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WHAT DETERMINES EXPECTED INTERNATIONAL ASSET RETURNS?

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ABSTRACT

This paper characterizes the forces that determine time-variation in expected international asset returns. We offer a number of innovations. By using the latent factor technique, we do not have to prespecify the sources of risk. We solve for the latent premiums and characterize their time-variation. We find evidence that the first factor premium resembles the expected return on the world market portfolio. However, the inclusion of this premium alone is not sufficient to explain the conditional variation in the returns. We find evidence of a second factor premium which is related to foreign exchange risk. Our sample includes new data on both international industry portfolios and international fixed income portfolios. We find that the two latent factor model performs better in explaining the conditional variation in asset returns than a prespecified two factor model. Finally, we show that differences in the risk loadings are important in accounting for the cross-sectional variation in the international returns.

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1. Introduction

International asset pricing models of Solnik (1974, 1983), Stulz (1981) and Adler and Dumas (1983) provide a framework to determine why expected asset returns differ across countries. Differential expected returns, in these models, are linked to differences in exposures to global risk factors.

Given the null hypothesis of world market integration, asset pricing theories typically start with a representative world investor maximizing expected utility. First-order conditions imply an Euler equation which says that the conditionally expected product of the total asset return times the marginal rate of substitution is equal to a constant. Linearization of the Euler equation shows that expected returns are linearly related to risk. However, there are many possible choices in the specification of the risk factors.

In Stulz (1981), expected returns are linear in a measure of world consumption risk. However, even in countries with the most sophisticated data collection procedures, consumption data suffers from a number of disadvantages.[1] As a result, it is problematic to estimate consumption risk of asset returns.

Solnik (1974) develops an international version of the Sharpe (1964) and Lintner (1965) capital asset pricing model where national investors differ in their consumption baskets and care about returns measured in their domestic currency. Adler and Dumas (1983) extend this model by allowing for stochastic national inflation. This approach does not suffer from the disadvantages that follow the use of consumption data, but requires stronger assumptions on consumption tastes. In these models, the common risk factor is the return on a value-weighted world equity market portfolio, hedged against currency risk. Unfortunately, the amount of currency hedging that enters this common factor depends on the individuals' util-

[1] For a description of the problems with U.S. consumption data, see Harvey (1988), Breeden, Gibbons and Litzenberger (1989) and Ferson and Harvey (1992). International consumption data is used in Braun, Constantinides and Ferson (1994). Wheatley (1988) uses the consumption framework to test the integration international capital markets.

ity function and relative wealth, and is not directly observable. Given the absence of observable market weights for the currencies entering the common risk factor, this model is empirically equivalent to a multi-risk factor model with a world equity market portfolio factor and currency risk factors. Under very restrictive (and unrealistic) assumptions about exchange rate uncertainty, this model reduces to a single observable risk factor model. For example, if purchasing power parity holds exactly at every instant, Grauer, Litzenberger and Stehle (1976) have shown that the world equity market portfolio would be the sole international risk factor.

A third route involves the specification of multivariate linear proxy for marginal utility. This representation, follows the work of Merton (1973), Ross (1976) and Solnik (1983), and suggests that expected returns are determined by exposures to many sources of risk. One difficulty with this approach is the identification of the set of factors.

While the asset pricing theories link average returns to average risk, they can also be used to study the time-variation in expected returns. Harvey (1991a), Solnik (1993), Campbell and Hamao (1992), Ferson and Harvey (1993) and Bansal, Hsieh and Viswanathan (1993) document that returns on many international equity portfolios are predictable. The asset pricing theories are required to explain both the changing cross-sectional differences in performance through time *and* the time-series predictability of the country equities.

Issues such as the integration of world capital markets and abnormal performance of any individual country cannot be answered without reference to an asset pricing theory. Indeed, there are a number of questions related to the asset pricing specification. How many factors are necessary to describe the time-variation in expected returns? What are the sources of risk? Can we characterize the time-variation in the reward per unit of sensitivity to the risk? Answers to these questions may help identify the most useful paradigm for international asset pricing. Identification of the forces that shape expected returns have immediate implications for dynamic portfolio strategies.

This paper uses the latent factors method developed by Hansen and Hodrick (1983) and Gibbons and Ferson (1985) to characterize conditionally expected international asset returns.[2] We apply this method to 18 country index returns as well as new data on 18 international industry portfolio returns and 8 bond portfolio returns. We offer important innovations. An advantage of the latent factor technique is that the researcher is not required to take a stand on the composition of the set of fundamental factors. In contrast to previous applications, our idea is to solve for the expected risk premiums from the latent factor estimation, characterize their time-series variation and try to understand what predetermined factors account for their movements.

To recover the latent premiums and risk loadings, it is necessary to assume that the risk loadings are constant. However, this assumption may not be unreasonable given that we study diversified portfolios of stocks rather than single issues. Our results indicate that the first risk premium resembles the expected return on a world market portfolio. However, this premium is not sufficient to characterize the variation in expected returns. A second premium, which is more complex to characterize, is also important. For our bond sample, this premium is related to foreign exchange returns. Our results indicate that expected returns are adequately characterized by two latent factors. Diagnostics and comparisons reveal that the latent factor model has distinct advantages over a prespecified two factor model.

Finally, we examine the ability of the model to account for the cross-section as well as the time-series of expected asset returns. Using the two latent factor model and the 44 international portfolios, differences in risk loadings across portfolios has some ability to explain the cross-sectional variation in expected returns. These results suggest that the asset pricing framework provides a useful paradigm to explain differences in expected returns.

The paper is organized as follows. Section two provides the econometric methodology that we use to extract the expected factor premiums from the asset

[2] This technique has been applied to U.S. and Japanese returns by Campbell and Hamao (1992), to 17 country returns by Harvey (1991a), G-7 equity and foreign exchange returns by Bekaert and Hodrick (1992) and daily G-7 returns by Chang, Pinnegar and Ravichandran (1991). Wheatley (1989) provides a critique of this method with reference to asset pricing tests.

returns. The data are described in the third section. The empirical results are presented in the fourth section. Some concluding remarks are offered in the final section.

2. Pricing models

2.1 Determinants of expected returns

Consider a general K-factor asset pricing model of the form:

$$E(R_{it}|\mathbf{Z}_{t-1}) = \lambda_0(\mathbf{Z}_{t-1}) + \beta_{i1}\lambda_1(\mathbf{Z}_{t-1}) + \cdots + \beta_{iK}\lambda_K(\mathbf{Z}_{t-1}), \qquad (1)$$

$$i = 1, \ldots, N, \quad t = 1, \ldots, T,$$

where

R_{it} = the return on asset i between period $t-1$ and t,

$\lambda_j(\mathbf{Z}_{t-1})$ = the expected risk premium on the j-th latent factor,

$\mathbf{Z}_{t-1}$ = the market-wide information available at t, an $L \times 1$ vector,

$\beta_{i1}, \ldots, \beta_{iK}$ = the constant conditional betas of asset i,

$N + 1$ = the number of assets ($N > K$), and

T = the number of periods.

Notice that the above K-factor model allows the conditional risk premiums, $\lambda_j(\mathbf{Z}_{t-1})$s, to vary over time as $\mathbf{Z}_{t-1}$ varies. The conditional betas, however, are assumed to be constant.

In terms of excess returns, the pricing relation (1) can be written:

$$E(r_{it}|\mathbf{Z}_{t-1}) = b_{i1}\lambda_1(\mathbf{Z}_{t-1}) + \cdots + b_{iK}\lambda_K(\mathbf{Z}_{t-1}), \qquad (2)$$

$$i = 1, \ldots, N, \quad t = 1, \ldots, T,$$

where $r_{it} = R_{it} - R_{0t}$ is the return on the i-th asset in excess of the return on the 0-th asset (the 0-th asset is arbitrarily ordered), and $b_{ij} = \beta_{ij} - \beta_{i0}$ is the 'excess' conditional beta. To simplify the presentation, we write (2) in matrix form. Define $\mathbf{r}$ as a $T \times N$ matrix of N excess returns over T periods, $\mathbf{Z}$ is a $T \times L$

matrix of instrumental variables, $\lambda(\mathbf{Z})$ is a $T \times K$ matrix of risk premiums on the K factors and B is a $K \times N$ matrix of excess conditional betas. The matrix form of the K-factor pricing theory (2) is:

$$E(\mathbf{r}|\mathbf{Z}) = \lambda(\mathbf{Z})\mathbf{B}. \tag{3}$$

To estimate the parameters, we assume the number of information variables is greater than the number of factors, i.e., $L > K$. Furthermore, we suppose throughout that $\lambda(\mathbf{Z})$ and $\mathbf{B}$ have full column rank K. Otherwise, (3) will be reduced to a pricing model with the number of factors being less than K.

As in most studies, we assume that the expected returns are governed by the multivariate regression model:

$$r_{it} = \theta_{1i}Z_{t-1,1} + \cdots + \theta_{Li}Z_{t-1,L} + \epsilon_{it}, \quad i = 1, \ldots, N, \quad t = 1, \ldots, T, \tag{4}$$

where ϵ_{it}'s are the disturbances which have zero means conditional on the instruments. Given the model (4), the pricing relationship (3) is valid if and only if the multivariate regression coefficient matrix Θ has rank K. In this case, we have:

$$H_0: \qquad \Theta = \mathbf{AB}, \tag{5}$$

where $\mathbf{A}$ is a $L \times K$ matrix of risk premium multipliers. Therefore, a test of (5) is a test of the factor pricing theory. As shown in section 2.2, both $\mathbf{A}$ and $\mathbf{B}$ can be estimated from (4) under the restriction (5) and asset pricing tests can then be constructed.

Notice that the K factors (latent variables) are unknown as are the risk premium multipliers. However, our goal is not just to report tests of the models restrictions. We also estimate the risk premium multipliers, $\mathbf{A}$, and the excess conditional betas, $\mathbf{B}$. Neither of the estimates is unique, since given estimates $\mathbf{A}$ and $\mathbf{B}$, any linear transformation of them, $\mathbf{AC}$ and $\mathbf{C}^{-1}\mathbf{B}$ gives rise to the same Θ and so the same behavior of the excess asset returns, where $\mathbf{C}$ is any $K \times K$ invertible matrix. However, the estimates of both $\mathbf{A}$ and $\mathbf{B}$ are determined up to a linear transformation. Furthermore, the estimation of Θ under the null is unique and the rank of Θ is uniquely determined.

To characterize the forces that determine the time-variation in the expected returns, we recover the risk premiums on the unknown factors, $\lambda(\mathbf{Z})$. Following Zhou (1993), consistent moment estimators of $\mathbf{A}$ can be analytically obtained, and hence $\lambda(\mathbf{Z})$ can also be analytically estimated as follows. Given an estimate of $\mathbf{A}$, $\widehat{\mathbf{A}}$, we obtain from (3) and (4) an estimate of the risk premiums:

$$\widehat{\lambda(\mathbf{Z})} = \mathbf{Z}\widehat{\mathbf{A}}. \tag{6}$$

Because $\hat{\mathbf{A}}$ is consistent, so is $\widehat{\lambda(\mathbf{Z})}$. Hence, we are able to estimate $\lambda(\mathbf{Z})$ and characterize the variation in the risk premiums.

2.2. Estimation and tests

We apply the generalized method of moments (GMM) procedure for the estimation and latent factors tests. The idea of this method is to use sample moment conditions to replace those of the model. Intuitively, given these moment conditions, the sample moments should be close to zero at the true parameters. As the GMM estimator is the solution that minimizes the weighted sample moments, it should be close to the true parameters. Indeed, as shown by Hansen (1982), the GMM estimator is consistent, i.e., converges to the true parameters with probability one as sample size gets large. In our case, the model implies the following moment conditions:

$$E(\mathbf{h}_t) = \mathbf{0}, \qquad \mathbf{h}_t \equiv \mathbf{u}_t \otimes \mathbf{Z}_{t-1}, \tag{7}$$

where $\mathbf{u}_t$ is the $N \times 1$ vector of model residuals from (3), $\mathbf{Z}_{t-1}$ is the $L \times 1$ vector of the instruments, $\otimes$ is the Kronecker product and $\mathbf{h}_t$ an $NL \times 1$ vector function of the residuals and instruments. Let $\mathbf{g}_T$ be the sample mean of $\mathbf{h}_t$:

$$\mathbf{g}_T = \frac{1}{T}\sum_{t=1}^{T} \mathbf{h}_t, \quad NL \times 1. \tag{8}$$

Hansen's (1982) GMM estimator is the solution of:

$$\min Q \equiv \mathbf{g}_T' \mathbf{W}_T \mathbf{g}_T, \tag{9}$$

where $\mathbf{W}_T$ is a positive definite $NL \times NL$ weighting matrix.

However, under the null that the rank of Θ is K, the unknown model parameters enter the quadratic form in a nonlinear way. It is not obvious, in general, how to analytically solve the GMM optimization problem (9). Moreover, the numerical optimization of (9) is a nontrivial task. Fortunately, based on Zhou (1993), we can solve the estimator analytically for a class of patterned weighting matrices:

$$\mathbf{W}_T \equiv \mathbf{W}_1 \otimes \mathbf{W}_2, \qquad \mathbf{W}_1 : N \times N, \quad \mathbf{W}_2 : L \times L.$$

The GMM estimator of Θ is explicitly given by:

$$\widehat{\Theta} = \widehat{\mathbf{A}}\widehat{\mathbf{B}}, \qquad \widehat{\mathbf{A}} : L \times K, \qquad \widehat{\mathbf{B}} : K \times N, \tag{10}$$

where

$$\widehat{\mathbf{A}} = (\mathbf{Z}'\mathbf{P}\mathbf{Z}\, T^2)^{-1/2}\mathbf{E}, \qquad \mathbf{P} \equiv \mathbf{Z}\mathbf{W}_2\mathbf{Z}',\ \mathbf{P} : T \times T,$$
$$\widehat{\mathbf{B}} = (\mathbf{Z}^{*\prime}\mathbf{P}\mathbf{Z}^*)^{-1}\mathbf{Z}^{*\prime}\mathbf{P}\mathbf{R}, \qquad \mathbf{Z}^* \equiv \mathbf{Z}\widehat{\mathbf{A}},\ \mathbf{Z}^* : T \times K,$$

and $\mathbf{E}$ is the $L \times K$ matrix stacked by the 'standardized' eigenvectors ($\mathbf{E}'\mathbf{E} = \mathbf{I}_K$) corresponding to the K largest eigenvalues of the $L \times L$ matrix:

$$(\mathbf{Z}'\mathbf{P}\mathbf{Z}\, T^2)^{-1/2\prime}(\mathbf{Z}'\mathbf{P}\mathbf{R}\, T^2)\mathbf{W}_1(\mathbf{Z}'\mathbf{P}\mathbf{R}\, T^2)'(\mathbf{Z}'\mathbf{P}\mathbf{Z}\, T^2)^{-1/2}. \tag{11}$$

Furthermore, the minimum of Q is given by:

$$Q^* = \mathrm{tr}\mathbf{W}_1(\mathbf{R}'\mathbf{P}\mathbf{R}\, T^2) - \gamma_1 - \cdots - \gamma_K, \tag{12}$$

where $\gamma_1, \ldots, \gamma_K$ are the K largest eigenvalues of the $L \times L$ matrix given in (11).

In practice, a consistent estimate of Θ is first analytically obtained as above by choosing the weighting matrix as the identity matrix. Then, a new weighting matrix can be computed:

$$\mathbf{W}_T = \left[\left(\frac{1}{T}\sum_{t=1}^{T}\mathbf{u}_t\mathbf{u}_t'\right) \otimes \left(\frac{1}{T}\sum_{t=1}^{T}\mathbf{Z}_{t-1}\mathbf{Z}_{t-1}'\right)\right]^{-1}. \tag{13}$$

and a new GMM estimator is obtained. Although both of the estimators are consistent, the latter is expected to be superior because the new weighting matrix will better capture the underlying model residual distribution.

In latent variables models, as shown in Hansen (1982), a consistent estimator of the covariance matrix of the model residuals is given by:

$$\mathbf{S}_T = \frac{1}{T}\sum_{t=1}^{T}(\mathbf{u}_t\mathbf{u}_t' \otimes \mathbf{Z}_{t-1}\mathbf{Z}_{t-1}'). \tag{14}$$

Recall our discussion in section 2.1 that the parameter estimates of **A** and **B** are unique up to an non-singular linear transformation. To obtain unique estimates, we follow the usual normalization by assuming the first $K \times K$ matrix of **B** be the identity matrix, $\mathbf{B} = (\mathbf{I}_K, \mathbf{B}_2)$. This is equivalent to choosing the first K assets as the reference assets [see Gibbon and Ferson (1985)]. After this normalization, there are $q = KL + K(N-K) = K(N-K+L)$ free parameters.

Let $\mathbf{D}_T$ be an $NL \times q$ matrix of the first order derivatives of $\mathbf{g}_T$ with respect to the free parameters. Based on (13) and (14), we can construct a GMM test:

$$H_Z \equiv T(\mathbf{M}_T\mathbf{g}_T)'\mathbf{V}_T(\mathbf{M}_T\mathbf{g}_T), \tag{15}$$

where $\mathbf{V}_T$ is a diagonal matrix, $\mathbf{V}_T = \mathrm{Diag}(1/v_1, \ldots, 1/v_d, 0, \ldots, 0)$, formed by $v_1 > \ldots > v_d > 0$, the positive eigenvalues of the following $NL \times NL$ semi-definite matrix:

$$\mathbf{\Omega}_T \equiv \left[\mathbf{I} - \mathbf{D}_T(\mathbf{D}_T'\mathbf{W}_T\mathbf{D}_T)^{-1}\mathbf{D}_T'\mathbf{W}_T\right]\mathbf{S}_T\left[\mathbf{I} - \mathbf{D}_T(\mathbf{D}_T'\mathbf{W}_T\mathbf{D}_T)^{-1}\mathbf{D}_T'\mathbf{W}_T\right]', \tag{16}$$

where $\mathbf{M}_T$ is an $NL \times NL$ matrix, of which the i-th row is the standardized eigenvector corresponding to the i-th largest eigenvalue of $\mathbf{\Omega}_T$ for $i = 1, \ldots, NL$. As shown in Zhou (1993), H_Z is asymptotically χ^2 distributed with degrees of freedom $(L-K)(N-K)$. This is the test of the model's overidentifying restrictions. The major advantage of using H_1 instead of the conventional GMM test is that H_Z is analytically available. In addition, the H_Z test delivers the same inference as the conventional GMM test, i.e., generating the same p-values.[3]

[3] This is numerically verified by Zhou (1993) in a smaller scale problem where the conventional GMM test is easy to compute.

2.3 Characterizing the variation in the premiums and diagnostics

With a set of prespecified variables, $\mathbf{F}$, which are likely candidates for the underlying factors in the economy, we can construct prespecified risk premiums by linearly projecting them on the information variables, $\mathbf{Z}$. We investigate whether this set of variables is correlated with $\lambda(\mathbf{Z})$, which are risk premiums on the latent factors. Since the estimation of $\mathbf{A}$ is only unique up to linear transformations, so are the estimated risk premiums $\widehat{\lambda(\mathbf{Z})}$. We also report the canonical correlation of the estimated risk premiums and the collection of prespecified factor premiums.

The estimation of both the model with constant conditional risk and the model with time-varying risk implies a disturbance or a pricing error matrix:

$$\mathbf{u} = \mathbf{r} - \lambda(\mathbf{Z})\mathbf{B}. \tag{17}$$

Disturbances will be affected by the number of factors that we allow in the estimation. The model implies that the conditional mean of the disturbance is zero.

One way to summarize the ability of the model to characterize the time variation in the expected returns is to study variance ratios. Let $E_M[\mathbf{r}]$ denote the model expected returns in (17). Following Ferson and Harvey (1991), we can compare the unconditional variance of these fitted returns to the unconditional variance of the fitted returns from the statistical projection model in (4) (denote as $E_Z[\mathbf{r}]$):

$$VR1 = \frac{\mathrm{Var}\{E_M[\mathbf{r}]\}}{\mathrm{Var}\{E_Z[\mathbf{r}]\}}. \tag{18}$$

If this ratio is close to one, then the expected returns from the model are closely mimicking the expected returns from the statistical model. As a result, the model 'explains' the time-variation in the expected returns.

We can also examine the variance of the part of the return that the model fails to explain. Let $E_M[\mathbf{u}]$ denote the fitted values of projecting the model residuals in (17) on the instrumental variables. If the variance of these fitted values is large, then the model is doing a poor job of setting the conditional mean of the disturbances equal to zero. A second variance ratio measures the ratio of the variance of these fitted values to the variance of the expected returns from the statistical projection in (4):

$$VR2 = \frac{\mathrm{Var}\{E_M[\mathrm{u}]\}}{\mathrm{Var}\{E_Z[\mathrm{r}]\}}. \tag{19}$$

If this ratio is close to zero, then the model pricing errors are not contributing to the predictable variation in the asset returns. These variance ratios are useful in determining not just how many premiums we need but the relative contribution of each additional premium.[4]

We also consider an additional diagnostic. The model implies that both the conditional and unconditional means of the disturbance matrix are zero. The unconditional mean is the average pricing error (APE). A large average pricing error indicates that the average return is much larger than the expected return implied by the model. Harvey's (1991a) implementation of the conditional CAPM resulted in large pricing errors for some international equity portfolios. We examine how these pricing errors are affected by increasing the number of risk factors.

Finally, we develop an analytical Wald test to examine whether or not there is structural change in the latent variables model. Suppose that the change occurs after T_1 periods. Let T_2 be the rest of the periods, $T_1 + T_2 = T$. Intuitively, we would like to compare the parameter estimates over the two subperiods. If there are substantially differences between the parameter estimates, we can reject the null that there is no structural change. Following Andrews and Fair (1988), a Wald test can be formed as follows:

$$W_T = T(\hat{\theta}_1 - \hat{\theta}_2)'(\mathrm{V}_1/\pi_{1T} + \mathrm{V}_2/\pi_{2T})^{-1}(\hat{\theta}_1 - \hat{\theta}_2), \tag{20}$$

where $\hat{\theta}_1$ and $\hat{\theta}_2$ are the analytical GMM estimators in the two subperiods, and $\pi_{1T} = T_1/T$ and $\pi_{2T} = T_2/T$. Let

$$\mathrm{V} = (\mathrm{D}_T'\mathrm{W}_T\mathrm{D}_T)^{-1}\mathrm{D}_T'\mathrm{W}_T\mathrm{S}_T\mathrm{W}_T\mathrm{D}_T(\mathrm{D}_T'\mathrm{W}_T\mathrm{D}_T)^{-1}, \tag{21}$$

[4] Ferson and Harvey (1993) provide a way to estimate the standard errors of the variance ratios. However, to get the standard errors, they are only able to consider one asset at a time. Our formulation requires the simultaneous examination of many assets. Furthermore, the variance ratios are only meant to be diagnostic measures.

then V_1 and V_2, the estimators of the asymptotic covariances of $\hat{\theta}_1$ and $\hat{\theta}_2$, are V valued at the two subperiods, respectively. In the Wald test, structural change is assessed by the stability of the parameters over two subperiods.

An alternative test may be developed that is based on the stability of the moments conditions over two subperiods. If there is no structural change, the sample moments in the second period should be close to zero even valued at the parameter estimator of the first period. This is the "predictive test" developed by Ghysels and Hall (1990). One advantage of the predictive test over the Wald test is that it uses only one estimator, making it useful in situations where it is difficult to obtain GMM estimators. However, in our case we have analytical solutions, so it is trivial for us to obtain $\hat{\theta}_1$ and $\hat{\theta}_2$. The predictive test has a much complex form when the weighting matrix is not the optimal one, so we will use only the Wald test to test the structural change in the latent variables model.

3. Data

3.1 Sources

The equity data in this study are drawn from Morgan Stanley Capital International (MSCI). Monthly data on equity indices for 16 OECD countries,[5] Hong Kong and Singapore/Malaysia are available from December 1969 to September 1991. These indices are value weighted and are calculated with dividend reinvestment. The equity indices are calculated from approximately 1500 stock returns which represents 83% of the total market value of the world's stock markets [see Schmidt (1990)]. Morgan Stanley also calculates a value-weighted world equity index which serves as the market portfolio. Returns are calculated in U.S. dollar terms.

[5] The 16 OECD countries are Australia, Austria, Belgium, Canada, Denmark, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, the United Kingdom, and the United States. Morgan Stanley also has data on Finland, Mexico and New Zealand but only from December 1987. These countries are omitted.

The MSCI international indices are composed of stocks that broadly represent stock composition in the different countries. For example, Harvey (1991a) reports a 99.1% correlation between the MSCI U.S. excess return and the New York Stock Exchange value-weighted return calculated by the Center for Research in Security Prices (CRSP) at the University of Chicago. There is a 95% correlation between the MSCI Japanese excess return and the Nikkei 225 return. An important difference between the MSCI indices and other national indices such as CRSP is the exclusion of investment companies and foreign domiciled companies. These stocks are excluded to avoid double counting.[6]

We introduce global industry indices which are also from Morgan Stanley Capital International.[7] 38 portfolios are available ranging from Aerospace and Military Technology to Wholesale and International Trade. As with the country portfolios, these indices are value weighted. In contrast to the country portfolios, the industry returns do not include dividends. However, later in the analysis we analyze an alternative set of industry portfolios that contain a dividend approximation based on an identical U.S. industry grouping.

We form 18 international industry portfolios from these 38 industries. These industry portfolios, which are documented in figure 1, resemble the SIC groupings used in the industry portfolios in Breeden, Gibbons and Litzenberger (1989).[8] The industry portfolios are formed by equally weighting the MSCI subindices in December 1969. This portfolio is held, without rebalancing, until the end of the

[6] There are disadvantages associated with the MSCI indices. First, the dividends included in the monthly return are 12-month moving averages. Second, there are no adjustments for cross-corporate ownership [see MacDonald (1989), French and Poterba (1991) and Fedenia, Hodder and Triantis (1991).]

[7] Industrial structure and international stock returns is examined in Roll (1992), Heston, Rouwenhorst, Wessels (1992) and Heston and Rouwenhorst (1993).

[8] However, Breeden, Gibbons and Litzenberger (1989) use only 12 portfolios. We form 18 portfolios by breaking up the Basic Industries category into separate portfolios for Aerospace and Military Technology, Chemicals, Forest Products, and Metals and Mining. We separate the Finance/Real Estate into two portfolios. Similarly, we separate Business Service industries from Personal Service industries. Finally, we add the Communications industry. In addition, we did not use the MSCI Multi-industry portfolio.

sample. Returns are calculated as the capital gain portion of this portfolio return. This produces a value-weighted return on an initially (December 1969) equally weighted investment.[9]

Our sample also includes bond returns from eight different countries: Canada, France, Germany, Japan, Netherlands, Switzerland, United Kingdom and United States. All of the bond indices, except for the U.S. index, are from Lombard Odier & Cie (1992) and are reported on a daily basis in the *Wall Street Journal Europe*. These bond indices are based on a small sample of plain-vanilla, actively traded, long-term government bonds in each country [see Solnik (1993)]. The U.S. bond index is from Ibbotson Associates. All eight bonds are available from January 1971 through September 1991.

Since our study focusses on expected returns, it is important to correctly specify the information environment. The set of predetermined instrumental variables follows Harvey (1991a) and includes: the world market return calculated in U.S. dollars (from Morgan Stanley Capital International), a dummy variable for the month of January, an exchange rate return index, the Standard and Poor's 500 dividend yield (from Standard and Poor's), the yield on a one-month Eurodollar deposit, the yield spread between Moody's Baa and Aaa rated bonds (from Moody's) and the excess return on a three month bill (from CRSP). The exchange rate return is based on the trade-weighted 10 countries' foreign exchange returns for the U.S. dollar investor. The exchange rate return is determined by the change in the exchange rate plus a local 30-day Eurocurrency deposit. The variable is measured in excess of the 30-day Eurodollar rate. All of the instrumental variables are available through September 1991.

We use instrumental variables that are common to all assets for a number of reasons. We are interested in characterizing the common components of expected returns across all assets. In our framework, this variation is being driven solely by global risk premiums. In addition, the evidence that local information variables influence expected returns is weak. Harvey (1991a) finds that 2 of 17 countries

[9] The value weights in December 1969 where not available to us. This is the reason that we initially equal weighted the portfolio. However, this is not very important since we can arbitrarily select portfolios for asset pricing tests.

are influenced by local information. Ferson and Harvey (1993) find that 7 of the 18 countries are influenced by local information. However, the median increase in explanatory power for these countries is only 3.1 percent. As a result, we focus on a common set of instrumental variables.

3.2 Summary statistics

Table 1 reports the means, standard deviations and autocorrelations of the asset returns, and the instrumental variables. Returns are presented in U.S. dollar terms. The sample contains 247 monthly observations extending from March 1971 through September 1991.

The first panel of table 1 examines the country equity returns. The average country equity returns range from 10.4% per annum in Italy to 26.6% per annum for Hong Kong. However, the highest standard deviation is found for Hong Kong, 43.5% per annum. Significant first-order autocorrelation is detected for five country returns: Austria, Denmark, Italy, Norway, and Singapore/Malaysia. These are fairly small portfolios compared to the capitalization of the world index[10] and may reflect infrequent trading of the stocks in these portfolios.

The next panel examines the global industry returns. These returns (as provided by MSCI) only contain the capital appreciation part of the equity return. The average annualized returns range from 7.3% for the Utilities industry to 13.5% for the Aerospace and Military Technology grouping. There is a wide range of volatility from 15.1% for Utilities to 26.7% for Metals and Mining. On a relative basis, there is less autocorrelation in these index returns than the country indices. Only 3 of 18 industries exhibit first-order autocorrelation coefficients that are greater than two standard errors from zero. This could reflect the fact that these portfolios are diversified over many markets.

The next panel presents the eight bond returns in U.S. dollar terms. The annualized returns range from 8.8% (Canada) to 14.1% (Japan). However, these

[10] The largest equity portfolio of this group, Italy, represents 1.4% of the MSCI world index as of the first quarter of 1989.

returns are greatly affected by the foreign exchange rate conversion. The volatility extends from 11.2% (Canada) to 17.6% (United Kingdom). No significant first-order autocorrelations are detected for the bond returns.

A number of the instrumental variables show a high degree of persistence. High autocorrelation is expected for the dividend yield variable because it is constructed as a 12-term moving summation. The one-month Eurodollar rate and the Baa-Aaa yield spread also exhibit very high autocorrelation. The mean world market return over the sample is 12.8% with a standard deviation of 14.9%. Interestingly, the average return exceeds the average U.S. equity return and the standard deviation is less than the U.S. return indicating that the U.S. equity portfolio is unconditionally dominated by the world portfolio over our sample.

Table 2 presents the results of linearly projecting the asset returns on the instrumental variables. The first panel considers the country index portfolios. The amount of variance explained for returns ranges from 2.1% for Italy to 12.2% for the United States. These results are consistent with those reported in Harvey (1991a). The heteroskedasticity-consistent multivariate test of predictability provides convincing evidence against the null hypothesis of no predictability.[11]

The amount of predictable variation in the industry portfolios is similar to the country index returns. Although, these industry portfolios are diversified across many different countries, each industry portfolio has a large U.S. component. Given that the instrumental variables are U.S. based, we expect to be able to predict these industry returns. Indeed, the statistical projection explains more than 8% of the variance in more than half of the industry portfolios. The highest R^2 is found for the Aerospace and Military Technology industry (14.2%) and the lowest is found for Textiles and Trade (5.6%). The multivariate test suggests that the null hypothesis of constant expected returns can be rejected at the 0.01% level.

The next panel examines the predictability of the fixed income returns. The statistical projection is able to account for on average 5% of the variance of the 8 countries' bond returns. The highest R^2 is found for the U.S. bond (7.9%) and

[11] This test is based on the Pillai trace statistic. For a description, see Kirby (1993).

the lowest for the U.K. bond (3.1%). Although the predictability of the bond returns is less than the equity returns, the multivariate test shows that the null hypothesis of no predictable variation is rejected at the 3.4% level.

Figure 2 plots the fitted values from the three groups of the regressions. Overlaid on each plot are the fitted values from regressing the world market return on the same instrumental variables. It is clear from the figure that the expected asset returns, to some degree, move together. This is the case for both the equity and fixed income portfolios. One also learns from the figures that the variation in the expected returns is related to the variation in the expected world market return. Both of these findings are important. The common movement in the expected return suggests that a global asset pricing model has some chance at identifying the determinants of the expected international returns. The coherence with the expected world market return suggests that the first factor premium may resemble the expected world market return – a premium implied by a world version of the capital asset pricing model.

4. Results

4.1 The number of factors

Table 3 considers the number of factors necessary to characterize the predictable variation in the equity returns using the latent factor model with constant conditional risk loadings. The returns are measured in excess of the 30-day U.S. Treasury bill rate. Estimation is separately carried out for the two equity groupings, country index returns and international industry returns.

For the country index returns, the results suggest a marginal rejection for the one to three factor models. The one factor results are consistent with the results of Harvey (1991a) who is unable to reject a conditional version of the Sharpe-Lintner model for 17 international equity portfolios.

For the industry returns, there is little evidence against the models' restrictions. This contrast with the country grouping could be due to the industry data only including the capital appreciation. As a result, we provide an alternative formulation of the industry portfolios which include a dividend approximation. The approximation is based on the dividend yields on U.S. stocks which fall into the same industry groupings detailed in figure 1.

The final part of table 3 examines the 8 fixed income portfolios. The test of the overidentifying conditions indicates that a one factor model is not rejected at conventional levels. However, the p-value jumps from 10.7% for the one factor model to 48.5% for the two factor model suggesting that more than one factor could be important.

4.2 Additional model diagnostics

While the statistical tests of the overidentifying restrictions were unable to unambiguously distinguish between the one and two factor models, a different picture emerges from the analysis of the pricing errors and variance ratios.

The first panel of table 4 presents average pricing errors and variance ratios for the country equity portfolios. Similar to the results in Harvey (1991a), the pricing errors of the one factor model are very large for some countries, particularly Hong Kong and Japan. The average pricing error, 0.431% per month, is about one third of the size of the average return. The average pricing error is reduced to only 0.181% with the two factor model.

A similar message is found in the variance ratios. With the one factor model, VR1 (explained by model) is 0.484 and VR2 (unexplained by model) is 0.589. This means that with the one factor model, the variance of the expected pricing errors is more than half of the predictable variance. However, with the two factor model, VR1 rises to 0.765 and VR2 falls to 0.303. With the three factor model, the VR1 and VR3 ratios are 0.845 and 0.226 respectively. This suggests that more than one factor is necessary to capture the country expected returns.

The second panel of table 4 carries out the same analysis for the 18 international industry portfolios (without dividends). From table 3, we were lead to believe that both the one and two factor models appear to fit these data better than they do for the country index returns – in that the p-values were higher. This appears to be confirmed by low relative pricing errors. The average error with the one factor model is 0.329% per month which compares to an average return of 0.885% per month. With the two factor model, the average error is reduced to 0.216% per month. However, the pricing error analysis is complicated by the lack of dividends in the data. One would expect lower or negative pricing error in returns which do not include dividends.

The variance ratio analysis indicates that the one factor model is describes 60% the time-variation in the expected returns. With the two factor model, the VR1 increases to 0.826. Not much is gained by going to the three factor model. The amount of variance explained increases by only 4%. The analysis on the industry returns with the dividend approximation reveals similar results. The one factor model explains 58% of the variation. When a second factor is introduced, the model explains 82% of the variation.

The final panel in table 4 examines how the model explains the variation in the international bond portfolios. The average pricing errors are small compared to the analysis of equities. The average bond returns from table 1 is .9% per month. The average pricing error reported in table 4 is 0.018% per month. The largest error is found for the Japanese bond. When a second factor is introduced, the pricing error is slightly reduced. The three factor model eliminates the average pricing error.

Similar to the equities, the first factor explains about 65% of the expected bond returns. When a second factor is introduced the proportion jumps to 83%. With three factors, 95% of the predictable variation is explained.

The pricing error and variance ratio analysis indicates that more than one factor is necessary to characterize the time-varying expected returns for all of the portfolios. This contrasts with the results reported in table 3 which suggested that one factor appeared to be enough (statistically) and provides motivation to explore other diagnostic measures.

The results of the stability tests reveal evidence against all of the specifications [not reported]. A popular assumption in most conditional asset pricing tests is that the factor premiums are linear in the instrumental variables and the coefficients are fixed through time.[12] Our tests suggest that the assumption of constant coefficients is rejected.

In our applications, we split the sample at the mid-point and let $T_1 = 123$, $T_2 = 124$ and $T = 247$. Coefficient stability is rejected for the one factor model for all the portfolios except the bond portfolio. For the two factor model, stability is rejected for the industry portfolios. The two factor bond model is marginally rejected. There is no evidence against stability for the country portfolios for the two factor model.

4.3 Characterizing the factor premiums

Given the assumptions of the econometric model, conditionally expected returns from the model are being driven by conditional variation in the risk premiums. There are two interesting questions that need to be addressed. First, do the model expected returns resemble the expected returns that result from the statistical projection of the asset returns on the instrumental variables. The variance ratios in table 4, indicate that the model fitted returns are indeed similar to the statistical fitted returns. Second, what are the model premiums? Do they have any economic interpretation?

The advantage of the technique of latent variables is that the researcher is not forced to take a stand on the specification of the proper set of factors. The model is estimated and the minimum number of premiums is extracted to characterize the time variation in the expected returns. We now investigate the economic interpretation of the latent premiums from our estimation.

Most asset pricing theories suggest that there is a role for a 'world' market portfolio as a factor. This is the international extension of the Sharpe (1964) and Lintner (1965) capital asset pricing model. The conditional version of this model suggests that the market premium is the conditionally expected excess return on

[12] For a recent example, see Dumas and Solnik (1993).

a world market portfolio.

There is some theoretical guidance in choosing a second factor. International asset pricing models suggest that deviations from purchasing power parity could induce a premium associated with foreign exchange risk. For example in the model of Adler and Dumas (1983) and Dumas and Solnik (1993), covariances with different foreign exchange investments are priced. We summarize the exchange risk factor by the return on a trade weighted FX portfolio in 10 countries. In contrast to the FX portfolio used in Ferson and Harvey (1993), our portfolio is a return in that it include both the exchange rate change and the local Eurocurrency deposit rate. The factor is measured in excess of the 30-day Treasury bill rate.

Three other prespecified factors are identified. These factors are motivated by Chen, Roll and Ross (1986). They include the change in the price of oil, the change in OECD industrial production and the OECD inflation rate. In contrast to the first two factors, these factors are not excess returns.

Table 5 presents the results of regressing these prespecified factors on the information set. The results indicate the that 13.6% of the variation in the excess market return can be predicted with this set of instruments. The results in table 5 suggest that 8.5% of the change in the FX index is predictable. The projections indicate that the three macroeconomic factors are, to some degree, predictable. While only 3.3% of the variation in the oil price change can be accounted for with the information set, over 27% of the variation in the OECD inflation rate is predictable. Industrial production has an R^2 of 1.41%.

In the lower panels of table 5, the coefficients associated with the instrumental variables representation of the latent premiums, $\hat{\mathbf{A}}$ from (6), are reported for the two factor specification. The patterns and magnitudes of the coefficients on the factor 1 premium for the international equity returns resemble the coefficients on the prespecified world excess returns regression. Specifically, the coefficients in the OLS regression on the four most significant variables DIV, E\$30, Baa-Aaa and 3-1BILL are 9.8, -5.6, 15.2 and 5.2 and from the latent factor estimation are 15.8, -8.0, 19.9 and 5.6. Similar patterns are found for the international industry returns and the bond portfolio returns. It is more difficult to characterize the second premium by examining the coefficients.

Table 6 shows the correlation between the expected values of these prespecified premiums and the latent premiums. In the two factor estimation, the first factor premium has 95% correlation with the world market expected return when the country indices are examined and about 90% correlation when the international industries are used in the estimation. For the fixed income portfolios, the first factor has 83% correlation with the expected excess market return.

Although the factor premiums are not constrained to be identical across the asset groups, the correlation of the premiums is very high. The premium from the country estimation has 95% correlation with the premiums from the industry estimation. The country risk premium has 80% correlation with the first premium from the bond return estimation.

Figure 3 provides plots of the conditionally expected excess world market return and the first factor premium for the country index returns, the international industry returns (without dividends) and the bond samples. The graphs provide three interesting insights.

First, the expected factor premiums from all the asset sets are similar. This suggests that the same forces are determining expected returns in both the equity and bond markets. Second, the closeness of the factor premiums from the latent variable model and the conditionally expected excess return on the world market portfolio is striking. Third, there is a distinct business cycle pattern in the expected values. While Fama and French (1989) and Ferson and Harvey (1991) have noted the business cycle patterns in U.S. expected returns, no one has documented any relation for international returns.

In figure 3, the NBER U.S. business cycle peaks and troughs are overlaid. Harvey (1991b) shows that there is an 88% correlation between the G-7 business cycle and the U.S. business cycle over the 1969–1989 period. Interestingly, the highest premiums occur around business cycle troughs and the lowest premiums are found around business cycle peaks. This is found for all the business cycles in the sample. The intuition follows from investors demanding a high premium at the trough of the business cycle to give up consumption in order to invest in equities. While these results are consistent with work on U.S. expected returns, the most recent business cycle provides some out-of-sample validation of these patterns.

Consistent with the analysis of the coefficients in table 5, the second factor premium is more difficult to characterize. For the bond sample, the second factor premium has a strong foreign exchange component (correlation 81%). However, the foreign exchange component is less important for the equity returns. For the country and industry returns, the second factor premium is related to the oil premium and the inflation premium. In the bond returns sample, the second premium is related to the inflation premium as well as the foreign exchange premium.

The fitted values of the second factor premium and the expected foreign exchange premium are presented in figure 4. Consistent with the correlation analysis, there is little relation between the second latent factor and the prespecified foreign exchange premium for the equity portfolio. However, the latent premium closely tracks the variation in the foreign exchange return for the bond returns.[13]

4.4 A comparison to a prespecified two factor model

We compare the performance of the two latent factor model to a conditional asset pricing model with two prespecified factors. Given the analysis in tables 5 and 6, we choose the excess world market return and the change in the U.S. dollar FX index as the prespecified factors.

Following Ferson (1990) and Harvey (1992), the following model is estimated:

$$(\boldsymbol{u}_f \quad \boldsymbol{e}) = (\boldsymbol{f} - \mathbf{Z}\boldsymbol{\delta}_f \quad \mathbf{r} - \mathbf{Z}\boldsymbol{\delta}_f(\mathbf{u}_f'\boldsymbol{u}_f)^{-1}\mathbf{u}_f'\mathbf{r}), \tag{22}$$

where $\boldsymbol{f}$ is a $T \times 2$ matrix of the prespecified factors, $\boldsymbol{u}_f$ is the factor innovation matrix, $\mathbf{r}$ are the asset excess returns, and $\boldsymbol{e}$ are the pricing errors. The model implies that $E[(\boldsymbol{u}_{ft} \quad \boldsymbol{e}_t)|\mathbf{Z}_{t-1}] = \mathbf{0}$. This model assumes that the factor premiums are linear in the information variables. In addition, $(\boldsymbol{u}_f'\boldsymbol{u}_f)^{-1}\boldsymbol{u}_f'\boldsymbol{r}$ is the

[13] The foreign exchange rate influence on the bond market premium is consistent with the results presented in Dumas and Solnik (1993).

conditional beta which is assumed to be constant. This system is estimated with Hansen's (1982) GMM. With 2 factors, 8 instruments and N assets, there will be $8 \times N$ overidentifying restrictions.

Table 7 presents the tests of the prespecified model as well as model diagnostics. For the country index portfolios, the model is not rejected at conventional levels (p-value is 0.120). However, this model does not appear to perform as well as the two factor latent variables model. Comparing the model diagnostics reported in tables 4 and 7, the average pricing error for the prespecified model is .240% per month for all the country returns compared to .181% for the latent factor model.

The prespecified model fails to explain many important portfolio expected returns such as Hong Kong which has an average error of 1.012% per month. More importantly, the VR2 ratio, which tells us the proportion of unexplained variance to the predictable variance, for the prespecified model is 52.6% for the country returns which is higher than the 30.3% reported in table 4.

A similar story emerges for the international industry portfolios (without dividends). The average pricing error for the prespecified model is -0.576% compared with 0.216% for the latent factor model.

The average pricing error across the 18 portfolios using the prespecified model is .123% compared to the .047% reported in table 4 for the latent factor model. The average pricing error for the Chemicals industry is -0.771% per month which is much different than the .080% per month with the latent factors model. Consistent with the country equity returns, the industry variance ratios are worse for the industry portfolios. The VR2 ratio is 36.7% compared to the 18.6% reported in table 4. However, the model's restrictions are not rejected at conventional levels with the prespecified factor model.

In the bond sample, the pricing errors are much higher with the prespecified factor model, 0.345% compared to 0.012% with the latent model. In addition, the variance ratio analysis indicates that little of the variation is explained by the two prespecified factors. In addition, there is evidence against the model's restrictions when the bond portfolios are examined. The p-value of the test of the overidentifying restrictions is .032.

4.5 The relative importance of the factor premiums

Another method of diagnosing the importance of the factor premiums is to measure the relative contribution of each premium to the conditionally expected returns. With the two factor model, the expected returns on asset i are determined by

$$r_{it} = b_{i1}\lambda_{1t} + b_{i2}\lambda_{2t}.$$

The proportions of predictable variance accounted for by the sources of risk are:

$$\text{Prop}_1 = \frac{b_{i1}^2 \text{Var}(\lambda_{1t})}{\text{Var}(\mathbf{B}_i\boldsymbol{\lambda}_t)} \qquad \text{Prop}_2 = \frac{b_{i2}^2 \text{Var}(\lambda_{2t})}{\text{Var}(\mathbf{B}_i\boldsymbol{\lambda}_t)}$$

where $\mathbf{B}_i\boldsymbol{\lambda}_t$ are the expected return generated by the asset pricing model, defined previously as $E_M[r_i]$. The variance ratios will not necessarily sum to unity because of a nonzero covariance between λ_{1t} and λ_{2t}.

Variance decompositions are presented for both the latent factor and prespecified factor models in table 8. The risk loadings are also reported in this table. For the equity returns, the first source of risk is most important. The first risk premium accounts for 69% of the model expected returns for the country index returns. There is very high correlation between the factor premiums with the industry portfolios. This is evident from the similarity of the **A** coefficients reported in table 5. As a result, only the one factor model is presented for the industry portfolios. In contrast to the equity portfolio, the first factor accounts for only 29% of the variation for the fixed income portfolios.

The second factor premium, while less overwhelming for the equities, plays an important role in the latent factor model. The second premium accounts for 28% of the variation in the model expected returns for the country indices and 70% of the variation of the bond portfolios.

The variance decomposition for the prespecified factor model exhibits some similarities to the latent factor model.[14] For example, with the country returns the first factor premium accounts for 80% of the expected return variation. The second factor accounts for 17%. For the industry portfolios, the first premium accounts for 97% of the variation and the second premium only 3.3%. Finally, in the analysis of the fixed income portfolios, more than one factor is needed. The first factor premium explains only 28% of the variation while the second premium accounts for 85% of the variation.

Overall, the results suggest a role for a second factor when portfolios are grouped by countries or with fixed income portfolios. This contrasts with results presented in Ferson and Harvey (1991) who find that the market premium is overwhelmingly important in explaining the conditionally expected returns using U.S. data. Our results are supportive of the recent prespecified factor models proposed by Dumas and Solnik (1993) and Ferson and Harvey (1993). Both of these models include a role for exchange risk. Our results suggest that exchange risk is related to the second latent factor. However, it is also clear that the second factor is more complex.

4.6 The cross-sectional behavior of asset returns

Most of our analysis has concentrated on explaining the time-variation in the expected returns for 44 different portfolios. Our results indicate that the two latent factor model, with constant conditional risk, can account for about 75% of the conditionally expected returns across these 44 portfolios. In this formulation, the time-variation is being driven by the latent premiums.

Asset pricing theories were originally developed to explain the cross-sectional behavior of expected returns. The model implies that assets with high risk should have high expected returns. Recently, Fama and French (1992) show "an absence of a relation between β and average returns for 1963–1990" using various U.S.

[14] The variance ratios of the latent factor and prespecified models cannot be directly compared because $E_M[r_i]$, the denominator, is different for the two models.

equity portfolios and assuming that the a U.S. equity market portfolio is the sole factor. These findings challenge the usefulness of the present asset pricing models.

However, as emphasized in Roll (1977), Ross (1977) and Roll and Ross (1993), the mean-variance inefficiency of the benchmark could lead to the finding of no significant relation between expected returns and β. Indeed, the results presented in table 1, suggest that the U.S. market portfolio is unconditionally dominated by the world market portfolio.

While our data and approach are not directly comparable with Fama and French (1992), some insight can be gained by examining the latent factor model's ability to explain the cross-sectional behavior of the average asset returns. Figure 5 plots the risk loadings from the latent two factor model against the average excess returns over the 1971–1991 period. In contrast to the previous results, the loadings are based on a latent factor estimation which simultaneously considers all 44 assets. This estimation is only feasible using the analytical method with patterned weighting matrices detailed in section 2.2. From this cross-sectional scatter plot, it is evident that the some of highest expected returns are found with the portfolios with the highest risk loadings.

If a regression of average returns on the risk loadings is estimated, the R^2 is 35% and the intercept is insignificantly different from zero. These results suggest that the asset pricing model provides a useful paradigm to explain both the cross-section and time-series behavior of expected asset returns.

5. Conclusions

This paper explores the sources of predictability in international bond and equity returns. While most research on international asset returns has relied upon either principal components analysis of the ex post asset returns or a prespecified factor approach, we investigate the usefulness of a latent factors technique. The advantage of this approach is that the factors need not be specified.

Our goal is not simply to test rank restrictions which determine the number of factors necessary to characterize the expected returns. Our idea is to solve for the factor premiums and explore their time-series patterns as well as the correlation with a set of prespecified variables.

We test our model on using 18 country index returns as well as new data on 18 international industry portfolio returns and 8 fixed income portfolios. Although the statistical tests cannot reject a one-factor model, our diagnostics indicate that at least one additional factor is necessary to characterize the expected returns for the country index returns and the bond returns. With only two factor premiums, 77% of the predictable variation in 18 country index returns can be explained. Using the 18 international industry portfolios or the 8 bond portfolios, the two factor model accounts for 83% of the predictable variation.

Our characterization of the factor premiums suggest that the first premium has a strong resemblance to the expected excess returns on the world market portfolio. Consistent with the findings in the U.S. data of Fama and French (1989), we find that the world market risk premium is highest at business-cycle troughs and lowest and business-cycle peaks. We find that the counter-cyclical behavior of the first risk premium also obtains in the most recent business cycle episode in 1990–1991.

The second premium is more difficult to characterize. For the bond returns, we find a high correlation between this premium and the conditionally expected change in a world foreign exchange returns index. This supports the role of foreign exchange risk proposed in Adler and Dumas (1983) and explored empirically in Ferson and Harvey (1993) and Dumas and Solnik (1993). However, the second latent factor appears to be characterized by more than a foreign exchange factor.

We also compare the performance of the latent factor model to a prespecified conditional factor model. The prespecified model assumes the existence of two factors: the excess returns on the world equity portfolio and the foreign exchange returns index. The model diagnostics suggest that the latent factor model has distinct advantages over the prespecified factor model in that the average pricing errors are smaller and the ability of the model to account for the expected returns is higher.

The relative importance of the risk premiums is also explored. Recent research, such as Ferson and Harvey (1991), suggests that the market factor is overwhelmingly important in explaining the time-series of expected asset returns. We find that the first factor premium is, indeed, the most important accounting for about 80% of the model's predictable variation. However, the second factor premium, is important for the country returns and very important for the bond returns.

Finally, we test the ability of the model to account for the cross-sectional behavior of expected returns. Recent work by Fama and French (1992) on U.S. equity data concludes that there is no significant relation between risk and return. Our results, which use international data and an international asset pricing framework, suggest that the cross-section of average returns is significantly related to the two risk loadings. The latent factor model appears to be a useful paradigm to help understand both the time-series and cross-sectional characteristics of expected returns.

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Table 1

Means, standard deviations and autocorrelations of international equity and bond returns calculated in U.S. dollars and based on data from March 1971 to September 1991 (247 observations).

Variable	Mean (arith.)	Mean (geo.)	Std. dev.	Autocorrelation ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	ρ_{24}
Country index returns									
Australia	14.260	10.156	27.745	-0.014	-0.053	-0.008	0.009	-0.042	0.041
Austria	15.907	13.396	22.495	0.167	0.040	0.033	0.083	0.025	0.026
Belgium	16.333	14.108	20.975	0.092	0.046	0.036	0.040	0.039	0.034
Canada	11.379	9.387	19.736	-0.013	-0.096	0.095	-0.025	-0.055	0.043
Denmark	17.734	15.724	19.757	0.018	0.132	0.092	0.102	-0.132	0.079
France	15.908	12.573	25.653	0.085	0.003	0.127	0.023	-0.045	-0.000
Germany	14.773	12.341	21.797	-0.007	-0.017	0.106	0.062	-0.054	0.002
Hong Kong	26.593	17.287	43.468	0.053	-0.036	-0.009	-0.055	-0.009	-0.017
Italy	10.400	6.760	27.140	0.145	-0.028	0.095	0.074	0.036	0.014
Japan	21.118	18.377	23.037	0.058	0.012	0.058	0.047	0.067	-0.000
Netherlands	16.987	15.123	18.902	0.033	-0.034	0.067	-0.100	0.056	0.003
Norway	16.319	12.117	28.836	0.168	-0.001	0.153	-0.073	0.031	0.014
Singapore/Malaysia	20.095	14.823	32.538	0.165	-0.011	-0.082	0.049	0.045	-0.002
Spain	11.787	9.069	23.171	0.124	-0.004	-0.043	0.080	-0.013	0.121
Sweden	18.480	15.942	22.154	0.080	-0.028	0.053	-0.014	0.031	0.003
Switzerland	13.883	11.821	20.131	0.048	-0.063	0.046	0.006	0.001	-0.016
United Kingdom	17.545	13.950	27.318	0.101	-0.093	0.059	0.004	-0.007	0.059
United States	11.500	10.181	15.988	0.022	-0.047	0.015	-0.022	0.052	-0.027
International industry returns (without dividends)									
Aerospace & Military Technolo	13.512	10.760	23.329	0.105	0.002	-0.036	-0.017	0.028	0.016
Capital Goods	10.563	8.783	18.640	0.050	-0.026	0.048	-0.054	-0.017	0.020
Chemicals	8.989	7.486	17.115	0.035	-0.058	0.135	-0.026	0.037	0.035
Communications	9.297	8.172	14.764	0.109	-0.018	0.028	-0.133	-0.003	0.017
Construction	12.434	10.055	21.852	0.065	0.037	-0.019	0.056	0.064	0.043
Consumer Durables	10.482	8.746	18.400	0.097	0.000	0.057	0.020	0.022	-0.006
Energy	10.832	8.284	22.585	0.021	-0.039	-0.014	0.037	0.063	-0.069
Finance	12.494	10.534	19.702	0.174	-0.033	0.010	-0.028	0.134	-0.024
Food & Tobacco	12.675	11.426	15.545	0.115	-0.000	0.093	-0.062	0.079	-0.012
Forest Products	7.569	5.430	20.646	0.038	-0.068	0.028	-0.007	-0.039	0.039
Leisure	11.198	8.957	20.902	0.177	0.058	0.033	-0.067	0.018	-0.102
Metals and Mining	10.014	6.427	26.702	0.038	-0.075	0.026	0.099	0.068	0.062
Real Estate	14.457	10.609	27.781	0.097	0.004	0.056	0.004	0.134	0.021
Services-Business	9.929	8.375	17.432	0.116	-0.060	0.057	-0.018	0.002	0.044
Services-Personal	11.190	9.702	17.134	0.030	0.022	-0.037	-0.003	0.100	0.008
Textiles & Trade	13.540	11.044	22.301	0.048	0.040	-0.013	-0.097	0.052	-0.037
Transportation	10.466	8.655	18.900	0.122	0.016	-0.017	-0.055	0.083	-0.014
Utilities	7.257	6.130	15.063	0.071	-0.075	-0.021	0.044	0.015	0.015

Table 1 (continued)

Variable	Mean (arith.)	Mean (geo.)	Std. dev.	Autocorrelation ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	ρ_{24}
International bond returns									
Canada	8.624	8.177	11.199	0.024	-0.068	0.019	-0.148	0.028	-0.012
France	11.319	10.296	14.037	0.035	0.055	0.099	0.094	-0.003	-0.043
Germany	12.416	11.316	14.556	0.042	0.062	-0.040	-0.005	-0.056	-0.032
Japan	14.147	12.886	15.575	0.090	0.010	0.054	0.048	0.081	-0.080
Netherlands	12.234	11.220	13.929	0.097	0.022	-0.011	0.013	-0.009	-0.011
Switzerland	10.873	9.899	13.735	0.093	0.086	0.018	0.080	0.033	-0.062
United Kingdom	10.551	9.000	17.596	0.059	0.015	-0.189	0.003	-0.007	0.019
United States	9.127	8.479	11.254	0.067	-0.038	-0.139	-0.004	-0.010	-0.060
Instrumental variables									
World return	12.768	11.605	14.891	0.092	-0.047	0.045	-0.018	0.060	0.015
G10 currency returns	1.647	1.200	9.464	0.016	0.127	0.056	0.050	0.023	-0.004
S&P 500 dividend yield	4.163	4.155	0.267	0.982	0.955	0.927	0.900	0.663	0.472
1 month Eurodollar	9.061	9.022	0.927	0.946	0.884	0.829	0.772	0.558	0.138
Moody's Baa-Aaa yield	1.274	1.273	0.127	0.950	0.881	0.831	0.795	0.437	0.079
3 month-1 month T. bill	0.908	0.907	0.473	0.277	0.018	-0.002	-0.008	-0.086	0.021

The industry portfolios are based on a aggregation of 37 Morgan Stanley Capital International industry indices.

Table 2

The predictability of international equity and bond returns calculated in U.S. dollars. Expected values are obtained by linearly projecting on the instrumental variables. The instrumental variables are: a constant, the lagged excess return on the Morgan Stanley Capital International world equity index, (WRD), the lagged return on the index of investments in 10 currencies, (XRG10), the lagged dividend yield on the Standard and Poor's 500 stock index (DIV), a dummy variable for the month of January (JAN), the return on a 30-day Eurodollar deposit (E$30), the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds (Baa-Aaa) and the return for holding a 90-day U.S. Treasury bill for one month less the return on a 30-day bill (90-30TB). Heteroskedasticity-consistent t-statistics are in brackets. Estimates are based on monthly data from 1971:3–1991:09 (247 observations).

Portfolio	Intercept	WRD_{t-1}	$XRG10_{t-1}$	JAN_t	DIV_{t-1}	$E\$30_{t-1}$	$Baa\text{-}Aaa_{t-1}$	$90\text{-}30TB_{t-1}$	$R^2/\bar{R}^2$
A. Country index returns									
Australia	0.004 [0.147]	0.266 [1.881]	-0.098 [-0.550]	0.023 [1.112]	13.375 [1.546]	-5.502 [-2.389]	-9.472 [-0.606]	11.563 [2.708]	0.103 0.077
Austria	0.051 [2.720]	0.175 [1.387]	-0.004 [-0.021]	-0.026 [-1.706]	-0.464 [-0.076]	-1.616 [-1.041]	-25.133 [-2.195]	3.446 [1.378]	0.052 0.024
Belgium	0.016 [0.862]	-0.069 [-0.695]	-0.061 [-0.345]	0.014 [1.195]	11.429 [2.047]	-6.651 [-3.600]	4.061 [0.355]	4.383 [1.551]	0.064 0.037
Canada	-0.006 [-0.306]	0.081 [0.986]	-0.121 [-1.082]	0.010 [0.602]	11.244 [1.822]	-3.829 [-2.115]	-4.583 [-0.384]	11.051 [3.295]	0.108 0.082
Denmark	0.033 [1.838]	-0.215 [-2.338]	-0.089 [-0.599]	0.018 [1.305]	-2.625 [-0.513]	-3.972 [-2.633]	19.701 [1.743]	1.110 [0.509]	0.054 0.027
France	-0.005 [-0.217]	0.010 [0.087]	-0.131 [-0.685]	0.014 [0.762]	24.223 [2.655]	-7.726 [-3.052]	-8.610 [-0.583]	0.748 [0.203]	0.054 0.027
Germany	0.005 [0.244]	0.038 [0.293]	-0.127 [-0.650]	-0.010 [-0.757]	12.709 [2.052]	-5.084 [-2.939]	0.753 [0.062]	2.288 [0.817]	0.036 0.007
Hong Kong	0.043 [0.864]	0.233 [1.406]	-0.089 [-0.294]	0.064 [2.534]	3.209 [0.221]	-6.846 [-1.887]	9.513 [0.398]	2.710 [0.432]	0.050 0.022
Italy	-0.014 [-0.643]	0.093 [0.837]	0.117 [0.579]	0.022 [1.215]	11.671 [1.419]	-1.764 [-0.623]	-7.306 [-0.466]	0.519 [0.159]	0.021 -0.007
Japan	0.027 [1.285]	0.086 [0.619]	0.083 [0.474]	0.003 [0.258]	1.632 [0.250]	-5.261 [-2.718]	23.461 [1.716]	-1.415 [-0.516]	0.053 0.026
Netherlands	-0.004 [-0.212]	-0.058 [-0.578]	-0.044 [-0.331]	0.021 [1.510]	13.885 [2.557]	-6.326 [-3.652]	12.485 [1.138]	4.256 [1.603]	0.091 0.065
Norway	0.035 [1.323]	0.063 [0.383]	-0.125 [-0.602]	0.042 [2.211]	2.339 [0.239]	-1.413 [-0.528]	-25.979 [-1.479]	5.774 [1.204]	0.035 0.007
Singapore/Malaysia	0.023 [0.659]	0.068 [0.425]	-0.086 [-0.431]	0.075 [2.726]	8.104 [0.454]	-5.383 [-1.864]	1.715 [0.100]	11.207 [2.888]	0.100 0.074
Spain	0.049 [2.697]	0.161 [1.332]	-0.015 [-0.090]	0.012 [0.726]	-14.188 [-1.854]	1.915 [0.971]	-10.080 [-0.788]	5.023 [1.779]	0.037 0.009
Sweden	-0.001 [-0.075]	0.149 [1.190]	-0.072 [-0.473]	0.026 [1.625]	-3.061 [-0.442]	-0.098 [-0.044]	21.608 [1.556]	2.027 [0.663]	0.048 0.020
Switzerland	0.011 [0.588]	-0.069 [-0.651]	-0.154 [-0.974]	0.008 [0.530]	13.230 [2.168]	-6.304 [-3.400]	-1.032 [-0.083]	5.056 [1.882]	0.064 0.037
United Kingdom	-0.019 [-0.801]	-0.045 [-0.309]	-0.057 [-0.316]	0.043 [1.572]	21.034 [2.535]	-7.825 [-3.246]	11.283 [0.735]	6.965 [1.981]	0.095 0.069
United States	-0.019 [-1.363]	-0.078 [-0.805]	-0.006 [-0.058]	0.015 [1.243]	10.551 [1.982]	-4.299 [-3.357]	17.832 [1.853]	6.299 [2.407]	0.122 0.096

Multivariate test of predictability

Portfolio	F-statistic	p-value
16 countries	1.5892	0.0001

Table 3 (continued)

Portfolio	Intercept	WRD_{t-1}	$XRG10_{t-1}$	JAN_t	DIV_{t-1}	$E\$30_{t-1}$	$Baa\text{-}Aaa_{t-1}$	$90\text{-}30TB_{t-1}$	$R^2/\bar{R}^2$
B. International industry returns (without dividends)									
Aerospace & Military Technology	-0.045 [-2.188]	-0.043 [-0.381]	-0.102 [-0.706]	0.044 [2.143]	33.983 [4.430]	-9.256 [-4.684]	4.242 [0.332]	1.290 [0.399]	0.142 0.117
Capital Goods	-0.011 [-0.688]	-0.035 [-0.370]	-0.000 [-0.003]	0.017 [1.494]	7.795 [1.380]	-5.477 [-3.882]	28.045 [2.755]	4.825 [2.001]	0.107 0.081
Chemicals	0.003 [0.177]	-0.084 [-0.943]	0.029 [0.257]	0.017 [1.493]	5.341 [1.063]	-4.374 [-5.097]	13.813 [1.568]	5.443 [2.726]	0.080 0.053
Communications	-0.012 [-0.928]	-0.045 [-0.557]	-0.079 [-0.785]	0.015 [1.315]	6.962 [1.539]	-3.348 [-3.108]	15.793 [1.902]	5.103 [2.613]	0.094 0.067
Construction	0.007 [0.367]	0.034 [0.266]	0.037 [0.235]	-0.001 [-0.073]	6.916 [1.169]	-5.399 [-3.103]	14.196 [1.227]	6.258 [2.286]	0.070 0.043
Consumer Durables	-0.014 [-0.851]	0.073 [0.837]	-0.041 [-0.317]	0.006 [0.507]	6.409 [1.218]	-5.334 [-3.368]	34.867 [3.636]	3.390 [1.649]	0.123 0.098
Energy	0.025 [1.265]	0.011 [0.083]	-0.012 [-0.082]	-0.013 [-0.768]	3.972 [0.565]	-3.675 [-1.691]	-8.266 [-0.540]	10.413 [2.400]	0.064 0.036
Finance	-0.013 [-0.738]	0.001 [0.015]	0.066 [0.463]	0.018 [1.130]	9.887 [1.842]	-5.158 [-3.385]	22.109 [2.017]	3.807 [1.855]	0.085 0.058
Food & Tobacco	-0.002 [-0.120]	-0.111 [-1.260]	0.008 [0.076]	0.007 [0.601]	6.106 [1.377]	-3.932 [-3.137]	15.553 [1.836]	6.270 [2.871]	0.094 0.068
Forest Products	-0.004 [-0.235]	-0.073 [-0.742]	0.051 [0.387]	0.017 [0.917]	2.864 [0.414]	-3.879 [-2.139]	20.672 [1.724]	9.370 [3.178]	0.099 0.072
Leisure	-0.017 [-0.974]	-0.007 [-0.058]	-0.008 [-0.062]	0.004 [0.247]	9.206 [1.301]	-4.696 [-2.445]	22.580 [1.859]	8.048 [2.999]	0.097 0.071
Metals and Mining	0.005 [0.194]	0.062 [0.498]	-0.268 [-1.577]	0.019 [0.944]	11.326 [1.456]	-4.222 [-1.644]	-12.841 [-0.713]	10.819 [2.558]	0.057 0.029
Real Estate	0.009 [0.322]	0.209 [1.636]	0.014 [0.071]	0.025 [1.586]	12.919 [1.475]	-6.274 [-2.430]	0.410 [0.027]	0.952 [0.238]	0.060 0.032
Services–Business	-0.016 [-1.008]	-0.000 [-0.003]	-0.063 [-0.560]	0.015 [1.421]	12.854 [2.469]	-5.919 [-3.927]	17.017 [1.755]	6.758 [2.795]	0.137 0.112
Services–Personal	-0.007 [-0.498]	-0.060 [-0.545]	0.089 [0.845]	-0.000 [-0.045]	7.251 [1.399]	-3.556 [-2.320]	12.329 [1.132]	6.694 [2.111]	0.080 0.053
Textiles & Trade	0.008 [0.394]	0.002 [0.016]	0.066 [0.437]	0.005 [0.385]	7.293 [1.142]	-5.621 [-3.129]	15.949 [1.362]	3.946 [1.561]	0.056 0.028
Transportation	0.003 [0.180]	0.022 [0.200]	-0.080 [-0.653]	0.016 [1.447]	8.731 [1.736]	-5.383 [-3.874]	10.330 [1.088]	4.807 [2.076]	0.075 0.048
Utilities	-0.004 [-0.324]	-0.121 [-1.431]	-0.020 [-0.168]	0.007 [0.868]	8.110 [2.001]	-4.126 [-3.444]	10.181 [1.279]	4.210 [2.172]	0.069 0.042

Multivariate test of predictability

Portfolio	F-statistic	p-value
18 industries	1.6389	0.0001

Table 2 (continued)

Portfolio	Intercept	WRD_{t-1}	$XRG10_{t-1}$	JAN_t	DIV_{t-1}	$E\$30_{t-1}$	$Baa\text{-}Aaa_{t-1}$	$90\text{-}30TB_{t-1}$	$R^2/\bar{R}^2$
C. Country bond returns									
Canada	0.004 [0.368]	-0.069 [-1.480]	-0.069 [-1.032]	-0.007 [-1.011]	1.178 [0.381]	-0.544 [-0.399]	0.520 [0.069]	5.830 [2.430]	0.068 0.041
France	0.017 [1.414]	-0.124 [-1.716]	0.000 [0.001]	-0.005 [-0.630]	6.522 [1.654]	-3.970 [-2.768]	-1.001 [-0.107]	2.905 [1.271]	[0.057 0.030
Germany	0.016 [1.253]	-0.166 [-2.261]	0.051 [0.410]	-0.011 [-1.241]	6.979 [1.637]	-3.765 [-2.724]	-1.216 [-0.136]	3.031 [1.449]	0.063 0.036
Japan	0.016 [1.099]	-0.075 [-0.944]	0.099 [0.756]	-0.015 [-1.453]	4.448 [0.880]	-4.064 [-2.633]	10.466 [0.980]	2.792 [0.991]	0.060 0.033
Netherlands	0.015 [1.209]	-0.163 [-2.151]	0.082 [0.748]	-0.011 [-1.341]	6.893 [1.692]	-3.645 [-2.846]	-2.260 [-0.255]	4.503 [2.263]	0.080 0.053
Switzerland	0.020 [1.656]	-0.146 [-2.140]	0.000 [0.003]	-0.013 [-1.639]	8.042 [1.830]	-4.045 [-3.071]	-6.235 [-0.728]	1.493 [0.780]	0.073 0.046
United Kingdom	0.002 [0.123]	-0.039 [-0.436]	-0.024 [-0.192]	0.007 [0.565]	10.072 [1.760]	-2.962 [-1.747]	-8.055 [-0.658]	3.825 [1.392]	0.031 0.002
United States	-0.003 [-0.303]	-0.165 [-3.116]	-0.036 [-0.437]	-0.008 [-1.196]	2.329 [0.814]	-0.486 [-0.394]	5.329 [0.878]	3.704 [1.858]	0.079 0.052

Multivariate test of predictability

Portfolio	F-statistic	p-value
8 countries	1.3822	0.0335
All assets	1.7037	0.0001

Table 3

Tests for the number of factors that determine expected international asset returns calculated in U.S. dollars in excess of the 30-day U.S. Treasury bill and based on monthly data from 1971:3–1991:09 (247 observations).

Assets	Number of Assets	Number of Factors	χ^2	P-value
Country index returns	18	1	144.09	0.059
	18	2	119.28	0.054
	18	3	99.26	0.032
International industry returns[a]	18	1	126.69	0.298
	18	2	100.91	0.345
	18	3	72.66	0.555
International industry returns[a] (with dividend approximation)	18	1	119.52	0.469
	18	2	91.90	0.599
	16	3	70.38	0.629
Bond returns[b]	8	1	61.56	0.107
	8	2	36.69	0.437
	8	3	19.69	0.763

[a]The industry porfolios are based on an aggregation of 37 Morgan Stanley Capital International industry porfolios. The MSCI indices do not include dividends. Results are presented using an approximation of the dividends based on the same industry groupings of NYSE and AMEX returns.

[b]Bond data are for Canada, France, Germany, Japan, The Netherlands, Switzerland, the United Kingdom and the United States. The bond data are from Lombard Odier & Cie.

Table 4

Variance ratios and average errors for factor model specifications assuming constant risk loadings using data from 1971:3-1991:09 (247 observations).

Portfolio	One factor model APE	One factor model VR1	One factor model VR2	Two factor model APE	Two factor model VR1	Two factor model VR2	Three factor model APE	Three factor model VR1	Three factor model VR2
Country index returns									
Australia	0.082	0.466	0.545	-0.182	0.595	0.369	0.162	1.002	0.016
Austria	0.676	0.006	1.010	0.258	0.724	0.356	0.366	0.712	0.374
Belgium	0.387	0.555	0.827	0.121	0.756	0.289	0.080	0.769	0.262
Canada	-0.035	0.542	0.461	-0.251	0.682	0.238	-0.096	0.868	0.110
Denmark	0.615	0.457	0.678	0.832	0.468	0.691	0.540	0.609	0.580
France	0.356	0.503	0.566	0.132	0.622	0.419	0.086	0.642	0.387
Germany	0.331	0.665	0.441	0.144	0.847	0.229	0.132	0.846	0.225
Hong Kong	1.198	0.370	0.781	0.562	1.039	0.146	0.646	1.009	0.183
Italy	0.152	0.062	0.950	-0.183	0.665	0.267	-0.216	0.727	0.187
Japan	0.623	0.516	0.664	0.515	0.861	0.360	0.493	0.850	0.369
Netherlands	0.363	1.027	0.090	0.236	1.068	0.032	0.238	1.062	0.035
Norway	0.626	0.147	0.956	0.450	0.386	0.800	0.617	0.747	0.364
Singapore/Malaysia	0.588	0.628	0.479	0.248	0.875	0.203	0.438	1.022	0.084
Spain	0.423	0.019	0.960	0.076	0.667	0.354	0.206	0.723	0.307
Sweden	0.760	0.181	0.933	0.405	0.865	0.326	0.353	0.942	0.240
Switzerland	0.172	0.795	0.256	0.114	0.811	0.228	0.158	0.831	0.217
United Kingdom	0.301	0.860	0.205	0.055	0.952	0.066	0.061	0.957	0.059
United States	-0.052	0.879	0.105	-0.097	0.887	0.080	-0.077	0.899	0.076
Average	0.431	0.484	0.589	0.161	0.765	0.303	0.232	0.845	0.226

Table 4 (continued)

Portfolio	One factor model APE	One factor model VR1	One factor model VR2	Two factor model APE	Two factor model VR1	Two factor model VR2	Three factor model APE	Three factor model VR1	Three factor model VR2
Industry returns (no dividends)									
Aerospace	0.549	0.435	0.658	0.408	0.669	0.347	0.132	0.963	0.056
Capital Goods	0.302	0.697	0.298	0.188	0.891	0.118	0.214	0.899	0.108
Chemicals	0.162	0.597	0.399	0.080	0.729	0.275	0.143	0.765	0.234
Communications	0.190	0.875	0.119	0.123	0.980	0.025	0.140	0.985	0.019
Construction	0.451	0.379	0.614	0.286	0.781	0.240	0.368	0.818	0.192
Consumer Durables	0.293	0.523	0.472	0.140	0.855	0.154	0.173	0.866	0.142
Energy	0.322	0.615	0.378	0.264	0.635	0.369	0.344	0.672	0.316
Finance	0.456	0.499	0.492	0.314	0.878	0.146	0.324	0.879	0.144
Food & Tobacco	0.474	0.840	0.146	0.417	0.890	0.120	0.444	0.895	0.104
Forest Products	0.055	0.699	0.300	-0.044	0.786	0.212	0.103	0.908	0.087
Leisure	0.358	0.681	0.312	0.216	0.933	0.078	0.284	0.959	0.047
Metals and Mining	0.254	0.533	0.462	0.132	0.733	0.274	0.191	0.747	0.257
Real Estate	0.604	0.071	0.926	0.382	0.780	0.255	0.354	0.780	0.258
Services-Business	0.252	0.726	0.269	0.119	0.969	0.037	0.127	0.969	0.037
Services-Persona	0.349	0.737	0.253	0.270	0.854	0.158	0.309	0.866	0.137
Textiles & Trade	0.543	0.504	0.465	0.427	0.763	0.264	0.434	0.763	0.263
Transportation	0.286	0.525	0.469	0.161	0.852	0.161	0.163	0.854	0.157
Utilities	0.019	0.874	0.125	-0.002	0.886	0.114	-0.022	0.891	0.109
Average	0.329	0.600	0.393	0.216	0.826	0.186	0.236	0.860	0.148

Table 4 (continued)

Portfolio	One factor model			Two factor model			Three factor model		
	APE	VR1	VR2	APE	VR1	VR2	APE	VR1	VR2
Industry returns (with dividend approximation)									
Aerospace	0.302	0.480	0.569	0.168	0.822	0.213	0.020	0.925	0.080
Capital Goods	0.092	0.656	0.367	0.014	0.848	0.156	0.091	0.933	-0.090
Chemicals	0.145	0.511	0.527	0.091	0.650	0.379	0.183	0.796	0.246
Communications	0.255	0.938	0.168	0.208	1.051	0.044	0.239	1.084	0.017
Construction	0.519	0.365	0.724	0.414	0.730	0.364	0.539	0.892	0.195
ConsumerDurables	0.116	0.489	0.534	0.010	0.823	0.180	0.099	0.919	0.101
Energy	0.271	0.611	0.454	0.237	0.627	0.435	0.333	0.743	0.323
Finance	0.516	0.506	0.609	0.415	0.902	0.219	0.486	0.977	0.142
Food & Tobacco	0.396	0.904	0.237	0.366	0.954	0.186	0.413	1.027	0.114
Forest Products	-0.088	0.584	0.397	-0.138	0.641	0.325	0.025	0.923	0.081
Leisure	0.118	0.600	0.426	0.029	0.821	0.186	0.149	0.972	0.058
Metals and Mining	0.147	0.387	0.638	0.060	0.624	0.390	0.184	0.727	0.301
Real Estate	0.674	0.065	0.979	0.520	0.777	0.294	0.585	0.808	0.254
Services-Busines	0.011	0.636	0.367	-0.081	0.887	0.090	-0.007	0.955	0.044
Services-Persona	0.195	0.727	0.335	0.144	0.831	0.222	0.216	0.930	0.134
Textiles & Trade	0.453	0.546	0.565	0.373	0.806	0.308	0.435	0.869	0.244
Transportation	0.302	0.527	0.550	0.217	0.867	0.203	0.276	0.927	0.149
Utilities	0.299	0.979	0.149	0.269	1.013	0.112	0.266	1.011	0.114
Average	0.262	0.584	0.477	0.184	0.815	0.239	0.251	0.912	0.149

Table 4 (continued)

Portfolio	One factor model APE	One factor model VR1	One factor model VR2	Two factor model APE	Two factor model VR1	Two factor model VR2	Three factor model APE	Three factor model VR1	Three factor model VR2
Bond returns									
Canada	-0.108	0.580	0.356	-0.106	0.612	0.326	-0.038	0.874	0.110
France	0.005	0.686	0.316	-0.004	0.944	0.054	-0.040	0.962	0.016
Germany	0.078	0.733	0.306	0.069	1.026	0.009	0.050	1.026	0.001
Japan	0.189	0.852	0.249	0.186	0.878	0.221	0.109	0.990	0.080
Netherlands	0.015	0.818	0.189	0.008	0.984	0.019	-0.013	0.984	0.009
Switzerland	0.027	0.443	0.568	0.014	0.964	0.042	-0.011	0.971	0.025
United Kingdom	-0.019	0.696	0.293	-0.023	0.765	0.222	-0.057	0.788	0.175
United States	-0.048	0.422	0.554	-0.048	0.450	0.525	0.031	0.990	0.023
Average	0.018	0.654	0.354	0.012	0.828	0.177	0.007	0.948	0.055

The industry porfolios are based on an aggregation of 37 Morgan Stanley Capital International industry porfolios. Returns are measured in excess of the 30-day U.S. Treasury bill. The bond data are from Lombard Odier & Cie. APE is the average pricing error (percent per month). VR1 is the ratio of the variance of the model expected returns (produced by the model estimation) to the variance of the expected returns generated by a linear regression of the asset returns on the instrumental variables. VR2 is the ratio of the variance of the expected model residuals (produced by a linear regression of the model residuals on the instrumental variables) to the variance of the expected returns generated by a linear regression of the asset returns on the instrumental variables.

Table 5

A comparision of the coefficients resulting from a regression of the prespecified macroeconomic factors on the instrumental variables and of the latent factors on the instrumental variables. These macroeconomic factors are: the excess return on the Morgan Stanley world market porfolio (in U.S. dollars), the excess return on foreign exchange investment in 10 countries (XRG10), the change in the price of oil, the change in OECD industrial production, and the change in OECD inflation. The instrumental variables are: a constant, the return on the world index (WRD), the excess return on foreign exchange investment in 10 countries (XRG10), the dividend yield on the Standard and Poor's 500 stock index (DIV), a dummy variable for the month of January (JAN), the return on a 30-day Eurodollar deposit (E$30), the yield on Moody's Baa rated bonds less the yield on Moody's Aaa rated bonds (Baa-Aaa) and the return for holding a 90 day U.S. Treasury bill for one month less the return on a 30 day bill (3-1BILL). Estimates are based on monthly data from 1971:3–1991:09 (247 observations).

Factors	Intercept	WRD_{t-1}	$XRG10_{t-1}$	JAN_t	DIV_{t-1}	$E\$30_{t-1}$	$Baa\text{-}Aaa_{t-1}$	$3\text{-}1BILL_{t-1}$	$R^2/\bar{R}^2$
Prespecified: OLS coefficients and t-statistics									
World excess market return	-0.009 [-0.651]	-0.027 [-0.350]	-0.009 [-0.091]	0.015 [1.477]	9.880 [2.323]	-3.584 [-4.756]	15.197 [1.955]	5.238 [2.770]	0.136 0.111
Excess return on G10 FX index	0.021 [2.523]	-0.087 [-1.843]	-0.008 [-0.101]	-0.012 [-2.071]	2.492 [0.919]	-2.286 [-2.862]	-8.450 [-1.475]	2.071 [1.629]	0.085 0.058
Change in oil price	0.054 [1.479]	-0.062 [-0.318]	0.028 [0.072]	0.041 [0.586]	-5.450 [-0.454]	4.749 [1.521]	-52.749 [-1.977]	-2.730 [-0.673]	0.033 0.005
Change in OECD industrial production	0.012 [5.486]	0.020 [1.454]	-0.057 [-2.954]	-0.000 [-0.236]	-1.300 [-1.494]	-0.373 [-1.803]	-1.877 [-1.130]	-0.581 [-1.951]	0.141 0.115
Change in OECD consumer prices	0.002 [2.188]	-0.002 [-0.323]	-0.002 [-0.273]	0.002 [2.036]	1.278 [4.325]	0.378 [4.090]	-2.827 [-4.443]	-0.128 [-0.860]	0.272 0.250
Latent coefficients using country returns									
Factor 1	-0.013	0.019	0.029	0.004	15.848	-7.987	19.910	5.616	-
Factor 2	0.009	0.177	0.114	0.036	-1.106	-0.346	-3.949	1.963	-
Latent coefficients using international industry returns (without dividends)									
Factor 1	-0.024	0.062	-0.137	0.002	13.519	-5.830	15.420	7.125	-
Factor 2	-0.021	0.034	-0.130	-0.001	11.952	-5.306	14.059	6.785	-
Latent coefficients using international industry returns (with dividend approximation)									
Factor 1	-0.033	0.092	-0.177	-0.005	20.706	-7.406	16.996	5.210	-
Factor 2	-0.022	0.024	-0.153	-0.009	15.201	-5.860	14.571	5.129	-
Latent coefficients using international bond returns									
Factor 1	0.003	-0.049	0.049	0.007	0.995	-1.436	3.766	4.059	-
Factor 2	0.012	-0.181	0.041	0.008	6.541	-4.139	-1.018	3.225	-

Heteroskedasticity-consistent t-ratios are in brackets. The industry porfolios are based on an aggregation of 37 Morgan Stanley Capital International industry porfolios. Bond data are for Canada, France, Germany, Japan, The Netherlands, Switzerland, the United Kingdom and the United States. The bond data are from Lombard Odier & Cie.

Table 6

Characterizing the factor premiums that determine expected international asset returns. Unconditional correlations of the factor premiums and the fitted expected values of five prespecified macroeconomic factors. These macroeconomic factors are: the excess return on the Morgan Stanley world market porfolio (in U.S. dollars), the excess return on foreign exchange investment in 10 countries (XRG10), the change in the price of oil, the change in OECD industrial production, and the change in OECD inflation. Expected values are obtained by projecting on the instrumental variables. Estimates are based on monthly data from 1971:3–1991:09 (247 observations).

Factor estimate	World market premium	FX return premium	Oil premium	OECD production premium	OECD inflation premium	Multiple correlation
			Country index returns			
Factor 1 premium	0.952	0.289	-0.710	-0.167	-0.512	0.9820
Factor 2 premium	0.307	-0.047	-0.253	0.079	-0.333	0.4316
			International industry returns (without dividends)			
Factor 1 premium	0.926	0.175	-0.746	-0.230	-0.396	0.9832
Factor 2 premium	0.932	0.205	-0.738	-0.248	-0.394	0.9846
			International industry returns (with dividend approximation)			
Factor 1 premium	0.895	0.138	-0.735	-0.204	-0.321	0.9916
Factor 2 premium	0.910	0.198	-0.725	-0.235	-0.326	0.9882
			International bond returns			
Factor 1 premium	0.834	0.428	-0.504	-0.354	-0.450	0.9584
Factor 2 premium	0.619	0.810	-0.179	0.019	-0.329	0.9518

The industry porfolios are based on an aggregation of 37 Morgan Stanley Capital International industry porfolios. Bond data are for Canada, France, Germany, Japan, The Netherlands, Switzerland, the United Kingdom and the United States. The bond data are from Lombard Odier & Cie.

Table 7

Variance ratios and average errors for a conditional asset pricing model with two prespecified factors and assuming constant risk loadings using data from 1971:3–1991:09 (247 observations).

Country index returns

Portfolio	APE	VR1	VR2
Australia	0.191	0.218	0.445
Austria	0.235	0.300	0.692
Belgium	0.166	0.430	0.291
Canada	0.020	0.407	0.348
Denmark	0.419	0.331	0.551
France	0.035	0.530	0.581
Germany	0.019	0.718	0.805
Hong Kong	1.012	0.228	0.489
Italy	-0.260	0.760	0.832
Japan	0.529	0.603	0.729
Netherlands	0.301	0.391	0.273
Norway	0.283	0.494	0.460
Singapore/Malaysia	0.663	0.222	0.331
Spain	-0.064	0.447	0.918
Sweden	0.479	0.487	0.863
Switzerland	-0.044	0.522	0.293
United Kingdom	0.233	0.336	0.272
United States	0.106	0.583	0.196
Average	0.240	0.445	0.520

χ^2	Degrees of freedom	P-value
164.12	144	0.120

International industry returns (without dividends)

Portfolio	APE	VR1	VR2
Aerospace & Military Technology	0.606	0.407	0.316
Capital Goods	0.314	0.915	0.127
Chemicals	0.167	1.179	0.212
Communications	0.525	0.999	0.225
Construction	0.320	0.730	0.102
Consumer Durables	0.257	0.634	0.207
Energy	0.594	0.989	0.661
Finance	0.500	1.060	0.172
Food & Tobacco	0.561	1.028	0.232
Forest Products	0.159	0.891	0.331
Leisure	0.429	0.958	0.176

Table 7 (continued)

International industry returns (continued)

Portfolio	APE	VR1	VR2
Metals and Mining	0.057	0.501	0.317
Real Estate	0.282	0.639	0.494
Services-Business	0.149	0.539	0.131
Services-Personal	0.352	1.002	0.202
Textiles & Trade	0.423	1.142	0.095
Transportation	0.339	0.988	0.066
Utilities	0.353	0.636	0.195
Average	0.355	0.847	0.237

χ^2	Degrees of freedom	P-value
152.43	144	0.299

International industry returns (with dividends)

Portfolio	APE	VR1	VR2
Aerospace & Military Technology	-0.323	0.416	0.255
Capital Goods	-0.698	1.085	0.245
Chemicals	-0.771	1.452	0.297
Communications	-0.542	1.085	0.246
Construction	-0.470	1.089	0.286
Consumer Durables	-0.612	0.801	0.413
Energy	-0.655	1.179	1.084
Finance	-0.571	1.315	0.301
Food & Tobacco	-0.348	1.175	0.220
Forest Products	-0.984	1.064	0.471
Leisure	-0.695	1.114	0.364
Metals and Mining	-0.571	0.752	0.517
Real Estate	-0.363	0.922	0.682
Services-Business	-0.590	0.683	0.193
Services-Personal	-0.492	1.205	0.328
Textiles & Trade	-0.475	1.437	0.329
Transportation	-0.647	1.253	0.295
Utilities	-0.561	0.865	0.072
Average	-0.576	1.050	0.367

χ^2	Degrees of freedom	P-value
163.39	144	0.128

Table 7 (continued)

International bond returns

Portfolio	APE	VR1	VR2
Canada	0.106	0.238	0.897
France	0.371	1.913	1.282
Germany	0.471	2.257	1.372
Japan	0.604	1.588	1.666
Netherlands	0.453	1.762	1.193
Switzerland	0.341	1.799	0.897
United Kingdom	0.286	1.511	1.414
United States	0.126	0.206	1.227
Average	0.345	1.409	1.244

χ^2	Degrees of freedom	P-value
86.49	64	0.032

The industry porfolios are based on an aggregation of 37 Morgan Stanley Capital International industry porfolios. The bond data are from Lombard Odier & Cie. APE is the average pricing error (percent per month). VR1 is the ratio of the variance of the model expected returns (produced by the model estimation) to the variance of the expected returns generated by a linear regression of the asset returns on the instrumental variables. VR2 is the ratio of the variance of the expected model residuals (produced by a linear regression of the model residuals on the instrumental variables) to the variance of the expected returns generated by a linear regression of the asset returns on the instrumental variables.

Table 8

The relative importance of the two sources for risk in the latent factor model as well as the model with two prespecified sources of risk. The prespecified sources of risk are the excess return on the Morgan Stanley Capital International world index and the excess return on an index of foreign currency investment in ten countries. The data are from 1971:3–1991:09 (247 observations).

Portfolio	Latent factor model: Factor 1 risk loading	Latent factor model: Factor 2 risk loading	Latent factor model: Proportion of variance due to factor 1	Latent factor model: Proportion of variance due to factor 2	Prespecified factor model: Factor 1 risk loading	Prespecified factor model: Factor 2 risk loading	Prespecified factor model: Proportion of variance due to factor 1	Prespecified factor model: Proportion of variance due to factor 2
Country index returns								
Australia	1.000	0.000	1.000	0.000	1.038	-0.156	1.007	0.011
Austria	0.000	1.000	0.000	1.000	0.309	0.931	0.178	0.761
Belgium	0.640	0.313	0.717	0.081	0.727	0.731	0.627	0.300
Canada	0.716	0.096	0.921	0.008	1.023	-0.324	0.989	0.047
Denmark	0.462	-0.292	1.212	0.211	0.526	0.598	0.574	0.351
France	0.635	0.219	0.797	0.045	0.950	0.735	0.729	0.206
Germany	0.520	0.175	0.801	0.043	0.656	0.861	0.508	0.415
Hong Kong	0.687	1.104	0.316	0.388	1.128	0.297	0.941	0.031
Italy	0.108	0.781	0.033	0.828	0.759	0.512	0.773	0.167
Japan	0.578	0.427	0.594	0.153	1.008	0.508	0.849	0.102
Netherlands	0.641	-0.154	1.098	0.017	0.855	0.369	0.879	0.078
Norway	0.238	0.256	0.462	0.253	0.948	0.161	0.968	0.013
Singapore/Malaysia	0.909	0.275	0.821	0.036	1.199	-0.304	0.999	0.030
Spain	-0.153	0.899	0.071	1.170	0.703	0.359	0.846	0.105
Sweden	0.242	0.745	0.138	0.617	0.774	0.318	0.887	0.071
Switzerland	0.730	-0.273	1.175	0.078	0.813	0.650	0.718	0.217
United Kingdom	1.025	0.038	0.979	0.001	1.173	0.311	0.940	0.031
United States	0.781	-0.330	1.189	0.100	1.022	-0.528	0.937	0.118
Average	0.555	0.293	0.685	0.279	0.867	0.335	0.797	0.170
International industry returns[a] (no dividends)								
Aerospace & Military Technology	1.000	-	1	0	0.891	-0.053	0.997	0.001
Capital Goods	0.985	-	1	0	1.028	-0.042	0.998	0.000
Chemicals	0.751	-	1	0	0.965	0.181	0.997	0.008
Communications	0.797	-	1	0	0.750	-0.140	0.986	0.008
Construction	0.781	-	1	0	0.948	0.716	0.898	0.122
Consumer Durables	0.924	-	1	0	0.929	0.095	1.000	0.003
Energy	0.906	-	1	0	1.008	-0.277	0.975	0.017
Finance	0.794	-	1	0	1.062	0.324	0.987	0.022
Food & Tobacco	0.870	-	1	0	0.842	0.116	1.000	0.005
Forest Products	1.051	-	1	0	1.068	-0.139	0.992	0.004
Leisure	1.064	-	1	0	1.083	-0.228	0.984	0.010
Metals and Mining	0.905	-	1	0	0.841	0.613	0.905	0.114
Real Estate	0.370	-	1	0	1.014	0.764	0.899	0.121
Services–Business	1.067	-	1	0	0.855	0.201	0.994	0.013
Services–Personal	0.834	-	1	0	0.864	0.048	1.001	0.001
Textiles & Trade	0.774	-	1	0	1.053	0.345	0.984	0.025
Transportation	0.758	-	1	0	0.965	0.213	0.995	0.011
Utilities	0.765	-	1	0	0.587	0.407	0.914	0.104
Average	0.855	-	1	0	0.931	0.174	0.973	0.033

Table 8 (continued)

Portfolio	Latent factor model				Prespecified factor model			
	Factor 1 risk loading	Factor 2 risk loading	Proportion of variance due to factor 1	Proportion of variance due to factor 2	Factor 1 risk loading	Factor 2 risk loading	Proportion of variance due to factor 1	Proportion of variance due to factor 2
	International bond returns							
Canada	1.000	0.000	1.000	0.000	0.251	0.159	1.001	0.380
France	0.000	1.000	0.000	1.000	0.105	0.971	0.013	1.060
Germany	-0.035	1.079	0.000	1.029	0.089	1.107	0.007	1.046
Japan	0.847	0.561	0.272	0.312	0.151	0.998	0.026	1.077
Netherlands	0.327	0.971	0.032	0.747	0.083	1.039	0.007	1.046
Switzerland	-0.646	1.273	0.151	1.535	0.044	1.002	0.002	1.026
United Kingdom	0.437	0.567	0.119	0.525	0.220	0.782	0.092	1.105
United States	0.352	0.345	0.173	0.435	0.251	0.060	1.099	0.060
Average	0.285	0.725	0.219	0.698	0.149	0.764	0.281	0.850

The proportion of variance due to the factor is calculated as the unconditional variance of the factor beta times the factor premium divided by the unconditional variance of the fitted returns from the model.

[a]The one-latent factor results are reported.

Number	Industry portfolio	MSCI composition
1	Aerospace & Military Technology	Aerospace & Military Technology
2	Capital Goods	Electronic Components & Instruments Industrial Components Machinery & Engineering
3	Chemicals	Chemicals
4	Communications	Broadcasting Telecommunications
5	Construction	Building Materials & Components Construction & Housing
6	Consumer Durables	Appliances & Household Durables Automobiles Electrical & Electronics
7	Energy	Energy Equipment & Services Energy Sources
8	Finance	Banking Financial Services Insurance
9	Food & Tobacco	Beverages & Tobacco Food & Household Products
10	Forest Products	Forest Products & Paper
11	Leisure	Leisure & Tourism Recreation, Other Consumer Goods
12	Metals & Mining	Gold Mines Metals (Non-Ferrous) Metals (Steel) Misc. Materials & Commodities
13	Real Estate	Real Estate
14	Services–Business	Business & Public Services Data Processing & Reproduction
15	Services–Personal	Health & Personal Care
16	Textiles & Trade	Mechandising Textiles & Apparel Wholesale & International Trade
17	Transportation	Transportation–Airlines Transportation–Road & Rail Transportation–Shipping
18	Utilities	Utilities–Electrical & Gas

Fig. 1. Composition of the international industry portfolios

Based on an aggregation of 37 Morgan Stanley Capital International industry porfolios. Each of the 37 MSCI portfolios are value weighted. The aggregated portfolios represent returns to a portfolio that starts with an equally-weighted investment in the MSCI categories in December 1969.

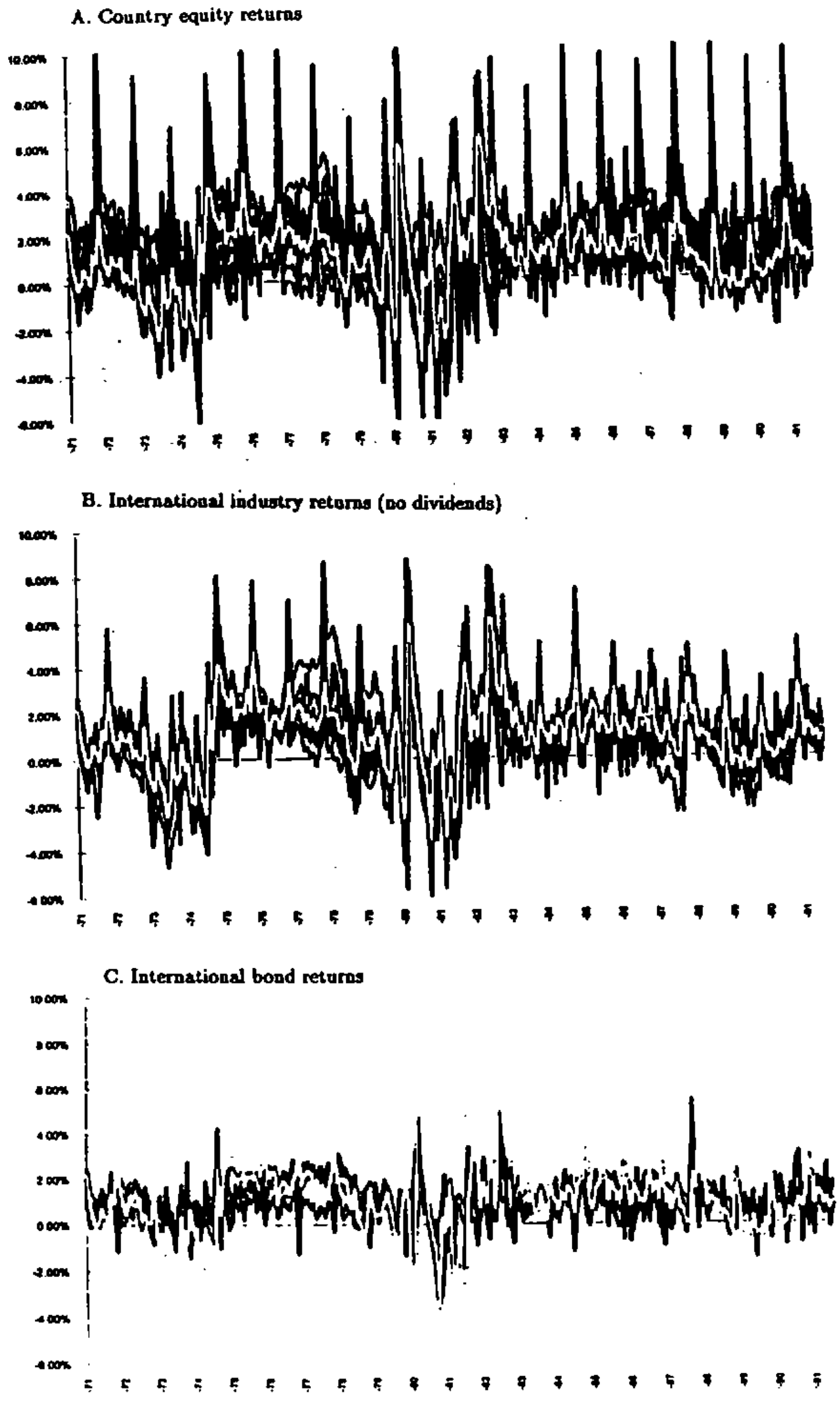

Fig. 2. The comovement of expected international asset returns.

The solid lines represent fitted values from regressions of the asset returns on the instrumental variables. The clear line represents the fitted value from regressing the MSCI world return on the instruments.

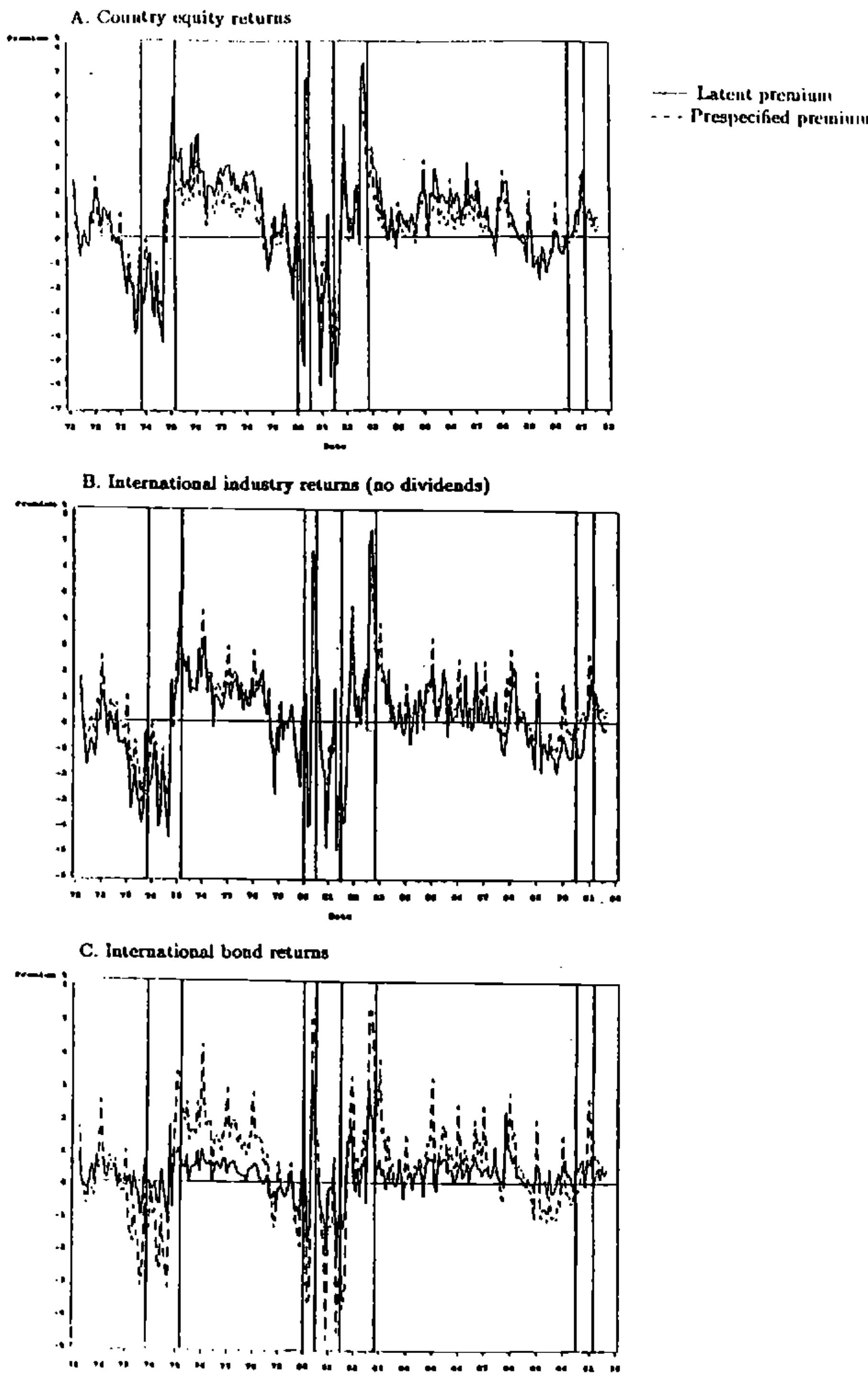

Fig. 3. Characterizing the first latent risk premium.

The solid line represent the premium associated with the first latent factor in a two factor model. The dashed line represents the fitted values from regressing the the Morgan Stanley Capital International (MSCI) world return in excess of the 30-day Treasury bill on the instrumental variables.

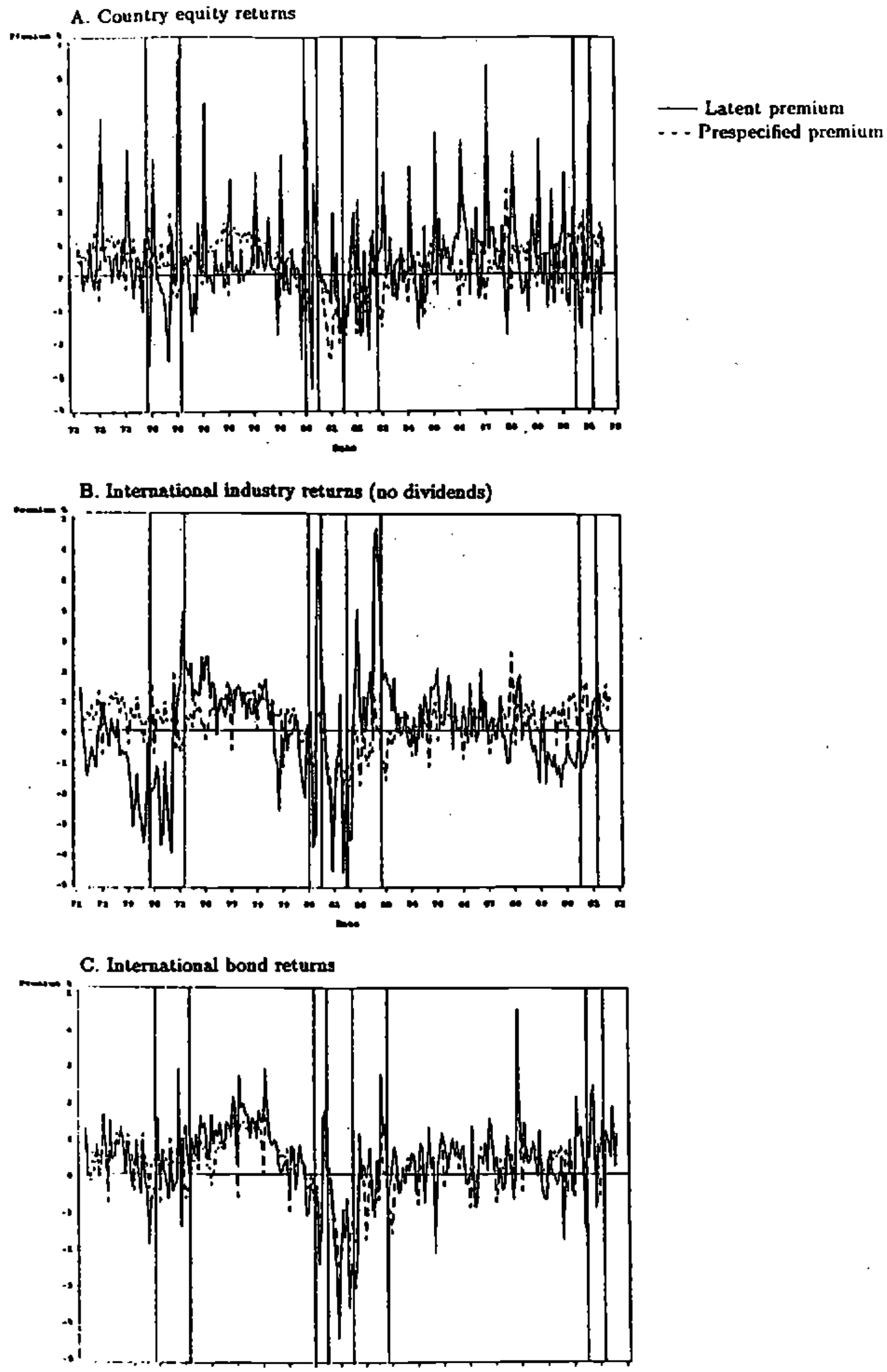

Fig. 4. Characterizing the second latent risk premium.

The solid line represent the premium associated with the second latent factor in a two factor model. The dashed line represents the fitted values from regressing the return on a trade weighted currency investment in 10 countries in excess of the 30-day Treasury bill on the instrumental variables.

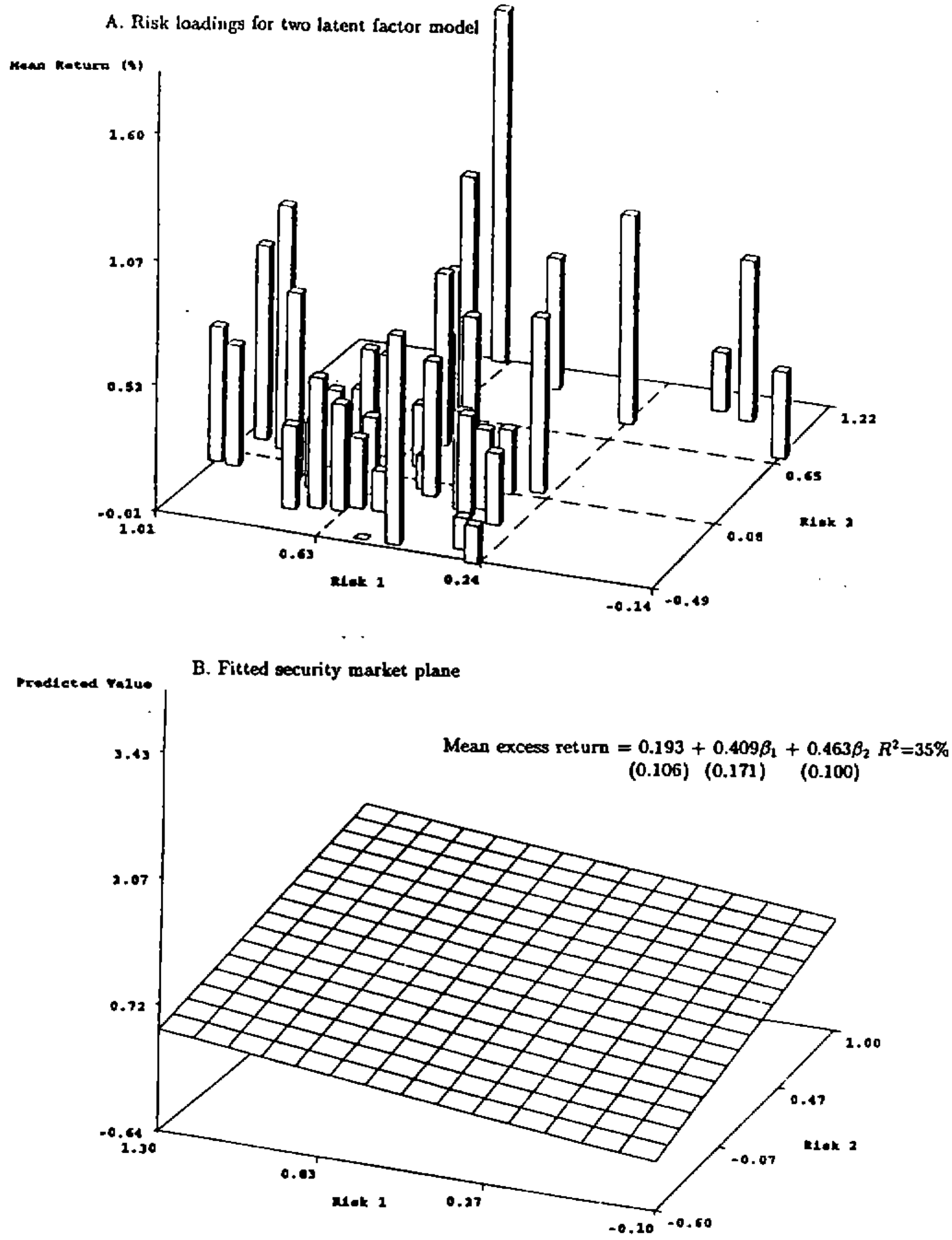

Fig. 5. The cross-section of average returns and risk.

The pillars represent the risk loadings for the first two factors in the two factor latent estimation. In contrast to the results presented in the paper, this estimation simultaneous considers all 44 assets. The average returns are in excess of the 30-day Treasury bill. The security market plane are the fitted values from the regression of the average returns on the betas.