TIME-VARYING WORLD
MARKET INTEGRATION

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ABSTRACT

We propose a conditional measure of capital market integration that allows us to characterize both the cross-section and time-series of expected returns in developed and emerging markets. Our measure, which arises from a conditional regime-switching model, 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. Our results suggest that a number of emerging markets exhibit time-varying integration. Interestingly, some markets appear to be 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.

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1. Introduction

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.

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 risks 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 which assumes market are segmented is one that "tests" a model like the Capital Asset Pricing Model of Sharpe (1964), Lintner (1965) and Black (1970), using one country's data. Indeed, all of the seminal U.S. asset pricing studies assume that the U.S. 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 1970's, in the 1980's the U.S. equity capitalization dropped below 50% of the world market capitalization. Indeed, Japan's market capitalization exceeded the U.S. (albeit briefly) in 1989.

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The second class of asset pricing studies assumes that world capital markets are perfectly integrated. These include studies of a world CAPM [see Harvey (1991a) and references therein], a world CAPM with exchange risk [see Dumas and Solnik (1993) and Dumas (1994)], world arbitrage pricing theory [see Solnik (1980) and Senbet et al. (1986)], world multibeta models [Ferson and Harvey (1993, 1994a,b)] 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 (1991a). 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 return. In a perfectly integrated market, only the covariance counts. In segmented markets, the variance is the relevant measure of country risk. Our inte-

gration measure is a time-varying weight which is applied to these two moments. The model allows for differing prices of variance risk across countries which depends on country-specific information and a world price of covariance risk which depends 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.

Our paper is organized as follows. In the second section, the asset pricing framework is presented. An outline of the econometric model is also detailed. The data on 33 national equity markets are described in the third section. In the fourth section, the results are analyzed. Some concluding remarks are offered in the final section.

2. Asset pricing with time-varying market integration

2.1 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 that:

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

(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 equity 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. All expectations are conditioned on Z_{t-1} – the information that investors use to set prices at time $t-1$. The risk-free rate has zero conditional variance because the return is determined at $t-1$. This model is tested in Harvey (1991a).

In the completely segmented market and under the same assumptions behind (1):

$$E_{t-1}[r_{i,t}^A] = \lambda_{i,t-1} \text{Cov}_{t-1}[r_{i,t}^A, r_{i,t}].$$

(2)

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Security A is now priced off its covariance with the return of the market portfolio in country i, r_i and λ_i is the local price of risk. At the national level,

$$E_{t-1}[r_{i,t}] = \lambda_{i,t-1}\text{Var}_{t-1}[r_{i,t}]. \tag{3}$$

Merton (1980) argues that λ_i is a measure of the representative investor's relative risk aversion. The model suggests that expected returns in a segmented market are determined by the variance of returns 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.

Econometrically, these models can be combined:

$$E_{t-1}[r_{i,t}] = \phi_{i,t-1}\lambda_{t-1}\text{Cov}_{t-1}[r_{i,t}, r_{w,t}] + (1 - \phi_{i,t-1})\lambda_{i,t-1}\text{Var}_{t-1}[r_{i,t}], \tag{4}$$

where $\phi_{i,t-1}$ is the conditional integration measure and falls in the interval $[0,1]$. If $\phi_{i,t-1}$ is one, it implies that only covariance with the world porfolio is priced and we can reject the null hypothesis of market segmentation. If $\phi_{i,t-1}$ is zero, then only variance is priced. This is consistent with a segmented capital market. In addition, the price of risks λ_{t-1} and $\lambda_{i,t-1}$ are allowed to differ. While (4) is not implied by any particular equilibrium asset pricing theory, it could be a reasonable approximation of the expected returns.[1] In addition, our methodology allows us to recover the fitted integration parameters and hence characterize the time path of integration in the 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 Errunza and Losq (1985) and Eun and Janakiramanan (1986). 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 U.S. and Japan. They use the definition of covariance to show, for example, that the conditionally

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[1] It is an approximation in at least two ways. First, a one factor model is imposed. Section 2.3.3 considers the implications of a multifactor formulation. Second, the hedging terms which are related to the covariances between local and world marginal rates of substitution are ignored.

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 modelling the U.S. and Japan, explicitly using the market share weights is less important for most of the markets in our sample since they are so small.

Following models like Stulz (1981), the returns in (1)–(4) should be real. Given that reliable inflation data in many of the countries that we study is 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.

2.2 Regime switching and integration

The model for expected returns in (4) may be considered in the class of regime-switching models. In the first regime, markets are integrated and expected returns adhere to (1). In the second regime, markets are segmented and expected returns are given by (2). Let S_t^i be an unobserved state variable which takes on the value of one when markets are integrated and a value of two when markets are segmented. Then the parameter $\phi_{i,t-1}$ can be interpreted as the conditional probability of being in regime 1, $\phi_{i,t-1} = \text{Prob}[S_t^i = 1 | \mathcal{Z}_{t-1}]$.

Several models are available to estimate $\phi_{i,t-1}$. In the regime switching model developed by Hamilton (1988, 1989, 1990), S_t^i follows a Markov process with constant transition probabilities. Recently, Diebold, Lee and Weinbach (1992), Ghysels (1992) and Gray (1993) have extended the Hamilton model to allow for time-varying transition probabilities. Gray (1993) shows that all these models are special cases of a general finite mixture distribution model with time-varying weights, i.e. $\phi_{i,t-1} = \phi_i(Z_{t-1}^*)$ with $\phi_i(\cdot)$ a functional form that constrains $\phi_{i,t-1}$ to be between zero and one and Z_{t-1}^* a set of variables in $\mathcal{Z}_{t-1}$.

The possibility of time-varying transition probabilities allows for an alternative interpretation of $\phi_{i,t-1}$. Whether a market is integrated with world capital

markets or segmented is greatly influenced by economic and financial market policies followed by its government or other regulatory institutions. Hence, $\phi_{i,t-1}$, can be interpreted as a policy weight, varying with policies affecting the degree of market integration. An obvious example is foreign ownership restrictions, often imposed by developing countries.

Barriers to investment (by foreigners in local markets or local participants in foreign markets) can take many forms. Moreover, not all barriers to foreign equity investment necessarily segment markets from the world capital market. For instance, Bekaert (1994) shows that the presence of country funds 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. However, given that the asset pricing model is correctly specified, the regime switching model allows us to infer the degree of market segmentation.

To infer $\phi_{i,t-1}$ from the data, we explore two different regime switching models. The first is the standard Hamilton model. Although the switching probabilities are time invariant, the regime probability 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 (1993) derives the following recursive representation for the regime probability:

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

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, $\mathcal{Z}_{t-1}$

In the second formulation, we allow the transition probabilities P and Q to

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be time varying. In particular, we model them as logistic functions of Z_{t-1}^*:

$$
\begin{aligned}
P_t &= \frac{\exp(\beta_1' Z_{t-1}^*)}{1 + \exp(\beta_1' Z_{t-1}^*)} \\
Q_t &= \frac{\exp(\beta_2' Z_{t-1}^*)}{1 + \exp(\beta_2' Z_{t-1}^*)}
\end{aligned}
\tag{6}
$$

where β_j, $j = 1, 2$, are vectors of parameters.

In implementing this model, we let Z_{t-1}^* be a subset of Z_{t-1}^i where Z_{t-1}^i is a collection of information variables specific to country i which 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. 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.

2.3 Estimation issues

2.3.1 The likelihood function

To complete the model described in (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 $R_{i,t} = [r_{i,t}, r_{w,t}]'$:

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

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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,

$$\begin{aligned} E[\mathbf{e}_t^I\mathbf{e}_t^{I\prime}|\mathcal{Z}_{t-1}] &= \mathbf{\Sigma}_t^I \\ E[\mathbf{e}_t^S\mathbf{e}_t^{S\prime}|\mathcal{Z}_{t-1}] &= \mathbf{\Sigma}_t^S \end{aligned} \tag{9}$$

The conditional variance dynamics are modelled as ARCH(k) following Baba, Engle, Kraft and Kroner (1989) (BEKK):[2]

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

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:

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

and

$$\mathbf{A}^S(2,1) = 0.$$

The first and second 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

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[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.

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 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, 1991a) suggests that the price of risk is time varying. In the most general version of the model, we let:

$$\lambda_{t-1} = exp(\delta' Z_{t-1})$$
$$\lambda_{i,t-1} = exp(\delta'_i Z^i_{t-1})$$

(12)

where Z represents global information variables and 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 (1993) 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:

$$log\, L(\mathbf{R}_{i,T}) = \sum_{t=1}^{T} log\{\phi_{i,t-1}g_{1,t} + (1 - \phi_{i,t-1})g_{2,t}\}$$

$$\text{with} \quad g_{1,t} = (2\pi)^{-1}|\Sigma^I_t|^{-1/2}exp\{-\frac{1}{2}(\mathbf{e}^{I\prime}_t(\Sigma^I_t)^{-1}\mathbf{e}^I_t)\} \ , \quad (13)$$

$$g_{2,t} = (2\pi)^{-1}|\Sigma^S_t|^{-1/2}exp\{-\frac{1}{2}(\mathbf{e}^{S\prime}_t(\Sigma^S_t)^{-1}\mathbf{e}^S_t)\}$$

$$\mathbf{R}_{i,T} = [R_{i,1}, R_{i,2}, \ldots, R_{i,T}]$$

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[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 currently being explored in Bekaert and Harvey (1994a).

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

$$\Theta = [\delta', \delta_i', \mathrm{Vech}(\mathbf{C}^I), \mathbf{C}^S(1,1), \mathbf{C}^S(1,2), \mathbf{A}^I(1,1), \mathbf{A}^I(2,1), \mathbf{A}^I(2,2), \beta']',$$

where β 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 normally 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 (13) directly, we proceed in two steps. First, we estimate $C^I(2,2)$, $A^I(2,2)$ and δ using the world market return and the world information variables, $\mathbf{Z}$. Second, we estimate (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.

2.3.2 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 price 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

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[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 (1993) 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.

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 where the time t disturbance $\mathbf{e}_t$ is orthogonal to information $\mathcal{Z}_{t-1}$ 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 $\mathcal{Z}_{t-1}$. 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 $\mathcal{Z} = Z$, $\mathcal{Z} = Z^i$, and $\mathcal{Z} = [Z, Z^i]$. Misspecification of the world market return equation is also possible. Therefore, we also regress $e_{w,t}$ onto Z_{t-1}.

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:

$$r_{i,t} = \zeta' \mathcal{Z}_{t-1} + \phi_{i;t-1}\lambda_{t-1}\mathrm{Cov}_{t-1}[r_{i,t}, r_{w,t}] + (1 - \phi_{i,t-1})\lambda_{i,t-1}\mathrm{Var}_{t-1}[r_{i,t}], \quad (14)$$

and we test whether $\zeta = 0$. The choices for $\mathcal{Z}$ 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.

We then estimate three alternative models embedded in the general specification (5) to (11). 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

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

[6] In the model with time-varying transition probabilities, this restriction cannot be imposed.

model is rejected, the constant prices of risk and no ARCH models are estimated using time-varying transition probabilities.

2.3.3 Other estimation approaches

Here we discuss some alternative approaches to estimation and extensions of the model which we leave for further research. A third version of the regime switching model allows the regime probabilities to be modelled directly as:

$$\phi_{i,t-1} = \frac{\exp(\gamma_i' Z_{t-1}^i)}{1 + \exp(\gamma_i' Z_{t-1}^i)}, \tag{15}$$

where γ_i is a vector of coefficients. In this formulation, it makes sense to condition the regime probabilities on local information variables.

Given this structure of the regime probabilities, it is possible to estimate a version of the time-varying integration model with the generalized method of moments. The following moment conditions can be explored:

$$u_{it} = r_{it} - \phi_{i,t-1}\varepsilon_{it}\varepsilon_{wt}(\delta' Z_{t-1}) - (1 - \phi_{i,t-1})\varepsilon_{it}^2(\delta_i' Z_{t-1}^i) \tag{16}$$

where the integration parameter, $\phi_{i,t-1}$, is defined as above and $\varepsilon_{it}, \varepsilon_{wt}$ are unexpected returns on country i's national index and the world index, respectively, which may be obtained in a first stage estimation. The world price of risk is $\delta' Z_{t-1}$ and the local price of risk is $\delta_i' Z_{t-1}^i$. These functions can be constrained to be positive using the exponential function.[7]

For each asset, γ_i and δ_i must be estimated. In addition, δ are the common coefficients which form the world price of risk. To identify this model, a number of countries must be examined simultaneously. This makes the estimation difficult.

Our framework can be extended to allow for multiple sources of risk. Indeed, an omitted risk factor could potentially mask itself through evidence of time-varying integration. One immediate extension, following Adler and Dumas (1983,

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[7] In this econometric specification, the dynamics of the conditional covariances and variances need not be explicitly modelled. This follows Campbell (1987), Harvey (1989), Dumas and Solnik (1992) and, recently, Kan and Zhang (1994).

equation 14), Dumas and Solnik (1994) and Dumas (1994), involves the addition of foreign exchange risk:

$$E_{t-1}[r_{i,t}] = \phi_{i,t-1}\lambda'_{t-1}\text{Cov}_{t-1}[r_{i,t}, \mathbf{f}_t] + (1 - \phi_{i,t-1})\lambda_{i,t-1}\text{Var}_{t-1}[r_{i,t}], \qquad (17)$$

where

$$\mathbf{f} = \begin{pmatrix} r_{w,t} \\ r_{c,i,t} \end{pmatrix} \qquad \lambda_t = \begin{pmatrix} \lambda_{w,t} \\ \lambda_{c,t} \end{pmatrix}$$

and $r_{c,i,t}$ is the currency excess return for country i, $\lambda_{w,t}$ is the world price of covariance risk, and $\lambda_{c,t}$ is the world price of foreign exchange rate risk. According to the Dumas and Solnik model, $K - 1$ currency returns should be priced, where K is the number of countries. This leads to an intractable number of parameters. Hence, it makes sense to follow Ferson and Harvey (1993, 1994a,b) and Bailey and Jagtiani (1994) and present the local currency returns computed against a trade-weighted basket of foreign currencies.

Implementing the Dumas and Solnik (1994) approach explicitly accounts for the role of foreign exchange rate risk. Indeed, strong assumptions (such as purchasing power parity) are needed in order to justify (1) [see Stulz (1981, 1993)]. The Dumas and Solnik model assumes perfect market integration and is invariant to the numeraire currency. However, if the market is segmented, the currency issue arises again. Our model is ideally suited to jointly address the role of currency risk and market integration.

3. Data and summary statistics

3.1 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. Our study focusses on twelve emerging markets: Chile, Colombia, Greece, India, Jordan, Korea, Malaysia, Mexico, Nigeria, Taiwan, Thailand and Zimbabwe are presented.[8]

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[8] The IFC tracks 20 emerging markets. Three of their countries, Indonesia,

The summary statistics are presented in Table 1 for the total available data for each country. Most of the MSCI data begins in December 1969 and earliest available data for 7 of the 12 emerging countries is December 1975. Our analysis will concentrate on the U.S. dollar returns.[9] 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 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% (Colombia) to 3% (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%. In the IFC emerging sample, 4 countries (Chile, Colombia, Mexico, and Taiwan) have average returns above 30%.

Emerging market returns are characterized by high volatility. Standard deviations range from 18% (Jordan) to 53% (Taiwan). In contrast for the MSCI countries, volatility ranges between 15% and 42%. There are 8 emerging countries with volatility higher than 30%.[10]

The emerging market returns are also more autocorrelated. In the MSCI sample of 18 countries with data from December 1969, there are only five countries with first-order autocorrelation that exceeds 10%. In contrast, 6 of the 12 emerging countries have with autocorrelations greater than 10%. There are two countries with autocorrelations above 20% (Colombia and Mexico). This suggests that the returns in many of these countries are predictable (to some extent) based on past returns alone.

Portugal and Turkey have very short samples.

[9] Calculating the returns in U.S. dollars eliminates the local inflation. However, the U.S. inflation remains in the returns.

[10] Bekaert and Harvey (1994a) explore the reasons why volatility is different in emerging versus developed markets and detail the time-series characteristics of emerging market volatility.

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3.2 Predictability

A number of studies have documented the existence of predictable variation in developed country returns.[11] Recently, evidence of predictable variation in emerging market returns has been documented in Bekaert (1994), Buckberg (1993) and Harvey (1993a,b, 1994).

In our econometric model, we separate the total information set, $\mathcal{Z}$, into local components, $\mathcal{Z}^i$ and global components, $\mathcal{Z}$. 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.[12]

The set of local information variables include: a constant, local equity returns, local exchange rate changes, and 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% average correlation among dividend yields in the MSCI countries.

Table 2 presents heteroskedasticity-consistent tests of the null hypothesis that expected returns are constant. In the first panel, tests are conducted on the developed markets. The multivariate test[13] of no predictability using the global infor-

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

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

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

mation variables for 18 markets (Finland, Ireland and New Zealand are excluded because their data begins 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.

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

4. Results

4.1 The world price of covariance risk

Table 3 presents the estimation of the ARCH-M model for the world price of covariance risk:

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

$$\lambda_{t-1} = exp(\delta' Z_{t-1})$$

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

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

and Z represents the global information variables. Consistent with the evidence presented in Harvey (1991a), the hypothesis that the world price of risk is constant is easily rejected. This is also seen in Figures 1a and 1b which plot the fitted prices of risk. Interestingly, there is 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.

There is less 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 rejected at conventional significance levels (p-value is 0.15). The fitted values of the full model and the no ARCH model are presented in figure 1. Both series exhibit the

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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.

4.2 Estimation

The results for estimating the standard Hamilton model are presented in table 4. 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 (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 10 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.[14]

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 4 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.

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[14] The nonlinear estimation, along with the large number of starting values, demanded two weeks of workstation time – per country.

4.2.1 Chile

The average value of the integration parameter for Chile is 0.59 and in recent years the value has dropped to .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% 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% 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, namely 25%. Bekaert (1994), 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 4.

4.2.2 Colombia

The results for Colombia also suggest that the market is more segmented than integrated. Over the entire sample, the ex ante probability of integration never exceeds 0.20. The P parameter is zero reflecting the difficulty in estimating the probability of integration given that the previous state was integrated.

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% of trading volume. The potential liquidity problems are also evident from the remarkable 49% serial correlation in returns reported in table 1.

While there are some recent positive developments in Colombia such as recent 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.

4.2.3 Greece

Greece is no longer an emerging market with US$5,500 GDP per capita 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, Greece fell within the emerging markets category. The evidence presented in table 4 suggests that the Greek market is integrated into world capital markets. The integration parameter in the 1990's 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 $US9.5 billion at the end of 1992. There is a large foreign participation in the stock market (about 20% of shares are owned by foreigners). Finally, there is reasonable liquidity with $9 million in average daily trading in 1992.

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

On the other hand, India is a very poor country with only US$300 of per

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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 tensions with Pakistan could also work against foreign investors participating in the Indian market.

The results in table 4 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.

4.2.5 Jordan

The estimates suggest that Jordan is not fully integrated into world capital markets. In table 4, the recent degree of integration is 79%. From figure 2, the degree of integration has fluctuated between 40% and 90% over the past five years. The Jordanian market is small with a market capitalization of US$3.2 billion at the end of 1992. Foreigners are restricted to owning up to 49% of equity (with the exception of tourism and agriculture where there are no limits). Importantly, 85% of equities is owned by Jordanians. The remaining 15% is thought to be owned mainly by investors from other Arab states. There are no 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.

4.2.6 Korea

Korea also fails to qualify as an emerging market with per capita GDP exceeding US$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 3 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% 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% in so-called unlimited industries and 8% in limited industries (which includes communications and defence). Recently, the 10% ceiling was raised to 25% for 45 firms which hit the 10% 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.[15]

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.

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[15] Errunza, Losq and Padmanabhan (1992) report unconditional tests of market integration (assuming the degree of integration is constant). They reject the polar cases of complete integration and complete segmentation for three of our countries, Chile, Greece and Korea. They fail to reject their hypothesis of mild segmentation for these same countries.

4.2.7 Malaysia

For Malaysia, our priors tilted towards integration. The equity market is large (US$94 billion at the end of 1992) with good trading volume (US$21.8 billion in 1992). Malaysia has experienced very mild inflation averaging only 4.6% 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%. Although foreign investment is limited by the Foreign Investment Committee to 30% of equity, it is not clear that this constraint is binding in our sample. In addition, foreigners can access 11 closed end funds, 7 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 problematic, the results in table 4 suggest that the market is integrated. The 1990's integration parameter is 0.79 and has been fairly stable from the start of our data.

4.2.8 Mexico

The results for Mexico are surprising. The model estimates suggest that Mexico's equity market is segmented. There is a slight upturn since 1991 Figure 2. Today, Mexico has one of the highest capitalized markets (US$139 billion at the end of 1992) with US$171 million in average daily trading volume. There are 36 ADRs and 6 US dollar based country funds. All of these factors point toward market integration.

While Mexican stocks get a lot of attention in the United States most observers don't realize that before 1991 only 1 Mexican ADR was trading. In the 1980s, there was only one country fund available. The major reforms are fairly recent. The Mexican stock market was made 100% investable (with the exception of certain key sectors such as banking) in May of 1989 and the dual exchange rate was abolished in November 1991. In addition, there has been a lot of economic turmoil. The debt crisis in 1982 deterred foreign investment. Mexico had the

fourth highest inflation rate over the past 6 years (behind Brazil, Argentina and Turkey). Given that most of the liberalizations occurred at the end of our sample, the results appear more credible.

4.2.9 Nigeria

We chose to examine Nigeria because we had a strong prior that this was the most likely market to be segmented. Per capita GDP is only US$295 in 1990 and over 80% of the economy is linked to petroleum. The results in table 4 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% investable and ranks Nigeria last among the emerging markets. Liquidity is extremely thin. Only 1% of market capitalization traded in 1992 (the average daily trading volume was only US$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.

4.2.10 Taiwan

Taiwan was another country where it was difficult to form a prior opinion about the degree of integration. Factors favoring integration included the high market capitalization (US$101 billion at the end of 1992) and the very large trading volume (US$214 billion). In addition, Taiwan no longer qualifies as an emerging market with 1992 GDP per capita of US$8815. 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.

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 4 approved investment funds. In

January 1991, direct investment by institutional investors 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% 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 8 closed-end funds, 9 open-end funds, and 4 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 begins in 1985) makes both estimation and inference difficult.

4.2.11 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 dramatic 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 [see Bailey and Jagtiani (1994)]. 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% 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% of the freely floating shares.[16] The market is large (US$58.3 billion in December 1992, 5th largest of the emerging markets) and very liquid (US$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.

4.2.12 Zimbabwe

We chose to examine Zimbabwe, as we did Nigeria, because of a strong prior that the country was not integrated into world capital markets. Zimbabwe is the second poorest country in our sample with per capita GDP of US$621 per year in 1991 (Nigeria is last with US$295). The market capitalization is the smallest in the sample at US$600 million and only US$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 2 confirms our prior that the market is not fully integrated. The average integration is 57% in the 1990's. 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 which 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 which remains unexplained.

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

4.3 Diagnostics

Table 5 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 (14). The test essentially adds a time-varying intercept to (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 risk premiums. Given a rejection of the specification, we need to exercise caution in interpreting the estimated degree of integration.

The specification test 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% 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 very 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 vari-

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ables. 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 local and world information. The residual R^2s are about the same (6%) – 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^2 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%. 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. Disturbingly, the local price of risk is imprecisely measured and large in magnitude.

Neither the Wald test or 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.

There is no hint 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.

4.4 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 integration. Table 6 presents tests of the following regression models:

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

$$\Delta s_{t+1} = \alpha_0^* + \alpha_1^* \Delta s_t + \alpha_2^* i_t + e_t^*$$

(18)

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_2 (coefficient on the estimated integration) which has one degree of freedom.

In (18), $\hat{\phi}_t$ is a generated regressor and as a result, the OLS standard errors are inappropriate. However, Pagan (1984) shows that the two stage estimators are consistent but may not have the same limiting distribution as the maximum likelihood estimator. In general, OLS will understate the true standard errors. Hence, if we fail to reject $\alpha_3 = 0$ with OLS standard errors, we would surely fail to reject with the adjusted standard errors. Hence, the tests we present in (18) will be conservative.[17]

The results in Table 6 suggest there is little evidence that exchange rate changes and the integration measure are interrelated. In four of 12 countries, do the tests rejected the hypothesis that $\alpha_3 = 0$ (Chile, Colombia, Korea and Malaysia). The p-values of these four countries' tests are all above 3%. Given that the standard errors are understated, it is unlikely that the adjusted standard

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[17] In some instances, testing $\alpha_3 = 0$ will yield correct inferences because under the null hypothesis, $\hat{\phi}_t = 0$. See Pagan (1984).

errors would present strong evidence against the null hypothesis in even these four countries.

4.5 Estimation of the constrained alternatives

Table 7 presents likelihood ratio tests of three specific alternative hypotheses: constant prices of risk, constant variance matrices and constant degree of integration. Some summary statistics on the estimated parameters are also presented.

The hypothesis that the price of local volatility is constant are rejected at the 5% 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, the estimation for this country was ill-behaved and we should be cautious in drawing conclusions.

Although the constant prices of risk assumption is rejected in most countries, the parameter estimates are presented. The estimation imposes positivity and forces the world price of risk to be identical across all countries. In 10 of the 12 countries, local prices of risk range from 43.59 in Jordan to 1.60 in Greece. The local prices of risk are more than two standard errors from zero in nine of these countries (Thailand's is 1.5 standard errors from zero).

The remaining two countries, Taiwan and Malaysia, have very low and very high local prices of risk. The low local price of risk in Taiwan is not necessarily problematic because, first, the restriction is rejected and, second, the evidence in table 4 suggests that Taiwan is integrated. As the degree of integration rises, it becomes more difficult to estimate the local price of risk parameters (because a very small weight is placed on them in the estimation).

Malaysia's constant local price of risk is far too high to be considered reasonable. The estimation for this country was ill-behaved and we are not certain that we obtained the global optimum. As a result, we need to be extra cautious in interpreting the results for Malaysia.

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The hypothesis that the variance matrices are constant is also tested with a likelihood ratio in Table 7. Constant variance matrices are rejected for eight of the ten countries for which this test was feasible. The hypothesis is rejected at the 10% 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 for the first three countries.

5. 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 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.

We do believe that information on investment regulations is useful. In fact, our asset pricing framework can be used to assess the effects of regulatory changes. It is possible to let the regime probabilities to 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 Campbell and Hamao (1992)]. A number of developing countries removed or relaxed restrictions on foreign equity ownership in the nineties [see Bekaert (1994) 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 (1994b).

References

Adler, Michael and Bernard Dumas, 1983, International portfolio selection and corporation finance: A synthesis, *Journal of Finance* 38, 925–984.

Baba, Yoshihisa, Robert F. Engle, Dennis F. Kraft and Kenneth F. Kroner, 1989, Multivariate simultaneous generalized ARCH, Working paper, University of California, San Diego, CA.

Bailey, Warren and Julapa Jagtiani, 1994, Foreign ownership restrictions and premiums for international investment: Some evidence from the Thai capital market, *Journal of Financial Economics*, forthcoming.

Bansal, Ravi, David Hsieh and S. Viswanathan, 1993, A new approach to international arbitrage pricing, *Journal of Finance* 48, 1719–1747.

Bekaert, Geert, 1994, Market integration and investment barriers in emerging equity markets, *World Bank Economic Review*, forthcoming.

Bekaert, Geert and Campbell R. Harvey, 1994a, Volatility control in emerging market equity portfolios, Working notes, Duke University and Stanford University.

Bekaert, Geert and Campbell R. Harvey, 1994b, The effect of liberalization policies on capital market integration and the cost of capital in emerging markets, Working notes, Duke University and Stanford University.

Bekaert, Geert and Robert Hodrick, 1992, Characterizing predictable components in excess returns on equity and foreign exchange markets, *Journal of Finance* 47, 467–509.

Black, Fischer, 1972, Capital market equilibrium with restricted borrowing, *Journal of Finance* 45, 444-455.

Bollerslev, Tim, and Jeffrey M. Wooldridge, 1992, Quasi-maximum likelihood estimation and inference in dynamic models with time-varying covariances, *Econometric Reviews* 11, 143–172.

Bonser-Neal, Catherine, Gregory Brauer, Robert Neal and Simon Wheatley, 1990, International investment restrictions and closed-end country fund prices, *Journal of Finance* 45, 523-548.

Buckberg, Elaine, 1993, Emerging stock markets and international asset pricing, in Stijn Claessens and Sudarshan Gooptu, eds., *Portfolio investment in developing countries*, (Washington: World Bank).

Campbell, John Y., 1987, Stock returns and the term structure, *Journal of Financial Economics* 18, 373–400.

Campbell, John Y. and Yasushi Hamao, 1992, Predictable bond and stock returns in the United States and Japan: A study of long-term capital market integration, *Journal of Finance* 47, 43–70.

Chan, K. C., G. Andrew Karolyi and René Stulz, 1992, Global financial markets and the risk premium on U.S. equity, *Journal of Financial Economics* 32, 137–168.

Cumby, Robert E. and Jack D. Glen, 1990, Evaluating the performance of international mutual funds, *Journal of Finance* 45, 497–521.

Cumby, Robert E. and Anya Khanthavit, 1992, A Markov switching model of market integration, Unpublished working paper, New York University.

Diebold, F. X., J-H. Lee and G. C. Weinbach, 1992, Regime switching with time-varying transition probabilities, in C. Hargreaves, ed., *Nonstationary time series analysis and cointegration*, (London: Oxford University Press) forthcoming.

Dumas, Bernard, 1994, A test of the international CAPM using business cycles indicators as instrumental variables, in Jeffrey Frankel, ed., *The Internationalization of Equity Markets*, (Chicago: University of Chicago Press), forthcoming.

Dumas, Bernard and Bruno Solnik, 1992, The world price of foreign exchange rate risk, Working paper, Group HEC, February.

Dumas, Bernard and Bruno Solnik, 1994, The world price of foreign exchange rate risk, Working paper, Group HEC.

Engel, Charles M. and Jeffrey A. Frankel, 1984, Do asset demand functions optimize over the mean and variance of real returns? A six currency test, *Journal of International Economics* 17, 309–323.

Engle, Robert F., David M. Lilien and Russell P. Robbins, 1987, Estimating time varying risk premia in the term structure: The ARCH-M model, *Econometrica* 55, 391–407.

Errunza, Vihang R. and Etienne Losq, 1985, International asset pricing under mild segmentation: Theory and test, *Journal of Finance* 40, 105–124.

Errunza, Vihang R., Etienne Losq and Prasad Padmanabhan, 1992, Tests of integration, mild segmentation and segmentation hypotheses, *Journal of Banking and Finance* 16, 949–972.

Eun, Cheol S. and S. Janakiramanan, 1986, A model of international asset pricing with a constraint on the foreign equity ownership, *Journal of Finance* 41, 897–914.

Everitt, B. S. and O. J. Hand, 1981, Finite mixture distributions, London: Chapman and Hall.

Ferson, Wayne E. and Campbell R. Harvey, 1993, The risk and predictability of international equity returns, *Review of Financial Studies* 6, 527–566.

Ferson, Wayne E. and Campbell R. Harvey, 1994a, Sources of risk and expected returns in global equity markets, *Journal of Banking and Finance*, forthcoming.

Ferson, Wayne E. and Campbell R. Harvey, 1994b, An exploratory investigation of the fundamental determinants of national equity market returns, in Jeffrey Frankel, ed., *The Internationalization of Equity Markets*, (Chicago: University of Chicago Press), forthcoming.

Frankel, Jeffrey A., 1982, In search of the exchange risk premium: A six currency test assuming mean-variance optimization, *Journal of International Money and Finance* 1, 255–274.

Gibbons, Michael R. and Wayne E. Ferson, 1985, Tests of asset pricing models with changing expectations and an unobservable market portfolio, *Journal of Financial Economics* 14, 217–236.

Gray, Stephen F., 1993, Capturing the nonlinearities in short-term interest rates: A new class of regime-switching models, Unpublished working paper, Stanford University.

Ghysels, Eric, 1993, A time-series model with periodic stochastic regime switches, Working

paper, University of Montréal.

Hamilton, J. D., 1988, Rational expectations econometric analysis of changes in regime: An investigation of the term structure of interest rates, *Journal of Economic Dynamics and Control* 12, 385–423.

Hamilton, J. D., 1989, A new approach of the economic analysis of nonstationary time series and the business cycle, *Econometrica* 57, 357–384.

Hamilton, J. D., 1990, Analysis of time series subject to changes in regime, *Journal of Econometrics* 45, 39–70.

Hansen, Lars P., 1982, Large sample properties of generalized method of moments estimators, *Econometrica* 50, 1029–1054.

Hansen, Lars P. and Robert J. Hodrick, 1983, Risk averse speculation in forward foreign exchange markets: An econometric analysis of linear models, in Jacob A. Frenkel, ed.,: *Exchange rates and international macroeconomics* (University of Chicago Press, Chicago, IL).

Harvey, Campbell R., 1989, Time-varying conditional covariances in tests of asset pricing models, *Journal of Financial Economics* 24, 289–317.

Harvey, Campbell R., 1991a, The world price of covariance risk, *Journal of Finance* 46, 111–157.

Harvey, Campbell R., 1991b, The term structure and world economic growth, *Journal of Fixed Income* 1, 4–17.

Harvey, Campbell R., 1993a, Portfolio enhancement with emerging markets and conditioning information, in Stijn Claessens, and Sudarshan Gooptu, eds., *Portfolio investment in developing countries*, (Washington: World Bank).

Harvey, Campbell R., 1993b, Conditional asset allocation in emerging markets, Working paper, Duke University, Durham, NC.

Harvey, Campbell R., 1994, Predictable risk and returns in emerging markets, Working paper, Duke University, Durham, NC.

Harvey, Campbell R., Bruno H. Solnik and Guofu Zhou, 1994, What determines expected international asset returns?, Working paper, Duke University, Durham, NC.

International Finance Corporation, 1993, IFC index methodology, Washington: World Bank.

Kan, Raymond and Chu Zhang, 1994, A test of conditional asset pricing models, Working paper, University of Alberta.

Kirby, Christopher M., 1994 A multivariate analysis of predictability in stock and bond returns, Working paper, Duke University, Durham, NC.

Lintner, John, 1965, The valuation of risk assets and the selection of risky investments in stock portfolios and capital budgets, *Review of Economics and Statistics* 47, 13–37.

Merton, Robert C., 1980, On estimating the expected return on the market: An exploratory investigation, *Journal of Financial Economics* 8, 323–361.

Pagan, Adrian, 1984, Econometric issues in the analysis of regressions with generated regressors, *International Economic Review* 25, 221–247.

Sharpe, William, 1964, Capital asset prices: A theory of market equilibrium under conditions of risk, *Journal of Finance* 19, 425–442.

Solnik, Bruno, 1974, An equilibrium model of the international capital market, *Journal of Economic Theory* 8, 500–524.

Stulz, René, 1981, A model of international asset pricing, *Journal of Financial Economics* 9, 383-406.

Stulz, René, 1993, International portfolio choice and asset pricing: An integrative survey, Working paper, Ohio State University, Columbus, OH.

White, Halbert, 1982, Maximum likelihood estimation of misspecified models, *Econometrica* 50, 1–26.

World Bank, 1993, Emerging stock markets factbook, International Finance Corporation, Washington, D.C.

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Table 1

Means, standard deviations and autocorrelations of international equity returns through December 1992

Country	Start	Arith. mean	Geo. mean	Std. dev.	Autocorrelation					
					ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	ρ_{24}

Morgan Stanley Capital International-Developed
U.S. dollar returns

Country	Start	Arith. mean	Geo. mean	Std. dev.	ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	ρ_{24}
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

International Finance Corporation-Emerging
U.S. dollar returns

Country	Start	Arith. mean	Geo. mean	Std. dev.	ρ_1	ρ_2	ρ_3	ρ_4	ρ_{12}	ρ_{24}
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

Table 2

Predictability using local and global information

Country	Exclude global χ^2	p-value	Exclude local χ^2	p-value	Exclude global+local χ^2	p-value
Morgan Stanley Capital International-Developed U.S. dollar returns						
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	-	-	-	-
International Finance Corporation-Emerging U.S. dollar returns						
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	-	-	-	-

All tests heteroskedasticity consistent. Country tests are based on univariate regressions.

Table 3

The world price of covariance risk

The model estimated is:

$$r_{w,t} = \lambda_{t-1} \text{Var}_{t-1}[r_{w,t}] + \epsilon_{w,t}$$
$$\lambda_{t-1} = exp(\delta' Z_{t-1})$$

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

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

A: Full model

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

B: Constant variance model

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

C: Constant price of risk

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

The χ^2 test in panel B has one degree of freedom. The test in panel C has 4 degrees of freedom.

Table 4

Estimation of the model with constant transition probabilities

Country	Transition probabilities		χ^2 [p-value]	Degree of integration full sample	Degree of integration post-1990
	P	Q			
International Finance Corporation-Emerging U.S. dollar returns					
Chile	0.9414 (0.0230)	0.8688 (0.0892)	12.198 [0.016]	0.59	0.26
Colombia	0.0000 (0.0000)	0.8322 (0.0202)	6.382 [0.172]	0.14	0.14
Greece	0.9868 (0.0011)	0.6244 (0.0155)	28.731 [0.000]	0.89	0.86
India	0.9962 (0.0054)	0.9941 (0.0081)	6.332 [0.176]	0.54	0.10
Jordan	0.9022 (0.0113)	0.1710 (0.0099)	14.570 [0.006]	0.85	0.79
Korea[a]	0.9573 (0.0028)	0.0000 (0.0000)	11.462 [0.022]	0.97	0.99
Malaysia[a]	0.8185 (0.0316)	0.3214 (0.0286)	4.123 [0.390]	0.79	0.79
Mexico	0.9363 (0.0427)	0.9839 (0.0113)	6.962 [0.155]	0.21	0.04
Nigeria	0.7402 (0.1486)	0.8941 (0.0847)	1.115 [0.892]	0.27	0.20
Taiwan	0.9309 (0.0090)	0.3086 (0.0312)	[b] -	0.89	0.90
Thailand[c]	0.9062 (0.0078)	0.9804 (0.0018)	14.369 [0.006]	0.77	1.00
Zimbabwe	0.9808 (0.0098)	0.9699 (0.0146)	10.387 [0.034]	0.47	0.52

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.

[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.

Table 5

Model diagnostics: Correlation of the country asset pricing errors with information

Country	World information Z			Local information Z'			World plus local Z=[Z, Z']		
	$\bar{R}^2$	W	LM	$\bar{R}^2$	W	LM	$\bar{R}^2$	W	LM

International Finance Corporation-Emerging
U.S. dollar returns

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]
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]

The asterisk indicates that the model is estimated with time-varying rather than constant transition probabilities. 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: Z global information, Z' local information or 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 Z, Z' or Z if the error distribution is correctly specified.

a The moment matrix of the scores was singular.

Table 6

The interaction between integration and foreign exchange regimes

Two regressions are estimated for each country:

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

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_2 (coefficient on the estimated integration).

Country	$\Delta \overline{R}^2$	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]
Malaysia	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]

The χ^2 test has one degree of freedom.

Table 7

Estimation results for the constrained alternatives

Country	Constant price of risk LR1	λ^S	Constant variance LR2	Constant degree of integration LR3
Chile	28.95 [0.000]	6.82[b] (0.23)	6.93 [0.031]	6.49 [0.011]
Colombia	23.28 [0.000]	4.58 (2.28)	4.83 [0.090]	0.533 [0.466]
Greece	49.05 [0.000]	1.60 (0.09)	52.34 [0.000]	8.99 [0.003]
India	11.30 [0.023]	9.92 (4.02)	14.78 [0.001]	8.89 [0.003]
Jordan	17.75 [0.001]	43.59 (14.15)	8.90 [0.012]	0.34 [0.560]
Korea	27.34 [0.000]	24.79 (1.70)	27.21 [0.000]	1.17 [0.281]
Malaysia	7.10 [0.131]	10,312.4[d] (553.8)	5.62 [0.060]	0.77 [0.380]
Mexico	12.73 [0.013]	3.01 (1.15)	10.28 [0.036]	18.54 [0.000]
Nigeria	6.22 [0.183]	10.87[c] (3.84)	[a] -	9.93 [0.002]
Taiwan	23.92 [0.000]	0.00002 (0.00004)	18.21 [0.000]	4.08 [0.043]
Thailand	4.30 [0.367]	1.97 (1.45)	[a] -	21.32 [0.000]
Zimbabwe	4.51 [0.342]	5.87 (2.30)	8.51 [0.014]	10.25 [0.001]

LRi (i=1,2,3) is the likelihood ratio statistic testing the restrictions in the panel heading and is distributed with 4,2, and 1 degree of freedom. λ^S is the local price of variance risk estimated under the restriction of constant prices of risk.

[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.

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

[c]Since the Hessian failed to invert, standard errors were computed from the cross-product of first derivatives.

[d]Estimation was ill-behaved and this optimum has not been confirmed as the global optimum.

Figure 1
Time-variation in the world price of risk

World Price of Risk
Constant Variance Model

NBER peaks and troughs labelled with arrows.

World Price of Risk
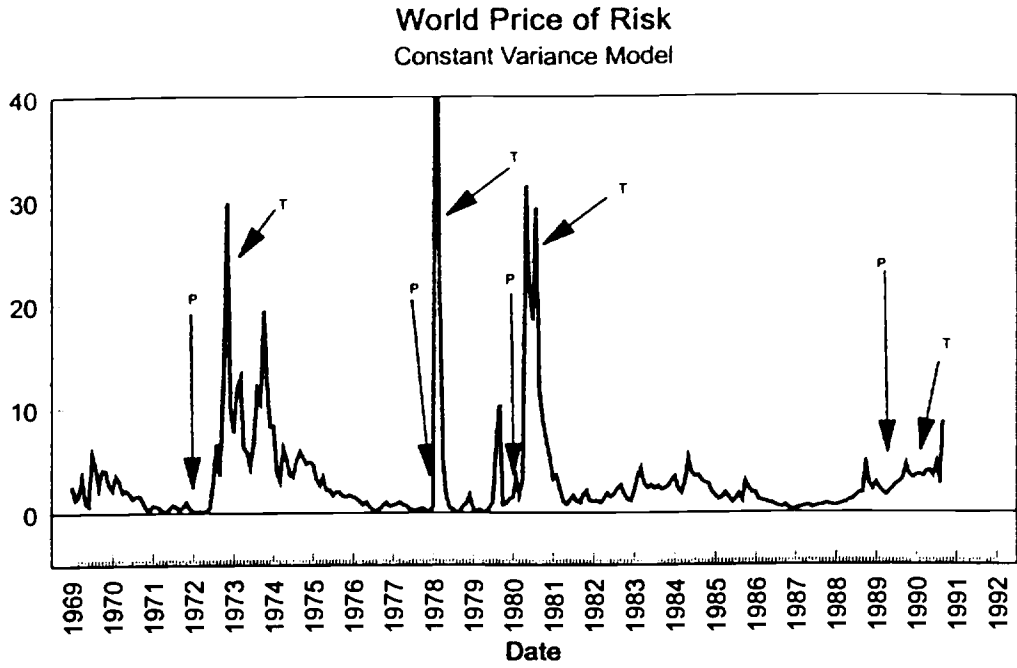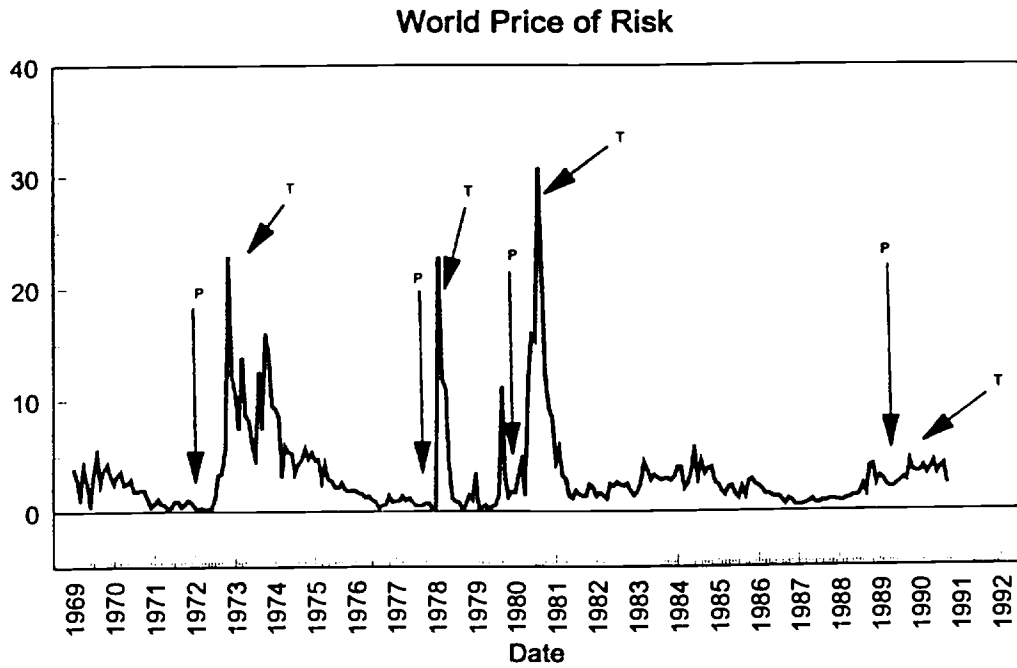
NBER peaks and troughs labelled with arrows.

Figure 2
Time-varying integration measures
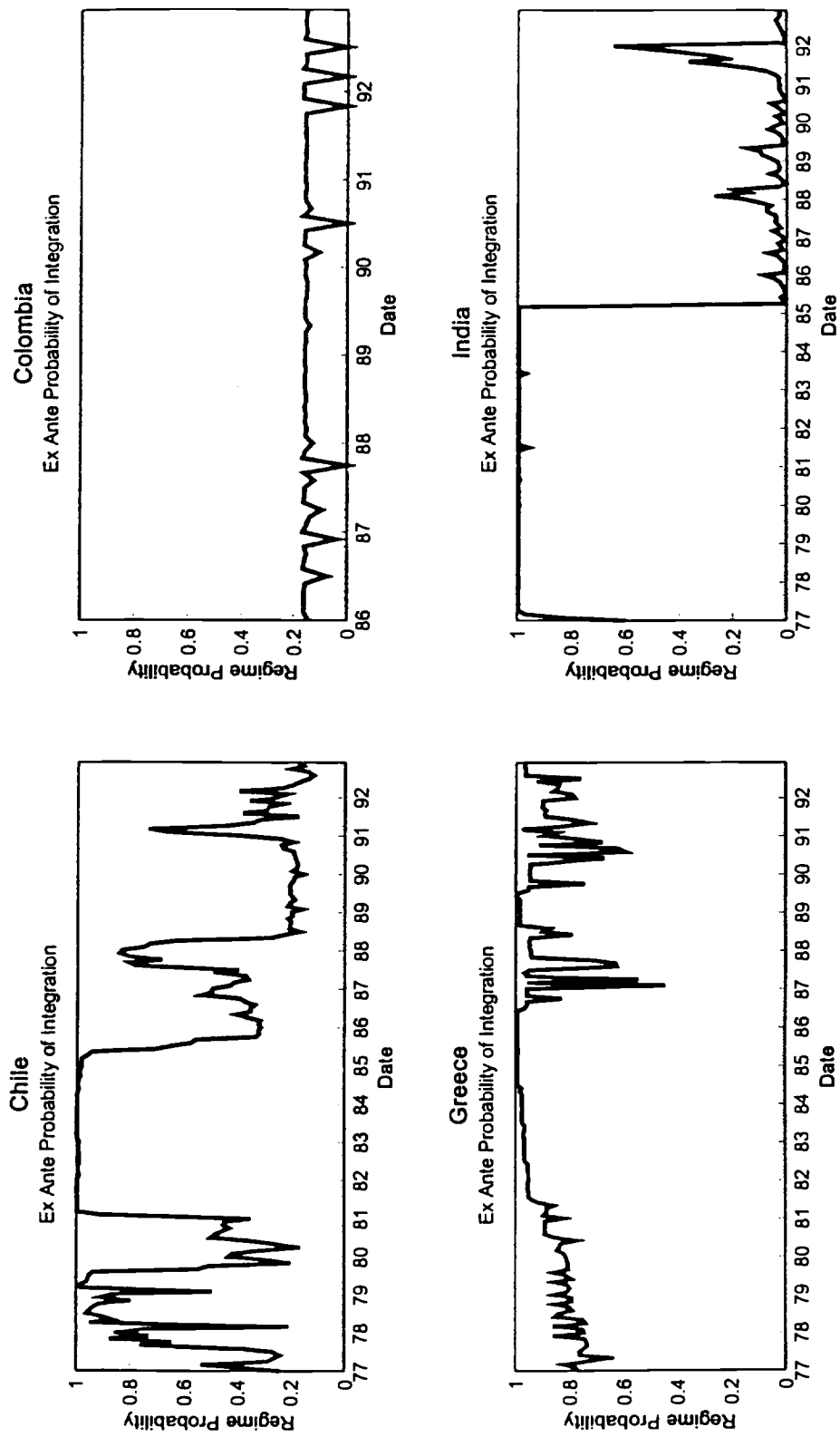
Figure 2 (continued)

Figure 2 (continued)

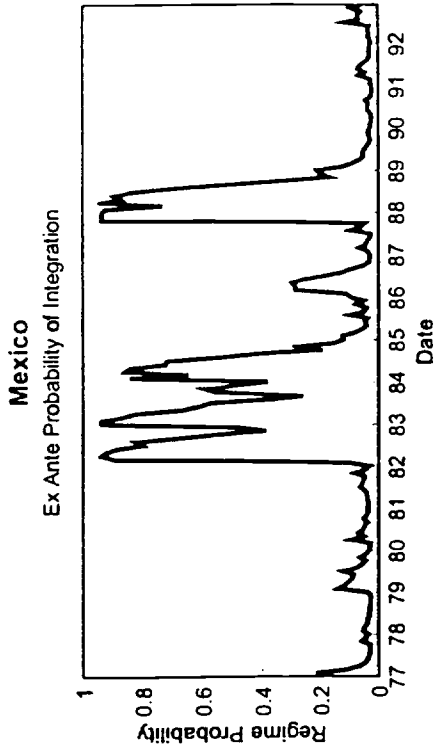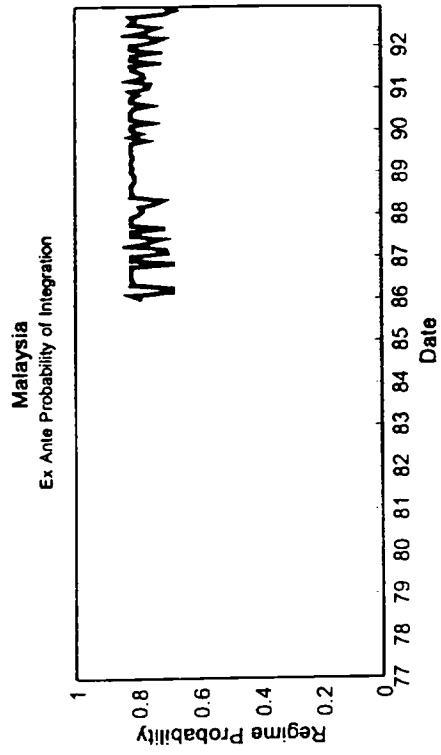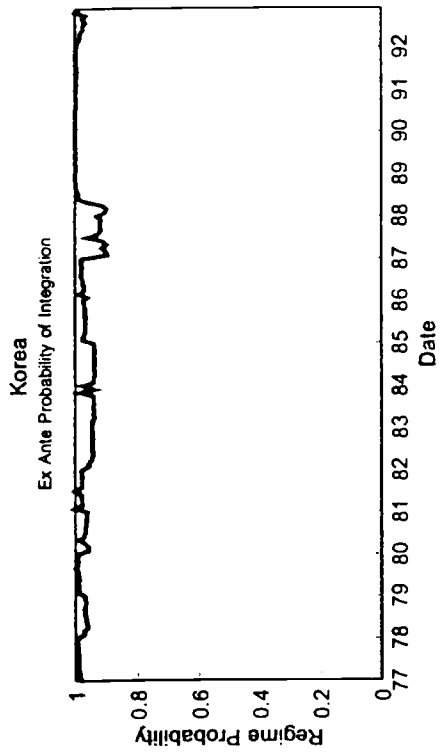Nigeria
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