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Russian and East European Finance and Trade, vol. 37, no. 5,
September–October 2001, pp. 45–63.

ISSN 1061–2009/2001 $9.50 + 0.00.

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Markowitz's (1959) approach to portfolio diversification indicates that today the global investor can earn potential gains from international diversification rather than domestic diversification, as long as returns in different countries are less correlated than those in domestic markets. Therefore, international correlations between stock returns are important for the global investor (Solnik 1991). In fact, low correlations have been reported among international returns supporting the benefits of international diversification (Granger and Morgenstern 1970; Speidell and Sappanfield 1992).

Following a correlation approach, Lee and Kim (1994) examine the effect of the October 1987 crash on the co-movements among national stock markets. Interrelationships among the price movements in different national stock markets are analyzed using correlation and exploratory factor analysis. The data on weekly returns of twelve national stock market indexes over the period August 1984 to December 1990 are used in both local currency and U.S. dollar terms for the analysis. This study finds that national stock markets became more interrelated after the crash, and the strengthening co-movements among national stock markets continued for a longer period after the crash. In addition, it is shown that the co-movements among national stock markets were stronger when the U.S. stock market was more volatile.

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Tang (1995) examines the inter-temporal stability in stock market co-movements. Contrary to previous findings, the empirical results show that for both domestic currency and U.S. dollar-based returns, the shorter the time period considered, the more stable the patterns of stock market co-movement, especially in the period before the 1987 stock crash when domestic currency returns are used. Darbar and Deb (1997) examine the co-movements of equity returns in major international markets by characterizing the time-varying cross-country covariances and correlations. Using a generalized positive definite multivariate generalized autoregressive conditional heteroscedasticity (GARCH) model, they find that the Japanese and U.S. stock markets have significant transitory covariance, but zero permanent covariance.

Another approach in investigating international co-movements is to focus on price discovery in world markets. Naturally, cointegration and error correction modeling provides a useful framework for analyzing price adjustments in internationally linked markets. Harris et al. (1995) for example, investigate New York, Pacific, and Midwest exchanges, and conclude that bidirectional price adjustments take place on all three exchanges. McInish and Wood (1992) also report that regional exchanges are not free riders on primary exchanges, and they contain information that is relevant for traders at primary markets (Garbade and Silber 1979).

It appears that previous empirical studies on the relationship between world stock markets do not provide consistent results. The reasons for the inconsistent results are numerous, including the choice of markets, different sample periods, different frequency of observations, and the different methodologies employed. The focus of previous studies also creates problems with interpretation of results. Most studies are concerned with integration versus segmentation of markets as indicators of the degree of international diversification for the global investor.

The major contributions of this paper are as follows. In this study, the degree of market integration is investigated in order to forecast national markets according to their international co-movements. The focus of the paper is different from previous research that investigates market integration for global diversification. Besides, we attempt to maintain a research framework, whereby a coherent database is used, to include all of the emerging markets as classified by the International Finance Corporation (IFC). The data frequency is weekly for all countries, and the cointegration methodology is employed to examine the interrelationship of the major world stock returns.

This paper aims at forecasting stock returns in emerging markets using their interrelations to other stock exchanges including world leaders and counterparts in their regions. For that purpose first, we examine international co-movements in stock prices by employing the Engle–Granger (1987) two-step cointegration technique. We determine the intra- and intercontinental co-movements of stock prices and group the national markets accordingly. Next, we forecast each national stock market according to the lead–lag structures and the transmission between the markets. Forecast performance of the error correction model (ECM) and vector autoregressive models (VAR) will be compared for all national markets.

Emerging Markets

The emerging markets are characterized by high returns and accompanying high volatilities (Harvey 1991). The growth rates in many emerging-market countries are higher than the growth rates of the economies of their developed counterparts (Greenwood 1993). Therefore, despite high volatilities, risk-adjusted returns may still be higher in emerging markets than in mature markets. Some authors even argue that investing in emerging markets can actually lead to lower portfolio risk for the global investor due to them being relatively uncorrelated with each other and the mature markets (Divecha et al. 1992).

Research on the linkages between national markets has been increasing extensively in recent years (Claessens 1995; Harvey 1995; Ma 1993). The issues that have been investigated are broad. A number of studies have examined co-movements in stock returns with reference to the expected return and diversification benefits of emerging-market investments (Harvey 1991; Wilcox 1992). A second group of studies examined the transmission of global shocks and the international spillover effects of specific news.

Another popular topic is inter-temporal stability, and research in this area is based on data sets including a limited number of countries. Cheung and Ho (1991) for example, investigated four Asia–Pacific countries, and Sinclair et al. (1997) examined nine large emerging markets for the stability of inter-temporal covariances between returns and demonstrated that they are quite unstable.

In a previous study, Meric and Meric (1989) provided empirical evidence to show that there was inter-temporal stability in the long-term co-movements of international stock markets before 1987. Meric and Meric (1996, 1997) provide new empirical evidence to show that inter-temporal stability in the long-term co-movement patterns of the world's major stock markets and the twelve largest European equity markets have changed significantly after the 1987 international equity market crash. Box's (1949) M statistic is used to test the long-term inter-temporal stability of the correlation matrix of the stock market index returns, and principal components analysis is used to study the long-term co-movement patterns of the stock markets. Co-movements between major and emerging market stock prices around the 1987 crash reveal a relationship between foreign-entry barriers and stock price transmission. For most countries, individual market return volatility and price spillovers among markets increase immediately after the crash. However, in markets with stiff entry barriers, volatility rises, but there are no price spillovers. The evidence that several emerging-market countries are poorly integrated financially with the industrialized countries (Rogers 1994).

Time-series analysis of the international co-movements of the stock markets (Shin 1993) is also another research area. In Jeon and von-Furstenberg (1990), the interrelationships among stock prices in major world stock exchanges have been investigated by applying the VAR approach to daily stock price indexes in Tokyo, Frankfurt, London, and New York for the period January 1986 through November

1988. Evidence of a significant structural change, with regard to the correlation structure and leadership, was found in the major world stock markets since the stock market crash of October 1987. The impulse response function analysis showed that the degree of international co-movements in stock price indexes has increased significantly since the crash.

Chaudhuri (1997) investigates the common trends in seven Asian markets by using the Johansen cointegration methodology and reports a single common trend. The issues addressed in the Cashin et al. (1995) study are closest to those investigated in this paper. They use the cointegration tests to assess the extent to which equity prices move similarly across countries and regions. Cashin et al. use seven industrial and six emerging-markets' data in weekly frequency for the six-year period from 1989 to 1995. They report increased integration of emerging equity markets since the beginning of 1990 via greater regionalization of national stock markets. Besides, if national stock markets are subject to a global shock that causes them to deviate from their long-run equilibrium relationship, it takes several months for the long-run relationship to reassert itself.

In this paper, our focus is forecasting stock returns in emerging markets using their interrelations to major world stock exchanges and regional counterparts. We examine international co-movements in stock prices as a basis of determining the intra- and intercontinental co-movements of stock prices and grouping the national markets accordingly. We forecast each national stock market according to the lead–lag structures and the transmission between the markets using ECM and VAR. In this framework, the New York, London, and Tokyo stock exchanges will be used to represent the world leaders. Stock returns of the sixteen emerging markets from different geographical locations are forecasted according to their inter- and intracontinental co-movements.

Data

The empirical analysis presented below is based on the stock returns of 16 emerging markets from three continents and three world leaders from those continents. Data is compiled from Data-Stream. London (FTSE All Share), Tokyo (NIKKEI 225), and New York (S&P 500) represent the leading stock markets in Europe, Asia, and the United States, respectively. The IFC indexes are used for the emerging markets from Europe, Asia, and the United States. European markets comprise Greece, Turkey, and Portugal; Asian markets comprise Jordan and India; Far Eastern markets comprise Korea, Malaysia, Philippines, Taiwan, and Thailand; and Latin American markets comprise Argentina, Brazil, Chile, Columbia, Venezuela, and Mexico. Except for London, all of the time series contain 475 weekly observations that cover the period between December 29, 1988 through January 29, 1998. The FTSE All Share index contains 390 weekly observations that cover the period September 8, 1990, through January 29, 1998.

The readers might note that we used the main national price indexes for the world leaders, and we used the IFC price indexes for the emerging markets. Our choice is based on the premise that the IFC provides a consistent dollar-based series, which is comparable across the countries, and also being highly correlated with the national indexes. Summary statistics about the data are presented in Table 1.

Table 1 reports the mean weekly return calculated as the log differences of the national indexes and the standard deviation of returns for the nineteen national indexes that constitute the sample of this study. The third and fourth moments are also given as the skewness and kurtosis coefficients. Six out of sixteen emerging markets have negative mean weekly returns. Standard deviations of emerging markets' returns are considerably higher than those in world leaders. Coefficient of variations up to 2,290 (Philippines) and 385 (Taiwan) are observed besides a minimum of nine (Chile and Columbia). Nine out of nineteen national stock returns have skewness coefficients of less than –0.5, indicating negative skewness. These indexes are from Jordan, Argentina, Brazil, Venezuela, Mexico, Malaysia, Philippines, Taiwan, and Thailand. As expected in most financial series, twelve of the nineteen national stock returns have kurtosis coefficients greater than 3 indicating leptokurtosis. These countries are Jordan, Argentina, Brazil, Columbia, Venezuela, Mexico, Korea, Malaysia, Philippines, Taiwan, Thailand, and Turkey. However, the Jarque–Bera (1980) test for normality indicates that all of the return series deviate significantly from normality, probably due to excess kurtosis and skewness.

Stochastic properties of the time series is investigated for each of the national stock markets by applying the Augmented Dickey–Fuller (ADF) unit root test (Dickey and Fuller 1981) at levels and first differences. ADF values for each national stock market are calculated by estimating regression equations for a random walk, a random walk with drift, and a random walk with drift and trend, respectively. For each estimation, Hsiao's (1981) final prediction error (FPE) model selection criteria is examined at lag lengths of one to four, and the one with the smallest FPE is selected. In all cases, the national stock returns have unit roots in levels, that is, they are not I(0) at 5 percent significance. However, the ADF test applied on the first differenced series does not exhibit a unit root, that is, are I(1) at one percent significance in all specifications. Table 2 reports ADF test results using Fuller's (1976) critical values for the I(0) and I(1) series, with the constant and trend specification. The possible existence of a long-run relationship between the non-stationary national stock prices indexes within each region and with the world leaders, can thus be tested by using the cointegration technique developed by Engle and Granger (1987) in the next region.

Cointegration Analysis

If x_t denotes an nx1 vector, and each of the national stock price series in x_t are I(d), and there exists an nx1 vector such that $x_t' \sim I$ (d-b), then $x_t' \sim CI$ (d,b), where a is called the *cointegrating vector.* An arbitrary linear combination of nonstationary

Table 1

Descriptive Statistics

Country	Mean	Standard deviations	Skewness	Kurtosis	Normality chi-squared
Jordan	0.001026	0.025241	−0.888636	11.174727	427.09 [0.0000]**
India	0.000478	0.040522	0.055362	2.533522	78.911 [0.0000]**
S&P 500	0.002661	0.016922	−0.184535	0.970039	16.794 [0.0002]**
Nikkei	−0.000513	0.028224	0.034878	2.640599	84.191 [0.0000]**
Argentina	0.004677	0.088899	−0.856808	12.967724	544.93 [0.0000]**
Brazil	0.003266	0.077766	−0.627736	3.360443	81.545 [0.0000]**
Chile	0.003450	0.030448	0.126374	1.247410	25.426 [0.0000]**
Columbia	0.004118	0.036064	1.221001	6.319911	110.52 [0.0000]**
Venezuela	0.002943	0.063775	−2.605020	32.152272	509.57 [0.0000]**
Mexico	0.003195	0.040020	−1.232360	7.878077	158.32 [0.0000]**
Korea	−0.002512	0.061128	−0.419486	25.649395	1,473.9 [0.0000]**

Malaysia	−0.000589	0.039078	−1.708737	11.504589	170.28 [0.0000]**
Philippines	0.000019	0.043527	−1.464878	8.447711	133.54 [0.0000]**
Taiwan	−0.000137	0.052815	−0.516211	5.690212	208.43 [0.0000]**
Thailand	−0.001195	0.052887	−0.509109	6.152581	235.19 [0.0000]**
Greece	−0.001461	0.039102	0.304948	2.610894	62.748 [0.0000]**
Turkey	−0.001096	0.075866	−0.394017	3.264223	82.894 [0.0000]**
Portugal	0.001489	0.025590	−0.165947	1.209664	20.317 [0.0000]**
London	0.002182	0.019014	0.260660	1.716535	33.612 [0.0000]**

Notes: (1) This table reports the descriptive statistics of the weekly stock returns of each country in the sample. (2) Stock returns are calculated as the first differences of the logarithm of national indexes representing continuously compounded returns. (3) The first four columns report the weekly mean return, its standard deviation, skewness and the kurtosis. Normality is tested by the Jarqua-Bera (1980) test for normality and p-values are obtained from the Chi-squared distribution with two degrees of freedom. (4) [*] Denotes significant at 5 percent and [**] denotes significant at 1 percent.

Table 2

Results of the ADF tests

Countries	I(0)	I(1)
Greece	–2.5709(2)	–10.657(2)**
Turkey	–3.0387(4)	–10.361(2)**
Portugal	–1.3545(4)	–9.9646(2)**
New York	–1.1529(1)	–12.274(1)**
Tokyo	–1.5573(1)	–9.8675(4)**
London	0.1817(4)	–11.193(4)**
Argentina	–2.8946(3)	–10.657(2)**
Brazil	–2.777(1)	–11.656(2)**
Chile	–0.3081(1)	–9.9646(2)**
Columbia	–0.55795(1)	–12.274(1)**
Venezuela	–1.3364(1)	–9.8675(3)**
Mexico	–2.0901(2)	–11.462(1)**
Korea	0.9256(1)	–11.716(2)**
Malaysia	–0.0429(4)	–8.92(3)**
Philippines	–1.2096(2)	–13.762(1)**
Taiwan	–2.4054(3)	–13.208(1)**
Thailand	0.8521(3)	–9.5485(3)**
Jordan	–2.6426(1)	–17.045(1)**
India	–2.3039(4)	–8.9811(3)**

Notes: (1) ADF test statistics reported here are based on regressions with constant and trend specification. (2) Each ADF regressions, initially includes four lagged differences to ensure that the residuals are empirically white noise. Then a sequential reduction procedure is applied to eliminate the insignificant lagged differences. Values in parentheses show the optimum number of lags used according to the FPE criteria. (3) [*] Denotes ADF test statistics significant at 5 percent and [**] significant at 1 percent.

time series $(y_t - ax_t)$ is expected to be nonstationary. However, if these series are cointegrated, a may take a value, such that $(y_t - x_t)$ is I(0), indicating a stationary relationship between the variables. The null hypothesis of no cointegration (against the alternative of cointegration) is tested using the Engle and Granger (1987) two-step procedure and the following equations:

$$y_t = \beta x_t + u_t \tag{1}$$

$$\Delta u_t = \delta u_{t-1} + \Sigma\delta_i \Delta_{t-i} + \varepsilon_t \tag{2}$$

The first step of this procedure involves regressing the log levels of the national stock price indexes on each other to obtain the ordinary least squares (OLS) regression residuals. Four lagged-difference terms are also used in this process. The second step is to test the existence of unit roots (that is, no cointegration) in the OLS residuals using the ADF test. The results of ADF test statistics on cointegrating regressions (without constant and trend specification) are presented in Table 3, both for all country pairs. The appropriate critical values are obtained from Engle and Granger (1987).

After establishing in the previous section that all of the individual time series are from the same data generating process, that is, same order of integration, we can proceed to test if the national equity market indexes form a cointegrating relationship with a stationary error term. We examine the long-run co-movements among the national equity markets by grouping them according to their geographical proximity. In Table 3, cointegration results are reported for the world leaders alone, for the European country markets, for the Latin American markets, for the Far Eastern markets, and finally for the Asian markets. For all groups based on regional proximity, the world leaders are also included.

Panel 1 of Table 3 shows that equity markets of the world leaders, namely New York, Tokyo, and London, are highly integrated. Panels 2 through 5 of Table 3 show that all of the national markets are integrated on a regional basis, as well as being integrated with the world leaders. This evidence is different from those of previous studies indicating low correlations among international returns (Divecha et al. 1992; Speidell and Sappanfield 1992). Still, we must mention that in previous studies evidence is reported for increased integration among national stock markets through time (Cashin et al. 1995; Choudry 1997).

The findings reported in Table 3 show that the emerging equity markets are linked to world markets and other emerging markets in their region through inter- and intraregional equilibrium relationships. On one hand, the results indicate that benefits from international portfolio diversification are no longer valid. On the other hand, they indicate that shocks to world leaders can affect emerging equity markets over the long run. Also, shocks to one emerging market can affect other equity markets in the same region.

The cointegration results presented above have important implications for the global investor. Accounting for the information embodied in the long-run equilibrium relationship, short-run dynamics can be examined to see the process by which the national indexes return to their equilibrium states. Thus, in today's global world, the national stock returns can be forecasted by using the error correction mechanisms implied by the cointegrating relationships. The short-run interaction between the national stock markets in a regional context and the interaction between the national stock markets and the world leaders can be used for improving forecasts regarding national equity markets.

Forecasting Using ECM and VAR Models

After the cointegrating vectors are determined in the previous section, first, an ECM that embodies both the short-run dynamics and the long run constraint is used to produce forecasts of the national stock returns. The significant inter- and

Table 3

Results of Cointegration Tests

Panel 1: World Leaders

Countries	Tokyo	London
New York	–9.0864 (4)**	–16.947(1)**
Tokyo	—	–13.047(1)**

Panel 2: Europe

Countries	E	Turkey	Portugal
Greece	—	—	—
Turkey	–12.785**(3)	—	—
Portugal	–13.534**(1)	–12.544**(2)	—
New York	–13.232**(1)	–13.076**(1)	–13.907**(1)
Tokyo	–13.096**(1)	–13.126**(1)	–13.515**(1)
London	–12.079**(1)	–12.204**(1)	–7.1912**(4)

Panel 3: Latin America

Countries	Argentina	Brazil	Chile	Columbia	Venezuela	Mexico
Argentina	—	—	—	—	—	
Brazil	−11.118**(3)	—	—	—	—	
Chile	−10.583**(2)	−9.9522**(4)	—	—	—	
Columbia	−10.660**(2)	−14.103**(1)	−12.639**(1)	—	—	—
Venezuela	−10.776**(2)	−11.572**(2)	−13.306**(1)	−9.6610**(3)	—	—
Mexico	−11.3877**(1)	−9.2654**(3)	−12.199**(1)	−11.500**(1)	−12.526**(1)	—
New York	−14.923**(1)	−14.893**(1)	−14.656**(1)	−14.420**(1)	−14.430**(1)	−14.259**(1)
Tokyo	−11.100**(1)	−10.888**(2)	−11.207**(2)	−13.329**(1)	−9.8821**(3)	−13.137**(1)
London	−8.9260**(4)	−14.692**(1)	−9.5027**(2)	−11.536**(1)	−9.3653**(3)	−10.456**(1)

Panel 4: Far East

Countries	Korea	Malaysia	Philippines	Taiwan	Thailand
Korea	—	—	—	—	
Malaysia	−11.662**(2)	—	—	—	
Philippines	−11.555**(2)	−9.374**(4)	—	—	—
Taiwan	−11.955**(2)	−8.6148**(3)	−8.4684**(4)	—	—
Thailand	−11.594**(2)	−17.619**(1)	−9.1559**(4)	−13.043**(1)	
New York	−11.423**(2)	−8.8706**(3)	−8.3718**(4)	−13.328**(1)	−9.9156**(3)
Tokyo	−11.474**(2)	−9.1122** (3)	−14.306**(1)	−13.231**(1)	−9.5084**(3)
London	−8.368 **(3)	−7.1761**(3)	−9.8751**(2)	−9.9060**(2)	−8.0590**(3)

(continues)

Table 3 (continued)

Results of Cointegration Tests

Panel 5: Asia

Countries	Jordan	India
Jordan	—	—
India	−11.289**(2)	—
New York	−10.458**(4)	−8.9117**(2)
Tokyo	−9.1952**(4)	−8.8143**(3)
London	−16.005**(1)	−8.0838**(3)

Notes: (1) The values reported here are the ADF test statistics based on regressions without constant and trend. (2) Each regression initially includes four lagged differences to ensure that the residuals are empirically white noise. Then a sequential procedure is applied to eliminate the insignificant lagged differences. Values in parentheses show the optimum number of lags according to the FPE criterion. (3) Critical values of the ADF test statistics are obtained from Engle and Granger (1987). (4) [*] Denotes ADF test statistics significant at 5 percent and [**] significant at 1 percent.

intracontinental co-movements reported in the previous section form the basis of the ECM forecasts presented in Table 4. The (nx1) vector x_t represents the time series of all the national indexes within a continent and has an error correction representation that can be expressed in the form:

$$\Delta x_t = \beta 0 + \beta_1 x_{t-1} + \beta_2 \Delta x_{t-1} + \ldots + \beta_p \Delta x_{t-p} + \lambda_i \Delta y_{t-i} + \varepsilon_t, \quad (3)$$

where β_0 is an (nx1) vector of intercept terms, β_i is an (nxn) coefficient matrix with i = 1...t–p, λ_i is an (nx3) coefficient matrix with i = 1...t–p, y_t is a (3x1) vector of the world leading indexes, ε_t is an (nx1) vector of error terms that are white noise and may be correlated with each other, and Δ stands for first differencing.

Next, national stock returns are forecasted by using vector autoregressions for forecast comparisons. The VAR model has the advantage of not having an underlying theory and does not need any assumptions about the values of the exogenous variables in the forecasting period. The significant inter- and intracontinental co-movements reported in the previous section form the basis of the VAR forecasts presented in Table 4. We employ the following VAR model defined in the first difference form:

$$\Delta x_t = \delta_0 + \delta_1 \Delta x_{t-1} + \delta_2 \Delta x_{t-2} + \ldots + \delta_p \Delta x_{t-p} + \lambda_i \Delta y_{t-i} + \varepsilon_t, \quad (4)$$

where δ_0 is an (nx1) vector of intercept terms, δ_i is an (nxn) coefficient matrix with i = 1...t–p, λ_i is an (nx3) coefficient matrix with i = 1...t–p, y_t is a (3x1) vector of the world leading indexes, ε_t is an (nx1) vector of error terms that are white noise and may be correlated with each other, and Δ stands for first differencing.

Both models are estimated by using the weekly data from the beginning of the sample and the last twenty-six weeks, twelve weeks, and four weeks are used as the out-of-sample period, respectively, to evaluate the forecasting performance of the ECM and VAR models. One-step-ahead forecasts are made on a weekly basis assuming that the forecaster, making a forecast for period t + 1, knows the realized values of the time series at time t. The forecast performance of the two models are evaluated and compared on the basis of parameter constancy and forecast accuracy.

We initially estimated the ECM and VAR models by including four lags of the national returns in each system. A constant and unrestricted world leaders returns are also included in both systems. For optimal lag selection we used the Schwartz (1978) criteria, which pointed to a single lag for all country groups. Then we estimated one-step-ahead forecasts for forecast horizons of twenty-six, twelve, and four weeks, respectively, and tested for parameter constancy. The first measure used for parameter constancy is the V[E] χ^2 (nH) for H forecasts and n equations. It represents the full variance matrix of all forecast errors E, which takes both parameter uncertainty and intercorrelations between forecast errors into account. The second measure employed in this paper for parameter constancy is the Forecast–Chow test based on forecast variance, F(nH, T-k), where T stands for the number of observations and k stands for the number of parameters to be estimated. This test ignores the correlation between forecast errors and is, thus, better cali-

Table 4

Results of ECM and VAR Forecasts

Panel 1: Europe

Countries	ECM	VAR
Greece	−0.0103(0.0342)	−0.0081(0.0365)
Turkey	0.0006(0.0454)	0.0047(0.0452)
Portugal	0.0249(0.0226)	0.0276(0.0228)
V[E] χ^2 (12)	10.533(0.56)	11.703(0.46)
F(12, 368:374)	0.8778(0.57)	0.9753(0.47)

Panel 2: Asia

Countries	ECM	VAR
Jordan	−0.0076(0.0083)	−0.0066(0.0084)
India	−0.0146(0.0414)	−0.0137(0.0421)
V[E] χ^2 (8)	3.9761(0.86)	3.93(0.86)
F(8,370:375)	0.49701(0.85)	0.4913(0.86)

Panel 3: Latin America

Countries	ECM	VAR
Argentina	−0.0273(0.0898)	−0.0226(0.0956)
Brazil	−0.0378(0.0623)	−0.0234(0.0608)
Chile	−0.0327(0.0771)	−0.0338(0.0788)
Columbia	−0.0438(0.0223)	−0.0386(0.0246)
Venezuela	−0.0774(0.0788)	−0.0668(0.0758)
Mexico	−0.0470(0.0589)	−0.0289(0.0603)
V[E] χ^2 (24)	46.975(0.00)**	43.507(0.01)**
F(24,362:371)	1.9573(0.01)**	1.8128(0.01)**

Panel 4: Far East

Countries	ECM	VAR
Korea	0.1227(0.1823)	0.0555(0.2008)
Malaysia	−0.0068(0.1895)	−0.0448(0.1989)
Philippines	−0.0165(0.1957)	−0.0397(0.2074)
Taiwan	0.0056(0.0585)	−0.0154(0.0583)
Thailand	0.1101(0.1857)	0.0181(0.1843)
V[E] χ^2 (20)	151.95(0.00)**	146.74(0.00)**
F(20,364:372)	7.5976(0.00)**	7.3371(0.00)**

Table 4

Results of ECM and VAR Forecasts

Notes: (1) The values reported in each cell are the mean forecast errors, and the values in parentheses are the related standard deviations. (2) The initial ECM and the VAR include four lags of the national returns, an unrestricted constant. World leaders' returns enter the equations unrestrictedly. (3) Our choice of one lag is based on the Schwartz and the Hannan-Quinn criteria, both of which pointed to a single lag for all country groups. (4) Forecast errors and related statistics reported in this table are based on one period ahead of static forecasts for a four-week forecast horizon. (5) V [E] χ^2 (nH) and the Chow's F-test F(nH, T-k) reported at the last two rows of each panel measure parameter constancy. There, n is the number of equations, H is the number of forecasts, T is the number of observations, and k is the number of parameters to be estimated. The table reports F statistics as F(nH, T-k_1:T-k_2), where k_1 is the number of parameters to be estimated by ECM, and k_2 is the number of parameters to be estimated by the VAR. (6) [*] Denotes test statistics significant at 5 percent and [**] significant at 1 percent.

brated (Chong and Hendry 1986). Forecast accuracy is measured by the mean forecast error (MFE), for each national stock return forecast.

Before reporting the results, we checked for the parameter constancy of each system, that is, whether the estimated parameters of the system remain constant during the forecast period as well. Parameter constancy was rejected for each country group and for both the ECM and the VAR models for the twenty-six-week forecast period. For the twelve-week forecast period, parameter constancy was rejected for all country groups except for the Far East. For the four-week forecast period estimated, parameters remained constant for both the Latin American and the Far Eastern markets. Results reported in Table 4 are for the four-week forecast horizon. The mean forecast errors and related standard deviations are reported for each national stock return forecast for both the ECM and the VAR models. The last two rows of each panel contain the parameter constancy test statistics and related *p*-values, for the ECM and the VAR models. Related forecast statistics are supplied by PCFIML.[1]

In panel 1 of Table 4, forecast statistics for European emerging-market stock returns are reported. Although the European emerging markets are cointegrated, neither the ECM nor the VAR forecasts pass the parameter constancy tests. For Turkey and Portugal, mean forecast errors are slightly smaller in ECM forecasts, and for Greece, the VAR model yields slightly better forecast errors. In panel 2 of Table 4, forecast statistics for Asian emerging markets are reported. These markets are also cointegrated. Still, both the ECM and the VAR forecasts fail to pass the parameter constancy tests. For both Jordan and India, mean forecast errors are slightly better in VAR forecasts.

In panel 3 of Table 4, forecast statistics for Latin American emerging markets are reported. For all of the Latin American markets, both ECM and VAR forecasts pass the parameter constancy tests. For all of the Latin American emerging markets, except for Chile, VAR model provides better mean forecast errors. In panel 4 of Table 4, forecast statistics for Far Eastern emerging markets are reported. For all of the Far Eastern markets, both the ECM and the VAR forecasts pass the parameter constancy tests. We must note here that, results not reported in this paper show that for this group of countries, ECM and VAR forecasts pass parameter constancy tests, also, for the twelve-week forecast horizon. For Korea and Thailand, the mean forecast errors are slightly better in VAR forecasts. For Malaysia, Philippines, and Taiwan, ECM forecasts provide better mean forecast errors.

Discussion and Conclusions

This paper attempts to maintain a research framework, whereby a coherent database is used to include all emerging markets as classified by the IFC. Besides, the data frequency is weekly for all countries. The focus of the paper is also different from previous research that investigates market integration for global diversification. In this paper, the degree of market integration is investigated in order to forecast national markets according to their international co-movements.

Before forecasting national stock returns, we first considered the descriptive statistics and the stationarity of national stock indexes. None of the stock return series could pass the normality tests, mainly due to leptokurtosis. Unconditional variances were much higher for emerging markets than their mature counterparts. The examination of the time-series properties of the national stock indexes revealed that all of the stock return series were stationary.

After determining that all of the stock index series are I(1), we conducted cointegration tests for all emerging markets with the world leaders and their regional counterparts. The results reveal that all national markets are cointegrated with the world leaders and with other emerging markets grouped according to their geographical proximity. This result is important in terms of its implications for the global investor. Accounting for the information embodied in the long-run equilibrium relationship, short-run dynamics can be examined to see the process by which the national indexes return to their equilibrium states.

Thus, the final step was to forecast national returns using the interaction between the national stock markets in a regional context and the interaction between the national stock markets and the world leaders. For that purpose, we used ECM and VAR models to forecast national markets. The results of the forecasting exercise were not very promising. For longer forecast horizons, none of the models could pass the parameter constancy tests. For shorter forecast horizons, only the Latin American and the Far Eastern markets could pass the parameter constancy tests. For those countries, mixed results were obtained as to better forecast errors from ECM and VAR models.

Latin American and Far Eastern markets have distinguishing characteristics among the emerging markets. They are more established in the sense that international awareness about those markets are high, and they have been attracting international investors for a longer time period. Also, in terms of listed companies, trading volumes, and market capitalization, these stock markets are in better terms than their counterparts in Europe and Asia. It can be argued that they are, thus, better integrated with the world and with each other, in terms of information and capital flows. Therefore their behavior could be better forecasted.

The implications of the poor parameter constancy performance of our models are various. First, it might be the case that both the VAR and the ECM were insufficient to model the underlying process. In this case, naive models such as ARIMA and various forms of single equation estimations, rather than system solutions, should be employed to see if better forecasts could be achieved. Next, the basic characteristics of emerging markets must be considered to improve forecasts of national returns. Emerging markets are characterized by rapid change (Muradoglu and Metin 1996) and high volatilities (Harvey 1991). Structural breaks must be determined for each national market, and the sample periods for forecasting must be specified accordingly. In order to incorporate high volatilities in emerging markets, conditional volatilities could also be incorporated into the mean equations via GARCH-M type models.

Note

1. PCFIML is a full-information maximum likelihood estimation package, which is developed by Doornik and Hendry (1997).

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