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International Integration of Equity Markets and Contagion Effects

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Abstract

This paper investigates empirically the degree of international integration of industrial and emerging country equity markets. It analyzes two issues: first, the extent to which equity prices have tended to move similarly across countries and regions in the long run; and second, the strength of cross-country "contagion" effects. The paper's findings suggest that both intra-regional and inter-regional linkages across national equity markets have strengthened in recent years. In addition, using impulse response functions, the paper shows that cross-country contagion effects of country-specific shocks dissipate in a matter of weeks while contagion effects of global shocks take several months to unwind themselves.

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Contents

Summary

This paper investigates empirically the degree of international integration of industrial and emerging equity markets as well as changes in the extent of this integration over time. It analyzes two key issues. First, the extent to which equity prices have tended to move similarly across countries and regions in the long run is analyzed using the Johansen cointegration methodology. Second, the strength of "contagion" effects is analyzed using deviations from the long-run equilibrium, and the paper presents estimates of the time taken for adjustment to these historical norms following global and country-specific shocks to national equity markets. The analysis is based on weekly data for the period January 1989-March 1995 for seven industrial country markets (United States, United Kingdom, Japan, France, Germany, Spain, and Australia) and six emerging markets (Brazil, Mexico, Korea, Malaysia, Thailand, and Jordan).

The paper first discusses some of the key issues in assessing the degree of integration of national stock markets and summarizes the results of the existing empirical studies. It then examines the structural characteristics of the equity markets in the sample and illustrates the degree of comovement in equity prices. Unit root tests of the national equity price indices are discussed and the results of multivariate cointegration analyses presented. The analysis is undertaken for the full period and a number of subperiods and for all countries as well as for countries included in various regional groupings. The paper also presents the results of an exercise to assess the strength of short-run contagion effects of country-specific shocks on emerging equity markets.

Unit root tests indicate that all national equity market indices follow a random walk, and hence a cointegration methodology is required to analyze comovements in these indices. The results from the cointegration approach suggest that the international integration of emerging equity markets has increased since the early 1990s. However, markets in industrial countries were already largely integrated at the start of the sample period. While there is a clear indication that the increasing integration of equity markets has occurred through greater regionalization of national stock markets, involving stronger linkages between emerging and industrial country markets sharing a common geographic region, cross-regional links between emerging and industrial equity markets have also strengthened. An examination of the short-run interaction between national market indices shows that cross-country contagion effects of country-specific shocks dissipate in a matter of weeks. However, 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 this relationship to reassert itself.

I. Introduction

This paper examines the degree of international integration of industrial and emerging country equity markets, and changes in this integration over time. Specifically, the paper analyzes the effect that the easing of controls on international portfolio capital flows (particularly into emerging markets), financial sector deregulation, and the relaxation of exchange controls since the late 1980s have had on the extent of equity market integration. It utilizes data on portfolio flows to individual countries to delineate a number of different breakpoints, and examines the differences in integration of national equity markets between the "control" and "post-control" subperiods.

The issues analyzed in this paper form part of a broader set of questions which have been examined extensively in the literature on integration of world capital markets, particularly whether or not financial assets issued in different countries yield the same risk-adjusted returns. If the yields differ, the issue is the extent to which these differences reflect controls on capital flows, or other informational and structural imperfections. To the extent that national equity markets are integrated, this has a bearing on investor perceptions concerning developments and policy changes in overseas markets, which in turn would affect the movement of portfolio capital. The degree of integration also has implications for the global cost of capital, the benefits of international asset diversification, and the likelihood of national financial disturbances spilling across international borders. 1/

While there have been a number of recent studies assessing integration among industrial country stock markets, the analysis of interdependence between emerging and industrial country markets has been limited. Moreover, there have been very few studies which have examined whether the degree of integration has been affected by the sharp increase in portfolio capital flows to emerging markets in recent years. In addition to examining these issues, this paper analyzes the extent to which any increased integration may have occurred on the basis of a regional grouping of industrial and emerging markets. For instance, it examines whether Latin American equity markets are relatively more integrated with the U.S. equity market, and whether Asian emerging equity markets are relatively more integrated with the Japanese equity market.

The above issues are examined empirically on the premise, well established in the literature, that the greater the international integration of equity markets, the higher the degree of correlation among national equity prices. The paper uses cointegration methodology, in particular the Johansen (1988) cointegration tests, to assess the extent to which equity prices have tended to move similarly across countries and regions in the long run. The assessment as to whether national equity

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1/ For a detailed survey of these issues, see Goldstein and Mussa (1993).

prices are cointegrated is equivalent to testing whether there are linear combinations of these indices which will converge to stationary long-run equilibrium relationships. Furthermore, using deviations from the long-run historical averages of the stock market price ratios--estimated from the cointegrating relationships--the paper calculates the time taken for adjustments to these historical norms to take place following a change in the equilibrium relationship. In addition, the paper analyzes the short-run interaction between national equity prices in a regional context, thus providing some indication of the strength of "contagion" effects.

The rest of the paper is organized as follows. Section II provides a discussion of the key issues in assessing the degree of integration of national stock markets, and summarizes the results of existing studies. Section III discusses the data, examines the structural characteristics of equity markets in our sample, and illustrates the degree of comovement in equity prices. It also notes the extent of portfolio capital flows to emerging markets between 1989-95, and delineates the time periods over which the flows appear to have increased sharply. This is followed in Section IV by a discussion of unit root tests of whether or not the national equity price series follow a random walk. Given that the equity price series are seen to follow a random walk, Section V reports the results of multivariate cointegration analyses for the full period, and for each of the two (control and post-control) subperiods. Section VI discusses the results of an exercise to assess the short-run "contagion" effects of global and country-specific shocks to national equity prices. The concluding section summarizes the main findings of the paper, and discusses their policy implications.

II. International Equity Market Integration: Conceptual Issues and Existing Studies

1. Key issues and methodological approaches

The issue of the linkages between national equity markets has been examined extensively in recent years, mainly with reference to industrial country markets (see, for instance, Jorian 1989, Obstfeld 1993, and Frankel 1994) but also increasingly with reference to emerging markets (Claessens 1995). As noted above, the empirical evidence on this subject has been generally in terms of the correlation of equity prices and returns across various groups of countries, and three distinct types of issues have been investigated.

First, studies have tried to examine the short- or long-run relationship between national equity markets (see, for example, Jorion 1989, Dwyer and Hafer 1988, Bekaert 1995 for the long-run, and Hamao et.al. 1990, for the short run). A sub-set of these studies has examined the effect of policy changes on the degree of integration (Taylor and Tonks 1989, Ma 1993, Harris and Smith 1995), or of the effect of particular economic events (Tang and Mak 1995).

Second, a number of studies have examined the transmission of particular shocks (such as the 1987 U.S. crash, and the mini-U.S. crash of October 1989) across international borders. The main issue here has been the extent and duration of "contagion" effects, where a shock in one national market is transmitted across national borders. A related aspect has been the spillover effects across international financial markets of specific new information or "news".

Third, a number of studies have examined whether comovement in equity prices across countries is sufficiently low to allow portfolio diversification benefits. There was an extensive analysis of this issue for industrial country markets in the 1970s and early 1980s (see Jorian 1989). In recent years, this issue has been examined with particular reference to the expected return and diversification benefits of investing in emerging markets (Divecha et.al. 1992, Harvey 1993, and Wilcox 1992).

There have been a variety of different model-based approaches to testing and measuring the degree of equity market integration. One approach assumes that markets are integrated and that a particular asset-pricing model holds (Campbell and Hamao 1992). This approach is hampered by the lack of a generally accepted international asset-pricing model, especially in view of the evidence on the intertemporal variation in expected excess returns. There is also the problem of joint hypothesis testing, whereby any test of the extent of market integration would be testing both the applicability of a particular asset pricing model as well as the degree of integration. Another approach explicitly models the existence of barriers to integration, and derives their effects on equilibrium returns (Cooper and Kaplanis 1994, Hietala 1989). As Bekaert (1995) suggests, this approach is restrictive in that it invariably limits the analysis to the effects of one particular barrier to investment.

There have also been a variety of statistical and econometric methodologies utilized for the empirical analysis. These range from: the computation of correlation coefficients between stock index returns across national stock markets; principal components analysis with the first component identified with common global factors; the application of statistical techniques to analyze the intertemporal variation in conditional variances (such as the ARCH processes); and the use of cointegration techniques.

2. Review of existing studies

The remainder of this section summarizes the methodology and results of a number of recent studies focussing on the first two broad issues noted above (the relationships among national equity markets and the international transmission of shocks), to provide a perspective on the empirical analysis undertaken in Sections IV to VI below.

Jorion (1989) presented correlation coefficients for 16 industrial country equity price indices for three periods (1959-70, 1971-78 and 1979-

86), using monthly data. 1/ He found that these correlations were relatively low--around 0.30--and that the magnitude actually decreased between 1971-78 and 1979-86. Given that a number of barriers to the movement of capital were lifted in the latter period, notably in the U.K. and Japan, it may appear surprising that this did not lead to closer movements among equity markets. A specific examination of the correlation over time between the U.S. and Japanese markets revealed virtual stability in the correlation coefficients.

In contrast to the span of the above study, von Furstenberg and Jeon (1989) analyze movements in the equity price indices of Japan, U.S., U.K., and Germany between 1986-88. Using principal components analysis of daily rates of changes, and assuming that the first and largest component can be identified with the most global of disturbances, the authors conclude that the global factor was consistently larger after the October 1987 crash than before. That is, after October 1987 the importance of country-specific shocks was shown to have declined relative to shocks common to all countries. In addition, using vector autoregression analysis to obtain impulse response functions, the authors show that in the post-crash period all markets had become more sensitive to innovations in Japan, but not more sensitive to innovations in the U.S. However, results before and after the 1987 crash show little difference in the degree to which foreign markets retained Japanese and U.S. stock market innovations. Developments in the U.K. index did have a consistently greater and more lasting effect on foreign markets after the October 1987 crash than before. This was attributed to the expansion and internationalization of the London market, which started the year prior to the crash.

Hamao et.al. (1991) also look at the effects of the October 1987 crash. They use a modified autoregressive conditional heteroskedasticity (ARCH) framework to assess volatility spillover effects in intraday returns among New York, Tokyo and London markets. Their data cover the five-year period from April 1, 1985 to February 28, 1990, and encompasses daily open and closing prices of major market indices on the three exchanges. They conclude that the international transmission of volatility does not occur evenly across the three markets--there were spillover effects of disproportionate size from one market to the next. They also find that volatility spillovers were relatively stable over the five-year period, even when the crash and post-crash periods are separated. However, there was

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1/ He also provides a detailed reference to earlier studies which examined the relationship between national equity markets.

some weak evidence that volatility spillover effects originating in Japan had become more pronounced following the 1987 crash. 1/

The studies which come closest to the basic issues being investigated in this paper are by Ng et.al. (1991) and Chou et.al. (1994). The first of these studies investigated the relationship between cross-country stock investment and volatility spillovers among five national stock markets including Japan, Korea, Taiwan, Thailand, and the U.S. The study, again using the ARCH model and daily stock returns for the five countries, finds that cross-country stock investment seems to be an important channel for the transmission of volatility across national stock markets. The four Pacific Basin countries all have substantial merchandise trade with the U.S., and therefore they would be expected to be affected by changes in U.S. market fundamentals. But the study finds no volatility spillover from the U.S. to Korea and Taiwan--markets with the most severe restrictions on cross-country investing. The authors conclude that this contradicts the contagion hypothesis, because if market fundamentals in the U.S., Korea and Taiwan are correlated, then prices in the New York market could convey information to these markets even without cross-country trading. The study also finds that the spillovers from the U.S. to Japan and Thailand increased substantially after institutional arrangements had been altered to facilitate cross-country investing in the latter two markets. It concludes that cross-country trading itself is a necessary condition for the spillover of market volatility, yet information received about the U.S. market is not sufficient to induce volatility spillover.

The study by Chou et.al. (1994) uses weekly data on stock prices for six industrial countries (U.S., Canada, U.K., France, Germany, and Japan) to examine the degree of comovement in prices over the period July 1976 to December 1989, and for various subperiods. The authors use both local currency price series and prices in U.S. dollars. Applying the multivariate cointegration test of Johansen (1988), they find that there are one to three cointegrating vectors in the six stock price indices, indicating long-run equilibrium relationships among the equity prices. Using subsets of countries, the authors also find that equity prices of the three European countries are cointegrated, and that equity prices of the U.S., Canada, and

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1/ Bertero and Mayer (1990) also analyze the effect of the 1987 stock market crash, mainly on industrial country equity markets. They find that the cross-country price correlations after the crash were in general higher than before the crash. Tang and Mak (1995) compare the degree of equity market integration amongst six industrial countries before and after the 1987 crash, and show that the correlation between stock market returns were significantly higher after the crash. There is a paucity of studies on the effect of the crash on integration among emerging equity markets, but Ekechi (1990) finds there was a positive and rising speculative bubble premium in many emerging markets before the 1987 crash.

Japan are also cointegrated. 1/ Finally, Harris and Smith (1995) use cointegration methodology to show that European equity markets are more highly integrated in the 1990s than a decade or more earlier. 2/ They also find that, despite Germany's significance for monetary and economic union, increased linkages between European equity markets have been associated with the U.K., rather than Germany, playing a more central role.

III. Data, and Structural Characteristics of Markets

The empirical analysis covers the six-year period from January 1989 to March 1995, and includes weekly closing prices of major market indices for seven industrial country markets and six emerging markets. The industrial country markets comprise the U.S., Japan, U.K., France, Germany, Spain and Australia, while the emerging markets comprise Brazil, Mexico, Korea, Malaysia, Thailand and Jordan. There were a number of different factors which determined the choice of countries in our samples: we wanted to analyze markets which varied in size (as measured by capitalization), trading activity and in location; and we wanted to include markets which have been particularly volatile in recent years.

It should be noted that for equity markets in industrial countries, we used the main national price indices, while for emerging markets we used the IFC price indices. The reason for this is chiefly that for emerging markets, unlike some of the national indices, the IFC provides a consistent series which can be regarded as comparable across countries, although the degree of comovement in the national and IFC indices tends to be very high. For the U.S. market the Standard & Poor's 500 composite index was used, while for the U.K. market we used the Financial Times 100 share index--both are equity value-weighted arithmetic indices, and the data for both indices were obtained from Bloomberg. For the Japanese market we used the Nikkei 225 Stock index, which is a price-weighted average stock price index. Data for markets in other industrial countries were obtained from Reuters equity price databanks.

Looking at the main characteristics of these markets, the U.S. is clearly the largest among our sample of 13 markets (Table 1). At end-1994 it had a total market capitalization of over US$ 5,000 billion, compared to US$ 3,700 billion for Japan and US$ 1,200 for the U.K. It is noticeable,

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1/ Additional studies of equity market integration, using other methodological approaches, have been carried out by King et.al. (1993), Ferson and Harvey (1993), Bekaert (1995), Kasa (1995) and Obstfeld (1993).

2/ It is noted that capital market integration was not high on the agenda of the European Community until the 1980s when agreement was reached on several policy measures to promote it, particularly the agreement to abolish foreign exchange controls on capital movements by July 1990 (with longer periods of adjustment permitted for four countries with small capital markets).

Table 1. Size of Equity Markets (End-Period)

	1988		1993		1994	
	Billions of US$	% of GDP	Billions of US$	% of GDP	Billions of US$	% of GDP
Emerging Markets						
Brazil	32.1	9.7	99.4	20.6	189.3	n.a.
Mexico	13.8	8.4	200.7	55.5	130.3	55.5
Korea	94.2	49.1	139.4	42.1	191.8	49.9
Malaysia	23.3	73.8	220.3	341.9	199.3	289.5
Thailand	8.8	14.5	130.5	105.5	131.5	91.6
Jordan	2.2	48.2	4.9	94.4	4.6	78.3
Industrial Markets						
France	244.8	26.0	456.1	36.4	451.3	32.9
Germany	251.8	21.3	463.5	26.9	470.5	24.7
Spain	91.1	26.0	119.3	28.2	154.9	31.5
United Kingdom	771.2	89.5	1151.6	121.7	1210.2	114.0
Australia	138.3	52.2	204.0	74.9	219.2	65.9
Japan	3906.7	131.6	3001.4	71.2	3719.9	78.4
United States	2793.8	56.9	5223.8	82.4	5081.8	75.5

Source: IFC (1995) and IMF (1994); n.a. denotes data not available.

however, that all industrial countries except Japan had a marked increase in capitalization over the sample period. In Japan, the collapse in land prices, flat economic activity, and serious weaknesses in the financial sector led to sharp fall in equity prices after 1988, which largely accounts for the observed decline in market capitalization.

Although the emerging markets were small compared to the industrial country markets at the beginning of the period, they have increased dramatically in capitalization in recent years. 1/ Prior to the weak activity in many of these markets (especially Mexico and Malaysia) during 1994, the capitalization of several of the markets had increased by a factor of 10 or more between 1989 and 1993. Even at end-1994, several of the emerging markets in our sample had capitalization-to-GDP ratios on a par with industrial countries, with the market in Malaysia clearly an outlier among the full sample. A striking feature of Table 1 is that even though the number of companies and trading is much smaller in Jordan than in other cases, the capitalization to GDP ratio is comparable to that in other markets.

Table 2 provides information on the trading volumes and the number of listed companies in each of the 13 countries. As the first column indicates, at the beginning of our sample period, the three largest industrial country markets were also the most active, with the annual trading volume of Japan exceeding markedly the trading volume in the U.S. However, by end-1994, while trading volumes had increased significantly in all other industrial markets, activity had been reduced by more than half in Japan. The sharpest increase in trading volumes occurred in emerging markets, where trading activity increased by a factor of 50 in Malaysia and by a factor of 15 in Mexico and Thailand.

As Table 2 indicates, while in general there is little correlation between capitalization and the number of listed companies across either emerging or industrial country markets, it is noticeable that over the sample period there was a marked increase in the number of listed companies in Asia. It is also clear that the number of listed companies in emerging markets is not that much smaller than in several of the industrial countries, indicating that the average size of listed companies in emerging markets is considerably smaller than in industrial markets.

Next consider changes in the magnitude of portfolio capital flows to emerging equity markets over recent years, in order to delineate different subperiods over which movements in equity prices across countries can be examined. As El-Erian and Kumar (1995) show, there was a marked increase in international portfolio equity flows to emerging markets from 1989

1/ For a discussion of the factors behind the increase in capitalization in emerging equity markets since the mid-1980s, and other issues related to the development of these markets, see Feldman and Kumar (1995).

Table 2. Trading Volume and Listed Companies

(In millions of U.S. dollars)

	1988		1994	
	Trading Volume	Listed Companies	Trading Volume	Listed Companies
Emerging Markets				
Brazil	18.0	589.0	109.5	554.0
Mexico	5.7	203.0	83.0	206.0
Korea	79.2	502.0	286.1	699.0
Malaysia	2.6	238.0	126.5	478.0
Thailand	5.6	141.0	80.2	389.0
Jordan	0.3	106.0	0.6	95.0
Industrial Markets				
France	65.5	646.0	174.3 1/	459.0
Germany	350.3	609.0	460.6	417.0
Spain	25.6	368.0	61.5	379.0
United Kingdom	579.2	2054.0	928.2	2070.0
Australia	37.4	1380.0	94.7	1144.0
Japan	2597.6	1967.0	1121.4	2205.0
United States	1719.7	6680.0	3592.7	7770.0

Source: IFC (1995)

1/ Data are for 1993.

onwards. 1/ For instance, from an annual average of equity flows of around US$ 4,000 million between 1986-88, equity flows to emerging markets jumped to US$ 11,000 million in 1989 and US$13,000 million in 1990, with the bulk of the increase going to Latin American markets (Table 3). It is clear that quite apart from economic fundamentals, such cross-country investment is affected by changes in restrictions on foreign ownership of domestic stocks and in foreign exchange controls. For instance, while many of the emerging markets have abolished all restrictions on foreign ownership of domestic stocks, considerable restrictions remain in a number of countries. 2/

An analysis of the data on portfolio equity flows, and more generally on aggregate capital flows, also suggests that by end-1990 or early 1991 there was a further increase in these flows to emerging markets. Table 3 provides some illustrative data on the current and capital accounts for the Asian and the Latin American regions as a whole. As these data indicate, for both regions there was a sharp increase in the balance on capital account from 1990 onwards. Of course, these capital account data include debt and long-term foreign direct investment, in addition to portfolio equity flows. However, it can also be argued that to the extent the non-portfolio equity flows have increased substantially, they would also have a bearing on the comovement in equity prices in emerging and industrial country markets.

Another relevant issue from the point of view of integration of markets is the source of portfolio equity flows. Here it appears that equity flows from Japan were destined mainly for Asian markets, while flows from the U.S. were more evenly divided between Latin America and Asia (see El-Erian and Kumar 1995). It is also worth noting that in terms of the structure of trade, several of the Asian countries, while having strong links with the U.S., are more closely tied to Japan. Elsewhere, portfolio flows among European countries are also substantial, and, of course, these economies are even more closely tied to each other by trade.

Finally, as an introduction to the econometric analysis provided in the following sections, Table 4 provides correlation coefficients of equity prices for the 13 markets in our sample, over the period January 1989 to March 1995. It shows that while the U.S. and U.K. markets were highly positively correlated, there was a marked negative correlation between the U.S. and Japanese markets. Reflecting in part the remarkable swing in asset prices that occurred in Japan in the late 1980s and early 1990s while U.S. equity prices remained buoyant. This negative correlation was also evident

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1/ For a broader discussion relating to the factors underlying the resumption of portfolio capital flows to developing countries, see El-Erian (1992) and Calvo et..al. (1993).

2/ Cashin and McDermott (1995) examine the influence of restrictions on ownership and foreign exchange movement for the emerging markets of Turkey, Jordan and Pakistan.

Table 3. Capital Flows to Emerging Markets in Asia and Latin America 1/

(In billions of U.S. dollars)

Year	Asia: Balance on goods, services and private transfers	Asia: Balance on capital accounts	Latin America: Balance on goods, services and private transfers	Latin America: Balance on capital accounts
1986	1.2	21.2	-17.8	10.0
1987	18.8	20.7	-12.4	15.0
1988	7.0	3.4	-13.8	12.3
1989	-0.9	9.7	-10.2	14.3
1990	-4.4	24.0	-9.3	25.8
1991	-4.7	44.8	-24.8	42.5
1992	-6.2	32.5	-41.1	61.9
1994	-29.2	65.8	-49.4	60.3
1995	-25.6	58.9	-50.9	60.0

Source: IMF, World Economic Outlook Database

1/ A minus sign indicates a deficit in the pertinent account.

Table 4. Price Index Correlations
(January 1989-March 1995)

France	1.0												
Germany	0.77	1.0											
U.K.	0.71	0.53	1.0										
Japan	-0.16	0.06	-0.62	1.0									
U.S.	0.60	0.37	0.94	-0.75	1.0								
Australia	0.76	0.72	0.86	-0.29	0.75	1.0							
Spain	0.60	0.72	0.39	0.37	0.17	0.69	1.0						
Brazil	0.24	0.45	0.50	-0.23	0.45	0.57	0.38	1.0					
Mexico	0.57	0.42	0.92	-0.76	0.93	0.77	0.20	0.61	1.0				
Jordan	0.69	0.54	0.92	-0.59	0.89	0.81	0.31	0.46	0.85	1.0			
Korea	0.29	0.53	0.31	0.30	0.11	0.55	0.75	0.60	0.18	0.25	1.0		
Malaysia	0.66	0.69	0.91	-0.47	0.81	0.90	0.52	0.67	0.86	0.89	0.50	1.0	
Thailand	0.61	0.66	0.88	-0.49	0.81	0.85	0.46	0.71	0.85	0.85	0.50	0.97	1.0
	France	Germany	U.K.	Japan	U.S.	Australia	Spain	Brazil	Mexico	Jordan	Korea	Malaysia	Thailand

Source: IFC (1995).

between Japan and Australia, which is somewhat surprising given that trade and capital flows between these countries are a large share of their total trade and capital flows. At the same time, while there was a positive correlation between the U.K. and other European markets (France, Germany and Spain), the magnitude of this relationship was considerably weaker than that between the U.S. and U.K. markets. With regard to emerging markets, it appears that markets in Jordan, Mexico, Malaysia and Thailand were highly correlated with the U.S. market, while the correlation of the Brazilian and Korean markets with the U.S. market was quite weak. The finding for Korea may be explained in terms of the relatively high restrictions on foreign equity investment; the finding for Brazil may be explained by the much greater role for domestic factors in determining movements in equity prices.

Figures 1 to 3 illustrate the comovements in regional national equity market indices, for the Americas, Europe and Asia, respectively. In particular, there appear to be common comovements in European indices; the indices in emerging Asian markets also exhibit similar movements, but only after 1990.

IV. Univariate Unit Root Tests

Prior to examining whether there are long-run comovements among national equity markets, it is necessary to establish the nonstationarity of the national indices of equity market prices. That is, to ensure that all the equity market indices form a cointegrating relationship with a stationary error term, it is necessary to establish that all the individual time series are of the same data-generating process (same order of integration) so as not to obtain spurious results (see Granger and Newbold 1974). 1/2/

1/ In this paper, use of the term integration in reference to econometric issues describes the representation of a process as a sum of past shocks. A process is said to be integrated of order d ($I(d)$) if, after differencing d times, the resulting process is stationary (denoted $I(0)$). On the other hand, use of the term integration in reference to economic issues describes the degree of interdependence across national equity markets.

2/ Nonstationary unit root processes are characterized by time paths that exhibit trend movements, and the paths of such series would usually be expected to diverge from their original value over time. Variables with unit roots are characterized by fluctuations around a stochastic trend, with shocks leading to permanent movements in the series away from trend. However, if there are strong long-run linkages between a group of individual series so that a linear combination of them is stationary (stable), then the series are said to be cointegrated.

In testing whether or not there is any long-run relationship between the national market indices, the cointegration technique makes use of the fact that nonstationary series generally evolve over time. A stationary error term means that the variables of interest have tended to move together over the long run, and if an exogenous shock drives the variables out of equilibrium, then there is a tendency for them to move together again. In this sense, cointegration of the national stock market price indices implies that these prices move together over time, and revert to an equilibrium relationship in response to shocks. A priori, we would expect that given the pronounced upward trend in national stock market indices, they would be nonstationary.

Based on Figures 1-3 and our discussion of portfolio capital flows, we argue that the increase in capital flows across national borders and the relaxation of exchange controls in 1990-91 divide the sample into two subperiods. We denote these as the "control" and "post-control" subperiods. The validity and usefulness of the sample breakpoints is examined empirically in this and the following section, using unit root tests and tests of cointegration.

To determine the order of integration of each country's weekly stock price series we use the univariate unit root tests of Phillips (1987), Perron (1988), and Phillips and Perron (1988). The construction of the tests and a discussion of their statistical properties are given in the above papers and in the Appendix. The results of the Phillips-Perron unit root tests are given in Tables A1-A3, first for the full period (which differs between markets in industrial and developing countries), then for various subperiods, with the variable breakpoint separating the control and post-control subperiods set at 1990 week 52. 1/ 2/

Clearly, the null hypothesis of a unit root in the series of equity market indices cannot be rejected at the five percent level for all national equity markets--all the statistics testing for one unit root versus no unit roots are insignificant. However, all the statistics testing for two unit

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1/ As noted in the data section above, the national stock market data for developing countries runs from 1989 week 1 to 1995 week 9; for developed countries the data runs from 1989 week 17 to 1995 week 10. The different starting dates is due to data being unavailable for Spain.

2/ All test statistics are calculated using both: (i) five lagged difference terms; and (ii) where the number of lagged difference terms is determined by the data-dependent lag truncation parameter of Andrews (1991). Both tests use the Fejer kernel and the prewhitening technique of Andrews and Monahan (1992). Results similar to those obtained for the 1990 week 52 breakpoint were derived using breakpoints at 1990 week 26 and 1991 week 26--these results are not reported, but are available from the authors upon request.

Figure 1: Americas--Equity Price Indices, 1989-95

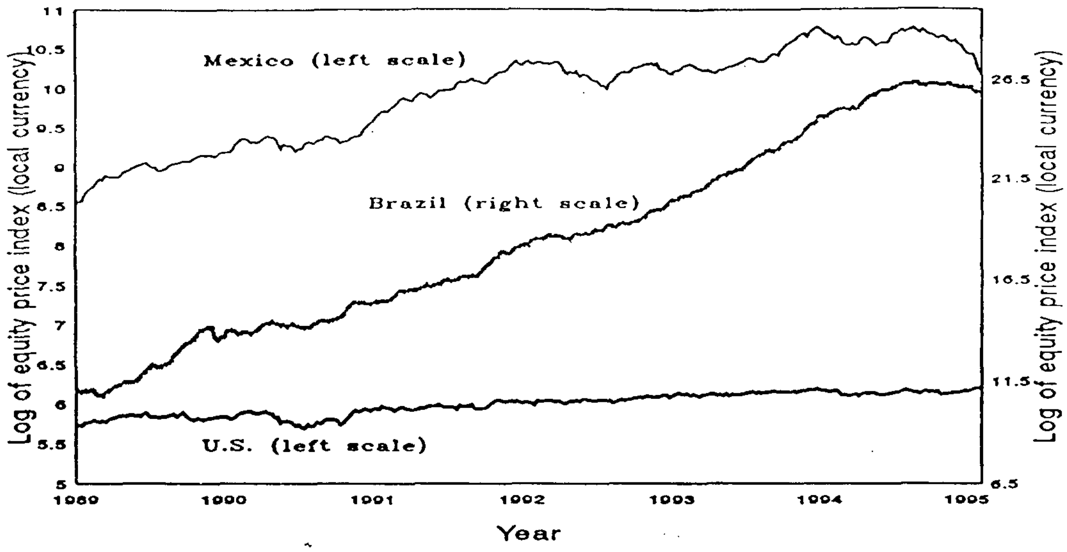


Figure 2: Europe and the Middle East--Equity Price Indices, 1989-95

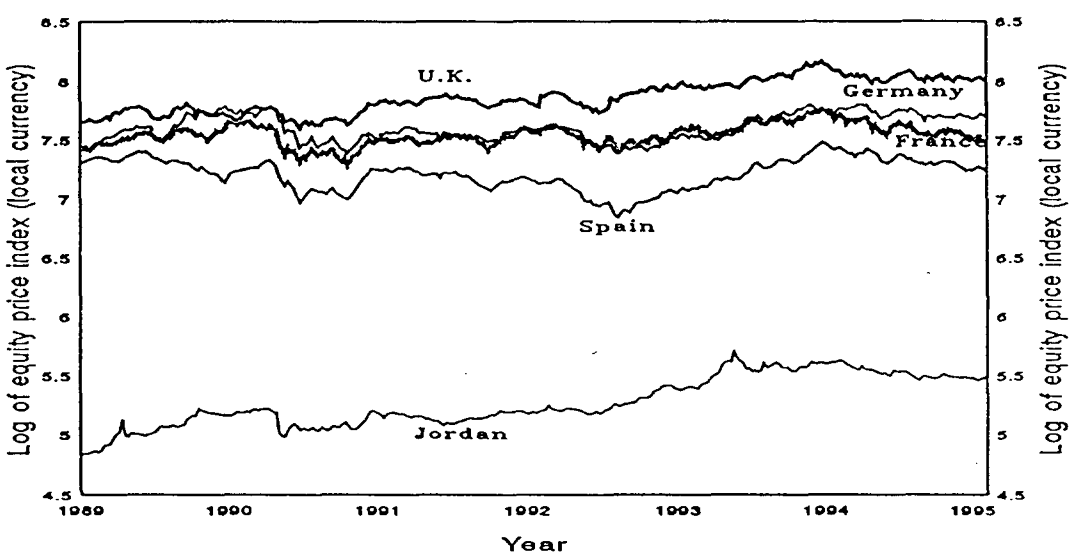
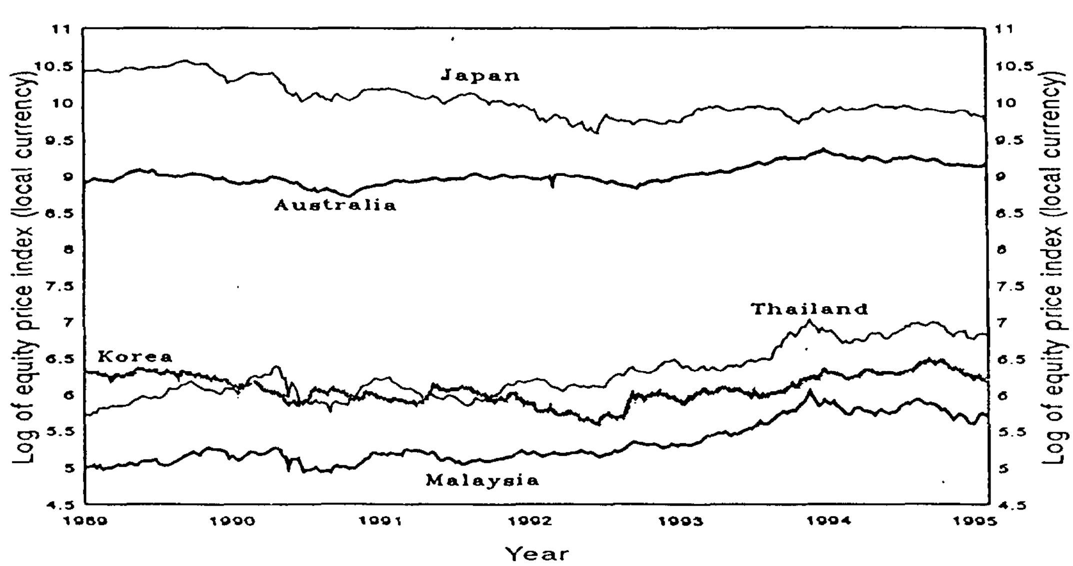


Figure 3: Asia Pacific--Equity Price Indices, 1989-95



roots versus at most one unit root are significant--evidence that all the national equity market indices are $I(1)$ series and thus nonstationary. 1/

V. Multivariate Cointegration Tests

To examine the extent of integration of national stock markets, that is whether there is a long-run commonality of cross-country movements in national indices of stock prices, we use Johansen's (1988) maximum likelihood estimator, derived from a time series canonical correlation analysis. In particular, we use Johansen's likelihood ratio trace statistic to estimate the number of linearly independent cointegrating vectors--long-run equilibrium relationships--(see also Johansen and Juselius 1990). An advantage of Johansen's procedure, which is of particular relevance for this paper, is that it allows the number of cointegrating vectors to be determined empirically. The procedure can be summarized as follows.

Let S_t be a 1 x p vector of the logarithm of international stock market prices ($s_{1t}, s_{2t}, \ldots, s_{pt}$), where s_{it} is the logarithm of country i's stock market index. S_t can be written as a q^{th} order vector autoregressive process

$$\Delta S_t = \sum_{i=1}^{q} \delta_i \Delta S_{t-i} + \pi S_{t-1} + \epsilon_t, \qquad (1)$$

where π is a p x p matrix. If the π matrix is of full rank p, then S_t is stationary. Alternatively, if the π matrix is of rank zero, then a vector autoregressive model solely in first differences is appropriate (as then the S_t are nonstationary). The most interesting situation is where r, the rank of π, lies between zero and p (the reduced rank case of Anderson 1951). In this situation there are p x r matrices γ and α such that $\pi=\gamma\alpha'$, where α is the matrix of r cointegrating vectors. Johansen's estimating method first concentrates the likelihood to purge it of the effects of short-run dynamics, by regressing ΔS_t on $\Delta S_{t-1}, \ldots, \Delta S_{t-p+1}$ and regressing S_{t-p} on $\Delta S_{t-i}, \ldots, \Delta S_{t-p+1}$. This yields a p x 1 residual vector R_{0t} from the ΔS_t regression, and a p x 1 residual vector R_{1t} from the S_{t-p} regression; the associated covariance matrices are:

$$C_{ij} = (1/T)\sum_{t=1}^{T} R_{it} R_{jt}' \quad i,j=0,1;\ t=1,\ldots,T. \qquad (2)$$

1/ The results of the tests for two unit roots versus at most one unit root are not reported, and are available from the authors on request.

The solution to the equation

$$|\lambda C_{11} - C_{10}C_{00}^{-1}C_{01}| = 0 \tag{3}$$

gives ranked eigenvalues $\lambda_1,\ldots,\lambda_p$; and a matrix of eigenvectors $V=(v_1,\ldots,v_p)$. The estimates of α that maximize the likelihood function under the hypothesis $\pi=\gamma\alpha'$ then form the set of eigenvectors $(v_1,\ldots,v_r)$. Accordingly, Johansen's likelihood-ratio trace test of the null hypothesis of no more than r stationary linear combinations of S_t is defined as:

$$J(r) = -T \textstyle\sum_{i=r+1}^{p} \ln(1-\lambda_i), \tag{4}$$

where T is the sample size. The distribution of this test statistic is reported by Johansen (1988) for $r=1,\ldots,5$; we base our inferences on the critical values of Osterwald-Lenum (1992). 1/

In summary, our examinations as to whether the national stock price indices are cointegrated is equivalent to testing whether there are linear combinations of these nonstationary variables which will converge to stationary long-run equilibrium relationships. If such a relationship does exist then the individual stock market indices have tended to move together over the long run, and will revert to an equilibrium relationship in response to exogenous shocks.

We report the cointegration results for the full period and for each of the subperiods, for emerging markets alone, industrial country markets alone, and all markets (Tables 5-7). In an attempt to bring out the underlying cointegrated relationships, we apply the Johansen cointegration test on the full set of industrial-country, developing-country and all-country market indices separately, then on smaller subsets of regionally-based national stock price indices. We pay particular attention to testing the possibility that increased integration of national stock markets, if any, has occurred on a regional basis. This is done by assessing the extent to which countries sharing a common geographic region have cointegrated national stock market indices.

For the six emerging markets (Brazil, Mexico, Thailand, Malaysia, Korea and Jordan) the regional groupings consist of: Brazil and Mexico (Latin American group); Thailand, Malaysia and Korea (Asia group); and Jordan alone. For markets in the seven industrial countries (Australia, Japan, U.K., U.S., Germany, France and Spain) the regional groupings are: France,

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1/ Accepting $J(0)$ implies there are no linearly independent cointegrating vectors. Rejecting $J(0)$ but failing to reject $J(1)$ implies there is one unique cointegrating vector.

Germany, U.K. and Spain (European group); Japan, U.S. and Australia (Pacific I group); Australia and U.S. (Pacific II group); Japan and Australia (Pacific III group); and U.S. alone. 1/ For all thirteen equity markets the regional groupings are: Jordan, France, Germany, U.K. and Spain (Europe+1 group); Japan, Australia, Thailand, Malaysia and Korea (Asia Pacific I group); Japan, Australia, Thailand, Malaysia, Korea and U.S. (Asia Pacific II); Brazil, Mexico and U.S. (Americas group). The cointegration test results are based on the logarithm of national equity price indices, and all use five lagged difference terms.

1. Analysis of emerging markets

For the six emerging markets the results are as follows. For the full period of analysis (1989 week 1 to 1995 week 9) there are four cointegrating vectors when analyzing all six markets with only one cointegrating vector for each of the Asia and Latin American groups, and of course none for Jordan. For the subperiod up to and including 1990 week 52 (control subperiod) there are three independent cointegrating vectors among the six markets; three for Asia and none for Latin America. For the subperiod after this date (post-control subperiod), the results differ in that now there is only one independent cointegrating vector for Asia and one for Latin America.

The result for Latin America is evidence of increasing integration across the emerging stock markets in this region, due in part to increasing liberalization of international capital flows and the relaxation of regulations governing national equity and financial markets (Table 5). In the control subperiod, both the Mexican and Brazilian indices moved independently of one another, which could be partly due to capital and exchange controls inhibiting arbitrage opportunities between these markets; in the post-control subperiod, the trends in the two market indices move together.

The result for Asia indicates that the time series properties of the national equity price indices differ between the control and post-control subperiods. In the control subperiod there is evidence that all Asian national stock price indices are stationary, and accordingly there is no information on the long-run equilibrium relationships between these series. 2/ In the post-control subperiod there is evidence of nonstationarity for each of the Asian equity price indices, with stochastic trends in the indices connected by at least one long-run equilibrium relationship.

1/ By definition, and given the results of our unit root tests, the stock market indices of Jordan and the U.S. are each $I(1)$.

2/ This result could be due to controls on foreign capital flows influencing the evolution of the national equity market indices.

Table 5. Johansen Cointegration Results for Six Emerging Equity Markets

Countries	Period 1989 W1 - 1995 W9 n	r	1989 W1 - 1990 W52 n	r	1991 W1 - 1995 W9 n	r
All	6	4	6	3	6	3
Asia	3	1	3	3	3	1
Latin America	2	1	2	0	2	1
Jordan	1	0	1	0	1	0

Source: Authors' calculations.

Notes: All results are derived from the Johansen (1998) technique, using five lagged difference terms. The number of countries is denoted by n; the number of independent cointegrating vectors is denoted by r. While the regional breakup of countries must sum to the total number of countries (n), the number of regional long-run equilibrium relationships (r) need not sum to the total, as there may be cross-regional equilibrium relationships. "All" denotes the six emerging equity markets: Jordan, Korea, Malaysia, Thailand, Mexico and Brazil. "Asia" denotes the three Asian emerging equity markets: Korea, Malaysia, and Thailand. "Latin America" denotes the two Latin American emerging equity markets: Mexico and Brazil. The full sample period runs from 1989 week 1 to 1995 week 9; the "control" subperiod runs from 1989 week 1 to 1990 week 52; and the "post-control" subperiod runs from 1991 week 1 to 1995 week 9.

When we examine the regions separately there are only two independent equilibrium relationships for the full period--one in Asia and one in Latin America. When all six emerging markets are examined together, we find four equilibrium relationships. The implication of this is that there must be at least two equilibrium relationships that connect the regions together. That is, shocks to national equity markets not only have intra-regional implications, but over a long enough period, can also affect national equity market indices in other regions. 1/

2. Analysis of industrial markets

For the seven industrial country markets the results are as follows. For the full period of analysis (1989 week 17 to 1995 week 10) there are four cointegrating vectors when analyzing all seven markets; no cointegrating vectors for the Pacific II and Pacific III groups; two for the European group; one for the Pacific I group and, of course, none for the United States. With a breakpoint of 1990 week 52, for the period up to and including this date (control subperiod) there are five independent cointegrating vectors among the seven countries, while the results for the other groups are as for the full sample. For the subperiod following the breakpoint (post-control subperiod) the results differ in that now there are only four independent cointegrating vectors among the seven markets; and only one cointegrating vector for the European group. 2/

Unlike the results for the emerging markets, for the seven industrial countries there is little evidence of increased integration of national equity markets in the post-control subperiod relative to the control subperiod. Equity markets in industrial countries were already highly integrated by the start of our sample period. This is evidenced by the relatively large number of intra-regional and inter-regional equilibrium relationships linking equity markets in industrial countries (Table 6).

3. Joint analysis of emerging and industrial markets

For the full period of analysis (1989 week 17 to 1995 week 9) there are eleven independent cointegrating vectors for the thirteen emerging and industrial country markets: three cointegrating vectors for the Asia Pacific II and Europe+1 groups; two cointegrating vectors for the Asia Pacific I group; and one for the Americas group. With a breakpoint of 1990 week 52, for the period up to and including this date (control subperiod) there are

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1/ The results for all six emerging markets indicate that there is an additional equilibrium relationship for the full period ($r=4$) than is evidenced in each of the subperiods (where $r=3$). This additional relationship is most likely weaker than the other three relationships which exist in both subperiods, and takes longer to assert itself on the data than the length of each subperiod.

2/ The result for Europe is sensitive to the number of lags--a reduced number of lags yields two cointegrating vectors.

Table 6. Johansen Cointegration Results for Seven Industrial Equity Markets

Countries	Period					
	1989 W17 - 1995 W9		1989W17 - 1990 W52		1991 W1 - 1995W10	
	n	r	n	r	n	r
All	7	4	7	5	7	4
Europe	4	2	4	2	4	1
Pacific I	3	1	3	1	3	1
Pacific II	2	0	2	0	2	0
Pacific III	2	0	2	0	2	0
U.S.	1	0	1	0	1	0

Source: Authors' calculations.

Notes: All results are derived from the Johansen (1998) technique, using five lagged difference terms. The number of countries is denoted by n; the number of independent cointegrating vectors is denoted by r. While the regional breakup of countries must sum to the total number of countries (n), the number of regional long-run equilibrium relationships (r) need not sum to the total, as there may be cross-regional equilibrium relationships. "All" denotes the seven industrial equity markets: U.S., Japan, U.K., France, Germany, Spain and Australia. "Europe" denotes the four European industrial equity markets: U.K., France, Germany, and Spain. "Pacific I" denotes the three Pacific industrial equity markets: U.S., Japan, and Australia. "Pacific II" denotes the two Pacific industrial equity markets: U.S., and Australia. "Pacific III" denotes the two Pacific industrial equity markets: Japan and Australia. The full sample period runs from 1989 week 17 to 1995 week 10; the "control" subperiod runs from 1989 week 17 to 1990 week 52; and the "post-control" subperiod runs from 1991 week 1 to 1995 week 10.

Table 7. Johansen Cointegration Results for Thirteen Equity Markets

Countries	Period 1989 W17 - 1995 W9 n	r	1989W17 - 1990 W52 n	r	1991 W1 - 1995W9 n	r
All	13	22	13	12	13	11
Americas	3	1	3	1	3	1
Europe+1	5	3	5	2	5	3
Asia Pacific I	5	2	5	3	5	2
Asia Pacific II	6	3	6	4	6	3

Source: Authors' calculations.

Notes: All results are derived from the Johansen (1998) technique, using five lagged difference terms. The number of countries is denoted by n; the number of independent cointegrating vectors is denoted by r. While the regional breakup of countries must sum to the total number of countries (n), the number of regional long-run equilibrium relationships (r) need not sum to the total, as there may be cross-regional equilibrium relationships. "All" denotes the thirteen emerging and industrial equity markets: Jordan, Korea, Malaysia, Thailand, Mexico, Brazil, U.S., Japan, U.K., France, Germany, Spain and Australia. "Americas" denotes the three American equity markets: U.S., Mexico and Brazil. "Europe+1" denotes the five European and Middle Eastern equity markets: U.K., France, Germany, Spain, and Jordan. "Asia Pacific I" denotes the five Asia Pacific equity markets: Japan, Australia, Korea, Malaysia, and Thailand. "Asia Pacific II" denotes the six Asia Pacific equity markets: U.S., Japan, Australia, Korea, Malaysia, and Thailand. The full sample period runs from 1989 week 17 to 1995 week 9; the "control" subperiod runs from 1989 week 17 to 1990 week 52; and the "post-control" subperiod runs from 1991 week 1 to 1995 week 9.

12 independent cointegrating vectors among the 13 countries; four cointegrating vectors for the Asia Pacific II group; three for the Asia Pacific I group; two for the Europe+1 group; and again one for the Americas group. For the subperiod following the breakpoint (post-control subperiod) the results differ considerably in that now there are only 11 independent cointegrating vectors among the thirteen countries; there are only two cointegrating vectors for the Asia Pacific I group; three for the Asia Pacific II group; and three for the Europe+1 group.

In the post-control subperiod there is evidence that the Jordanian equity market has tied itself to movements in European markets--in contrast, for the control subperiod there was no change in the number of equilibrium relationships when Jordan was included with the European countries (Tables 6 and 7). In moving from the Asia Pacific I to the Asia Pacific II group, the addition of the U.S. increases the number of long-run equilibrium relationships, and demonstrates the importance of accounting for movements in the U.S. equity market as a key determinant of movements in the Asia Pacific equity markets. 1/

When we examine separately the regions (Asia Pacific I, Americas and Europe+1) which embrace all 13 markets, we find there are only 6 independent equilibrium relationships for the full period. However, when all 13 markets are examined together, we find 11 equilibrium relationships. Accordingly, there are nearly as many inter-regional equilibrium relationships as there are intra-regional equilibrium relationships. An important implication of this is that there are only two independent trends in the equity markets of our 13 country sample. This highlights the importance of inter-regional linkages (most likely the strong effect of the U.S. market on other equity markets), and demonstrates that limiting any analysis of integration across national equity markets to those involving intra-regional linkages will ignore key effects flowing from cross-regional links between emerging and industrial equity markets.

In summary, for both industrial and emerging country markets, and across regions containing diverse sets of countries, we find evidence of increased integration among national stock markets in the post-control subperiod of the sample, in comparison with the control subperiod. The linkages across national stock markets in the European and the Americas groups appear to be relatively stronger than those across national stock markets in the Asia Pacific groups (as the number of independent cointegrating vectors is lower in the former than in the latter). However, the links among national stock markets in the Asia Pacific groups have definitely strengthened in the post-controls subperiod, in comparison with the controls subperiod. It is also clear that emerging equity markets have become more integrated with equity markets in regional industrial countries

1/ The results for the Asia Pacific I and II are sensitive to the number of lags--a reduced number of lags yields two and three cointegrating vectors, respectively, in the control subperiod.

over the sample period--a finding consistent with the lowering of barriers to international capital flows into, and the presence of foreigners trading in, equities in emerging markets. Our analysis thus confirms: (i) that there has been increasing integration of national stock markets in recent years, and (ii) this has occurred not only through greater regional integration of national stock markets, involving stronger linkages among emerging and industrial equity markets sharing a common geographic region, but also through strengthened cross-regional linkages between emerging and industrial markets.

VI. Short-Run Dynamics--Contagion Effects

The cointegration results discussed above have provided important insights into the long-run equilibrium relationship between prices in various national stock markets. This section considers the short-run dynamics of this relationship. In addition, it examines an issue which is of primary interest to policymakers: the dynamic process by which the stock market indices return to their equilibrium states following a shock to equilibrium.

Lütkepohl and Reimers (1992), Mellander et.al. (1992), and Pesaran and Shin (1994) have used impulse response functions in the context of cointegrated systems, to examine the process by which the system returns to its equilibrium state. In this section we essentially follow the approach of Lütkepohl and Reimers (1992). Accordingly, we analyze the time paths of the response of individual stock market indices to shocks using the orthogonalized impulse response approach, after controlling for the long-run restrictions implied by the cointegrated relations among the national stock markets. This is an important advance on current practice in empirical time series analysis, as we explicitly account for information embodied in the long-run equilibrium relationships. That is, our approach to examining short-run dynamics is a multivariate version of the typical single-equation error correction model.

This section looks specifically at the short-run interaction between the national stock market indices in a regional context, and tries to isolate the presence of so-called "contagion" effects, which have received attention in recent years. Here we highlight the effects of shocks on emerging equity markets by considering two cases: (i) the effects on emerging equity markets of global shocks to the long-run equilibrium relationship of regional equity markets; and (ii) the effects of country-specific shocks on other emerging markets and on the regional equilibrium relationship itself. It is important to be clear about these two different effects. A global shock represents a shift away from a given long-run equilibrium relationship among the national stock market indices, while a country-specific shock represents a short-lived movement in a national price index. As an example of each type of shock, compare the post-December 1994 fall in demand for peso-denominated bonds and stocks (a large proportion of which were held outside Mexico) with a country-specific shock in the Mexican corporate sector (such as, for instance, a takeover bid for a Mexican company by a U.S. company). The former is a global shock which moved the Americas stock markets out of its long-run equilibrium; the latter has only a localized effect with little spillover (contagion effects) across Mexico's borders.

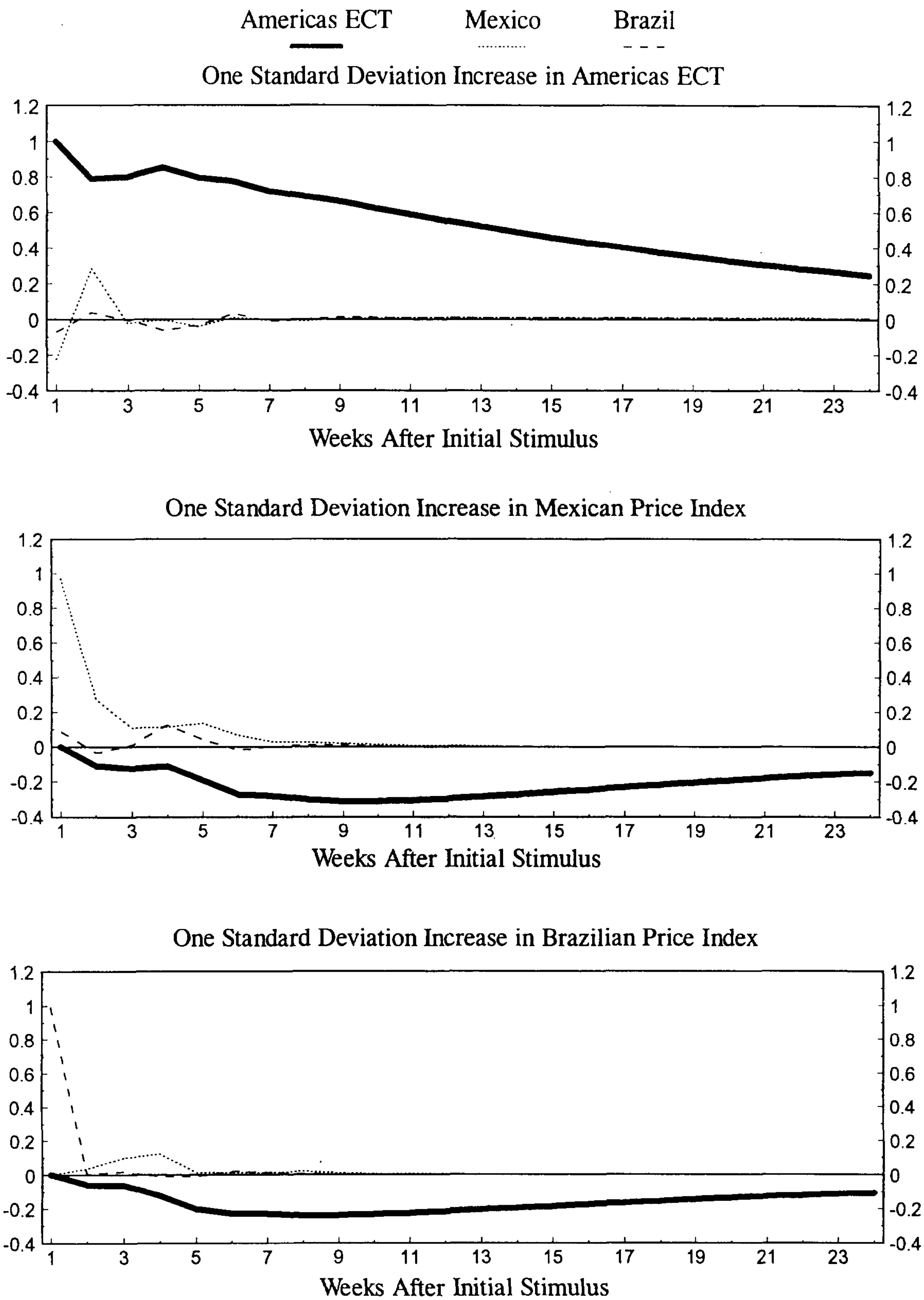
Our methodology for undertaking this investigation is as follows: taking the Americas region as our example, we first regressed the U.S. stock market index on a constant term and the stock market indices of Mexico and Brazil; the residuals from this regression then become the error correction term (ECT) for the Americas. This ECT can be looked upon as the deviations from the historical norm of the evolution of the three national stock market price indices (given that our earlier cointegration analysis indicated the three indices were cointegrated over the sample period). A vector error correction model (VECM) was specified such that the temporal ordering imposed was from the Americas ECT to Mexico and Brazil. That is, Brazil and Mexico are tied to the U.S. in equilibrium, and the former two national markets will adjust to movements in the latter market, not the other way around (all VECMs were run at a lag length of order five).

From the VECM we obtained impulse response functions for the interaction between the Americas' ECT and the national market indices. We then trace out the plots of the effect of a one standard deviation positive shock to the Americas' ECT and see its effect through time on the Mexican and Brazilian market indices; similarly for a one standard deviation shock to the Mexican index, and to the Brazilian index (Figure 4).

The Figures reveal that an unanticipated change in the historical norm of the stock market indices would cause market participants to revise their forecasts of expected deviations from the long-run equilibrium by 25 percent, even 20 weeks into the future. Hence, once the long-run equilibrium relationship is disturbed, any forces that restore this relationship between the three national stock markets tend to operate relatively slowly. In other words, controlling for the cointegration which exists across the three national stock markets, it takes about 24 weeks for the long-run equilibrium to be regained following such a global shock. On the other hand, market movements in Brazil and Mexico have little impact on deviations from the long-run historical relationship between the national price indices. The short-run dynamics of country-specific shocks (contagion effects) appear to dissipate after a period of (at most) two weeks following an unanticipated shock to Brazilian or Mexican stock prices (Figure 4). It is also interesting to note that the effect of a Mexico-specific shock on the Americas' ECT is somewhat greater than that of a Brazil-specific shock, as the ECT is 0.3 standard deviations below zero up to eight weeks after the Mexican shock; it is only 0.25 standard deviations below zero at the same time after an equivalent Brazilian shock (Figure 4).

Figure 4

Americas Responses to Exogenous Shocks



There could be at least two reasons for the relatively slow adjustment of equity prices to their new long run equilibrium following global shocks. The first reason is the possibility that during times of financial stress, uncertainty about market fundamentals and policy responses increases. This can create an incentive for some investors to wait until the uncertainty has been resolved, thus disturbing historical relationships in the stock market prices across countries. The second reason is that real variables in different national economies may respond with differing speeds to any given global shock. This difference in the speed of adjustment could disturb the historical trends in the market fundamentals and thus again the historical relationship in stock market prices across countries. This temporary change in the historical relationship need not, of course, yield any unexploited profitable opportunities.

The results for the emerging markets of the Asia group (Korea, Thailand and Malaysia) and Europe/Middle East (Jordan) are similar to those illustrated above for the Americas. Global shocks to Asian stock markets result in persistent deviations from the long-run relationship between the national stock market indices--it takes about 25 weeks for this equilibrium to be restored, while again the effects of country-specific shocks have been largely exhausted after about three weeks (Figure 5). A similar picture emerges for Jordan with respect to deviations from the long-run equilibrium relationship between it and the European stock market indices (these last about 24 weeks), although it takes longer (up to six weeks) for any contagion effects to be exhausted (Figure 6).

In summary, the above analysis suggests that contagion effects, operating across national borders within a given geographic region, dissipate quickly--such disturbances are typically eliminated in a few weeks. However, if national stock markets are subject to global shocks which cause them to deviate from their long-run equilibrium relationship, it takes at least 24 weeks for the long-run relationship to reassert itself. That is, the dynamics of the long-run process for the evolution of national stock market indices take about six months to work through the system. Clearly, the most important channel through which adjustment to shocks occurs is the error correction mechanism, which captures deviations from the long-run relationship between national stock market indices emanating from global shocks. In contrast, in our sample country-specific shocks have minimal contagion effects on the evolution of stock prices across national borders.

VII. Summary and Conclusions

This paper has investigated empirically the degree of international integration of industrial and emerging country equity markets, and changes in this integration over time. Using weekly data over the six year period from January 1989 to March 1995 for seven industrial and six emerging markets, two key issues were analyzed. First, the extent to which equity prices have tended to move similarly across countries and regions in the long-run--this issue was analyzed using the Johansen cointegration methodology. Second, whether there is any evidence of "contagion" effects--this issue was analyzed using deviations from the long-run cointegrating relationships among the national equity markets, and involved estimating the time taken for adjustments to these historical norms, following global or country-specific shocks to these markets. The empirical analysis yielded several key findings, which are summarized below.

The application of Phillips-Perron unit root tests showed that for the full sample period (and for the various subperiods) the null hypothesis of a unit root in the national stock market indices could not be rejected--all indices were nonstationary (integrated of order one). The results from the cointegration analysis suggest that there has been increased international integration of emerging equity markets, beginning in the early 1990s. However, markets in industrial countries were already largely integrated at the start of our sample period. While this increasing integration has occurred through greater regionalization of national stock markets, involving stronger linkages between emerging and industrial country markets sharing a common geographic region, there has also been strengthened cross-regional links between emerging and industrial equity markets.

An examination of the short-run interaction between national market indices shows that any cross-country contagion effects emanating from country-specific shocks dissipate in a matter of weeks. However, if the national stock markets are subject to a global shock which causes them to deviate from their long-run equilibrium relationship, it takes several months for this long-run relationship to reassert itself.

These findings have a number of implications. The cointegration results, which indicate that the globalization of equity markets has increased in recent years, suggest that investors need to more closely monitor developments in emerging markets. While this does not suggest that there are no benefits of risk reduction through portfolio diversification, it does suggest that to reap these benefits investors may have to be more selective. It is also noteworthy that globalization of national equity markets has accompanied increased regional links between industrial and emerging equity markets. This result mirrors trends in global trade integration, which has also proceeded apace in the context of regional-based trading arrangements. These developments suggest powerful economic (and some non-economic) factors which continue to exercise a dominant regional influence.

Finally, the results assessing the magnitude and duration of contagion effects are interesting from a policy perspective. They suggest that it is appropriate for policymakers to monitor developments occurring in other national equity markets, so as to respond to contagion effects spilling over to their own equity markets. In particular, if the shocks affecting national equity markets are global in nature (such as the October 1987 fall in U.S. equity prices or the December 1994 Mexican devaluation and fall in the equity prices), their intra-regional and inter-regional effects

Figure 5

Asian Responses to Exogenous Shocks

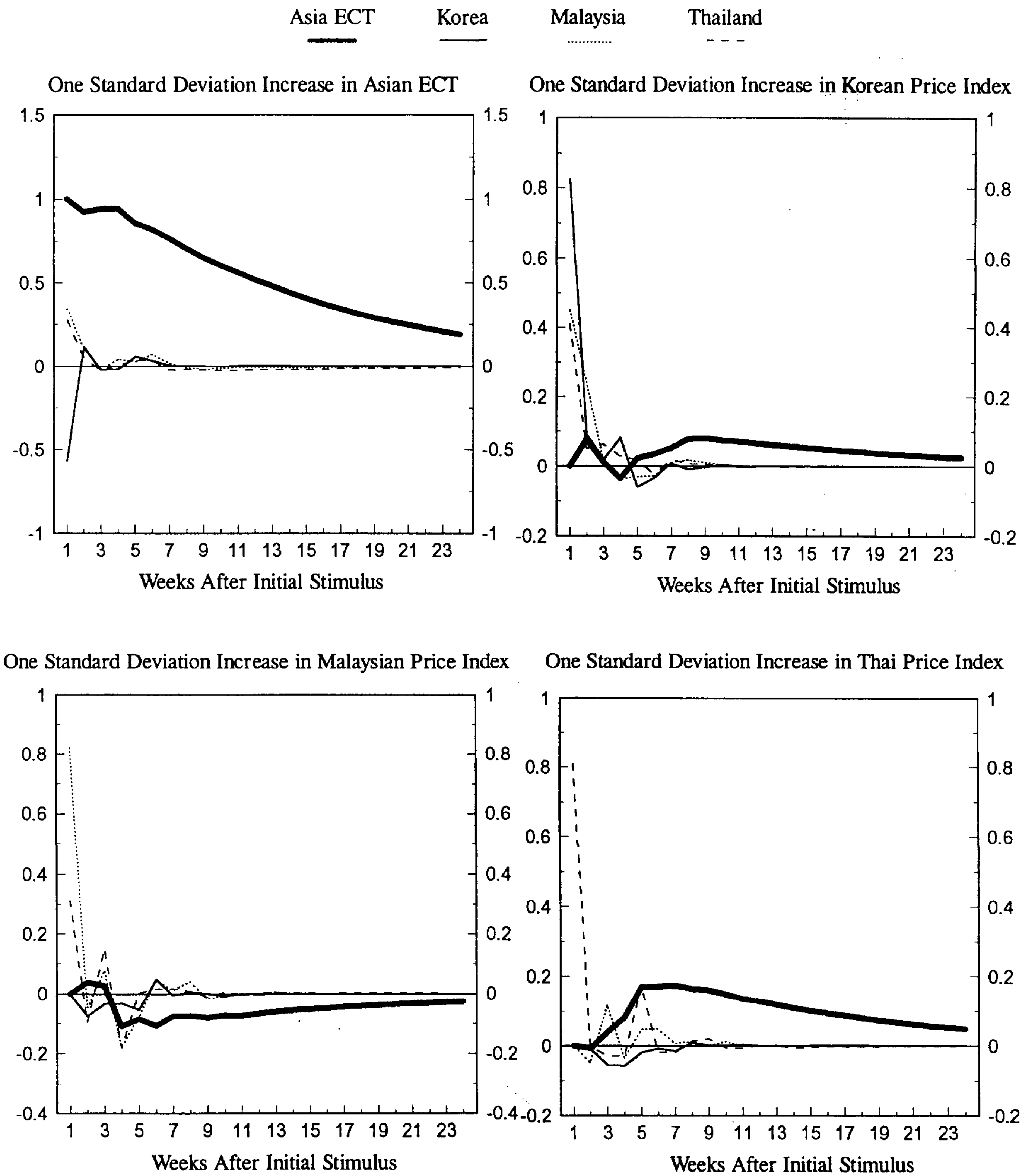
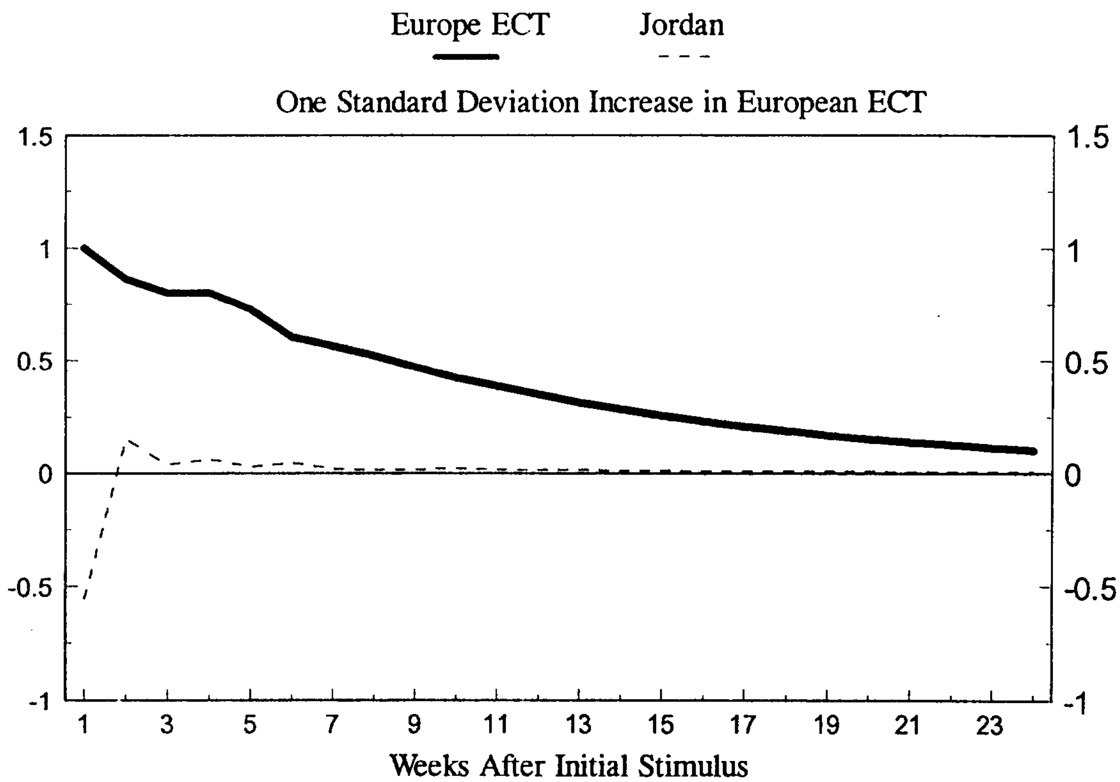
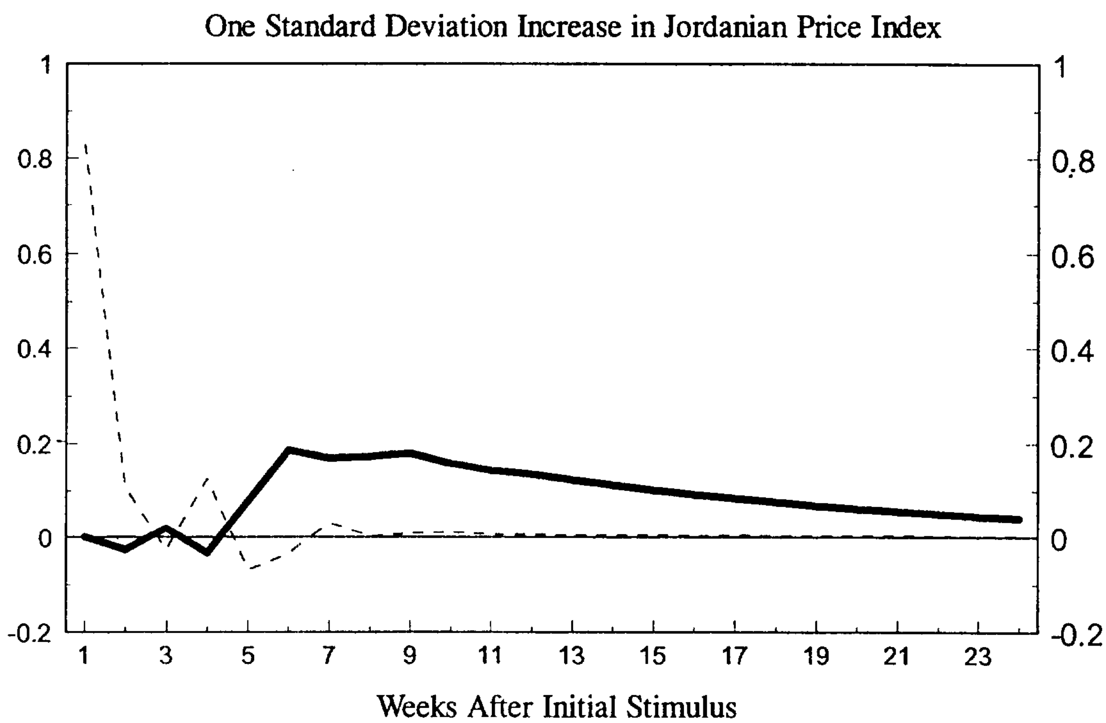


Figure 6

European Responses to Exogenous Shocks





could last a considerable length of time. Accordingly, in the absence of appropriate action by policymakers, such global shocks could have serious implications for the behavior of prices in individual national equity markets. However, if the shocks affecting national equity markets are localized or country-specific in nature, then any subsequent disturbance will be quickly eliminated. In such circumstances, there may be little role for intervention by policymakers in the equity markets of other countries.

Univariate Unit Root Tests on National Indices of Stock Market Prices

The first step in examining the data is to determine the time series properties of the national indices of stock market prices for the 13 countries in our sample. The augmented Dickey-Fuller (ADF) test of Dickey and Fuller (1979, 1981) and the $Z(\alpha)$ and $Z(t)$ tests of Phillips (1987), Phillips and Perron (1988) and Perron (1988) are used to determine the order of integration of the series. The Fejer kernel is used, as is the prewhitening technique of Andrews and Monahan (1992) to aid in estimating the long-run variance of the error term in the unit root regression. Andrew's (1991) data-dependent automatic bandwidth selector is also invoked to determine the optimal lag truncation order.

In undertaking the test an ordinary least squares regression of $x_t = \mu + \beta t + \alpha x_{t-1} + \epsilon_t$ was run, where x is the natural log of the national stock market index of each of the 13 countries. As described in the text, the sample ranges from 1989 week 17 to 1995 week 9 (305 weeks); with a breakpoint at 1990 week 52, the first subperiod comprises 87 weeks (1989 week 17 to 1990 week 52) and the second subperiod comprises 217 weeks (1991 week 1 to 1995 week 9). The results of the stationarity tests are presented below for the full period (Table A1), the first subperiod (Table A2) and the second subperiod (Table A3).

The results are presented for unit root tests based on five lagged difference terms (given the Johansen (1988) cointegration test results) and for the optimal bandwidth (using Andrew's (1991) technique). They indicate that all of the series are nonstationary (Table A1). The statistics testing for one unit root versus no unit roots are insignificant, indicating we cannot reject the null hypothesis of a unit root--the national stock market indices are integrated of order one ($I(1)$). Moreover, the results for the subperiods generated using the other breakpoints (1991 week 26 and 1991 week 52) are similar to those reported for a breakpoint of 1990 week 52.

Table A1. Univariate Unit Root Tests on Stock Prices, 1989-95 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
FRANCE (Five Lagged Difference Terms)					
None	1.000	0.080	1.000	0.009	0.153
Constant	0.956	-2.544	0.957	-13.876	-2.745
Constant and Time Trend	0.948	-2.609	0.951	-16.139	-2.801
(Optimal Band Width)					
None	...	...	1.000	0.009	0.151
Constant	...	...	0.957	-13.461	-2.708
Constant and Time Trend	...	...	0.951	-15.440	-2.738
GERMANY (Five Lagged Difference Terms)					
None	1.000	0.539	1.000	0.034	0.583
Constant	0.973	-2.082	0.975	-8.538	-2.128
Constant and Time Trend	0.970	-2.171	0.972	-9.553	-2.226
(Optimal Band Width)					
None	...	...	1.000	0.034	0.605
Constant	...	...	0.975	-7.809	-2.041
Constant and Time Trend	...	...	0.972	-8.680	-2.126
UNITED KINGDOM (Five Lagged Difference Terms)					
None	1.000	0.866	1.000	0.046	0.969
Constant	0.986	-1.656	0.988	-4.289	-1.588
Constant and Time Trend	0.930	-3.456 4/	0.948	-20.187	-3.167
(Optimal Band Width)					
None	...	...	1.000	0.046	1.023
Constant	...	...	0.988	-3.893	-1.527
Constant and Time Trend	...	...	0.948	-17.290	-2.930

Table A1 (cont'd). Univariate Unit Root Tests on Stock Prices, 1989-95 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
JAPAN					
(Five Lagged Difference Terms)					
None	1.000	-1.312	1.000	-0.071	-1.242
Constant	0.990	-1.397	0.992	-3.128	-1.334
Constant and Time Trend	0.975	-2.014	0.979	-8.405	-2.059
(Optimal Band Width)					
None	...	...	1.000	0.071	-1.289
Constant	...	...	0.992	-2.913	-1.294
Constant and Time Trend	...	...	0.979	-7.622	-1.962
UNITED STATES					
(Five Lagged Difference Terms)					
None	1.000	1.646	1.000	0.075	1.661
Constant	0.992	-1.125	0.989	-2.868	-1.405
Constant and Time Trend	0.940	-2.712	0.934	-19.848	-3.233
(Optimal Band Width)					
None	...	...	1.000	0.075	1.615
Constant	...	...	0.989	-2.821	-1.398
Constant and Time Trend	...	...	0.934	-17.826	-3.073
AUSTRALIA					
(Five Lagged Difference Terms)					
None	1.000	0.561	1.000	0.030	0.780
Constant	0.990	-1.303	0.991	-3.469	-1.339
Constant and Time Trend	0.978	-2.061	0.983	-6.475	-1.801
(Optimal Band Width)					
None	...	...	1.000	0.031	0.860
Constant	...	...	0.991	-2.867	-1.222
Constant and Time Trend	...	...	0.983	-5.299	-1.629

Table A1 (cont'd). Univariate Unit Root Tests on Stock Prices, 1989-95 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
SPAIN					
(Five Lagged Difference Terms)					
None	1.000	-0.261	1.000	-0.010	-0.143
Constant	0.982	-2.077	0.987	-7.646	-1.983
Constant and Time Trend	0.981	-2.142	0.987	-7.767	-2.016
(Optimal Band Width)					
None	...	...	1.000	-0.010	-0.143
Constant	...	...	0.987	-7.741	-1.995
Constant and Time Trend	...	...	0.987	-7.885	-2.030
BRAZIL					
(Five Lagged Difference Terms)					
None	0.999	-0.338	0.999	-0.256	-0.335
Constant	0.982	-1.699	0.984	-5.193	-1.673
Constant and Time Trend	0.981	-0.519	0.992	-3.090	-0.466
(Optimal Band Width)					
None	...	...	0.999	-0.248	-0.328
Constant	...	...	0.984	-4.969	-1.640
Constant and Time Trend	...	...	0.992	-2.610	-0.398
MEXICO					
(Five Lagged Difference Terms)					
None	1.000	1.254	1.001	0.159	2.084
Constant	0.994	-2.124	0.990	-3.269	-2.777
Constant and Time Trend	0.997	-0.275	1.002	-2.375	-0.616
(Optimal Band Width)					
None	...	...	1.001	0.160	2.276
Constant	...	...	0.990	-3.171	-2.906
Constant and Time Trend	...	...	1.002	-1.408	-0.392

Table A1 (cont'd). Univariate Unit Root Tests on Stock Prices, 1989-95 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
JORDAN					
(Five Lagged Difference Terms)					
None	1.000	1.447	1.000	0.116	1.453
Constant	0.988	-1.817	0.988	-3.566	-1.818
Constant and Time Trend	0.967	-2.156	0.968	-10.272	-2.260
(Optimal Band Width)					
None	...	...	1.000	0.116	1.452
Constant	...	...	0.988	-3.573	-1.818
Constant and Time Trend	...	...	0.968	-9.945	-2.223
KOREA					
(Five Lagged Difference Terms)					
None	1.000	-0.222	1.000	-0.019	-0.188
Constant	0.982	-1.704	0.983	-5.709	-1.751
Constant and Time Trend	0.981	-1.823	0.982	-5.945	-1.834
(Optimal Band Width)					
None	...	...	1.001	-0.019	-0.190
Constant	...	...	0.983	-5.359	-1.700
Constant and Time Trend	...	...	0.982	-5.653	-1.794
MALAYSIA					
(Five Lagged Difference Terms)					
None	1.000	1.229	1.000	0.130	1.240
Constant	0.995	-0.866	0.995	-1.610	-0.837
Constant and Time Trend	0.977	-1.924	0.978	-7.749	-2.005
(Optimal Band Width)					
None	...	...	1.000	0.130	1.194
Constant	...	...	0.995	-1.727	-0.870
Constant and Time Trend	...	...	0.978	-8.051	-2.042

Table A1 (cont'd). Univariate Unit Root Tests on Stock Prices, 1989-95 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
THAILAND					
(Five Lagged Difference Terms)					
None	1.000	1.188	1.001	0.177	1.448
Constant	0.991	-1.211	0.992	-2.716	-1.255
Constant and Time Trend	0.964	-2.504	0.969	-10.510	-2.321
(Optimal Band Width)					
None	...	...	1.001	0.178	1.541
Constant	...	...	0.992	-2.436	-1.200
Constant and Time Trend	...	...	0.969	-9.044	-2.157

1/ The augmented Dickey-Fuller (1979, 1981) and Phillips-Perron (1988) tests are based on the following model for any series x: $x_t = \mu + \beta t + \alpha x_{t-1} + \epsilon_t$. The tests rely on rejecting a unit root ($\alpha = 1$) in favor of stationarity. The H_0: $\alpha = 1$ is tested by the ADF, $Z(\alpha)$ and $Z(t)$ tests. If the null hypothesis cannot be rejected (the test statistics are not significant), then the series has a unit root. The value of α is also presented when using either test. The Phillips-Perron $Z(\alpha)$ and $Z(t)$ tests are preferable to the ADF test more commonly reported in the literature, as they are robust to a wide variety of serial correlation and time-dependent heteroscedasticity. The ADF test result is also reported here, as it is the univariate version of the Johansen method used in this paper to test for cointegration (see Section V).

2/ All series are in logs. Consistent with our approach in testing for cointegration (see Section V), stationarity tests were carried out using five lagged difference terms, and are denoted as "Five Lagged Difference Terms". Andrew's (1991) data-dependent automatic bandwidth selector is also used to determine the optimal lag truncation order, and these test results are denoted as "Optimal Band Width". Both tests were carried out using the Fejer kernel and the prewhitening technique of Andrews and Monahan (1992).

3/ The critical values for, respectively, the $Z(\alpha)$ test, and (both the) ADF and $Z(t)$ tests at the 5 percent level of significance are: -8.0 and -1.95 for no deterministic components (μ=β=0), -14.0 and -2.88 for the constant term only (μ=0, β≠0) and -21.4 and -3.43 for both deterministic components (μ≠0, β≠0), and are taken from Fuller (1976).

4/ The ADF test results for the U.K. (in contrast to the $Z(\alpha)$ and $Z(t)$ test results) reject the null hypothesis of a unit root--however, as noted above, this test is less robust than the Phillips-Perron tests.

Table A2. Univariate Unit Root Tests on Stock Prices, 1989 week 17 to 1990 week 52 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
FRANCE					
(Five Lagged Difference Terms)					
None	1.000	-0.446	1.000	-0.011	-0.302
Constant	0.957	-1.122	0.969	-4.043	-1.235
Constant and Time Trend	0.946	-1.411	0.959	-4.175	-1.389
(Optimal Band Width)					
None	...	...	1.000	-0.011	-0.299
Constant	...	...	0.969	-3.990	-1.225
Constant and Time Trend	...	...	0.959	-4.526	-1.451
GERMANY					
(Five Lagged Difference Terms)					
None	1.000	-0.081	1.000	-0.000	-0.005
Constant	0.947	-1.456	0.958	-4.278	-1.487
Constant and Time Trend	0.958	-1.166	0.967	-2.785	-1.071
(Optimal Band Width)					
None	...	...	1.000	-0.000	-0.004
Constant	...	...	0.958	-3.893	-1.420
Constant and Time Trend	...	...	0.967	-2.862	-1.088
UNITED KINGDOM					
(Five Lagged Difference Terms)					
None	1.000	0.061	1.000	0.003	0.112
Constant	0.840	-3.009	0.910	-11.799	-2.533
Constant and Time Trend	0.806	-3.503 4/	0.895	-12.580	-2.739
(Optimal Band Width)					
None	...	...	1.000	0.003	0.122
Constant	...	...	0.910	-10.028	-2.353
Constant and Time Trend	...	...	0.895	-11.161	-2.612

Table A2 (cont'd). Univariate Unit Root Tests on Stock Prices, 1989 week 17 to 1990 week 52 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
JAPAN					
(Five Lagged Difference Terms)					
None	1.000	-1.017	1.000	-0.033	-0.981
Constant	0.989	-0.470	0.999	-0.972	-0.443
Constant and Time Trend	0.929	-1.928	0.947	-6.484	-1.964
(Optimal Band Width)					
None	...	...	1.000	-0.033	-1.102
Constant	...	...	0.999	-0.457	-0.235
Constant and Time Trend	...	...	0.947	-5.378	-1.824
UNITED STATES					
(Five Lagged Difference Terms)					
None	1.000	0.099	1.000	0.012	0.381
Constant	0.885	-2.246	0.900	-9.785	-2.479
Constant and Time Trend	0.873	-2.432	0.895	-9.609	-2.547
(Optimal Band Width)					
None	...	...	1.000	0.012	0.408
Constant	...	...	0.900	-8.298	-2.337
Constant and Time Trend	...	...	0.895	-8.550	-2.456
AUSTRALIA					
(Five Lagged Difference Terms)					
None	1.000	-0.791	1.000	-0.014	-0.610
Constant	0.982	-0.656	1.007	-0.723	-0.270
Constant and Time Trend	0.916	-2.372	0.935	-7.958	-2.377
(Optimal Band Width)					
None	...	...	1.000	-0.014	-0.660
Constant	...	...	1.007	-0.215	-0.087
Constant and Time Trend	...	...	0.935	-7.109	-2.305

Table A2 (cont'd). Univariate Unit Root Tests on Stock Prices, 1989 week 17 to 1990 week 52 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
SPAIN					
(Five Lagged Difference Terms)					
None	1.000	-1.056	1.000	-0.037	-0.892
Constant	0.977	-0.986	0.997	-1.940	-0.726
Constant and Time Trend	0.877	-2.837	0.918	-12.796	-2.698
(Optimal Band Width)					
None	...	...	1.000	-0.037	-0.816
Constant	...	...	0.997	-2.671	-0.910
Constant and Time Trend	...	...	0.918	-14.752	-2.871
BRAZIL					
(Five Lagged Difference Terms)					
None	1.002	2.135	1.003	0.258	2.373
Constant	0.987	-1.285	0.990	-1.150	-1.081
Constant and Time Trend	0.967	-1.059	0.978	-4.470	-1.344
(Optimal Band Width)					
None	...	...	1.003	0.257	2.353
Constant	...	...	0.990	-1.172	-1.081
Constant and Time Trend	...	...	0.978	-4.614	-1.369
MEXICO					
(Five Lagged Difference Terms)					
None	1.001	1.821	1.001	0.092	2.558
Constant	0.964	-2.074	0.959	-3.662	-2.904
Constant and Time Trend	0.883	-2.425	0.901	-9.713	-2.862
(Optimal Band Width)					
None	...	...	1.001	0.092	2.707
Constant	...	...	0.959	-3.656	-2.909
Constant and Time Trend	...	...	0.901	-9.675	-2.860

Table A2 (cont'd). Univariate Unit Root Tests on Stock Prices, 1989 week 17 to 1990 week 52 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
JORDAN					
(Five Lagged Difference Terms)					
None	1.001	0.787	1.000	0.042	0.759
Constant	0.910	-2.327	0.931	-5.344	-2.182
Constant and Time Trend	0.930	-1.514	0.933	-4.775	-1.164
(Optimal Band Width)					
None	...	...	1.000	0.042	0.692
Constant	...	...	0.931	-5.940	-2.215
Constant and Time Trend	...	...	0.933	-5.702	-1.751
KOREA					
(Five Lagged Difference Terms)					
None	0.999	-0.903	1.000	-0.040	-0.810
Constant	0.968	-1.058	0.969	-2.558	-1.127
Constant and Time Trend	0.877	-2.006	0.893	-10.244	-2.336
(Optimal Band Width)					
None	...	...	1.000	-0.040	-0.808
Constant	...	...	0.969	-2.444	-1.102
Constant and Time Trend	...	...	0.893	-9.062	-2.206
MALAYSIA					
(Five Lagged Difference Terms)					
None	1.000	-0.048	1.000	-0.001	-0.022
Constant	0.928	-1.543	0.928	-6.113	-1.785
Constant and Time Trend	0.939	-1.264	0.932	-5.321	-1.597
(Optimal Band Width)					
None	...	...	1.000	-0.001	-0.023
Constant	...	...	0.928	-5.779	-1.738
Constant and Time Trend	...	...	0.932	-5.250	-1.586

Table A2 (cont'd). Univariate Unit Root Tests on Stock Prices, 1989 week 17 to 1990 week 52 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
THAILAND					
(Five Lagged Difference Terms)					
None	1.000	0.207	1.000	0.031	0.409
Constant	0.916	-1.988	0.920	-6.366	-2.184
Constant and Time Trend	0.936	-1.313	0.932	-4.668	-1.504
(Optimal Band Width)					
None	...	...	1.000	0.031	0.431
Constant	...	...	0.920	-5.639	-2.125
Constant and Time Trend	...	...	0.932	-4.112	-1.409

1/ The augmented Dickey-Fuller (1979, 1981) and Phillips-Perron (1988) tests are based on the following model for any series x: $x_t = \mu + \beta t + \alpha x_{t-1} + \epsilon_t$. The tests rely on rejecting a unit root ($\alpha = 1$) in favor of stationarity. The H_0: $\alpha = 1$ is tested by the ADF, $Z(\alpha)$ and $Z(t)$ tests. If the null hypothesis cannot be rejected (the test statistics are not significant), then the series has a unit root. The value of α is also presented when using either test. The Phillips-Perron $Z(\alpha)$ and $Z(t)$ tests are preferable to the ADF test more commonly reported in the literature, as they are robust to a wide variety of serial correlation and time-dependent heteroscedasticity. The ADF test result is also reported here, as it is the univariate version of the Johansen method used in this paper to test for cointegration (see Section V).

2/ All series are in logs. Consistent with our approach in testing for cointegration (see Section V), stationarity tests were carried out using five lagged difference terms, and are denoted as "Five Lagged Difference Terms". Andrew's (1991) data-dependent automatic bandwidth selector is also used to determine the optimal lag truncation order, and these test results are denoted as "Optimal Band Width". Both tests were carried out using the Fejer kernel and the prewhitening technique of Andrews and Monahan (1992).

3/ The critical values for respectively, the $Z(\alpha)$ test, and (both the) ADF and $Z(t)$ tests at the 5 percent level of significance are: -7.8 and -1.95 for no deterministic components (μ=β=0), -13.5 and -2.90 for the constant term only (μ=0, $\beta \neq 0$) and -20.2 and -3.47 for both deterministic components ($\mu \neq 0$, $\beta \neq 0$), and are taken from Fuller (1976).

4/ The ADF test results for the U.K. (in contrast to the $Z(\alpha)$ and $Z(t)$ test results) reject the null hypothesis of a unit root--however, as noted above, this test is less robust than the Phillips-Perron tests.

Table A3. Univariate Unit Root Tests on Stock Prices, 1991 week 1 to 1995 week 9 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
FRANCE					
(Five Lagged Difference Terms)					
None	1.000	0.373	1.000	0.023	0.512
Constant	0.944	-2.628	0.942	-12.004	-2.997
Constant and Time Trend	0.939	-2.228	0.935	-13.542	-2.673
(Optimal Band Width)					
None	...	...	1.000	0.023	0.498
Constant	...	...	0.942	-11.970	-2.995
Constant and Time Trend	...	...	0.935	-13.279	-2.648
GERMANY					
(Five Lagged Difference Terms)					
None	1.000	0.820	1.000	0.038	0.959
Constant	0.981	-1.477	0.980	-4.591	-1.667
Constant and Time Trend	0.966	-1.882	0.967	-7.811	-2.011
(Optimal Band Width)					
None	...	...	1.000	0.038	0.988
Constant	...	...	0.980	-4.310	-1.626
Constant and Time Trend	...	...	0.967	-7.109	-1.922
UNITED KINGDOM					
(Five Lagged Difference Terms)					
None	1.000	0.936	1.000	0.044	1.153
Constant	0.972	-2.180	0.974	-5.897	-2.252
Constant and Time Trend	0.920	-3.060	0.938	-15.317	-2.855
(Optimal Band Width)					
None	...	...	1.000	0.044	1.201
Constant	...	...	0.974	-5.647	-2.240
Constant and Time Trend	...	...	0.938	-13.724	-2.712

Table A3 (cont'd). Univariate Unit Root Tests on Stock Prices, 1991 week 1 to 1995 week 9 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
JAPAN					
(Five Lagged Difference Terms)					
None	1.000	-0.839	1.000	-0.036	-0.762
Constant	0.973	-1.678	0.977	-5.908	-1.681
Constant and Time Trend	0.966	-1.912	0.969	-7.993	-1.993
(Optimal Band Width)					
None	...	...	1.000	-0.036	-0.766
Constant	...	...	0.977	-5.679	-1.646
Constant and Time Trend	...	...	0.969	-7.619	-1.946
UNITED STATES					
(Five Lagged Difference Terms)					
None	1.000	1.894	1.000	0.064	1.912
Constant	0.975	-1.950	0.968	-5.942	-2.767
Constant and Time Trend	0.878	-3.419	0.872	-21.758	-4.198
(Optimal Band Width)					
None	...	...	1.000	0.064	2.007
Constant	...	...	0.968	-5.851	-2.780
Constant and Time Trend	...	...	0.872	-20.978	-4.170
AUSTRALIA					
(Five Lagged Difference Terms)					
None	1.000	1.235	1.000	0.045	1.498
Constant	0.984	-1.734	0.982	-3.967	-2.130
Constant and Time Trend	0.974	-1.571	0.973	-6.109	-1.782
(Optimal Band Width)					
None	...	...	1.000	0.045	1.618
Constant	...	...	0.982	-3.746	-2.153
Constant and Time Trend	...	...	0.973	-5.252	-1.657

Table A3 (cont'd). Univariate Unit Root Tests on Stock Prices, 1991 week 1 to 1995 week 9 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
SPAIN					
(Five Lagged Difference Terms)					
None	1.000	0.292	1.000	0.029	0.494
Constant	0.986	-1.537	0.987	-5.099	-1.722
Constant and Time Trend	0.982	-1.627	0.987	-5.785	-1.689
(Optimal Band Width)					
None	...	...	1.000	0.029	0.518
Constant	...	...	0.987	-4.731	-1.669
Constant and Time Trend	...	...	0.987	-5.295	-1.615
BRAZIL					
(Five Lagged Difference Terms)					
None	0.998	-0.485	0.999	-0.316	-0.457
Constant	0.971	-1.572	0.975	-6.062	-1.608
Constant and Time Trend	0.992	-0.170	0.992	-1.680	-0.305
(Optimal Band Width)					
None	...	...	0.999	-0.310	-0.454
Constant	...	...	0.975	-5.622	-1.540
Constant and Time Trend	...	...	0.992	-1.629	-0.296
MEXICO					
(Five Lagged Difference Terms)					
None	1.000	0.652	1.000	0.077	1.167
Constant	0.980	-2.969	0.980	-5.084	-2.788
Constant and Time Trend	0.976	-1.626	0.996	-3.968	-1.026
(Optimal Band Width)					
None	...	...	1.000	0.078	1.276
Constant	...	...	0.980	-4.897	-2.850
Constant and Time Trend	...	...	0.996	-3.162	-0.865

Table A3 (cont'd). Univariate Unit Root Tests on Stock Prices, 1991 week 1 to 1995 week 9 1/ 2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
JORDAN					
(Five Lagged Difference Terms)					
None	1.000	1.263	1.000	0.075	1.277
Constant	0.988	-1.682	0.990	-2.482	-1.467
Constant and Time Trend	0.977	-1.503	0.984	-4.893	-1.360
(Optimal Band Width)					
None	...	...	1.000	0.075	1.392
Constant	...	...	0.990	-2.261	-1.453
Constant and Time Trend	...	...	0.984	-3.838	-1.166
KOREA					
(Five Lagged Difference Terms)					
None	1.000	0.436	1.000	0.024	0.267
Constant	0.986	-1.199	0.987	-3.421	-1.262
Constant and Time Trend	0.968	-1.891	0.968	-8.334	-2.197
(Optimal Band Width)					
None	...	...	1.000	0.024	0.275
Constant	...	...	0.987	-3.213	-1.220
Constant and Time Trend	...	...	0.968	-7.822	-2.139
MALAYSIA					
(Five Lagged Difference Terms)					
None	1.000	1.343	1.000	0.125	1.425
Constant	0.992	-1.179	0.993	-1.839	-1.128
Constant and Time Trend	0.971	-1.747	0.977	-7.093	-1.786
(Optimal Band Width)					
None	...	...	1.001	0.124	1.345
Constant	...	...	0.993	-1.992	-1.157
Constant and Time Trend	...	...	0.977	-7.787	-1.880

Table A3 (cont'd). Univariate Unit Root Tests on Stock Prices, 1991 week 1 to 1995 week 9 1/2/

Deterministic Components	Tests 3/				
	α	ADF	α	$Z(\alpha)$	$Z(t)$
THAILAND (Five Lagged Difference Terms)					
None	1.000	1.047	1.001	0.143	1.449
Constant	0.993	-0.964	0.993	-2.041	-1.127
Constant and Time Trend	0.944	-2.927	0.965	-11.686	-2.383
(Optimal Band Width)					
None	...	...	1.001	0.144	1.520
Constant	...	...	0.993	-1.893	-1.096
Constant and Time Trend	...	...	0.965	-10.351	-2.238

1/ The augmented Dickey-Fuller (1979, 1981) and Phillips-Perron (1988) tests are based on the following model for any series x: $x_t = \mu + \beta t + \alpha x_{t-1} + \epsilon_t$. The tests rely on rejecting a unit root ($\alpha = 1$) in favor of stationarity. The H_0: $\alpha = 1$ is tested by the ADF, $Z(\alpha)$ and $Z(t)$ tests. If the null hypothesis cannot be rejected (the test statistics are not significant), then the series has a unit root. The value of α is also presented when using either test. The Phillips-Perron $Z(\alpha)$ and $Z(t)$ tests are preferable to the ADF test more commonly reported in the literature, as they are robust to a wide variety of serial correlation and time-dependent heteroscedasticity. The ADF test result is also reported here, as it is the univariate version of the Johansen method used in this paper to test for cointegration (see Section V).

2/ All series are in logs. Consistent with our approach in testing for cointegration (see Section V), stationarity tests were carried out using five lagged difference terms, and are denoted as "Five Lagged Difference Terms". Andrew's (1991) data-dependent automatic bandwidth selector is also used to determine the optimal lag truncation order, and these test results are denoted as "Optimal Band Width". Both tests were carried out using the Fejer kernel and the prewhitening technique of Andrews and Monahan (1992).

3/ The critical values for respectively, the $Z(\alpha)$ test, and (both the) ADF and $Z(t)$ tests at the 5 percent level of significance are: -8.0 and -1.95 for no deterministic components ($\mu=\beta=0$), -13.8 and -2.89 for the constant term only ($\mu=0$, $\beta\neq0$) and -21.1 and -3.43 for both deterministic components ($\mu\neq0$, $\beta\neq0$), and are taken from Fuller (1976).

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