The Risk and Predictability of International Equity Returns

Wayne E. Ferson
University of Washington

Campbell R. Harvey
Duke University

We investigate predictability in national equity market returns, and its relation to global economic risks. We show how to consistently estimate the fraction of the predictable variation that is captured by an asset pricing model for the expected returns. We use a model in which conditional betas of the national equity markets depend on local information variables, while global risk premia depend on global variables. We examine single- and multiple-beta models, using monthly data for 1970 to 1989. The models capture much of the predictability for many countries. Most of this is related to time variation in the global risk premia.

We investigate the sources of risk and predictability of international equity market returns. We examine several global economic risk factors, including a world market portfolio, exchange rate fluctuations, mea-

Part of this research was conducted while the authors were at the University of Chicago, Graduate School of Business. We thank Warren Bailey, Eugene Fama, Stephen Foerster, Kenneth French, Allan Kleidon, Bruce Lehmann (a referee), Johnny Liew, John D. Martin, Lars Nielson, Bruno Solnik, an anonymous referee, the editor Richard Green, and participants in workshops at the Universities of California at Berkeley, Chicago, Florida, Iowa, Manitoba, Stanford, Texas at Austin, Texas at Dallas, Tel Aviv, Queens, Washington, at the August, 1991 National Bureau of Economic Research Summer Institute's International Studies/Macroeconomics week, the 1992 Utah Winter Finance Conference, the 1992 Western Finance Association, the 1992 French Finance Association and the September 1992 Berkeley Program in Finance for helpful discussions. Ferson acknowledges financial support from the Pigott–Paccar professorship at the University of Washington. Address correspondence to Wayne E. Ferson, Department of Finance and Business Economics, DJ-10, University of Washington, Seattle, WA 98195.

The Review of Financial Studies 1993 Volume 6, number 3, pp. 527–566

sures of global inflation, world interest rates, international default risk, and world industrial production. We formulate an empirical beta pricing model, where country-specific conditional betas measure sensitivity to the global risk factors. Both the betas and the expected risk premia can vary over time.

Most tests of conditional asset pricing models ask the models to explain 100 percent of the predictability of the asset returns. Since a model can be useful even if it does not account for all of the variance, we estimate directly the fraction of the predictability that is explained by the model and the fraction that is left unexplained. We develop an approach for consistently estimating such fractions, using the generalized method of moments.

We study the predictability in 18 national equity market returns, using regressions on predetermined variables. Such regressions have been examined before, but our focus is unique. We concentrate on the marginal impact of local market variables, given a common set of instruments representing the state of the global economy. The local information is often important, and our regressions suggest that its effect on country returns is related to the country-specific betas.

Estimating the beta pricing models, we find that they can capture substantial fractions of the predictability for many of the countries. Single-beta models, using the world market index, are compared with multiple-beta models, which better explain the predictability in returns. Movements in the betas, while statistically significant, contribute only a small fraction to the predicted variation in expected returns. The global risk premia appear to be the dominant source of the predictability.

The paper is organized as follows. Section 1 reviews the models. Section 2 describes our empirical methodology, and Section 3 introduces the data. The empirical results are presented in Section 4, and Section 5 offers some concluding remarks.

1. The Models

The usual objective of empirical work on international asset pricing models is to explain differences in average returns. Average returns are estimates of unconditional expected returns, formed using no information about the current state of the economy. However, asset pricing models may also be interpreted as statements about expected returns *conditional* on currently available information. We focus on the ability of beta pricing models to capture the predictability of international equity market returns through conditional expected risk premia and conditional betas.

Beta pricing models to describe expected returns across countries have been developed by a number of authors, who show that the models require strong assumptions. We assume that the national equity markets are perfectly integrated in a global economy, with no barriers to extranational equity investments, no transactions or information costs, and no taxes. Such extreme assumptions are unlikely to provide a good approximation to the actual complexity of international investments. Our approach is to see how far one can go in capturing equity market predictability by using such a simple framework. The results are encouraging, and we expect that further refinements of the models should produce even better explanatory power.

If we assume rational expectations, actual returns differ from their conditional expected values in the model by an error term that is orthogonal to the conditioning information. The conditioning information, Ω_{t-1}, is assumed to be public knowledge at time $t-1$. Predictability of returns is attributed to the correlation between expected returns and the current information. Following previous studies, the information, Ω_{t-1}, is persistent over time, and the expected returns inherit this persistence. We model expected returns as functions of betas and risk premia. Therefore, predictability should arise because betas or risk premia are correlated with the information variables. We assume that conditional expected returns can be written as

$$E(R_{it} \mid \Omega_{t-1}) = \lambda_0(\Omega_{t-1}) + \sum_{j=1}^{K} b_{ij}(\Omega_{t-1})\lambda_j(\Omega_{t-1}), \qquad (1)$$

where the $b_{ij}(\Omega_{t-1})$ are the conditional regression betas of the returns, R_{it}, measured in a common currency, on K global risk factors, $j = 1,\ldots,K$. The expected risk premia, $\lambda_j(\Omega_{t-1})$, $j = 1,\ldots,K$, are the expected excess returns on *mimicking portfolios* for the risk factors, similar to the static models of Huberman, Kandel, and Stambaugh (1987) and Lehmann and Modest (1988).[1] The intercept $\lambda_0(\Omega_{t-1})$ is the expected return of portfolios with all of their betas equal to zero. If there is a risk-free asset available at time $t-1$, then its rate of return equals $\lambda_0(\Omega_{t-1})$. Equation (1) implies an expression for the expected *excess* returns:

$$E(r_{it} \mid \Omega_{t-1}) = \sum_{j=1}^{K} \beta_{ij}(\Omega_{t-1})\lambda_j(\Omega_{t-1}), \qquad (2)$$

where $\beta_{ij}(\Omega_{t-1}) = b_{ij}(\Omega_{t-1}) - b_{fj}(\Omega_{t-1})$ are the conditional betas of the excess returns and $b_{fj}(\Omega_{t-1})$, $j = 1,\ldots,K$, are the conditional betas of a Treasury bill.

[1] Mimicking portfolios are defined as portfolios that may be substituted for the factors in a factor model regression, to measure the betas, and whose expected excess returns are the risk premiums.

1.1 Choosing the factors

The mean-variance mathematics [e.g., Roll (1977)] implies that some portfolio can always serve as a single factor, such that Equations (1) and (2) are satisfied. Therefore, the choice of factors determines the empirical content of the models. Although we do not attempt to test specific international asset pricing theories, our choice of factors follows previous theoretical and empirical work on international asset pricing. We study both single- and multiple-factor models. The single-factor model is similar to the capital asset pricing model (CAPM) of Sharpe (1964) and Lintner (1965). Stulz (1981b, 1984) and Adler and Dumas (1983) provide conditions under which a single-beta CAPM based on the world market portfolio holds globally. They assume no exchange risk and a constant investment opportunity set, in addition to the general assumptions we have described. Our investigation of the CAPM should therefore be viewed as extending static model restrictions to a conditional setting, similar to Harvey (1991).

When purchasing power parity does not hold, then consumers face exchange risk for investing internationally. Solnik (1974) showed that exchange risks should be "priced" even in a world otherwise similar to that of the static CAPM. Adler and Dumas [1983, Equation (14)] present a model in which a combination of the world market and measures of exchange risk is mean-variance efficient. The exchange risk can be broken down into a separate factor for each currency, as in Dumas and Solnik (1992), or approximated by a single variable. We study a two-beta model in the spirit of the latter approach, using the world market portfolio and an aggregate of exchange risks as the two factors.

International equilibrium and arbitrage pricing (APT) models with several risk factors are described by Stulz (1981a), Hodrick (1981), Ross and Walsh (1983), and Bansal, Hsieh, and Viswanathan (1992), among others. The central intuition of such models is that only the pervasive sources of common variation should be priced. Korajczyk and Viallet (1989) and Heston, Rouwenhorst, and Wessels (1991) find evidence for several common sources of variation in U.S. and European stocks, which suggests that a number of worldwide risk factors may be important. We therefore study models with several global risk factors.

1.2 Modeling conditioning information

We study the *predictable variation,* which we define as the unconditional variance of the conditional expected excess returns. We measure to what extent various specifications of Equation (2) can capture the predictable variation, using time-varying country-specific conditional betas and global risk premia. Previous studies find that the conditional second moments of national equity market returns move over time in association with lagged variables [e.g., King, Sentana, and Wadhwani (1990), Harvey (1991)]. Other studies find evidence of time-varying betas for international asset returns [e.g., Giovannini and Jorion (1987, 1989), Mark (1985)]. We therefore allow for time variation in both the expected risk premia and the conditional betas.

Let $\Omega_{t-1} = \{Z_{t-1}, Z^i_{t-1}, i = 1,...,n\}$, where Z_{t-1} represents our global information variables and Z^i_{t-1}, our local information variables for country i. We assume globally integrated capital markets, which implies that the risk premia should not be country-specific. We therefore restrict the risk premia in (2) to depend only on the global variables, Z_{t-1}. Exploratory regressions, described here, suggest that the local market information variables are related to country-specific betas. In the interest of parsimony, therefore, we assume that the betas are functions only of the local market information and model the predictable variance, using Equation (2), as

$$\mathrm{Var}\{E(r_{it} \mid \Omega_{t-1})\} = \mathrm{Var}\left\{\sum_{j=1}^{K} \beta_{ij}(Z^i_{t-1})\lambda_j(Z_{t-1})\right\}. \qquad (3)$$

Some informal intuition for the impact of the restrictions in (3) for country i can be obtained by assuming that $E(r_{it} \mid \Omega_{t-1})$ is a function $f(Z^i_{t-1}, Z_{t-1})$. Dropping the subscripts, consider an example where there is a single factor ($K = 1$), where β, λ, Z^i, and Z are scalars and where Z^i is uncorrelated with Z. Writing $f(Z^i, Z) = \beta(Z^i, Z)\lambda(Z^i, Z)$ and taking a first-order Taylor series about the means, we have

$$\begin{aligned}\mathrm{Var}(f) \approx &\left[\lambda(\cdot)\frac{\partial\beta}{\partial Z^i} + \beta(\cdot)\frac{\partial\lambda}{\partial Z^i}\right]^2 \mathrm{Var}(Z^i) \\ &+ \left[\lambda(\cdot)\frac{\partial\beta}{\partial Z} + \beta(\cdot)\frac{\partial\lambda}{\partial Z}\right]^2 \mathrm{Var}(Z),\end{aligned} \qquad (4)$$

where $\lambda(\cdot)$ and $\beta(\cdot)$ are evaluated at the means. The first term captures the contribution of the local information to the predictable variance of country i's return, and the second term captures the contribution of the global information. Market integration can be interpreted as implying that $\partial\lambda/\partial Z^i = 0$ in the first term. The assumption that the betas depend only on the local market information implies that $\partial\beta/\partial Z = 0$ in the second term. By setting $\partial\beta/\partial Z = 0$, we are ignoring what should be the smaller of the coefficients that scale the variance in the second term of (4). This occurs because the square of an average risk premium is a small number compared with the square of an average beta. The term that we retain should capture the dominant effect of the global information variables on the predictable variation.

2. Methodology

To estimate the fraction of the predictable variation that a beta pricing model captures, we use a regression of the excess country return, r_{it}, on the information variables as a base case. Returns are measured in a common currency, which we choose to be the U.S. dollar. (Later, we investigate the sensitivity of the results to the currency of denomination.) With a linear regression model for the conditional expected return given Z_{t-1}, $E(r_{it} \mid Z_{t-1}) = Z'_{t-1}\delta_i$, where δ_i is the coefficient vector. The predictable variance of the return, using Z_{t-1}, is $\mathrm{Var}[E(r_{it} \mid Z_{t-1})] = \mathrm{Var}[Z'_{t-1}\delta_i]$.[2]

The predictable variation captured by the model depends on the conditional betas and the risk premia. We use a linear regression to model the expected risk premia, following much of the literature on conditional asset pricing. That is, we assume that $\lambda(Z_{t-1}) = E(F_t \mid Z_{t-1}) = \gamma' Z_{t-1}$, where γ is an $L \times K$ matrix of coefficients and the F_t are mimicking portfolio excess returns for K risk factors. We approximate the conditional betas as linear functions of the local information variables: $\beta_{ij}(Z^i_{t-1}) = \kappa'_i Z^i_{t-1}$, where κ_i is an $L \times K$ matrix of coefficients that describe the conditional betas for country i as a linear function of the lagged, local market variables.[3] With these assumptions, the predictable variance of the return captured by the beta pricing model is $\mathrm{Var}[\Sigma_j E(F_{jt} \mid Z_{t-1})\beta_{ij}(Z^i_{t-1})] = \mathrm{Var}[Z'_{t-1}\gamma\kappa'_i Z^i_{t-1}]$. We express this as a proportion, defining the following variance ratio:

$$\mathrm{VR1}_i = \frac{\mathrm{Var}\left[E\left(\sum_{j=1}^{K} F_{jt} \mid Z_{t-1}\right)\beta_{ij}(Z^i_{t-1})\right]}{\mathrm{Var}[E(r_{it} \mid Z_{t-1})]} = \frac{\mathrm{Var}[Z'_{t-1}\gamma\kappa'_i Z^i_{t-1}]}{\mathrm{Var}[Z'_{t-1}\delta_i]}. \tag{5}$$

The variance ratio VR1 measures the fraction of the predictable variance in the return attributed to the model.

We estimate the model by first defining the following error terms for each country i:

$$u1_{it} = (r_{it} - Z'_{t-1}\delta_i), \tag{6a}$$

[2] We also report results where the conditional variance of the expected return is formed by regressing the country return on both the global and the local information variables.

[3] Linear approximations for betas are used by Campbell (1987) and Shanken (1990), among others. A problem common to all such approaches, including ours, is that the information set used in the empirical work is implicitly assumed to represent all publicly available information. Our "unrestricted" regression for the predictable variation does not nest the expected return predicted by the model, as would normally be the situation for hypothesis tests. The large number of product terms would make such an approach unwieldy here.

$$u2_{it} = (F'_t - Z'_{t-1}\gamma)', \tag{6b}$$

$$u3_{it} = [(u2_{it}u2'_{it})(\kappa'_i Z^i_{t-1}) - (F_t u1'_{it})], \tag{6c}$$

$$u4_{it} = (Z'_{t-1}\delta_i - \theta_i), \tag{6d}$$

$$u5_{it} = (Z'_{t-1}\gamma)(\kappa'_i Z^i_{t-1}) - \theta_i + \alpha_i, \tag{6e}$$

$$u6_{it} = (u4^2_{it})\mathrm{VR1}_i - u5^2_{it}. \tag{6f}$$

The parameters are $\{\theta_i, \alpha_i, \mathrm{VR1}_i, \gamma, \delta_i, \kappa_i\}$, where the first three parameters are scalars. The parameter α_i is the difference between the unconditional mean return and the unconditional mean of the model fitted return. It therefore measures an "average pricing error," analogous to the traditional α measure of performance. If the model is well specified, α_i should be zero.

The model implies the orthogonality conditions[4]

$$E(u1_{it}Z_{t-1},\ u2_{it}Z'_{t-1},\ u3_{it}Z^{i\prime}_{t-1},\ u4_{it},\ u5_{it},\ u6_{it}) = 0.$$

The number of orthogonality conditions and the number of parameters in the system are $2LK + L + 3$, and the system is exactly identified. The model is estimated for each country by using Hansen's (1982) generalized method of moments (GMM).[5]

We also modify system (6) to obtain a complementary variance ratio:[6]

$$\mathrm{VR2}_i = \frac{\mathrm{Var}\left[E(r_{it} \mid Z_{t-1}) - \sum_{j=1}^{K} \beta_{ij}(Z^i_{t-1})E(F_{jt} \mid Z_{t-1})\right]}{\mathrm{Var}[E(r_{it} \mid Z_{t-1})]}$$

$$= \frac{\mathrm{Var}[Z'_{t-1}\delta_i - Z'_{t-1}\gamma\kappa'_i Z^i_{t-1}]}{\mathrm{Var}[Z'_{t-1}\delta_i]}. \tag{7}$$

The ratio VR2 measures the predictable variation in the return that is not captured by the model.

The difference between the returns and the model expected returns should have the property that their expected values, given *all* of the

[4] As a check, we reestimated the model for a number of cases, forcing the error term in the conditional beta equation (6c) to be orthogonal to the global instruments, Z_{t-1}, instead of the local instruments. The results were broadly similar.

[5] The system is estimated separately for each country in order to keep the size of the problem tractable. As the system is exactly identified, the point estimates of the parameters are the same as they would be if the same system was estimated jointly across the countries. We use $L = 7$ global information variables, so the number of moment conditions in a model with $K = 5$ factors is 80. We also have seven local market information variables for each country, and 239 monthly return observations. We are unable to estimate system (6) using all of the variables, as the number of orthogonality conditions would exceed the number of observations.

[6] To estimate $\mathrm{VR2}_i$, we replace the model-fitted part of the return $Z'_{t-1}\gamma\kappa'_i Z^i_{t-1}$ in (6e) with $Z'_{t-1}\delta_i - Z'_{t-1}\gamma\kappa'_i Z^i_{t-1}$.

conditioning information, Ω_{t-1}, are zero. We have imposed many restrictions on the model, requiring the local information variables to enter only through the betas and the world information to enter only through the risk premia. Although we are forced to estimate the models by using subsets of the information, we conduct diagnostics on the model "pricing errors," which are

$$\epsilon_{it} = r_{it} - Z'_{t-1}\gamma\kappa'_{t} Z^{i}_{t-1}. \quad (8)$$

The pricing errors are the sum of the unexpected part of the returns and any specification error in the model for the expected returns. If the model is well specified, we should find that $E(\epsilon_{it} \mid Z^{i}_{t-1}) = E(\epsilon_{it} \mid Z_{t-1}) = 0$; that is, the pricing errors should be unpredictable. (The unconditional mean of the pricing error is α_i, which should also equal zero.)

As a check on the sensitivity of our results, we estimate variance ratios by the cross-sectional regression (CSR) methods of Fama and MacBeth (1973), as employed by Ferson and Harvey (1991) in a domestic context. This approach, which we review in the Appendix, is a multistep procedure that allows for time-varying covariances, variances, betas, and expected returns.

3. The Data

3.1 Country returns

We study equity returns for 18 national markets as provided by Morgan Stanley Capital International (MSCI). Total monthly returns are used for 1970 to 1989. The U.S. dollar returns are measured in excess of the U.S. Treasury bill that is the closest to 30 days to maturity, as provided by the Center for Research in Security Prices (CRSP) at the University of Chicago. To convert from local currency values to U.S. dollar values, we use the closing European interbank currency rates from MSCI on the last trading day of the month.

3.2 Global economic risk variables

We construct a set of variables to represent global economic risks. Our approach is to choose variables a priori and to investigate their importance with simple, factor model regressions. Then we study the pricing of the most important risks. Summary statistics for the variables are presented in Table 1; details about the data sources and definitions are provided in the Appendix.

WDRET is the U.S. dollar return of the MSCI world equity market in excess of a short-term interest rate. Asset pricing models usually include a role for a "market portfolio" as a measure of risk. Harvey

(1991) studies the risk premium associated with conditional covariances of returns with the world stock return index and concludes that it partially explains the differential performance of the U.S. and Japanese stock markets. The world equity market index is a value-weighted combination of the country returns.[7]

dG10FX is the log first difference in the trade-weighted U.S. dollar prices of the currencies of 10 industrialized countries. The G-10 countries are defined as the G-7 (not including the United States), plus the Netherlands, Belgium, Sweden, and Switzerland. The G-7 countries are Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States. This series is from the International Monetary Fund (IMF), as reported by Citibase. A positive change (dG10FX > 0) indicates a depreciation of the dollar. In Adler and Dumas (1983, Equation 14), the theoretical exchange risk factor depends on exchange rates, consumer price index changes, and risk tolerance in each country. This factor is difficult to replicate empirically, so Dumas and Solnik (1992) break it down into separate variables for each country. We use a single aggregate measure as a parsimonious alternative to the approach of Dumas and Solnik (1992). Previous studies examine the pricing of exchange risks in national equity markets. They find little evidence that exchange risks are priced on average, but there is some evidence for time-varying currency risk premia.[8] A unique feature of our study is to use a global measure of exchange risk in a multicountry asset pricing model.

G7UI is the unexpected component of a monthly global inflation measure. The G-7 inflation rate is a weighted average of the percentage changes in the consumer price indices (CPI) in the G-7 countries, using the relative shares of the total real, gross domestic product (GDP) as the weights. An inflation state variable could arise in a multibeta model if inflation has real effects, in the general sense that global inflation is correlated with marginal utility. For example, higher inflation may signal higher levels of economic uncertainty, which make consumers worse off. If national equity market returns differ in their exposure to changes in the global inflation outlook, there may be an inflation risk premium in global equity markets.[9]

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

[8] See Hamao (1988), Bodurtha, Cho, and Senbet (1989), Brown and Otsuki (1990a,b), and Dumas and Solnik (1992).

[9] If higher inflation makes consumers worse off and therefore is associated with higher real marginal utility, we would expect a negative inflation premium.

Table 1
Summary statistics for the world risk factors and instrumental variables: 1970:2–1989:12 (239 observations)

Variable	Symbol	Mean	Std. dev.	ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	ρ_{24}
			World risk factors						
World excess return	wdret	0.545	4.189	.15	−.03	.05	.03	.09	.00
Change in Eurodollar–Treasury yield	dted	−0.046	3.988	−.08	−.09	−.07	−.13	.02	.09
Log change in G-10 foreign exchange rate	dG10fx	0.104	2.099	.31	.06	.09	.08	.06	−.07
Unexpected G-7 inflation	dG7ui	−0.005	0.204	.00	.04	−.02	−.09	.01	.08
Change in long-term G-7 expected inflation	dG7elt	−0.039	1.275	−.34	−.12	.09	−.16	−.08	.01
Change in price of oil	doil	0.062	0.861	.56	.22	.08	.03	−.02	−.02
Change in G-7 industrial production	dG7ip	0.215	0.817	.11	.27	.24	.15	−.08	−.17
G-7 real interest rate	G7rtb	0.132	0.316	.67	.52	.53	.52	.67	.55
			World instrumental variables						
Lagged world excess return	wr	0.523	4.211	.14	−.02	.07	.04	.08	−.00
Lagged world dividend yield	wrddiv	0.320	0.080	.98	.96	.95	.93	.76	.57
Lagged Eurodollar–Treasury yield spread	ted	0.121	0.084	.74	.57	.40	.44	.26	−.01
Lagged slope of U.S. term structure	term	0.111	0.118	.87	.76	.70	.64	.28	−.08
30-day U.S. Treasury bill rate	tb1	0.597	0.224	.93	.87	.82	.78	.59	.28

Correlations of the world risk factors

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

Correlations of the world instruments

Variable	wr	jan	divwrd	ted	term	tb1
wr	1.000					
jan	0.091	1.000				
divwrd	−0.156	0.016	1.000			
ted	−0.347	0.078	0.350	1.000		
term	0.138	−0.006	−0.271	−0.414	1.000	
tb1	−0.230	−0.059	0.518	0.390	−0.530	1.000

World risk factors

The world excess return is the arithmetic return on the Morgan Stanley Capital International world equity index (including dividends) minus the Ibbotson Associates one-month bill rate. The change in the Eurodollar–Treasury yield spread is the first difference of the spread between the 90-day Eurodollar yield and the 90-day Treasury-bill yield (from the Federal Reserve Bulletin). The log change in the G-10 foreign exchange rate is based on the trade-weighted dollar per foreign exchange

dG7ELT is the monthly change in a measure of long-term inflationary expectations. Chen, Roll, and Ross (1986) include a measure of unexpected inflation and a measure of changes in expected inflation in their study for the United States. dG7ELT is formed by regressing a 48-month moving average of the G-7 inflation rate on our predetermined global information variables and taking the first difference of the fitted values.

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

G7RTB is a weighted average of short-term interest rates in the G-7 countries, using the shares of G-7 GDP as the weights, minus the G-7 inflation rate. Real interest rates are often used in economic models to capture the state of investment opportunities. For example, Merton (1973) and Cox, Ingersoll, and Ross (1985) develop models in which interest rates are state variables. Chen, Roll, and Ross (1986) and Ferson and Harvey (1991) include real interest rate risk in empirical models for the U.S. market.[10]

dOIL is the change in the monthly average U.S. dollar price per barrel of crude oil. Chen, Roll, and Ross (1986) propose oil prices as a measure of economic risk in the U.S. market, and Hamao (1988)

[10] Although the correlation between G7RTB and G7UI is relatively high (at −.56), it is not perfect because the G-7 nominal interest rates are not part of the conditioning information used to form G7UI and because G7RTB is not prewhitened.

←

rate of 10 industrialized countries (G-7 plus the Netherlands, Belgium, Sweden, and Switzerland), from the International Monetary Fund. The unexpected inflation for the G-7 countries is derived from a time-series model applied to an aggregate G-7 inflation rate where the (varying) weights in the aggregate are determined by country weights in total G-7 gross domestic product. The change in long term G-7 expected inflation is found by regressing a 48-month moving average of the G-7 inflation rate on the lagged instrumental variables and taking the first difference of the fitted values. The change in the price of oil is the log change in the average U.S. dollar price per barrel at the wellhead from 1974 to 1989 and the posted West Texas Intermediate price from 1969 to 1973. The change in G-7 industrial production is calculated by weighting local industrial production index levels by the following weights: Canada, .04314; France, .09833; Germany, .05794; Italy, .13093; Japan, .07485; U.K., .11137; U.S., .48343, which are the weights in G-7 gross domestic product in the third quarter of 1969. The growth rate is the logarithmic difference in the aggregate industrial production index. The G-7 real interest rate is calculated by aggregating individual countries' short-term interest rates minus inflation rates using (varying) weights based on quarterly shares in G-7 gross domestic product.

World instrumental variables

The instrumental variables are the lagged return of the Morgan Stanley Capital International world index in excess of the CRSP 30-day bill, a dummy variable for the month of January, the dividend yield (based on the past 12 months' dividends) on the Morgan Stanley Capital International world equity index, the difference between the 90-day Eurodollar rate and the CRSP three-month Treasury-bill yield, the difference between the U.S. 10-year Treasury-bond yield and the CRSP three-month bill yield, and the CRSP 30-day Treasury-bill yield.

and Brown and Otsuki (1990b) study oil prices in the Japanese equity market.[11]

dG7IP is a weighted average of industrial production growth rates in the G-7 countries, where a measure of relative production shares is used as the weights. Chen, Roll, and Ross (1986) and Shanken and Weinstein (1990) examine the average pricing of U.S. industrial production in the U.S. market. Hamao (1988) examines domestic industrial production risk in the Japanese equity market, and Bodurtha, Cho, and Senbet (1989) estimate the average risk premia for domestic industrial production risk in several countries. No previous study has examined the pricing of global industrial output risks in a conditional asset pricing model.

Our application of beta pricing models requires mimicking portfolios for the risk factors. When a factor is an excess return, the best way to estimate a mimicking portfolio is to use the excess return of the asset directly [Shanken (1992)]. Among our global risk factors, only WDRET is an excess return. We estimate mimicking portfolios for the other variables. One way to estimate mimicking portfolios is to use a large cross section of asset returns [Connor and Korajczyk (1986), Lehmann and Modest (1988), Korajczyk and Viallet (1989)], an approach not available to us. We estimate mimicking portfolios by two common methods, which we describe, and compare them to check the sensitivity of the results.

3.3 The predetermined instruments

We include a list of predetermined instrumental variables similar to previous studies of predictability in country returns. The global information variables, Z_{t-1}, are (1) the yield of a one-month U.S. Treasury bill, (2) the dividend yield of the MSCI world stock market index, (3) a spread between the yields to maturity of 10-year U.S. Treasury bonds and 90-day U.S. Treasury bills, (4) the lagged value of the Eurodollar–U.S. Treasury (TED) spread, (5) the lagged return on the MSCI world market index, and (6) a dummy variable for the month of January. These variables represent readily available, global information that may influence expectations about future equity returns.

For our country-specific instruments, Z^i_{t-1}, we replace the U.S. Treasury bill with a short-term interest rate from the specific country. The world dividend yield is replaced with the dividend yield for the national stock market. The term spread is replaced with a yield spread

[11] We used a spliced series of the posted West Texas intermediate crude and the average U.S. wellhead price, as described in the Appendix. These are not the best indicators of market prices, but they are the best available to us for this period. Futures markets for crude oil did not develop until 1983 (heating oil futures began trading in 1978). Chen, Roll, and Ross (1986) used the energy component of the Producer Price Index. Given the prevalence of long-term oil price contracts over much of the sample, this measure is not likely to better reflect current oil market conditions.

of domestic long-term over short-term, low-risk bonds. The lagged world index return is replaced with the lagged return of the national stock market index. These variables represent information specific to the domestic markets, to the extent that the global aggregates are not sufficient for the local market information. Of course, the distinction between national market information and global information is artificial, because the information sets of investors overlap in more complicated ways. We choose this design on the basis of data availability, parsimony, and empirical tractability. Our data sources and definitions are provided in the Appendix.[12] Table 1 presents monthly summary statistics of the world information variables.

Since the predetermined variables follow previous empirical work, there is a natural concern about predictability uncovered through collective "data snooping" by a series of researchers. Solnik (1993) uses a set of country-specific instruments similar to a subset of ours and argues that step-ahead tests provide evidence that the predictive ability of the instruments is economically significant. Such results increase our confidence that the predictability is an economic phenomenon.[13]

4. Empirical Results

4.1 Preliminary regressions

Table 2 summarizes factor models, where each national equity market return is regressed over time on the eight global risk factors. We use 60-month rolling regressions as a simple way to approximate a factor model with time-varying betas.[14] The right-hand column of Table 2 presents the average of the adjusted R^2's of the rolling regressions for each country. By this measure the global risk factors explain, ex post, 14 to 80 percent of the variance over the 1975–1989 period. In separate regressions, we found that the world market portfolio is by

[12] We studied one other variable, a lagged measure of volatility for the S&P 500 stock market index, constructed from daily returns in the fashion of French, Schwert, and Stambaugh (1987). This variable was also studied by Cutler, Poterba, and Summers (1990). We found that the lagged volatility had no marginal explanatory power for our sample of monthly returns.

[13] The direction of any bias due to data snooping is not clear. On the one hand, the ability of beta pricing models to explain the predictability has not been a criterion for the choice of the lagged information variables. Spurious predictability of the returns, as would be implied by data snooping, should therefore be difficult to "explain" using the models. On the other hand, we choose the global risk variables following previous studies. Most of the previous studies used the factors in unconditional models and did not focus on predictability. Data snooping biases in the risk factors should therefore not be strongly correlated with those in the predetermined variables. However, there must be some correlation between the returns and the factors, which implies that any data-snooping bias may not be independent across the two.

[14] See Braun, Nelson, and Sunier (1991) for evidence that such betas are similar to EGARCH models of conditional betas.

Table 2
The proportion of times that the right-tail probability value was less than 10% for the statistic testing whether the beta coefficients are equal to zero or equal across all countries, based on rolling time series regressions on eight world risk factors. The sample is 1975:2–1989:12 (179 observations)

	Source of risk, j								
Country, i	wdret	dted	dG10fx	G7ui	dG7elt	doil	dG7ip	G7rtb	$\bar{R}^2$
Australia	0.972	0.358	0.229	0.140	0.553	0.112	0.061	0.173	0.406
Austria	0.475	0.279	0.838	0.073	0.162	0.078	0.000	0.151	0.294
Belgium	1.000	0.128	0.777	0.011	0.168	0.050	0.028	0.145	0.444
Canada	1.000	0.061	0.145	0.285	0.553	0.307	0.000	0.196	0.566
Denmark	1.000	0.011	0.274	0.073	0.369	0.017	0.000	0.212	0.233
France	1.000	0.229	0.514	0.184	0.173	0.017	0.341	0.078	0.408
Germany	0.989	0.380	0.804	0.034	0.000	0.291	0.006	0.106	0.358
Hong Kong	0.899	0.291	0.162	0.000	0.156	0.341	0.257	0.017	0.190
Italy	0.894	0.469	0.212	0.112	0.330	0.274	0.056	0.106	0.241
Japan	1.000	0.095	0.503	0.207	0.318	0.201	0.251	0.173	0.430
Netherlands	1.000	0.140	0.570	0.017	0.045	0.101	0.039	0.263	0.555
Norway	0.972	0.034	0.291	0.251	0.430	0.453	0.240	0.179	0.313
Singapore/Malaysia	0.961	0.050	0.391	0.101	0.240	0.469	0.006	0.162	0.324
Spain	0.760	0.486	0.285	0.106	0.307	0.475	0.000	0.084	0.138
Sweden	0.944	0.006	0.458	0.112	0.212	0.073	0.218	0.263	0.249
Switzerland	1.000	0.084	0.626	0.017	0.391	0.173	0.011	0.436	0.530
United Kingdom	1.000	0.039	0.419	0.078	0.106	0.134	0.140	0.156	0.444
United States	1.000	0.034	0.972	0.196	0.324	0.402	0.168	0.050	0.790
$\beta_{ij} = 0$ for $i = 1,\ldots,18$	1.000	0.425	0.978	0.123	0.939	0.670	0.095	0.425	
$\beta_{ij} = \beta_i$ for $i = 1,\ldots,18$	0.832	0.374	0.983	0.112	0.816	0.665	0.061	0.402	

Proportions are based on heteroskedasticity consistent test statistics. The standard error of the fraction rejected (adjusted for overlapping observations) is .143. The $\bar{R}^2$'s are average time-series R^2's adjusted for degrees of freedom. The values in rows 1–18 represent the proportion of times that the probability value for the statistic, testing whether the beta coefficient is zero, was less than 10 percent. The values in rows 19 and 20 represent the proportion of times that the probability value was less than 10 percent for the tests that the beta associated with each source of risk is zero across 18 country portfolios or equal across 18 country portfolios, respectively. The world excess return, wdret, is the arithmetic return on the Morgan Stanley Capital International world equity index (including dividends) minus the Ibbotson Associates one-month bill rate. The change in the Eurodollar–Treasury yield spread, dted, is the first difference of the spread between the 90-day Eurodollar yield and the 90-day Treasury-bill yield (from the Federal Reserve Bulletin). The log change in the G-10 foreign exchange rate, dG10fx, is based on the trade-weighted dollar per foreign exchange rate of 10 industrialized countries (G-7 plus the Netherlands, Belgium, Sweden, and Switzerland). The unexpected inflation for the G-7 countries, G7ui, is derived from a time-series model applied to an aggregate G-7 inflation rate where the (varying) weights in the aggregate are determined by country weights in total G-7 gross domestic product. The change in long-term expected G-7 inflation, dG7elt, is a result of projecting the four-year moving average of G-7 inflation on the set of lagged instrumental variables. The change in the price of oil, doil, is the log change in the average U.S. dollar price per barrel at the wellhead from 1974 to 1989 and the posted West Texas Intermediate price from 1969 to 1973. dG7ip is the change in G-7 industrial production. The G-7 real interest rate, G7rtb, is calculated by aggregating individual countries' short-term interest rates minus inflation rates using (varying) weights based on quarterly shares in G-7 gross domestic product.

far the most important factor in this sense. It alone explains 5 to 71 percent of the ex post variance, depending on the country.

We use the regressions in Table 2 to delete a subset of our initial risk factors from the subsequent analysis. If there is a variable whose betas are not different across the countries or different from zero, then that variable will not be priced. The bottom row of Table 2 presents tests of the hypothesis that the betas for each global risk variable are zero in all of the countries and of the hypothesis that they are equal across the countries. The first number is the fraction of the 60-month regressions in which a Wald test rejects the hypothesis that the betas are equal to zero for all countries, using a 10 percent significance level. The second number is the fraction of the 60-month regressions in which the test rejects the hypothesis that the betas are jointly equal across the countries but not necessarily equal to zero. If the null hypotheses are true, then we expect to reject in 10 percent of the cases. We calculate an approximate standard error for the fraction rejected, given 179 trials, as equal to 0.143.[15] To include a risk variable in our model, we require that the fraction rejected in Table 2 be at least two standard errors above the expected fraction of 0.10. This leads us to drop the variables G7UI, G7IP, and dTED.

Table 3 summarizes regressions that use the lagged world information variables to predict the excess country returns. The apparent predictable variation measured by the adjusted R^2's ranges across the countries, from virtually zero to over 10 percent. Table 4 shows the marginal explanatory power of additional lagged variables. The first three columns report R^2's for regressions of the returns on the global information variables, augmented with either the local versions of the information variables or with the lagged values of the rolling regression betas from Table 2.[16] The fourth column of the table presents F-tests for the incremental explanatory power of the local information variables, given the global variables. They are significant at the 5 percent level for 7 of the 18 countries, which provides some evidence that local information is important.[17] A joint heteroskedasticity-consistent Wald test for all 18 countries produces a test statistic of 180.1. The right-tail p value from the χ^2 distribution is less than .001.

The beta pricing model assumes that expected returns are determined by conditional betas, which are country specific, and by expected risk premia, which are global measures. If conditional betas

[15] Each joint test based on one rolling regression, using a test of size α, is viewed as generating a binomial trial. If these trials were independent, then the variance of the fraction rejected in n trials is approximately given by $\alpha(1 - \alpha)/n$. But the rolling regressions are not independent, because of the overlapping data. The variance of the fraction rejected, p, is adjusted for the overlap as follows. Assuming that the underlying data are independent across the months, then the autocovariance of the p_t's that is induced by overlapping data is $\mathrm{Cov}(p_t, p_{t-j}) = [(60 - j)/60]\mathrm{Var}(p_t)$ if $j < 60$, and zero otherwise. We construct the covariance matrix of the vector of the p_t's and find the standard error of p for 179 trials, using the 60-month regression approach and $\alpha = .10$, to be .143.

[16] The rolling regression betas for time $t - 1$ are not strictly predetermined to the extent that publication lags and data revisions imply that the economic series were not actually available to market participants at time $t - 1$. We therefore estimated these regressions with betas lagged back two months. Also, for the first 60 months of the sample, the betas are constant.

[17] This is so in spite of the fact that in some of the more regulated economies (e.g., Sweden) the interest rates used as instruments are not competitively determined market rates.

Table 3
Regression of asset returns on world instrumental variables. The sample is 1970:2–1989:12 (239 observations)

Country	Instrument							
	intercept	wr_{t-1}	jan_t	$divwrd_{t-1}$	ted_{t-1}	$term_{t-1}$	$tb1_t$	$\bar{R}^2$
Australia	0.040	0.190	0.027	12.985	−18.619	−11.911	−7.482	0.068
	(0.025)	(0.142)	(0.020)	(8.826)	(7.642)	(5.642)	(3.566)	
Austria	0.066	0.069	−0.033	−9.252	−9.797	−3.182	−1.850	0.074
	(0.021)	(0.084)	(0.011)	(5.454)	(4.055)	(4.171)	(1.974)	
Belgium	0.039	−0.042	0.022	1.669	−9.274	0.785	−4.481	0.042
	(0.019)	(0.090)	(0.012)	(5.416)	(5.181)	(3.596)	(2.278)	
Canada	0.030	0.077	0.019	11.999	−9.768	−10.176	−7.226	0.064
	(0.020)	(0.086)	(0.016)	(5.615)	(5.898)	(4.143)	(2.443)	
Denmark	0.039	−0.196	0.020	−3.129	−9.065	1.605	−2.041	0.032
	(0.020)	(0.087)	(0.015)	(5.213)	(4.615)	(3.387)	(2.047)	
France	0.027	−0.001	0.021	9.103	−16.422	−3.642	−4.442	0.026
	(0.023)	(0.099)	(0.019)	(7.402)	(7.146)	(4.829)	(5.061)	
Germany	0.023	0.073	−0.005	5.338	−9.273	−1.842	−3.473	0.013
	(0.023)	(0.112)	(0.015)	(5.754)	(5.629)	(4.057)	(2.224)	
Hong Kong	0.069	0.127	0.074	9.510	−31.837	−10.652	−6.656	0.057
	(0.046)	(0.164)	(0.026)	(11.681)	(10.542)	(7.544)	(4.587)	
Italy	0.009	0.124	0.030	0.889	−12.733	−1.112	0.726	0.011
	(0.026)	(0.112)	(0.018)	(6.864)	(6.281)	(4.469)	(2.967)	
Japan	0.038	0.206	0.008	0.874	−7.769	1.399	−3.582	0.063
	(0.022)	(0.110)	(0.012)	(5.926)	(5.968)	(3.876)	(2.309)	
Netherlands	0.021	−0.030	0.031	10.720	−16.574	−2.796	−4.373	0.076
	(0.019)	(0.091)	(0.014)	(5.186)	(5.242)	(3.742)	(2.318)	

Table 3
Continued

Country	intercept	wr_{t-1}	jan_t	$divwrd_{t-1}$	ted_{t-1}	$term_{t-1}$	$tb1_t$	$\bar{R}^2$
	Instrument							
Norway	0.070	−0.041	0.048	−5.766	−13.773	−11.186	−2.801	0.032
	(0.027)	(0.154)	(0.018)	(7.863)	(8.189)	(5.493)	(3.350)	
Singapore/Malaysia	0.072	0.051	0.081	11.648	−14.650	−17.227	−11.381	0.105
	(0.030)	(0.152)	(0.027)	(9.406)	(8.083)	(5.409)	(3.520)	
Spain	0.048	0.117	0.019	−17.163	−1.116	−0.172	1.586	0.030
	(0.019)	(0.099)	(0.016)	(6.496)	(5.583)	(3.718)	(2.313)	
Sweden	0.018	0.154	0.025	−0.861	−3.416	−2.423	−0.318	0.005
	(0.020)	(0.114)	(0.015)	(5.857)	(5.782)	(4.144)	(2.524)	
Switzerland	0.025	−0.073	0.012	9.226	−12.698	−2.580	−5.306	0.039
	(0.019)	(0.097)	(0.016)	(5.699)	(5.487)	(3.735)	(2.262)	
United Kingdom	0.009	−0.023	0.050	20.683	−19.979	−5.257	−6.881	0.074
	(0.022)	(0.142)	(0.029)	(8.430)	(7.038)	(5.153)	(3.297)	
United States	0.006	−0.041	0.024	11.497	−11.277	−2.537	−4.134	0.061
	(0.015)	(0.086)	(0.013)	(5.231)	(4.601)	(3.309)	(1.948)	
World	0.017	0.033	0.022	9.140	−12.056	−2.877	−4.175	0.095
	(0.014)	(0.065)	(0.011)	(4.265)	(3.869)	(2.848)	(1.682)	

$\bar{R}^2$'s are adjusted for degrees of freedom. Standard errors in parentheses are heteroskedasticity consistent. The instrumental variables are the following: a constant, a dummy variable for the month of January, the lagged excess return of the Morgan Stanley Capital International world index minus the CRSP 30-day bill, the dividend yield (over the past 12 months) on the Morgan Stanley Capital International world equity index, the difference between the 90-day Eurodollar rate and the CRSP three-month Treasury-bill yield, the difference between the U.S. 10-year Treasury-bond yield and the CRSP three-month bill yield, and the CRSP 30-day Treasury-bill yield.

Table 4
The incremental explanatory power of local information variables in predicting 18 countries' equity returns. The sample is 1970:2–1989:12 (239 observations)

Country	R^2 world	R^2 world + local	R^2 world + 5 betas	World + local F-test: exclude local p value	World + local + 1 beta F-test: exclude local p value	World + local + 5 betas F-test: exclude local p value
Australia	0.091	0.133	0.143	2.642 0.035	1.915 0.109	1.861 0.118
Austria	0.097	0.118	0.139	1.368 0.246	1.854 0.119	1.885 0.114
Belgium	0.067	0.089	0.142	1.385 0.240	0.552 0.698	1.398 0.236
Canada	0.087	0.128	0.182	2.563 0.039	2.566 0.039	1.798 0.130
Denmark	0.056	0.079	0.090	1.402 0.234	1.550 0.189	1.403 0.234
France	0.051	0.056	0.104	0.320 0.864	0.579 0.678	1.286 0.276
Germany	0.037	0.041	0.067	0.200 0.938	0.734 0.569	0.320 0.864
Hong Kong	0.081	0.081	0.103	0.013 0.987	0.027 0.974	0.475 0.623
Italy	0.036	0.061	0.085	1.504 0.202	1.413 0.230	0.991 0.413
Japan	0.086	0.092	0.112	0.349 0.844	0.427 0.789	1.001 0.408
Netherlands	0.099	0.133	0.172	2.202 0.070	1.639 0.165	2.979 0.020
Norway	0.056	0.102	0.101	2.822 0.026	2.860 0.024	1.935 0.105
Singapore/Malaysia	0.128	0.182	0.147	3.626 0.007	3.601 0.007	5.021 0.001
Spain	0.055	0.085	0.086	1.880 0.115	1.804 0.129	2.775 0.028
Sweden	0.030	0.068	0.056	2.269 0.063	2.780 0.028	3.502 0.008
Switzerland	0.063	0.103	0.103	2.445 0.047	2.048 0.089	2.116 0.080
United Kingdom	0.097	0.202	0.152	6.702 0.000	4.153 0.003	4.504 0.002
United States	0.084	0.148	0.202	4.061 0.003	2.078 0.084	0.576 0.680

The "world" regressions are time-series regressions of each country's excess return on the set of lagged world instruments. The "world" instrumental variables are the following: a constant, a dummy variable for the month of January, the lagged Morgan Stanley Capital International world return minus the CRSP 30-day bill, the dividend yield on the Morgan Stanley Capital International world equity index, the difference between the 90-day Eurodollar rate and the CRSP three-month bill yield, the difference between the U.S. 10-year Treasury-bond yield and the CRSP three-month bill yield, and the CRSP 30-day Treasury-bill yield.

The "world + local" regressions include the additional instrumental variables, which are the lagged excess return of the local equity market, the dividend yield for the local equity market, the difference between the long-term and short-term interest rates in the country and the local short-term interest rate.

The "world + local + 1 beta" regressions use the same regressors as the "world + local" are approximately constant, then predictable variation should be captured by global variables. If time variation in the betas is important, then local information may enter through the betas. Table 4 shows that the lagged betas deliver an increase in the explanatory power of the regressions. Their incremental forecast power is comparable to the local versions of the information variables. This suggests that the betas may capture information about the future returns, similar to the local variables.

In the fifth and sixth columns of Table 4 the incremental explanatory power of the local information variables is illustrated, in regressions which include both the lagged betas and the global variables. In the fifth column the lagged beta for each country with respect to the world market index is used. The sixth column introduces betas for all five of the global risk factors. An F-test examines the marginal explanatory power of the local variables. Their marginal impact is reduced, although not completely eliminated, when the lagged betas are included. Overall, the regressions provide some support for our specification of the empirical asset pricing model, in which we assume that local information variables enter through the betas. Of course, the regressions are only suggestive of how such a model will actually perform.

4.2 Explaining predictability using global economic risk factors

Table 5 addresses the extent to which the models can explain predictable variation in the country returns. The table reports for each country the average pricing error, α_i, its standard error, the variance ratios and their standard errors, and some analysis of the predictability that remains in the model pricing errors.

Panel A of Table 5 summarizes the single-factor model, in which the world market portfolio is the factor. The average pricing error is smaller than the average excess return for all countries and is more than two standard errors from zero in only three cases. However, the standard errors are large. Regressing the pricing errors over time on the lagged global information variables, the adjusted R^2's are negative for 10 of the 18 countries. Regressing the pricing errors on the local

---

←

regressions, in addition to the beta coefficient from a regression of the asset return on the excess world market return from $t-62$ to $t-2$.

The "world + local + 5 betas" regressions use the same regressors as the "world + local" regressions, in addition to the beta coefficients from a regression of the asset return on five world risk factors from $t-62$ to $t-2$. The risk factors are the excess world market return, the log change in a U.S. dollar versus G-10 currency index, the change in long-term expected G-7 inflation, the log change in the price of oil and the G-7 real interest rate.

Table 5
A decomposition of the predictable variation in international equity returns

Country	Average return	Average pricing error α_i	VR1	VR2	χ^2 constant betas	$\bar{R}^2$ pricing errors on Z	$\bar{R}^2$ pricing errors on Z^i
		A: 1-factor model[1]					
Australia	0.468	0.251	0.523	0.267	5.959	−0.004	0.003
		(0.477)	(0.240)	(0.134)	[0.428]		
Austria	0.756	0.633	0.091	0.809	2.703	0.055	0.070
		(0.361)	(0.068)	(0.164)	[0.845]		
Belgium	0.897	0.646	0.644	0.310	12.287	−0.007	−0.020
		(0.316)	(0.387)	(0.179)	[0.056]		
Canada	0.451	0.157	1.169	0.402	17.096	0.003	0.000
		(0.284)	(0.496)	(0.236)	[0.009]		
Denmark	0.816	0.473	0.494	0.708	1.790	0.015	−0.009
		(0.330)	(0.321)	(0.337)	[0.938]		
France	0.729	0.233	0.925	0.079	3.798	−0.023	−0.020
		(0.379)	(0.520)	(0.124)	[0.704]		
Germany	0.651	0.300	0.841	0.151	2.930	−0.021	−0.020
		(0.325)	(0.529)	(0.182)	[0.818]		
Hong Kong	1.630	1.397	0.274	0.334	7.143	0.001	0.004
		(0.820)	(0.170)	(0.149)	[0.308]		
Italy	0.296	−0.053	0.790	0.906	18.022	−0.000	0.023
		(0.472)	(0.545)	(0.642)	[0.006]		
Japan	1.313	0.784	0.510	0.386	16.769	0.006	−0.011
		(0.306)	(0.224)	(0.204)	[0.010]		
Netherlands	0.830	0.444	0.696	0.097	6.709	−0.018	0.011
		(0.257)	(0.242)	(0.074)	[0.349]		
Norway	0.932	0.889	0.726	0.647	16.299	0.004	0.006
		(0.493)	(0.434)	(0.344)	[0.012]		
Singapore/Malaysia	1.114	0.646	0.407	0.389	18.910	0.019	0.032
		(0.519)	(0.192)	(0.166)	[0.004]		
Spain	0.361	0.105	0.464	1.118	21.545	0.029	−0.004
		(0.390)	(0.359)	(0.443)	[0.001]		
Sweden	0.964	0.810	1.169	0.970	8.733	−0.001	0.007
		(0.384)	(1.070)	(0.831)	[0.189]		
Switzerland	0.548	0.224	1.025	0.212	6.373	−0.016	−0.013
		(0.286)	(0.499)	(0.177)	[0.383]		
United Kingdom	0.761	0.060	0.708	0.252	13.573	−0.012	0.066
		(0.403)	(0.311)	(0.147)	[0.035]		
United States	0.380	−0.052	1.218	0.179	20.141	−0.016	0.005
		(0.183)	(0.403)	(0.133)	[0.003]		
Average		0.442	0.704	0.456			
		B: 2-factor model with BGL mimicking portfolio[2]					
Australia	0.468	−0.008	0.602	0.402	24.250	0.001	0.011
		(0.482)	(0.263)	(0.175)	[0.019]		
Austria	0.756	−0.110	0.465	0.394	19.094	0.012	0.014
		(0.250)	(0.181)	(0.143)	[0.086]		
Belgium	0.897	0.293	1.074	0.227	20.974	−0.013	−0.013
		(0.250)	(0.428)	(0.179)	[0.051]		

[1] The excess world market return is the single factor.

[2] In this two-factor model, the second factor is the change in the log of the U.S. versus G-10 countries exchange rate index. A mimicking portfolio is formed for the exchange rate index using the technique of Breeden, Gibbons, and Litzenberger (1989).

Table 5
Continued

Country	Average return	Average pricing error α_i	VR1	VR2	χ^2 constant betas	$\bar{R}^2$ pricing errors on Z	$\bar{R}^2$ pricing errors on Z^f
Canada	0.451	0.247	1.174	0.361	23.581	−0.003	−0.003
		(0.268)	(0.445)	(0.189)	[0.023]		
Denmark	0.816	0.046	0.868	0.504	10.672	0.002	−0.013
		(0.284)	(0.476)	(0.300)	[0.557]		
France	0.729	−0.185	1.266	0.290	10.382	−0.014	−0.013
		(0.351)	(0.619)	(0.275)	[0.582]		
Germany	0.651	0.019	1.250	0.237	16.373	−0.020	−0.017
		(0.291)	(0.640)	(0.257)	[0.175]		
Hong Kong	1.630	0.647	0.428	0.419	13.050	0.002	0.006
		(0.789)	(0.235)	(0.195)	[0.365]		
Italy	0.296	−0.492	1.228	1.086	21.511	0.001	0.024
		(0.444)	(0.775)	(0.827)	[0.043]		
Japan	1.313	0.438	0.751	0.277	41.392	−0.009	−0.022
		(0.255)	(0.235)	(0.157)	[0.000]		
Netherlands	0.830	0.284	0.793	0.148	17.597	−0.012	0.018
		(0.251)	(0.268)	(0.103)	[0.128]		
Norway	0.932	0.779	0.856	0.727	24.087	0.005	0.005
		(0.490)	(0.500)	(0.381)	[0.020]		
Singapore/Malaysia	1.114	0.663	0.418	0.395	26.163	0.019	0.030
		(0.542)	(0.190)	(0.173)	[0.010]		
Spain	0.361	−0.310	0.704	0.763	29.886	0.009	−0.008
		(0.376)	(0.471)	(0.381)	[0.003]		
Sweden	0.964	0.636	1.453	1.173	12.164	0.003	0.014
		(0.389)	(1.275)	(1.020)	[0.433]		
Switzerland	0.548	−0.091	1.357	0.405	21.769	−0.003	0.007
		(0.244)	(0.520)	(0.247)	[0.040]		
United Kingdom	0.761	−0.144	0.768	0.367	17.122	−0.006	0.069
		(0.417)	(0.332)	(0.264)	[0.145]		
United States	0.380	0.116	1.252	0.151	23.457	−0.017	0.001
		(0.151)	(0.260)	(0.084)	[0.024]		
Average		0.157	0.928	0.463			
		C: 2-factor model with FM mimicking portfolio[a]					
Australia	0.468	0.410	0.539	0.308	19.954	−0.002	0.006
		(0.469)	(0.247)	(0.143)	[0.068]		
Austria	0.756	0.455	0.135	0.784	29.191	0.049	0.060
		(0.337)	(0.082)	(0.190)	[0.004]		
Belgium	0.897	0.450	0.672	0.337	20.959	−0.005	−0.020
		(0.300)	(0.387)	(0.202)	[0.051]		
Canada	0.451	0.351	1.186	0.510	49.514	0.008	0.004
		(0.275)	(0.536)	(0.296)	[0.000]		
Denmark	0.816	0.421	0.544	0.712	5.514	0.015	−0.007
		(0.326)	(0.340)	(0.357)	[0.939]		
France	0.729	0.144	0.923	0.156	12.567	−0.021	−0.017
		(0.382)	(0.549)	(0.209)	[0.401]		
Germany	0.651	−0.061	1.258	0.449	37.869	−0.016	−0.015
		(0.315)	(0.760)	(0.410)	[0.000]		
Hong Kong	1.630	1.042	0.748	0.224	38.665	−0.010	−0.008
		(0.548)	(0.298)	(0.129)	[0.000]		

[a] In this two-factor model, the mimicking portfolio for the exchange rate index is formed using the method of Fama and MacBeth (1973).

Table 5
Continued

Country	Average return	Average pricing error α_i	VR1	VR2	χ^2 constant betas	$\bar{R}^2$ pricing errors on Z	$\bar{R}^2$ pricing errors on Z^i
Italy	0.296	−0.227	0.921	0.954	38.695	0.001	0.022
		(0.479)	(0.636)	(0.688)	[0.000]		
Japan	1.313	0.816	0.645	0.381	48.863	0.005	−0.014
		(0.300)	(0.261)	(0.223)	[0.000]		
Netherlands	0.830	0.339	0.746	0.061	30.669	−0.020	0.013
		(0.248)	(0.236)	(0.058)	[0.002]		
Norway	0.932	0.748	0.724	0.588	24.670	0.000	0.003
		(0.497)	(0.433)	(0.332)	[0.016]		
Singapore/Malaysia	1.114	0.922	0.457	0.593	78.739	0.045	0.059
		(0.497)	(0.318)	(0.225)	[0.000]		
Spain	0.361	0.105	0.427	1.101	22.996	0.027	−0.007
		(0.388)	(0.321)	(0.422)	[0.028]		
Sweden	0.964	0.814	1.308	0.994	12.792	−0.002	0.006
		(0.388)	(1.160)	(0.833)	[0.384]		
Switzerland	0.548	0.029	1.266	0.269	28.797	−0.011	−0.010
		(0.292)	(0.578)	(0.253)	[0.004]		
United Kingdom	0.761	0.148	0.753	0.432	75.157	0.006	0.089
		(0.421)	(0.358)	(0.259)	[0.000]		
United States	0.380	0.028	1.305	0.258	34.513	−0.013	0.008
		(0.195)	(0.447)	(0.178)	[0.001]		
Average		0.382	0.809	0.506			
	D: 5-factor model with BGL mimicking portfolios[a]						
Australia	0.468	0.115	0.951	0.458	52.191	−0.009	0.009
		(0.442)	(0.406)	(0.300)	[0.007]		
Austria	0.756	−0.114	0.610	0.292	68.107	−0.005	0.002
		(0.212)	(0.184)	(0.123)	[0.000]		
Belgium	0.897	0.418	1.172	0.352	48.992	−0.015	−0.010
		(0.258)	(0.481)	(0.248)	[0.016]		
Canada	0.451	0.130	1.489	0.388	48.977	−0.005	−0.001
		(0.296)	(0.476)	(0.207)	[0.016]		
Denmark	0.816	−0.032	1.198	0.667	66.355	0.010	−0.012
		(0.274)	(0.572)	(0.419)	[0.000]		
France	0.729	0.122	1.405	0.505	72.931	−0.009	−0.007
		(0.356)	(0.715)	(0.354)	[0.000]		
Germany	0.651	−0.013	1.432	0.224	79.956	−0.023	−0.017
		(0.283)	(0.708)	(0.222)	[0.000]		
Hong Kong	1.630	0.560	0.701	0.650	63.312	0.003	0.003
		(0.796)	(0.378)	(0.318)	[0.000]		
Italy	0.296	−0.234	1.122	1.262	72.547	0.012	0.023
		(0.392)	(0.687)	(0.992)	[0.000]		
Japan	1.313	0.390	0.744	0.360	77.714	−0.006	−0.020
		(0.244)	(0.257)	(0.196)	[0.000]		
Netherlands	0.830	0.195	0.975	0.204	29.562	−0.012	0.013
		(0.250)	(0.359)	(0.148)	[0.488]		

[a] The risk factors are the excess world market return, the log change in a U.S. dollar versus G-10 currency index, the change in long-term expected G-7 inflation minus the Treasury-bill return, the change in the price of oil minus the Treasury-bill return and the G-7 real interest rate. Mimicking portfolios for the last four factors are formed with the technique of Breeden, Gibbons, and Litzenberger (1989).

Table 5
Continued

Country	Average return	Average pricing error α_i	VR1	VR2	χ^2 constant betas	$\bar{R}^2$ pricing errors on Z	$\bar{R}^2$ pricing errors on Z^i
Norway	0.932	0.830	1.736	0.718	92.306	−0.009	0.007
		(0.426)	(0.802)	(0.503)	[0.000]		
Singapore/Malaysia	1.114	0.115	0.718	0.409	70.742	0.006	0.023
		(0.488)	(0.279)	(0.218)	[0.000]		
Spain	0.361	0.013	1.149	0.545	50.345	−0.006	−0.010
		(0.321)	(0.674)	(0.367)	[0.011]		
Sweden	0.964	0.703	1.276	1.141	50.396	−0.003	0.004
		(0.387)	(1.094)	(0.976)	[0.011]		
Switzerland	0.548	−0.095	1.619	0.347	56.521	−0.012	0.009
		(0.226)	(0.515)	(0.282)	[0.002]		
United Kingdom	0.761	0.036	1.349	0.955	55.936	0.004	0.045
		(0.413)	(0.941)	(0.944)	[0.003]		
United States	0.380	0.124	1.507	0.273	100.423	−0.015	0.006
		(0.168)	(0.349)	(0.187)	[0.000]		
Average		0.181	1.175	0.542			
	E: 5-factor model with FM mimicking portfolios[5]						
Australia	0.468	−0.441	1.837	1.427	153.978	0.028	0.012
		(0.554)	(1.640)	(1.515)	[0.000]		
Austria	0.756	0.168	0.301	0.651	69.248	0.028	0.029
		(0.263)	(0.155)	(0.203)	0.001		
Belgium	0.897	0.382	0.674	0.457	101.254	−0.001	−0.009
		(0.296)	(0.369)	(0.294)	[0.000]		
Canada	0.451	0.491	2.103	0.878	134.101	0.018	0.024
		(0.333)	(0.849)	(0.467)	[0.000]		
Denmark	0.816	0.559	0.471	0.770	102.177	0.009	−0.008
		(0.359)	(0.303)	(0.386)	[0.000]		
France	0.729	0.150	0.958	0.702	110.318	−0.014	−0.014
		(0.419)	(0.577)	(0.434)	[0.000]		
Germany	0.651	−0.108	1.458	0.578	87.854	−0.019	−0.015
		(0.309)	(0.810)	(0.510)	[0.000]		
Hong Kong	1.630	0.336	1.629	0.510	129.590	−0.020	−0.018
		(0.534)	(0.779)	(0.453)	[0.000]		
Italy	0.296	−0.122	1.957	2.059	152.785	−0.009	−0.001
		(0.419)	(1.375)	(1.407)	[0.000]		
Japan	1.313	1.136	1.038	0.640	177.925	−0.004	−0.012
		(0.311)	(0.467)	(0.401)	[0.000]		
Netherlands	0.830	0.308	0.799	0.171	102.242	−0.017	0.009
		(0.268)	(0.330)	(0.108)	[0.000]		
Norway	0.932	0.610	2.185	1.357	166.421	−0.007	0.006
		(0.551)	(1.289)	(0.972)	[0.000]		
Singapore/Malaysia	1.114	0.244	0.873	0.693	101.276	0.017	0.022
		(0.595)	(0.403)	(0.377)	[0.000]		
Spain	0.361	0.367	0.874	0.824	97.672	0.001	−0.016
		(0.352)	(0.584)	(0.480)	[0.000]		
Sweden	0.964	0.684	1.252	1.374	81.206	−0.004	−0.008
		(0.427)	(1.101)	(1.117)	[0.000]		
Switzerland	0.548	0.163	1.142	0.444	76.119	−0.012	−0.012
		(0.328)	(0.552)	(0.339)	[0.000]		

[5] In this model, Fama and MacBeth (1973) mimicking portfolios are used.

Table 5
Continued

Country	Average return	Average pricing error α_i	VR1	VR2	χ^2 constant betas	$\bar{R}^2$ pricing errors on Z	$\bar{R}^2$ pricing errors on Z^i
United Kingdom	0.761	−0.177	1.856	1.311	104.485	−0.014	0.014
		(0.391)	(2.089)	(1.916)	[0.000]		
United States	0.380	−0.022	1.510	0.415	101.549	−0.016	0.010
		(0.258)	(0.641)	(0.338)	[0.000]		
Average		0.262	1.293	0.848			

The following system is estimated for each asset i:

Disturbance	Orthogonal to
$u1_{it} = (r_{it} - Z'_{t-1}\delta_i)$	Z_{t-1}
$u2_t = (F'_t - Z'_{t-1}\gamma)'$	Z_{t-1}
$u3_{it} = [(u2_t u2'_t)(\kappa'_i Z^i_{t-1}) - (F_t u1'_{it})]$	Z^i_{t-1}
$u4_{it} = (Z'_{t-1}\delta_i - \theta_i)$	1
$u5_{it} = (Z'_{t-1}\gamma)(\kappa'_i Z^i_{t-1}) - \theta_i + \alpha_i$	1
$u6_{it} = (u4^2_{it})\text{VR1}_i - u5^2_{it}$	1

where r_i represents the return on asset i, Z is the common world predetermined information, Z^i is the local information, δ are coefficients from a linear projection of the asset returns on the information, $\kappa'_i Z^i$ are the fitted conditional betas, F are the factor returns, γ are the coefficients from a linear projection of the factor returns on the information, θ_i is the mean asset return, α_i is the difference between the mean asset return and the model-fitted mean asset return, and VR1 is the ratio of the variance of the asset pricing model's fitted values to the variance of the statistical model's fitted values. In a separate estimation, the last two equations are replaced with

$u5_{it} = (Z'_{t-1}\delta_i - Z'_{t-1}\gamma\kappa'_i Z^i_{t-1}) - \theta_{2i}$	1
$u6_{it} = (u4^2_{it})\text{VR2}_i - u5^2_{it}$	1

where VR2 is the ratio of the variance of the model's unexplained expected returns to the variance of the statistical model's fitted values. χ^2 is the Wald test for the hypothesis that the conditional betas are constant over time (κ_i's zero except for the intercept). The $\bar{R}^2$ are for regressions of the model pricing errors, defined as $\epsilon_{it} = r_{it} - Z'_{t-1}\gamma\kappa'_i Z^i_{t-1}$ on the lagged instruments. The sample is 1970:2–1989:12 (239 observations).

versions of the information variables, 7 of the 18 adjusted R^2's are negative. The largest of the 36 adjusted R^2's for the pricing errors are 7 percent (Austria) and 6.6 percent (United Kingdom). The variance ratios VR1 are larger than the VR2's in 13 of the 18 countries, which suggests that the model captures more of the predictability than it leaves in the residuals. The average VR1 is 0.704, and the average VR2 is 0.456. However, the large standard errors preclude precise inferences.[18]

Panels B and C of Table 5 show results for two-factor models, in which the exchange risk variable is a second factor. Since the exchange risk variable is not an excess return, we construct a mimicking port-

[18] Joint tests would be preferred in order to account for correlation across the countries. However, with 80 orthogonality conditions per country in the five-factor model and only 236 time-series observations, we are unable to provide joint tests. We leave it to the reader to make these judgments informally.

folio.[19] Mimicking portfolios for the factors are constructed in two ways. The first is a variation of the approach in Breeden, Gibbons, and Litzenberger (BGL, 1989). BGL construct a maximum correlation portfolio by regressing the factor over time on the test assets. The slope coefficients are proportional to the portfolio weights. Our modification of the BGL approach is to include the world information variables in the regressions, thereby approximating a maximum conditional correlation portfolio. The results using this portfolio for the exchange risk factor are reported in panel B.

The BGL approach assumes that the mimicking portfolio weights are fixed parameters over the sample, which is a potential weakness. An alternative method uses cross-sectional regressions of the country returns each month on the lagged, rolling regression betas for dG10FX, which is similar to Fama and MacBeth (1973). The cross-sectional regression coefficients are the excess returns on a mimicking portfolio and are used in panel C. This approach allows the mimicking portfolio weights to vary month by month. The details of the approach are described in the Appendix.[20]

The results for the two-factor model show modest improvement over the single-factor model. The average pricing error α_i is reduced, relative to the single-factor model for 11 of the countries. The estimates of the α_i are more than two standard errors from zero in only 3 of 36 cases. The adjusted R^2's from regressing the pricing errors on the lagged variables present a similar pattern to the one-factor model. Twelve to 17 of the 18 VR1's are larger than the VR2's.

Panels D and E of Table 5 summarize the five-factor models. The world excess return WDRET is used directly as a factor, while mimicking portfolios are used for the variables G7RTB, dOIL, dG7ELT, and dG10FX. In panel D the modified BGL mimicking portfolios are used, and in panel E the Fama–MacBeth portfolios are used. The statistics point to a fairly dramatic improvement in the fit of the model relative to the single-factor models. Only 1 of the 36 average pricing errors, α_i, is more than two standard errors from zero. Thirty-one of the 36 VR1's are larger than the corresponding VR2's, and only 3 of 36 VR2's are more than two standard errors greater than zero. The regressions of the model residuals on the lagged world and on the

[19] The variable dG10FX approximates an excess return when the trade weights are known and a trade-weighted combination of foreign currency deposit rates is close to the U.S. bill rate. We therefore estimated a two-factor model in which we used dG10FX directly instead of a mimicking portfolio. The results were broadly similar.

[20] Both the BGL and the cross-sectional regression approach have the disadvantage that the estimation of system (6) does not account for the fact that mimicking portfolios were formed in a previous step. See Wheatley (1989) for an analysis of this problem. We experimented with GMM systems in which mimicking portfolio weights were estimated simultaneously with the other model parameters, but we found the systems empirically intractable.

lagged local market variables show little evidence of remaining predictability. These results show that the five-factor models can capture much of the predictable variation in most of the country returns. There are a few countries, however, where the models have difficulties. Austria and Italy are two cases where there is apparent predictability that the models do not capture well.[21]

Finally, Table 5 shows a Wald test of the hypothesis that the conditional betas may be regarded as constant over time, where the alternative is the linear model.[22] The test rejects constant betas in the one-factor model at the 5 percent level, for 8 of the 18 countries. For the two-beta models, the tests reject in 9 to 13 countries. In the five-beta models, constant betas are rejected in all but 1 of the 36 cases. Therefore, time variation in conditional betas appears to be statistically significant in our model.[23]

4.3 The importance of changing betas

Although statistical tests reject the hypothesis that the conditional betas are constant, this does not provide a measure of how important movements in the betas are for explaining return predictability. We investigate this question by estimating the following decomposition:

$$\begin{aligned}\mathrm{Var}\{E(\beta'\lambda \mid Z)\} = {} & E(\beta)'\mathrm{Var}\{E(\lambda \mid Z)\}E(\beta) \\ & + E(\lambda)'\mathrm{Var}\{\beta(Z)\}E(\lambda) + \phi. \end{aligned} \tag{9}$$

The left-hand side of (9) is the predictable variation that is captured by the model. The first term on the right-hand side is the part attributed to movements in expected risk premia. The second term is the part attributed to time variation in the betas. The term ϕ represents interaction effects that arise because the expected risk premia and betas may be correlated through time. Ferson and Harvey (1991) used a similar decomposition in domestic data, which they estimated with a multistep regression procedure.

We employ the GMM to consistently estimate the decomposition (9). We start with the first three equations of system (6). Two additional equations are added to the system to identify parameters for the unconditional means of the betas and of the risk premia. A third equation identifies the unconditional means of the products of the

[21] We estimated versions of the models in which the variance ratios used projections of returns on both the global and the local information variables in the denominator. Not surprisingly, these larger models produced less precise results. There was evidence that a five-factor model performs better than a one- or two-factor model, but the overall performance of the models was worse.

[22] The statistic is a quadratic form in the coefficients that model the betas as functions of the lagged Z_{t-1}^{i}, where the matrix is the inverse of a heteroskedasticity-consistent estimate of their covariance matrix. The statistic is asymptotically a χ^2 variable under the null hypothesis.

[23] One should be cautious about interpreting the Wald tests, because the number of restrictions is large relative to the sample size. The tests are also likely to be correlated across the countries.

betas and risk premia. The variances are constructed as the means of the products minus the products of the means. A fourth equation defines a parameter equal to the ratio of the first term on the right-hand side of (9) to the left-hand side of (9). This is the fraction of the model predicted variance of return that is attributed to variation in expected risk premia. A complementary ratio, calculated in a separate estimation, measures the fraction that is attributed to variation in betas. The notes to Table 6 display the equations in detail.

Table 6 shows that there is only a small direct contribution of time-varying betas to the model variation in expected country returns. Most of the predictable variation that is captured by our model is attributed to movements in the global risk premia.[24] There are, however, sizable interaction effects. The sum of the direct beta and risk premia effects is less than 1.0 for most of the countries. This implies that the betas and the expected risk premia are positively related for those countries, which has an interesting interpretation. If the expected risk premia are countercyclical in the aggregate, the estimates suggest that the sensitivity to the global risk factors are higher for most of the countries in a weak global economy.[25] Apparent exceptions are Japan and Germany, where the point estimates suggest that the betas are lower in a weak global economy.

4.4 Diagnostics

Table 7 shows some regressions to further check the specification of the models. In the first two panels, the pricing errors for each country are regressed on dummy variables, indicating one of three currency market regimes. They are the 1970:2–1973:2 period of fixed exchange rates, the "dirty float" period from 1973:3 to 1980:12, and the subsequent period of more flexible exchange rates. Of course, the use of three fixed regimes for each country is a dramatic simplification, but it could still be informative to see if the average pricing errors are significantly different in these three regimes. The first panel shows results for the one-factor model, in which the coefficient on a dummy variable exceeds two standard errors for five of the countries. The second panel summarizes the five-factor model, using the BGL mimicking portfolios. There are only two cases of coefficients that are more than two standard errors from zero, and none exceed 2.5 standard errors. There is little evidence of misspecification associated with the currency regimes.

---

[24] Ferson and Harvey (1991) find similar results for portfolios of U.S. stocks. To assess the importance of the functional form of the betas for this result, we estimated single-factor models in which the squares of the local information variables are included in the beta equations. The results are similar to the first panel of Table 6.

[25] See Harvey, Solnik, and Zhou (1992) for evidence that the expected risk premium on the world market index, which we use in the one-factor model, is countercyclical.

Table 6
The role of changing risk and changing risk premia in the predictable variation in international equity returns.

Country	Proportion of variance due to changing risk premia (Γ_1)	Proportion of variance due to changing betas (Γ_2)
	A: 1-factor model[1]	
Australia	0.610 (0.229)	0.010 (0.016)
Austria	0.957 (0.343)	0.023 (0.045)
Belgium	0.909 (0.224)	0.010 (0.012)
Canada	0.552 (0.129)	0.011 (0.013)
Denmark	0.931 (0.270)	0.004 (0.009)
France	0.862 (0.227)	0.005 (0.006)
Germany	1.154 (0.183)	0.005 (0.009)
Hong Kong	0.610 (0.263)	0.011 (0.015)
Italy	0.701 (0.208)	0.040 (0.048)
Japan	1.240 (0.193)	0.018 (0.022)
Netherlands	0.801 (0.154)	0.004 (0.007)
Norway	0.524 (0.227)	0.024 (0.030)
Singapore/Malaysia	0.441 (0.183)	0.027 (0.033)
Spain	0.881 (0.352)	0.043 (0.055)
Sweden	0.812 (0.290)	0.016 (0.022)
Switzerland	0.735 (0.193)	0.006 (0.008)
United Kingdom	0.603 (0.209)	0.019 (0.025)
United States	0.762 (0.148)	0.007 (0.007)
Average	0.783	0.016
	B: 5-factor model[2]	
Australia	0.495 (0.199)	0.032 (0.035)
Austria	0.731 (0.209)	0.020 (0.022)
Belgium	0.544 (0.138)	0.013 (0.013)

[1] The excess world market return is the single factor.

[2] The risk factors are the excess world market return, the log change in a U.S. dollar versus G-10 currency index, the change in long-term expected G-7 inflation minus the Treasury-bill return, the change in the price of oil minus the Treasury-bill return and the G-7 real interest rate. Mimicking portfolios for the last four factors are formed with the technique of Breeden, Gibbons, and Litzenberger (1989).

Table 6
Continued

Country	Proportion of variance due to changing risk premia (Γ_1)	Proportion of variance due to changing betas (Γ_2)
Canada	0.606	0.005
	(0.155)	(0.008)
Denmark	0.910	0.011
	(0.232)	(0.017)
France	0.663	0.021
	(0.210)	(0.021)
Germany	1.036	0.012
	(0.183)	(0.021)
Hong Kong	0.333	0.068
	(0.175)	(0.066)
Italy	0.669	0.042
	(0.236)	(0.046)
Japan	1.107	0.038
	(0.280)	(0.033)
Netherlands	0.738	0.010
	(0.207)	(0.035)
Norway	0.448	0.034
	(0.168)	(0.035)
Singapore/Malaysia	0.558	0.040
	(0.223)	(0.041)
Spain	0.516	0.023
	(0.210)	(0.057)
Sweden	0.537	0.042
	(0.222)	(0.057)
Switzerland	0.660	0.007
	(0.205)	(0.009)
United Kingdom	0.277	0.040
	(0.302)	(0.036)
United States	0.674	0.008
	(0.169)	(0.011)
Average	0.639	0.026

The following system is estimated for each asset i:

Disturbance	Orthogonal to
$u1_{it} = (r_{it} - Z'_{t-1}\delta_i)$	Z_{t-1}
$u2_t = (F'_t - Z'_{t-1}\gamma)'$	Z_{t-1}
$u3_{it} = [(u2_t u2'_t)(\kappa'_i Z^i_{t-1}) - (F_t u1'_{it})]$	Z^i_{t-1}
$u4_{it} = (Z'_{t-1}\gamma)(\kappa'_i Z^i_{t-1}) - \mu_{1i}$	1
$u5_{it} = (\kappa'_i Z^i_{t-1}) - \mu_{2i}$	1
$u6_t = (\gamma' Z_{t-1}) - \mu_3$	1
$u7_{it} = (u4^2_{it})\Gamma_{1i} - [\mu'_{2i}(u6_t u6'_t)\mu_{2i}]$	1

where r_i represents the return on asset i, Z is the common world predetermined information, Z^i is the local information, F are the factor returns, δ are coefficients from a linear projection of the asset returns on the information, γ are coefficients from a linear projection of the factor returns on the information, $\kappa'_i Z^i_{t-1}$ are the fitted conditional betas, μ_1 is the mean fitted value from the model, μ_2 are the mean conditional betas, μ_3 are the mean conditional risk premiums, Γ_1 is the ratio of the predictable variance due to the risk premiums to the variance of the model-fitted returns. In a separate estimation, the last equation is replaced with

$$u7_{it} = (u4^2_{it})\Gamma_{2i} - [\mu'_3(u5_{it}u5'_{it})\mu_3] \qquad 1$$

where Γ_2 is the ratio of the predictable variance due to changing conditional betas to the variance of the model-fitted returns. The sample is 1970:2–1989:12 (239 observations).

Table 7
The performance of the asset pricing models during different exchange regimes and under different capital controls.

Country	Intercept	Fixed Exchange Rates 1970:2–1973:2	Dirty Float Exchange Rates 1973:3–1980:12	$\bar{R}^2$
	A: 1-factor model pricing errors[1]			
Australia	0.621	−1.297	−0.434	−0.005
	(0.797)	(1.297)	(1.130)	
Austria	1.030	0.396	−1.177	0.004
	(0.693)	(0.920)	(0.803)	
Belgium	1.427	−0.229	−1.919	0.016
	(0.620)	(0.857)	(0.844)	
Canada	0.250	0.292	−0.357	−0.007
	(0.547)	(0.887)	(0.815)	
Denmark	0.975	0.992	−1.682	0.024
	(0.578)	(1.073)	(0.758)	
France	0.853	−1.002	−1.192	−0.002
	(0.683)	(1.042)	(1.055)	
Germany	0.947	−0.877	−1.314	0.002
	(0.640)	(1.028)	(0.853)	
Hong Kong	0.775	6.215	−0.873	0.031
	(1.009)	(2.902)	(1.564)	
Italy	0.903	−1.773	−1.752	0.005
	(0.729)	(1.047)	(1.141)	
Japan	1.423	1.080	−2.071	0.034
	(0.606)	(1.060)	(0.812)	
Netherlands	1.026	−0.791	−1.181	0.003
	(0.523)	(0.893)	(0.752)	
Norway	0.933	0.631	−0.364	−0.007
	(0.737)	(1.257)	(1.185)	
Singapore/Malaysia	0.271	3.128	−0.278	0.010
	(0.757)	(1.672)	(1.232)	
Spain	1.189	−0.168	−2.721	0.034
	(0.659)	(0.912)	(0.922)	
Sweden	1.800	−1.246	−2.048	0.015
	(0.657)	(1.045)	(0.874)	
Switzerland	0.593	−0.139	−0.893	−0.003
	(0.537)	(0.904)	(0.802)	
United Kingdom	0.642	−0.449	−1.315	−0.001
	(0.604)	(1.047)	(1.088)	
United States	0.465	−0.539	−1.114	0.005
	(0.427)	(0.721)	(0.638)	
	B: 5-factor model (with BGL mimicking portfolios) pricing errors[2]			
Australia	0.479	−1.466	−0.352	−0.004
	(0.790)	(1.299)	(1.106)	
Austria	−0.034	0.498	−0.405	−0.005
	(0.648)	(0.870)	(0.763)	
Belgium	0.870	−0.036	−1.149	0.001
	(0.612)	(0.841)	(0.844)	

[1] The excess world market return is the single factor.

[2] The risk factors are the excess world market return, the log change in a U.S. dollar versus G-10 currency index, the change in long-term expected G-7 inflation minus the Treasury-bill return, the change in the price of oil minus the Treasury-bill return and the G-7 real interest rate. Mimicking portfolios for the last four factors are formed with a technique similar to Breeden, Gibbons, and Litzenberger (1989).

Table 7
Continued

Country	Intercept	Fixed Exchange Rates 1970:2–1973:2	Dirty Float Exchange Rates 1973:3–1980:12	$\bar{R}^2$
Canada	0.563	0.005	−1.115	0.001
	(0.546)	(0.879)	(0.800)	
Denmark	0.333	0.654	−1.197	0.008
	(0.577)	(1.028)	(0.767)	
France	0.239	−0.801	0.020	−0.007
	(0.672)	(1.030)	(1.071)	
Germany	0.484	−0.868	−0.931	−0.003
	(0.630)	(1.022)	(0.852)	
Hong Kong	0.230	4.525	−0.953	0.015
	(1.011)	(2.873)	(1.568)	
Italy	0.170	−1.568	−0.413	−0.003
	(0.731)	(1.041)	(1.138)	
Japan	1.024	0.051	−1.648	0.011
	(0.605)	(1.053)	(0.817)	
Netherlands	0.748	−0.959	−1.038	0.001
	(0.523)	(0.887)	(0.751)	
Norway	0.596	0.647	0.342	−0.008
	(0.733)	(1.249)	(1.153)	
Singapore/Malaysia	0.456	1.531	−1.485	0.007
	(0.757)	(1.695)	(1.199)	
Spain	0.231	0.800	−0.876	0.001
	(0.650)	(0.879)	(0.919)	
Sweden	1.416	−0.791	−1.516	0.004
	(0.657)	(1.043)	(0.875)	
Switzerland	0.255	−0.354	−0.759	−0.004
	(0.525)	(0.900)	(0.789)	
United Kingdom	0.609	−0.483	−1.278	−0.001
	(0.620)	(1.048)	(1.036)	
United States	0.831	−0.672	−1.549	0.018
	(0.425)	(0.718)	(0.622)	
	C: 1-factor model pricing errors[1]			
Japan	1.423	0.095	−1.870	0.015
	(0.606)	(1.041)	(0.824)	
	D: 5-factor model pricing errors[2]			
Japan	1.024	−0.793	−1.368	0.003
	(0.605)	(1.025)	(0.827)	

The following system is estimated for each asset i:

Disturbance	Orthogonal to
$u1_{it} = (r_{it} - Z_{t-1}\delta_i)$	Z_{t-1}
$u2_t = (F_t' - Z_{t-1}\gamma)'$	Z_{t-1}
$u3_{it} = [(u2_t u2_t')(\kappa_i' Z_{t-1}^i) - (F_t u1_{it}')]$	Z_{t-1}^i

where r_i represents the return on asset i, Z is the common world predetermined information, Z^i is the local information, δ are coefficients from a linear projection of the asset returns on the information, $\kappa_i' Z_{t-1}^i$ are the fitted conditional betas, γ are coefficients from a linear projection of factor returns on the information, and F are the factor returns. The pricing errors are $\epsilon_{it} = r_{it} - Z_{t-1}\gamma\kappa_i' Z_{t-1}^i$. The pricing errors are regressed on two dummy variables and a constant. The table shows the regression coefficients and their standard errors. The first dummy is set equal to one during 1970:2–1973:2, the period of fixed exchange rates. The second dummy is set equal to one

Table 7
Continued

during 1973:3–1980:12, the period of active central bank intervention in the foreign currency markets (the so-called dirty float). In the analysis of Japanese capital controls, three regimes are examined. The first dummy variable is set equal to one during 1970:2–1973:12, the period when no foreign corporation could invest in Japanese securities. The second dummy variable is set equal to one during 1974:1–1980:12, a period of severe capital controls. The sample is 1970:2–1989:12 (239 observations).

The currency regime periods are similar to periods of different capital control restrictions in Japan. In the third panel of Table 7 the pricing errors for Japan are regressed on dummy variables, indicating different capital control regimes. The first is the 1970:2–1973:12 period, in which most capital flows were not officially allowed. The second is the 1974:1–1980:12 period, when capital flows were severely restricted.[26] A dummy variable coefficient is significant in the one-factor model but not in the five-factor model. This is additional evidence that systematic errors in a one-factor asset pricing model can be reduced by moving to a five-factor model.

We conducted a number of further experiments to check the sensitivity of our results to the econometric methods. We estimated predictable variance ratios using cross-sectional regression techniques similar to Ferson and Harvey (1991), as described in the Appendix. We found that the results were broadly similar. For example, in only 3 of the 18 countries were the fractions of the predictable variance explained by the five-factor model smaller than the fraction unexplained. The average value of the ratio VR1 across the countries was 0.67 in the one-factor model and 0.93 in the five-factor model. Repeating this analysis while using the first and second halves of the sample provided no strong evidence that the models perform better in the second half. The VR1's in the one-factor model were slightly higher in the first half of the sample.[27]

We also examined the time series of the adjusted R^2's from the rolling regressions of the country returns on the global risk variables, and we saw no tendency for them to increase over the sample period. The correlation between the ratios VR1 from the five-factor model

[26] Capital controls in Japan were actually relaxed in a series of steps, which raises the possibility that a more detailed analysis could detect their effects [see, for example, Bonser-Neal et al. (1990)].

[27] We repeated this analysis, using the local instrument set to capture the predictable variation in the returns, and the overall impressions were similar. In the one-factor (five-factor) model the ratio VR2 was larger than the ratio VR1 for 11 (15) of the countries, and the average of the VR2's was greater than the VR1's. We also repeated the analysis, using the world instrument set augmented by the lagged betas, and the results were similar. We checked the sensitivity to using an alternative beta estimation technique. We estimated the betas as the slope coefficients in rolling regressions that included both the risk factors and the lagged instruments Z_{t-1} on the right-hand side. When we used these rolling betas conditioned on Z_{t-1}, we found a slight decline in the average of the VR1's in the single-beta model. With five factors, however, the variance ratios were slightly more favorable to the models.

and the average adjusted R^2's is 0.7. On average, a country for which the factor model regressions have higher explanatory power is a country for which the beta pricing model explains more of the predictability in returns. When we examined the relation between the VR1's and the R^2's of the predictive regressions, we found virtually no relation.

To assess the sensitivity of the results to the currency of denomination, we reestimated system (6) for a number of the cases, using returns denominated in local currency units, in excess of a local short-term interest rate, as described in the Appendix. The overall results for those cases are not dramatically different.

5. Concluding Remarks

Using global risk factors to model returns across countries implies some strong assumptions. Such a model ignores, for example, the costs of extranational investment and information problems. Our model assumes that expected returns are determined by country-specific betas and global risk premia. We allowed the betas to vary over time with local market information variables. Assuming market integration, we forced the risk premia to depend only on global information variables. Despite these restrictions, our evidence suggests that the models can capture much of the predictable variation in a sample of returns for 18 countries. Models that incorporate additional considerations should produce even better explanatory power.

We showed how to estimate the predictable variance of returns that is explained by an asset pricing model jointly with the other parameters of the model. This approach avoids many of the econometric problems of multistep procedures and is flexible enough to address other research questions. We used the approach to estimate the contributions of time-varying betas and time-varying risk premia to the predictability in returns. We found that the largest component is the time-varying risk premia.

Appendix

This appendix describes the cross-sectional regression methods and records our data sources and definitions.

Cross-sectional regression methods

The cross-sectional regression (CSR) methods of Fama and MacBeth (1973) have typically been used to investigate the average pricing of economic risks. Ferson and Harvey (1991) use a multistep CSR methodology to estimate conditional asset pricing models. In the first step, instruments for the conditional betas in month t are obtained by regressing the excess country returns on the risk factors and using the time series for months $t-60$ to $t-1$. The second step is a cross-sectional regression for each month t of the asset returns on the predetermined betas:

$$r_{it} = \gamma_{0t} + \sum_{j=1}^{K} \gamma_{jt}\beta_{ij,t-1} + \epsilon_{it}, \qquad i = 1,\ldots,N, \tag{A1}$$

where the $\beta_{ij,t-1}$ are the betas of the excess returns for month t. The slope coefficient, γ_{jt}, $j = 1,\ldots,K$, is a portfolio excess return. The portfolio has maximum conditional correlation with the factor, as measured by the betas and the cross-sectional regression residuals.[28] The cross-sectional regression provides a decomposition of each excess return for each month. The first component, $\Sigma_{j=1}^{K} \gamma_{jt}\beta_{ij,t-1}$, represents the part of the return that is related to the cross-sectional structure of risk, as measured by the betas. The predictability of returns should be due to this component. The remaining component of return is the sum of the residual for the asset and the intercept for month t, $\epsilon_{it} + \gamma_{0t}$. This is the part of the return that is uncorrelated with the measures of risk. The part of the return that is unrelated to risk should be unpredictable.

We regress each excess return on the lagged instruments and calculate the time-series variance of the fitted values. The objective is to see how much of this predictable variance is "explained" by the model. The part captured by the model is the variance of the projection of the model fitted values [$\Sigma_{j=1}^{K} \gamma_{jt}\beta_{ij,t-1}$, from Equation (A1)] on the instruments. We calculate a ratio, VR1 for each i, dividing this variance by the variance of the fitted values of the excess return. The predictable component of a return that is not captured by the model is measured as the variance of the projection of $u_t = \gamma_{0t} + \epsilon_{it}$ on the lagged variables. This is summarized, using Z_{t-1} as the lagged variable in the following variance ratios:

$$\mathrm{VR1} = \frac{\mathrm{Var}\left\{P\left(\sum_{j=1}^{K} \gamma_{jt}\beta_{ij,t-1} \mid Z_{t-1}\right)\right\}}{\mathrm{Var}\{P(r_t \mid Z_{t-1})\}},$$

$$\mathrm{VR2} = \frac{\mathrm{Var}\{P(\epsilon_{it} + \gamma_{0t} \mid Z_{t-1})\}}{\mathrm{Var}\{P(r_t \mid Z_{t-1})\}}, \tag{A2}$$

[28] The correlation for a given factor is maximized among all portfolios with zero betas on the other factors, if the betas are the true conditional betas and a GLS cross-sectional regression is used. We report results using the simpler OLS cross-sectional regressions.

where $P(\cdot\,|Z_{t-1})$ stands for the linear projection onto Z_{t-1} and $\mathrm{Var}\{\cdot\}$ is the variance.

The CSR approach presents certain econometric problems. One problem is measurement errors in the betas, which can bias the second step, cross-sectional regressions. Shanken (1992) provides a review and analysis of the large-sample issues, assuming that the betas are constant parameters. Amsler and Schmidt (1985) provide evidence on the small-sample properties of the time-series averages of cross-sectional regression estimators. Little is known, however, about the finite-sample properties of CSR approaches for conditional asset pricing.

Connor and Uhlaner (1989) show that an iterated version of the CSR methodology can deliver consistent estimates of the risk premia under certain assumptions.[29] We use a two-stage version of the Fama–MacBeth regressions. Specifically, we use the estimated risk premia from the first-stage, cross-sectional regressions in a second stage as proxies for the risk factors. We calculate a new set of rolling regression betas for the country returns on these factors, and we use these second-round betas in a second stage of cross-sectional regressions to estimate a new set of risk premia. Our results are based on these second-round cross-sectional regressions.[30]

The world risk factors

WDRET is the arithmetic return on the Morgan Stanley Capital International world equity index, including dividends, minus the Ibbotson Associates one-month U.S. Treasury bill rate. dTED is the difference between the 90-day Eurodollar yield (Citibase FYUR3M) and the 90-day Trêasury-bill yield (Citibase FYGM3 secondary market, converted from discount to true yield to maturity). dG10FX is the difference in the trade-weighted dollar prices of foreign exchange for 10 industrialized countries (Citibase FXG10).

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

29 They assume there is a factor structure with constant loadings.

30 Connor and Uhlaner (1989) show that iterated Fama–MacBeth estimates suffer from the same rotational indeterminacy as does factor analysis. Therefore, a cost of our approach is that we are unable to isolate the pricing effects of specific factors in a multiple-beta model.

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

The parameters are estimated with 250 monthly observations. The χ^2 test for significance of the first six residual autocorrelations has a p value of .111, and the corresponding statistic for the first 12 autocorrelations has a p value of .275.

dG7ELT is the result of projecting the four-year moving average of G-7 inflation on the lagged global information variables specified below. dOIL is the natural log of the average U.S. dollar price of per barrel at the wellhead from 1974 to 1989 and the posted West Texas Intermediate price from 1969 to 1973. Since the West Texas price is consistently higher than the average wellhead price, the 1969–1973 data is grossed down by 65 percent. This represents the average premium of West Texas over the average wellhead during 1974–1976. dG7IP is calculated by weighting local industrial production indices by the following (fixed) factors: Canada .04314, France .09833, Germany .05794, Italy .13093, Japan .07485, U.K. .11137, U.S. .48343, which are the weights in G-7 gross domestic product in the third quarter of 1969. The logarithmic difference in this aggregate index is the growth in G-7 industrial production. G7RTB is calculated by aggregating individual countries' short-term interest rates. The following interest rates are used (Citibase FYCA3M–Canada 90-day Treasury bill, FYFR3M–France 90-day bill, FYGE3M–Germany 90-day bill, FYIT6M–Italy 180-day bill, FYCMJP–Japan commercial paper 1969–1976 and FYJP3M–Japan Gensaki rate 1977–1989, FYUK3M–United Kingdom 90-day bill, FYUS3M–United States 90-day bill). The aggregated G-7 interest rate is calculated by using the countries' previous quarter's shares in G-7 gross domestic product. The real G-7 interest rate is calculated by subtracting the G-7 inflation rate.

The global information variables

The Eurodollar–Treasury yield spread is the difference between the 90-day Eurodollar rate (Citibase FYUR3M) and the CRSP–Fama 90-day bill yield. The slope of the term structure is the difference between the U.S. 10-year Treasury-bond yield (Citibase FYGT10) and the CRSP Fama 90-day-bill yield. The U.S. Treasury-bill yield is the 30-day yield from the CRSP–Fama files. This variable is not lagged, because the nominal one-month yield is known at the end of the previous month.

The country-specific information variables

The lagged country returns are 18 Morgan Stanley Capital International equity indices. These returns are in excess of the CRSP–Fama 30-day bill. The lagged value of the dividend yields for 18 MSCI equity indices are used in place of the MSCI world dividend yield. The numerator is a 12-month moving sum of the dividends, and the denominator is the current index level. The short-term interest rates for the various countries are listed together with their series codes from IFS or Citibase. These are as follows: Australia, 13-week bill (IFS 61C); Austria, money market rate (IFS 60B); Belgium, 3-month bill (Citibase FYBE3M); Canada, 3-month bill (IFS 60C); Denmark, discount rate 1969–1971 (IFS 60A); call money rate 1972–1989 (IFS 60B); France, 3-month interbank (Citibase FYFR3M); Germany, Frankfurt 90-day rate (Citibase FYWG3M); Hong Kong, no data, U.S. 3-month bill used; Italy, 6-month bill (Citibase FYIT6M); Japan, call money rate 1969–1976 (Citibase FYCMJP); Gensaki rate, 1977–1989 (Citibase FYJP3M); Netherlands, call money rate 1969–1978:11 (IFS 60B), 3-month bill 1979:12–1989; Norway, prime rate 1969–1971:1, call money rate 1971:12–1989 (IFS60B); Singapore/Malaysia, no data, U.S. bill; Spain, prime rate 1969–1973:12, call money rate 1974–1976 (IFS 60B), 3-month bill 1977–1989 (IFS 60C); Sweden, 3-month bill (IFS 60C); Switzerland, 3-month deposit rate (Citibase FYSW3M); United Kingdom, 3-month bill (Citibase FYUK3M); United States, 3-month bill (Citibase FYUS3M).

TERM = the lagged term premium: The difference between long term interest rates and the above short-term rates: Australia, 15-year Treasury bond (IFS 61C); Austria, government bond (IFS 61); Belgium, government bond (Citibase FYBEGB); Canada, government bond (IFS 61); Denmark, government bond (IFS 61); France, government bond (Citibase FYFRGB); Germany, government bond (Citibase FYWGGB); Hong Kong, no data, U.S. Treasury bond; Italy, government bond (Citibase FYITGB); Japan, government bond (Citibase FYJPGB); Netherlands, government bond (IFS 61); Norway, government bond (IFS 61); Singapore/Malaysia, no data, U.S. Treasury bond; Spain, government bond (IFS 61); Sweden, government bond (Citibase FYSDGB); Switzerland, government bond (Citibase FYSWGB); United Kingdom, government bond (Citibase FYUKGB), United States, government bond (Citibase FYUSGB).

References

Adler, Michael, and Bernard Dumas, 1983, "International Portfolio Selection and Corporation Finance: A Synthesis," *Journal of Finance*, 38, 925–984.

Amsler, Christine, and P. Schmidt, 1985, "A Monte Carlo Investigation of the Accuracy of Multivariate CAPM Tests," *Journal of Financial Economics*, 14, 359–375.

Bansal, Ravi, David Hsieh, and S. Viswanathan, 1992, "A New Approach to International Arbitrage Pricing Theory," working paper, Duke University.

Bodurtha, James N., D. Chinhyung Cho, and Lemma W. Senbet, 1989, "Economic Forces and the Stock Market: An International Perspective," *Global Finance Journal,* 1, 21–46.

Bonser-Neal, Catherine, Greggory Brauer, Robert Neal, and Simon Wheatley, 1990, "International Investment Restrictions and Closed-end Country Fund Prices," *Journal of Finance,* 45, 523–548.

Braun, Phillip, Dan Nelson, and Alan Sunier, 1991, "Good News, Bad News, Volatility and Betas," working paper, University of Chicago.

Breeden, Douglas T., Michael R. Gibbons, and Robert H. Litzenberger, 1989, "Empirical Tests of the Consumption-Oriented CAPM," *Journal of Finance,* 44, 231–262.

Brown, Stephen J., and Toshiyuki Otsuki, 1990a, "Macroeconomic Factors and the Japanese Equity Markets: The CAPMD Project," in Edwin J. Elton and Martin J. Gruber (eds.), *Japanese Capital Markets,* Harper and Row, New York, pp. 175–192.

Brown, Stephen J., and Toshiyuki Otsuki, 1990b, "A Global Asset Pricing Model," working paper, New York University.

Campbell, John Y., 1987, "Stock Returns and the Term Structure," *Journal of Financial Economics,* 18, 373–400.

Chen, Nai-fu., Richard R. Roll, and Stephen A. Ross, 1986, "Economic Forces and the Stock Market," *Journal of Business,* 59, 383–403.

Cho, David, C. Eun, and Lemma Senbet, 1986, "International Arbitrage Pricing Theory: An Empirical Investigation," *Journal of Finance,* 41, 313–329.

Connor, Gregory, and Robert Korajczyk, 1986, "Performance Measurement with the Arbitrage Pricing Theory: A New Framework for Analysis," *Journal of Financial Economics,* 15, 373–394.

Connor, Gregory, and Robert Uhlaner, 1989, "A Synthesis of Two Approaches to Factor Estimation," working paper, University of California at Berkeley.

Cox, John C., Jonathan E. Ingersoll, Jr., and Stephen A. Ross, 1985, "A Theory of the Term Structure of Interest Rates," *Econometrica,* 53, 385–407.

Cutler, David M., J. Poterba, and L. Summers, 1990, "International Evidence on the Predictability of Stock Returns," working paper, Massachusetts Institute of Technology.

Dumas, Bernard, and Bruno Solnik, 1992, "The World Price of Exchange Rate Risk," working paper, Wharton School and HEC.

Fama, Eugene F., and James D. MacBeth, 1973, "Risk, Return and Equilibrium: Empirical Tests," *Journal of Political Economy,* 81, 607–636.

Ferson, Wayne E., and Campbell R. Harvey, 1991, "The Variation of Economic Risk Premiums," *Journal of Political Economy,* 99, 385–415.

French, Kenneth R., G. William Schwert, and Robert F. Stambaugh, 1987, "Expected Stock Returns and Volatility," *Journal of Financial Economics,* 19, 3–29.

French, Kenneth R., and James Poterba, 1991, "Were Japanese Stock Prices Too High?," *Journal of Financial Economics,* 29, 337–364.

Giovannini, Alberto, and Phillipe Jorion, 1987, "Interest Rates and Risk Premia in the Stock Market and in the Foreign Exchange Market," *Journal of International Money and Finance,* 6, 107–123.

Giovannini, Alberto, and Phillipe Jorion, 1989, "Time Variation of Risk and Return in the Foreign Exchange and Stock Markets," *Journal of Finance,* 44, 307–325.

Hamao, Yasushi, 1988, "An Empirical Examination of Arbitrage Pricing Theory: Using Japanese Data," *Japan and the World Economy,* 1, 45–61.

Hansen, Lars P., 1982, "Large Sample Properties of the Generalized Method of Moments Estimators," *Econometrica,* 50, 1029–1054.

Harvey, Campbell R., 1991, "The World Price of Covariance Risk," *Journal of Finance,* 46, 111–157.

Harvey, Campbell R., Bruno Solnik, and Guofu Zhou, 1992, "What Determines Expected Asset Returns?," working paper, Duke University.

Heston, Steven, Geert Rouwenhorst, and Roberto E. Wessels, 1991, "The Structure of International Stock Returns," working paper, Yale School of Organization and Management.

Hodrick, Robert J., 1981, "Intertemporal Asset Pricing with Time-varying Risk Premia," *Journal of International Economics,* 11, 573–587.

Huberman, Gur, Shmuel A. Kandel, and Robert F. Stambaugh, 1987, "Mimicking Portfolios and Exact Arbitrage Pricing," *Journal of Finance,* 42, 1–10.

Jorion, Phillipe, 1991, "The Pricing of Exchange Risk in the Stock Market," *Journal of Financial and Quantitative Analysis,* 26, 363–376.

King, Mervyn, Enrique Sentana, and Sushil Wadhwani, 1990, "A Heteroskedastic Factor Model of Asset Returns and Risk Premia with Time-varying Volatility: An Application to Sixteen World Stock Markets," working paper, London School of Economics.

Korajczyk, Robert A., and Claude J. Viallet, 1991, "Equity Risk Premia and the Pricing of Foreign Exchange Risk," forthcoming in *Journal of International Economics.*

Korajczyk, Robert A., and Claude J. Viallet, 1989, "An Empirical Investigation of International Asset Pricing," *Review of Financial Studies,* 2, 553–586.

Lehmann, Bruce N., and David M. Modest, 1988, "The Empirical Foundations of the Arbitrage Pricing Theory," *Journal of Financial Economics,* 21, 213–254.

Lintner, John, 1965, "The Valuation of Assets and the Selection of Risky Investments in Stock Portfolios and Capital Budgets," *Review of Economics and Statistics,* 47, 13–37.

Lo, Andrew W., and A. Craig MacKinlay, 1990, "Data Snooping Biases in Tests of Financial Asset Pricing Models," *Review of Financial Studies,* 3, 431–468.

Mark, Nelson C., 1985, "On Time-Varying Risk Premia in the Foreign Exchange Market: An Econometric Analysis," *Journal of Monetary Economics,* 16, 3–18.

Merton, Robert C., 1973, "An Intertemporal Capital Asset Pricing Model," *Econometrica,* 41, 867–887.

Roll, Richard R., 1977, "A Critique of the Asset Pricing Theory's Tests. I: On Past and Potential Testability of the Theory," *Journal of Financial Economics,* 4, 349–357.

Ross, Stephen A., and Michael Walsh, 1983, "A Simple Approach to the Pricing of Risky Assets with Uncertain Exchange Rates," in R. Hawkins, R. Levich, and C. Wihlborg (eds.), *The Internationalization of Financial Markets and National Economic Policy,* JAI Press, Greenwich, CT.

Shanken, Jay, 1990, "Intertemporal Asset Pricing: An Empirical Investigation," *Journal of Econometrics,* 45, 99–120.

Shanken, Jay, 1992, "On the Estimation of Beta Pricing Models," *Review of Financial Studies,* 5, 1–34.

Shanken, Jay, and Mark I. Weinstein, 1990, "Macroeconomic Variables and Asset Pricing: Estimation and Tests," working paper, University of Rochester.

Sharpe, William F., 1964, "Capital Asset Prices: A Theory of Market Equilibrium under Conditions of Risk," *Journal of Finance,* 19, 425–442.