Time-Varying World Market Integration

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ABSTRACT

We propose a measure of capital market integration arising from a conditional regime-switching model. Our measure allows us to describe expected returns in countries that are segmented from world capital markets in one part of the sample and become integrated later in the sample. We find that a number of emerging markets exhibit time-varying integration. Some markets appear more integrated than one might expect based on prior knowledge of investment restrictions. Other markets appear segmented even though foreigners have relatively free access to their capital markets. While there is a perception that world capital markets have become more integrated, our country-specific investigation suggests that this is not always the case.

WHY DO DIFFERENT COUNTRIES' market indices command different expected returns? This question lies at the foundation of international finance. The answer follows from another question: What makes international finance different from finance in general? In studying assets in the United States, we would say that differing expected returns are due to differing risk exposures. In international markets, the answer is more difficult. Aside from the obvious complications arising from country-specific exchange rates, "risk" is hard to quantify if a country is not fully integrated into world capital markets.

Markets are completely integrated if assets with the same risk have identical expected returns irrespective of the market. Risk refers to exposure to some common world factor. If a market is segmented from the rest of the world, its covariance with a common world factor may have little or no ability to explain its expected return.

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The reward to risk is also an important consideration. In integrated world capital markets, there are common rewards to risk associated with risk exposures. In explaining the cross-section of expected returns, the reward to risk is not important because it is common to all the integrated countries. However, in segmented markets, the rewards to risk may not be the same because the sources of risk are different.

Asset pricing studies can be classified in three broad categories: segmented markets, integrated markets, or partially segmented markets. An example of an asset pricing study that assumes markets are segmented is one that "tests" a model like the Capital Asset Pricing Model (CAPM) of Sharpe (1964), Lintner (1965) and Black (1972), using one country's data. Indeed, all of the seminal U.S. asset pricing studies assume that the United States is a completely segmented market—or that the market proxy represents a broader world market return. While this might have been a reasonable working assumption through the 1970s, in the 1980s the U.S. equity capitalization dropped below 50 percent of the world market capitalization. Indeed, Japan's market capitalization exceeded the United States (albeit briefly) in 1989.

The second class of asset pricing studies assumes that world capital markets are perfectly integrated. These include studies of a world CAPM (see Harvey (1991) and references therein), a world CAPM with exchange risk (see Dumas and Solnik (1995) and Dumas (1994)), a world consumption-based model (Wheatley (1988)), world arbitrage pricing theory (see Solnik (1983) and Cho, Eun, and Senbet (1986)), world multibeta models (Ferson and Harvey (1993, 1994a, 1994b)) and world latent factor models (Campbell and Hamao (1992), Bekaert and Hodrick (1992) and Harvey, Solnik, and Zhou (1994)). Rejection of these models can be viewed as a rejection of the fundamental asset pricing model, inefficiency in the market, or rejection of market integration.

A good example of the difficulty in interpreting the joint hypotheses is presented in Harvey (1991). Using data through May 1989, Harvey finds that the conditionally expected returns in Japan are too high to be explained by asset pricing theory, or that the risk exposure was too small. In multivariate tests, the asset pricing model is not rejected. Is the rejection in Japan a result of using a one factor model, a function of Japanese stock prices deviating from their fundamental values (inefficiency), or an implication of imposing the null hypothesis of complete market integration?

Yet another strand of the literature falls in between segmentation and integration—the so-called mild segmentation model (see Errunza, Losq, and Padmanabhan (1992) and references therein). The advantage of these models is that the polar segmented/integrated cases are not assumed. The disadvantage of these models is that the degree of segmentation is fixed through time. This runs counter to the intuition (as do the polar cases) that some markets have become more integrated through time.

Our contribution is to propose a methodology that allows for the degree of market integration to change through time. While this method can be applied to a general multifactor model, the intuition can be readily obtained in a one factor setting. We allow conditionally expected returns in any country to be affected by their covariance with a world benchmark portfolio *and by* the variance of the country returns. In a perfectly integrated market, only the covariance counts. In segmented markets, the variance is the relevant measure of market risk. While our approach is not directly implied by any current asset pricing theory, it has the appeal of nesting, as special cases, two previous approaches to international asset pricing: complete segmentation and complete integration.

Our integration measure is a time-varying weight that is applied to the covariance and the variance. The model allows for a differing price of variance risk across countries, which depends on country-specific information, and a world price of covariance risk, which depends only on global information. The model is conditional in the sense that predetermined information is allowed to affect the expected returns, covariances, variances, and the integration measure.

Our procedure allows us to recover fitted values for the integration measure so that the degree and trend of a particular market's integration can be depicted through time. However, caution must be exercised in interpreting our results. First, our tests use the simplest asset pricing framework—the one factor model. Omitted factors may induce variation in the integration measure that is not related to market integration. Second, we conduct a battery of specification tests which suggest that the model is rejected in most countries. However, no one has ever attempted to estimate the degree and variation through time of capital market integration. In addition, the test rejections do not necessarily undermine the interpretation of our integration measure. Indeed, we think that our results are useful and interesting. In many countries, variation in the integration measure coincides with capital market reforms. In contrast to general perceptions that markets are becoming more integrated, our results suggest that some countries are becoming less integrated into the world market.

Our article is organized as follows. In Section I, the asset pricing framework is presented. An outline of the econometric model is also detailed. The data on 12 emerging equity markets are described in Section II. In Section III, the results are analyzed. Specification tests are conducted in Section IV. The final section explores the linkages between financial market integration and economic growth and offers some concluding remarks.

I. Asset Pricing with Time-Varying Market Integration

A. The Model

In completely integrated markets and in the absence of exchange risk, a conditional CAPM of Sharpe (1964) and Lintner (1965) imposes the restrictions:

$$E_{t-1}[r_{i,t}^A] = \lambda_{t-1}\text{cov}_{t-1}[r_{i,t}^A, r_{w,t}], \tag{1}$$

where $E_{t-1}[r_{i,t}^A]$ is the conditionally expected excess return on security A's equity (in country i), r_w is the return on a value-weighted world portfolio, cov_{t-1} is the conditional covariance operator and λ_{t-1} is the conditionally expected world price of covariance risk for time t. The risk-free rate has zero conditional variance because the return is determined at $t-1$. This model is tested in Harvey (1991).

In completely segmented markets and under the same assumptions as equation (1):

$$E_{t-1}\left[r_{i,t}^A\right] = \lambda_{i,t-1}\text{cov}_{t-1}\left[r_{i,t}^A, r_{i,t}\right]. \quad (2)$$

Security A is now priced with respect to its covariance with the return on the market portfolio in country i, r_i, and λ_i is the local price of risk. Aggregating (2) at the national level,

$$E_{t-1}[r_{i,t}] = \lambda_{i,t-1}\text{var}_{t-1}[r_{i,t}]. \quad (3)$$

Merton (1980) argues that λ_i is a measure of the representative investor's relative risk aversion. The model suggests that the expected return in a segmented market is determined by the variance of return in that market times the price of variance. The price of variance will depend on the weighted relative risk aversions of the investors in country i.

Equations (1) and (3) focus on the conditions of complete market integration and segmentation, respectively. Suppose that markets are either fully integrated or fully segmented. When there is a change from market segmentation to market integration (or vice versa), the valuation of the payoffs and, hence, the stochastic process governing returns changes. The switch may be a complete surprise or it may be partially expected.

When market participants expect a switch from a segmented to integrated market in the future, or vice versa, equilibrium expected returns may reflect hedging demands against the switch. In that case, neither equation (1) nor equation (3) describes equilibrium expected returns. Furthermore, it may be incorrect to associate this switch with changes in investment restrictions, since the restrictions may not have been binding.

We approach this difficult problem empirically using a regime-switching model. Let S_t^i be an unobserved state variable that takes on the value of one when markets are integrated and a value of two when markets are segmented. In the first regime, returns are drawn from a distribution with conditional mean given by equation (1). In the second regime, the conditional mean of returns is given by equation (3). At each point in time, there may be a positive probability of a regime switch that is governed by switching probabilities.

From the viewpoint of the econometrician with information set $\mathscr{Z}_{t-1}$, the conditional mean return is given by:

$$E_{t-1}[r_{i,t}] = \phi_{i,t-1}\lambda_{t-1}\text{cov}_{t-1}[r_{i,t}, r_{w,t}] + (1-\phi_{i,t-1})\lambda_{i,t-1}\text{var}_{t-1}[r_{i,t}], \quad (4)$$

where the parameter $\phi_{i,t-1}$, which falls in the interval $[0,1]$, is the econometrician's time-varying assessment of the likelihood that the market is integrated. It can be interpreted as the conditional probability of being in regime 1, $\phi_{i,t-1} = \text{prob}[S_t^i = 1|\mathscr{Z}_{t-1}]$. Although equation 4 does not necessarily reflect equilibrium expected returns of the market participants, it may provide a reasonable approximation to expected returns in this setting.

To infer $\phi_{i,t-1}$ from the data, we explore two different regime switching models. The first is the standard Hamilton (1989, 1990) model. Here, S_t^i follows a Markov process with constant transition probabilities. Although the switching probabilities are time invariant, the regime probability, $\phi_{i,t-1}$, and hence the degree of market integration, varies through time as new information changes the econometrician's inference on the relative likelihood of the two regimes. Gray (1995a) derives the following recursive representation for the regime probability:

$$\phi_{t-1} = (1 - Q) + (P + Q - 1)\left[\frac{f_{1,t-1}\phi_{t-2}}{f_{1,t-1}\phi_{t-2} + f_{2,t-1}(1 - \phi_{t-2})}\right]. \tag{5}$$

where the country i subscript has been suppressed and

$$P = \text{prob}[S_t = 1|S_{t-1} = 1]$$
$$Q = \text{prob}[S_t = 2|S_{t-1} = 2]$$

and $f_{j,t}$ is the likelihood at time t conditional on being in regime j and time $t - 1$ information, $\mathscr{Z}_{t-1}$.

Diebold, Lee, and Weinbach (1992), Ghysels (1993), and Gray (1995a, 1995b) extended the Hamilton model to allow for time-varying transition probabilities. In the second formulation, we allow the transition probabilities P and Q to be time varying, modeling them as logistic functions of $\mathbf{Z}_{t-1}^*$:

$$P_t = \frac{\exp(\boldsymbol{\beta}_1'\mathbf{Z}_{t-1}^*)}{1 + \exp(\boldsymbol{\beta}_1'\mathbf{Z}_{t-1}^*)}$$
$$Q_t = \frac{\exp(\boldsymbol{\beta}_2'\mathbf{Z}_{t-1}^*)}{1 + \exp(\boldsymbol{\beta}_2'\mathbf{Z}_{t-1}^*)} \tag{6}$$

where $\boldsymbol{\beta}_j$, $j = 1, 2$, are vectors of parameters.

In implementing this model, we let $\mathbf{Z}_{t-1}^*$ be a subset of $\mathbf{Z}_{t-1}^i$ where $\mathbf{Z}_{t-1}^i$ is a collection of information variables specific to country i. $\mathbf{Z}_{t-1}^*$ includes lagged dividend yields and lagged equity market capitalization as a proportion of GDP. Since all of these variables might be influenced by a change in policies affecting market integration, they should influence the switching probabilities. For example, dividend yields typically decrease and market capitalization to GDP typically increases when markets become integrated. Although it is possible that global information variables are also important in determining the switching probabilities, we only allow the global variables to influence the probabilities indirectly—through their correlation with the local information variables.

Cumby and Khanthavit (1992) also investigate a standard Hamilton model for equity returns in Korea, Taiwan, and Thailand. Although they do not formulate an explicit model of time-varying integration, they attempt to relate their results to the capital market policies followed in these countries. Below, we will compare our results to theirs.

Following models like Stulz (1981b), the returns in equations (1) to (4) should be real. Given that reliable inflation data in many of the countries that we study are not available and given a lack of short-term interest rate data (to form local excess returns), we choose to calculate the local market volatility in U.S. dollar terms. The excess return should approximate a real return.

B. An Alternative Interpretation

There is another interpretation of equation (4). In addition to being used by an econometrician attempting to infer whether a certain capital market is integrated or segmented, (4) could be viewed within the context of a one-factor partially-segmented world asset pricing model. As Gray (1994a) emphasizes, the regime switching model in equation (5) is a special case of a general finite mixture distribution model with time-varying weights, i.e. $\phi_{i,t-1} = \phi_i(\mathbf{Z}^*_{t-1})$ with $\phi_i(\cdot)$ a functional form that constrains $\phi_{i,t-1}$ to be between zero and one and $\mathbf{Z}^*_{t-1}$ a set of variables in $\mathscr{Z}_{t-1}$. Rather than the outcome of a regime switching model, equation (4) may be viewed as an imperfect approximation of expected returns in partially segmented markets.

Whether a market is integrated with world capital markets or segmented is greatly influenced by the economic and financial market policies followed by its government or other regulatory institutions. Barriers to investment (by foreigners in local markets or local participants in foreign markets) can take many forms. An obvious example is foreign ownership restrictions, often imposed by developing countries. However, not all barriers to foreign equity investment necessarily segment markets from the world capital market. For instance, Bekaert (1995) shows that the presence of country funds and/or cross-listed securities might serve to effectively integrate markets with the world capital market despite the existence of severe restrictions on direct foreign equity ownership. In general, it is hard to infer the actual degree of market segmentation from the complex set of capital market restrictions in place in a particular country at any one time.[1] However, the regime switching model allows us to infer the degree of market integration. Indeed, $\phi_{i,t-1}$ may be interpreted as a policy weight, varying with policies affecting the degree of market integration.

The nature of the approximation depends on the nature of the model of partial integration that one has in mind. Stulz (1981a), for example, assumes that world investors face a proportional tax on the (absolute) holdings of foreign equities. He shows that expected returns for individual stocks depend

[1] A similar argument is made in the development economics literature. A number of papers examine the relation between financial markets and economic growth. See, for example, King and Levine (1993), Pagano (1993), and Obstfeld (1994). The link between market integration and economic development is sketched in the conclusions.

on the covariance with the world market portfolio, and an intercept that depends on the world beta and on whether the local securities are held short or long by the world investors. Moreover, the model implies that some securities will not be traded internationally.

In models that incorporate foreign ownership restrictions, unrestricted stocks may be priced globally, whereas for restricted stocks both the covariance with the world market portfoilo and the covariance with part of the local market is priced. It is tempting to conclude that the parameter $\phi_{i,t-1}$ will reflect the relative market capitalization of unrestricted versus restricted shares. However, this is not necessarily true. The $\phi_{i,t-1}$ weight will depend on the covariance structure of the local stocks, since globally priced stocks might have spillover effects on correlated restricted stocks. Similarly, the importance of nontraded stocks in Stulz's (1981a) model, which he shows to be low world beta stocks, will affect the magnitude of $\phi_{i,t-1}$, since these stocks are likely to covary more with the local than with the global market. Importantly, our model lacks an intercept as is implied by the Stulz model. In our specification tests, however, we incorporate an intercept.

In general, the presence of unequal integration of individual shares makes it difficult to apply our model to the pricing of individual stocks. However, our methodology enables us to recover the fitted integration parameters for the market as a whole. This allows us to characterize the path of integration through time for each emerging equity market.

The idea that both the covariance with the world return and the covariance with the local market return affect securities' expected returns reaches back to Stehle (1977). He devises a test of local versus global pricing of individual stocks by modelling expected returns as a function of the covariance with the local market portfolio and of the covariance with the component of the world market portfolio which is orthogonal to the local market portfolio. A recent example of covariance and variance influencing conditionally expected national returns is proposed in Chan, Karolyi, and Stulz (1992) in their study of the United States and Japan. They use the definition of covariance to show, for example, that the conditionally expected U.S. market return is affected by both the covariance with other countries and its own variance. The weights they place on the second moments are derived from actual market shares of the U. S. and Japan in the world market portfolio. While this intuition is critical for modeling the United States and Japan, explicitly using the market share weights is less important for the countries in our emerging markets sample since they are so small.

C. Estimation Issues

C.1. The Likelihood Function

To complete the model described in equation (4), we need an auxiliary assumption on the movement of expected returns on the world equity portfolio. Consequently, we estimate a series of bivariate models for

$\mathbf{R}_{i,t} = [r_{i,t}, r_{w,t}]'$:

$$
\begin{aligned}
r_{i,t} &= \phi_{i,t-1}\lambda_{t-1}\mathrm{cov}_{t-1}[r_{i,t}, r_{w,t}] + (1-\phi_{i,t-1})\lambda_{i,t-1}\mathrm{var}_{t-1}[r_{i,t}] + e_{i,t} \\
r_{w,t} &= \lambda_{t-1}\mathrm{var}_{t-1}[r_{w,t}] + e_{w,t}
\end{aligned}
\tag{7}
$$

Let $\mathbf{e}_t = [e_{i,t}, e_{w,t}]'$ and define $\mathbf{e}_t^I$ ($\mathbf{e}_t^S$) as the disturbance vector under integration (segmentation):

$$
\mathbf{e}_t = \phi_{i,t-1}\mathbf{e}_t^I + (1-\phi_{i,t-1})\mathbf{e}_t^S. \tag{8}
$$

We assume that the residuals are heteroskedastic and that the variance processes under integration and segmentation differ:

$$
\begin{aligned}
E\left[\mathbf{e}_t^I\mathbf{e}_t^{I'}|\mathscr{Z}_{t-1}\right] &= \boldsymbol{\Sigma}_t^I \\
E\left[\mathbf{e}_t^S\mathbf{e}_t^{S'}|\mathscr{Z}_{t-1}\right] &= \boldsymbol{\Sigma}_t^S.
\end{aligned}
\tag{9}
$$

To relate equation (9) to our previous notation, $\mathrm{cov}_{t-1}[r_{i,t}, r_{w,t}]$ is the off-diagonal element of $\boldsymbol{\Sigma}_t^I$ while $\mathrm{var}_{t-1}[r_{it}]$ is the first diagonal element of $\boldsymbol{\Sigma}_t^S$. The conditional variance dynamics are modeled as ARCH(k) following Baba, Engle, Kraft, and Kroner (1989) (BEKK):[2]

$$
\begin{aligned}
\boldsymbol{\Sigma}_t^I &= \mathbf{C}^I + (\mathbf{A}^I)'\left[\sum_{k=1}^{K} w_k(\mathbf{e}_{t-k}\mathbf{e}'_{t-k})\right]\mathbf{A}^I \\
\boldsymbol{\Sigma}_t^S &= \mathbf{C}^S + (\mathbf{A}^S)'\left[\sum_{k=1}^{K} w_k(\mathbf{e}_{t-k}\mathbf{e}'_{t-k})\right]\mathbf{A}^S,
\end{aligned}
\tag{10}
$$

where $\mathbf{C}^I$ and $\mathbf{C}^S$ are symmetric 2×2 matrices, $\mathbf{A}^I$ and $\mathbf{A}^S$ are 2×2 matrices. An advantage of this model of conditional variances is that it guarantees positive definite conditional variance matrices under weak conditions. In addition, the model imposes restrictions across equations and thereby economizes on parameters relative to other multivariate ARCH models.

To further limit parameter proliferation, we impose the additional restrictions:

$$
\begin{aligned}
&\mathbf{C}^I(2,2) = \mathbf{C}^S(2,2), \\
&\mathbf{A}^I(j,j) = \mathbf{A}^S(j,j) \quad \text{for } j = 1,2, \\
&\mathbf{A}^I(1,2) = \mathbf{A}^S(1,2) = 0, \\
&\qquad \text{and} \\
&\mathbf{A}^S(2,1) = 0.
\end{aligned}
\tag{11}
$$

The first, second, and third restrictions make the conditional variance of the world market return independent of the regime. The restriction $\mathbf{A}^I(1,2) = \mathbf{A}^S(1,2) = 0$ ensures that country-specific shocks do not affect the conditional variance of the world market return. The restriction $\mathbf{A}^S(2,1) = 0$ ensures that

[2] Frankel (1982) and Engel and Frankel (1984) are examples of ARCH-M models that impose similar restrictions to ours. However, these models assume perfect capital market integration.

the world market shocks do not affect the conditional variance of the country return when the market is segmented.[3] The dynamics of the conditional variances are constrained to be the same in both regimes. In the estimation, we set $K = 3$ and let $w_k = 2(K + 1 - k)/(K(K + 1))$ as in Engle, Lilien, and Robbins (1987). The resulting weights on the three past residual vectors are 1/2, 1/3, and 1/6, respectively.

The evidence presented in Campbell (1987) and Harvey (1989, 1991) suggests that the price of risk is time varying. In the most general version of the model, we let:

$$\begin{aligned} \lambda_{t-1} &= \exp(\boldsymbol{\delta}'\mathbf{Z}_{t-1}) \\ \lambda_{i,t-1} &= \exp(\boldsymbol{\delta}_i'\mathbf{Z}^i_{t-1}) \end{aligned} \tag{12}$$

where $\mathbf{Z}$ represents global information variables and $\mathbf{Z}^i$ represents a set of local information variables. A similar assumption underlies much of the latent variables literature (Hansen and Hodrick (1983) and Gibbons and Ferson (1985)) and has recently been imposed by Dumas and Solnik (1995) and Dumas (1994). The exponentiation imposes one of the necessary conditions of the asset pricing theory—that the price of risk is positive.

The model is estimated by maximum likelihood assuming normally distributed error terms. The log-likelihood function, apart from some initial conditions, can be written:

$$\begin{aligned} \log L(\mathbf{R}_{i,T}) &= \sum_{t=1}^{T} \log\{\phi_{i,t-1} g_{1,t} + (1 - \phi_{i,t-1}) g_{2,t}\} \\ \text{with} \quad g_{1,t} &= (2\pi)^{-1} |\boldsymbol{\Sigma}_t^I|^{-1/2} \exp\left\{ -\frac{1}{2}\left(\mathbf{e}_t^{I\prime} (\boldsymbol{\Sigma}_t^I)^{-1} \mathbf{e}_t^I\right)\right\}, \\ g_{2,t} &= (2\pi)^{-1} |\boldsymbol{\Sigma}_t^S|^{-1/2} \exp\left\{ -\frac{1}{2}\left(\mathbf{e}_t^{S\prime} (\boldsymbol{\Sigma}_t^S)^{-1} \mathbf{e}_t^S\right)\right\} \\ \mathbf{R}_{i,T} &= [R_{i,1}, R_{i,2}, \ldots, R_{i,T}] \end{aligned} \tag{13}$$

where T is the sample size and $\phi_{i,t-1}$ is the integration measure previously specified. The parameter vector is given by

$$\Theta = \left[\boldsymbol{\delta}', \boldsymbol{\delta}_i', \mathrm{vech}(\mathbf{C}^I)', \mathbf{C}^S(1,1), \mathbf{C}^S(1,2), \mathbf{A}^I(1,1), \mathbf{A}^I(2,1), \mathbf{A}^I(2,2), \boldsymbol{\beta}'\right]',$$

where $\boldsymbol{\beta}$ summarizes the parameters needed to estimate $\phi_{i,t-1}$. Under very weak assumptions, including misspecification of the error distribution function (see White (1982)), the vector of parameters, Θ, is asymptotically nor-

[3] The assumption that the world shocks do not affect the local variance in the segmented market is far stronger than the restriction that local shocks do not affect the world variance process. The plausibility of this restriction is explored in Bekaert and Harvey (1995a).

mally distributed with covariance matrix $\mathbf{A}^{-1}\mathbf{B}\mathbf{A}^{-1}$, where $\mathbf{A}$ is the Hessian form and $\mathbf{B}$ the outer product form of the information matrix. Below, we report "robust" standard errors.[4]

Rather than estimating the likelihood function in equation (13) directly, we proceed in two steps. First, we estimate $C^I(2,2)$, $A^I(2,2)$, and $\boldsymbol{\delta}$ using the world market return and the world information variables, $\mathbf{Z}$. Second, we estimate equation (13) country by country imposing the parameter estimates from the first stage. This procedure imposes the restriction that the price of world market risk is the same in each country, which leads to more powerful tests. A disadvantage to this approach is that the usual standard errors are likely to be understated since we ignore the sampling error in the first-stage parameter estimates.

C.2. Specification Tests and Diagnostics

Many of the markets in our sample show predictable variation in returns. In contrast to previous work, our model has *three* sources of time-variation in expected returns: variation in the prices of risk (λ_{t-1}, $\lambda_{i,t-1}$), variation in the conditional risk measures (covariance with world and local market variances) and variation in the degree of market integration ($\phi_{i,t-1}$). Our estimation technique allows us to recover the time path of all three components. To gauge the ability of the model to capture the observed predictability of returns, we test whether the time t disturbance, $\mathbf{e}_t$, is orthogonal to information $\mathscr{Z}_{t-1}$ available at time $t-1$. The first set of diagnostics reports the R^2 and a heteroskedasticity-consistent Wald test of the joint significance of the coefficients of a linear regression of $e_{i,t}$ onto a set of information variables $\mathscr{Z}_{t-1}$. If the model fails to replicate the observed time-variation of expected returns, it is useful to track the source of the rejection. Hence, we set $\mathscr{Z} = \mathbf{Z}$, $\mathscr{Z} = \mathbf{Z}^i$, and $\mathscr{Z} = [\mathbf{Z}, \mathbf{Z}^i]$.

In addition to these informal diagnostics, we also perform a number of formal Lagrange Multiplier (LM) tests.[5] The alternative model that we consider is:

$$r_{i,t} = \boldsymbol{\zeta}'\mathscr{Z}_{t-1} + \phi_{i,t-1}\lambda_{t-1}\mathrm{cov}_{t-1}[r_{i,t}, r_{w,t}] + (1 - \phi_{i,t-1})\lambda_{i,t-1}\mathrm{var}_{t-1}[r_{i,t}], \tag{14}$$

and we test whether $\boldsymbol{\zeta} = \mathbf{0}$. The choices for $\mathscr{Z}$ are the same as above. We report the standard LM test computed as the uncentered R^2 from a regression of the unit vector on the matrix of scores under the null.

[4] The estimator is the quasi-maximum likelihood estimator (QMLE). For GARCH models, Bollerslev and Wooldridge (1992) show that the QMLE is generally consistent and has a limiting normal distribution as long as the first two conditional moments are correctly specified. Gray (1995a) has extended these results to standard regime switching models. Note that for ARCH-in-mean models the asymptotic properties of the maximum likelihood estimators have not been worked out.

[5] Computational difficulty in estimating even larger models prevents us from considering Wald or likelihood ratio tests.

We then estimate three alternative models embedded in the general specification equations (5) to (12). In the first alternative, we assume constant prices of risk and provide a likelihood ratio test of this restriction. The second alternative constrains the conditional variances to be constant over time (no ARCH). This produces a second likelihood ratio. Finally, in the third alternative, the degree of integration is constrained to be constant over time. In the standard Hamilton model, this alternative is nested by setting $1 - Q = P$.[6] It corresponds to a standard mixture of normals model (see Everitt and Hand (1981)). This delivers the final likelihood ratio.

Finally, we also report a likelihood ratio test of the standard Hamilton model versus a model with time-varying transition probabilities. When the Hamilton model is rejected, the constant prices of risk and no ARCH models are estimated using time-varying transition probabilities.

II. Data and Summary Statistics

A. The Data

Our sample of national equity markets includes data for both developed markets from Morgan Stanley Capital International (MSCI) and emerging markets from the International Finance Corporation (IFC) of the World Bank. The IFC provides value-weighted indices of a representative sample of equities in each country covering at least 60 percent of the market's capitalization. Our study focusses on twelve emerging markets: Chile, Colombia, Greece, India, Jordan, Korea, Malaysia, Mexico, Nigeria, Taiwan, Thailand, and Zimbabwe. These markets account for over 80 percent of the capitalization of all of the markets followed by the IFC.[7]

The summary statistics are presented in Table I for the total available data for each country. Most of the MSCI data begin in December 1969 and earliest available data for seven of the 12 emerging countries is December 1975. Our analysis will concentrate on the U.S. dollar returns.[8] The statistics include the average (annualized) arithmetic and geometric return, standard deviation, and autocorrelations. The developed market summary statistics are presented over different samples by other authors and appear for the purpose of comparison with the emerging market returns.

The range of average returns is much greater for the emerging than the developed markets. The mean U.S. dollar returns for the emerging markets vary from 43 percent (Colombia) to 3 percent (Nigeria). This sharply contrasts with the range of average returns in the developed markets. In the MSCI sample, no country has an average arithmetic return that exceeds 30 percent.

[6] Although we did not do this, in the model with time-varying transition probabilities, this restriction can be imposed by setting $\beta_1 = -\beta_2$.

[7] The appendix provides the details of the index construction and compares the IFC methodology to the MSCI methodology.

[8] Calculating the returns in U.S. dollars eliminates the location inflation. However, the U.S. inflation remains in the returns.

Table I

Summary Analysis of International Equity Returns

Means, standard deviations, and autocorrelations coefficients of 21 developed market returns based on the Morgan Stanley Capital International (MSCI) indices and 12 emerging market returns based on the International Finance Corporation (IFC) indices. Both arithmetic averages and geometric averages are reported. Both means and standard deviations are in annualized percentage terms. All returns are calculated in U.S. dollar terms. The sample ends in December 1992.

Country	Start	Arithmetic Mean	Geometric Mean	Std. Dev.	Autocorrelation ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	ρ_{24}
Panel A: Morgan Stanley Capital International										
Australia	70.01	11.40	7.56	26.95	0.00	−0.05	−0.00	−0.00	−0.04	0.04
Austria	70.01	14.20	11.81	21.97	0.15	0.03	0.05	0.10	0.03	0.02
Belgium	70.01	16.08	14.01	20.19	0.09	0.04	0.03	0.04	0.04	0.04
Canada	70.01	10.66	8.75	19.35	0.00	−0.09	0.07	−0.03	−0.04	0.05
Denmark	70.01	14.37	12.44	19.52	0.04	0.14	0.09	0.10	−0.11	0.07
Finland	88.01	−10.25	−13.18	24.05	0.14	−0.36	−0.17	−0.04	0.01	0.42
France	70.01	14.62	11.55	24.64	0.08	−0.00	0.12	0.03	−0.03	−0.00
Germany	70.01	12.78	10.47	21.36	0.01	−0.02	0.10	0.07	−0.05	0.02
Hong Kong	70.01	27.21	18.66	41.58	0.06	−0.04	−0.01	−0.05	−0.01	−0.02
Ireland	88.01	8.03	5.24	23.95	−0.14	−0.09	−0.10	0.19	−0.23	0.41
Italy	70.01	8.08	4.59	26.62	0.11	−0.01	0.09	0.08	0.03	0.04
Japan	70.01	18.00	15.32	22.97	0.06	0.00	0.10	0.04	0.06	0.01
Netherlands	70.01	15.69	13.91	18.53	0.04	−0.04	0.05	−0.10	0.06	−0.00
New Zealand	88.01	−2.20	−5.38	26.01	−0.05	−0.09	−0.11	−0.14	−0.10	0.07
Norway	70.01	14.92	10.88	28.30	0.16	−0.01	0.14	−0.05	0.02	0.04
Singapore/ Malaysia	70.01	18.84	14.08	30.95	0.17	−0.01	−0.08	0.05	0.04	0.00
Spain	70.01	9.80	7.19	22.75	0.12	−0.02	−0.03	0.07	−0.01	0.13
Sweden	70.01	15.67	13.13	22.31	0.08	−0.02	0.04	−0.01	0.03	0.05
Switzerland	70.01	13.36	11.38	19.73	0.05	−0.07	0.04	−0.00	0.01	−0.01
United Kingdom	70.01	15.42	12.04	26.52	0.09	−0.10	0.06	0.01	−0.01	0.05
United States	70.01	11.54	10.24	15.90	0.02	−0.04	0.00	−0.01	0.04	−0.01

Table I—*Continued*

Country	Start	Arithmetic Mean	Geometric Mean	Std. Dev.	Autocorrelation ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	ρ_{24}
Panel B: International Finance Corporation										
Chile	76.01	36.67	29.06	39.58	0.17	0.26	−0.01	−0.03	0.08	0.06
Colombia	85.01	43.64	38.48	32.14	0.49	0.15	−0.03	−0.16	−0.06	−0.08
Greece	76.01	7.47	1.54	36.22	0.13	0.18	0.03	−0.06	−0.04	0.06
India	76.01	20.20	16.54	27.23	0.08	−0.10	−0.03	−0.11	−0.08	−0.01
Jordan	79.01	10.75	9.15	17.89	0.00	0.02	0.18	0.00	−0.00	−0.00
Korea	76.01	21.26	16.27	32.34	−0.00	0.08	0.02	−0.02	0.09	0.05
Malaysia	85.01	13.84	10.24	26.35	0.05	0.06	−0.07	−0.01	−0.10	0.09
Mexico	76.01	30.39	19.20	44.56	0.25	−0.07	−0.04	0.04	−0.01	0.02
Nigeria	85.01	2.70	−5.49	36.50	0.09	−0.13	−0.21	0.03	−0.06	−0.02
Taiwan	85.01	34.02	20.21	52.90	0.07	0.05	−0.04	0.07	0.14	−0.08
Thailand	76.01	22.33	18.87	25.76	0.11	0.15	0.01	−0.12	0.04	−0.07
Zimbabwe	76.01	7.77	1.99	34.17	0.14	0.16	0.25	0.17	−0.01	−0.03

In the IFC sample, four countries (Chile, Colombia, Mexico, and Taiwan) have average returns above 30 percent.

Emerging market returns are characterized by high volatility. Standard deviations range from 18 percent (Jordan) to 53 percent (Taiwan). In contrast for the MSCI countries, volatility ranges between 15 percent and 42 percent. There are eight emerging markets with volatility higher than 30 percent.[9]

The emerging market returns are also more autocorrelated. In the MSCI sample of 18 countries with data from December 1969, there are only 5 countries with first-order autocorrelation that exceeds 10 percent. In contrast, six of the 12 emerging countries have autocorrelations greater than 10 percent. There are two countries with autocorrelations above 20 percent (Colombia and Mexico). This suggests that the returns in many of these countries are predictable (to some extent) based on past returns alone.

B. Predictability

A number of studies have documented the existence of predictable variation in developed country returns.[10] Recently, evidence of predictable variation in emerging market returns has been documented in Bekaert (1995) and Harvey (1993a, b, 1995).

In our econometric model, we separate the total information set, $\mathscr{Z}$, into local components, $\mathbf{Z}^i$, and global components, $\mathbf{Z}$. It is also necessary to be parsimonious with respect to the number of information variables presented. The global information variables include: a constant, the world market dividend yield in excess of the 30-day Eurodollar rate, the default spread (Moody's Baa minus Aaa bond yields), the change in the term structure spread (U.S. 10-year bond yield minus 3-month U.S. bill), and the change in the 30-day Eurodollar rate. These variables are designed to capture fluctuations in expectations of the world business cycle.[11]

The set of local information variables include a constant, local equity returns, local exchange rate changes, local dividend yields, and the ratio of equity market capitalization to GDP. These variables are designed to capture expectations about local economic conditions. Obviously, some of these variables will be correlated with the global variables—just as local economic growth may be correlated with world economic growth. However, the degree of correlation is small. For example, Ferson and Harvey (1994b) find less than 40 percent average correlation among dividend yields in the MSCI countries.

[9] Bekaert and Harvey (1995a) explore the reasons why volatility is different in emerging versus developed markets and detail the time-series characteristics of emerging market volatility.

[10] See Harvey (1991a), Bekaert and Hodrick (1992), Campbell and Hamao (1992), Ferson and Harvey (1993, 1994b), and Harvey, Solnik and Zhou (1994).

[11] While some of the variables are U.S. based, Harvey (1991b) shows that U.S. economic growth has 89 percent correlation with G-7 economic growth. He also finds that measures of the U.S. term structure have 87 percent correlation with GDP weighted measures of the world term structure.

Table II presents heteroskedasticity-consistent tests of the null hypothesis that expected returns are constant. In Panel A, tests are conducted on the developed markets. The multivariate test[12] of no predictability using the global information variables for 18 markets (Finland, Ireland, and New Zealand are excluded because their data begin in 1988) provides evidence against the null hypothesis. In addition, the table shows that the combination of global and local variables enhances the degree of predictability.

Panel B considers the 12 emerging markets. In more than half of these countries, the null hypothesis of no predictability is rejected at the 10 percent level. A multivariate test using the global information variables also provides a rejection of the null hypothesis at the 10 percent level of significance.

III. Empirical Results

A. The World Price of Covariance Risk

Table III presents the estimation of the ARCH-M model for the world price of covariance risk:

$$r_{w,t} = \lambda_{t-1}\text{var}_{t-1}[r_{w,t}] + \varepsilon_{w,t}$$

$$\lambda_{t-1} = \exp(\boldsymbol{\delta}'\mathbf{Z}_{t-1})$$

where $\text{var}_{t-1}[r_{w,t}]$, is given by:

$$h_t = c^2 + \alpha^2 \sum_{k=1}^{K} w_k \varepsilon^2_{w,t-k}$$

and $\mathbf{Z}$ represents the global information variables. Consistent with the evidence presented in Harvey (1991), the hypothesis that the world price of risk is constant is easily rejected. This is also seen in Figure 1 which plots the fitted prices of risk.

The price of risk displays a distinct business cycle pattern (NBER peaks and troughs of the U.S. business cycle are denoted by arrows). The price of risk is highest at economic troughs and lowest at economic peaks. This variation is consistent with the U.S. results of Fama and French (1989) who argue that risk premiums should be highest in recessions to lure investors into the market. Using the latent variables methodology and a sample of 18 MSCI country equity returns and eight fixed income returns, Harvey, Solnik, and Zhou (1994) also document business cycle variation in the world equity market risk premium.

The spikes in the world price of risk coincide with the deepest recessions: 1973–74, 1979, and 1981. Notably, there is no spike in March 1991 (NBER cyclical trough for the United States). However, the most recent recession was unusual in a number of respects. First, there was considerable debate about the nature of this recession. This is evidenced by the NBER waiting an unusually long time (over a year) before dating the business cycle trough.

[12] For a detailed analysis of this test and other multivariate tests of predictability, see Kirby (1994).

Table II

Predictability Using Local and Global Information

Heteroskedasticity-consistent tests of the null hypothesis that expected returns are constant. Tests are conducted on both the developed and emerging markets. Multivariate tests of no predictability using the global information variables for 18 developed markets with the longest samples and for the 9 emerging markets with the longest samples are presented. Individual country regressions are then estimated with both local and global information variables. Heteroskedasticity-consistent exclusion tests of global, local and global/local information are presented. The global information variables include a constant, the world market dividend yield in excess of the 30-day Eurodollar rate, the default spread (Moody's Baa minus Aaa bond yields), the change in the term structure spread (U.S. 10-year bond yield minus 3-month U.S. bill), and the change in the 30-day Eurodollar rate. The set of local information variables includes a constant, local equity returns, local exchange rate changes, local dividend yields, and the ratio of equity market capitalization to GDP. All returns are calculated in U.S. dollars.

	Exclude Global		Exclude Local		Exclude Global + Local	
Country	χ^2	p-value	χ^2	p-value	χ^2	p-value
Panel A: Morgan Stanley Capital International						
Australia	11.22	0.024	5.07	0.280	8.07	0.427
Austria	14.82	0.005	6.85	0.144	16.77	0.033
Belgium	7.62	0.107	4.86	0.302	8.73	0.366
Canada	6.32	0.177	6.70	0.153	18.46	0.018
Denmark	4.82	0.306	3.82	0.431	8.29	0.405
Finland	4.48	0.344	6.34	0.175	14.08	0.080
France	3.39	0.495	3.83	0.430	6.82	0.556
Germany	7.04	0.134	3.88	0.422	16.92	0.031
Hong Kong	1.35	0.853	4.52	0.341	6.24	0.620
Ireland	5.38	0.250	3.88	0.422	14.00	0.082
Italy	0.79	0.939	8.01	0.091	9.60	0.294
Japan	7.95	0.093	17.70	0.001	28.06	0.001
Netherlands	6.51	0.164	6.42	0.170	21.75	0.005
New Zealand	4.30	0.368	2.22	0.696	22.50	0.004
Norway	2.85	0.583	5.37	0.252	23.73	0.003
Singapore/Malaysia	10.39	0.034	5.41	0.248	8.63	0.375
Spain	9.88	0.043	2.62	0.624	23.40	0.003
Sweden	8.43	0.077	3.89	0.421	11.00	0.202
Switzerland	5.88	0.208	4.24	0.374	17.89	0.022
United Kingdom	3.63	0.459	3.68	0.451	4.95	0.763
United States	13.85	0.008	4.89	0.299	14.55	0.069
18 countries	—	0.086	—	—	—	—
Panel B: International Finance Corporation						
Chile	15.99	0.003	4.09	0.394	12.55	0.128
Colombia	3.01	0.556	5.43	0.246	15.10	0.057
Greece	8.29	0.082	16.28	0.003	181.63	0.000
India	5.96	0.202	3.13	0.537	6.38	0.604
Jordan	9.16	0.057	4.79	0.309	6.59	0.581
Korea	8.06	0.090	1.88	0.599	13.76	0.056
Malaysia	0.97	0.914	2.83	0.586	3.44	0.904
Mexico	15.83	0.003	3.09	0.542	18.55	0.018
Nigeria	2.16	0.707	0.96	0.916	6.96	0.541
Taiwan	21.24	0.000	3.18	0.529	16.06	0.042
Thailand	4.80	0.309	5.13	0.274	13.51	0.095
Zimbabwe	4.83	0.305	5.43	0.246	22.62	0.004
9 countries	—	0.049	—	—	—	—

Table III

The World Price of Covariance Risk

The model estimated is:

$$r_{w,t} = \lambda_{t-1}\text{var}_{t-1}[r_{w,t}] + \varepsilon_{w,t}$$
$$\lambda_{t-1} = \exp(\delta'\mathbf{Z}_{t-1})$$

where r_w is the world market return, λ is the world price of covariance risk, ε_w is the return disturbance, and the conditional variance, $h_t = \text{var}_{t-1}[r_{w,t}]$, is given by:

$$h_t = c^2 + \alpha^2 \sum_{k=1}^{K} w_k \varepsilon^2_{w,t-k}$$

The conditioning information, **Z**, is a set of global information variables which include a constant, the world market dividend yield in excess of the 30-day Eurodollar rate, the default spread (Moody's Baa minus Aaa bond yields), the change in the term structure spread (U.S. 10-year bond yield minus 3-month U.S. bill), and the change in the 30-day Eurodollar rate. In the estimation, we set $K = 3$ and let $w_k = 2(K + 1 - k)/(K(K + 1))$ as in Engle, Lilien, and Robbins (1987). The resulting weights on the three past residual vectors are 1/2, 1/3 and 1/6, respectively. The χ^2 statistic in Panel B provides a test of whether the conditional variance is constant. It has one degree of freedom. The χ^2 statistic in Panel C provides a test of whether the price of risk is constant. It has 4 degrees of freedom.

Panel A: Full Model

δ_1	δ_2	δ_3	δ_4	δ_5	c	α
0.217	−0.280	0.291	1.564	−0.123	0.039	0.345
(0.930)	(0.112)	(0.116)	(0.494)	(0.202)	(0.003)	(0.161)

Panel B: Constant Variance Model

δ_1	δ_2	δ_3	δ_4	δ_5	c	χ^2
0.154	−0.351	0.343	1.656	−0.063	0.042	2.078
(0.712)	(0.124)	(0.163)	(0.474)	(0.176)	(0.003)	[0.149]

Panel C: Constant Price of Risk

λ	c	α	χ^2
1.876	0.040	0.361	11.166
(1.535)	(0.004)	(0.152)	[0.025]

Second, the recession was less severe than the three previous ones. Third, and most importantly, the U.S. recession did not coincide with a world recession. Recessions in most European countries and Japan followed the U.S. recession. Presumably, variation in the world price of risk should reflect the world business cycle. The fact that the U.S. business cycle was out of phase with other major economies during the last U.S. recession could account for the lack of variation in the world price of risk.

There is only weak evidence that the variance dynamics follow an ARCH process. The α parameter is significant at standard levels (t-ratio of 2.1); however the χ^2 test of the null hypothesis that the variance is constant is not

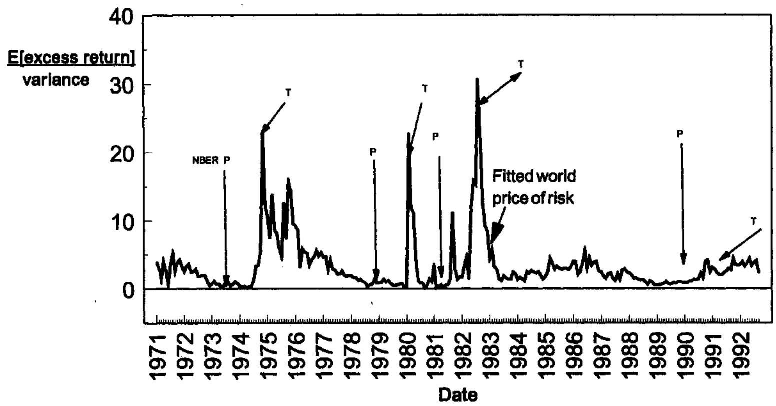


Figure 1. Time-variation in the world price of risk. The model estimated is:

$$r_{w,t} = \lambda_{t-1}\text{var}_{t-1}[r_{w,t}] + \varepsilon_{w,t}$$
$$\lambda_{t-1} = \exp(\delta'\mathbf{Z}_{t-1})$$

where r_w is the world market return, λ is the world price of covariance risk, ε_w is the return disturbance, and the conditional variance, $h_t = \text{var}_{t-1}[r_{w,t}]$, is given by:

$$h_t = c^2 + \alpha^2 \sum_{k=1}^{K} w_k \varepsilon_{w,t-k}^2$$

The conditioning information, **Z**, is a set of global information variables which include a constant, the world market dividend yield in excess of the 30-day Eurodollar rate, the default spread (Moody's Baa minus Aaa bond yields), the change in the term structure spread (U.S. 10-year bond yield minus 3-month U.S. bill), and the change in the 30-day Eurodollar rate. In the estimation, we set $K = 3$ and let $w_k = 2(K + 1 - k)/(K(K + 1))$ as in Engle, Lilien, and Robbins (1987). The resulting weights on the three past residual vectors are 1/2, 1/3 and 1/6, respectively. The fitted values of λ_{t-1} are presented. P and T denote the National Bureau of Economic Research U.S. business cycle peaks and troughs, respectively.

rejected at conventional significance levels (p-value is 0.15). We examined the fitted values of the full model and the no ARCH model (full model in Figure 1). Both series exhibit the same time-series characteristics. However, the no ARCH model price of risk has some extreme values (over 100) at the beginning of 1980. Given the significance of the α coefficient and the higher volatility (and unreasonable values) implied by the no ARCH model, we choose to use the model with ARCH and time-varying prices of risk in the subsequent analysis.

B. Estimation

The results for estimating the standard Hamilton model are presented in Table IV. In this model, the transition probabilities are constant. The first column reports the probability of being in the integrated state given that the previous state was integration (P). The second column reports the probability of being in the segmented state given that the previous state was segmented

(Q). These transition probabilities along with the lagged regime probabilities and the likelihood form the conditional measure of integration in equation (4). The table also reports a likelihood ratio test of the null hypothesis that the transition probabilities are constant.

Both the standard Hamilton model and the model with time-varying transition probabilities are highly nonlinear and, as a result, special care must be taken in the estimation. We first estimated the standard Hamilton model and confirmed the optimum with at least ten different sets of starting values. We use the parameters from the Hamilton model as a starting point for the time-varying transition probability or full model. This model has the most parameters, and up to 25 different sets of starting values were used to confirm the global optimum.

For Chile, Greece, Jordan, Korea, Thailand, and Zimbabwe, the model with constant transition probabilities is clearly rejected. For Colombia and Mexico, there is some weak evidence against the constant transition probability model. Table IV also reports the mean levels of the integration parameter over the entire sample as well as over the last three years (post 1990). We will now examine, in more detail, the results for each country.

B.1. Chile

The average value of the integration parameter for Chile is 0.59 and in recent years the value has dropped to 0.26. The trend towards segmentation is evident in Figure 2, which plots the ex ante probability of integration based on the model with time-varying transition probabilities. The integration parameter is equal to 1.0 between 1981 and 1984 and then drops sharply.

There are a priori reasons to expect some degree of segmentation in the Chilean market. Foreign equity investors must pay a 35 percent tax on both dividends and capital gains. Most importantly, there are currency controls (see World Bank (1993)). The official rate often diverges from the market rate, and most foreign investment flows are required to use the official rate. The market is illiquid and dominated by only a few stocks (the top five stocks account for over 50 percent of the market capitalization). To make things worse, for most of the sample, capital repatriation was not allowed for five years. This has recently been changed to one year.

Chile has one of the lowest percentages of equity that is investable among the emerging markets, namely 25 percent. Bekaert (1995), who provides detailed evidence of barriers to entry in emerging markets, ranks Chile 17th out of 20 in terms of investability. The institutional barriers to investment are consistent with the estimates of the degree of integration reported in Table IV.

B.2. Colombia

The results for Colombia suggest that the market is more segmented than integrated. Over the entire sample, the ex ante probability of integration never exceeds 0.17. There is little variation in the integration measure for a number of reasons. First, since a Hamilton model is estimated, the transition

Table IV

Estimation of the Model With Constant Transition Probabilities

The unobserved state variable, S_t^i, takes on the value of one when markets are integrated and a value of two when markets are segmented. Then, $\phi_{i,t-1}$ is the conditional probability of being integrated (regime 1), $\phi_{i,t-1} = \text{prob}[S_t^i = 1|\mathscr{Z}_{t-1}]$. To infer $\phi_{i,t-1}$ from the data, we explore two different regime switching models. In the Hamilton (1989, 1990) model, the switching probabilities are time invariant but the regime probability and the degree of market integration varies through time as new information changes the econometrician's inference on the relative likelihood of the two regimes. The transition probabilities are:

$$P = \text{prob}[S_t = 1|S_{t-1} = 1]$$
$$Q = \text{prob}[S_t = 2|S_{t-1} = 2]$$

where $\mathscr{Z}_{t-1}$ is the conditioning information. In our model with time-varying transition probabilities, we allow P and Q to be logistic functions of $\mathbf{Z}_{t-1}^*$ (includes lagged dividend yields and lagged equity market capitalization as a proportion of GDP).

The χ^2 statistic is from a likelihood ratio test for constant transition probabilities and has 4 degrees of freedom. The transition probabilities are from the simple Hamilton model (constant transition probabilities). The mean degrees of integration are from the model with time-varying transition probabilities, unless the simple Hamilton model is not rejected. In the latter case, the degrees of integration are based on the Hamilton model estimates. All returns are measured in U.S. dollars.

Country	Transition Probabilities P	Transition Probabilities Q	χ^2 [p-value]	Degree of Integration Full Sample	Degree of Integration Post-1990
Chile	0.9414	0.8688	12.198	0.59	0.26
	(0.0230)	(0.0892)	[0.016]		
Colombia	0.0000	0.8322	6.382	0.14	0.14
	(0.0000)	(0.0202)	[0.172]		
Greece	0.9868	0.6244	28.731	0.89	0.86
	(0.0011)	(0.0155)	[0.000]		
India	0.9962	0.9941	6.332	0.54	0.10
	(0.0054)	(0.0081)	[0.176]		
Jordan	0.9022	0.1710	14.570	0.85	0.79
	(0.0113)	(0.0099)	[0.006]		
Korea[a]	0.9573	0.0000	11.462	0.97	0.99
	(0.0028)	(0.0000)	[0.022]		
Malaysia[a]	0.8185	0.3214	4.123	0.79	0.79
	(0.0316)	(0.0286)	[0.390]		
Mexico	0.9363	0.9839	6.962	0.21	0.04
	(0.0427)	(0.0113)	[0.155]		
Nigeria	0.7402	0.8941	1.115	0.27	0.20
	(0.1486)	(0.0847)	[0.892]		

Table IV—*Continued*

Country	Transition Probabilities P	Q	χ^2 [p-value]	Degree of Integration Full Sample	Degree of Integration Post-1990
Taiwan	0.9309	0.3086	[b]	0.89	0.90
	(0.0090)	(0.0312)	—		
Thailand[c]	0.9062	0.9804	14.369	0.77	1.00
	(0.0078)	(0.0018)	[0.006]		
Zimbabwe	0.9808	0.9699	10.387	0.47	0.52
	(0.0098)	(0.0146)	[0.034]		

[a] The estimation for Korea was extremely ill-behaved and the full model for Malaysia failed to satisfy all convergence criteria. The results for the full model have not been confirmed as the global optimum.

[b] We failed to find an optimum with a likelihood value higher than that of the Hamilton model.

[c] Due to an ill-conditioning problem, standard errors were computed from the inverse of the Hessian.

probabilities are not time-varying. Second, the estimated P parameter is very close to zero. As a result, ϕ cannot exceed 0.17.[13] Given that P is zero, the system will never spend more than one period in the integrated state. According to the estimates, Colombia looks very segmented. However, the pricing of world factors does have some influence on the equity returns.

The evidence of segmentation in Colombia is consistent with the investment environment. The Colombian market is one of the most illiquid among the emerging markets. It ranks third last (just ahead of Chile and Nigeria) in terms of value traded divided by market capitalization. In addition, four securities account for 50 percent of trading volume. The potential liquidity problems are also evident from the remarkable 49 percent serial correlation in returns reported in Table I.

While there are some recent positive developments in Colombia such as announcements of privatization programs, there is no evidence yet of increased integration. Colombia is a good example of why integration cannot be accurately measured by regulatory standing. The degree of investability is quite high in Colombia. However, the lack of liquidity, combined with the political risk induced by the ongoing war with the drug cartels, has kept this market largely segmented.

B.3. Greece

Greece is no longer an emerging market, with U.S. $5,500 GDP per capital in 1990 (the World Bank definition of emerging market is less than $2,200 per capita in 1990). However, when the IFC indices were formed in 1981,

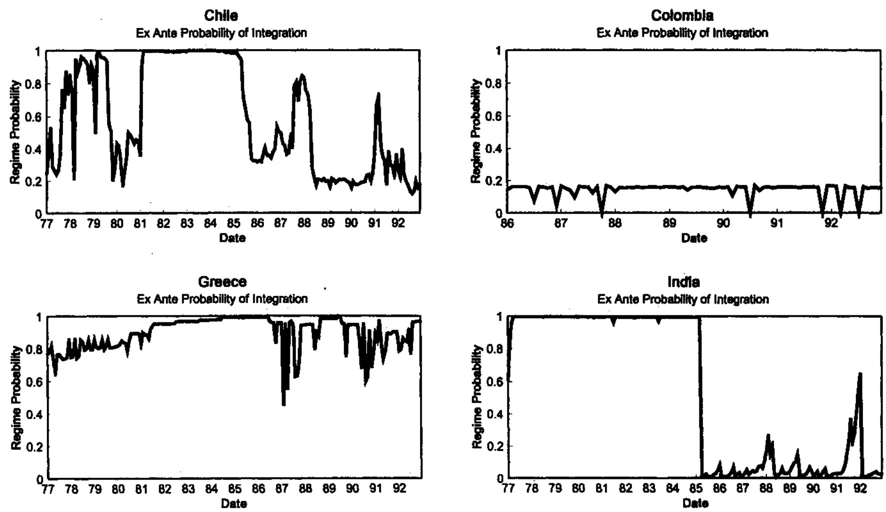
[13] This is the case because $\phi_{t-1} = (1 - Q)(1 - W_{t-1})$ where W_{t-1} is implicitly defined in equation (5). Since $Q = 0.83$, and $0 < W_{t-1} < 1$, the maximum value ϕ can take is 0.17.

Greece fell within the emerging markets category. The evidence presented in Table IV suggests that the Greek market is integrated into world capital markets. The integration parameter in the 1990s is 0.86.

The integration of Greece is consistent with the investment environment. Outside certain industries, such as banking, shipping, and insurance, there are no foreign investment restrictions. The market capitalization is $U.S. 9.5 billion at the end of 1992. There is a large foreign participation in the stock market (about 20 percent of shares are owned by foreigners). Finally, there is reasonable liquidity with $9 million in average daily trading in 1992.

B.4. India

It was difficult to develop a prior assessment of the degree of integration of the Indian market. Factors favoring integration include the long history of

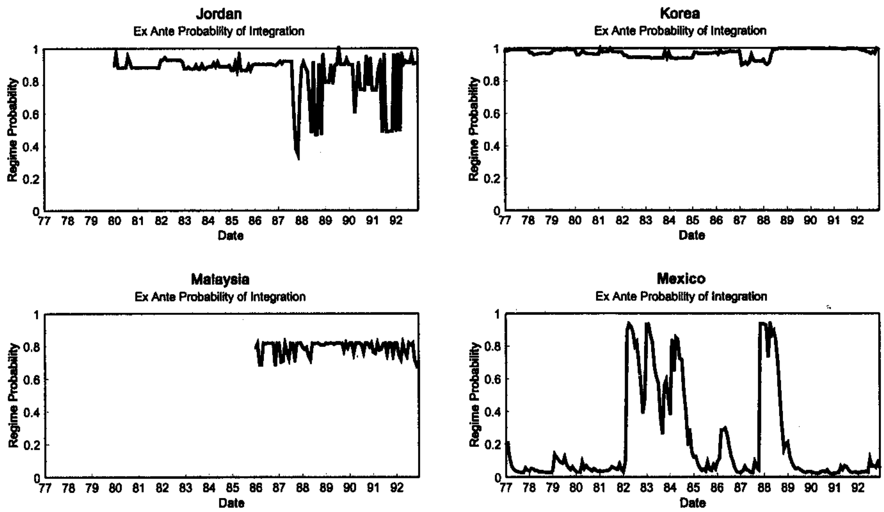


(a)

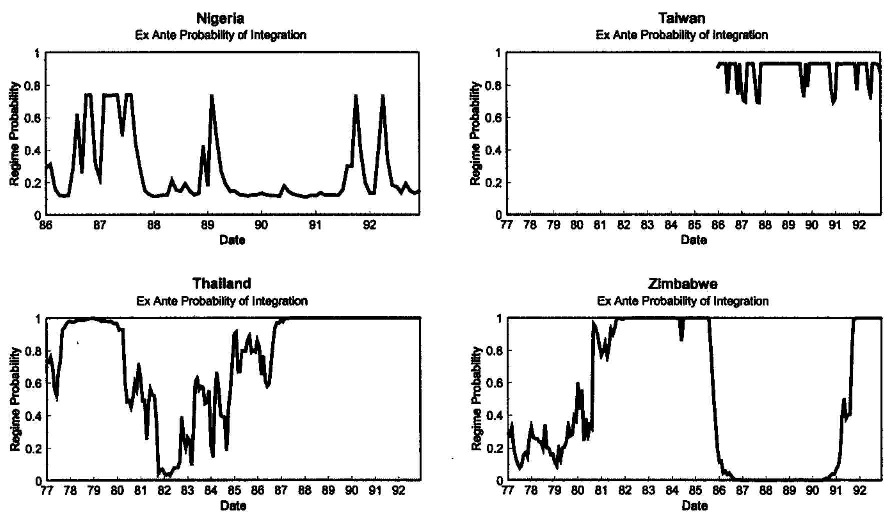
Figure 2. Time-varying integration measures. We use a regime-switching framework to estimate the ex ante probability of integration. The unobserved state variable, S_t^i, takes on the value of one when markets are integrated and a value of two when markets are segmented. Then, $\phi_{i,t-1}$ is the conditional probability of being integrated (regime 1), $\phi_{i,t-1} = \text{prob}[S_t^i = 1|\mathscr{Z}_{t-1}]$. To infer $\phi_{i,t-1}$ from the data, we use two different regime switching models. In the Hamilton (1989, 1990) model, the switching probabilities are time invariant but the regime probability and the degree of market integration vary through time as new information changes the econometrician's inference on the relative likelihood of the two regimes. The transition probabilities are:

$$P = \text{prob}[S_t = 1|S_{t-1} = 1]$$
$$Q = \text{prob}[S_t = 2|S_{t-1} = 2]$$

where $\mathscr{Z}_{t-1}$ is the conditioning information. In our model with time-varying transition probabilities, we allow P and Q to be logistic functions of $\mathbf{Z}_{t-1}^*$ (includes lagged dividend yields and lagged equity market capitalization as a proportion of GDP). The ex ante probabilities of integration for Chile, Greece, Jordan, Korea, Thailand, and Zimbabwe are based on the model with time-varying transition probabilities.



(b)



(c)

Figure 2.–*Continued.*

equity trading (the Bombay exchange is 115 years old) and the large number of stocks that trade (2556 securities were listed in 1991 on 19 exchanges). The capitalization at the end of 1992 was U.S. $65.1 billion with reasonable trading volume (U.S. $13.2 billion).

On the other hand, India is a very poor country, with only U.S. $300 of per capita GDP. Stock market investment is limited to authorized investors only. That is, foreigners need permission of the Reserve Bank of India to purchase shares. However, once approved there is complete freedom to repatriate.

Other factors such as political and religious strife and the the tensions with Pakistan could also work against foreign investors participating in the Indian market.

The results in Table IV suggest that India is not fully integrated into world capital markets. The average degree of integration has decreased. In Figure 1, the time-series patterns in the degree of integration are striking. The model suggests that India was fully integrated into world capital markets until the end of 1984. The integration parameter then plummets to close to zero. Interestingly, this closely coincides with the assassination of Prime Minister Indira Gandhi on October 31, 1984. There is some recent evidence of a movement towards higher levels of integration.

B.5. Jordan

The estimates suggest that Jordan is not fully integrated into world capital markets. In Table IV, the recent degree of integration is 79 percent. From Figure 2, the degree of integration has fluctuated between 40 percent and 90 percent over the past five years. The Jordanian market is small, with a market capitalization of U.S. \$3.2 billion at the end of 1992. Foreigners are restricted to owning up to 49 percent of equity (with the exception of tourism and agriculture where there are no limits). Importantly, 85 percent of equities is owned by Jordanians. The remaining 15 percent is thought to be owned mainly by investors from other Arab states. There are no American Depository Receipts (ADRs) and no country funds. The only way to access the Jordanian market is to trade there directly. These factors are consistent with our evidence.

B.6. Korea

Korea also fails to qualify as an emerging market with per capita GDP exceeding U.S. \$5,000 in 1990. The evidence suggests that this market is integrated. The ex ante probability of integration lies between 0.85 and 1.00 through the entire sample. Over the past three years, the integration parameter is 0.99.

The Korean market definitely clears the liquidity hurdle. It is the third most active emerging market (behind Taiwan and Thailand) with over 100 percent of its market capitalization turning over each year. In terms of total capitalization, Korea is also the third largest emerging market (behind Mexico and Taiwan) and the 15th largest equity market in the world.

However, if integration is measured by looking at the investment regulations, one would probably conclude that the market was segmented for most of our sample. Regulations on foreign participation prohibited direct access to the Korean market until January 1992. Even the recent liberalization does not seem that impressive. Foreign ownership is limited to 10 percent in so-called unlimited industries and 8 percent in limited industries (which includes communications and defense). Recently, the 10 percent ceiling was raised to 25 percent for 45 firms that hit the 10 percent cap.

But there are other ways for foreigners to access the Korean market. At last count, Korea has 17 U.S. dollar denominated country funds and 17 non-U.S. dollar country funds. Many of these country funds have a long history (Korea Trust began in 1981) and have allowed foreigners to participate, to some degree, in the Korean market.

Cumby and Khantavit (1992) also study a regime-switching model for the Korean stock market jointly with the world market. They allow a different mean and variance in each regime, but there is no time-variation allowed in either. Hence, it is difficult to compare their results to ours. Unlike our results, they find clearly distinguishable regimes in the Korean equity market, but find it difficult to attribute the regime switches to policies concerning capital market integration. However, consistent with our results, their graphs of the regime probabilities suggest that the regime associated with capital market integration dominates during the sample. Bae (1993) also studies the Korean equity market. He finds that both domestic and international factors are important in pricing Korean equities and that the Korean market has become more integrated with world capital markets in recent years.

B.7. Malaysia

For Malaysia, our priors tilted towards integration. The equity market is large (U.S. $94 billion at the end of 1992) with good trading volume (U.S. $21.8 billion in 1992). Malaysia has experienced very mild inflation averaging only 4.6 percent over the past 25 years. In addition, the currency is a free float and foreigners can have Ringgit accounts.

Importantly, foreigners play a large role in the Malaysian market. At the end of 1992, foreign participation in Malaysian equities was 27 percent. Although foreign investment is limited by the Foreign Investment Committee to 30 percent of equity, it is not clear that this constraint is binding in our sample. In addition, foreigners can access 11 closed-end funds, seven open-end funds, and 13 ADRs.

All of these factors suggest that the market is integrated. This is confirmed in the data. Although the estimation for Malaysia was difficult, the results in Table IV suggest that the market is integrated. The 1990s integration parameter is 0.79 and has been fairly stable from the start of our data.

B.8. Mexico

The results for Mexico seem surprising. The model estimates suggest that Mexico's equity market was segmented during most of our sample with the exception of the 1982–1985 period and the late 1980s. There is also a slight upturn since 1991 (see Figure 2). Today, Mexico has one of the highest capitalized markets (U.S. $139 billion at the end of 1992), with U.S. $171 million in average daily trading volume. There are 36 ADRs and six U.S. dollar based country funds. All of these factors point toward market integration.

While Mexican stocks get much attention in the United States, most observers do not realize that before 1991 only two Mexican ADRs were trading. In fact, the Mexican market was effectively closed to foreign investors until 1981. The upturn in our integration measure in Figure 2 occurs shortly after the initial public offering of the Mexican Fund on the New York Stock Exchange. While the debt crisis, starting in 1982, may have deterred foreign investment at first, debt conversion programs subsidized and encouraged foreign equity investments. From 1986 onwards, Mexico ran a large debt-equity swap problem while dismantling several important obstacles to foreign investment.[14] Nevertheless, the Mexican country fund remained the main vehicle of access to Mexican stocks for most of the eighties.

The major reforms are fairly recent. The Mexican stock market was made 100 percent investable (with the exception of certain key sectors such as banking) in May 1989, and the dual exchange rate was abolished in November 1991. In addition, there has been a lot of economic turmoil. Mexico had the fourth highest inflation rate over the past six years (behind Brazil, Argentina, and Turkey). Given that most of the liberalizations occurred at the end of our sample, the results appear more reasonable.

B.9. Nigeria

We chose to examine Nigeria because we had a strong prior that this is the most likely market to be segmented. Per capita GDP is only U.S. $295 in 1990 and over 80 percent of the economy is linked to petroleum. The results in Table IV confirm that this market is more segmented than integrated. Over the past three years, the ex ante probability of integration is only 0.20.

The evidence of segmentation is consistent with the investment environment. The IFC categorizes the market as 0 percent investable and ranks Nigeria last among the emerging markets. Liquidity is extremely thin. Only 1 percent of market capitalization traded in 1992 (the average daily trading volume was only U.S. $55,000). All direct investment must be preapproved by the government. There are no Nigerian country funds and no ADRs. While there was some reason for optimism about reform after Nigeria's first democratic elections in late 1992, the military changed their mind and decided not to recognize the results of the election.

B.10. Taiwan

Taiwan is another country where it was difficult to form a prior opinion about the degree of integration. Factors favoring integration included the high market capitalization (U.S. $101 billion at the end of 1992) and the very large trading volume (U.S. $214 billion in 1992). In addition, Taiwan no longer qualifies as an emerging market with 1992 GDP per capita of U.S. $8,815. The NT dollar is technically floating, but the Central Bank of China keeps close control, i.e, it is not freely traded. Foreign investors are allowed to repatriate once per quarter.

[14] We thank Jorge Calderon-Rossell for useful information on the development of Mexican capital markets.

Factors that work against integration are the regulations controlling the amount of foreign equity ownership. Foreign ownership was first allowed in 1983 (our sample begins in 1985) but restricted to four approved investment funds. In January 1991, direct investment by institutional investors was allowed. Foreign individuals cannot invest directly. In addition, some industries are not investable, others have investment limits. Furthermore, no single investor can own more than 5 percent of a firm's equity.

The model suggests that Taiwan is integrated. The average degree of integration in the 1990s is 0.89. Though foreign direct participation is limited, there are eight closed-end funds, nine open-end funds, and four investment trusts. These alternative ways to access the market along with the direct institutional participation could explain the estimated degree of integration.

Cumby and Khantavit (1992) also study the degree of integration in Taiwan and detail a stronger covariance between local returns and world returns in the integrated state than in the nonintegrated state. Similar to our experience, the short period of data availability (data begin in 1985) makes both estimation and inference difficult.

B.11. Thailand

The model estimates for Thailand show a dramatic increase in the ex ante probability of integration beginning in 1986. Using a different methodology, Cumby and Khantavit (1992) also show a large shift in the degree of integration in 1986 (from 0.1 to 1.0). This change coincides with the beginning of trading on the Alien board. Most Thai stocks have foreign ownership limits. When these limits are met, identical shares (in terms of dividends and voting rights) are traded on two exchanges, the Main board—for resident Thais, and the Alien board—for nonresidents (see Bailey and Jagtiani (1994)).

The existence of the Alien and Main boards implies some direct access barriers for foreigners. In addition, foreigners are not allowed to own property in Thailand. As a result of the property restrictions, a corporation cannot have greater than 49 percent foreign ownership. Although there are ownership restrictions, the foreigners have a long history of participation in the Thai market.

In addition, there are many ways to access the Thai market. As of December 1992, there were 26 closed-end and 11 open-end Thai funds trading world wide. Direct investment, even with the ownership restrictions, is also relatively easy. Foreigner holdings are estimated to represent up to 60 percent of the freely floating shares.[15] The market is large (U.S. $58.3 billion in December 1992, 5th largest of the emerging markets) and very liquid (U.S. $72.1 billion in 1992) with the second highest turnover ratio among the emerging markets. All of these factors increase the probability that the market is integrated.

[15] See Asiamoney (1994). The free float excludes the large blocks of shares owned by family groups and banks.

B.12. Zimbabwe

We chose to examine Zimbabwe, as we did Nigeria, because of a strong prior that the country is not integrated into world capital markets. Zimbabwe is the third poorest country in our sample with per capita GDP of U.S. $621 per year in 1991 (Nigeria is last with U.S. $295). The market capitalization is the smallest in the sample at U.S. $600 million and there is only U.S. $85,000 in average daily trading volume. While foreign investors are allowed in all but certain key sectors, the market is classified as uninvestable because of strict foreign exchange controls.

The evidence in Table IV confirms our prior that the market is not fully integrated. The average integration is 57 percent in the 1990s. However, much more information can be obtained from Figure 2. There is a sharp increase in the integration parameter in the late 1970s which coincides with the optimism leading to independence that was officially achieved on April 18, 1980.

In the mid 1980s, the integration parameter falls to zero. This coincides with the time that the strict exchange controls are implemented. Recently, there has been a sharp increase in the integration parameter that remains unexplained.

IV. Diagnostics

A. Robustness of the Specification

Table V presents three sets of model diagnostics. First, we regress the model errors (returns minus the model fitted values) for each country on the three sets of information variables. This produces an adjusted R^2 and a heteroskedasticity-consistent χ^2 test. The χ^2-statistic tests the hypothesis that the regression coefficients on the instruments are equal to zero. Finally, we present a Lagrange multiplier test of the alternative specified in equation (14). The test essentially adds a time-varying intercept to equation (4). The coefficient on the constant is the analogue to the Jensen (1969) "alpha." However, our alternative also tests for predictability of the pricing errors.

These tests are important for the interpretation of our results. There are many reasons why the model diagnostics might present evidence against the specification. Foremost on this list of reasons is that we choose to examine a single factor specification. Missing risk premiums could mask themselves in time-varying integration. Given a rejection of the specification, we need to exercise caution in interpreting the estimated degree of integration.

The specification tests suggest that the model specification is rejected for Chile, Greece, Korea, Mexico, and Zimbabwe. There is mixed evidence for India, Malaysia, Nigeria, Taiwan, and Thailand. We fail to reject the model for Colombia and Jordan.

First, consider the countries where the model is rejected. Chile's model errors are strongly correlated with local information. The R^2 is close to 10 percent when the errors are regressed on predetermined local information variables. While the R^2 is small on the world information, both the Wald and Lagrange multiplier tests present evidence against the specification with the common world information. A similar pattern is found for Greece. The errors are highly correlated with local information. However, the model is rejected by the Wald test with the common world information variables.

The rejections for Korea and Zimbabwe follow similar patterns. Parallel to Chile and Greece, model errors are more correlated with local information variables. But the correlations are much smaller with R^2s averaging only three percent. Consistent with the other countries, both the Wald and Lagrange multiplier test provide convincing evidence against the specification.

In contrast to the previous four countries, the Mexican rejection appears to be equally driven by both local and world information. The residual R^2s are about the same (6 percent) as are the p-values for the more formal statistical tests.

There is mixed evidence against the model for India and Taiwan. In all cases, the model R^2s are zero when measured against local information, world information or the combined world and local information sets. For both countries, the Wald test fails to reject the null hypothesis. However, when the time-varying intercept is injected into the estimation, the Lagrange multiplier test detects a misspecification.

The evidence for Thailand depends on the information set used. With the local information set, the Wald test fails to reject the null hypothesis and the Lagrange Multiplier test delivers a p-value of 3.6 percent. More convincing evidence against the model is furnished with the world information set.

We classify Malaysia as mixed because of the estimation problems that we encountered. Although the Wald tests do not reject the null hypothesis for any of the information specifications, we could not confirm that we achieved the global optimum.

Neither the Wald test nor Lagrange multiplier test provide any evidence against the null hypothesis for Nigeria. However, we classify the evidence as mixed in this country because of the large model residual R^2 with the local information.

Our diagnostic tests do not uncover evidence of misspecification for Colombia and Jordan. The residual R^2s are low in every case. Furthermore, both the Wald and Lagrange multiplier tests fail to reject the model specification.

These diagnostics suggest evidence against the model specification for a number of our sample countries. The strength of rejection and the source of the rejection generally differs across countries. A rejection does not imply that the model yields no useful information. Nevertheless, extreme caution should be exercised when interpreting the integration measure, $\phi_{i,t-1}$, in those countries where there is evidence against the model.

B. Integration and Foreign Exchange Regimes

It is possible that the estimated degree of integration is capturing changes in foreign exchange regimes rather than the broader notion of capital market

Table V

Model Diagnostics: Correlation of the Country Asset Pricing Errors with Information

The R^2 statistics are adjusted for degrees of freedom and result from a regression of the error term $e_{i,t}$ on the set of instruments: $\mathbf{Z}$ global information, $\mathbf{Z}^i$ local information, or $\mathscr{Z}$ global and local information. The W-statistics are heteroskedasticity-consistent Wald tests on the joint significance of the coefficients in that regression. The p-values are based on a χ^2 distribution with degrees of freedom equal to the number of included regressors. The LM tests are standard Lagrange multiplier tests of the alternative specified in equation 14. They are asymptotically distributed χ^2 with degrees of freedom equal to the number of elements in $\mathbf{Z}$, $\mathbf{Z}^i$, or $\mathscr{Z}$ if the error distribution is correctly specified. The asterisk indicates that the model is estimated with time-varying rather than constant transition probabilities. All returns are measured in U.S. dollars.

	World Information $\mathbf{Z}$			Local Information $\mathbf{Z}^i$			World Plus Local $\mathscr{Z} = [\mathbf{Z}, \mathbf{Z}^i]$		
Country	$\bar{R}^2$	W	LM	$\bar{R}^2$	W	LM	$\bar{R}^2$	W	LM
Chile*	0.0097	38.64 [0.000]	21.32 [0.000]	0.0945	61.39 [0.000]	25.41 [0.000]	0.1004	83.51 [0.000]	28.33 [0.001]
Colombia	−0.0262	4.94 [0.423]	3.02 [0.697]	−0.0362	3.31 [0.652]	6.49 [0.261]	0.0234	8.48 [0.487]	10.53 [0.309]
Greece*	−0.0069	29.83 [0.000]	a —	0.2726	42.06 [0.000]	a —	0.2681	57.64 [0.000]	a —
India	−0.0012	6.55 [0.257]	25.80 [0.000]	−0.0163	1.27 [0.938]	31.08 [0.000]	−0.0111	11.16 [0.265]	33.74 [0.000]
Jordan*	−0.0171	3.33 [0.648]	8.34 [0.138]	0.0444	3.74 [0.587]	9.99 [0.075]	0.0280	4.08 [0.906]	11.38 [0.251]
Korea*	0.0074	17.49 [0.004]	a —	0.0243	30.71 [0.000]	a —	0.0443	42.03 [0.000]	a —
Malaysia	−0.0452	1.84 [0.871]	a —	−0.0060	6.45 [0.265]	a —	−0.0523	7.73 [0.562]	a —
Mexico	0.0580	21.82 [0.001]	29.64 [0.000]	0.0657	13.38 [0.020]	33.72 [0.000]	0.1266	42.80 [0.000]	44.53 [0.000]

Table V—*Continued*

Country	World Information $\mathbf{Z}$			Local Information $\mathbf{Z}^i$			World Plus Local $\mathscr{Z} = [\mathbf{Z}, \mathbf{Z}^i]$		
	$\bar{R}^2$	W	LM	$\bar{R}^2$	W	LM	$\bar{R}^2$	W	LM
Nigeria	−0.0355	4.05 [0.543]	5.24 [0.388]	0.3830	7.11 [0.213]	3.28 [0.657]	0.3900	9.48 [0.395]	9.85 [0.363]
Taiwan	−0.0232	5.99 [0.307]	26.59 [0.000]	0.0290	4.10 [0.535]	20.81 [0.001]	0.0006	14.03 [0.121]	30.27 [0.000]
Thailand*	0.0394	19.55 [0.002]	13.23 [0.021]	0.0627	6.33 [0.275]	11.94 [0.036]	0.0793	25.86 [0.002]	20.58 [0.015]
Zimbabwe*	0.0106	9.44 [0.093]	19.44 [0.002]	0.0281	17.75 [0.003]	19.55 [0.002]	0.0720	28.86 [0.001]	25.16 [0.003]

[a] The moment matrix of the scores was singular.

Table VI

The Interaction between Integration and Foreign Exchange Regimes

Two regressions are estimated for each country:

$$\Delta s_{t+1} = \alpha_0 + \alpha_1 \Delta s_t + \alpha_2 i_t + \alpha_3 \hat{\phi}_t + e_{t+1}$$
$$\Delta s_{t+1} = \alpha_0^* + \alpha_1^* \Delta s_t + \alpha_2^* i_t + e_{t+1}^*$$

where s_t is exchange rate versus the U.S. dollar, i_t is the interest rate, and $\hat{\phi}_t$ is the estimated degree of integration. We report the difference between the adjusted R^2s of the two models as well as the χ^2 and p-value associated with α_3 (coefficient on the estimated integration). The χ^2 test has one degree of freedom. All returns are measured in U.S. dollars.

Country	$\Delta \bar{R}^2$	χ^2
Chile	0.0197	4.081 [0.043]
Colombia	0.0093	4.344 [0.037]
Greece	0.0015	1.779 [0.182]
India	0.0072	2.964 [0.085]
Jordan	−0.0066	0.0013 [0.971]
Korea	0.0125	4.536 [0.033]
Malasia	0.0618	4.576 [0.032]
Mexico	0.0012	0.8513 [0.356]
Nigeria	0.0328	2.683 [0.101]
Taiwan	−0.0116	0.0436 [0.835]
Thailand	0.0005	2.512 [0.113]
Zimbabwe	0.0076	2.798 [0.094]

integration.[16] Table VI presents tests of the following regression models:

$$\begin{aligned} \Delta s_{t+1} &= \alpha_0 + \alpha_1 \Delta s_t + \alpha_2 i_t + \alpha_3 \hat{\phi}_t + e_{t+1} \\ \Delta s_{t+1} &= \alpha_0^* + \alpha_1^* \Delta s_t + \alpha_2^* i_t + e_{t+1}^* \end{aligned} \tag{15}$$

[16] We are grateful to Burton Hollifield for suggesting this possibility.

where s_t is U.S. dollar per local currency exchange rate, i_t is the interest rate, and $\hat{\phi}_t$ is the estimated degree of integration. We report the difference between the adjusted R^2s of the two models as well as the χ^2 and p-value associated with α_3 (coefficient on the estimated integration) which has one degree of freedom. To mitigate the generated regressor problem in equation (15) (see Pagan (1984)), we report heteroskedasticity-consistent standard errors.

The results in Table VI do not show strong evidence that exchange rate changes and the integration measure are interrelated. In only four of 12 countries, do the tests reject the hypothesis that $\alpha_3 = 0$ at the 5 percent level of significance (Chile, Colombia, Korea, and Malaysia). Moreover, the increases in the adjusted R^2 are small, except for Malaysia and Nigeria.

C. Estimation of the Constrained Alternatives

Table VII presents likelihood ratio tests of three specific alternative hypotheses: constant prices of risk, constant variance matrices, and constant degree of integration.

The hypothesis that the price of local volatility is constant is rejected at the 5 percent level in Chile, Colombia, Greece, India, Jordan, Korea, Mexico, and Taiwan. There is no evidence against the hypothesis for Nigeria, Thailand, or Zimbabwe. We also fail to reject the constant local price of risk for Malaysia. However, as mentioned earlier, the estimation for this country was ill-behaved and we should be cautious in drawing conclusions.

The hypothesis that the variance matrices are constant is also tested with a likelihood ratio in Table VII. Constant variance matrices are rejected for eight of the ten countries for which this test was feasible. The hypothesis is rejected at the 10 percent level for the remaining two countries.

The third likelihood ratio provides a test of the hypothesis that the degree of integration is constant. This hypothesis is rejected for Chile, Greece, India, Mexico, Nigeria, Taiwan, Thailand, and Zimbabwe. The rejection is informally confirmed by noticing the time variation in fitted integration measures in Figure 2. Constant integration is not rejected for Colombia, Jordan, Korea, and Malaysia. This can be confirmed by viewing the fitted integration measures.[17]

V. Conclusions and Further Research

Most would agree that the degree to which many countries are integrated into world capital markets has changed over time. However, all previous research has made one of three assumptions: all markets are perfectly integrated, individual markets are perfectly segmented, or local markets are

[17] In an other diagnostic exercise, we estimated our model for a developed market, Germany, which is most likely completely integrated into world capital markets. Our results (available on request) indicate that Germany is integrated throughout the sample.

Table VII

Estimation Results for the Constrained Alternatives

LR1 is the likelihood ratio statistic testing the restriction that the price of risk is constant and has 4 degrees of freedom. LR2 provides a test of constant variances and has 2 degrees of freedom. LR3 presents a test of whether the degree of integration is constant and has 1 degree of freedom.

Country	Constant Price of Risk LR1	Constant Variance LR2	Constant Degree of Integration LR3
Chile	28.95	6.93	6.49
	[0.000]	[0.031]	[0.011]
Colombia	23.28	4.83	0.533
	[0.000]	[0.090]	[0.466]
Greece	49.05	52.34	8.99
	[0.000]	[0.000]	[0.003]
India	11.30	14.78	8.89
	[0.023]	[0.001]	[0.003]
Jordan	17.75	8.90	0.34
	[0.001]	[0.012]	[0.560]
Korea	27.34	27.21	1.17
	[0.000]	[0.000]	[0.281]
Malaysia	7.10	5.62	0.77
	[0.131]	[0.060]	[0.380]
Mexico	12.73	10.28	18.54
	[0.013]	[0.036]	[0.000]
Nigeria	6.22	[a]	9.93
	[0.183]	—	[0.002]
Taiwan	23.92	18.21	4.08
	[0.000]	[0.000]	[0.043]
Thailand	4.30	[a]	21.32
	[0.367]	—	[0.000]
Zimbabwe	4.51	8.51	10.25
	[0.342]	[0.014]	[0.001]

[a] The likelihood value is higher in the constrained model. This is possible since the first-stage estimation prevents a complete nesting of the two models.

partially integrated with the degree of integration being constant. We provide a framework which allows for time-varying conditional market integration.

The degree that a national capital market is integrated into world capital markets is notoriously difficult to measure. Some have suggested that the correlation of the local market return with the world return is a measure of integration. However, this is flawed because a country could be perfectly integrated into world markets but have a low or negative correlation because its industry mix is much different from the average world mix.

Others have looked to investment restrictions as an indicator of integration. This measure is problematic because there are numerous types of restrictions, with some being more important than others across different countries. Importantly, the investment restrictions may not be binding. That is, investors may be able to access the national market in other ways. As a result, it may be a mistake to conclude that the market is segmented based on statutory investment restrictions.

We measure the degree of integration directly from the returns data. Our model nests the polar cases of complete integration and complete segmentation. The econometric method allows for the degree of integration to change through time. Our results indicate time-varying integration for a number of countries.

There are a number of possible extensions of our research. Our framework can be modified to allow for multiple sources of risk. An omitted risk factor could potentially mask itself through evidence of time-varying integration. One immediate modification, following Adler and Dumas (1983) and Dumas and Solnik (1995), involves the addition of foreign exchange risk. Indeed, strong assumptions (such as purchasing power parity) are needed in order to justify our basic model in equation (1) (see Stulz (1981b, 1993)).

Our modeling approach can be used to assess the effects of regulatory changes. It is possible to let the regime probabilities be functions of indicator variables that capture policy changes. For instance, Japan abolished many of its capital market restrictions in the 1980s (see Bonser-Neal, Brauer, Neal, and Wheatley (1990) and Gultekin, Gultekin, and Penati (1989)). A number of developing countries removed or relaxed restrictions on foreign equity ownership in the nineties (see Bekaert (1995) and Harvey (1993a)). However, we do not find overwhelming evidence pointing to increased integration (only four of the 12 countries have higher integration measures in the 1990s). Our framework will allow us to test directly whether these policy changes had a discernable affect on the degree of market integration and whether the cost of capital was altered. This research is currently being pursued in Bekaert and Harvey (1995b).

Finally, measuring the degree of financial market integration has implications beyond explaining why expected returns differ across different countries. There is a strong interest in development economics in models that relate capital market restrictions and the stage of financial market development to economic growth.

Economic growth is fundamentally linked to financial integration. A number of recent models show how improved risk sharing leads to higher economic growth.[18] Capital market integration provides the opportunity for better diversification. In a segmented economy, a consumer or a firm may only select low-risk low-expected return investments. With integrated mar-

[18] Examples are Levine (1991) and Saint-Paul (1992). Pagano (1993) presents a detailed review of the literature relating financial markets to economic growth.

kets, individuals shift to high-risk high-expected return projects because they are able to diversify their overall risk (see Obstfeld (1994)).

There is an expanding body of empirical work on the relation between capital market restrictions and economic growth. King and Levine (1993) detail a significant cross-sectional correlation between variables that proxy for both the depth of the financial sector and its development and economic growth. Atje and Jovanovic (1992) find significant correlations bewteen the ratio of stock market trading volume to GDP and economic growth. A problem with this empirical work, which is recognized by the authors, is that it is difficult to specify a set of variables that proxy for capital market restrictions (or capital market openness). Our article provides a new approach to assessing the degree of market integration. The empirical relation between integration and economic development is explored in Bekaert and Harvey (1995c).

Appendix

A. *IFC Emerging Market Equity Indices*

The International Finance Corporation (IFC) began calculating emerging market indices in 1981. The indices, known as the IFC Global (IFCG) Indices, do not take into consideration restrictions on foreign ownership. Recently, the IFC has introduced a second set of indices, the IFC Investable (IFCI) Indices. The IFCI Indices reflect restrictions on ownership limits. For example, if a firm had a market capitalization of U.S. $300 million and the national law restricts foreign ownership to 50 percent of any company, the IFCG uses the full $300 million as the market capitalization while the IFCI uses $150 million.

Since our article studies the integration of the emerging markets in world capital markets, we have chosen to use the IFCG. An additional reason for using the IFCG is the limited sample size of the IFCI (data begin in 1988). However, it is important to understand the restrictions in each market and the methodology used to construct the indices. The following description follows the International Finance Corporation (1993). (See Table AI.)

A.1. Selection Criteria

The IFC selects stocks for inclusion in the indices based on three criteria: size, liquidity, and industry. The indices include the largest and most actively traded stocks in each market, targeting 60 percent of total market capitalization at the end of each year. As a second objective, the index targets 60 percent of trading volume during the year. Size is measured by market capitalization and liquidity is measured by the total value of shares traded during the year.

Only stocks that are listed on one of the major exchanges in the emerging markets are included in the index. The index will not include stocks whose issuing company is headquartered in an emerging market but listed only on foreign markets.

Appendix Table AI

Market Weights in the IFC Indices at the End of March 1993

Total market capitalization is the number of shares multiplied by the end of March 1993 share value for each stock listed in 20 markets. The International Finance Corporation (IFC) creates value weighted indices of a smaller number of stocks within each country. The number of stocks included in each IFC index is also reported along with the market capitalization of the stocks included in the IFC index. Finally, we report the share that each country commands in the IFC emerging market composite index. The source of the data is IFC (1993).

Country	Total Market Capitalization (Millions US$)	Number of Stocks	Market Capitalization (Millions US$)	Weight in IFC Composite
Panel A: Latin America				
Argentina	19,101.8	30	14,994.8	2.9
Brazil	59,488.5	70	37,245.8	7.2
Chile	33,510.6	35	21,658.8	4.2
Colombia	6,571.0	20	4,156.1	0.8
Mexico	132,574.8	74	83,683.6	16.1
Venezuela	6,228.8	17	3,982.6	0.8
Panel B: East Asia				
Korea	105,929.0	134	71,016.7	13.7
Philippines	16,340.6	37	11,528.8	2.2
Taiwan	148,487.9	78	95,244.0	18.4
Panel C: South Asia				
India	59,793.1	108	29,987.8	5.8
Indonesia	14,385.0	41	9,469.1	1.8
Malaysia	100,142.6	66	68,153.6	13.1
Pakistan	7,198.9	64	4,607.5	0.9
Thialand	58,909.0	58	37,271.8	7.2
Panel D: Europe/Mideast/Africa				
Greece	9,928.9	36	6,304.7	1.2
Jordan	3,788.4	29	2,146.3	0.4
Nigeria	871.9	24	552.7	0.1
Portugal	9,988.2	38	7,255.3	1.4
Turkey	13,470.2	36	9,568.2	1.8
Zimbabwe	613.9	21	376.9	0.1
Panel E: IFC Regional Indices				
Composite	806,710.1	995	518,828.2	100.0
Latin America	257,475.4	246	165,721.7	31.9
Asia	511,187.1	586	327,299.3	63.1
Europe/Mideast/Africa	38,661.5	184	26,204.1	5.1

If many stocks meet the liquidity and size criteria, but only one or two are needed, IFC selects the stocks that represent industries that are not well represented in the IFC index.

In a few instances, particularly where multiple classes of stocks are common (e.g., Brazil and Mexico), IFC may include more than one class of stock for the same company even though they are not necessarily actively traded. The purpose is to give a balanced view of the capitalization of companies that have other classes of stock that are actively traded.

It is useful to compare and contrast the criteria used by Morgan Stanley Capital International (MSCI) (see Schmidt (1990)) and the IFC. In constructing the MSCI indices, 60 percent coverage of the total market capitalization of each market is also the first objective. In contrast to the IFC, there is no secondary objective regarding the volume of trading. The second MSCI criteria is that the companies included in the index replicate the industrial composition of the local market. In addition, the MSCI index tries to take a representative sample of large, medium, and small capitalization stocks (instead of just concentrating on the largest capitalization companies). MSCI uses liquidity as a consideration in choosing among the medium and small capitalization stocks. MSCI, like IFC, excludes nondomiciled companies and investment funds. MSCI excludes companies with restricted float due to dominant shareholders or cross-ownership. Similar to IFC, MSCI also publishes "Free" indices that exclude companies whose shares are not readily available to foreign investors.

A.2. Survivorship Bias and Time-Varying Inclusion Criteria

The IFC does not select stocks based on financial history or future expected performance. Nevertheless, any size or liquidity screen will tend to select stocks that have done well (or avoid poor performers). This is the case for all stock market indices. This selection criteria is not a problem if it is done on an ex ante basis. However, the IFC started their indices in 1981 and nine indices began in December 1975. This reconstruction induces an obvious survivorship bias. Harvey (1995) details the survivorship problem induced by the backtracking of the indices. He argues the survivorship problem is not that serious. In many countries, the very high return periods follow the lookback data in 1981.

Another issue focusses on time-varying criteria for including stocks in the indices. There is little publicly available information on how the original stocks were selected for index construction. However, from conversations with people that either worked at the World Bank, or still work at the World Bank, it is clear that the joint size, liquidity, and industry criteria were not used consistently through the years. Our conversations suggest that in 1981 size was the single criterion used in index construction.

A.3. Index Methodology

The IFC indices represent value-weighted portfolios of the stocks in each market. That is, each stock is weighted by its market capitalization in the same way that the Morgan Stanley Capital International (MSCI) country indices are formed (chained Paasche method).

Adjustments in the index divisor are initiated if new shares are issued or rights are declared. The change in the divisor neutralizes the effects of these two issues. The divisor will also change when a stock is added to or deleted from the index.

A.4. Currency Conversion

The IFC indices are calculated in local currencies as well as in U.S. dollars. For most markets, the indices use exchange rates taken from the *Wall Street Journal* or the *Financial Times*. When a multiple exchange rate system exists, the IFC uses the nearest equivalent "free market" rate or the rate that would apply to the repatriation of capital and income. In a few cases, the newspaper rates are inappropriate and the IFC uses rates provided by the IFC's correspondents in each market.

A.5. Price Information

The principal source of prices and changes in capitalization used for the IFC indices is a network of correspondents in each market, including local IFC brokers, investment banks, stock exchanges, and regulatory authorities.

When a stock is not traded on the date of the index, the last traded price is used. When a stock is traded on more than one local exchange, the price used by the IFC is taken from the exchange where the trading was most active.

Some markets, notably Thailand, maintain "Alien" boards to ensure that the total foreign ownership does not exceed a certain share of the total. The IFC has found that trading in stocks on the Alien board is generally thin and stocks often trade at different prices. The IFC asserts that price movements on the Alien board often lag behind those of the "Main" market. In the index construction, the IFC uses the Main Board prices. Additional analysis of the Thai market is presented in Bailey and Jagtiani (1994).

A.6. Index Revisions

Once a year, the individual component stocks are reviewed to see if the index meets the objective criteria. Additions and deletions are made as necessary. Although the target for the global indices is to attain 60 percent of market value, the actual coverage will vary. To ensure consistency, the IFC will not generally add or delete stocks unless the coverage of the global index drops below 50 percent or rises about 70 percent of the total market capitalization.

On a quarterly basis, the IFC will drop stocks that have been suspended, delisted, merged, dropped, or otherwise made irrelevant. Newly listed stocks that result from a merger or split of a stock already in the index will be added. The IFC may add a newly listed stock to the index between the regular quarterly revisions if it is unusually large. Such changes are announced one week prior to their actual inclusion in the indices.

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