 | .017 | .016 | .017 | .018 | .023 | −.002 | - |
| Japan 70:2–89:5 (232) | .067 | .045 | .006 | .043 | .003 | .025 | .029 | .024 | .005 | .065 |
| Netherlands 70:2–89:5 (232) | .076 | .076 | .029 | .076 | .025 | .068 | .087 | .076 | .001 | - |
| Norway[c] 71:9–89:5 (213) | .019 | .052 | .056 | .051 | .062 | .051 | .045 | .040 | −.002 | - |
| Spain[d] 74:1–89:5 (184) | .004 | .013 | .014 | .010 | .010 | .018 | - | - | −.003 | - |
| Sweden 70:2–87:1 (204) | .022 | .022 | .013 | .017 | .034 | .012 | .020 | .007 | −.004 | - |
| Switzerland 75:9–89:5 (175) | .046 | .044 | −.009 | .041 | −.015 | .000 | .000 | .010 | .002 | - |
| United Kingdom 72:1–89:5 (208) | .078 | .079 | .082 | .115 | .078 | .077 | .071 | .076 | .006 | - |
| United States 70:2–89:5 (232) | .125 | .124 | .077 | .127 | .076 | .127 | .076 | .052 | .010 | .100 |

[a] Both short term and long term interest rates are not available.
[b] Excludes 1986:3–1986:6 when the interest rate data are not available.
[c] Excludes 1980:8–1980:9 when the interest rate data are not available.
[d] Long term interest rates are not available.

dividend yields and lagged returns) to columns 6–8, the explanatory power is marginally increased in only three countries: Spain (+0.8%), the Netherlands (+1.1%), and Italy (+0.6%). The local interest rate variables decrease the explanatory power of the regressions for the other countries.

The two local information variables that are the most important are the lagged own-country returns and the local dividend yields. The inclusion of the lagged own-country returns increases the explanatory power of the regressions (columns 1 and 2) in 8 of 17 countries and has a neutral effect on two other countries. The inclusion of the local dividend yields (in addition to the U.S. dividend yield) increases the explanatory power of the regressions (columns 1 and 4) in 9 of 17 countries and has a neutral effect on two countries.

While the lagged own-country return and the local dividend yield increase the explanatory power of some of the regressions, the overall improvement is small. In the countries that experience increased adjusted $R^2$ values, the average increment is only 1.7%. Most of the explanatory power is driven by the common variables. Cutler, Poterba and Summers (1990) state that "it seems unlikely that similar processes generate required returns . . . " in the international markets. The results in Table IV provide evidence against this claim. Expected returns in the individual countries appear to be generated by common world factors.

### *C. Conditional Asset Pricing with Time-Varying Moments*

Table V provides tests of the general model that allows for time-varying expected returns, covariances, and variances. Tests of the asset pricing restrictions are provided for individual countries as well as multiple country systems.[21]

The $\chi^2$ statistic provides a test of the model's restrictions. This statistic summarizes the departures from the null hypothesis—that the world market portfolio is conditionally mean variance efficient. Beside the $\chi^2$ is the related but more intuitive $R^2$ statistic which is the adjusted coefficient of determination from a regression of the model errors on the common information variables. If the model fits well, the errors should be unrelated to the information, and the $\chi^2$ and the $R^2$ should be small.

The asset pricing model in equation (6) can be tested using individual countries as well as multiple countries.[22] The test at the individual country level may not be powerful because the cross-asset restriction (identical conditionally expected world market return divided by conditional variance of the world market for each country) is not imposed. That is, if the model is *not* rejected at the individual country level, some caution should be exercised in

[21]For ease of exposition, beginning in Table V the countries are ordered by unconditional mean return.

[22]It is not computationally feasible to test equation (6) with all 17 countries. The dimensionality of the weighting matrix would be 210 which is much larger than any previously estimated GMM system. Furthermore, Hansen and Singleton (1982) warn that the quality of the consistent estimator of the weighting matrix may deteriorate with high dimensions.

interpreting the results—because not all of the CAPM's restrictions have been imposed. However, a *rejection* at the individual country level may provide valuable information about the model's failings. Using the common instrumental variables, Table V indicates that the model's restrictions are rejected at the 5% level for three of the 17 countries: Japan, Norway, and Austria. The model is rejected at the 10% level for the United States.

In the multiple country test using the returns of the Group of 7 countries, the model's restrictions are not rejected at standard levels of significance. The lack of rejection in this multivariate test reinforces the importance of testing at the individual portfolio level. The multivariate test is not powerful enough to reject the hypothesis of conditional mean-variance efficiency. However, the evidence at the individual country level (in particular for Japan) suggests that the world portfolio is not conditionally mean-variance efficient.

The last two columns provide test statistics for the model estimated using some local information variables. In the single country estimation, the inference is generally robust to the choice of the information set. Using the first local information set, the same four countries are rejected at the 10% level. Using the second local information set, three of those four countries are rejected at the 10% level. With this information set, Japan is no longer rejected. For Japan, the local information variables provide a less powerful test. This is perhaps not surprising given the results in Table IV which show that the maximum explanatory power for Japanese returns derives from the common information set.

Similar patterns arise when the local information variables are used in the multivariate test. Using the first local information set, the probability value of the test statistic is 13.7% which is somewhat lower than the 21.9% reported using the common instruments. With the second local information set, the *p*-value is 12.7%. Consistent with the previous results, this local information set fails to provide evidence to reject the null hypothesis at standard levels of significance.

Some additional information is provided in Table V. The average pricing errors and the average absolute pricing errors based upon estimation with the common instrumental variables are provided in the fourth and fifth columns. For Japan, the pricing error is positive, indicating that the actual return is on average higher than the expected return given the level of risk. A large positive pricing error is also found with the Hong Kong portfolio. The Austrian pricing error is negative, indicating that the average return is less than what is expected given the country risk. Interestingly, over this time period, both the mean error and the absolute pricing error for the United States are small.

The average conditional covariance is provided in the second column.[23] It is clear that the ordering of the average conditional covariances is not the

[23]This is not the unconditional covariance. It is the average value of the product of the innovations in the conditional mean of the country return and the world market return. This covariance is conditional on the common information set.

Table V

## Estimates of a Conditional CAPM with Time Varying Expected Returns, Conditional Covariances, and Conditional Variances

Results based on monthly data from 1970:2–1989:5 (232 observations). The country returns **r** are calculated in U.S. dollars in excess of the holding period return on the Treasury bill that is closest to 30 days to maturity. The equity data are from Morgan Stanley Capital International. The following system of equations is estimated with the generalized method of moments:

$$\varepsilon_t = (\mathbf{u}_t \quad u_{mt} \quad \mathbf{h}_t) = \begin{pmatrix} [\mathbf{r}_t - \mathbf{Z}_{t-1}\boldsymbol{\delta}]' \\ [r_{mt} - \mathbf{Z}_{t-1}\boldsymbol{\delta}_m]' \\ [u_{mt}^2 \mathbf{Z}_{t-1}\boldsymbol{\delta} - u_{mt}\mathbf{u}_t \mathbf{Z}_{t-1}\boldsymbol{\delta}_m]' \end{pmatrix}', \quad (6)$$

where $r_m$ is the excess return on the world portfolio, $\boldsymbol{\delta}$ represents the coefficients associated with the instrumental variables, **u** is the forecast error for the country returns, $u_m$ is the forecast error for the world market return, and **h** represents the deviation of the country return from the model's expected return. There are three sets of instrumental variables **Z** that are used in the estimation. The common set of predetermined instrumental variables are: a constant, the excess return on the world index, a dummy variable for the month of January, the 1-month return for holding a 90-day U.S. Treasury bill less the return on a 30-day bill, the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds, and the dividend yield on the Standard and Poor's 500 stock index less the return on a 30-day bill. Local instrument set A is the common instrument set augmented with the country-specific dividend yield. Local instrument set B includes the country-specific dividend yield and the country-specific excess return in place of the world excess return.

| Portfolio | Average Return | Average Conditional Covariance[a] | Average Error[b] | Average Absolute Error[c] | $\bar{\bar{R}}^{2\,d}$ | Common Instruments $\chi^{2\,e}$ [P-value] | Local Instruments A: $\chi^2$ [P-value] | Local Instruments B: $\chi^2$ [P-value] |
|---|---|---|---|---|---|---|---|---|
| Hong Kong | 0.0168 | 1.8000 | 0.0037 | 0.0202 | .022 | 8.14 [.228] | 8.25 [.311] | 7.91 [.340] |
| Japan | 0.0134 | 1.4007 | 0.0013 | 0.0104 | −.012 | 14.96 [.021] | 13.95 [.052] | 10.47 [.163] |
| Sweden | 0.0094 | 1.0978 | 0.0014 | 0.0100 | −.008 | 8.49 [.204] | 8.60 [.282] | 9.06 [.248] |

**Table V**—*Continued*

| Portfolio | Average Return | Average Conditional Covariance[a] | Average Error[b] | Average Absolute Error[c] | $\tilde{R}^{2}$[d] | Common Instruments $\chi^{2}$[e] [P-value] | Local Instruments A: $\chi^2$ [P-value] | Local Instruments B: $\chi^2$ [P-value] |
|---|---|---|---|---|---|---|---|---|
| Norway | 0.0093 | 1.6264 | −0.0009 | 0.0162 | .067 | 13.15 [.041] | 12.49 [.085] | 15.96 [.025] |
| Belgium | 0.0087 | 1.3071 | 0.0004 | 0.0110 | −.018 | 9.47 [.149] | 10.74 [.150] | 9.80 [.200] |
| Netherlands | 0.0077 | 1.4351 | 0.0005 | 0.0081 | −.020 | 6.30 [.391] | 8.64 [.280] | 8.43 [.296] |
| United Kingdom | 0.0074 | 1.8155 | −0.0007 | 0.0157 | −.022 | 1.10 [.981] | 4.97 [.663] | 3.73 [.810] |
| Denmark | 0.0072 | 0.9425 | 0.0006 | 0.0092 | −.018 | 9.28 [.159] | 9.93 [.193] | 7.40 [.389] |
| France | 0.0065 | 1.6901 | −0.0011 | 0.0122 | −.002 | 10.31 [.112] | 11.80 [.107] | 13.58 [.059] |
| Austria | 0.0055 | 0.6339 | −0.0033 | 0.0098 | .163 | 20.50 [.002] | 21.47 [.003] | 22.57 [.002] |
| Germany | 0.0050 | 1.1866 | −0.0004 | 0.0112 | −.020 | 3.47 [.748] | 3.45 [.841] | 4.29 [.746] |
| Switzerland | 0.0046 | 1.3832 | −0.0012 | 0.0093 | −.004 | 10.25 [.115] | 10.92 [.142] | 9.91 [.194] |
| Australia | 0.0044 | 1.5857 | 0.0014 | 0.0161 | −.008 | 5.78 [.449] | 6.87 [.442] | 5.11 [.647] |
| Canada | 0.0044 | 1.5231 | −0.0004 | 0.0097 | −.012 | 3.14 [.791] | 3.13 [.873] | 2.89 [.895] |
| United States | 0.0037 | 1.4250 | −0.0000 | 0.0043 | −.016 | 10.75 [.096] | 12.83 [.076] | 11.19 [.131] |
| Spain | 0.0036 | 1.0274 | −0.0004 | 0.0138 | .005 | 10.41 [.109] | 11.28 [.127] | 8.28 [.308] |

Table V—*Continued*

| Portfolio | Average Return | Average Conditional Covariance[a] | Average Error[b] | Average Absolute Error[c] | $\bar{R}^{2}$[d] | Common Instruments $\chi^{2}$[e] [P-value] | Local Instruments A: $\chi^2$ [P-value] | Local Instruments B: $\chi^2$ [P-value] |
|---|---|---|---|---|---|---|---|---|
| Italy | 0.0022 | 1.2197 | −0.0003 | 0.0125 | .016 | 9.82 [.132] | 9.89 [.198] | 11.09 [.135] |
| G-7[f] | | | | | | 48.77 [.219] | 59.91 [.137] | 60.43 [.127] |

[a]The average value of $u_i \times u_m$ multiplied by 1000 for country $i$ based on single country estimation with the common instrument set.

[b]The average value of $e_i$ for country $i$ based on single country estimation with the common instrument set divided by the average conditional variance of the world market return.

[c]The average absolute value of $e_i$ for country $i$ based on single country estimation with the common instrument set divided by the average conditional variance of the world market return.

[d]The adjusted coefficient of determination from a regression of the model errors ($e_{it}$) on the common instrumental variables.

[e]The minimized value of the GMM criterion function. *P*-value is the probability that a $\chi^2$ variate exceeds the sample value of the statistic. For single country systems with the common instrument set, there are 12 parameters and 18 orthogonality conditions leaving 6 overidentifying conditions. For the single country systems with local instrument sets, there are 14 parameters and 21 orthogonality conditions leaving 7 overidentifying restrictions. In the multiple equation system with the common instrument set, there are 48 parameters and 90 orthogonality conditions; this implies that there are 42 overidentifying restrictions to be tested. In the multiple equation system with the local instrument sets, there are 56 parameters and 105 orthogonality conditions; this implies that there are 49 overidentifying restrictions to be tested. The degrees of freedom in the test statistic correspond to the number of overidentifying restrictions.

[f]Japan, United Kingdom, France, Germany, Canada, United States, and Italy. These countries comprise 90% of the MSCI world index. In the tests with the local information variables, the local variables are used to get the forecasted country returns (equations 1 through 7). The common variables are used for the world expected return (equation 8). Orthogonality conditions are formed with the local information variables.

same as the ordering of the average returns. Nevertheless, it is interesting to note that Hong Kong has one of the highest average conditional covariances as well as the highest average return. However, the conditional asset pricing model does not restrict the 'average' conditional covariance to be positively related to the 'average' return. The average conditional covariance is provided only as bridge to unconditional asset pricing.

The results for the general formulation provides evidence against the asset pricing model's restrictions. Consistent with the results of tests of unconditional mean-variance efficiency, when many countries are examined there is little evidence against the model's restrictions. However, a country by country examination detects some significant departures from the null hypothesis. When Japan is examined, the restrictions are strongly rejected.

If some of the moments are constant, then it may be possible to construct more powerful tests. Two other formulations are examined: one that assumes constant conditional betas and another that assumes a constant world price of covariance risk. The constant conditional beta formulation is closely linked to unconditional formulations of the Sharpe-Lintner model. The constant world price of covariance risk is often interpreted as a measure of aggregate relative risk aversion.

### *D. Conditional Asset Pricing with Constant Conditional Betas*

Table VI presents tests of the conditional version of the original Sharpe-Lintner formulation which implies that expected asset returns are proportional to the expected (mean-variance efficient) world market portfolio returns. Beta is the coefficient of proportionality. As with the previous table, single country as well as multiple country tests are presented. In most cases with the common set of instruments, the proportionality coefficients are more than two standard errors from zero. The highest beta is found with the Hong Kong portfolio.[24] The two smallest betas are estimated for Spain and Italy. These two countries have the smallest average returns.

The United States has a beta of 0.97 while Japan has a beta of 1.42. However, the difference in the betas does not explain the difference in the average excess returns of 0.37% per month for the United States and 1.34% per month for Japan. As with the general model presented in the previous table, the model's restrictions are rejected when the Japanese returns are examined. The results in Table VI indicate stronger evidence against the restrictions. The model is also rejected at the 5% level for Denmark as well as Austria. There is evidence against the model at the 10% level for Norway. The probability value for the United States is 0.103 which is slightly higher than the general model. Consistent with the results in Table V, there is little evidence against the restrictions in the multiple country system conditioning on the common information set.

[24] These betas are estimated under the null hypothesis that the model is true, i.e., no intercepts are included. An alternative formulation would specify intercepts in equation (7) and test whether they are significantly different from zero.

Table VI

## Estimates of a Conditional CAPM with Time Varying Expected Returns and Constant Conditional Betas

Results are based on monthly data from 1970:2–1989:5 (232 observations). The country returns **r** are calculated in U.S. dollars in excess of the holding period return on the Treasury bill that is closest to 30 days to maturity. The equity data are from Morgan Stanley Capital International. Standard errors in parentheses are heteroskedasticity consistent. Generalized method of moments is used to estimate:

$$\mathbf{k}_t = \mathbf{r}_t - r_{mt}\boldsymbol{\beta} \qquad (7)$$

where $r_m$ is the excess return on the world market portfolio, $\boldsymbol{\beta}$ is the proportionality coefficient that relates the expected world excess return to the expected country return, and **k** represents the deviations from the country returns and the model's expected returns. There are three sets of instrumental variables **Z** that are used in the estimation. The common set of predetermined instrumental variables are: a constant, the excess return on the world index, a dummy variable for the month of January, the 1-month return for holding a 90-day U.S. Treasury bill less the return on a 30-day bill, the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds, and the dividend yield on the Standard and Poor's 500 stock index less the return on a 30-day bill. Local instrument set A is the common instrument set augmented with the country-specific dividend yield. Local instrument set B includes the country-specific dividend yield and the country-specific excess return in place of the world excess return.

| Portfolio | $\beta_j$ | Average Return | Average Error[a] | Average Absolute Error[b] | $\bar{R}^2$[c] | Common Instruments $\chi^2$[d] [P-value] | Local Instruments A: $\chi^2$ [P-value] | Local Instruments B: $\chi^2$ [P-value] |
|---|---|---|---|---|---|---|---|---|
| Hong Kong | 2.0802 | .0168 | .0058 | .0841 | .011 | 4.73 | 4.95 | 4.22 |
| | (0.4450) | | | | | [.450] | [.551] | [.647] |
| Japan | 1.4178 | .0134 | .0059 | .0386 | .096 | 18.34 | 18.27 | 17.50 |
| | (0.2195) | | | | | [.003] | [.006] | [.008] |
| Sweden | 0.7281 | .0094 | .0055 | .0425 | −.003 | 5.01 | 5.24 | 4.46 |
| | (0.2086) | | | | | [.414] | [.514] | [.614] |

**Table VI—*Continued***

| Portfolio | $\beta_j$ | Average Return | Average Error[a] | Average Absolute Error[b] | $\bar{R}^{2c}$ | Common Instruments $\chi^{2d}$ [P-value] | Local Instruments A: $\chi^2$ [P-value] | Local Instruments B: $\chi^2$ [P-value] |
|---|---|---|---|---|---|---|---|---|
| Norway | 0.5742 (0.2859) | .0093 | .0062 | .0564 | .021 | 11.02 [.051] | 11.22 [.082] | 16.61 [.011] |
| Belgium | 1.0573 (0.1974) | .0087 | .0030 | .0357 | .003 | 6.06 [.300] | 7.76 [.256] | 7.23 [.300] |
| Netherlands | 1.0061 (0.1349) | .0077 | .0023 | .0291 | −.013 | 2.64 [.755] | 4.74 [.578] | 4.73 [.579] |
| United Kingdom | 1.3840 (0.2531) | .0074 | −.0000 | .0432 | −.011 | 1.11 [.953] | 2.93 [.817] | 3.16 [.788] |
| Denmark | 0.4908 (0.1411) | .0072 | .0046 | .0361 | .013 | 12.56 [.028] | 12.74 [.047] | 7.79 [.254] |
| France | 0.6937 (0.2426) | .0065 | .0028 | .0468 | −.006 | 4.37 [.497] | 5.60 [.470] | 6.49 [.340] |
| Austria | 0.1851 (0.1794) | .0055 | .0046 | .0353 | .065 | 19.69 [.001] | 19.99 [.003] | 19.40 [.004] |
| Germany | 0.7035 (0.2159) | .0050 | .0013 | .0388 | −.001 | 4.78 [.443] | 4.78 [.573] | 4.93 [.579] |
| Switzerland | 0.8293 (0.1624) | .0046 | .0002 | .0332 | .003 | 5.86 [.320] | 8.39 [.211] | 7.89 [.246] |
| Australia | 1.3949 (0.2622) | .0044 | −.0030 | .0532 | −.001 | 5.78 [.328] | 7.48 [.279] | 6.42 [.378] |
| Canada | 1.0466 (0.1539) | .0044 | −.0012 | .0300 | .020 | 6.76 [.239] | 6.98 [.323] | 7.41 [.284] |
| United States | 0.9656 (0.0939) | .0037 | −.0014 | .0176 | .038 | 9.17 [.103] | 9.53 [.146] | 5.89 [.439] |

**Table VI**—*Continued*

| Portfolio | $\beta_j$ | Average Return | Average Error[a] | Average Absolute Error[b] | $\bar{R}^{2c}$ | Common Instruments $\chi^{2d}$ [P-value] | Local Instruments A: $\chi^2$ [P-value] | Local Instruments B: $\chi^2$ [P-value] |
|---|---|---|---|---|---|---|---|---|
| Spain | 0.3003 (0.2483) | .0036 | .0020 | .0460 | .004 | 8.66 [.124] | 8.90 [.179] | 8.90 [.179] |
| Italy | 0.4499 (0.2414) | .0022 | −.0002 | .0549 | −.003 | 5.73 [.334] | 5.73 [.454] | 5.26 [.511] |
| G-7[e] | | | | | | 39.72 [.268] | 45.96 [.312] | 48.42 [.229] |

[a]The average value of $k_i$ for country $i$ based on single country estimation with the common instrument set.

[b]The average absolute value of $k_i$ for country $i$ based on single country estimation with the common instrument set.

[c]The minimized value of the GMM criterion function. P-value is the probability that a $\chi^2$ variate exceeds the sample value of the statistic. For single country systems with the common instrument set, there is one parameter and 6 orthogonality conditions leaving 5 overidentifying conditions. For single country systems with the local instrument sets, there is one parameter and 7 orthogonality conditions leaving 6 overidentifying conditions. In the multiple equation system with the common instrument set, there are 7 parameters and 42 orthogonality conditions; this implies that there are 35 overidentifying restrictions to be tested. In the multiple equation system with the local instrument sets, there are 7 parameters and 49 orthogonality conditions; this implies that there are 42 overidentifying restrictions to be tested. The degrees of freedom in the test statistic correspond to the number of overidentifying restrictions.

[d]The adjusted coefficient of determination from a regression of the model errors ($k_{it}$) on the common set of instrumental variables.

[e]Japan, United Kingdom, France, Germany, Canada, United States, and Italy. These countries comprise 90% of the MSCI world index. In the estimation with local instruments, orthogonality conditions are formed using the local instrumental variables.

Re-estimation of the models using the local information sets has virtually no impact on the inference. With the first set of local instruments, the restrictions are rejected at the 5% level for Japan, Denmark, and Austria and at the 10% level for Norway. These results are identical to those using the common information variables. With the second set of local instruments, the model's restrictions are rejected at the 5% level for Japan, Norway, and Austria. In contrast to the results in Table V, the inclusion of the local instruments does not increase the evidence against the model's restrictions when the multiple country system is estimated.

Table VI also reports pricing errors based upon the estimation with the common set of instrumental variables. The average error for Japan is 0.59%. This implies that the model is delivering an average expected return of 0.75% per month, and the average realized return is 1.34%. For the United States, the model is predicting a 0.51% return while only 0.34% is realized on average. The model appears to fit quite well for the United Kingdom with a less than 0.01% pricing error.

### *E. Conditional Asset Pricing with a Constant World Price of Covariance Risk*

Table VII presents test of the formulation that allows for time-varying conditional covariances. The constant in the estimation is the expected compensation for world market volatility—or the world price of covariance risk. However, this parameter is not restricted to be the same across countries in the single country estimation.

The results in Table VII reveal more evidence against this formulation than the previous two tables. Using the common information set, the model's restrictions are rejected at the 5% level for Hong Kong, Japan, Sweden, Belgium, the Netherlands, the United Kingdom, Denmark, Austria, Switzerland, Australia, Canada, and United States.[25] The estimates with the local information sets provide similar evidence against the model's restrictions. The average pricing error for Japan is of the same magnitude as the constant conditional beta model. The pricing error for the United States is three times the size of the pricing error in the constant conditional beta formulation.

There is wide variation in the magnitude of the reward to risk ratio. For example, the expected compensation for world market volatility in the United States is 5.4. The same measure in Japan in 13.1. In a financially integrated global market with time-invariant reward to risk, this ratio should be the same across all countries. If the financial markets are not perfectly integrated or if the asset pricing model is misspecified, then there is no reason that the reward to risk ratios should be the same.

A formal examination of the differences in the reward to risk ratio across different countries is presented in the last two lines of Table VII. First, a seven country system is estimated with the reward to risk ratio constrained to be constant across all countries. This measure can be interpreted as the

[25]The results for the United States are consistent with the formulation tested in Campbell (1987) and Harvey (1989, 1990).

Table VII

## Estimates of a Conditional CAPM with Time Varying Expected Returns and a Constant Price of Covariance Risk

The results are based on data from 1970:2–1989:5 (232 observations). The country returns **r** are calculated in U.S. dollars in excess of the holding period return on the Treasury bill that is closest to 30 days to maturity. The equity data are from Morgan Stanley Capital International. Standard errors in parentheses are heteroskedasticity consistent. Generalized method of moments is used to estimate the system:

$$\eta_t = (u_{mt} \quad \mathbf{e}_t) = \begin{pmatrix} [r_{mt} - \mathbf{Z}_{t-1}\boldsymbol{\delta}_m]' \\ [\mathbf{r}_t - \lambda(u_{mt}\mathbf{r}_t)]' \end{pmatrix}', \tag{10}$$

where $r_m$ is the excess return on the world portfolio, $\boldsymbol{\delta}_m$ values are the coefficients associated with the instrumental variables for estimating the conditional mean of the world return, $u_m$ is the forecast error in the conditional mean of the world return, and $\lambda$ is the world price of covariance risk. In the estimation, the world price of risk (expected world excess return divided by the variance of world excess returns) is held constant through time. However, when (10) is estimated country by country, the cross-country restriction that the world price of risk is the same *in each country* is not imposed. The final line of the table tests whether the world price of risk is statistically different across the group of seven countries. There are three sets of instrumental variables **Z** that are used in the estimation. The common set of predetermined instrumental variables are: a constant, the excess return on the world index, a dummy variable for the month of January, the 1-month return for holding a 90-day U.S. Treasury bill less the return on a 30-day bill, the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds, and the dividend yield on the Standard and Poor's 500 stock index less the return on a 30-day bill. Local instrument set A is the common instrument set augmented with the country-specific dividend yield. Local instrument set B includes the country-specific dividend yield and the country-specific excess return in place of the world excess return.

| Portfolio | $\lambda_j$ | Average Conditional Covariance[a] | Average Return | Aveage Error[b] | Average Absolute Error[c] | $\bar{R}^2$[d] | Common Instruments $\chi^2$[e] [*P*-value] | Local Instruments A: $\chi^2$ [*P*-value] | Local Instruments B: $\chi^2$ [*P*-value] |
|---|---|---|---|---|---|---|---|---|---|
| Hong Kong | 11.6735 | 1.9589 | .0168 | −.0060 | .0945 | .029 | 18.05 | 18.24 | 10.62 |
| | (4.4172) | | | | | | [.003] | [.006] | [.101] |
| Japan | 13.0803 | 1.4792 | .0134 | −.0059 | .0464 | .045 | 16.18 | 15.66 | 11.66 |
| | (3.0075) | | | | | | [.006] | [.016] | [.070] |

**Table VII—Continued**

| Portfolio | $\lambda_j$ | Average Conditional Covariance[a] | Average Return | Aveage Error[b] | Average Absolute Error[c] | $\bar{R}^{2d}$ | Common Instruments $\chi^{2e}$ [P-value] | Local Instruments A: $\chi^2$ [P-value] | Local Instruments B: $\chi^2$ [P-value] |
|---|---|---|---|---|---|---|---|---|---|
| Sweden | 8.3814 (4.0166) | 1.1494 | .0094 | −.0003 | .0507 | .037 | 11.40 [.044] | 11.54 [.073] | 11.28 [.080] |
| Norway | 4.3466 (3.2699) | 1.6953 | .0093 | .0019 | .0653 | .022 | 10.45 [.064] | 11.04 [.087] | 15.71 [.015] |
| Belgium | 9.7415 (3.1100) | 1.3983 | .0087 | −.0050 | .0444 | .067 | 17.76 [.003] | 17.93 [.006] | 16.50 [.011] |
| Netherlands | 9.3435 (2.6487) | 1.5527 | .0077 | −.0068 | .0429 | .070 | 15.26 [.009] | 16.54 [.011] | 14.55 [.024] |
| United Kingdom | 8.8913 (2.1730) | 1.9757 | .0074 | −.0102 | .0566 | .065 | 13.22 [.021] | 13.52 [.035] | 13.87 [.031] |
| Denmark | 9.6631 (3.7602) | .9888 | .0072 | −.0024 | .0419 | .029 | 13.11 [.022] | 13.13 [.041] | 12.70 [.048] |
| France | 3.6508 (2.7151) | 1.7456 | .0065 | .0001 | .0563 | .017 | 8.82 [.116] | 10.40 [.109] | 9.96 [.126] |
| Austria | −0.2664 (3.9388) | .7291 | .0055 | .0057 | .0367 | .057 | 19.41 [.002] | 19.93 [.003] | 17.72 [.007] |
| Germany | 1.4478 (2.9738) | 1.2424 | .0050 | .0032 | .0462 | .025 | 10.91 [.053] | 11.24 [.081] | 10.61 [.101] |
| Switzerland | 3.6684 (2.5421) | 1.4785 | .0046 | −.0008 | .0433 | .059 | 14.66 [.012] | 15.13 [.019] | 14.85 [.021] |
| Australia | −3.7285 (2.6900) | 1.7379 | .0044 | .0109 | .0585 | .063 | 20.62 [.001] | 22.06 [.001] | 20.47 [.002] |
| Canada | 1.4885 (2.3548) | 1.6587 | .0044 | .0019 | .0438 | .104 | 18.41 [.003] | 14.68 [.023] | 13.68 [.033] |
| United States | 5.3716 (2.0917) | 1.6159 | .0037 | −.0049 | .0362 | .086 | 18.20 [.003] | 18.93 [.004] | 18.97 [.004] |

**Table VII**—*Continued*

| Portfolio | $\lambda_j$ | Average Conditional Covariance[a] | Average Return | Aveage Error[b] | Average Absolute Error[c] | $\bar{R}^{2}$[d] | Common Instruments $\chi^{2}$[e] [P-value] | Local Instruments A: $\chi^2$ [P-value] | Local Instruments B: $\chi^2$ [P-value] |
|---|---|---|---|---|---|---|---|---|---|
| Spain | 0.9848 | 1.0625 | .0036 | .0025 | .0485 | .009 | 8.37 | 8.32 | 9.85 |
| | (3.8771) | | | | | | [.137] | [.216] | [.131] |
| Italy | 0.6396 | 1.2604 | .0022 | .0014 | .0593 | .006 | 7.23 | 7.23 | 8.37 |
| | (3.8879) | | | | | | [.204] | [.300] | [.212] |
| G-7[f] | 11.4716 | | | | | | 48.42 | 50.32 | 48.47 |
| | (1.8252) | | | | | | [.198] | [.382] | [.454] |
| G-7[g] | $\lambda_j = \lambda$ | | | | | | 23.76 | 32.72 | 22.74 |
| | $j = 1, \ldots, 7$ | | | | | | [.001] | [< .001] | [.001] |

[a]The average value of $u_i \times u_m$ multiplied by 1000 for country $i$ based on single country estimation.

[b]The average value of $e_i$ for country $i$ based on single country estimation.

[c]The average absolute value of $e_i$ for country $i$ based on single country estimation.

[d]The adjusted coefficient of determination from a regression of the model errors ($e_{it}$) on the common instrumental variables.

[e]The minimized value of the GMM criterion function. *P*-value is the probability that a $\chi^2$ variate exceeds the sample value of the statistic. For single country systems with the common instrument set, there are 7 parameters and 12 orthogonality conditions leaving 5 overidentifying conditions. For single country systems with the local instrument sets, there are 6 parameters and 14 orthogonality conditions leaving 6 overidentifying conditions. In the multiple equation system with common instruments, there are 7 parameters and 48 orthogonality conditions; this implies that there are 41 overidentifying restrictions to be tested. In the multiple equation system with local instruments, there are 8 parameters and 56 orthogonality conditions; this implies that there are 48 overidentifying restrictions to be tested. The degrees of freedom in the test statistic correspond to the number of overidentifying restrictions.

[f]Japan, United Kingdom, France, Germany, Canada, United States, and Italy. These countries comprise 90% of the MSCI world index. In the estimation with the local instruments, the common variables are used for the conditional mean of the world return (equation 1). The orthogonality conditions are formed with the local variables in equations 2 through 8 and with the common variables in equation 1.

[g]The test statistic has 6 degrees of freedom.

world price of covariance risk. The estimate is 11.5 which is closer to the single country estimates of Japan and the U.K. than the other $G$-7 countries. The $\chi^2$ test indicates that there is little evidence against the model's restrictions using common or local instrumentation. However, there is evidence at the single country level against the restrictions for four of the $G$-7 countries.

Intuitively, one would expect rejection in the multiple country system given strong rejections in three of the single country tests and a marginal rejection for another country. It has been argued that the multiple country test may lack power. An alternative test is presented in the final line of Table VII. In this estimation, the equality of the reward to risk ratios across the seven countries is explicitly tested. The test proceeds in two steps. Initially, a seven country system is estimated with country specific reward to risk ratios. The weighting matrix is saved from this unrestricted estimation as well as the final $\chi^2$ statistic. Second, a seven country system is estimated with the reward to risk ratios restricted to be the same across the seven countries. However, in the estimation the weighting matrix is the saved matrix from the unconstrained estimation. The difference in the final $\chi^2$ statistics is distributed $\chi^2$ with six degrees of freedom.[26] The results in the final line of Table VII provide convincing evidence against the model's null hypothesis that the ratios are the same. This multivariate test is powerful enough to reject the model's restrictions.

### *F. Diagnostics*

The asset pricing model implies that the coefficient $\lambda$ which transforms conditional covariance with the world market portfolio into conditionally expected returns is the same for all countries. The evidence presented in Table VII provides sharp evidence against this hypothesis. In the country by country estimation, the routine was fitting different world prices of covariance risk in order to match the conditional covariances with the conditionally expected returns. For example, the Japanese conditional covariance was not high enough to account for the large conditionally expected returns. The routine fit a $\lambda$ coefficient of 13.1, which is much higher than the average, to accommodate the higher expected returns.

Some authors including Merton (1980) have related the price of risk to the coefficient of relative risk aversion. It is tempting to interpret the results in Table VII as evidence of higher risk aversion in Japan. However, there are two important qualifications. First, the reward to risk ratio can only be linked to risk aversion if international markets are completely segmented. That is, in completely segmented markets, a country whose residents are more risk-averse will have a higher reward to risk ratio than other countries. However, few would argue that Japan is completely segmented. If a country is not completely segmented, then the relation between the reward to risk

[26]See Gallant and Jorgenson (1979), Newey and West (1987), Eichenbaum, Hansen, and Singleton (1988), and Eichenbaum and Hansen (1990) for discussions of this multivariate test.

ratio and risk aversion will depend on how domestic residents can access foreign markets and how foreign investors can access the domestic market.

Second, in a world of complete segmentation, the relevant reward to risk ratio is the conditionally expected own-country return divided by the own-country variance. This is not what is estimated in Table VII. To get an idea of the magnitude of the local reward to risk ratios, these are presented in Table VIII.

Two of the smallest reward to risk ratios are found in Italy and Spain—which have the smallest average returns. Interestingly, the highest reward to risk ratio is found in Japan. The magnitude of the measure is double that of the one found in the United States. Under the null hypothesis of complete segmentation, the differences in these ratios may account for the higher expected returns that Japan has experienced relative to the United States.

The table also provides tests of whether the ratio is constant through time. In 10 of the 17 countries, there is evidence at the 5% level of significance against the hypothesis that the ratio is constant through time. There is evidence against the hypothesis at the 10% level for two other countries. The inference and the parameter estimates are not sensitive to whether the common or local instruments are used.

The final line of Table VIII provides estimates of the world price of covariance risk. The evidence strongly suggests that the world price of covariance risk is not constant. These results may explain the rejections of the model tested in Table VII. That is, rejection of the model tested in Table VII could be caused by the inefficiency of the world portfolio and/or by incorrectly specifying the world price of covariance risk to be a constant. The results in the last line of Table VIII indicate that the world price of covariance risk is not constant.

The results in Table VIII consider reward to risk ratios for countries under the hypothesis of segmented markets. However, under complete segmentation, the U.S. dollar returns are no longer the relevant metric. In addition, excess returns should be calculated in excess of local short-term interest rates. Table IX re-estimates the reward to risk ratios using both own currency returns and local interest rates.

The results in Table IX are similar to those presented in Table VIII. Only four countries (Belgium, the Netherlands, Denmark, and Switzerland) have reward to risk ratios that are more than one standard error different from those presented in Table VIII. However, it should be noted that the sample is different for some of these countries because data on short-term interest rates are not available back to 1970 in some countries, e.g., the sample used for Switzerland contains 175 observations (75:9–89:5) in Table IX compared to 232 observations (70:2–89:5) in Table VIII. Also, there are two countries for which I could not obtain the short-term interest rates: Belgium and Hong Kong.

The $\chi^2$ test of the constancy of the reward to risk ratio generally tells a similar story. However, when own-currency returns are considered, Italy,

Table VIII

## Test of Whether the Price of Risk is Constant

The results are based on monthly data 1970:2–1989:5 (232 observations). The country returns $r_j$ are calculated in U.S. dollars in excess of the holding period return on the Treasury bill that is closest to 30 days to maturity. The equity data are from Morgan Stanley Capital International. Standard errors in parentheses are heteroskedasticity consistent. Generalized method of moments estimation of the following system of equations:

$$\eta_t = (u_{jt} \quad e_{jt}) = \begin{pmatrix} [r_{jt} - \mathbf{Z}_{t-1}\delta_j]' \\ [r_{jt} - \lambda_j^* u_{jt}^2]' \end{pmatrix}',$$

where $\delta_j$ are coefficients associated with the instrumental variables that are used to obtain the conditional mean return for country $j$, $u_j$ is the forecast error in the conditional mean of the country return, $e_j$ is the deviation from the return and the model's expected return, and $\lambda_j$ is the country-specific price of risk (expected return divided by variance of the returns). There are three sets of instrumental variables $\mathbf{Z}$ that are used in the estimation. The common set of predetermined instrumental variables are: a constant, the excess return on the world index, a dummy variable for the month of January, the 1-month return for holding a 90-day U.S. Treasury bill less the return on a 30-day bill, the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds, and the dividend yield on the Standard and Poor's 500 stock index less the return on a 30-day bill. Local instrument set A is the common instrument set augmented with the country-specific dividend yield. Local instrument set B includes the country-specific dividend yield and the country-specific excess return in place of the world excess return.

| Portfolio | $\lambda^*$ | Common Instruments $\chi^{2\text{a}}$ [$P$-value] | Local Instruments A: $\chi^2$ [$P$-value] | Local Instruments B: $\chi^2$ [$P$-value] |
|---|---|---|---|---|
| Hong Kong | 1.2539 | 18.13 | 18.49 | 14.25 |
| | (0.4926) | [.003] | [.005] | [.027] |
| Japan | 5.0597 | 18.10 | 17.98 | 13.63 |
| | (1.0934) | [.003] | [.006] | [.073] |
| Sweden | 2.9443 | 8.97 | 9.01 | 9.02 |
| | (1.0896) | [.110] | [.173] | [.173] |
| Norway | 1.6027 | 10.18 | 10.79 | 14.67 |
| | (0.8033) | [.070] | [.095] | [.023] |
| Belgium | 3.1584 | 16.48 | 16.67 | 16.90 |
| | (0.9994) | [.006] | [.011] | [.010] |
| Netherlands | 4.2864 | 16.00 | 18.47 | 15.96 |
| | (1.1884) | [.007] | [.005] | [.014] |
| United Kingdom | 2.5569 | 11.60 | 14.19 | 14.44 |
| | (0.6435) | [.041] | [.028] | [.025] |
| Denmark | 3.4518 | 13.66 | 13.63 | 11.89 |
| | (1.1364) | [.018] | [.034] | [.064] |
| France | 1.4355 | 7.35 | 8.53 | 8.41 |
| | (0.8928) | [.196] | [.202] | [.209] |
| Austria | 2.7478 | 14.32 | 14.91 | 14.51 |
| | (1.0645) | [.014] | [.021] | [.024] |
| Germany | 1.6569 | 9.08 | 9.56 | 9.06 |
| | (1.1074) | [.106] | [.144] | [.170] |

**Table VIII**—*Continued*

| Portfolio | $\lambda^*$ | Common Instruments $\chi^{2a}$ [*P*-value] | Local Instruments A: $\chi^2$ [*P*-value] | Local Instruments B: $\chi^2$ [*P*-value] |
|---|---|---|---|---|
| Switzerland | 1.9780 | 13.78 | 14.41 | 14.25 |
| | (1.1413) | [.017] | [.025] | [.025] |
| Australia | 1.1108 | 17.17 | 18.00 | 17.32 |
| | (0.8608) | [.004] | [.006] | [.008] |
| Canada | 1.9765 | 15.11 | 15.06 | 14.02 |
| | (1.1897) | [.010] | [.020] | [.029] |
| United States | 2.6655 | 19.26 | 15.59 | 19.57 |
| | (1.5090) | [.002] | [.030] | [.003] |
| Spain | 0.8077 | 7.80 | 7.96 | 9.65 |
| | (1.0175) | [.168] | [.241] | [.140] |
| Italy | 0.4034 | 7.10 | 7.10 | 8.40 |
| | (0.8384) | [.214] | [.312] | [.210] |
| World | 5.7238 | 21.06 | – | – |
| | (1.8272) | [< .001] | | |

[a]The minimized value of the GMM criterion function. *P*-value is the probability that a $\chi^2$ variate exceeds the sample value of the statistic. In the estimation with the common instrumental variables, there are one parameter and 6 orthogonality conditions, leaving 5 overidentifying conditions. In the estimation with the local information variables, there are one parameter and 7 orthogonality conditions, leaving 6 overidentifying conditions.

Germany, and Sweden are added to the list of countries that have significant variation in the reward to risk ratios. With the local currency returns, we can no longer reject the constancy of the ratio of expected returns to volatility for Switzerland. There is little difference in the test results across the different sets of conditioning information.

### *G. Risk and Return in October 1987*

Although the evidence suggests departures from conditional mean-variance efficiency, the asset pricing formulation may still be useful in explaining cross-sectional variation in returns. One phenomena that would be a challenge for the asset pricing model to explain is the cross-country variation of returns in October 1987. Some countries were hit much harder than others. For example, the excess return in the Australian market was $-45\%$ and in Hong Kong $-44\%$. In contrast, the Danish return loss was only 8%.

For a given world price of covariance risk in October 1987, the asset pricing model suggests that the most severe losses should be associated with countries with the highest risk in October 1987. Using data for 1981–1987, Roll (1988) shows that the unconditional betas are important in explaining the cross-sectional returns in October. However, no one has examined the

**Table IX**

**Test of Whether the Price of Risk is Constant Using Returns Calculated in Local Currency and Country-Specific Short-Term Interest Rates**

The country returns $lr_j$ are calculated in local currency in excess of the holding period return on the country's Treasury bill or the call money rate. The equity data are from Morgan Stanley Capital International. Standard errors in parentheses are heteroskedasticity consistent. Generalized method of moments is used to estimate the following system of equations:

$$\eta_t = (u_{jt} \quad e_{jt}) = \begin{pmatrix} [lr_{jt} - \mathbf{Z}_{t-1}\delta_j]' \\ [lr_{jt} - \lambda_j^* u_{jt}^2]' \end{pmatrix}',$$

where $u_j$ is the forecast error in the conditional mean return for the country portfolio, and $\lambda_j$ is the country-specific price of risk (or ratio of the expected return to the variance of returns). There are three sets of instrumental variables $\mathbf{Z}$ used in the estimation. The common set of predetermined instrumental variables are: a constant, the excess return on the world index calculated in U.S. dollars, a dummy variable for the month of January, the 1-month return for holding a 90-day U.S. Treasury bill less the return on a 30-day bill, the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds, and the dividend yield on the Standard and Poor's 500 stock index less the return on a 30-day bill. Local instrument set A is the common instrument set augmented with the country-specific dividend yield. Local instrument set B includes the country-specific dividend yield and the country-specific excess return in local currency in place of the world excess return.

| Portfolio and (number of observations) | $\lambda^*$ | Common Instruments $\chi^{2a}$ [*P*-value] | Local Instruments A: $\chi^2$ [*P*-value] | Local Instruments B: $\chi^2$ [*P*-value] |
|---|---|---|---|---|
| Hong Kong[b] | 1.9082 | 19.56 | 20.04 | 17.08 |
| 70:2–89:5 (232) | (0.5739) | [.002] | [.003] | [.009] |
| Japan | 5.7076 | 17.57 | 17.57 | 11.89 |
| 70:2–89:5 (232) | (1.4382) | [.004] | [.007] | [.064] |
| Sweden | 2.7887 | 11.86 | 14.20 | 14.97 |
| 70:2–87:1 (204) | (1.2409) | [.037] | [.027] | [.021] |
| Norway[d] | 0.8359 | 8.36 | 8.98 | 13.20 |
| 71:9–89:5 (213) | (0.8624) | [.138] | [.175] | [.040] |
| Belgium[b] | 7.4042 | 14.25 | 13.51 | 13.99 |
| 70:2–89:5 (232) | (1.4294) | [.014] | [.036] | [.030] |
| Netherlands | 6.1320 | 22.50 | 22.38 | 21.63 |
| 70:2–89:5 (232) | (1.3633) | [< .001] | [.001] | [.001] |
| United Kingdom | 3.5292 | 10.67 | 14.38 | 14.72 |
| 72:1–89:5 (208) | (0.8786) | [.058] | [.026] | [.023] |
| Denmark | 1.8502 | 14.96 | 15.14 | 16.94 |
| 72:2–89:1 (204) | (1.3639) | [.010] | [.019] | [.012] |
| France[c] | 1.1956 | 8.02 | 9.45 | 8.66 |
| 70:2–89:5 (229) | (1.0487) | [.155] | [.150] | [.193] |
| Austria | 2.3982 | 16.88 | 18.09 | 8.49 |
| 71:12–89:1 (216) | (1.2800) | [.005] | [.006] | [.204] |
| Germany | 1.4855 | 11.92 | 12.65 | 12.31 |
| 70:2–89:5 (232) | (1.3246) | [.036] | [.049] | [.055] |

**Table IX**—*continued*

| Portfolio and (number of observations) | $\lambda^*$ | Common Instruments $\chi^{2^a}$ [*P*-value] | Local Instruments A: $\chi^2$ [*P*-value] | Local Instruments B: $\chi^2$ [*P*-value] |
|---|---|---|---|---|
| Switzerland | 6.1072 | 5.98 | 6.38 | 7.05 |
| 75:9–89:5 (175) | (2.4113) | [.309] | [.382] | [.316] |
| Australia | 1.0063 | 19.88 | 22.13 | 20.67 |
| 70:2–89:1 (228) | (0.9840) | [.001] | [.001] | [.002] |
| Canada | 1.4874 | 15.17 | 15.52 | 13.76 |
| 70:2–89:5 (232) | (1.2706) | [.010] | [.017] | [.032] |
| United States | 2.6655 | 19.26 | 15.59 | 19.57 |
| 70:2–89:5 (232) | (1.5090) | [.002] | [.030] | [.003] |
| Spain | −0.1027 | 6.93 | 6.95 | 8.64 |
| 74:1–89:5 (184) | (1.1953) | [.226] | [.325] | [.195] |
| Italy | 0.4415 | 11.15 | 11.62 | 12.02 |
| 71:2–89:5 (220) | (0.9056) | [.048] | [.071] | [.061] |

[a]The minimized value of the GMM criterion function. *P*-value is the probability that a $\chi^2$ variate exceeds the sample value of the statistic. In the estimation with the common instrumental variables, there are one parameter and 6 orthogonality conditions, leaving 5 overidentifying conditions. In the estimation with the local instrumental variables, there are one parameter and 7 orthogonality conditions, leaving 6 overidentifying conditions.

[b]Short term interest rates are not available. The returns are not excess returns.

[c]Excludes 1986:3–1986:6 when the interest rate data are not available.

[d]Excludes 1980:8–1980:9 when the interest rate data are not available.

conditional risk in October 1987.[27] To do this, conditional covariances were estimated for October 1987 based on the common information variables available in September 1987. Hence, the fitted conditional covariance is an out-of-sample forecast of the risk.

The fitted conditional covariances are plotted against the returns in October 1987 in Figure 2. Consistent with the asset pricing theory, there is roughly a linear relation between the returns and the conditional covariances. The slope of the relation is *negative*. However, this is reasonable because the out-of-sample world price of covariance risk is negative in October 1987. To ensure that this result is not a fluke, Figure 3 provides a plot of the same variables for October 1986. In that month, the forecasted world price of covariance risk is positive, and the relation between the returns and the conditional covariances is roughly positive.

In any given month, the world price of covariance risk is fixed. A cross-sectional regression of the October 1987 returns on the estimated conditional risk explains 41% of the variation. Hence, the differences in the countries'

[27]Ferson and Harvey (1991) argue that the rolling unconditional beta may be interpreted as a conditional risk measure. In fact, for October 1987 the correlation of Roll's 6-year betas and my conditional covariances is 67.3%.

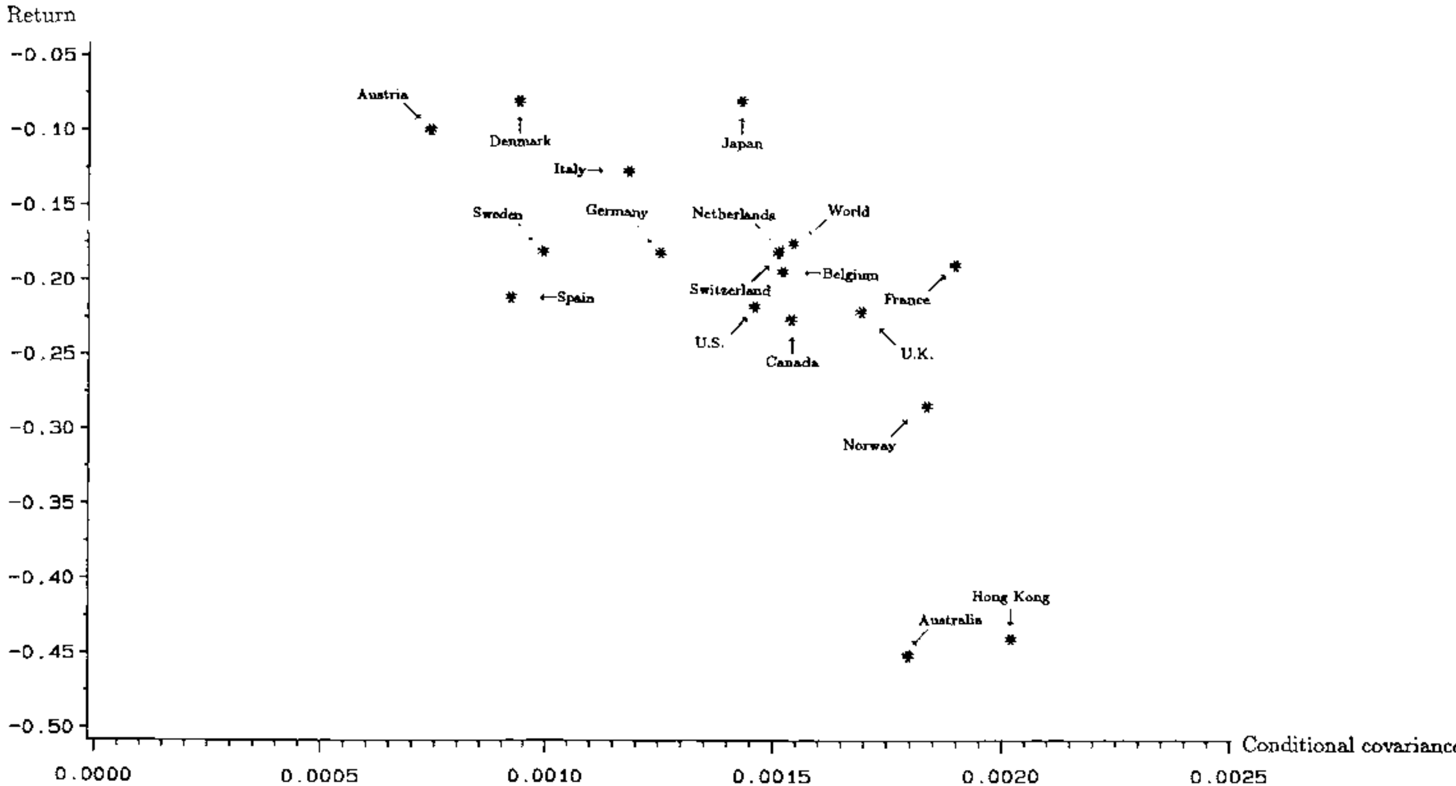


**Figure 2. Realized returns and conditional covariances in October 1987.** Conditional covariances of 17 countries' returns with the world portfolio and the conditional variance of the world portfolio are forecasted based on data available through September 1987. First, the excess returns are forecasted through September 1987. The instrumental variables used in the forecasting model are: a constant, the lagged excess return on the world index, a dummy variable for the month of January, the return for holding a 90-day U.S. Treasury bill for 1 month less the return on a 30-day bill, the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds, and the dividend yield on the Standard and Poor's 500 stock index less the return on a 30-day bill. Second, the residuals from each country's forecast are multiplied by the residuals from the forecasting model for the world portfolio. The product of the residuals are projected on the same instrumental variables. Also, the squares of the residuals from the world forecasting model are projected on the same instrumental variables. Finally, the coefficients from these regressions are applied to the values of the instruments in September 1987 to obtain out-of-sample forecasts of the conditional covariances and the conditional variance for October 1987. The returns are in U.S. dollars and are from Morgan Stanley Capital International.

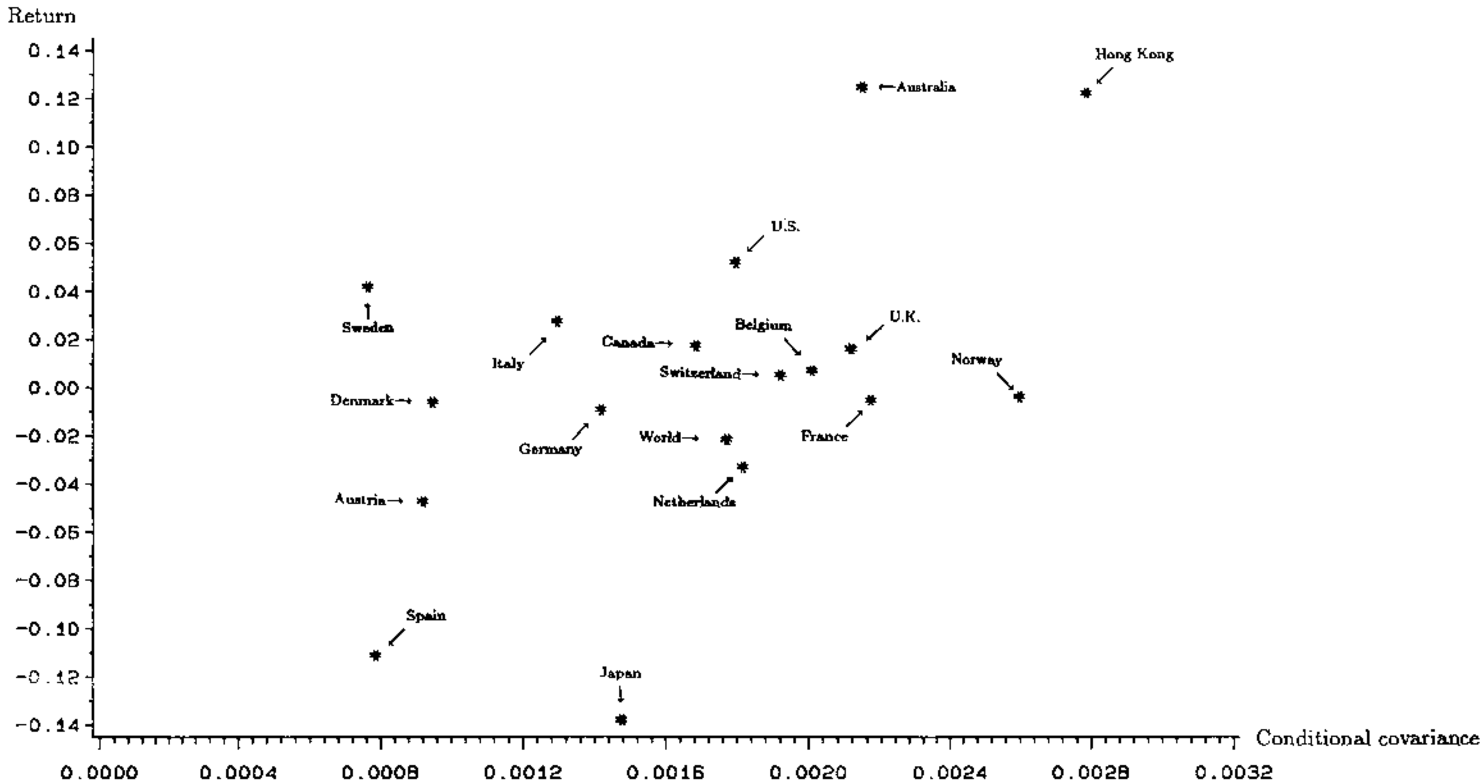


**Figure 3.** Realized returns and conditional covariances in October 1986. Conditional covariances of 17 countries' returns with the world portfolio and the conditional variance of the world portfolio are forecasted based on data available through September 1986. First, the excess returns are forecasted through September 1986. The instrumental variables used in the forecasting model are: a constant, the lagged excess return on the world index, a dummy variable for the month of January, the return for holding a 90-day U.S. Treasury bill for 1 month less the return on a 30-day bill, the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds, and the dividend yield on the Standard and Poor's 500 stock index less the return on a 30-day bill. Second, the residuals from each country's forecast are multiplied by the residuals from the forecasting model for the world portfolio. The product of the residuals are projected on the same instrumental variables. Also, the squares of the residuals from the world forecasting model are projected on the same instrumental variables. Finally, the coefficients from these regressions are applied to the values of the instruments in September 1986 to obtain out-of-sample forecasts of the conditional covariances and the conditional variance for October 1986. The returns are in U.S. dollars and are from Morgan Stanley Capital International.

conditional covariances appear to account for a large portion of the differences in country performance.

## IV. Conclusions

Tests of the conditional version of the Sharpe-Lintner capital asset pricing model are executed with country-specific stock portfolios. The tests assume that the representative investor only cares about U.S. dollar returns. Capital markets are also assumed to be fully integrated.

The tests allow for time-varying conditional moments. For most countries, a single source of risk appears to adequately describe the cross-sectional variation in returns across different countries. In an example, the differences in conditional covariances are able to account for a large portion of the different losses that countries experienced in October 1987. However, the model's restrictions are consistently rejected for Japan. Japan's covariance risk explains some—but not all—of its performance in the 1970:2–1989:5 sample.

However, all tests are joint tests of many hypotheses. An alternative hypothesis for the Japanese performance is that the market is not fully integrated. In this case, the Japanese covariance with the world market portfolio is not the relevant risk measure. Furthermore, the world price of risk is not the appropriate price of covariance risk. Evidence is presented that supports the hypothesis that the price of risk may be higher in Japan. This is consistent with the Japanese market performance over the sample.

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