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Market Integration and Contagion*

Contagion is usually defined as correlation between markets in excess of that implied by economic fundamentals; however, there is considerable disagreement regarding the definition of the fundamentals, how they might differ across countries, and the mechanisms that link them to asset returns. Our research starts with a two-factor model with time-varying betas that accommodates various degrees of market integration. We apply this model to stock returns in three different regions: Europe, Southeast Asia, and Latin America. In addition to examining contagion during crisis periods, we document historic patterns and measure the volatility driven by global, regional, and local factors.

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I. Introduction

Contagion in equity markets refers to the notion that markets move more closely together during periods of crisis. One of the most interesting aspects of the contagion debate is the disagreement over a precise definition. Forbes and Rigobon (2001) declare that "there is no consensus on exactly what constitutes contagion or how it should be defined." Rigobon (2001) states "paradoxically, . . .there is no accordance on what contagion means."

What is clear is that contagion is not simply revealed by increased correlation of market returns during a crisis period. From a completely statistical perspective, one would expect higher

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correlations during periods of high volatility.[1] Forbes and Rigobon (2001) present a statistical correction for this conditioning bias and Q3 argue that there was no contagion during the three most recent crises.

We define contagion as excess correlation, that is, correlation over and above what one would expect from economic fundamentals. Unfortunately, there is disagreement on the definitions of the fundamentals, the potential country-specific nature of the fundamentals, and the mechanism that links the fundamentals to asset correlation.

Our paper takes an asset pricing perspective to the study of contagion. For a given factor model, increased correlation is expected if the volatility of a factor increases. The size of the increased correlation depends on the factor loadings. Contagion is simply defined by the correlation of the model residuals.

By defining the factor model, we avoid a problem with the bias correction for correlations that Forbes and Rigobon (2001) propose, a Q4 bias correction that does not work in the presence of common shocks. Defining the factor model does mean that we effectively take a stand on the global, regional, and country-specific fundamentals, as well as the mechanism that transfers fundamentals into correlation. Of course, any statements on contagion are contingent on the correct specification of the factor model; therefore, we start with a model that has the maximum flexibility.

We apply a two-factor model with time-varying loadings to “small” stock markets in three regions: Europe, Southeast Asia, and Latin America. The two factors are the U.S. equity market return and a regional equity portfolio return. Our framework nests three models: a world capital asset pricing model (CAPM), a CAPM with the U.S. equity return as the benchmark asset, and a regional CAPM with a regional portfolio as the benchmark. We test the asset pricing specifications by adding local factors.

Segmentation and integration play a critical role in our tests. If the countries in a particular region are globally integrated for most of the sample period but suddenly see their intraregional correlations rise dramatically during a regional crisis, our test rejects the null hypothesis of no contagion. If, however, these countries do not follow a global CAPM but rather a regional CAPM, the increased correlations may simply be a consequence of increased factor volatility.

Our volatility model is related to Bekaert and Harvey (1997) and Ng (2000) in that equity return volatilities follow univariate generalized

1. See Stambaugh (1995); Boyer, Gibson, and Loretan (1999); Loretan and English (2000); Forbes and Rigobon (2001); and early work by Pindyck and Rotemberg (1990, 1993). Work linking news, volatility, and correlation includes King and Wadhawani (1990); Hamao, Masulis, and Ng (1990); and King, Sentana, and Wadhwani (1994).

autoregressive conditional heteroscedasticity (GARCH) processes with asymmetry. Hence, negative news regarding the world or regional market may increase volatility of the factor more than positive news and lead to increased correlations between stock markets.[2] Moreover, our model incorporates time-varying betas, where the betas are influenced by the trade patterns. Chen and Zhang (1997) find that the cross-market correlations of stock returns are related to external trade among countries.

Previous studies on international market linkages focused mainly on one source of risk or on the effects of a single international market (often the U.S. or world market) on other stock markets.[3] In fact, a contemporaneous paper (Tang 2001) uses a definition of contagion similar to the one we propose but restricts the model to a world CAPM. Our structure, which allows world market integration to be a special case, bears some resemblance to the setup in Chan, Karolyi, and Stulz (1992). However, the existing literature has focused primarily on world market integration, and regional integration has been scarcely discussed. Exceptions include Engle and Susmel (1993), who group the data according to time zones and search for common regional news factors, and Cheung, He, and Ng (1997), who examine common predictable components in returns within a region.

Our main contribution is to examine periods of crises and investigate whether our model can generate sudden increases in correlations across countries. Our approach, however, produces many other useful empirical tests and implications. Indeed, our framework provides a natural test for world and regional market integration. In addition, we analyze the time variation and cross-sectional patterns in regional versus world market correlations, addressing the question of whether correlations between country returns and the world or regional market have increased over time. Finally, we measure the proportions of variance driven by global, regional, and local factors and how these proportions change through time.

The remainder of the paper is organized as follows. Section II presents the empirical model specifications and several testable hypotheses. Section III describes the data and presents the empirical results. Some conclusions are offered in the final section.

2. Longin and Solnik (1995) report an increase in cross-country correlation during volatile periods. Other empirical studies (for example, Erb, Harvey, and Viskanta 1994 and De Santis and Gerard 1997) find different correlations in up and down markets, while Longin and Solnik (2001), Ang and Bekaert (2002), and Das and Uppal (2001) document higher correlations in bear markets. Q13

3. See, for example, Hamao, Masulis, and Ng (1990); Bekaert and Hodrick (1992); Bekaert and Harvey (1995, 1997); Karolyi (1995); Karolyi and Stulz (1996); Hartmann, Straetmans, and de Vries (2001); and Connolly and Wang (2002).

II. Framework

The international version of the conditional CAPM of Sharpe (1964) and Lintner (1965) under the assumption of purchasing power parity (PPP) predicts the excess return on a world market portfolio, with the factor of proportionality being the country-specific conditional beta. Rather than focusing on currency risk (see Ferson and Harvey 1993 and Dumas and Solnik 1995), we extend the traditional CAPM from a one-factor to a two-factor setting, by dividing the world market into the United States and a particular region and allow for local factors to be priced. Q5

A. The Model

Let $R_{i,t}$ be the excess return on the national equity index of country i in U.S. dollars. The model has the following form:

$$\boldsymbol{R}_{i,t} = \delta_i^{'}\mathbf{Z}_{i,t-1} + \beta_{i,t-1}^{\mathrm{us}}\mu_{\mathrm{us},t-1} + \beta_{i,t-1}^{\mathrm{reg}}\mu_{\mathrm{reg},t-1} + \beta_{i,t-1}^{\mathrm{us}}e_{\mathrm{us},t} + \beta_{i,t-1}^{\mathrm{reg}}e_{\mathrm{reg},t} + e_{i,t}, \tag{1}$$

$$\boldsymbol{e}_{i,t} \mid \boldsymbol{I}_{t-1} \sim \boldsymbol{N}\left(0, \sigma_{i,t}^2\right), \tag{2}$$

$$\sigma_{i,t}^2 = a_i + b_i\sigma_{i,t-1}^2 + c_ie_{i,t-1}^2 + d_i\eta_{i,t-1}^2, \tag{3}$$

where $\mu_{\mathrm{us},t-1}$ and $\mu_{\mathrm{reg},t-1}$ are the conditional expected excess returns on the U.S. and regional markets, respectively, based on information available at time $t-1$, $e_{i,t}$ is the idiosyncratic shock of any market i, including the U.S. and regional portfolio, $\eta_{i,t}$ is the negative return shock of country i, that is, $\eta_{i,t} = \min\left\{0, e_{i,t}\right\}$, and I_{t-1} includes all the information available at time $t-1$. The vector $Z_{i,t-1}$ contains a constant and the local dividend yield, which help estimate the expected return of market i. The dividend yields are lagged by 1 month. The variance of the idiosyncratic return shock of market i follows a GARCH process in eq. (3) with asymmetric effects in conditional variance, as in Glosten, Jagannathan, and Runkle (1993) and Zakoian (1994).

The sensitivity of equity market i to the foreign news factors is measured by the parameters $\beta_{i,t-1}^{\mathrm{us}}$ and $\beta_{i,t-1}^{\mathrm{reg}}$. Following Bekaert and Harvey (1995, 1997) and Ng (2000), we model these risk parameters to be time-varying as

$$\beta_{i,t-1}^{\mathrm{us}} = \mathrm{p}_{1,i}^{'}\mathbf{X}_{i,t-1}^{\mathrm{us}} + \mathrm{q}_i^{'}\mathbf{X}_{i,t-1}^{w} \cdot w_{us,t-1}, \tag{4}$$

$$\beta_{i,t-1}^{\mathrm{reg}} = \mathrm{p}_{2,i}^{'}\mathbf{X}_{i,t-1}^{\mathrm{reg}} + \mathrm{q}_i^{'}\mathbf{X}_{i,t-1}^{w} \cdot \left(1 - w_{us,t-1}\right), \tag{5}$$

where $w_{us,t-1}$ denotes the market capitalization of the United States, relative to the total world market capitalization, at time $t-1$. In eqq. (4) and (5), we further introduce three different sets of local instruments, $\boldsymbol{X}^{\text{us}}_{i,t-1}$, $\boldsymbol{X}^{\text{reg}}_{i,t-1}$, and $\boldsymbol{X}^{w}_{i,t-1}$. The set $\boldsymbol{X}^{\text{us}}_{i,t-1}$ $\left(\boldsymbol{X}^{\text{reg}}_{i,t-1}\right)$ consists of information variables that capture the covariance risk of market i with the United States (the region). We use a constant and the sum of exports to and imports from the United States (the rest of the world) divided by the sum of total exports and total imports. We try to capture within the region trade by all trade minus the United States. While this is only a proxy, it imposes a clean relation between $\mathbf{X}^{\text{us}}_{i,t-1}$, $\mathbf{X}^{\text{reg}}_{i,t-1}$, and $\boldsymbol{X}^{w}_{i,t-1}$. Note that our structure allows the conditional betas to be affected by trade. Chen and Zhang (1997) study the relation between cross-market return correlation and bilateral trade and find that countries with heavier external trade to a region tend to have higher return correlations with that region. Similarly, the information set $\boldsymbol{X}^{w}_{i,t-1}$ consists of local instruments that should capture the covariance risk of market i with a world portfolio. Here, we include a constant and the country's total size of trade as a percentage of GDP. All the trade variables are lagged by 6 months.[4]

The U.S. and regional markets models are special cases of eqq. (1)–(5). For the U.S. market (with i = us), $p_{1,\text{us}} = p_{2,\text{us}} = q_{\text{us}} = 0$ (that is, $\beta^{\text{us}}_{\text{us},t-1} = \beta^{\text{reg}}_{\text{us},t-1} = 0$) and $\mathbf{Z}_{\text{us,t}-1}$ contains a set of world information variables, including a constant, the world market dividend yield, the spread between the 90-day Eurodollar rate and the 3-month Treasury-bill yield, the difference between the U.S. 10-year Treasury bond yield and the 3-month bill yield, and the change in the 90-day Treasury-bill yield. For the regional portfolio (with i = reg), $p_{2,\text{reg}} = q_{2,\text{reg}} = 0$ (that is, $\beta^{\text{us}}_{\text{us},t-1} = \mathbf{p}'_{1,\text{reg}}\mathbf{X}^{\text{us}}_{\text{reg},t-1}$ and $\beta^{\text{reg}}_{\text{reg},t-1} = 0$), $\mathbf{Z}_{\text{reg},t-1}$ includes a constant and the regional market dividend yield (weighted by market capitalization), and $\mathbf{X}^{\text{us}}_{\text{reg,t}-1}$ contains a constant and the sum of the region's total exports to and imports from the United States divided by the sum of total exports and imports of the region.

As shown in eq. (1), the expected excess return on market i is a linear function of some local information variables and the expected excess returns on the U.S. and regional markets: that is,

$$\begin{aligned}\mu_{i,t-1} &= \boldsymbol{E}\left[\boldsymbol{R}_{i,t-1} \mid \mathbf{I}_{t-1}\right] = \boldsymbol{\delta}'_i\mathbf{Z}_{i,t-1} + \beta^{\text{us}}_{i,t-1}\mu_{\text{us},t-1} + \beta^{\text{reg}}_{i,t-1}\mu_{\text{reg},t-1} \\ &= \boldsymbol{\delta}'_i\mathbf{Z}_{i,t-1} + \left[\beta^{\text{us}}_{i,t-1} + \beta^{\text{reg}}_{i,t-1}\beta^{\text{us}}_{\text{reg},t-1}\right]\left(\boldsymbol{\delta}'_{\text{us}}\mathbf{Z}_{us,t-1}\right) + \beta^{\text{reg}}_{i,t-1}\left(\boldsymbol{\delta}'_{\text{reg}}\mathbf{Z}_{\text{reg},t-1}\right).\end{aligned} \tag{6}$$

Hence, the effect of world market information originating from the United States on market i's expected return has two components: a

4. The appendix provides a detailed discussion of the construction of the information variables.

direct impact, as measured by $\beta^{us}_{i,t-1}$, and an indirect effect via its influence on the regional market, as measured by $\beta^{reg}_{i,t-1}\beta^{us}_{reg,t-1}$.

Similarly, the unexpected portion of the market return is driven not only by shocks from the local market but also by two foreign shocks originating in the United States and the region: that is,

$$\varepsilon_{i,t} = \beta^{us}_{i,t-1} e_{us,t} + \beta^{reg}_{i,t-1} e_{reg,t} + e_{i,t}, \tag{7}$$

where $\varepsilon_{i,t}$ denotes the return residual of market i. To complete the model, we further assume that the idiosyncratic shocks of the United States, regional market, and country i are uncorrelated. As a result, the model implies the following variance and covariance expressions:

$$h_{i,t} = \boldsymbol{E}\left[\varepsilon^2_{i,t} \mid \mathbf{I}_{t-1}\right] = \left(\beta^{us}_{i,t-1}\right)^2 \sigma^2_{us,t} + \left(\beta^{reg}_{i,t-1}\right)^2 \sigma^2_{reg,t} + \sigma^2_{i,t}, \tag{8}$$

$$h_{i,us,t} = \boldsymbol{E}\left[\varepsilon_{i,t}\varepsilon_{us,t} \mid \mathbf{I}_{t-1}\right] = \beta^{us}_{i,t-1}\sigma^2_{us,t}, \tag{9}$$

$$h_{i,reg,t} = E\left[\varepsilon_{i,t}\varepsilon_{reg,t} \mid \mathbf{I}_{t-1}\right] = \beta^{us}_{i,t-1}\beta^{us}_{reg,t-1}\sigma^2_{us,t} + \beta^{reg}_{i,t-1}\sigma^2_{reg,t}, \tag{10}$$

$$h_{i,j,t} = \boldsymbol{E}\left[\varepsilon_{i,t}\varepsilon_{j,t} \mid \mathbf{I}_{t-1}\right] = \beta^{us}_{i,t-1}\beta^{us}_{j,t-1}\sigma^2_{us,t} + \beta^{reg}_{i,t-1}\beta^{reg}_{j,t-1}\sigma^2_{reg,t}. \tag{11}$$

Equation (8) shows that the return volatility of market i is positively related to the conditional variances of the U.S. and regional markets. Consequently, we can investigate whether potential asymmetric effects in the U.S. or regional markets induce asymmetry in the conditional return volatility of any equity market.

The conditional covariance dynamics given in eqq. (9)–(11) have several important implications. First, a market's covariance with the U.S. (regional) market return is positively related to its country-specific beta with the United States, $\beta^{us}_{i,t-1}$(region, $\beta^{reg}_{i,t-1}$). Second, provided that the country-specific beta parameter $\beta^{us}_{i,t-1}$ is positive, higher volatility in the U.S. market induces higher return covariance between the United States and market i. Third, the covariance with the regional market or any other national market j within the same region increases in times of high return volatility in the United States or the regional market or both. This natural implication of any factor model, coupled with asymmetric volatility, could lead to the appearance of "contagious bear markets." Whereas these points, which follow immediately from eqq. (9)–(11), apply to covariances, they are also true for

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correlations.[5] Note, that increased trade integration (changes in $\beta^{j}_{i,t-1}$ for j = us,reg) may also increase the correlation with the U.S. or regional market and between countries.

In the empirical section, we study the time variation and cross-sectional patterns in regional versus U.S. market correlations of market i. The U.S. and regional correlations are given by

$$\rho_{i,\mathrm{us},t} = \frac{\beta^{\mathrm{us}}_{i,t-1}\sigma_{\mathrm{us},t}}{\sqrt{h_{i,t}}} \tag{12}$$

$$\rho_{i,\mathrm{reg},t} = \frac{\beta^{\mathrm{us}}_{i,t-1}\beta^{\mathrm{us}}_{\mathrm{reg},t-1}\sigma^{2}_{\mathrm{us},t} + \beta^{\mathrm{reg}}_{i,t-1}\sigma^{2}_{\mathrm{reg},t}}{\sqrt{h_{i,t}h_{\mathrm{reg},t}}} \tag{13}$$

where $h_{\mathrm{reg},t} = \left(\beta^{\mathrm{us}}_{\mathrm{reg},t-1}\right)^{2}\sigma^{2}_{\mathrm{us},t} + \sigma^{2}_{\mathrm{reg},t}$ is the conditional variance of the regional market return. We also examine the (relative) proportions of conditional return variance that are accounted for by the United States and the region. The following variance ratios are computed:

$$\mathrm{VR}^{\mathrm{us}}_{i,t} = \frac{\left(\beta^{\mathrm{us}}_{i,t-1}\right)^{2}\sigma^{2}_{\mathrm{us},t}}{h_{i,t}}, \tag{14}$$

$$\mathrm{VR}^{\mathrm{reg}}_{i,t} = \frac{\left(\beta^{\mathrm{reg}}_{i,t-1}\right)^{2}\sigma^{2}_{\mathrm{reg},t}}{h_{i,t}}, \tag{15}$$

Clearly, these variance ratios increase when the "factor" variance (U.S. or regional market) increases. We are specifically interested in the crisis periods and investigate whether the model can generate sudden increase in correlations across markets following a crisis.

B. Testable Hypotheses Regarding Market Integration

The specification presented in eqq. (1)–(5) is a general two-factor model that allows us to examine several testable hypotheses. First, if the two-factor model holds (that is, if the two foreign risk factors are sufficient in explaining the expected return on market i), the local instruments should have no explanatory power on their own market return and, therefore, $\boldsymbol{\delta}_i = \mathbf{0}$. Consequently, we interpret this test as a test of market integration, where integration can be global or regional.

5. It is straightforward to show this formally, but it is also clear intuitively. If the variance of the common factor goes to zero, then the returns are driven only by idiosyncratic shocks. At the other extreme, if the variance of the common factor goes to infinity, the idiosyncratic shocks become irrelevant and the fluctuations in returns are fully explained by the common shock. Hence, the correlation becomes one.

Second, the model nests the one-factor CAPM as a special case. Under the constraint that $\mathbf{p}_{2,i} = \mathbf{0}(\mathbf{p}_{1,i} = \mathbf{0})$ and $\mathbf{q}_i = \mathbf{0}$, together with $\boldsymbol{\delta}_i = \mathbf{0}$, the model reduces to the traditional CAPM, with the United States (the region) being the benchmark market and $\beta^{us}_{i,t-1}(\beta^{reg}_{i,t-1})$ equal to the conditional beta of market i with the U.S. (regional) market. The model then implies that market i is fully integrated with the U.S. (regional) market. Under this setting, we should be able to detect deviations from the one-factor integrated model. We interpret the test, $\mathbf{p}_{1,i} = \mathbf{q}i = \boldsymbol{\delta}_i = \mathbf{0}$, as a test of regional market integration.

Third, our model encompasses a world market integration model. Suppose that the world market is separated into the United States and a particular region and that each market i is fully integrated with the world capital market. This happens when $\mathbf{p}_{1,i} = \mathbf{0}$ and $\mathbf{p}_{2,i} = \mathbf{0}$; that is, $\beta^{us}_{i,t-1} = \beta^{w}_{i,t-1} \cdot w_{us,t-1}$ and $\beta^{reg}_{i,t-1} = \beta^{w}_{i,t-1} \cdot (1 - w_{us,t-1})$. Hence, $\beta^{w}_{i,t-1} = \mathbf{q}_i' \mathbf{X}^{w}_{i,t-1}$ is the conditional beta of market i with the world market portfolio in the traditional CAPM. Hence, our framework encompasses the world market integration model presented in Chan, Karolyi and Stulz (1992).

C. Contagion Definitions and Tests

We measure contagion by measuring the correlation of the model's idiosyncratic shocks or unexpected returns. We establish a baseline level of contagion by examining shock correlations estimated over the full sample period; however, we are most interested in the shock correlations during particular periods. Our tests involve the time-series cross-section regression model:

$$\hat{e}_{i,t} = w_i + v_{i,t}\hat{e}_{g,t} + u_{i,t}$$

$$v_{i,t} = v_0 + v_1 \boldsymbol{D}_{i,t}$$

where $\hat{e}_{i,t}$ and $\hat{e}_{g,t}$ are the estimated idiosyncratic return shocks of market i and region g, respectively. For the regional residuals, three cases are considered: $\hat{e}_{g,t} = \hat{e}_{us,t}$, $\hat{e}_{g,t} = \hat{e}_{reg,t}$, and $\hat{e}_{g,t} = \sum_{j \in G}^{j \neq i} \hat{e}_{j,t}$ where G denotes a particular country group. In studying the market residuals, countries are categorized into the following country groups: Europe, Europe excluding Turkey, Asia, and Latin America. $D_{i,t}$ is a dummy variable that represents five sample periods: the second subsample period, the Mexico crisis period from November 1994 to December 1995, the Asian crisis period from April 1997 to October 1998, abnormally negative U.S. unexpected market returns (i.e., the unexpected returns are one standard deviation below zero), and abnormally negative regional unexpected market returns. Our tests determine whether v_0 and v_1 are zero (overall contagion), and whether v_1 is significantly different from zero (contribution of particular periods to contagion).

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D. Estimation Method

The model presented in the first section can be expressed in a multivariate setting. Let $\mathbf{R}_t = \left[R_{\mathrm{us},t}, R_{\mathrm{reg},t}, R_{1,t}, \ldots, R_{N,t}\right]'$, $\boldsymbol{\mu}^*_{t-1} = \left[\boldsymbol{\delta}'_{\mathrm{us}}\mathbf{Z}_{\mathrm{us},t-1}, \boldsymbol{\delta}'_{\mathrm{reg}}\mathbf{Z}_{\mathrm{reg},t-1}, \boldsymbol{\delta}'_1\mathbf{Z}_{1,t-1}, \ldots, \boldsymbol{\delta}'_N\mathbf{Z}_{N,t-1}\right]'$ and $\mathbf{e}_t = \left[e_{\mathrm{us},t}, e_{\mathrm{reg},t}, e_{1,t}, \ldots, e_{N,t}\right]'$, where N is the number of countries within the particular region. The general $(N+2)$ multivariate model has the following form:

$$\begin{aligned}\mathbf{R}_t &= \boldsymbol{\Phi}_{t-1}\boldsymbol{\mu}^*_{t-1} + \boldsymbol{\beta}_{t-1}\mathbf{e}_t \\ \mathbf{e}_t \mid \mathbf{I}_{t-1} &\sim N\left(0, \boldsymbol{\Sigma}_t\right)\end{aligned} \tag{16}$$

$$\boldsymbol{\Sigma}_t = E\left[\mathbf{e}_t\mathbf{e}'_t \mid \mathbf{I}_{t-1}\right] = \mathrm{diag}\left\{\boldsymbol{\sigma}^2_{j,t}\right\}, \quad \text{where } \boldsymbol{j} = \mathrm{us}, \mathrm{reg}, 1, \ldots, N,$$

$$\text{with} \quad \boldsymbol{\Phi}_{t-1} = \begin{bmatrix} 1 & 0 & 0 & \ldots & 0 \\ \beta^{\mathrm{us}}_{\mathrm{reg},t-1} & 1 & 0 & \ldots & 0 \\ \phi_{1,t-1} & \beta^{\mathrm{reg}}_{1,t-1} & & & \\ \vdots & \vdots & \mathbf{I}_{(N)} & & \\ \phi_{N,t-1} & \beta^{\mathrm{reg}}_{N,t-1} & & & \end{bmatrix}$$

$$\text{and} \quad \boldsymbol{\beta}_{t-1} = \begin{bmatrix} 1 & 0 & 0 & \ldots & 0 \\ \beta^{\mathrm{us}}_{\mathrm{reg},t-1} & 1 & 0 & \ldots & 0 \\ \beta^{\mathrm{us}}_{1,t-1} & \beta^{\mathrm{reg}}_{1,t-1} & & & \\ \vdots & \vdots & \mathbf{I}_{(N)} & & \\ \beta^{\mathrm{us}}_{N,t-1} & \beta^{\mathrm{reg}}_{N,t-1} & & & \end{bmatrix}, \text{where}$$

$_{i,t-1} = \beta^{\mathrm{us}}_{i,t-1} + \beta^{\mathrm{reg}}_{i,t-1}\beta^{\mathrm{us}}_{\mathrm{reg},t-1}$ and $\mathbf{I}_{(N)}$ is a $(N \times N)$ identity matrix.

We estimate the joint multivariate likelihood function for the returns in three stages. Given that the density of the U.S. return, conditional on $\boldsymbol{I}_{t-1}$, depends only on $\boldsymbol{\theta}_{\mathrm{us}} = \left[\boldsymbol{\delta}'_{\mathrm{us}}, a_{\mathrm{us}}, b_{\mathrm{us}}, c_{\mathrm{us}}, d_{\mathrm{us}}\right]'$, in the first stage, we estimate the (univariate) model in eqq. (1)–(5) for the U.S. market. In the second stage, based on the U.S. estimates from stage 1, we examine the model for the regional market portfolio. Conditional on $\boldsymbol{I}_{t-1}$ and $\boldsymbol{R}_{\mathrm{us},t}$, the density function of the regional market return depends only on $[\boldsymbol{\theta}'_{\mathrm{us}}, \boldsymbol{\theta}'_{\mathrm{reg}}]'$, where $\boldsymbol{\theta}'_{\mathrm{reg}} = \left[\mathbf{D}'_{\mathrm{reg}}, \mathbf{p}'_{1,\mathrm{reg}}, a_{\mathrm{reg}}, b_{\mathrm{reg}}, c_{\mathrm{reg}}, d_{\mathrm{reg}}\right]'$ and, thus, consistent estimates of $\boldsymbol{\theta}_{\mathrm{reg}}$ are obtained by maximizing the univariate likelihood for the regional market return. Finally, in the third stage, we estimate the univariate model in eqq. (1)–(5) country by country, conditioning on the U.S. and regional market model estimates. This methodology is similar to the one proposed by Bekaert and Harvey (1997).

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E. Model Selection and Specification Tests

Because asymmetric and symmetric GARCH models produce very different conditional variances, our model selection focuses on that issue. Moreover, it is conceivable that asymmetric effects at the U.S. or regional market level make the country-specific asymmetry superfluous. We first carry out a likelihood ratio (LR) test for the null hypothesis of no asymmetry in the conditional variance of the local return residual. If the LR test rejects the null hypothesis at the 5% level, the model with asymmetry is chosen; if the *p*-value of the test statistic is greater than 15%, the model without asymmetry is selected. For the intermediate case, we regress the squared return residuals for both models on the estimated conditional variances, as in Pagan and Schwert (1990), and select the model with the higher R^2.

We conduct specification tests on the estimated standardized idiosyncratic shocks, $\hat{z}_{i,t} = \hat{e}_{i,t}/\hat{\sigma}_{i,t}$ for all i, using the generalized method of moments. Under the null hypothesis that the model is correctly specified,

$$E\left[\hat{z}_{i,t}\right] = 0, \tag{17a}$$

$$E\left[\hat{z}_{i,t}\hat{z}_{i,t-s}\right] = 0, \text{ for } s = 1, \ldots, \tau, \tag{17b}$$

$$E\left[\hat{z}_{i,t}^2 - 1\right] = 0, \tag{17c}$$

$$E\left[\left(\hat{z}_{i,t}^2 - 1\right)\left(\hat{z}_{i,t-s}^2 - 1\right)\right] = 0, \text{ for } s = 1, \ldots, \tau, \tag{17d}$$

$$E\left[\hat{z}_{i,t}^3\right] = 0, \tag{17e}$$

$$E\left[\hat{z}_{i,t}^4 - 3\right] = 0. \tag{17f}$$

Equations (17b) and (17d) are a consequence of the correct specification for the conditional mean and variance, and these two constraints are tested separately by a χ^2 test with τ degrees of freedom. The unconditional moments in the other four constraints are tested jointly by calculating a χ^2 statistic with four degrees of freedom. We also carry out a joint test of all six restrictions, which has $2\tau + 4$ degrees of freedom. In all of the specification tests, τ is set equal to 4. The Monte Carlo analysis in Bekaert and Harvey (1997), in a similar setting, confirms that the small sample distribution of the test statistics is relatively well described by χ^2 distributions, despite the multistage estimation.

III. Results

A. Equity Market Data

Our sample of national equity markets includes data for both developed markets, as compiled by Morgan Stanley Capital International (MSCI), and emerging markets from the International Finance Corporation (IFC) of the World Bank. The sample period begins in January 1980 for most of the MSCI data and January 1986 for the IFC data. The sample ends in December 1998. We study a total of 22 countries grouped into three geographical regions: Asia, Europe, and Latin America. The regional equity indices we examine are the MSCI Europe index, as well as our own Asia and Latin America emerging market indices. The Asia (Latin America) emerging market index is a weighted average of all the Asian (Latin American) emerging markets, excluding the country under investigation. Hence, we compute the Asia or Latin America emerging market index, $R_{\text{reg}/i,t}$, as

$$R_{\text{reg}/i,t} = \sum_{k \neq i} w_{k,t} R_{k,t} \Big/ \sum_{k \neq i} w_{k,t},$$

with k indexing the Asian or Latin American markets, except market i and w_k, denoting the market capitalization of market k. Country returns are not always more highly correlated with a regional index than with either the U.S. or MSCI world index returns. For example, all European countries are more highly correlated with the MSCI World and Europe indices than with the United States, but Indonesia, Korea, and the Philippines have higher unconditional correlations with the United States than with their regional index. Within the Latin American group, Brazil, Chile, Colombia, and Mexico have higher correlations with the United States than with their regional index. Detailed summary statistics regarding correlations are available on request.

As section II.A indicated, we use a substantial number of both economic and financial information variables, which are detailed in the appendix.

B. U.S. and Regional Models

Table 1 details the U.S. and regional market model estimation. The first row (Wald test I) shows that, consistent with previous research, there is significant variation in the conditional mean for the U.S. returns. Our results strongly reject the hypothesis of no asymmetry in the conditional variances. All three specification tests fail to reject the U.S. model specification. The joint test fails to reject the specification at the 5% level but provides some evidence against the specification at the 10% level.[6]

6. When we do not explicitly mention the test level, we use 5% tests to judge significance.

TABLE 1 **The U.S. and Regional Market Return Model**

	Model	Specification Tests				Wald Tests (p-value)			Implied Statistics					
									$\hat{\beta}^{us}_{i,t-1}$		$\hat{\rho}_{i,us,t}$		$v\hat{R}^{us}_{i,t}$	
Market	(Asy/Sym)	Mean	Moment	Variance	Joint	I	II	III	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
United States	Asym	1.555 [.817]	4.962 [.291]	4.808 [.308]	20.90 [.052]	<.001								
Europe	Sym	6.774 [.148]	4.705 [.319]	2.738 [.603]	15.96 [.193]	.390	.856	<.001	.657	.016	.587	.102	.355	.133
Asia	Sym	3.593 [.464]	11.12 [.025]	1.718 [.788]	20.15 [.064]	<.001	.174	.114	.296	.263	.146	.138	.040	.058
Latin America	Sym	2.935 [.569]	3.027 [.553]	2.134 [.711]	13.46 [.336]	.365	.603	<.001	.977	.112	.432	.114	.199	.120

NOTE.—The following model is estimated:

$$R_{i,t} = \delta_i' \mathbf{Z}_{i,t-1} + \beta^{us}_{i,t-1}\hat{\mu}_{us,t-1} + \beta^{us}_{i,t-1}\hat{e}_{us,t} + e_{i,t}$$

$$\beta^{us}_{i,t-1} = p'_{1,i}\mathbf{X}^{us}_{i,t-1}$$

$$e_{i,t} \mid I_{t-1} \sim N\left(0, \sigma^2_{i,t}\right)$$

$$\sigma^2_{i,t} = a_i + b_i\sigma^2_{ri,t-1} + c_i e^2_{i,t-1} + d_i\eta^2_{i,t-1}$$

$$\eta_{i,t-1} = \min\left\{0, e_{i,t}\right\}$$

where $\hat{\mu}_{us,t-1}$ and $\hat{e}_{us,t}$ are the conditional expected excess return and residual of the U.S. market. For the U.S. market (i.e., i = us), $\mathbf{p}_{1,us} = \mathbf{0}$, and $\mathbf{Z}_{us,t-1}$ represents a set of U.S. or world information variables, which includes a constant, the world market dividend yield, the spread between the 90-day Eurodollar rate and the 3-month Treasury-bill yield, the difference between the U.S. 10-year Treasury bond yield and the 3-month Treasury-bill yield, and the change in the 90-day Treasury-bill yield. All these U.S. information variables are lagged by 1 month. For the regional market (i.e., i = reg), $\mathbf{Z}_{reg,t-1}$ represents a set of regional information variables, which includes a constant and the regional market dividend yield weighted by the market capitalization, and $\mathbf{X}^{us}_{reg,t-1}$ a constant, and the sum of the region's total exports to and imports from the United States divided by the sum of total exports and imports of the region. The dividend yield is lagged by 1 month and the trade variable is lagged by 6 months.

All monthly returns are calculated in excess of the U.S. 1-month Treasury-bill rate and in U.S. dollars. The sample covers the period from January 1980 to December 1998 for the United States and Europe, while the data for Asia and Latin America start from January 1986. Return data for the United States and Europe are from Morgan Stanley Capital International (MSCI), whereas Asia and Latin America data are from the International Finance Corporation (IFC). The Asia or Latin America emerging market index is a value-weighted average of all the Asian or Latin American emerging markets in the sample.

To test for model specification, the mean test is based on the first four autocovariances of the scaled residuals, eq. (17b); the variance test is based on the first four autocovariances of the squared scaled residuals, eq. (17d); the moment test is based on four moments, eqq. (17a, 17c, 17e, 17f); and the joint test is based on all the restrictions in 17. Three hypotheses are tested. Wald I is a test of the significance of the regional information in the mean, i.e., $\delta_i = \mathbf{0}$; Wald II is a test of the significance of the trade variable in $\beta^{us}_{i,t-1}$; and Wald III is a test of the significance of the regional variables on $\beta^{us}_{i,t-1}$, i.e., $\mathbf{p}_{1,i} = \mathbf{0}$. Sample means and standard deviations of the implied $\hat{\beta}^{us}_{i,t-1}$, the conditional correlation between the U.S. and regional market ($\hat{\rho}_{i,us,t}$), and the variance ratio of conditional variance of the regional portfolio accounted for by the U.S. factors ($v\hat{R}^{us}_{i,t}$) are reported. p-values are given in brackets.

While we are constrained by data beginning in the 1980s for Asian and Latin American emerging markets, the U.S. and European data are available earlier. We have independently conducted alternative estimations using, in particular, the U.S. market model estimated over a longer sample. These results are available on request. Over the longer sample, the U.S. market exhibits the same strong asymmetry. Furthermore, the three specification tests and the joint test fail to provide evidence against the specification. Finally, all the results we report regarding integration and contagion are qualitatively robust to the use of these U.S. residuals.

The next part of table 1 presents the regional model estimation. We find little evidence of asymmetric volatility outside the United States. We fail to reject symmetry in the European, Asian, and Latin American regional portfolios. The three specification tests and the joint test fail to provide evidence against the specifications at conventional significance levels. The local instruments have significant explanatory power in Asia but not in Latin America or Europe.

Table 1 also presents a test of whether the coefficients on the trade variable account for variation in the beta with respect to the United States (Wald test II), and we find no significant effect. However, do find that the beta with respect to the United States (Wald test III) is significantly different from zero for Europe and Latin America. For Asia, the p-value is 0.11. We also report the average conditional betas and correlations of the three regional portfolios with the United States. Europe has the highest average conditional correlation with the United States. (0.587), followed by Latin America (0.432), and Asia (0.146). Latin America's high β (around one) translates into lower correlation and a low proportion of variance explained by U.S. shocks because of its relatively high return variability. In Europe, more than 30% of the conditional return variance can be attributed to U.S. shocks.

These conditional betas and correlations are the cornerstone of our tests of contagion and market integration. We are interested in whether these betas and correlations increase during crisis periods. Our framework gives us the ability to decompose the increased correlation of returns into two components: the part the asset pricing model explains and the part the model does not explain. The explained part provides potential insights about market integration through the movements in the betas. We define contagion as the correlation of the unexplained portion.

C. Country Models and Integration

Our framework nests at least three distinct models: an asset pricing model with a single factor (a regional portfolio return), an asset pricing model with a single U.S. factor, and a world capital asset pricing model.

Detailed country-by-country results are available on request. Here, we summarize the main findings.

In Europe, most country residuals still display asymmetric volatility, but in the other regions, only half the countries do. Our residual specification tests typically fail to reject. The joint test is the most powerful, rejecting at the 1% level in three countries (Greece, Turkey, and Colombia) and at the 5% level in four other countries.

We also test whether lagged local information enters the mean equation (test of $\boldsymbol{\delta}_i = \mathbf{0}$). If the asset pricing model is properly specified, these lagged instruments should not enter the model. That is, the asset pricing model through its time-varying risk and risk premiums should capture variation in the country's conditional mean. This test can be thought of as a test of whether the conditional alpha (or pricing error) is zero and, under the null hypothesis of the regional or world CAPM, as a test of market integration.

The hypothesis that local information is unrelated to the pricing errors is rejected in 7 of 10 European countries. In Asia, local information is important for explaining the pricing errors in Korea, Malaysia, the Philippines, and Taiwan. Local information is also important for the pricing errors in Argentina, Chile, Colombia, and Mexico.

Similar to our regional analysis, we are interested in whether the beta with respect to the U.S. is influenced by trade with the U.S. and, more generally, whether the beta is equal to zero (to test whether $\mathbf{p}_{1,i} = \mathbf{0}$). We find that U.S. trade affects the conditional betas in 8 of 10 European countries (the exceptions are Austria and Portugal), 5 of 6 Asian countries (the exception is the Philippines), and 2 of 6 Latin American countries. The tests of whether the betas are equal to zero closely mimic the tests of whether trade is important. The beta with respect to the United States is significantly different from zero in 8 of 10 European sample countries. The beta with respect to the United States is not zero at the 5% level for all Asian countries except for the Philippines (where it is significant at the 10% level). In Latin America, three of the countries (Chile, Colombia, and Venezuela) have statistically significant nonzero betas with respect to the United States.

At the country level, we can also examine how trade with the rest of the world affects the regional beta and, more generally, whether the beta with respect to the regional benchmark is equal to zero (to test whether $\mathbf{p}_{2,i} = \mathbf{0}$). Trade with the rest of the world is important for 9 of 10 European markets' regional betas. In Asia, five of the six countries have regional betas that are significantly influenced by trade (at the 10% level). The regional beta of three Latin American countries (Chile, Colombia, and Mexico) is influenced by trade. In the more general tests of whether the regional betas equal zero, we find that 8 of 10 European countries have nonzero betas and that Austria does when the test is conducted at the 10% level. In Asia, the betas are nonzero (at the 5%

level) for all countries except for the Philippines, where the test rejects at the 10% level. In contrast, only two of Latin American countries' regional betas are statistically nonzero: Chile and Colombia.

We also test the significance of the total trade size as a percentage of GDP in the U.S. and regional betas. Here, we find that trade affects the betas of 9 of 10 European countries (the exception is Austria), 5 of 6 Asia countries (the exception is Thailand) but only 1 of 6 Latin American countries (Chile).

The more interesting tests restrict two sets of parameters. If both $\mathbf{p}_{1,i} = \mathbf{0}$ and $\mathbf{q}_i = \mathbf{0}$, then the model reduces to a CAPM with a single regional factor. This regional factor model is rejected at the 5% level for all countries except for Venezuela. If both $\mathbf{p}_{2,i} = \mathbf{0}$ and $\mathbf{q}_i = \mathbf{0}$, the model reduces to a single factor model with the U.S. market return as the relevant benchmark. This model is rejected for 20 of 22 countries at the 5% significance level, with Mexico and Venezuela being the only exceptions. If both $\mathbf{p}_{1,i} = \mathbf{0}$ and $\mathbf{p}_{2,i} = \mathbf{0}$, the model reduces to a standard world CAPM model. The simple world CAPM is rejected in 20 of 22 countries at the 10% level and 21 of 22 countries at the 5% level. The countries adhering to the world CAPM are Portugal and Venezuela (at the 10% level).

These Wald tests reveal that the special cases are usually rejected. Consequently, a regional international model is not a good description of the data by itself, but the covariance with one regional benchmark is a significant determinant of expected returns in most markets.

Table 2 reports average betas, correlations, and variance ratios for all the countries with respect to the U.S. and regional markets. Note that our model produces time-varying betas, correlations, and variance ratios, but we report only the sample average of these conditional variables.

First, let us focus on the small European markets. The betas and correlations with respect to the U.S. market are surprisingly small for most markets and even negative for Turkey. The exceptions are Finland (dominated by Nokia, a very international, U.S. listed company) with a beta of 0.883 and Norway, an oil sensitive economy, with a beta of 0.784. However, with the exception of Greece, betas and correlations with the regional market (the European index) are always larger than with the U.S. market. Given the small size of these markets and their correspondingly small weight in the index, this is not spuriously accounted for by index composition. Not surprisingly, this implies that the fraction of the return shock variance explained by U.S. factors is small, mostly in the 15–22% range. The regional market accounts for 25–35% of total shock variance, with the exceptions being Greece (close to 0%) and Turkey (3.4%). The qualitative nature of the results is definitely in line with what we would expect given the relative idiosyncratic nature of various markets.

TABLE 2 Implied Statistics of the Country-Specific Model

Market	$\hat{\beta}_i^{us}$	$\hat{\beta}_i^{reg}$	$\hat{\rho}_{i,\mathrm{us},t}$	$\hat{\rho}_{i,\mathrm{reg},t}$	$\mathrm{V}\hat{\mathrm{R}}_{i,t}^{us}$	$\mathrm{V}\hat{\mathrm{R}}_{i,t}^{reg}$
European Countries						
Austria	.224	.954	.153	.527	.038	.308
	(.155)	(.242)	(.121)	(.105)	(.055)	(.125)
Belgium	.509	.868	.398	.721	.174	.362
	(.081)	(.103)	(.126)	(.119)	(.119)	(.122)
Denmark	.459	.724	.345	.596	.136	.243
	(.125)	(.154)	(.129)	(.093)	(.112)	(.106)
Finland	.883	.976	.416	.573	.196	.168
	(.297)	(.277)	(.150)	(.136)	(.143)	(.084)
Greece	.248	−.048	.110	.053	.026	.002
	(.231)	(.100)	(.117)	(.061)	(.040)	(.006)
Norway	.784	.799	.448	.604	.228	.173
	(.277)	(.071)	(.166)	(.093)	(.146)	(.052)
Portugal	.675	.971	.357	.594	.150	.251
	(.249)	(.292)	(.150)	(.179)	(.126)	(.156)
Spain	.606	.963	.370	.642	.171	.275
	(.264)	(.177)	(.184)	(.168)	(.151)	(.100)
Sweden	.643	.903	.409	.658	.186	.260
	(.143)	(.057)	(.137)	(.093)	(.136)	(.067)
Turkey	−.241	.795	−.046	.100	.087	.034
	(1.273)	(.511)	(.293)	(.168)	(.119)	(.034)
Asian Countries (with the Asia emerging market index being the regional market)						
Indonesia	.849	.448	.251	.289	.100	.255
	(.615)	(.964)	(.194)	(.426)	(.140)	(.237)
Korea	.139	.169	.077	.220	.009	.056
	(.049)	(.047)	(.053)	(.108)	(.012)	(.064)
Malaysia	.875	.334	.443	.372	.237	.134
	(.333)	(.269)	(.204)	(.247)	(.182)	(.118)
Philippines	.767	.442	.284	.351	.104	.192
	(.398)	(.464)	(.152)	(.314)	(.116)	(.184)
Taiwan	−.055	.558	.031	.398	.059	.177
	(.696)	(.278)	(.242)	(.110)	(.080)	(.098)
Thailand	.723	.650	.302	.521	.132	.278
	(.365)	(.343)	(.201)	(.176)	(.162)	(.146)
Asian Countries (with MSCI Pacific being the regional market)						
Indonesia	.449	−.156	.173	−.035	.052	.032
	(.296)	(.276)	(.148)	(.147)	(.063)	(.050)
Korea	.425	.524	.190	.397	.048	.139
	(.175)	(.221)	(.111)	(.160)	(.082)	(.101)
Malaysia	.906	.232	.432	.311	.250	.040
	(.508)	(.160)	(.253)	(.193)	(.211)	(.042)
Philippines	.688	.452	.290	.336	.120	.120
	(.381)	(.536)	(.190)	(.225)	(.151)	(.151)
Taiwan	.157	.397	.075	.253	.038	.072
	(.419)	(.183)	(.180)	(.172)	(.068)	(.064)
Thailand	.780	.413	.302	.312	.124	.090
	(.412)	(.452)	(.181)	(.212)	(.138)	(.089)

TABLE 2 **(*Continued*)**

Market	$\hat{\beta}_i^{us}$	$\hat{\beta}_i^{reg}$	$\hat{\rho}_{i,us,t}$	$\hat{\rho}_{i,reg,t}$	$V\hat{R}_{i,t}^{us}$	$V\hat{R}_{i,t}^{reg}$
Latin American Countries						
Argentina	.927	.781	.263	.474	.100	.185
	(.203)	(.233)	(.178)	(.209)	(.135)	(.137)
Brazil	1.205	.825	.324	.475	.131	.154
	(.257)	(.174)	(.162)	(.186)	(.127)	(.113)
Chile	.537	.090	.293	.254	.097	.058
	(.093)	(.139)	(.104)	(.206)	(.091)	(.102)
Colombia	.216	−.015	.123	.033	.021	.009
	(.075)	(.069)	(.074)	(.109)	(.033)	(.014)
Mexico	.907	.217	.354	.272	.143	.052
	(.203)	(.224)	(.134)	(.197)	(.114)	(.050)
Venezuela	.413	.017	.139	.073	.048	.009
	(.445)	(.131)	(.170)	(.126)	(.084)	(.013)

NOTE.—The following model is estimated for the country portfolios:

$$R_{i,t} = \delta_i'\mathbf{Z}_{i,t-1} + \beta_{i,t-1}^{us}\hat{\mu}_{us,t-1} + \beta_{i,t-1}^{reg}\hat{\mu}_{reg,t-1} + \beta_{i,t-1}^{us}\hat{e}_{us,t} + \beta_{i,t-1}^{reg}\hat{e}_{reg,t} + e_{i,t}$$

$$e_{i,t} \mid I_{t-1} \sim N(0, \sigma_{i,t}^2)$$

$$\sigma_{i,t}^2 = a_i + b_i\sigma_{i,t-1}^2 + c_i e_{i,t-1}^2 + d_i\eta_{i,t-1}^2$$

$$\eta_{i,t-1} = \min\{0, e_{i,t-1}\}$$

where $e_{i,t}$ is the idiosyncratic shock of market i, $\hat{\mu}_{us,t-1}$ and $\hat{e}_{us,t}$($\hat{\mu}_{reg,t-1}$ and $\hat{e}_{reg,t}$) are the conditional expected excess return and residual on the U.S. (regional) market. $\mathbf{Z}_{i,t-1}$ represents a set of local information variables. The region market is represented by the MSCI Europe index, and the Asia and Latin America emerging market indices, a value-weighted average of all the Asian (Latin American) emerging markets, excluding the country under consideration.

Sample average of standard deviation of the implied beta parameters ($\hat{\beta}_i^{us}$ and $\hat{\beta}_i^{reg}$), correlations with the U.S. and regional markets ($\hat{\rho}_{i,us,t}$ and $\hat{\rho}_{i,reg,t}$), and variance ratios accounted for by the United States and region ($V\hat{R}_{i,t}^{us}$ and $V\hat{R}_{i,t}^{reg}$) are reported. Standard deviations are given in parentheses.

The results for the Asian markets are somewhat surprising. The betas with respect to the U.S. market factor are quite high, exceeding 0.7 in four of the six markets. Only Korea and Taiwan display very small betas. The correlations are lower, because of the higher idiosyncratic volatility of these markets. Except for Korea and Taiwan, the beta with respect to the United States is always larger than the beta with respect to the regional market. In terms of variance ratios, Taiwan and Korea are similar to Greece and Turkey: The U.S. and regional factors do not account for very much of the total variation of return shocks. However, the other markets are closer to what we see for the European markets, with the regional and U.S. factors jointly accounting for over 30% of the variance of return shocks.

For the Latin-American countries, high betas with respect to the U.S. market are no surprise, but there is substantial cross-country variation, which ranges from 0.413 for Venezuela to 1.205 for Brazil. The regional betas are always much smaller than the U.S. betas. This is also true, to a lesser extent, for the correlations, with the exceptions being

Argentina and Brazil. Overall, this analysis suggests that regional integration may not be as strong a phenomenon as previously thought. Examining the variance ratios, in four of the six countries we explain less than 20% of the shock variance with both the U.S. and regional factors. Only in Argentina and Brazil do we explain around 28% of the variance, which is still lower than what we observe for most European and the Asian markets. These results also help us calibrate the results on changes in betas, correlations and variance ratios during crises times.

D. Patterns in Regional and Global Integration

We investigate patterns in regional and global integration by examining how the estimated betas and correlations change during particular periods. We also examine the patterns in the variance ratios (amount of variance in the country's unexpected return accounted for by the United States or region). We consider five sample periods: the second half of the sample (or subsample), the Mexican crisis, the Asian crisis, periods of abnormally negative U.S. unexpected returns, and periods of abnormally negative regional unexpected returns. *Abnormal* is defined as more than one standard deviation below zero. We run panel regressions of each of our measures on a constant and on a dummy variable that takes a value of one during these designated periods.

The first panel of table 3 compares the first half of the sample to the second half, which is dominated by the 1990s. For most countries, the betas, correlations, and variance ratios with respect to the United States and the region increase, leading to positive-slope coefficients. This increase suggests increased linkages among the various countries. In Asia, there is a sharp increase in the regional betas, correlations, and variance ratios in the second half of the sample. In general, the regional correlations, betas, and variance ratios increase by more than their U.S. counterparts. There are some exceptions. In Europe, the betas with respect to the United States increase somewhat more than the regional betas. This is somewhat surprising, given that the second subsample is a time when Europe was moving further toward unification and a single currency. However, it is probably best to interpret these results as showing increased correlation both within the region and with respect to the United States, given that there is little economic difference between the increases.

Panel B of table 3 examines the Mexican crisis. For Latin America, there is no significant increase in the regional beta or correlation during the crisis. At only 0.004, the increment to regional correlation is not even one standard error from zero. Indeed, the regional beta decreases while the beta with respect to the United States increases, but neither change is significantly different from zero. The regional variance ratio change is not significantly above zero. Overall, the model suggests no change in correlation during this crisis period.

TABLE 3 Patterns in Regional Integration

Market	$\beta^{us}_{i,t-1}$	$\beta^{reg}_{i,t-1}$	$\beta^{reg}_{i,t-1} - \beta^{us}_{i,t-1}$	$\rho_{i,\text{us},t}$	$\rho_{i,\text{reg},t}$	$\rho_{i,\text{reg},t} - \rho_{i,\text{us},t}$	$VR^{us}_{i,t}$	$VR^{reg}_{i,t}$	$VR^{reg}_{i,t} - VR^{us}_{i,t}$
Panel A. Second Subsample Dummy									
Europe	.054	.044	−.075	.065	.061	.001	.016	.017	.003
	(.010)	(.007)	(.008)	(.008)	(.006)	(.005)	(.005)	(.003)	(.005)
Europe/Turkey	.042	.043	−.063	.061	.058	.003	.018	.021	.003
	(.009)	(.007)	(.009)	(.008)	(.006)	(.005)	(.005)	(.004)	(.005)
Asia	−.036	.261	.164	−.026	.182	.170	−.007	.084	.079
	(.013)	(.022)	(.025)	(.009)	(.017)	(.022)	(.004)	(.010)	(.014)
Latin America	.130	.063	−.011	.089	.144	.063	.033	.016	−.007
	(.013)	(.009)	(.014)	(.010)	(.012)	(.009)	(.006)	(.004)	(.005)
Panel B. Mexico Crisis Dummy									
Europe	.003	.023	−.040	.023	.031	.006	−.004	.006	.010
	(.019)	(.013)	(.020)	(.015)	(.011)	(.010)	(.008)	(.004)	(.010)
Europe/Turkey	.006	.019	−.046	.027	.034	.003	−.003	.005	.009
	(.019)	(.013)	(.018)	(.016)	(.011)	(.010)	(.009)	(.004)	(.010)
Asia	.015	.016	.005	.034	.073	.049	.005	.028	.018
	(.021)	(.018)	(.033)	(.016)	(.025)	(.026)	(.004)	(.016)	(.018)
Latin America	.013	−.025	.003	−.006	−.003	.004	−.008	−.005	.008
	(.020)	(.018)	(.020)	(.016)	(.022)	(.017)	(.008)	(.003)	(.009)
Panel C. Asia Crisis Dummy									
Europe	.141	.058	−.063	.139	.074	−.040	.077	−.003	−.054
	(.018)	(.010)	(.018)	(.013)	(.010)	(.009)	(.008)	(.002)	(.011)
Europe/Turkey	.121	.059	−.049	.134	.068	−.037	.084	−.003	−.065
	(.017)	(.010)	(.016)	(.013)	(.010)	(.009)	(.009)	(.002)	(.012)
Asia	−.096	.287	.280	−.075	.067	.065	−.010	.028	.026
	(.014)	(.036)	(.042)	(.012)	(.022)	(.031)	(.005)	(.015)	(.020)
Latin America	.092	.085	−.003	.104	.158	.051	.071	.008	−.030
	(.018)	(.015)	(.017)	(.015)	(.018)	(.014)	(.010)	(.005)	(.009)

TABLE 3 (*Continued*)

Market	$\beta^{us}_{i,t-1}$	$\beta^{reg}_{i,t-1}$	$\beta^{reg}_{i,t-1} - \beta^{us}_{i,t-1}$	$\rho_{i,us,t}$	$\rho_{i,reg,t}$	$\rho_{i,reg,t} - \rho_{i,us,t}$	$VR^{us}_{i,t}$	$VR^{reg}_{i,t}$	$VR^{reg}_{i,t} - VR^{us}_{i,t}$
Panel D. Abnormally Negative U.S. Unexpected Return Dummy									
Europe	−.002	−.002	−.0004	.003	.002	−.003	.006	.001	−.008
	(.013)	(.008)	(.012)	(.011)	(.008)	(.007)	(.006)	(.001)	(.007)
Europe/ Turkey	−.002	−.002	−.0002	.002	.001	−.003	.007	.001	−.008
	(.013)	(.008)	(.012)	(.011)	(.008)	(.007)	(.006)	(.001)	(.007)
Asia	.002	−.010	−.013	.010	−.006	−.017	.004	−.012	−.012
	(.015)	(.014)	(.027)	(.013)	(.021)	(.022)	(.003)	(.014)	(.015)
Latin America	−.002	−.001	.002	.010	.013	.002	.011	.001	−.007
	(.016)	(.014)	(.016)	(.013)	(.018)	(.014)	(.007)	(.003)	(.007)
Panel E. Abnormally Negative Regional Unexpected Return Dummy									
Europe	−.010	−.005	−0.003	.018	.005	−.016	.014	−.001	−.013
	(.013)	(.009)	(.013)	(.011)	(.008)	(.007)	(.006)	(.002)	(.007)
Europe/ Turkey	−.009	−.005	−.003	.020	.006	−.017	.015	−.002	−.016
	(.013)	(.009)	(.013)	(.011)	(.008)	(.007)	(.006)	(.002)	(.007)
Asia	.017	−0.026	−.035	.025	−.012	−.054	.005	−.017	−.030
	(.021)	(.018)	(.034)	(.017)	(.024)	(.028)	(.005)	(.016)	(.018)
Latin America	−.024	−.018	.001	.020	.013	−.006	.018	.006	−.007
	(.018)	(.016)	(.018)	(.014)	(.020)	(.016)	(.008)	(.003)	(.008)

NOTE.—The following time-series cross-section regression model is estimated:

$$S_{i,t} = \kappa_i + \phi D_{i,t} + u_{i,t}$$

where $S_{i,t}$ denotes the implied statistic being examined, such as $\beta^{us}_{i,t-1}$, $\beta^{reg}_{i,t-1}$, $\beta^{reg}_{i,t-1} - \beta^{us}_{i,t-1}$, market correlations, $\rho_{i,us,t}$, $\rho_{i,reg,t}$, $\rho_{i,reg,t} - \rho_{i,us,t}$, as given in eqq. (12) and (13), and variance ratios, $VR^{us}_{i,t}$, $VR^{reg}_{i,t}$, $VR^{reg}_{i,t} - VR^{us}_{i,t}$, as in eqq. (14)and(15). $D_{i,t}$ is a dummy variable that represents (A) the second subsample period, (B) the Mexico crisis period from November 1994 to December 1995, (C) the Asia crisis period from April 1997 to October 1998, (D) abnormally negative U.S. unexpected market returns (i.e., the unexpected returns are one standard deviation below zero), and (E) abnormally negative regional unexpected market returns. In studying the implied statistics, countries are categorized into four country groups: Europe, Europe excluding Turkey, Asia, and Latin America. The estimation results correct for groupwise heteroscedasticity with Newey-West correction for serial correlation (with one lag). The parameter estimates of ϕ are reported, with standard errors given in parentheses.

The Asian crisis presents a completely different story. The regional correlations, betas, and variance ratios for Asia in panel C of table 3 increase by economically meaningful magnitudes in Asia, and the change appears statistically significant. The fact that the effect is economically much smaller for correlations and variance ratios than for betas is due to higher overall volatility during this period. Comovements with the United States appear to have decreased during the period, suggesting increased economic regional integration during this time. Interestingly, the same regional effects are evident in both Latin America and Europe during this period, but they are smaller and, in fact, are dominated by increased comovement with the U.S. market.

The last two panels of table 3 examine periods of large negative returns. In periods of negative abnormal returns, we would expect the asymmetric model to generate higher correlations. Although these negative abnormal returns are usually associated with higher correlation, the increment in correlation is substantially smaller than that experienced during the Asian crisis. Indeed, the last two panels help calibrate the importance of Asian crisis with respect to our comovement variables.

E. Contagion

The increased correlation detected during the Asian crisis is not itself evidence of contagion. The hypothesis of contagion would be supported if the model's idiosyncratic shocks exhibit significant correlation. Table 4 provides a baseline estimate of contagion over the full sample. We examine the correlation of the country's idiosyncratic shocks with the U.S. residuals, the regional residuals, and every other country's idiosyncratic residuals.

To assess the statistical significance of the residual correlations, we perform a bootstrap exercise based on 5,000 draws of the actual return residual set, with the same number of observations as in our sample. The bootstrap experiment is constructed as follows. First, we compile all the idiosyncratic shocks from all markets, including the U.S. and regional indices, into one grand vector of return shocks. Second, in each replication, we draw from the grand return shock vector to construct a matrix of return shocks with the same number of observations as in the sample (rows) and number of countries (columns), then compute the bivariate correlation and cross-country correlation matrix. We use 5,000 replications in all. Finally, we record the 95% values for the bivariate correlation and the cross-country correlation matrix.

In the first panel of table 4, we find that there is no evidence of excess correlation between the European countries and the United States. One country, Greece, has excess correlation with the regional residual. However, most interestingly, we find evidence in all but one country (Belgium) of contagion among the European countries. That is, the

Table 4 Correlations of Market Residuals

Market	Correlations of $e_{i,t}$ and $e_{us,t}$	$e_{i,t}$ and $e_{reg,t}$	$e_{i,t}$ and $e_{j,t}$ Mean	Maximum	Minimum
European Countries					
Austria	.039	.060	.070+	.257	−.036
Belgium	.028	−.015	−.009	.110	−.132+
Denmark	−.005	.008	.042+	.146	−.036
Finland	−.046	−.043	.081+	.327+	−.132+
Greece	.104	.212+	.134+	.350+	−.006
Norway	.052	.071	.132+	.249	.006
Portugal	−.034	.014	.114+	.350+	−.074
Spain	−.001	−.015	.071+	.166	−.049
Sweden	.041	−.019	.097+	.327+	−.125+
Turkey	.011	−.002	.131+	.283+	.013
Asian Countries					
Indonesia	−.114	.018	.097+	.197	−.154+
Korea	.105	−.056	.123+	.314+	−.116+
Malaysia	.118	.173	.171+	.316+	.042
Philippines	−.054	−.066	.091+	.218	−.114+
Taiwan	.136	.013	−.105	.042	−.183+
Thailand	.049	−.032	.158+	.316+	−.183+
Latin American Countries					
Argentina	−.036	−.092	−.028	.056	−.099+
Brazil	−.134	−.286	−.072	.039	−.205+
Chile	.028	.011	−.001	.115	−.099+
Colombia	.0005	−.005	.068+	.180	−.012
Mexico	.135+	.057	−.034	.017	−.205+
Venezuela	−.038	.002	.039+	.180	−.064

NOTE.—The following model is estimated for the country portfolios:

$$R_{i,t} = \delta_i' \mathbf{Z}_{i,t-1} + \beta^{us}_{i,t-1}\hat{\mu}_{us,t-1} + \beta^{reg}_{i,t-1}\hat{\mu}_{reg,t-1} + \beta^{us}_{i,t-1}\hat{e}_{us,t} + \beta^{reg}_{i,t-1}\hat{e}_{reg,t} + e_{i,t}$$

$$e_{i,t} \mid I_{t-1} \sim N\left(0, \sigma^2_{i,t}\right)$$

$$\sigma^2_{i,t} = a_i + b_i\sigma^2_{i,t-1} + c_i e^2_{i,t-1} + d_i\eta^2_{i,t-1}$$

$$\eta_{i,t-1} = \min\left\{0, e_{i,t-1}\right\}$$

where $\hat{\mu}_{us,t-1}$ and $\hat{\mu}_{reg,t-1}$ are the conditional expected excess returns on the U.S. and regional markets, respectively, and $e_{i,t}$ is the idiosyncratic shock of any market i, including the U.S. and regional portfolio. $\mathbf{Z}_{i,t-1}$ represents a set of local information variables.

The correlations of market residuals across different markets are computed over the longest possible overlapping sample between the two markets. The + symbol indicates 5% rejection of the null hypothesis of zero correlation according to the appropriate small sample distribution. The small sample distribution is computed based on 5,000 draws of the actual return residual set, $\{\hat{e}_{j,t}\}_{\forall j}$, with the same number of observations as the markets in our sample.

residual correlations among the different countries are significantly above zero.

In the second panel of table 4, we find no evidence of excess correlation with the U.S. residual or the regional residual for any of the Asian countries; however, there is sharp evidence of average contagion within Asia. Every country except Taiwan has significant excess correlation with other Asian countries. Indeed, the magnitude of these correlations is roughly double what we documented for Europe.

The Latin American countries are presented in the final panel of table 4. Only one country, Mexico, has excess correlation with the U.S. portfolio. No country has excess correlation with the regional portfolio. Only two countries, Colombia and Venezuela, show significant excess correlation with other Latin American countries.

This analysis measures the correlation of idiosyncratic residuals over the entire sample. We are most interested in the time-series patterns of these residuals. Table 5 uses the five time periods introduced in table 3 to examine patterns in the residuals. We use a panel regression of the country's idiosyncratic shocks onto a country-specific constant and a relevant benchmark residual with the slope coefficient allowed to change during the time periods of interest. We estimate the panel model accommodating groupwise heteroscedasticity.[7]

In panel A, the v_1 coefficient measures the additional correlation in the second half of the sample. Regardless of the benchmark or region, v_1 is positive, suggesting that the idiosyncratic residuals are better correlated in the second half of the sample. The correlation with respect to the U.S. index residuals is significantly higher only for Asia; however, the correlation with the regional residuals is significantly higher for all regions in the second half of the sample. Considering the sum of the country-specific residuals, we find that the correlation jumps significantly in the second half of the period for both Asia and Latin America.

The joint test of $v_0 = v_1 = 0$ is an overall test of contagion. We reject at the 5% level for Asia with respect to the U.S. index, for Latin America with respect to regional return residuals, and for all regions with respect to the "sum of other residuals" benchmark. In the last case, v_0 is by itself also significant for three regions. Clearly, country residuals within a region are correlated beyond what is captured in our model, suggesting evidence of contagion. Overall, contagion worsened in the second half of the sample, but it is only economically and statistically significant for Asia and Latin America.

Panel B of table 5 examines the Mexican crisis. Our results show that there is no significant increase in residual correlations within Latin

7. We also conducted the estimation assuming an AR(1) model for the residual. These alternative estimates yield qualitatively similar results.

TABLE 5 Cross-Section Analysis of Market Residuals

	U.S. Return Residuals ($\hat{e}_{us,t}$)				Regional Return Residuals ($\hat{e}_{reg,t}$)				Sum of Return Residuals ($\sum_{j\in G}^{j\neq i}\hat{e}_{j,t}$)			
			Wald Test				Wald Test				Wald Test	
Country Group	v_0	v_1	$\{w_i=0\}_\forall$	$v_0 = v_1 = 0$	v_0	v_1	$\{w_i = 0\}_\forall$	$v_0 = v_1 = 0$	v_0	v_1	$\{w_i = 0\}_\forall$	$v_0 = v_1 = 0$
Panel A. Second Subsample Dummy												
Europe	.015	.035	3.634	1.305	−.022	.136	3.702	4.743	.028	−.0002	3.708	42.58
	(.038)	(.061)	[.962]	[.521]	(.043)	(.069)	[.960]	[.093]	(.007)	(.009)	[.960]	[<.001]
Europe/Turkey	.016	.032	3.636	1.239	−.022	.136	3.701	4.706	.028	.011	3.646	37.44
	(.038)	(.061)	[.934]	[.538]	(.044)	(.069)	[.930]	[.095]	(.008)	(.011)	[.933]	[<.001]
Asia	−.020	.346	8.232	6.094	−.044	.160	8.271	5.868	−.017	.132	10.36	66.02
	(.096)	(.164)	[.222]	[.048]	(.040)	(.067)	[.219]	[.053]	(.020)	(.025)	[.110]	[<.001]
Latin America	−.066	.254	8.560	2.173	−.073	.202	10.07	9.024	−.037	.100	11.10	18.12
	(.100)	(.174)	[.200]	[.337]	(.036)	(.069)	[.122]	[.011]	(.013)	(.024)	[.085]	[<.001]
Panel B. Mexico Crisis Dummy												
Europe	.025	.190	3.588	1.807	.031	−.003	3.647	.847	.029	−.010	3.702	42.86
	(.030)	(.207)	[.964]	[.405]	(.034)	(.245)	[.962]	[.655]	(.004)	(.026)	[.960]	[<.001]
Europe/Turkey	.025	.197	3.584	1.855	.032	−.037	3.637	.876	.033	.010	3.669	36.49
	(.030)	(.208)	[.937]	[.395]	(.034)	(.247)	[.934]	[.645]	(.006)	(.033)	[.932]	[<.001]
Asia	.099	.036	8.163	1.663	.005	.380	8.255	2.922	.069	.071	10.13	37.57
	(.079)	(.472)	[.226]	[.436]	(.033)	(.229)	[.220]	[.232]	(.012)	(.074)	[.119]	[<.001]
Latin America	.040	−.781	9.428	2.433	−.021	.039	9.169	.446	−.010	.048	9.580	1.517
	(.083)	(.506)	[.151]	[.296]	(.032)	(.126)	[.164]	[.800]	(.012)	(.044)	[.143]	0.468]

Panel C. Asia Crisis Dummy												
Europe	.020	.058	3.656	1.459	.010	.272	3.747	5.484	.027	.011	3.726	43.53
	(.032)	(.083)	[.962]	[.482]	(.035)	(.127)	[.958]	[.064]	(.005)	(.013)	[.959]	[<.001]
Europe/Turkey	.020	.058	3.655	1.442	.011	.264	3.743	5.177	.032	.009	3.672	36.83
	(.032)	(.083)	[.933]	[.486]	(.035)	(.127)	[.928]	[.075]	(.006)	(.017)	[.932]	[<.001]
Asia	−.001	.504	8.268	8.526	−.013	.155	8.336	3.348	.018	.111	10.28	59.26
	(.087)	(.193)	[.219]	[.014]	(.036)	(.087)	[.215]	[.187]	(.016)	(.024)	[.113]	[<.001]
Latin America	−.083	.525	8.701	6.540	−.057	.436	10.13	16.95	−.023	.201	12.94	24.24
	(.090)	(.206)	[.191]	[.038]	(.032)	(.107)	[.119]	[<.001]	(.011)	(.041)	[.044]	[<.001]
Panel D. Abnormally Negative U.S. Unexpected Return Dummy												
Europe	.016	.026	3.794	1.121	−.003	.200	3.874	5.764	.029	−.004	3.697	42.93
	(.046)	(.067)	[.956]	[.571]	(.037)	(.090)	[.953]	[.056]	(.005)	(.012)	[.960]	[<.001]
Europe/Turkey	.019	.018	3.740	1.031	−.002	.196	3.870	5.578	.036	−.012	3.642	37.25
	(.046)	(.067)	[.928]	[.597]	(.037)	(.090)	[.920]	[.061]	(.006)	(.014)	[.933]	[<.001]
Asia	−.185	.504	10.07	9.770	−.044	.187	8.522	7.157	.049	.088	10.07	46.65
	(.127)	(.177)	[.122]	[.008]	(.039)	(.071)	[.202]	[.028]	(.013)	(.028)	[.122]	[<.001]
Latin America	−.274	.512	13.11	7.580	−.067	.169	11.39	5.768	−.002	−.033	8.922	1.362
	(.135)	(.187)	[.041]	[.023]	(.037)	(.072)	[.077]	[.056]	(.012)	(.032)	[.178]	[.506]

Table 5 (*Continued*)

	U.S. Return Residuals ($\hat{e}_{us,t}$)				Regional Return Residuals ($\hat{e}_{reg,t}$)				Sum of Return Residuals ($\sum_{j \in G}^{j \neq i} \hat{e}_{j,t}$)			
			Wald Test				Wald Test				Wald Test	
Country Group	v_0	v_1	$\{w_i = 0\}_\forall$	$v_0 = v_1 = 0$	v_0	v_1	$\{w_i = 0\}_\forall$	$v_0 = v_1 = 0$	v_0	v_1	$\{w_i = 0\}_\forall$	$v_0 = v_1 = 0$
Panel E. Abnormally Negative Regional Unexpected Return Dummy												
Europe	−.007	.086	4.160	2.827	.032	−.002	3.629	.850	.028	.002	3.713	42.71
	(.040)	(.063)	[.940]	[.243]	(.043)	(.078)	[.963]	[.654]	(.005)	(.013)	[.959]	[<.001]
Europe/Turkey	−.003	.076	4.091	2.443	.031	.002	3.644	.864	.034	−.007	3.634	36.57
	(.040)	(.063)	[.905]	[.295]	(.043)	(.078)	[.933]	[.649]	(.006)	(.018)	[.934]	[<.001]
Asia	−.086	.498	8.520	11.29	−.036	.150	8.876	3.913	.073	−.009	10.22	36.76
	(.098)	(.161)	[.202]	[.004]	(.041)	(.078)	[.181]	[.141]	(.013)	(.029)	[.116]	[<.001]
Latin America	−.164	.417	10.29	6.286	−.099	.214	13.64	9.350	−.013	.053	9.800	3.056
	(.110)	(.167)	[.113]	[.043]	(.041)	(.071)	[.034]	[.009]	(.012)	(.032)	[.133]	[.217]

NOTES.—The following time-series cross-section regression model is estimated:

$$\hat{e}_{i,t} = w_i + v_{i,t}\hat{e}_{g,t} + u_{i,t}$$

$$v_{i,t} = v_0 + v_1 D_{i,t}$$

where $\hat{e}_{i,t}$ and $\hat{e}_{g,t}$ are the estimated idiosyncratic return residuals of market i and region g, respectively. For the regional residuals, three groups are considered: $\hat{e}_{g,t} = \hat{e}_{us,t}$, $\hat{e}_{g,t} = \hat{e}_{reg,t}$, and $\hat{e}_{g,t} = \sum_{j \in G}^{j \neq i} \hat{e}_{j,t}$ where G denotes a particular country group. In studying the market residuals, countries are categorized into nine country groups: Europe, Europe excluding Turkey, MSCI Pacific, Asia (IFC composite), Asia emerging markets, IFC Latin America, Latin America (IFC composite), Latin America emerging markets, and all markets. $D_{i,t}$ is a dummy variable that represents (A) the second subsample period, (B) the Mexico crisis period from November 1994 to December 1995, (C) the Asia crisis period from April 1997 to October 1998, (D) abnormally negative U.S. unexpected market returns (i.e., the unexpected returns are one standard deviation below zero), and (E) abnormally negative regional unexpected market returns. The estimation results correct for groupwise heteroscedasticity. The parameter estimates of v_0 and v_1 are reported, and standard errors are given in parentheses, while p-values are reported in brackets.

America. Insignificant effects are also found for Europe and Asia during the Mexican crisis, suggesting little evidence of contagion resulting specifically from the Mexican crisis. However, the overall contagion tests confirm the results of panel B, indicating contagion across countries in the region.

Panel C of table 5 presents the results for the Asian crisis. Here we see significantly higher residual correlations among Asian countries for all residual benchmarks. The increase in correlation for Asia is many times larger than the increase in correlation for either Latin America or Europe when investigating comovements with U.S. return residuals or the sum of the idiosyncratic residuals, but the increase in correlation is on the same order of magnitude when examining regional residuals. For Latin America and Europe, statistical significance is reached only in this last case. Hence, the Asian crisis worsened contagion.

Panels D and E of table 5 put the historical crisis periods in perspective. Compare the increases in regional excess correlation in panel E (abnormal negative regional unexpected returns) with those reported for the Asian crisis. The increase in residual correlations when regional returns are negative is on the same order of magnitude as the increase observed during the Asian crisis in panel C, except for the sum of idiosyncratic residuals, where the increase during the Asian crisis is much larger. The U.S. unexpected returns in panel D yield an increase in correlations for Asia on the same order of magnitude as observed for all of panel C, even for the sum of idiosyncratic residuals. One interpretation of this result is that our model fails to capture fully asymmetric volatility (higher volatility in bear markets) and the potential effects it has on correlations during crisis periods.[8] If this is the case, what we call *contagion* here for Asia may no longer be considered contagion vis-à-vis a richer model.

IV. Conclusions

Contagion is a level of correlation over what is expected. Considerable research has investigated whether the crises in both Mexico and Asia in the 1990s resulted in contagion. Our research directly addresses the issue of "what is expected." We present a two-factor asset pricing model and define *contagion* as correlation among the model residuals. It is important here to operate within the framework of a model. Indeed, increased return correlation between two countries during a period of crisis could simply be the consequence of their exposure to a common factor. That is, it is necessary to undo the natural changes in correlation

8. Ang and Bekaert (2002) report that GARCH models (although simpler than the ones we study here) do not capture asymmetric correlations very well and propose a regime switching model.

that result from an asset pricing model, before making statements about contagion.

Our framework allows for time-varying expected returns as well as time-varying risk loadings for the countries we examine. Our results suggest that there is no evidence of additional contagion caused by the Mexican crisis. However, we find economically meaningful increases in residual correlation, especially in Asia, during the Asian crisis. Dungey and Martin (2001), using a different methodology, find similar results for Asia and explore the role of currency risk in equity market contagion.

One useful extension of our methodology could be to investigate contagion in currency markets and link equity to currency contagion. In fact, our framework is very different from the typical empirical strategy used in the international economics literature, where crisis indicators in one country (e.g., the probability of a speculative attack or the magnitude of a crisis indicator) are directly linked to indicators in other countries (see De Gregario and Valdés 2001 and Eichengreen, Rose, and Wyplosz 1997). As Rigobon (1999) also stresses, this approach is problematic in Q6 the presence of common unobservable shocks and increased variances during crisis periods. Our asset pricing approach, which directly models the shock and correlation structure and uses crisis and noncrisis periods for identification, does not suffer from these problems. Of course, it is possible that our model of correlations is incorrect and contagion could simply be a result of model misspecification. Nevertheless, we believe that it is more desirable to frame statements about excess correlation in the context of an asset-pricing model.

Appendix

Information Variable Specification

In estimating the time-varying beta model in eqq. (1)–(5), we introduce several sets of information variables in the empirical model. This appendix provides a detailed discussion of these information variables.

Stage 1. U.S. Model

The U.S. instrument set, $\mathbf{Z}_{\mathrm{us},t-1}$, includes a constant, the lagged world market dividend yield, the lagged spread between the 90-day Eurodollar rate and the 3-month Treasury-bill yield; the difference between the U.S. 10-year Treasury-bond yield and the 3-month Treasury-bill yield, and finally, the change in the 90-day Treasury-bill yield.

Stage 2. Regional Model

The regional instrument set, $\mathbf{Z}_{\mathrm{reg},t-1}$, includes a constant and regional market dividend yield. Note that, for the MSCI Europe, the market-capitalization weighted dividend yield includes the large, developed markets, such as France, Germany, Italy, the Netherlands, Switzerland, and the United Kingdom. The Asian or Latin American emerging market indices are market-capitalization weighted over all the markets in Asia or Latin America except for the one under examination.

The trade data set, $\mathbf{X}^{us}_{reg,t-1}$, from the World Bank CD-ROM, includes a constant and the adjusted trade with the United States (i.e., sum of exports to and imports from the United States divided by the sum of total exports and imports) lagged 6 months. For the MSCI Europe, we aggregate over France, Germany, Italy, the Netherlands, Switzerland, and the United Kingdom. For our Asian or Latin American emerging market indices, we aggregate over all the markets in Asia or Latin America except for the one under examination.

Stage 3. Individual Country Models

The individual country instrument set, $\mathbf{Z}_{i,t-1}$, includes a constant and the local dividend yield (Source: IFC).

The data set for trade with the United States, $\mathbf{X}^{us}_{i,t-1}$, includes a constant as well as adjusted trade with the United States (i.e., sum of exports to and imports from the United States divided by the sum of total exports and imports) lagged 6 months. Note, for Belgium, there are many missing values in its trade with the United States; therefore, we replace the data for Belgium with the average of France, Germany, and the Netherlands.

The data set for trade with the region, $\mathbf{X}^{reg}_{i,t-1}$, includes a constant and the sum of exports to and imports from the rest of the world, except to and from the United States, divided by the sum of total exports and imports, lagged 6 months.

The trade with the rest of the world data set, $\mathbf{X}^{w}_{i,t-1}$, includes a constant and total trade by GDP, lagged 6 months.

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