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INTRODUCTION

Given the recent developments in the liberalization and internationalization of the Philippine capital market, it is important to re-examine the issue of Philippine equity market integration with the international equity markets. In particular, this study will investigate the impact of the domestic capital market liberalization on the degree of integration of the equity market of the Philippines with those of its top sources of foreign investments and major trading partners in the Asia-Pacific region, namely, Taiwan, Japan, Hong Kong, Singapore and the United States.
One notion of capital market integration is as follows. Capital markets are fully integrated when financial assets with the same risk but traded in different markets have identical expected returns in some common currency. However, it is possible that the degree of integration of a particular market may not be complete because of the presence of barriers to international capital flows. The degree of integration depends on factors that determine the ease of arbitrage or cross-border trades in financial assets across international markets. This is because the barriers, if substantial and effective, can prevent cross-border arbitrage activities that eliminate any excess return relative to the equilibrium return dictated by some common global factor. Consequently, financial assets in these markets can have different expected returns even when their risk characteristics are the same. There is also a greater likelihood of transmission of market turbulence from one country to another if international capital markets are integrated. Intuitively, if international capital markets are integrated, the resultant ease of internationally diversifying an investor's portfolio reduces the sensitivity of the portfolio's return to local events. That is, unfavorable news in one country may be neutralized by positive news in another country. However, if markets are segmented, the investor's portfolio may be significantly more sensitive to local information than global information and as such required returns can be different across markets. This study attempts to find evidence on the degree of integration of the Philippine capital market and international capital markets using national equity markets data; i.e., equity market integration.

Bekaert (1995) distinguishes among three types of barriers that can affect the degree of equity market integration. The first group is comprised of direct restrictions on foreign ownership as well as foreign exchange and capital controls. The second group includes indirect barriers related to the regulatory and accounting environment such as lack of information on these markets and on the health of the companies, the inefficiency and slowness of settlement systems, poor accounting standards, and the fear of expropriation due to minimal investor protection. The third group consists of barriers arising from market-specific risks that discourage foreign investment including liquidity risk, political risk, macroeconomic instability, and currency risk. However, Bekaert argues that the presence of country funds and/or cross-listed securities might promote equity market integration despite the existence of severe restrictions on direct foreign equity ownership.
The lowering of international investment barriers due to liberalization of direct and indirect impediments to foreign investor participation in the Philippine equity market in recent years leads one to infer that the linkage and resultant integration of the Philippine equity market with international equity markets has increased. However, because of remaining barriers to international investment, there is reason to believe that it is not fully integrated with the world capital markets.

II. LITERATURE REVIEW

Recent literature examining the issue of capital market integration has focused on developing or emerging country equity markets. Indeed, there are studies that provide evidence that the Philippine equity market as well as a number of developing countries that have liberalized their equity markets have become integrated with the global capital market. The existing literature that has investigated the degree of integration of the Philippine stock market with international stock markets can be classified into two categories. The first group examines equity market integration assuming an underlying asset pricing model. The second category of studies investigates the degree of integration based on correlations.

Asset Model Integration Studies

Buckberg (1995), using monthly U.S. dollar stock market index excess return data obtained from the International Finance Corporation (IFC)-constructed emerging stock market indexes for the period 1985 to 1991, investigates the extent of integration of 20 emerging stock markets into the global financial market via tests of a conditional International Capital Asset Pricing Model (ICAPM) with time-varying expected returns and constant conditional proportionality (beta). The ICAPM posits that if emerging markets are part of a global market, then each market's expected returns should be proportional to that market's covariance with the world portfolio. The results of the tests of the conditional ICAPM revealed that the model cannot be rejected in 18 of the 20 emerging markets in her sample, including the Philippines. She concludes that these markets were integrated into the global capital market during the period 1985 to 1991. Moreover, since only two of the emerging markets banned or severely restricted foreign investment during part of this period, Buckberg conjectures
that rising capital flows from industrial economies was evidently the mechanism of integration.

Note, however, that Buckberg's results depend on the validity of the implicit assumption that world markets are perfectly integrated and that the ICAPM is sufficient to explain the cross-section of expected returns in both emerging and industrialized equity markets. Harvey (1995a) suspects that not all emerging markets are fully integrated into the world capital markets. He documents the varying degrees of direct and indirect barriers that confront foreign investors in the 20 emerging markets in Buckberg's sample. Based on the stylized facts, he concludes that the degree of integration varies across different countries. As such, he doubts whether any asset pricing model that assumes complete integration of capital markets would be able to completely explain the behavior of security prices in these emerging markets.

To illustrate his point, Harvey (1995b) tested a single-factor international asset pricing model with constant expected returns and risks, using monthly U.S. dollar return (not excess returns) data on the same 20 emerging stock market indices as constructed by IFC covering the period 1985 to 1992. The single factor he uses is the excess returns on a world market portfolio constructed by Morgan Stanley Capital International (MSCI). Among his findings is that for the Philippines, the single-factor model cannot be rejected. However, while the estimate of the beta on the MSCI world market portfolio is significantly different from zero, the coefficient is less than one. Furthermore, the $R^2$ of the regression is quite low suggesting that the model is inadequate in characterizing this market's returns. Taken together, the evidence indicates that the Philippine stock market is not well integrated into the global capital market. However, he finds evidence that the Philippine stock market's risk exposure to the world market portfolio changes over time. His estimates of a five-year rolling correlation measure of the local market return and the MSCI excess returns indicate that correlations increased progressively reaching 40 percent in the case of the Philippines. This suggests that this market may be becoming increasingly integrated. Overall, he finds that the model can be rejected for 13 out of the 20 emerging markets in his sample implying that these markets are segmented from the global economy. Harvey (1995b) concludes from his evidence that the single-factor model such as the world beta model
which assumes complete integration is not sufficient to characterize expected returns in emerging markets.

Harvey (1995c) further explores the hypothesis that the emerging markets are not completely integrated into the global capital markets by examining the influence of local information and global information on the predictability of returns in emerging equity markets. Using the same data as in Harvey (1995b), he finds that in the case of the Philippines, the results of separate bivariate linear regressions of the market’s excess returns on the world information variables and on the set of information specific to the Philippine market reveal that both sets of information significantly influence the market’s excess returns. Similarly, when local information is combined with the world information, the combined information significantly influences the Philippine market’s excess returns. The results of the test of the null hypothesis of exclusion of the local information variables in the Philippine regression equation indicate that the null can be rejected, suggesting that Philippine stock market returns are importantly influenced by local information. Moreover, the regression results indicate that the proportion of variance due to local information is greater than the proportion of variance due to world information suggesting that slightly more than one half of the predictable variance in this market’s returns is induced by local information. In contrast, a similar study by Harvey (1991) finds that most of the variation in the developed country expected returns is being driven by global information variables. The evidence that the predictability in Philippine stock market returns is almost equally influenced by local information and global information suggests that the Philippine stock market is partially integrated with the world capital markets. The result is similar for most of the other emerging stock markets, except that the variation explained by local information are even greater in magnitude. Harvey concludes that these findings further put into question the results of studies using any asset-pricing model that assumes complete market integration.

**Correlation Studies**

Bekaert (1995), using monthly dollar index excess return data from 1985 to 1992, examines the degree of integration of 19 emerging equity markets using an expected return-based correlation measure that does not assume an asset pricing model nor complete world capital market integration. Bekaert’s approach
is similar to the world latent factors model used by Campbell and Hamao (1992) in studying the long-term capital market integration of the U.S. and Japan. In his study, the world factor is assumed to be captured by stock market and interest rate variables of the United States.

Bekaert finds that among the emerging markets, the Philippines exhibits one of the highest significant correlations, with the magnitude of the correlation being comparable with those of the industrialized markets. Overall, the null hypothesis of no correlation cannot be rejected. He concludes from his results that there must exist global news factors affecting these markets simultaneously. Moreover, his evidence suggests that these markets are integrated into the world market although the degree of integration varies across these markets. Bekaert further provides some evidence of a trend toward increasing integration of equity markets by estimating the correlations from an earlier sample (1976 to 1985). However, this does not include the Philippines as data was not available prior to 1985.

Although Bekaert’s measure does not suffer from the pitfalls of tests of market integration that assume an equilibrium asset pricing model, he cites that one limitation of his measure of market integration is that it works as a perfect measure of integration only in a one-factor world with constant risk exposures. This is because if there was only one source of risk and markets were perfectly integrated, expected returns would be perfectly correlated. However, it is unlikely that only one risk factor explains all of the cross-section and time variation in equity returns. But then, Bekaert argues that it is equally unlikely that expected return correlations are low in perfectly integrated markets. Hence, high expected return correlations estimated based on his methodology may convey information, indirectly, about the degree of market integration. Another problem with Bekaert’s approach is an implicit assumption that he makes in his estimation of the correlations of expected returns. He assumes that the Vector Autoregression (VAR) framework generates the expected returns correctly. Consequently, if there is measurement error in the resulting expected return estimates that is uncorrelated across the U.S. and the emerging markets, the estimated correlations will overestimate the true degree of expected return correlation. In order to address these issues, Harvey and Bekaert (1995) examine the extent of emerging market integration using a pricing model which allows for time-
varying market integration. They find that a number of emerging markets exhibit time-varying integration with some markets appearing more integrated than one might expect based on prior knowledge of investment restrictions. On the other hand, they find that other markets appear segmented even though foreigners have relatively free access to their capital markets. However, the Philippine equity market is not in their sample of emerging markets, possibly owing to the lack of data required for their tests.

It should be noted that although the above studies indicate some evidence that the Philippine equity market has become integrated with the world capital market, the results appear to point out that the degree of integration is weak; i.e., the Philippine equity market is not fully integrated into the world market. Specifically, the results of these studies seem to point out the increase in the sensitivity of the returns on Philippine equities to global factors in the latter 1980s and early 1990s and that capital market liberalization reforms instituted during this period may have been partly responsible for this. However, stylized facts pointing to remaining barriers to international investment in the Philippines preclude us from concluding that the integration is complete.

III. SCOPE AND LIMITATIONS OF THE STUDY

In order to avoid the problems and the difficulties of measuring equity market integration that were encountered in the previous studies, a notion of integration that does not necessitate an asset pricing model nor assume complete world market integration for its empirical implementation and testing, was utilized. This concept of integration involves investigating whether there exists a long-run equilibrium relationship among international stock markets that is reflected in long-run common movement among the levels of the each country's stock market indexes (not returns). In turn, this long-run relationship is given by the cointegration properties of the national stock market price indexes. Under this concept, finding a cointegration relationship among the national stock market indexes is taken to be suggestive of market integration. A detailed explanation of this measure of market integration is provided in the following section.

While the subset of international markets included in the study is smaller compared to the previous studies, it is more defined in the sense that it considers only the countries with which the Philippines has close political and economic ties and which are
The effect of capital market liberalization measures
g eographically proximal. For example, all of these economies are
members of the Asia Pacific Economic Cooperation (APEC); the
Philippines and Singapore are both members of the Association
of Southeast Asian Nations (ASEAN); Hong Kong, Singapore and
Japan are the major financial centers in Asia; and, Hong Kong,
Singapore and Taiwan are three of the four Newly Industrializing
Economies (NIEs) in Asia. The U.S., Japan, Hong Kong, and
Singapore have relatively more open, developed, and mature equity
markets. Although Taiwan still has a relatively restrictive foreign
investment policy regime, it has embarked on some reforms aimed
at opening up its domestic equity markets to foreign investors since
the latter part of the 1980s and it is one of the Asia-Pacific region's
fastest growing stock markets in terms of market capitalization
and turnover. Thus, it may be reasonable to expect stronger
results, either in favor of or against the hypothesis about the extent
of Philippine equity market integration. However, it must be
emphasized at the outset that the conclusions reached in this study
are confined to the stock markets in the sample. For example, a
finding that the Philippine stock market index is not integrated
with the other five international stock markets does not imply that
the other stock markets are not integrated into the world market in
general.

Unlike the previous works, the paper utilizes market data that
cover a longer span of time, January 1980 to December 1995.
The previous studies’ empirical evidence on the degree of
integration of the Philippine equity market into the global market
has been inferred from monthly returns data covering a relatively
short period of time, from 1985 to 1992. It should be noted that
further opening up of the market occurred after 1992 and the impact
of this is not reflected in their results. Moreover, since data prior
to 1985 was not available for the Philippines in these studies, not
much can be said about the impact of liberalization on the degree
of Philippine equity market integration. This issue will be addressed
in this paper by investigating the impact of the removal and/or
relaxation of capital controls in the Philippines on the degree of
integration of the domestic stock market with the other five
international stock markets in the sample.
IV. THEORETICAL FRAMEWORK: COINTEGRATION AND EQUITY MARKET INTEGRATION

A conventional measure of the degree of integration among international stock markets is the simple cross-country correlation of stock market index returns. Under this approach, an increase in return correlation is interpreted as evidence of an increase in the degree of market integration as well as an indicator of a reduction in the benefits to international diversification. Aside from this economic rationale, there is also a statistical motivation for the use of returns. A source of nonstationarity of equity prices is that the series contains a unit root (or stochastic trend). A statistical implication of a unit root in the series is that most of the distribution moments are undefined. Consequently, the conventional hypothesis tests cannot be performed. Given that the stock prices series has a single unit root, then first differencing can render the series stationary. Thus, providing another reason for the use of returns or first difference of the (natural log) of stock prices.

However, long-run information on the interrelationship of the international stock markets that are reflected in the price levels is lost by differencing. This is because return correlations are influenced by both independent short-run trading noises as well as by long-run fundamental relationships among the markets which are induced by their internationalization and liberalization. Consequently, the short-run noise can possibly make the markets appear more independent than they truly are.

However, econometric developments now enable one to determine whether or not a long-run statistical relationship among nonstationary variables exists. In particular, it is now possible to examine the hypothesis of a long-run relationship among levels of national stock market indexes based on the concepts of cointegration and its relationship to error correction as introduced by Engle and Granger (1987). The intuition behind their concept is as follows. If two national stock market indexes each follow a stochastic trend over time, then in general they will wander apart. However, cointegration means that the two nonstationary indexes are tied together by some long-run equilibrium relationship, and thus, they cannot drift apart indefinitely. This in turn implies that the deviations from their equilibrium relationship must be stationary; i.e., the deviations have bounded fluctuations about a fixed level. Mathematically, the long-run equilibrium relationship is given by some linear combination of the two nonstationary variables such
that the combination of the two nonstationary series is stationary. If this were the case, then, following Engle and Granger (1987), the two national stock market indexes are said to be cointegrated.

Since it is possible that two stock markets have a long-run relationship with other stock markets as well, then the equilibrium relationship must involve several stock market indexes. If a long-run relationship among these nonstationary price series truly exists, then, the (stationary) deviation from the long-run relationship can only be constructed from a combination of the stock market indexes involved. This is known as multivariate cointegration.

In this study, a finding of a cointegration relationship among the Philippine stock market index and the national stock market indexes of Taiwan, Japan, Hong Kong, Singapore and the U.S. is interpreted as evidence that these equity markets are integrated. This is because cointegration of national stock market indexes implies that a long-run equilibrium relationship exists among these indexes. In turn, the long-run equilibrium relationship drives the national stock market indexes to move together in the long-run even if they don’t in the short-run and even if these indexes are individually nonstationary. Given that long-run co-movement of stock market prices of these countries is suggestive of equity market integration, then cointegration implies market integration. This reasoning is similar to that used by Campbell and Hamao (1992) and Bekaert (1995) except that they conjecture on the co-movements in expected returns instead of co-movements in the levels of the stock market indexes.

It seems reasonable to infer that a finding of common movement in prices suggests integration. Common movement in prices possibly arises from the response of each country to some economic force, perhaps a common world growth factor, which systematically affects the equilibrium prices of the stocks of the different countries. On the other hand, the strength of the co-movement depends on the degree of openness of the international capital markets. This is because the factors that promote market integration, such as liberalization of capital markets and the globalization of securities, also promote international capital mobility. Ripley (1973) argues that, with freer capital mobility, capital flows will tend to reduce interest rate differentials between countries by increasing the supply of capital in the country with the high interest rate, and reducing it in the country with the low interest rate. Given that stock prices are affected by interest rate
movements, assuming that the present value model of stock prices holds, the equalization of national interest rates via the interest rate parity condition will result in co-movement between national stock price indexes.

Following Taylor and Tonks (1989), let $S_i$ and $S_j$ be the (natural) logarithm of the stock market price indexes of country $i$ and country $j$ at time $t$, respectively. Suppose that the stock market price indexes of countries $i$ and $j$ are perfectly correlated in the long-run. This implies a linear relationship between $S_i$ and $S_j$, i.e., $S_i = a + bS_j$ for some scalars $a$ and $b$. However, this relationship may be distorted in the short-run by the joint effect of country-specific factors, given by the stationary disturbance $e_i$; i.e., $S_i = a + bS_j + e_i$. By translating the stock market price indexes into short-run stock market returns, the noise in the system is probably increased. Consequently, even though a long-run relationship exists between the levels of the stock market price indexes, only a weak degree of return correlation may be statistically observed. Therefore, using co-movements of short-term (expected) returns on the national stock market indexes, instead of co-movements of the levels of these indexes, may understate the true degree of integration of these national stock markets.

Additionally, because liberalization and globalization of securities can facilitate cross-country investing or arbitrage activities, then the increased participation of foreigners in domestic equity markets can potentially strengthen the linkage between local and foreign markets and cause international stock prices to move together. For example, negative events in foreign countries can affect stock prices in these countries and therefore the liquidity of investors residing in these countries. Consequently, this affects the relative returns of these foreign investors' investment in other countries and induce them to adjust their portfolios. When foreign investors own a substantial proportion of the tradable stocks in the domestic equity markets, then their transactions can significantly influence the domestic equity prices and thus cause co-movements in the stock prices of the foreign and domestic markets. Additionally, it can be the case that domestic investors have no information on why the foreign investors are adjusting their portfolios and so domestic investors will tend to react to such changes. Such reactions can amplify the effect of foreign
disturbance on the domestic equity market and thus give rise to co-movement in the foreign and domestic stock markets' prices.

However, a weakness of cointegration as a measure of equity market integration is that the underlying co-movement of stock prices in different countries may be due to factors other than those which facilitate the cross-country investing or arbitrage activities. That is, there may be some forces that make stock prices in different countries positively correlated even if there are no international financial transactions. Likewise, national stock markets could be segmented but subject to common shocks that move stock prices in these countries in similar fashion. For example, Ripley (1973) cites that "countries whose incomes move in similar manner may have stock prices that also move together." This is because movements in income affect expectations about future economic developments and affect investors' abilities to purchase equities. Thus, similar movement in incomes may result in an indirect link between the stock prices in the two countries even if no foreigner can buy stocks in either country.

Furthermore, there has been recent debate on the implication of cointegration on financial market efficiency. The controversy arises from the error-correction mechanism implied by cointegration and the claim that stock price changes should not be predictable in an efficient market. Note that, following Engle and Granger (1987), if two national stock market indexes are cointegrated, then the relationship between these two variables can instead be expressed in terms of an error-correction model (ECM). An ECM for these stock market indexes relates the changes in the variables to lagged changes in both variables and a lagged linear combination of the levels. The linear combination of levels that enters the ECM is just that combination which is stationary in levels. The ECM can be thought of as the process by which the variables in the system being analyzed respond to the long-run equilibrium error in order to eliminate it. However, this implies that the stock market price changes of at least one of the stock market indexes are predictable from the previous period's linear combination of the two indexes. This is contradictory to the semi-strong form efficient markets hypothesis with constant expectations.

For example, Richards (1995) argues that if markets are efficient, then national stock market indexes cannot be cointegrated. He shows this by imposing the cointegration condition on the stock market indexes of two countries (assumed to contain
a unit root) and then assumes that excess returns on the stock market indexes are generated by the CAPM with time-varying risk. However, it is possible that the results of Richards are driven by his assumption of CAPM as the underlying asset pricing model. It is well known that any test of market efficiency is a joint test of the underlying equilibrium asset pricing model and market efficiency. Bossaerts (1988), for example, demonstrates that asset prices can be cointegrated in an economy where a representative agent is governed by rational expectations. He shows that in such an economy, asset returns are predictable, yet each agent is behaving optimally. Furthermore, Dwyer and Wallace (1992), using absence of arbitrage profits as a definition of market efficiency instead of unpredictability of returns, show that there is no general equivalence between market inefficiency and cointegration, or for that matter, a lack of cointegration. That is, the absence of cointegration is neither necessary nor sufficient for market efficiency when efficiency is not defined by the lack of predictability but by the absence of arbitrage profits, i.e., when one cannot realize abnormal profits from the forecast. They illustrate this by using different models and their analysis shows that different models give rise to different implications and that either finding cointegration or a failure to find it can be consistent with market efficiency. A similar argument is used by Baffes (1994) in the case of currency markets. He shows that cointegration between exchange rates does not necessarily imply market inefficiency since market efficiency does not rule out predictable exchange rate movements, but only rules out arbitrage opportunities from predictable exchange rate movements.

V. DATA

The empirical analysis in this paper is based on monthly data on national stock market price indexes of the Philippines, Japan, Hong Kong, Singapore, Taiwan, and the U.S. covering the period January 1980 to December 1995. The following value-weighted indexes are used: Philippine Stock Exchange Commercial and Industrial Index (Philippines), Tokyo Stock Exchange Price Index (Japan), Hang Seng Index (Hong Kong), Stock Exchange of Singapore All-Shares Index (Singapore), the Taiwan Stock Exchange Weighted Price Index (Taiwan), and the Standard & Poors' 500 Index (United States). All of these indexes are denominated in local currency. Foreign exchange rate data were
collected for the Asian countries in order to transform the local stock market price index series into prices denominated in U.S. dollars. All analysis in this study is based on the natural logarithm of the stock market price indexes in U.S. dollars, denoted PHL (for the Philippines), TWN (for Taiwan), JPN (for Japan), HK (for Hong Kong), SIN (for Singapore) and US (for the U.S.). The stock market price indexes data are month-end values and exchange rate data are month-end market mid-point rates. Appendix A shows the sources of the data for the various countries in the sample.

Month-end stock price index data are used in order to reduce the possible biases arising from infrequent trading and nonsynchronous trading of some of the component stocks of the stock market index as well as day-of-the-week effects. Such biases may be more pronounced when daily or weekly data are used (Lo and MacKinlay, 1988). Shiller and Perron (1985) and Perron (1991) point out that the power of unit root tests and cointegration tests is mainly a function of the span of the data (i.e., length of time period) and only slightly depends on the number of observations. More specifically, they show that for a given span, additional observations obtained by increasing the frequency of the data (e.g., moving from monthly to weekly or daily) increases the power of the test only marginally, with the rate of increase declining as the sampling interval is decreased. In addition, Hakkio and Rush (1991) stress that since cointegration is essentially a long-run concept, then, it requires long spans of data to give the tests for cointegration much power. They showed that increasing the sampling frequency without increasing the length of the span yields little additional information when testing whether two series are cointegrated, but that the converse is true when the span of the data is increased using the same sampling frequency.

U.S. dollar-denominated stock market price indexes are used not just to make the indexes directly comparable but in order to account for exchange rate fluctuations or risk. Apart from the local stock market returns, exchange rate risk is also important for overseas investors who wish to diversify internationally since they have to transact through the exchange rate mechanism both at the point of time they purchase foreign securities and at the point of time that they sell these same securities. Moreover, we adjust for exchange rate fluctuations since exchange rate fluctuations possibly reflects the impact of reforms aimed at liberalizing foreign exchange regulations, which in turn contribute to capital market
integration. Also, the results of this study can be directly compared
with the previous studies that have used U.S. dollar-denominated
data in examining the degree of integration of the Philippine equity
market with the world capital market.

VI. METHODOLOGY

Cointegration Tests

In this study, we use the maximum likelihood estimation
procedure based on Johansen (1988, 1991) and Johansen and
Juselius (1990) in testing for cointegration and estimating
cointegration relationships among the six national stock market
indexes. Cheung and Lai (1993) and Gonzalo (1994) show that
this procedure has good finite sample properties. We will refer to
this approach simply as the Johansen procedure in the remainder
of the paper.

Johansen’s procedure examines cointegration based on the
technique of reduced rank regression in the VAR framework. Let
$S_t$ be the $(nx1)$ vector system of the (natural logarithm) national
stock market price indices in levels, $(S_{1t}, S_{2t}, \ldots, S_{nt})'$, where each
national stock market price index is possibly nonstationary. The
maintained assumption in this approach is that the vector system,$S_t$, is generated by a VAR($j$) in levels with $j$ lags being sufficient to
summarize all the dynamic correlations between the elements of
the vector system. That is,

$$S_t = \mu + \Pi_1 S_{t-1} + \Pi_2 S_{t-2} + \ldots + \Pi_j S_{t-j} + \epsilon_t, \quad (1)$$

where $\mu$ is an $(nx1)$ vector of constants or drift terms, the
$(nxn)$ matrix $\Pi_k$ represents the matrix of autoregressive coefficients
for $k = 1, 2, \ldots, j$, and the innovation sequence is assumed to be an
i.i.d. Gaussian process, i.e., $\epsilon_t \sim \mathcal{N}(0, \Omega)$.

Johansen (1988) and Johansen and Juselius (1990) show
that, without loss of generality, any VAR of the form given by
Equation (1) can be written as

$$\Delta S_t = \Gamma_1 \Delta S_{t-1} + \Gamma_2 \Delta S_{t-2} + \ldots + \Gamma_j \Delta S_{t-j} + \mu + \Pi S_{t-1} + \epsilon_t, \quad t = 1, 2, \ldots, T. \quad (2)$$
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where, $\Delta$ is the difference operator, $T$ is the number of observations per variable exclusive of the lags, $\Pi$ is an $(n \times n)$ matrix of coefficients with $\Pi = (I_n - \Pi_1 - \Pi_2 - \ldots - \Pi_T)$, $\Gamma(1)$ is the autoregressive matrix polynomial $\Gamma(L) = I_n - \Gamma_1 L - \Gamma_2 L^2 - \ldots - \Gamma_L L^L$ evaluated at $L = 1$ where $L$ is the lag operator, and the $(n \times n)$ matrices $\Gamma_s = [-\Gamma_{s+1} + \Gamma_{s+2} + \ldots + \Gamma_L]$, for $s = 1, 2, \ldots, j-1$, of unknown parameters.

Equation (2) is just the traditional VAR in first differences except for the term $\Pi S_{t-1}$. Within this parameterization, the short-run dynamics are described by the matrices $\Gamma_s$ while all the information on the long-run relationships or cointegration vectors among the $n$ national stock market price indexes in the vector system $S_t$ is given by the so-called long-run impact matrix $\Pi$. In this study, the appropriate lag order, $j$, is determined using a sequential likelihood ratio statistic with an adjustment for small sample bias (see J. Hamilton, 1994, p. 297), while ensuring that the residuals are uncorrelated.

Johansen's procedure is essentially one that finds the rank of the matrix $\Pi$. Johansen shows that the hypothesis of cointegration can be formulated as the hypothesis of reduced rank of the matrix $\Pi$. Specifically, Johansen shows that the number of distinct cointegration vectors or stationary relationships which exists among the variables of the vector system $S_t$ will be given by the rank of the matrix $\Pi$, $r$. There are three possible cases:

1. If $r = 0$, then the vector system $S_t$ is fully nonstationary. In this case the elements in the vector system $S_t$ are not cointegrated and thus there is no long-run relationship among the six national stock market price indexes. In this case, the traditional VAR in differences (Equation (2) without the term $\Pi S_{t-1}$) is well-specified.

2. If $r = n$, then the vector system $S_t$ is fully stationary. This is the case when each series in the vector system is stationary.

3. If the $0 < r < n$, then there will be only $r$ linear combinations of $S_t$ that are stationary and thus there are $r$ cointegration relationships among the six national stock market indexes. In this case, there exists an $n \times r$ matrix $\alpha$ such that $\Pi = \alpha \beta'$, where $\beta$ is an $n \times r$ matrix of full rank $r$ such that $\beta' S_t$ is stationary even though $S_t$ is nonstationary. Therefore, the cointegrating relations $\beta' S_t$ can be interpreted as stationary relationships among the nonstationary
levels of the variables. The $r$ columns of matrix $\beta$ are the
cointegration vectors $(b_1, b_2, \ldots, b_r)$, which contain information on
the equilibrium relationships that dictate the long-run movement
of the variables in the system $S_r$.

Given that $r$ cointegration relationships are found among the
variables in $S_r$, Equation (2) can be considered as a vector error-
correction model and can be written as

$$
\Delta S_t = \Gamma_1 \Delta S_{t-1} + \Gamma_2 \Delta S_{t-2} + \ldots + \Gamma_r \Delta S_{t-r} + \mu + \alpha \beta' S_{t-1} + \epsilon_t, \quad t=1,2,\ldots,T \quad (3)
$$

or

$$
\Delta S_t = \Gamma_1 \Delta S_{t-1} + \Gamma_2 \Delta S_{t-2} + \ldots + \Gamma_r \Delta S_{t-r} + \mu + \alpha z_{t-1} + \epsilon_t, \quad t=1,2,\ldots,T. \quad (4)
$$

In Equation (4), if $z_{t-1} = \beta' S_{t-1} < 0$, then it is interpreted as the
long-run equilibrium error that describes the short-run deviations
of the variables from the $r$ distinct stationary or long-run equilibrium
relationships. The $n \times r$ coefficients in the matrix
$\alpha = (a_1, a_2, \ldots, a_r)$ can be interpreted as the average rate of
reversion or speed of adjustment of the variables in the vector
system toward the long-run equilibrium relationships underlying
the cointegration vectors. These coefficients measure the current
period's correction of the last period's deviation in order to maintain
the long-run equilibrium relationship. A low magnitude indicates
slow adjustment while a large coefficient indicates rapid
adjustment. The intuition behind the error correction equations
is that if the six stock market price indexes in the vector system
are cointegrated, then the short-run changes in each country's
stock market prices are due to the effects of the lagged changes
in own and other countries' stock market prices, the previous
period's equilibrium error, and random factors. The last period's
equilibrium error enters the error-correction equations to capture
the effect of short-term deviations from the long-run equilibrium.
As such, the error-correction equations can be interpreted as the
process by which the national stock market indexes in the
cointegrated system respond to the long-run equilibrium error in
order to eliminate it.

Johansen and Juselius (1990), Dickey et al. (1991) and
DeFusco et al. (1996) suggest that the results of cointegration
test are stronger and more robust if more than one significant
cointegration vector is found. For example, DeFusco et al. (1996)
argue that cointegration vectors can be conceptually interpreted
as constraints that an economic system imposes on the movement of the variables in the long run. Thus, for an $n$-variable system, if there is only one cointegration vector (or $n - 1$ unit roots), then the system can deviate in $n - 1$ independent directions and is stable in only one direction (i.e., there is only one direction where the variance is finite).

This study employs two alternative likelihood ratio tests of cointegration suggested by Johansen (1988) and Johansen and Juselius (1990). The first is the trace test statistic which tests the null hypothesis that there are at most $r$ cointegration relationships against the general alternative that there are $n$ cointegration relationships. An alternative likelihood ratio test statistic is the maximum eigenvalue test statistic, which tests the null hypothesis of $r$ cointegrating relations against the alternative of at most $r + 1$ cointegrating relations. The distribution of these test statistics follow a $(n-r)$ dimensional standard Brownian motion and the simulated asymptotic critical values are provided by Johansen and Juselius (1990) and Osterwald-Lenum (1992).

In this study, a finding of a cointegration relationship among the Philippine stock market index and the national stock market indexes of Taiwan, Japan, Hong Kong, Singapore and the U.S. is interpreted as evidence that these equity markets are integrated.

**Pre- and Post-Liberalization Subsample Analysis**

Since liberalization of the domestic capital market helps promote its integration with other international stock markets, we expect a priori that the Philippine stock market will be more integrated in the post-liberalization period than the period prior to its market opening. Given that cointegration suggests integration, the change in the degree of integration of the Philippine stock market can be examined by comparing the number of cointegration relationships or vectors in subsample periods. Because various policy changes and events relating to the internationalization and liberalization of the Philippine market came into effect during the period January 1980 to December 1995, it is quite difficult to pinpoint an exact date for the regime shift. However, Buckberg (1995), De Santis and Imrohoroglu (1995), and Kim and Singal (1993) identify October 1989 as the initial equity market opening up date for the Philippines. On this date, the first closed-end country fund devoted to Philippine securities was admitted in London’s stock exchange. Henceforth, more major reforms to open up the
domestic capital market have been implemented by the Philippine government. As such, changes in the degree of integration of the Philippine stock market into the five other international stock markets will be investigated by examining the changes in the number of cointegration relationships between two subsample periods: the pre-liberalization subsample covering January 1980 to September 1989 and the post-liberalization subsample covering the period October 1989 to December 1995. A finding of more cointegration vectors in the second subsample is then taken as an indication that the Philippine stock market has become more integrated with the other five international stock markets.

The approach used here is similar to that of Chou, Ng and Pi (1994) except that in this study, the break in the overall sample is determined solely by events relating to one stock market. However, there are limitations that make this approach imprecise. First, it is difficult to isolate the impact of liberalization because there may be other confounding events that can potentially contribute to a change in the degree of integration, if ever there is one. Therefore, we cannot attribute any observed change solely to market opening. Secondly, cointegration is a long-run property of the data. Since the process of liberalization in the Philippine capital market since 1989 has been gradual and consequently the impact of liberalization may take time, the results obtained from a relatively limited post-liberalization subsample may not be entirely conclusive. Thirdly, recent studies have suggested that inferences based on the asymptotic critical values may be misleading in small samples. Cheung and Lai (1993), using Monte Carlo simulations, find that Johansen’s cointegration tests are biased toward finding cointegration more often than what asymptotic theory suggests. Similarly, Gregory (1994) shows that for small samples and a high number of explanatory variables, the test size (frequency of rejecting the null hypothesis of no cointegration when it is true) of Johansen’s cointegration tests is significantly higher than the test size of cointegration tests based on other methods. Specifically, he finds that as the VAR sample size \( T \) falls, or the number of variables \( n \) or lags \( j \) in the system increases, the tests are biased toward finding cointegration more often when the asymptotic critical values are used. As such, they suggested that the asymptotic critical values be adjusted upward. One adjustment method proposed by Reinsel and Ahn (1988, 1992) involves multiplying Johansen’s test statistics by a scaling factor given by \( (T-nj)/T \) and
comparing these adjusted values with their asymptotic critical values. Cheung and Lai (1993) show that an equivalent way is to multiply the asymptotic critical values by a scaling factor given by $T/(T-n_J)$. Cheung and Lai (1993) indeed find that the finite-sample critical values obtained using the Reinsel and Ahn method are a very significant improvement over the asymptotic critical values of Johansen's tests. Since the time series of national stock market indexes in both pre- and post-liberalization subsample periods are not relatively long, the finite sample critical values using the adjustment factor suggested by Reinsel and Ahn (1988) is used in this study in order to minimize the potential small-sample bias.

However imprecise it may be, there seems to be some valid reason to believe that this approach should be able to capture possible significant change in the long-run equilibrium relationship of the Philippine stock market with the other international stock markets. This is because the markets of the United States, Hong Kong, Singapore and Japan have been relatively more open during the entire sample period.\textsuperscript{12} On the other hand, despite the opening up of its equity market in January 1991 (see Buckberg, 1995; and, Kim and Singal 1993), access to Taiwanese equity market for foreign investors remains heavily restricted (see, e.g., Euromoney, 1994, 1995; and, International Society of Securities Administrators, 1994). Most importantly, the market opening event in October 1989 is not the sole event that increased the access of foreign investors into the Philippine domestic equity market. It must be emphasized at the outset that there is no claim that this is the sole reason for the change in degree of Philippine equity market integration, should there be any. Given that more reforms to liberalize and internationalize the domestic capital market and to reduce the barriers to international capital flows have been implemented by the Philippine government beyond this date, it is expected that these events can possibly have an impact on the degree of Philippine equity market integration as well.

**Other Hypothesis Tests**

Assuming that at least one cointegration vector is found in the above analysis, the final step in Johansen's cointegration procedure involves the estimation of the error-correction equations given by Equation (3) or alternatively Equation (4), using the maximum likelihood method. Accordingly, it is possible to test the following hypotheses on the parameters of the estimated error-
correction equations. A detailed discussion of the estimation procedure and the following tests is provided in Hamilton (1994).

To illustrate these tests, consider the trivariate Philippine-Japan-US system with one cointegration vector \((r=1)\) and two lags \((j=2)\). For this case, the error-correction equations for the stock market indexes of the Philippines (PHL), Japan (JPN) and the US are given by

\[
\begin{align*}
\Delta S_{PHL} &= \Gamma_{11} \Delta S_{PHL,-1} + \Gamma_{12} \Delta S_{JPN,-1} + \Gamma_{13} \Delta S_{US,-1} + \mu_{PHL} + \alpha_1 (\beta_1 S_{PHL,-1} + \beta_2 S_{JPN,-1} + \beta_3 S_{US,-1}) + \epsilon_{PHL}, \\
\Delta S_{JPN} &= \Gamma_{21} \Delta S_{PHL,-1} + \Gamma_{22} \Delta S_{JPN,-1} + \Gamma_{23} \Delta S_{US,-1} + \mu_{JPN} + \alpha_1 (\beta_1 S_{PHL,-1} + \beta_2 S_{JPN,-1} + \beta_3 S_{US,-1}) + \epsilon_{JPN}, \\
\Delta S_{US} &= \Gamma_{31} \Delta S_{PHL,-1} + \Gamma_{32} \Delta S_{JPN,-1} + \Gamma_{33} \Delta S_{US,-1} + \mu_{US} + \alpha_1 (\beta_1 S_{PHL,-1} + \beta_2 S_{JPN,-1} + \beta_3 S_{US,-1}) + \epsilon_{US},
\end{align*}
\]

\[\text{(5)}\]

**Test of the significance of the estimates of the short run parameters in the matrix } \Gamma_{ij}.

The elements of the matrix \(\Gamma_{ij}, \Gamma_{jk}\), summarize the short-run price dependence of the stocks traded in the national stock markets in the system. Each coefficient measures the impact on the short-run movements of a particular country's stock market index of lagged changes in own and other countries' stock market price indexes. For example, in the trivariate case above, \(\Gamma_{12}\) and \(\Gamma_{13}\) measure the impact of lagged changes in the stock market index of Japan and U.S. stock market index, respectively, on the current period's change in the Philippine stock market index. On the other hand, \(\Gamma_{11}\) measures own lagged changes on the current period's change in the Philippine stock market index. When the cross-price effects are found to be significant in a given country's error-correction equation, they indicate that the short run changes in a given country's stock market prices are influenced by independent or local information contained in the short-run changes in the stock prices of the other national stock markets. The statistical significance of each of the coefficients of the lagged first differences that appear in each error-correction equation can be tested using the conventional \(t\) - test.

**Long-Run Exclusion Test: Test of significance of the coefficients of the cointegration vectors in the matrix } \beta.

As mentioned above, the cointegration vectors contain information on the equilibrium relationships that dictate the long-run movement of the national stock market indexes in the system.
Although there is no straightforward interpretation of the coefficients of the cointegration vectors, their relative magnitudes can shed light on the importance of each of the national stock market indexes in obtaining the underlying long-run equilibrium relationships.

Consider the trivariate case. A finding that $\beta_1$ is not significantly different from zero together with $\beta_2$ and $\beta_3$ being statistically significantly different from zero is interpreted as evidence that the Philippine national stock market index is not important in obtaining the long-run relationships underlying the cointegration vector in the system and that only the Japanese and U.S. stock market indexes are needed to get the cointegration relationship. Alternatively, a zero cointegration coefficient for the Philippine stock market index implies that the long-run movements of this stock market index is not influenced by the equilibrium relationship underlying the cointegration vector that was found and that this vector is simply picking up the bivariate cointegration relationship between the Japanese and U.S. stock market indexes. The significance of the cointegration coefficient corresponding to a given national stock market index is determined by testing the null hypothesis that the coefficient is not significantly different from zero using a likelihood ratio test.

**Weak Exogeneity Test: Tests of restrictions on the adjustment speed coefficients in the matrix $\alpha$.**

As previously mentioned, the error-correction term $z_{t-1}$ describes the short-run deviations from the long-run equilibrium relationships underlying the significant cointegration vectors while the coefficients in the matrix $\alpha$ can be interpreted as the average rate of reversion or speed of adjustment of the national stock market indexes in the system toward the long-run equilibrium relationships.

Consider the trivariate example above. Suppose that all coefficients of the cointegration vector are statistically significant and so $z_{t-1} = \beta_1 S_{PHL,t-1} + \beta_2 S_{JPN,t-1} + \beta_3 S_{US,t-1}$. Now, $\alpha_1$ is the rate of reversion of the Philippine stock market index toward the equilibrium relationship governing the long-run movements of the stock market price indexes of the Philippines, Japan, and the U.S. If $\alpha_1$ is not statistically significantly different from zero, the Philippine stock market index is said to be weakly exogenous with respect to the cointegration vector; i.e., the Philippine stock market index is not influenced by the information on the long-run relationship.
underlying the cointegration vector. However, if $\alpha_1$ is statistically significant, then this implies that short-run changes in the Philippine stock market index partly reflect error-correcting price adjustments that maintain the long-run equilibrium relationship among the three national stock market indexes. In the Johansen framework, the test of weak exogeneity, which is equivalent to the test of the null hypothesis that the coefficient of the error correction term $z_{t-1}$ in a particular country's error correction equation is not significantly different from zero, is performed using a likelihood ratio test.

Another interpretation of a statistically significant $\alpha_1$ is that current changes in the Philippine stock market price index are predictable from the previous period's linear combination of the three national stock market indexes, $z_{t-1}$. As discussed in the theoretical framework, such predictability is not contradictory with market efficiency given that efficiency is defined by the absence of arbitrage profits. Moreover, Eckbo and Liu (1993) argue that predictability of returns is consistent with the general notion of market efficiency in a setting with time-varying expected returns. In a survey article, Fama (1991) argues that time-variation in expected returns may be a response to innovations which are common across different securities and markets. Among others, empirical studies by Keim and Stambaugh (1986), Fama and French (1989, 1990), Turtle (1991), Turtle, Buse and Korkie (1994), Whitelaw (1994), and Cheung et al. (1995) provide evidence of time-varying expected returns which are induced by predictable components of aggregate real economic activity. Along this line, it is therefore possible that the error-correction term, $z_{t-1}$, represents the changing market expectations.

VII. EMPIRICAL RESULTS

Univariate Unit Root Tests

The initial condition for the system of stock market price index levels to be cointegrated is that each stock market price index series be nonstationary. For each country's stock market price index, the approaches of Dickey and Fuller (1979, 1981) [Augmented Dickey-Fuller Test (ADF)] and Phillips and Perron (1988) [Phillips-Perron Test (PP)] are used to test the null hypothesis that the series contains a single unit root. The ADF test assumes that the series follows an autoregressive process with Gaussian i.i.d. innovations. On the other hand, the PP test
allows for a wide variety of heterogeneously distributed and weakly dependent disturbances which makes it robust to the presence of both serial correlation and heteroscedasticity. The critical values of the test statistics of both approaches are, however, identical.

Two forms of the ADF and PP tests are considered in this study. The first form, labeled as constant and trend, allows for possible deterministic time trend and non-zero constant mean under the alternative hypothesis. The second form, labeled constant no trend, allows for a non-zero constant mean but not a deterministic time trend under the alternative hypothesis. This test is appropriate when no evidence in favor of the time trend is found.

Table 1 shows the results of the unit root tests conducted on the levels of the six national stock market indexes measured in U.S. dollars. The results of the ADF and PP tests which include both a constant and time trend suggest that for both subsamples and for all national stock market indexes, the null hypothesis of a unit root cannot be rejected at the 0.05 level of significance and that there is no evidence of a trend in the first difference series. Since the time trend is found to be insignificant, we then tested the unit root null hypothesis against the alternative that the levels of a given stock market index is stationary around a linear trend. The results of the ADF and PP tests which include only a constant but no time trend indicate that the null of a unit root in the levels of all national stock market indexes cannot be rejected at the 0.05 level of significance in both subsample periods. Overall, these results suggest that, in both the pre-liberalization and post-liberalization subsample periods, all national stock market indexes are nonstationary.

**Multivariate Cointegration Test Results**

Given the preceding results, we then conducted Johansen's tests of cointegration for both subsamples. As an initial step, the unrestricted VAR in error-correction form (Equation (2)) was estimated for each subsample using various lag lengths \(j\), beginning with a lag length of one. An unrestricted vector of constants was included in the estimation. Based on the sequential likelihood ratio test procedure, lag lengths of \(j = 3\) for the pre-liberalization subsample and \(j = 2\) for the post-liberalization subsample were found to be sufficient. Table 2 summarizes the results of the multivariate and univariate diagnostic tests on the residuals of the
estimated unrestricted VAR in error correction form for both subsamples.

For the pre-liberalization period, the results of the multivariate tests of autocorrelation indicate that for the entire system, the null hypotheses of zero first order autocorrelation and of zero fourth order autocorrelation cannot be rejected at the 0.05 level of significance. The evidence of serial correlation-free residuals are also found using the univariate Ljung-Box portmanteau test statistics. However, the multivariate normality test statistic indicates that for the entire system, the null hypothesis of normality can be rejected at the 0.05 level. Likewise, the univariate normality test statistic indicates that the null hypothesis of normality can be rejected at the 0.05 level of significance for each of the six equations. The magnitude and significance of the coefficients on skewness and excess kurtosis suggest that nonnormality of the residuals arises mainly because the innovations are leptokurtic. However, a study by Gonzalo (1994) finds that for cointegration analysis involving more than two variables, the Johansen procedure provides results that are more robust to various deviations from classical assumptions (e.g., nonnormality of errors due to skewness and excess kurtosis) than other methods of estimating cointegration relationships.\footnote{\textsuperscript{13}}

For the post-liberalization subsample, both multivariate and univariate tests of the null hypothesis of no serial correlation, first order and higher orders, suggest that the residuals are free of serial correlation. The univariate normality tests suggest that the null hypothesis of normality cannot be rejected at the 0.05 level of significance for three of the six equations, including that for the Philippines. However, the multivariate test indicates that for the entire system, the null hypothesis of normality can be rejected at the 0.05 level of significance. Since all skewness coefficients are not significantly different from zero for each equation, the rejection of normality of the residuals for three of the six equations can be attributed to excess kurtosis. Overall, the error term diagnostic test results suggest that the deviations from the Gaussian i.i.d. innovations assumption underlying the Johansen procedure are less serious in the post-liberalization subsample than in the pre-liberalization. Therefore, we expect the distortions in the test size due to nonnormal innovations to be less serious in the second subsample.
The results of performing the Johansen’s cointegration tests for the two subsamples are shown in Table 3. Based on both the trace and maximum eigenvalue test statistics, the null hypothesis of no cointegration cannot be rejected at the 0.05 level of significance for the pre-liberalization subsample. In contrast, for the post-liberalization subsample, the null hypothesis of no cointegration is rejected at the 0.05 level of significance while the null hypothesis of at most 1 cointegration relationship is not rejected. The results for the post-liberalization subperiod point to the existence of one cointegration vector. Considering that we have controlled for the small-sample bias in Johansen’s cointegration tests and that the test size bias due to deviations from normality is possibly less serious in the second subsample, then, these results clearly indicate that there is one cointegration relation among the Philippine stock market index and the other five international stock market indexes during the period October 1989 to December 1995.

The results of the cointegration tests imply that the cointegration relation among the Philippine stock market index and those of the other five international stock markets has changed over the two subsample periods. In particular, these results are consistent with the Philippine stock market being cointegrated with the stock markets of Taiwan, Japan, Hong Kong, Singapore, and the United States after, but not before, October 1989. A detailed analysis of the cointegration relationship should provide additional evidence for this assertion. In turn, the cointegration relationship implies an underlying equilibrium relationship among these national stock price indexes which dictates their common movement in the long run. However, following DeFusco et al. (1996), the existence of only one cointegration vector, or equivalently five common unit roots, among the system of stock market indexes in this sample implies that the long-run relationships is stable in only one direction. That is, the system can deviate in five independent directions suggesting that the degree of cointegration relationship is weak.

The weak relationship may be explained by the existence of direct and indirect impediments to the flow of international capital that remain in the Philippine market. These include, among others, restrictions on foreign ownership in some sectors specified in the country's Foreign Investment Act negative list; some degree of political instability which discourages foreign investors from participating in the economy; the relative smallness of the equity
market which can have negative implications on market liquidity; and, the few internationally cross-listed securities. Considering that cointegration and market integration are long-run concepts and that the effect of the liberalization takes time, we do not expect an immediate dramatic impact. This weak integration evidence is consistent with the findings of Buckberg (1995), Bekaert (1995) and Harvey (1995a, 1995b).

The finding of a cointegration vector among the stock market price series in this study allows us to estimate a vector error-correction model in the form of Equation (3) or Equation (4). Since this formulation includes both levels and first differences, it allows us to examine both the short-run and long-run dynamics of the individual national stock market indexes in the system. In particular, one can gain insight on which countries are important to obtaining this long-run equilibrium relationship and whether this relationship plays an important role in explaining the short-run movement of each country's stock market prices. These insights should provide a clearer picture on the strength of the long-run relationship underlying the cointegration vector. The details of the estimated error-correction model with the restriction of one cointegration vector imposed [Equation (4)] are summarized in part A of Table 4, including the short-run parameters, the normalized coefficients of the estimated cointegration vector, and the coefficients of speed of adjustment toward the long-run equilibrium relationship underlying the cointegration vector.15

First, examine the long-run dynamics of the estimated cointegration relationship. The equilibrium error relationship implied by the cointegration vector is given by

\[ z_{t+1} = 1.00S_{PH, t} + 0.485S_{TM, t} + 0.7675S_{JP, t} + 0.595S_{HK, t} - 1.2425S_{SE, t} + 0.293S_{US, t} \]  

As mentioned earlier, there is no straightforward interpretation of the coefficients of the cointegration vector. However, their relative magnitude can be used to determine which national stock market indexes are important in obtaining the long-run relationship underlying the cointegration vector, or equivalently, which national stock market indexes' long run movements are significantly influenced by the underlying long-run equilibrium relationship. This is confirmed by performing a likelihood ratio test to test the statistical significance of each national stock market's cointegration coefficient. The null hypothesis is that the coefficient of a particular
country's stock market index is zero in the cointegration vector. This is equivalent to testing the restriction that only the remaining five stock market indexes are important in obtaining the long-run equilibrium relationship. Under the null, the likelihood ratio test statistic is distributed as $\chi^2(1)$. The results of these tests are shown in Table 5.

From part A of Table 5, one may see that the null hypothesis of zero cointegration coefficient can be rejected at the 0.05 level of significance for all countries except Hong Kong and the U.S.. In the case of Hong Kong, the null hypothesis can only be rejected at the 0.10 level of significance. A test of the joint hypothesis that the cointegration coefficients are zero in the equations of Hong Kong and the U.S. yielded a likelihood ratio test statistic of 5.15. Under the null hypothesis, this test statistic is distributed as $\chi^2(2)$ with critical value of 5.99 at the 0.05 level of significance. Thus, the joint null hypothesis cannot be rejected. These findings suggest that the equilibrium relationship underlying the cointegration vector exerts significant influence on the long-run movements of the stock market prices of the Philippines, Taiwan, Japan, and Singapore. However, the same equilibrium relationship only plays a minor influence on the long-run movement of the stock market prices of Hong Kong and appears unimportant to the U.S. stock market index. That the long-run relationship is not important to the U.S. market suggests that the cointegration relationship is confined to the Asian stock markets. This may be due to the geographical separation of the Asian markets and that of the U.S. market and may be suggestive of the notion of separated financial markets in the Pacific Basin (Asian versus Pacific). This finding is similar to the multivariate cointegration results of Chung and Liu (1994). They find that the Johansen's cointegration tests indicate that cointegration relationships are shared by the stock markets of U.S., Japan, Hong Kong, Singapore and South Korea. However, their results suggest that the U.S. stock market does not appear to belong to the "common" stock region of the four Asian countries since the cointegration vector coefficient for the U.S. is not found to be significantly different from zero.

Next, we investigate the importance of the long-run relationship underlying the cointegration vector to the short-run movements of the national stock markets in the sample. Part B of Table 4 shows the results of estimating a VAR in first differences with 2 lags using data for the post-liberalization subsample period.
The coefficients on the lagged values of the first differences are generally insignificant. As discussed earlier, the VAR in first differences is similar to the vector error-correction model except for the lagged error-correction terms defined by $z_{t-1} = \beta' S_{t-1}$. The estimates of the VAR in first differences are presented, as well, in order to determine the incremental impact of including the error-correction terms. The parameters are estimated using OLS to make it comparable to the estimates of the vector error-correction model. However, as pointed out earlier, given the finding of a cointegration relationship among the six price series, the traditional VAR in first differences is a misspecified model (see Engle and Granger, 1987).

We find that only 3 out of the 36 coefficients on the lagged values of $\Delta S_{PH}$, $\Delta S_{TW}$, $\Delta S_{JP}$, $\Delta S_{HK}$, $\Delta S_{SI}$, and $\Delta S_{US}$ are significant at the 0.05 level. In particular, one may see that own lagged changes and lagged changes in the non-Philippine stock market indexes do not significantly influence the current period's changes in the Philippine stock market index. Likewise, lagged changes in the Philippine stock market index do not significantly affect changes in any of the other countries' stock market indexes. The generally insignificant cross-price effects suggest that the independent or local information contained in individual country's stock market prices does not appear to influence the stock market indexes of the other countries. Now, examine to what extent the information on the long-run equilibrium relationship, that is captured by the cointegration relationship, helps explain the short-run movements in the stock market prices of the six countries.

Observe from part A of Table 4 that, similar to the results of the VAR in first differences, the coefficient estimates on the lagged values of $\Delta S_{PH}$, $\Delta S_{TW}$, $\Delta S_{JP}$, $\Delta S_{HK}$, $\Delta S_{SI}$, and $\Delta S_{US}$ are generally insignificant with only 4 of the 32 coefficients significant at the 0.05 level. Note, however, that the results suggest that the lagged stock market index return of Singapore has a significant effect on the short-run changes in the Philippine stock market when the error-correction term is included in the Philippine equation but it is insignificant under the traditional VAR in first differences. Meanwhile, based on the $t$-values, the coefficient of adjustment speed in the equation of the Philippines is significant at the 0.05 level. This implies that short-run movements in the stock market prices are significantly influenced by information about the long-run relationship underlying the estimated cointegration vector or a
faster speed of adjustment in the Philippine stock market index toward the long-run equilibrium relationship. However, based on the t-values, the coefficients on the five remaining international stock markets are insignificant. Consequently, we perform a likelihood ratio test to test the null hypothesis that the adjustment speed coefficient is zero in each of the equations of the non-Philippine stock markets. If the adjustment speed coefficient is zero, this implies that short-run movements in the stock market prices of these countries are not influenced by the information about the long-run equilibrium relationship underlying the estimated cointegration vector. The results are reported in part B of Table 5. We observe that for each of these stock market indexes, the null hypothesis of zero coefficient of speed of adjustment cannot be rejected at the 0.05 level of significance. These results imply that the long-run relationship does not influence the short-run movement of these stock market indexes. A likelihood ratio test of the joint hypotheses that the adjustment speed coefficients are zero in all of the equations of these five international stock markets yielded a test statistic of 9.85. Under the null, this test statistic is distributed as $\chi^2(5)$ with critical value of 11.10 at the 0.05 level of significance. On the other hand, the critical value at the 0.10 level of significance is 9.24, suggesting that the joint hypotheses can only be weakly rejected.

Next, a comparison of the two models' $R^2$s reveals that adding the error-correction terms produces a substantial increase in the explained variation of $\Delta S_{PHL}$ from 0.127 to 0.300. This is true as well for $\Delta S_{TWN}$ and $\Delta S_{JPN}$ (from 0.099 to 0.138 and from 0.055 to 0.080, respectively). On the other hand, adding the error-correction terms results to modest increases in the explained variation of $\Delta S_{HK}$, $\Delta S_{SIN}$, and $\Delta S_{US}$ (from 0.096 to 0.105, from 0.038 to 0.040, and from 0.061 to 0.073, respectively).

The significant coefficient of adjustment speed and the substantial increase in the explained variation as a result of adding the lagged long-run equilibrium error provide evidence on the important influence of the information on the long-run equilibrium relationship on the short-run movements of the Philippine stock market index. Moreover, there appears to be evidence that in the short-run, this long-run equilibrium information is most important to the Philippine stock market. In contrast, the overall evidence indicates that this same relationship plays only a minor role in
explaining the short-run movements of the non-Philippine stock market indexes in the sample.

Given that cointegration implies market integration, the results of the cointegration tests and analysis of the significance of the long-run relationship underlying the cointegration vector lead us to believe that the Philippine stock market has become integrated with the stock markets of its top sources of foreign investment and major trading partners during the subsample period October 1989 to December 1995. However, since there is only one significant cointegration relationship which seems not to be vital to the Hong Kong and U.S. markets and because of the minor influence of the information on the long-run relationship on the short-run changes in the remaining international stock markets, the degree of integration of the Philippine equity market is taken to be weak. We also conducted a test of the joint hypotheses of zero adjustment coefficients for the equations of Taiwan, Japan, Hong Kong, Singapore and the U.S. and zero cointegration coefficients for Hong Kong and the U.S. This yielded a likelihood ratio test statistic of 23.38. Under the null, this test statistic has a $\chi^2(7)$ distribution with critical value of 14.10 at the 0.05 level. Therefore, the joint hypotheses can be significantly rejected. The last result indicates that imposing these restrictions on the system might be unreasonable.

Considering that more policy changes which eased access by foreigners into the domestic market as well as some reforms that made investing abroad by Filipinos easier were initiated beyond 1989 and that liberalization helps promote market integration, it may be reasonable to deduce from the evidence that market liberalization reforms are partly responsible for its integration into the international stock markets. Similarly, this result implies that the barriers which existed prior to the opening up of the Philippine market may have been effective in restricting the mobility of international capital flows which in turn prevented the market from being integrated. More importantly, these results indicate that the impact of liberalization takes time and therefore its effect on market integration is not immediately substantial.

It must be stressed, however, that the finding of no cointegration among the Philippine stock market index and those of Japan, Hong Kong, Singapore, Taiwan and the U.S. prior to October 1989 does not necessarily imply that these non-Philippine stock market indexes are not integrated into the international stock
markets in general. One argument that can be raised is that if the restrictiveness of barriers to international investment is what is preventing the Philippine stock market from being integrated with the other international stock markets, then it should be the case that the stock market indexes of countries which do not have strong restrictions and are freely open should be cointegrated prior to October 1989. However, since the purpose of this study is to examine the potential impact of the liberalization of the Philippine equity market on its major economic partners in the Asia-Pacific region, we did not pursue this issue at it is beyond our focus. Instead, existing literature involving the stock market of Japan, Hong Kong, Singapore, and the U.S. provide evidence toward this end. For example, Kasa (1992) finds a strong cointegration relationship among the equity markets in the U.S., Japan, England, Germany and Canada using data for the period 1974-1990. Corhay et al. (1995), find one cointegration relationship among the stock markets of Japan, Hong Kong, Singapore, Australia and New Zealand are cointegrated using data for the period 1972 to 1992. Hung and Cheung (1995), using data for the period 1981 to 1991, find some cointegration relationships among the equity markets of Hong Kong, Korea, Malaysia, Singapore, and Taiwan.

In summary, our results indicate that:

(1) The Philippine stock market is not cointegrated with the markets of its major trading partners and investing countries prior to liberalization. However, after its initial market opening, there is one significant cointegration vector which appears not to include the U.S. market. This suggest that capital market liberalization has resulted in the integration of the Philippine stock market with the Asian markets.

(2) For the post-liberalization subperiod, the return on the Philippine stock market index is significantly related to deviations from the long-run equilibrium relationship underlying the cointegration vector. This implies that the long-run equilibrium information underlying the cointegration vector significantly influences the short-run movements of the Philippine stock market index during this period.

(3) For the post-liberalization subperiod, adding the error-correction term in the traditional VAR in returns model
substantially increases the explanatory power for the returns on the Philippine stock market index. This implies that the deviations from the long-run equilibrium relationship between the Philippine stock market and the markets of its major economic partners provide incremental information on the movements of Philippine stock market returns. Analogously, useful incremental information is omitted when the long-run stock market price dynamics as reflected in the error-correction term is not incorporated in the model that is used to investigate the short-run dynamics of the Philippine stock market index.

VIII. CONCLUSIONS

This study investigated the impact of liberalization and internationalization on the degree of integration of the Philippine stock market into the stock markets of its major economic partners in the Asia-Pacific region. It employed a measure of the degree of integration based on the concept of cointegration and its relationship with error-correction, as introduced by Engle and Granger (1987). Based on Johansen's maximum likelihood cointegration tests, the null hypothesis of no cointegration cannot be rejected in the subperiod January 1980 to September 1989 while a single cointegration vector was found in the subperiod October 1989 to December 1995. Given that cointegration implies market integration, these results indicate that the Philippine stock market has become integrated with the stock markets of its major economic partners after its initial market opening on October 1989. This finding can partially be attributed to the major reforms aimed at liberalizing and internationalizing the Philippine capital market that were instituted after October 1989. This is because liberalization and internationalization encourages international capital flows which in turn promotes capital market integration. However, the fact that only one cointegration vector was found to be significant and that the influence of the underlying long-run equilibrium relationship on the non-Philippine stock market indexes is minor imply a weak degree of integration of the Philippine stock market into the stock markets of Taiwan, Japan, Hong Kong, Singapore and the United States. The weak integration of the Philippine stock market can possibly be explained by the fact that the market opening is not yet complete. There are barriers to international investment that continue to prevail in this country.
These include, among others, remaining restrictions on foreign equity participation in some sectors of economic activity, some degree of political instability that can discourage participation in the domestic market by overseas investors, the relatively small equity market base, and the dearth of internationally cross-listed domestic securities. The weak evidence of integration in the post-liberalization subperiod also provides an indication that the impact of liberalization takes time and therefore its effect on market integration is not immediately substantial.

Given that cointegration suggests long-run co-movements of national stock market prices, we expect long-run horizon return correlations among the cointegrated international stock market indexes to be higher than the correlations reflected in short-run stock market returns. In turn, this implies limited diversification benefits from investing in the stocks of the cointegrated markets. This is because the presence of common influences limits the amount of independent variation in the national stock market prices. However, since the above results indicate that the Philippine stock market is weakly integrated into the international stock markets in the sample, the potential long-run diversification benefits offered by Philippine equities to overseas investors is still significant.

Finally, the post-liberalization subperiod results indicate that the information contained in the long-run equilibrium relationship between the Philippine stock market and the markets of its major economic partners is helpful in explaining the short-run movements of the Philippine stock market index. This suggests that an appropriate model that seeks to explain the predictable component of Philippine stock market returns should take into account the existence of this cointegration relationship.
ENDNOTES

1. This paper is based on a chapter of the author's Ph.D. in Finance dissertation at the University of Alberta, Canada. Thanks are due to the Canadian International Development Agency and the Association of Deans of Southeast Asian Graduate Schools of Management for funding.

2. Among the major financial market liberalization reforms in the Philippines in recent years are as follows [International Monetary Fund (1986-1996)]. On 6/6/91 the Philippine Congress approved Republic Act No. 7042 or the Foreign Investments Act of 1991 which allows foreigners to own up to 100% of the common stock in a domestic Philippine company for many industries, where the previous limit was 40%. On 1/3/92 the Central Bank of the Philippines issued rules and regulations that liberalized foreign exchange transactions, effectively increasing the ease of capital outflows. On 8/10/92 President Ramos proposed measures to eliminate most remaining foreign exchange restrictions. On 6/22/94 President Ramos issued an executive order that further increased the number of industries in which foreigners can own up to 100% of outstanding common shares. On 5/18/94 Republic Act No. 7721 liberalized the entry of foreign banks to operate in the Philippines.


4. Taiwan's stock market capitalization stood at US$187.206 billion and market turnover was 174.9% as of 1995. In this same year, the capitalization of the developed and relatively open markets of Hong Kong and Singapore were US$303.705 and US$148.004 billion, respectively. Their corresponding market turnovers were 37.3% and 42.2%. (International Finance Corporation, 1996)

5. A competing hypothesis is that the series is stationary around a deterministic trend. If this is the case, the series is rendered stationary by detrending.

6. It can be shown that when the price indexes are translated into returns, the variance of the corresponding disturbance term,
(\(e_t - e_{t,\tau}\)), is greater than the variance of \(e_t\) whenever the autocorrelation of the original noise term, \(\rho_t\), is less than 0.5. The proof is as follows:

Let \(\text{Var}(e_t) = \gamma_0\) and \(\text{Cov}(e_t - e_{t,\tau}) = \gamma_1\).

Then \(\text{Var}(e_t - e_{t,\tau}) = 2(\gamma_0 - \gamma_1) = V\).

Now \(V > \gamma_0\) when \(2(\gamma_0 - \gamma_1) > \gamma_0\).

This holds when \(\rho_t < 0.5\) since \(2(\gamma_0 - \gamma_1) > \gamma_0 \Rightarrow \gamma_0 > 2\gamma_1 \Rightarrow 1 > 2\rho_t\) as \(\rho_t = (\gamma_1/\gamma_0)\).

7. Although it would have been more appropriate to use the Philippine Stock Exchange Composite Index, the available data begins only in 1987. However, it may be reasonable to use the Commercial and Industrial index since the component stocks in this subindex account for approximately 92% of the total market capitalization of the stocks in the Composite Index (see Rodrigo, 1993). Moreover, the commercial-industrial sector accounted for 75% of market turnover in 1995 (see Euromoney, 1996). All of the indexes in the sample do not include dividends. It would be ideal to use indexes which include dividends since investors are concerned about total returns. However, due to limitations on dividend data, all analyses are performed on indexes whose component stock prices are not adjusted for dividends. DeFusco et al. (1996) theoretically show that, assuming the discounted cash flow model of asset prices is valid and if expected discount rates and dividend growth rates are stationary processes, the results of the cointegration tests will be the same whether one uses price indexes excluding dividends or price indexes adjusted for dividends.

8. We estimate a model with a constant term in order to capture the upward trending behavior of the national stock market price indexes in the sample. The cointegration estimation and testing procedures are performed using CATS-PC Version 1.00 and RATS Version 4.2.

9. The Monte Carlo simulation results of Cheung and Lai (1993) indicate that serial correlation introduces a serious problem for the Johansen procedure. On the other hand, Cushman, Lee, and Thorgeirsson (1996) point out that the normality assumption is not as serious as it appears. They argue that for nonnormal \(e_t\),
the Johansen estimators become quasi-ML estimators. Moreover, the asymptotic results for the cointegration test statistics do not change, as long as the distribution for $e_i$ satisfies the conditions necessary to invoke the functional central limit theorem (e.g., i.i.d. with finite covariance matrix). Gonzalo (1994) finds that the Johansen procedure provides results that are more robust to deviations from normality than those obtained from other methods of estimating cointegration relationships.


11. These results have an implication on the specification of VAR models whose variables are nonstationary in the levels. If the nonstationary time series in the vector system of the VAR are cointegrated, a VAR model in levels is inefficient and may lead to spurious regression results, while that in first differences is misspecified. (see Engle and Granger, 1987)

12. It should be noted that, starting from the late 1970s until the mid 1980s, the Japanese government instituted a wave of major reforms that liberalized the Japanese capital market (see, e.g., Korkie and Nakamura, 1997).

13. On the other hand, Cheung and Lai (1993), using a simple Monte Carlo experiment, examine the potential effects of nonnormal innovations, including nonsymmetric and leptokurtic ones, on the size of Johansen’s cointegration tests. Among other things, they find that skewness and excess kurtosis produce a statistically significant effect on the test sizes of both the trace and maximum eigenvalue tests. However, they find that while both tests are reasonably robust to nonnormal disturbances, the trace test shows more robustness to both skewness and excess kurtosis in disturbances than the maximum eigenvalue test.

14. The analysis was also performed with a dummy variable for the international stock market crash of October 1987. The results are similar to the ones reported in this study in which a dummy variable is not included. A problem with including this dummy
variable, however, is that it changes the limit distribution of the cointegration test statistics and therefore it is not appropriate to compare them with the critical values reported in Osterwald-Lenum (1992) which were simulated without such dummy variable.

15. Since their values are not uniquely defined, Johansen (1988) suggested normalizing the coefficients of the cointegration vectors before any inferences about them can be made. An advantage of the Johansen procedure is that the results and implications are invariant to the chosen normalizing variable. In this study, normalization is performed using the coefficient on the Philippine stock market index.

BIBLIOGRAPHY


Campbell, J. and Y. Hamao. “Predictable stock returns in the

Cheung, Y., J., He, J. and L. Ng. "What are the Global Sources of Rational Variation in International Equity Returns?" Working paper, City University of Hong Kong and University of California, Santa Cruz.


de los Angeles, E. "Philippine Stock Exchange: After the Unification." Speech delivered at the 7th annual conference of the Pacific Basin capital markets held on July 6-8, 1995, Philippines.

De Santis, G. and S. Imrohoroglu "Stock Returns and Volatility in Emerging Financial Markets." Working paper, Department of Finance and Business Economics, School of Business Administration, University of Southern California.


Financial Times (London), various issues.


Johansen, S. "Estimation and hypothesis testing of


Ripley, D. Systematic elements in the linkage of national stock


Wall Street Journal, various issues.

Table 1. Results of Unit Root Tests on the Levels of the National Stock Market Indexes for Two Subsamples*.

<table>
<thead>
<tr>
<th></th>
<th>Pre-Liberation Sub-Sample: January 1980 to September 1989</th>
<th>Post-Liberation Sub-Sample: October 1989 to December 1995</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>ADF Test</td>
<td>PP Test</td>
</tr>
<tr>
<td></td>
<td>constant</td>
<td>constant</td>
</tr>
<tr>
<td></td>
<td>no trend</td>
<td>no trend</td>
</tr>
<tr>
<td>Philippines</td>
<td>(Φ)</td>
<td>(Φ)</td>
</tr>
<tr>
<td></td>
<td>(τ)</td>
<td>(τ)</td>
</tr>
<tr>
<td></td>
<td>(Φ)</td>
<td>(Φ)</td>
</tr>
<tr>
<td></td>
<td>(Φ)</td>
<td>(Φ)</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>3.39</td>
<td>0.47</td>
</tr>
<tr>
<td></td>
<td>4.56</td>
<td>-1.00</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Taiwan</td>
<td>3.82</td>
<td>1.88</td>
</tr>
<tr>
<td></td>
<td>4.37</td>
<td>-2.87</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td>1.77</td>
<td>0.33</td>
</tr>
<tr>
<td></td>
<td>5.25</td>
<td>-2.74</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Hong Kong</td>
<td>1.48</td>
<td>-1.31</td>
</tr>
<tr>
<td></td>
<td>2.26</td>
<td>-1.07</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Singapore</td>
<td>3.16</td>
<td>-2.43</td>
</tr>
<tr>
<td></td>
<td>4.33</td>
<td>-0.77</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>U. S.</td>
<td>3.15</td>
<td>-0.42</td>
</tr>
<tr>
<td></td>
<td>4.60</td>
<td>0.09</td>
</tr>
</tbody>
</table>

*The reported ADF test statistics, Φ, are F-statistics on the joint significance of β₁ and β₂ in the regression ΔSₜ = β₀ + β₁Sₜ₋₁ + β₂S₂Sₜ₋₁ + Σᵢ=₁ β₃ΔSₜ₋ᵢ, while the ADF test statistics, τ, are t-statistics on β₁ in the regression ΔSₜ = β₀ + β₁Sₜ₋₁ + Σᵢ=₁ β₃ΔSₜ₋ᵢ. The critical values at the 0.05 level of significance are 6.47 and -2.89, respectively for the sample period January 1980 to September 1989 and 6.61 and -2.91, respectively for the sample period October 1989 to December 1995. The reported PP test statistics, Z(Φ), are F-statistics on the joint significance of β₁ and β₂ in the regression ΔSₜ = β₀ + β₁Sₜ₋₁ + β₂S₂Sₜ₋₁ corrected for serial correlation and/or heteroscedasticity using the Newey West (1987) method while the PP test statistics, Z(τ), are t-statistics on β₁ in the regression ΔSₜ = β₀ + β₁Sₜ₋₁ corrected for serial correlation and/or heteroscedasticity using the Newey West (1987) method. The critical values at the 0.05 level of significance are the same as those of the ADF test statistics.

** denotes rejection at the 0.05 level of significance of the null hypothesis of a unit root.
Table 2. Error Term Diagnostics\(^{\text{a,b}}\)

<table>
<thead>
<tr>
<th>Equation for</th>
<th>Pre-Liberization Subsample: January 1980 to September 1989</th>
<th>Post-Liberization Subsample: October 1989 to December 1995(^{\text{a}})</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td><strong>LM(1)</strong></td>
<td><strong>LM(4)</strong></td>
</tr>
<tr>
<td>Philippines</td>
<td>0.10</td>
<td>0.16</td>
</tr>
<tr>
<td>Taiwan</td>
<td>0.09</td>
<td>0.90</td>
</tr>
<tr>
<td>Japan</td>
<td>0.10</td>
<td>2.92</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>0.00</td>
<td>2.72</td>
</tr>
<tr>
<td>Singapore</td>
<td>0.13</td>
<td>1.26</td>
</tr>
<tr>
<td>U.S.</td>
<td>0.18</td>
<td>0.17</td>
</tr>
</tbody>
</table>

a The residuals are based on the estimates of the unrestricted VAR in error-correction form [Equation (2)] with lags \(j=3\) and an unrestricted constant. The lag length was determined using a sequential likelihood ratio procedure. The null hypothesis that system of variables \(x\), is generated by a Gaussian VAR with \(j\) lags against the alternative of \(j+1\) lags is given by

\[
L = \frac{1}{n} \left(1 + \frac{p}{n(j+1)} \right) \left(\ln \frac{\hat{\Sigma}_{x}}{\hat{\Sigma}_{x,0}}\right)
\]

where \(T\) is the VAR sample size, \(p=1+[(j+1)j]/2\) is the correction for small sample bias. \(\hat{\Sigma}_{x,0}\) is the ML estimate of the variance-covariance matrix of the VAR with \(s\) lags, \(s=j,j+1\). This test statistic has a \(\chi^2\) distribution.

b The residuals are based on the estimates of the unrestricted VAR in error-correction form [Equation (2)] with lags \(j=2\) and an unrestricted constant with the lag length determined using the sequential likelihood ratio procedure discussed in note a.

c \(LM(1)\) and \(LM(4)\) are multivariate Lagrange Multiplier tests for first- and fourth-order autocorrelation as proposed by Godfrey (1988). The null hypothesis is that the \(k\)-th order autocorrelation is zero in all the equations in the VAR system. Both test statistics are distributed as \(\chi^2(36)\) with critical value of 51 at the 0.05 level of significance.

d D-H Normality is the multivariate normality test suggested by Hansen and Juselius (1995). The null hypothesis is that the estimated residuals in all of the equations in the VAR system are randomly drawn from a multivariate normal distribution. This test statistic is distributed as \(\chi^2(12)\) with critical value of 21 at the 0.05 level of significance.

e LB(1) and LB(4) are the univariate test statistics for testing first-order and fourth order autocorrelations, respectively, as proposed by Ljung and Box (1978). The null hypothesis is that the first \(k\) autocorrelations are zero in a given equation in the VAR system. These statistics are distributed as \(\chi^2(1)\) and \(\chi^2(4)\), respectively. The critical values at the 0.05 level of significance are 3.84 and 9.49, respectively. Tests for higher-order autocorrelation (beyond four) were conducted but none are significant. Only the first and fourth are reported for parsimony.

f Skewness is the coefficient of skewness and Kurtosis is the coefficient of excess kurtosis as proposed by Kendall and Stuart (1958). The null hypothesis tested here is that each coefficient is zero. Both statistics have a standard normal distribution.

g J-B Normality is the univariate normality test suggested by Jarque and Bera (1980). The null hypothesis is that the estimated residuals in a given equation in the VAR system are randomly drawn from a normal distribution. This test statistic is distributed as \(\chi^2(2)\) with critical value of 5.99 at the 0.05 level of significance.

h ** denotes rejection at the 0.05 level of significance of the null hypothesis.
Table 3. Johansen's Likelihood Ratio Tests Statistics for Cointegration Among the National Stock Market Indexes for Two Subsamples

<table>
<thead>
<tr>
<th>Null Hypothesis</th>
<th>Test Statistics</th>
<th>Small-Sample Critical Values (0.05 level of significance)</th>
<th>Test Statistics</th>
<th>Small-Sample Critical Values (0.05 level of significance)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Trace (Maximum Eigenvalue)</td>
<td>Trace (Maximum Eigenvalue)</td>
<td>Trace (Maximum Eigenvalue)</td>
<td>Trace (Maximum Eigenvalue)</td>
</tr>
<tr>
<td>n=0</td>
<td>107.55 (41.03)</td>
<td>111.80 (46.75)</td>
<td>133.77** (88.42**)</td>
<td>112.67 (47.11)</td>
</tr>
<tr>
<td>n=1</td>
<td>65.52 (29.99)</td>
<td>81.37 (38.73)</td>
<td>88.05 (43.71)</td>
<td>82.00 (40.04)</td>
</tr>
<tr>
<td>n=2</td>
<td>39.64 (21.37)</td>
<td>56.06 (32.15)</td>
<td>54.65 (20.37)</td>
<td>59.00 (32.40)</td>
</tr>
<tr>
<td>n=3</td>
<td>15.17 (8.81)</td>
<td>35.25 (24.90)</td>
<td>14.28 (7.29)</td>
<td>35.92 (25.10)</td>
</tr>
<tr>
<td>n=4</td>
<td>8.36 (5.61)</td>
<td>16.30 (16.71)</td>
<td>8.99 (6.81)</td>
<td>18.44 (16.84)</td>
</tr>
<tr>
<td>n=5</td>
<td>0.54 (0.54)</td>
<td>4.47 (4.47)</td>
<td>0.18 (0.18)</td>
<td>4.50 (4.50)</td>
</tr>
</tbody>
</table>

* The test statistics are based on the estimates of the unrestricted VAR in error-correction form [Equation (2)] with lags \( j = 3 \) and an unrestricted constant. Eigenvalues are \( (0.2735, 0.2235, 0.1935, 0.0867, 0.0644, 0.0118) \).

* The test statistics are based on the estimates of the unrestricted VAR in error-correction form [Equation (2)] with lags \( j = 2 \) and an unrestricted constant. Eigenvalues are \( (0.5919, 0.3698, 0.2435, 0.0950, 0.0891, 0.0024) \).

* The small-sample critical values are the asymptotic critical values from Osterwald-Lenum (1992, Table 1 Case 1) for a system of six variables and a constant in the VAR multiplied by \( T/(T-n) \), based on the suggestion by Reinsel and Ahn (1988, 1992). The correction factor, \( T/(T-n) \), is 1.1875 for the subsample period January 1980 to September 1989 and 1.1967 for the subsample period October 1989 to December 1995.

* ** denotes rejection at the 0.05 level of significance of the null hypothesis of at most \( r \) cointegration vectors against the alternative of six cointegration vectors.

* ** denotes rejection at the 0.05 level of significance of the null hypothesis of \( r \) cointegration vectors against the alternative of \( r+1 \) cointegration vectors.
### Table 4. Estimation Results: October to December 1995\(^{a-c}\)

<table>
<thead>
<tr>
<th>A. Vector Error Correction Model</th>
<th>B. VAR in First Differences</th>
</tr>
</thead>
<tbody>
<tr>
<td>(\Delta S_{\text{Phil}})</td>
<td>(\Delta S_{\text{Phil}})</td>
</tr>
<tr>
<td>(0.064)</td>
<td>(-0.076)</td>
</tr>
<tr>
<td>(0.58)</td>
<td>(-0.75)</td>
</tr>
<tr>
<td>(0.30)</td>
<td>(0.18)</td>
</tr>
<tr>
<td>(1.82)</td>
<td>(1.44)</td>
</tr>
<tr>
<td>(0.02)</td>
<td>(0.04)</td>
</tr>
<tr>
<td>(0.20)</td>
<td>(0.55)</td>
</tr>
<tr>
<td>(-0.04)</td>
<td>(-0.02)</td>
</tr>
<tr>
<td>(-0.47)</td>
<td>(-0.31)</td>
</tr>
<tr>
<td>(-0.03)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>(-0.41)</td>
<td>(0.17)</td>
</tr>
<tr>
<td>(0.01)</td>
<td>(0.04)</td>
</tr>
<tr>
<td>(0.24)</td>
<td>(1.18)</td>
</tr>
<tr>
<td>(0.03)</td>
<td>(0.19)</td>
</tr>
<tr>
<td>(0.43)</td>
<td>(1.92)</td>
</tr>
</tbody>
</table>

**Long-run Parameters:**

- Vector of adjustment speeds, \(\alpha\):
  - \(\alpha_1\) = \(-0.368\)
  - \(\alpha_2\) = \(0.202\)
  - \(\alpha_3\) = \(-0.100\)
  - \(\alpha_4\) = \(-0.053\)
  - \(\alpha_5\) = \(0.020\)
  - \(\alpha_6\) = \(-0.026\)
  - \(\alpha_7\) = \(4.28\) ***
  - \(\alpha_8\) = \(1.83\) ***

- Cointegration vector (normalized), \(\beta\):
  - \(\beta_1\) = \(0.001\)
  - \(\beta_2\) = \(0.006\)
  - \(\beta_3\) = \(0.16\)
  - \(\beta_4\) = \(0.26\)

- **R²** = 0.500

---

* Estimate \(t\)-values with the rank restriction \(r=1\) imposed on the vector error-correction model (Equation (2)). The estimates for the subsample January 1989 to September 1989 are not reported since no significant cointegration vector was found.

* \(S_{\text{Phil}}, S_{\text{TMY}}, S_{\text{JAP}}, S_{\text{REST}}, S_{\text{JM}}, \) and \(S_{\text{U.S.}}\) are the levels of the (natural logarithm) national stock market price indexes measured in U.S. dollars for the Philippines, Taiwan, Japan, Hong Kong, Singapore, and the United States, respectively.

* \(D\) is the first-difference operator.

* The coefficients of the cointegration vector reported above are simply the coefficients of the corresponding eigenvector normalized by the eigenvector coefficient of the Philippine stock market index. In the Johansen framework, the short-run parameter estimates and test statistics as well as the cointegration test statistics are invariant to the choice of the normalizing variable. The eigenvector is \((7.792, -3.780, 5.975, -4.634, -9.861, 2.282)\).

** ** denotes significance the 0.05 level.
Table 5. Tests of Restrictions on the Cointegratin Vector and the Vector of Adjustment Speed coefficients Post-Liberalization Subsample: October 1989 to December 1985$^{a,b}$

(A) Test of Restrictions on the Cointegration Vector, $\beta$

$H_0$: The cointegration coefficient is zero in the error correction equation of the individual national stock market index.

<table>
<thead>
<tr>
<th>Equation for</th>
<th>Likelihood Ratio Test Statistic</th>
</tr>
</thead>
<tbody>
<tr>
<td>Philippines</td>
<td>19.87***</td>
</tr>
<tr>
<td>Taiwan</td>
<td>3.45**</td>
</tr>
<tr>
<td>Japan</td>
<td>11.01***</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>3.85**</td>
</tr>
<tr>
<td>Singapore</td>
<td>7.48**</td>
</tr>
<tr>
<td>U.S.</td>
<td>1.31</td>
</tr>
</tbody>
</table>

(B) Tests of Restrictions on the Vector of Adjustment Speed Coefficients, $\alpha$

$H_0$: The adjustment speed coefficient is zero in the error correction equation of an individual national stock market index.

<table>
<thead>
<tr>
<th>Equation for</th>
<th>Likelihood Ratio Test Statistic</th>
</tr>
</thead>
<tbody>
<tr>
<td>Taiwan</td>
<td>2.54</td>
</tr>
<tr>
<td>Japan</td>
<td>1.50</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>0.98</td>
</tr>
<tr>
<td>Singapore</td>
<td>0.13</td>
</tr>
<tr>
<td>U.S.</td>
<td>0.87</td>
</tr>
</tbody>
</table>

* The likelihood ratio test statistics are based on Johansen (1988, 1991) and Johansen and Juselius (1990). Under the null hypothesis, both statistics have a $\chi^2 (1)$ distribution with critical values of 3.84 and 2.71 at the 0.05 and 0.10 levels of significance, respectively.

$^{a}$ denotes that the null hypothesis can be rejected at the 0.05 level of significance.

$^{b}$ denotes that the null hypothesis can be rejected at the 0.10 level of significance.
### Appendix A: Data Sources

<table>
<thead>
<tr>
<th>Country/Stock Market Index</th>
<th>Stock Market Price Index</th>
<th>Exchange Rate</th>
</tr>
</thead>
</table>