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Evidence of financial integration in Asia: An empirical application of panel unit root tests and multivariate cointegration and causality procedures

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Abstract. This paper measures the extent of financial integration and interdependence among Asian equity markets over the period January 1993 to June 2006 using daily data. The analysis includes three developed markets (Hong Kong, Japan and Singapore) and eight emerging markets (China, India, Indonesia, Korea, Malaysia, the Philippines, Taiwan and Thailand). The study uses panel unit root tests to test for non-stationarity, and conducts multivariate cointegration, Granger causality and level VAR procedures and variance decomposition are conducted to examine the equilibrium and causal relationships between these markets. The results indicate that there is a stationary long-run equilibrium relationship among, and significant and substantial short and long-run causal linkages between, these Asian equity markets. This evidence suggests that a high level of financial integration currently exists in the Asian region, notwithstanding the absence of extensive formal regional agreements aimed at promoting financial integration as found elsewhere, especially in the European Union.

Keywords: financial integration, portfolio diversification, international capital allocation, economic development and growth, market efficiency.

1. Introduction

Financial integration is the process by which a country's financial markets - including money, bond, bank credit and equity markets - become more closely integrated with those in other countries. Three widely-accepted and interrelated benefits accrue from this process: more opportunities for risk sharing and diversification, the better allocation of capital across investment opportunities, and the potential for higher economic growth. First, sharing risk across regions enhances specialisation, increases the set of financial instruments available, and thereby provides additional possibilities for portfolio diversification. Second, the elimination of barriers to trading, clearing and settlement allows firms to choose the most efficient location for their financing activities. Investors too are free to invest their funds where they will be allocated to their most productive end-use. The improvement in capital allocation also enhances financial development, thereby assisting the process of economic growth with additional funds flowing to (often less-developed) countries that have more and better productive opportunities.

Financial integration arises in two main ways. One is from formal efforts to integrate financial markets with particular partners, typically those that share membership in some wider regional agreement. Integration in this sense involves the elimination of cross-border restrictions on the activities of firms and investors within the region, as well as the harmonisation of rules, taxes and regulations between member countries. The European Union is an obvious example. It is generally expected that financial integration should follow from these developments. However, financial integration may also emerge less formally, very often but not always as a precursor to explicit regional agreements. Several factors contribute to this means of financial integration. These include foreign bank entry into domestic markets, direct borrowing by firms in international markets, bilateral financial and trade agreements, strengthening finance and trade relationships between countries, and the convergence of business and investor practices. Financial integration such as this is relatively more common in the developing world, especially in geographically close regions.

Financial markets in Asia are a pertinent context in which to consider financial integration, especially in terms of equity markets. Within-Asia cross-border investment and financing opportunities benefited from a succession of booming economies, most recently in China and India. Taking a lesson from the 1997 financial crisis, many Asian countries have now restructured and reformed their economic and financial systems to attract equity capital to assist financial development and economic growth. Simultaneously, the formation of monetary and trade unions in Europe and elsewhere encouraged similar developments and their benefits are now considered in Asia. Key factors in these developments include the increasing shares of intra-regional trade and investment, the large number of emerging markets in the region and their strong growth potential.

To inform policy and provide guidance for investing and financing in Asia, empirical work is needed which reflects, recognizes and appropriately measures the complex market interrelationships that exist in this globally important region. The key requirements are that the measures should assess the current level of financial integration, and indicate whether integration is progressing, stable or regressing. One possible approach is the use of cointegration, causality and variance decomposition methods to assess the equilibrium and causal relationships between financial markets. Unfortunately, despite more than a decade of work, relatively little empirical evidence still exists concerning financial integration among Asian equity markets, notwithstanding valuable contributions by Bailey and Stulz (1990), Cheung and Mak (1992), Lai *et al.* (1993), Chung and Liu (1994), Kwan *et al*. (1995), Solnik *et al*. (1996), Darbar and Deb (1997), Yuhn (1997), Janakiramanan and Lamba (1998), Ramchand and Susmel (1998), Masih and Masih (1999), Roca (1999), Worthington *et al*. (2003) and Worthington and Higgs (2004a; 2004b), amongst others.

This study employs a quantitative method for measuring financial integration in Asian equity markets. The chapter itself is divided into four main areas. The second section presents the data employed in the analysis. The third section explains the methodology and the fourth section deals with the results. The chapter ends with some brief concluding remarks.

2.2 Data

The data employed in the study is composed of value-weighted equity market indices for eleven Asian markets namely, China (CHN), Hong Kong (HKG), India (IND), Indonesia (INA), Japan (JPN), Korea (KOR), Malaysia (MLY), the Philippines (PHL), Singapore (SNG), Taiwan (TWN) and Thailand (THA). Three of these markets are categorised as 'developed' (Hong Kong, Japan, and Singapore) while the remainder are 'emerging' or 'developing'. All index data specified is obtained from Morgan Stanley Capital International (MSCI) (2007) in US dollar terms and encompasses the period 1 January 1993 to 31 June 2006.

> In constructing a country index every listed security in the market is identified. Securities are free float adjusted, classified in accordance with the Global Industry Classification Standard (GICS®), and screened by size and liquidity. MSCI then constructs its indices by targeting for index inclusion 85% of the free float adjusted market capitalization in each industry group, within each country. By targeting 85% of each industry group, the MSCI Country Index captures 85% of the total country market capitalization while it accurately reflects the economic diversity of the market.

MSCI indices are widely employed in the financial integration literature because of the degree of comparability, the avoidance of dual listing and the breadth and reflectivity of index coverage [see, for instance, Meric and Meric (1997), Yuhn (1997), Cheung and Lai (1999), Roca (1999), Worthington *et al*. (2003), Worthington and Higgs (2004a; 2004b)]. The daily data employed comprise the longest continuous common-time series available for the eleven Asian equity markets, encompassing 3,522 observations for each market. The eleven markets together constitute a total of 38,742 observations.

3. Empirical methodology

This paper investigates the integration and interdependence among Asian equity markets as follows. Panel unit root tests are first conducted as a means of informing subsequent methods.

Multivariate cointegration, Granger causality, level VAR and variance decomposition methods are then employed to examine the integration and interrelationships among markets.

3.1 Panel unit root tests

Panel unit root tests comprise a multivariate analogue to standard univariate unit root tests, including the Augmented Dickey-Fuller (ADF), Phillips-Perron (PP) and Kwiatkowski, Phillips, Schmidt and Shin (KPSS) tests. The main purpose in extending the application of purely time-series unit root tests to panel unit root tests is to use the increase in sample size from pooling cross-sectional data to improve the power of the tests. Three panel unit root tests are examined namely: the Levine, Lin and Chu (2002), Im, Pesaran and Shin (2003) and Hadri (2000) tests.

(i) A basic model

Assume the time series $\{y_{i,0}, \ldots, y_{i,T}\}$ on the cross-section units (or markets) $i = 1, 2, \ldots, M$ over T time periods are generated for each i by a simple first-order autoregressive, AR(1), process:

$$y_{i,t} = (1-\rho_i)\mu_i + \rho_i y_{i,t-1} + \varepsilon_{i,t} \qquad i = 1, 2, ..., M, \quad t = 1, 2, ..., T \tag{1}$$

where $y_{i,t}$ denotes the observed cross-section for the i-th unit at time t and $\varepsilon_{i,t}$ is white noise for the i-th unit at time t. The errors $\varepsilon_{i,t}$ are identically and independently distributed (*i.i.d*) across i and t with $E(\varepsilon_{i,t}) = 0$, $E(\varepsilon_{i,t}^2) = \sigma_i^2 < \infty$ and $E(\varepsilon_{i,t}^4) < \infty$. Under the null hypothesis of a unit root, $\rho_i = 1$ for all i, equation (1) can be rewritten as the following basic ADF specification:

$$\Delta y_{i,t} = \alpha_i + \phi_i y_{i,t-1} + \sum_{j=1}^{q_i} \gamma_{i,j} \Delta y_{i,t-j} + \varepsilon_{i,t} \tag{2}$$

where $\alpha = (1 - \rho_i)\mu_i$, $\phi_i = (\rho_i - 1)$ and γ_i are coefficients to be estimated for the i-th unit, q_i is the number of lagged terms for the i-th unit $\Delta y_{i,t} = y_{i,t} - y_{i,t-1}$ and all other parameters are as previously defined.

(ii) Levine, Lin and Chu test

One of the first panel unit root tests was proposed by Levine and Lin (1992) and subsequently formalised in Levine et al. (2002) (hereafter LLC). The LLC test permits the intercept, time trend, residual variance and higher-order autocorrelations to vary across individual markets.

The LLC test is based on a pooled panel estimator which assumes a common $\phi_i = \phi$ but allows q_i to vary across the cross-sections. It also requires the independently generated time series to have a common sample size. The LLC test may then be viewed as a pooled ADF test potentially with different lag lengths across the cross-sections of the panel. The main limitation of this test is that it imposes a cross-equation restriction on the first-order autocorrelation coefficients. Under the LLC, the null and alternative hypotheses are given as:

$H_{0,LLC}$: $\phi_1 = \phi_2 = \ldots = \phi_M = 0$

$H_{1,LLC}$: $\phi_1 = \phi_2 = \ldots = \phi_M = \phi < 0$

Under the null hypothesis, each cross-section has a unit root (or is non-stationary) while under the alternative each cross-section unit is stationary. The LLC test statistic under the null hypothesis is a modified *t*-statistic.

(iii) Im, Pesaran and Shin test

The Im, Pesaran and Shin (2003) test (herafter IPS) is introduced to take account of the major weakness of the LLC test where it is assumed that all individual AR(1) series have a common autocorrelation coefficient. It allows for individual processes by permitting ϕ_i to vary across the cross-sections. The IPS test begins by specifying a separate ADF regression for each cross-section unit specified by equation (2). The null and alternative hypotheses for the IPS test are:

$H_{0,IPS}$: $\phi_i = \phi = 0$ for $\forall\ i$

$H_{1,IPS}$: $\phi_i < 0$ for $i = 1, 2, \ldots, M_1$ and $\phi_i = 0$ for $i = M_1 + 1, \ldots, M$

Under the null hypothesis, all cross-section units in the panel are non-stationary. The IPS test assumes that under the alternative at least one cross-section unit, but not all cross-section units is stationary. This differs from the LLC test which presumes all cross-section units are stationary under the alternative hypothesis.

The IPS test is based on M independent tests on M cross-section units while the LLC test combines the test statistics. The random errors, $\varepsilon_{i,t}$, are assumed to be serially correlated with different serial correlation properties and different variances across each cross-section unit. The core of the IPS test is based on a group-mean *t*-bar statistic where the *t*-statistics are drawn from each ADF test and averaged across the panels. Adjustment factors are used to standardise the *t*-bar statistic into a standard normal IPS *W*-statistic under the null hypothesis.

(iv) Hadri test

The Hadri (2000) panel unit root test parallels the well-known KPSS unit root test with the null hypothesis of no unit root in any of the cross-section units in the panel. As with the KPSS unit root test, the Hadri test is based on the residuals from individual OLS regressions of $y_{i,t}$ on a constant or a constant and a trend. The test statistics is distributed as standard normal under the null. The error process may be assumed to be homoskedastic across the panel or heteroskedastic across the cross-section units. Two *Z*-statistics are presented. One Z-statistic is derived from the Lagrange Multiplier (LM) statistic where the residuals from the ADF regression are associated with the homoskedasticity assumption across the panel and the other using the LM statistic that is heteroskedasticity consistent.

3.2 Multivariate cointegration

Following Engle and Granger (1987), suppose the set of M market index series $y_t = [y_{1t,}\ y_{2t}, \cdots, y_{Mt}]'$ are all I(1) and $\beta' y_t = u_t$ is I(0), then β is said to be a cointegrated vector and $\beta' y_t = u_t$ is called the cointegrating regression. The components of y_t are said to be cointegrated of order *d*, denoted by $y_t \sim CI(d, b)$ where $d > b > 0$, if (i) each component of y_t is integrated of order *d*, and (ii) there exists at least one vector $\beta = (\beta_1, \beta_2, \ldots, \beta_M)$, such that the linear combination is integrated of $(d - b)$. By Granger's theorem, if the indices are cointegrated, they can be expressed in an Error Correction Model (ECM) encompassing the notion of a long-run equilibrium relationship and the introduction of past disequilibrium as explanatory variables in the dynamic behaviour of current variables.

In order to implement the ECM, the order of cointegration must be known. A useful statistical test for determining the cointegration order proposed by Johansen (1991) and Johansen and Juselius (1990) is the trace test. For example, to test for no cointegrating relationship, *r* is set to zero and the null hypothesis is $H_0 : r = 0$ and the alternative is $H_1 : r > 0$. However, the Johansen (1991) test can be affected by the lag order. The lag order is determined by using both the likelihood ratio test and information criteria in VAR. The optimum number of lags to be used in the VAR models is determined by the likelihood ratio test statistic:

$$LR = (T - K)\ln(|\Sigma_0| / |\Sigma_A|) \qquad (3)$$

where *T* is the number of observations, *K* denotes the number of restrictions, Σ denotes the determinant of the covariance matrix of the error term, and subscripts *0* and *A* denote the restricted and unrestricted VAR, respectively. LR is asymptotically distributed χ^2 with degrees of freedom equal to the number of restrictions. The test statistic in (3) is used to test the null hypothesis of the number of lags being equal to k-1 against the alternative hypotheses that k = 2, 3, … and so on. The test procedure continues until the null hypothesis fails to be rejected, thereby indicating the optimal lag corresponds to the lag of the null hypothesis.

3.3 Multivariate Granger causality and level VAR tests

To examine the short-run relationships among the markets, Granger (1969) causality tests are specified. Essentially tests of the prediction ability of time series models, a market index causes another index in the Granger sense if past values of the first index explain the second, but past values of the second index do not explain the first. When the indices in question are cointegrated, Granger causality is tested using the ECM:

$$\Delta y_t = \gamma_0 + \sum_{i=1}^{r} \psi_i \Theta_{t-1} + \sum_{i=1}^{m} \gamma_i \Delta y_{t-i} + \varepsilon_t \tag{4}$$

where Θ contains r individual error-correction terms, r are long-term cointegrating vectors via the Johansen procedure, ψ and γ are parameters to be estimated, and all other variables are as previously defined.

One problem with a Granger causality test based on (4) is that it is affected by the specification of the model. ECM is estimated under the assumption of a certain number of lags and cointegrating equations, which means that the actual specification depends on the pre-test unit root and cointegration (Johansen) tests. To avoid possible pre-test bias, Toda and Yamamoto (1995) propose the level VAR procedure. Essentially, the level VAR procedure is based on VAR for the level of variables with the lag order p in the VAR equations given by $p=k+d_{max}$, where k is the true lag length and d_{max} is the possible maximum integration order of variables. Therefore, the estimated VAR is expressed as:

$$y_t = \hat{\gamma}_0 + \hat{\gamma}_1 t + \cdots + \hat{\gamma}_q t^q + \hat{J}_1 y_{t-1} + \cdots + \hat{J}_k y_{t-k} + \cdots + \hat{J}_p y_{t-p} + \hat{\varepsilon}_t, \tag{5}$$

where $t = 1, \ldots, T$ is the trend term and $\hat{\gamma}_i, \hat{J}_j$ are parameters estimated by OLS. Note that d_{max} does not exceed the true lag length k. Equation (5) can be written as:

$$Y' = \hat{\Gamma}\Lambda + \hat{\Phi}X + \hat{\Psi}Z' + \hat{\mathrm{E}}' \tag{6}$$

where $\hat{\Gamma} = (\hat{\gamma}_0, \ldots, \hat{\gamma}_q)$, $\Lambda = (\tau_1, \ldots, \tau_T)$ with $\tau_t = (1, t, \ldots, t^q)$, $\hat{\Phi} = (\hat{J}_1, \ldots, \hat{J}_k)$, $\hat{\Psi} = (\hat{J}_{k+1}, \ldots, \hat{J}_p)$, $X = (x_1, \cdots, x_T)$ with $x_t = (y'_{t-1}, \ldots, y'_{t-k})'$, $Z = (z_1, \cdots, z_T)$ with

$z_t = (y'_{t-k-1}, \ldots, y'_{t-p})'$ and $\mathrm{E} = (\hat{\varepsilon}_1, \ldots, \hat{\varepsilon}_T)$. As restrictions in parameters, the null hypothesis $H_0 : f(\phi) = 0$ where $\phi = vec(\Phi)$ is tested by a Wald statistic defined as:

$$W = f(\hat{\phi})' \left[F(\hat{\phi}) \left\{ \hat{\Sigma}_\varepsilon \otimes (X'QX)^{-1} \right\} F(\hat{\phi})' \right]^{-1} f(\hat{\phi}) \tag{7}$$

where $F(\phi) = \partial f(\phi) / \partial \phi', \hat{\Sigma}_\varepsilon = T^{-1} \hat{\mathrm{E}}' \hat{\mathrm{E}},\ Q = \hat{Q}_\tau - \hat{Q}_\tau Z (Z' \hat{Q}_\tau Z)^{-1} Z' \hat{Q}_\tau$ and $\hat{Q}_\tau = I_T - \hat{\Lambda} (\hat{\Lambda}' \hat{\Lambda})^{-1} \hat{\Lambda}'$ where I_T is a $T \times T$ identity matrix. Under the null hypothesis, the Wald statistic (7) has an asymptotic Chi-square distribution with m degrees of freedom that corresponds to the number of restrictions. Although Toda and Yamamoto (1995) present this method principally for the purpose of Granger-causality testing, tests based on level VAR equations can also be used to examine long-run relationships. Test results based on the ECM can then be regarded as an indicator of short-run causality, while the causality tests by the level VAR can complement the result of the cointegration tests in terms of long-run information.

3.4 Variance decomposition

One limitation of these tests is that while they indicate which markets Granger-cause another, they do not indicate whether yet other markets can influence a market through other equations in the system. Likewise, Granger causality does not provide an indication of the dynamic properties of the system, nor does it allow the relative strength of the Granger-causal chain to be evaluated. However, decomposition of the variance of forecast errors allows the relative importance of the variance in causing fluctuations in that market to be ascertained. The decomposition process therefore allows the variance of the forecast errors to be divided into percentages attributable to innovations in all other markets and a percentage attributable to innovations in the market of interst. One problem here is that the decomposition of variances is sensitive to the assumed origin of the shock and the order it is transmitted to other markets. To overcome this problem, a generalised impulse response analysis, which is not subject to any arbitrary othogonalisations of innovations in the system, is applied (Masih and Masih, 1999).

The variance decomposition analysis illustrates the system dynamics by decomposing the random variation of one market into component shocks and analysing how these shocks in turn affect prices in other markets. Consider the following VAR model of m market indices proposed by Eun and Shim (1989: 243):

$$y_t = \alpha + \sum_{S=1}^{n} A(S) y_{t-S} + e_t \tag{8}$$

where y_t is a $m\times1$ vector of indices, α and $A(S)$ are respectively $m\times1$ and $m\times m$ coefficients, n is the lag length, and e_t is a $m\times1$ column of forecast errors of the best linear predictor of y_t using past values of y. By construction, if the forecast error e_t is uncorrelated with all past values of y and is also a linear combination of current and past y_t, then e_t is serially uncorrelated. The i,j component of $A(S)$ measures the direct effect of the jth market on the ith market in S periods. As shown by Sim (1980), by the successive substitution of $e_t s$ into y_{t-S}, the VAR model becomes the following moving average representation where the price of each market is a function of past innovations of other markets:

$$y_t = \sum_{S=0}^{\infty} B(S) e_{t-S} \tag{9}$$

Since e_t is serially uncorrelated, the components of e_t may be contemporaneously correlated. To observe the structure of the response of each market to a unit shock in another market within S periods, the error term is transformed by the triangular orthogonalisation procedure. Let $e = Vu$ where V is a lower triangle matrix and u is an orthogonalised innovation from e such that $Eee' = S$ and $VV' = S$ and the transformed innovation u_t has an identity covariance matrix. Equation (9) can then be re-written as:

$$y_t = \sum_{S=0}^{\infty} B(S) V u_{t-S} = \sum_{S=0}^{\infty} C(S) u_{t-S} \tag{10}$$

where $C(S) = B(S)V$. The $i.j$th component of $C(S)$ represents the impulse response of the ith market in S periods to a shock of one standard error in the jth market. From the orthogonalised innovations, the forecast variance of each market can also be decomposed into portions accounted by shocks or innovations from other markets. The orthogonalisation generates the quantity $\sum_{S=0}^{T} C_{ij}^2(S)$, which is the proportion of forecast error variance of y_i due to innovations in y_j. This variance decomposition provides a measure of the overall relative importance of the markets in generating fluctuations in their own and other markets.

4. Empirical results

Table 1 illustrates the panel unit root tests comprising statistics for the LLC t, IPS W and Hadri homoskedastic and heteroskedastic Z tests and their corresponding p-values at price levels and first-differences for the eleven Asian markets. The LLC t test statistic and p-value for the price level series are 0.5582 and 0.7117, respectively. This implies that the sample evidence on the whole panel of eleven Asian markets does not provide sufficient evidence to reject $H_{0,LLC}$. This suggests that there is insufficient evidence to conclude that each individual price level series is stationary. The LLC *t*-test for the first-differenced price series on the whole panel produced a *t*-statistic of -231.7150 and a *p*-value of 0.0000, which concludes the rejection of $H_{0, LLC}$ at the 5 percent level of significance. The rejection of the null hypothesis implies that each price differenced series is stationary. According to the IPS test at price levels across the eleven Asian markets, the ISP *W*-statistic of 1.8923 and *p*-value of 0.9708 suggest that the null hypothesis, $H_{0,IPS,}$ that all cross-section units in the panel are non-stationary cannot be rejected. The ISP panel unit root test then indicates that at price levels all eleven Asian markets are non-stationary. The first differenced series across all eleven Asian markets gave a ISP *W*-statistic of -197.8160 and a *p*-value of 0.0000 thus rejecting the null, $H_{0,IPS}$ which concludes that at least one of the price differenced series in the eleven Asian markets is stationary. Turning to the Hadri homoskedastic and heteroskedastic *Z* tests of the null hypothesis that all series in the panel are stationary; for the price level series, the null hypothesis is rejected with a homoskedastic *Z*-statistic of 29.5510 and a *p*-value of 0.0000 and a heteroskedastic *Z*-statistic of 58.0591 and a *p*-value of 0.0000. This suggests that the price level series for all Asian markets tend to be non-stationary. With respect to the first-differenced series, the Hadri homoskedastic *Z*-statistic of -1.7010 and *p*-value of 0.9555 and the heteroskedastic *Z*-statistic of 0.5902 and *p*-value of 0.2775 fail to reject the required null thus indicating that all price differenced series are stationary.

According to the panel unit root tests, analysis of the price level series indicates non-stationarity in all eleven Asian markets while the first-differenced price series exhibit stationarity. The finding of non-stationarity in levels and stationarity in differences suggests that each index price series is integrated of order I(1). The finding of non-stationarity in levels and stationarity in first differences provides comparable Asia-Pacific evidence to Elyasiani et al. (1998), Masih and Masih (1999) and Worthington and Higgs (2004), amongst others. As a result, the differenced series are used to carry out lag length selection, causality tests and decomposition of the forecast error variance for the markets to be analysed.

<TABLE 1 HERE>

Johansen cointegration tests are used in order to obtain the cointegration rank. The eigenvalues and trace test statistics are detailed in Table 2 for the various null and alternative hypotheses. As multivariate cointegration tests cover all eleven markets rather than simple bivariate combinations they consider the wide range of options available to Asian investors and financiers, as well as the scope of financial integration that may not be reflected in pairwise combinations. The trace test statistic is greater than the critical value for the null hypotheses of r = 0 thereby rejecting the null hypothesis. However, the null hypothesis of r ≤ 1 fails to be rejected in favour r > 1 indicating the order of cointegration is 1. However, similar hypothesis are rejected up to, but not including, r ≤ 4 thereby suggesting an order of integration of four. The primary finding obtained from the Johansen cointegration tests is that a stationary long-run relationship exists between the eleven Asian equity markets over the period 1993-2006.

<TABLE 2 HERE>

Since cointegration exists, Granger causality tests are performed on the basis of equation (4). *F*-statistics are calculated to test the null hypothesis that the first market index series does not Granger-cause the second, against the alternative hypothesis that the first index Granger-causes the second. The calculated statistics and *p*-values for the markets are provided in Table 3. Among the eleven Asian markets, forty significant causal links are found at the 0.10 level or lower. For example, column 1 shows that the Indian, Japanese, Malaysian, Taiwanese and Thai markets Granger-cause the Chinese market. In turn, Thailand (column 10) is found to have a Granger causal relationship with Hong Kong, Japan and Malaysia.

Further insights are gained by examining the rows in Table 3 indicating the effects of a particular market on all markets. It is evident that the Malaysian and Thai markets are among the most influential in the eleven Asian markets. Malaysia influences six markets namely, China, India, the Philippines, Singapore, Taiwan and Thailand. Thailand also Granger causes six markets, namely, China, Hong Kong, India, Korea, the Philippines and Taiwan. The least influential market in terms of Granger-causality is the Philippines.

<TABLE 3 HERE>

There is also an indication that there is feedback at play in several combinations. For example, Thailand Granger-causes Hong Kong and Hong Kong Granger-causes Thailand. Gnereally, high levels of market linkages arise because of the presence of common investor and financing groups. One implication in Table 3 is that there may be no gains from portfolio diversification between those countries where a significant causal relationship exists. Another is that since we have a finding of causality these markets must be seen as violating weak-form efficiency since one of the markets can help forecast the other. A final implication is that the law of one price holds in part: often used as a simple indicator of financial integration. All in all, the presence of Granger causality implies that there are sufficient short-run interrelationships between the markets to believe that some form of financial integration is present.

<TABLE 4 HERE>

The long-run causality Wald test statistics and *p*-values based on Toda and Yamamoto's (1995) level VAR procedure are provided in Table 4. The model is estimated for the levels, such that a significant Wald test statistic indicates a long-term relationship. This serves to supplement the findings obtained from the Granger causality (short run) results in Table 3. Among the eleven markets, fifty-four significant causal links are found at the 10 percent level or lower. For example, column 3 shows that the stock markets in China, Hong Kong, Japan, Korea, Malaysia, the Philippines, Singapore, Taiwan and Thailand affect the Indian market. The rows in Table 4 indicate the effects of a particular market on all markets. It is evident that markets in Japan and Taiwan are the most influential markets in the long run among the eleven Asian equity markets: together these markets impact upon China, Hong Kong, India, Korea, Malaysia Singapore and Thailand. The least influential markets are China and India.

<TABLE 5 HERE>

However, these results should be interpreted with the qualification that Granger causality tests only indicate the most significant direct causal relationship. For example, it may be that some markets influence non-Granger caused markets indirectly through other markets. In order to address this concern, Table 5 presents the decomposition of the forecast error variance for the 2-day, 5-day, 10-day and 15-day ahead horizons for the eleven Asian member equity markets. Each row indicates the percentage of forecast error variance explained by the market indicated in the first column. For example, at the 2-day horizon, the variance in the Chinese market explains 84.39 percent of its own innovations, whereas 6.75 percentage of variance is explained by innovations in the Malaysian market, 3.01 percent by the Singaporean market, 2.44 percent by the Thai market and 1.38 percent by the Japanese market. All Asian home markets explain at least 75 percent of their own innovations with the exception of Singapore and Hong Kong. Singapore influences some 68 percent of its own innovations and Hong Kong only 54 percent. The Singaporean market significantly influences the Malaysian market by 19 percent, even after 15 days. It is readily apparent from the decomposition of the forecast error variance in Table 5 that sizeable differences in the percentage of variance explained by domestic and international markets prevail across Asian stock markets.

5. Concluding Remarks

This paper investigates financial integration and interdependence in Asian equity markets during the period 19938 to 2006. Three of these markets are regarded as developed (Hong Kong, Japan and Singapore) and the majority are viewed as emerging (namely, China, India, Indonesia, Korea, Malaysia, the Philippines, Taiwan and Thailand). Panel unit root tests are used to test for non-stationarity, and multivariate cointegration, Granger causality and level VAR procedures and variance decomposition are conducted to examine the equilibrium and causal relationships among these markets. The results indicate that there is a stationary long-run equilibrium relationship among, and significant and substantial short and long-run causal linkages between, these Asian equity markets. Possible reasons in the absence of explicit regional agreements aimed at financial integration include long-standing trends in trade and investment interaction, the more recent convergence in monetary policies and the almost universal process of economic reform.

The findings obtained in this paper would indicate that three main benefits thought to accrue from financial integration – more opportunities for risk sharing and diversification, the better allocation of capital across investment opportunities, and the potential for higher economic growth – are present in Asian regional markets. Three caveats apply. First, it would appear that the level of financial integration is relatively higher in economies like Japan, Indonesia, Malaysia, Singapore and Taiwan that share many interdependent relationships. The evidence for financial integration is less convincing for economies such as China and India. Second, financial integration in this paper has only been examined in the context of equity markets. Accordingly, no comment can be made on the extent of integration in the bond, money and bank credit markets, or realistically on the financial sector as a whole. Finally, despite the relatively high number of interdependencies and the overall level of integration, Asian domestic markets are relatively isolated. All Asian home markets explain at least 75 percent of their own innovations with the exception of Singapore and Hong Kong: Singapore influences some 68 percent of its own innovations and Hong Kong only 54 percent. As a point of comparison, recent work by Worthington et al. (2003a) found that non-domestic markets explained 48.1 percent of the variance for France, 64.9 percent for Germany, 38.7 percent for Italy, 60.1 for the Netherlands and 65 for Spain. This would indicate, in line with the policy emphasis and interests of the European Union and the European Central Bank, that financial integration is more advanced in at least one other regional setting.

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TABLE 1. *Panel unit root tests*

	Level series		First-differenced series	
	Statistic	*p*-value	Statistic	*p*-value
Levin, Lin & Chu *t**	0.5582	0.7117	-231.7150	0.0000
Im, Pesaran and Shin *W*-stat	1.8923	0.9708	-197.8160	0.0000
Hadri Homoskedastic *Z*-stat	29.5510	0.0000	-1.7010	0.9555
Hari Heteroskedastic *Z*-stat	58.0591	0.0000	0.5902	0.2775

Notes: Period 1/1/1993 – 30/6/2006; hypotheses $H_{1,LLC}$: each series is stationary, $H_{1,ISP}$: at least one series is stationary, H_1 (Hadri homoskedastic and heteroskedastic *Z*-stat) each series is stationary**;** the lag orders are determined by the significance of the coefficient for the lagged terms; for the price levels series intercepts and tends are included; for the first-differenced price series only intercepts are included.

TABLE 2. *Johansen cointegration tests*

H_0	H_1	Eigen-value	Trace test	Critical value
$r = 0$	$r > 0$	0.0209	**384.1003	310.8100
$r \leq 1$	$r > 1$	0.0168	**309.8591	263.4200
$r \leq 2$	$r > 2$	0.0162	**250.4497	222.2100
$r \leq 3$	$r > 3$	0.0146	**193.1358	182.8200
$r \leq 4$	$r > 4$	0.0115	141.4251	146.7600
$r \leq 5$	$r > 5$	0.0088	100.7337	114.9000
$r \leq 6$	$r > 6$	0.0075	69.7911	87.3100
$r \leq 7$	$r > 7$	0.0046	43.4837	62.9900
$r \leq 8$	$r > 9$	0.0037	27.3488	42.4400
Accepted			4	

Notes: Period 1/1/1993 – 30/6/2006; 0.05 percent level critical values from Osterwald-Lenum (1992); the optimal lag order of each VAR model selected using LR tests for the significance of the coefficient for maximum lags and Schwarz's Bayesian Information Criterion; in each cointegrating equation, the intercept and trend are included.

TABLE 3. *Granger causality tests*

Market	CHN	HKG	INA	IND	JPN	KOR	MLY	PHL	SNG	THA	TWN	Causes
CHN	-	6.4837	0.3089	0.3086	0.9569	1.8707	0.0053	4.7128	1.8842	0.6716	0.1241	2
		0.0109	*0.5784*	*0.5786*	*0.3280*	*0.1715*	*0.9420*	*0.0300*	*0.1699*	*0.4126*	*0.7247*	
HKG	0.0179	-	2.3446	0.0399	2.4283	0.0613	0.1052	0.0802	1.9150	9.6039	3.5959	2
	0.8937		*0.1258*	*0.8416*	*0.1192*	*0.8045*	*0.7457*	*0.7770*	*0.1665*	*0.0020*	*0.0580*	
INA	5.1989	0.5544	-	0.1164	1.4946	1.3141	4.5660	36.8903	0.0151	0.5507	0.2066	3
	0.0227	*0.4566*		*0.7329*	*0.2216*	*0.2517*	*0.0327*	*0.0000*	*0.9022*	*0.4581*	*0.6495*	
IND	0.0344	2.8171	5.8639	-	5.7946	22.7973	1.3130	1.0648	0.3068	1.2852	4.3124	5
	0.8530	*0.0934*	*0.0155*		*0.0161*	*0.0000*	*0.2519*	*0.3022*	*0.5797*	*0.2570*	*0.0379*	
JPN	3.0604	9.9014	1.7360	1.0134	-	2.9116	1.6957	0.8648	3.5976	3.6398	2.3474	5
	0.0803	*0.0017*	*0.1877*	*0.3142*		*0.0880*	*0.1929*	*0.3525*	*0.0579*	*0.0565*	*0.1256*	
KOR	2.5248	5.4020	0.5165	1.6619	2.5356	-	0.4042	0.0005	0.2310	0.0015	4.4486	2
	0.1122	*0.0202*	*0.4724*	*0.1974*	*0.1114*		*0.5250*	*0.9820*	*0.6308*	*0.9695*	*0.0350*	
MLY	8.9985	2.6095	12.0606	0.7438	0.4809	1.5631	-	22.8239	6.1288	10.9525	7.6491	6
	0.0027	*0.1063*	*0.0005*	*0.3885*	*0.4881*	*0.2113*		*0.0000*	*0.0133*	*0.0009*	*0.0057*	
PHL	0.5752	0.3425	0.5439	0.6855	0.0934	0.0128	0.7407	-	0.5523	2.3024	1.2172	0
	0.4483	*0.5585*	*0.4609*	*0.4078*	*0.7599*	*0.9098*	*0.3895*		*0.4574*	*0.1293*	*0.2700*	
SNG	0.1470	6.1749	0.0866	0.7330	7.2328	11.7719	0.8190	3.9567	-	2.5915	8.4511	5
	0.7015	*0.0130*	*0.7685*	*0.3920*	*0.0072*	*0.0006*	*0.3655*	*0.0468*		*0.1075*	*0.0037*	
THA	17.9150	7.7108	15.9398	0.0052	0.1038	3.3021	1.6074	19.3920	2.1455	-	3.2638	6
	0.0000	*0.0055*	*0.0001*	*0.9426*	*0.7473*	*0.0693*	*0.2049*	*0.0000*	*0.1431*		*0.0709*	
TWN	10.1054	7.1444	4.1313	0.0120	1.0297	1.3537	1.4657	0.4235	11.0216	0.0286	-	4
	0.0015	*0.0076*	*0.0422*	*0.9128*	*0.3103*	*0.2447*	*0.2261*	*0.5152*	*0.0009*	*0.8656*		
Caused	5	7	4	0	2	4	1	5	3	3	6	40

Notes: Granger causality tests are conducted by adjusting the long-term cointegrating relationship by the ECM; figures in italics are *p*-values; tests indicate Granger causality by row to column and Granger caused by column to row. For example, in the period 1/1/1993 – 30/6/2006, China (row) Granger-causes two markets (Hong Kong and The Philippines) and is Granger-caused by India, Japan, Malaysia and Thailand (using a critical value of 0.10).

TABLE 4. *Long-run causality test by level-VAR*

	CHN	HKG	INA	IND	JPN	KOR	MLY	PHL	SNG	THA	TWN	Causes
CHN	-	13.6576	11.2449	4.1512	7.7310	4.0344	16.9985	5.4987	9.1039	4.2394	5.6179	3
		0.0179	*0.0467*	*0.5279*	*0.1717*	*0.5445*	*0.0045*	*0.3581*	*0.1050*	*0.5155*	*0.3452*	
HKG	0.3806	-	21.6250	5.9600	5.3959	0.9769	14.9117	4.9712	22.4738	17.8617	6.9906	4
	0.9958		*0.0006*	*0.3101*	*0.3695*	*0.9644*	*0.0107*	*0.4194*	*0.0004*	*0.0031*	*0.2213*	
INA	30.8035	18.2633	-	8.8321	8.1790	3.2816	21.9163	41.7116	20.4798	4.1911	7.3631	5
	0.0000	*0.0026*		*0.1160*	*0.1466*	*0.6567*	*0.0005*	*0.0000*	*0.0010*	*0.5222*	*0.1950*	
IND	1.0726	2.1004	7.9991	-	13.8424	34.5463	7.6759	3.7273	5.9462	5.9705	10.8892	3
	0.9565	*0.8351*	*0.1563*		*0.0166*	*0.0000*	*0.1750*	*0.5893*	*0.3115*	*0.3091*	*0.0536*	
JPN	10.2425	14.8819	10.5388	3.4696	-	12.6766	11.4971	2.3142	9.8576	12.9290	6.7797	7
	0.0686	*0.0109*	*0.0613*	*0.6280*		*0.0266*	*0.0424*	*0.8042*	*0.0794*	*0.0241*	*0.2375*	
KOR	5.7747	7.0044	12.7065	13.8854	5.6482	-	4.9194	8.3836	11.7652	3.2207	10.1354	4
	0.3288	*0.2203*	*0.0263*	*0.0164*	*0.3420*		*0.4258*	*0.1363*	*0.0382*	*0.6660*	*0.0715*	
MLY	13.7947	5.9693	31.7159	3.3616	8.2616	3.0022	-	32.4753	18.6217	34.6479	10.0389	6
	0.0170	*0.3092*	*0.0000*	*0.6444*	*0.1424*	*0.6996*		*0.0000*	*0.0023*	*0.0000*	*0.0741*	
PHL	5.2818	2.7629	11.3782	7.4790	4.9019	1.8515	22.3550	-	4.2412	12.5524	10.7344	4
	0.3825	*0.7365*	*0.0444*	*0.1874*	*0.4280*	*0.8693*	*0.0004*		*0.5152*	*0.0280*	*0.0569*	
SNG	7.2173	8.9781	10.3704	11.6187	7.7511	12.8661	3.9583	11.9517	-	3.9503	15.2032	5
	0.2050	*0.1099*	*0.0654*	*0.0404*	*0.1705*	*0.0247*	*0.5554*	*0.0355*		*0.5566*	*0.0095*	
THA	26.5184	13.0652	16.6506	0.3531	7.3727	7.8946	14.2183	32.3982	6.2501	-	11.1922	6
	0.0001	*0.0228*	*0.0052*	*0.9965*	*0.1944*	*0.1621*	*0.0143*	*0.0000*	*0.2826*		*0.0477*	
TWN	27.3577	17.5491	27.8066	3.8022	4.4826	9.9676	10.4432	2.2454	21.8159	10.3081	-	7
	0.0000	*0.0036*	*0.0000*	*0.5782*	*0.4822*	*0.0762*	*0.0636*	*0.8142*	*0.0006*	*0.0670*		
Caused	5	5	9	2	1	4	7	4	6	5	6	54

Notes: Regular figures are Wald statistics for Granger causality tests. Figures in italics are p-values. The level VARs are estimated with a lag order of $p = k + d_{max}$; k is selected by the LR test and d_{max} is set to one. Tests indicate Granger causality by row to column and Granger caused by column to row. For example, Hong Kong (row) Granger-causes four markets (India, Malaysia, Singapore and Thailand) and is Granger-caused by China, India, Japan, Thailand and Taiwan (using a critical value of 0.10).

TABLE 5. *Generalised variance decomposition*

	Period	S.E.	CHN	HKG	INA	IND	JPN	KOR	MLY	PHL	SNG	THA	TWN
CHN	2	1.0111	84.3921	0.0037	0.8844	0.4764	1.3842	0.0677	6.7506	0.0148	3.0086	2.4398	0.5776
	5	1.0134	84.0631	0.0058	0.9324	0.4815	1.3799	0.0683	6.9341	0.0207	3.0114	2.5245	0.5783
	10	1.0134	84.0628	0.0058	0.9324	0.4815	1.3799	0.0683	6.9342	0.0207	3.0114	2.5246	0.5783
	15	1.0134	84.0628	0.0058	0.9324	0.4815	1.3799	0.0683	6.9342	0.0207	3.0114	2.5246	0.5783
HKG	2	75.9325	10.5356	54.0989	0.7064	1.9142	4.6970	0.1504	11.9271	0.0092	11.1174	3.7888	1.0550
	5	75.9480	10.5361	54.0774	0.7086	1.9180	4.6963	0.1506	11.9394	0.0103	11.1159	3.7928	1.0546
	10	75.9480	10.5361	54.0774	0.7086	1.9180	4.6963	0.1506	11.9394	0.0103	11.1159	3.7928	1.0546
	15	75.9480	10.5361	54.0774	0.7086	1.9180	4.6963	0.1506	11.9394	0.0103	11.1159	3.7928	1.0546
INA	2	5.5859	0.0003	0.0730	80.9512	1.8130	0.6846	0.0138	10.6578	0.0139	3.2730	2.4297	0.0897
	5	5.5999	0.0009	0.0799	80.6723	1.8420	0.6840	0.0146	10.8298	0.0179	3.2718	2.4958	0.0910
	10	5.5999	0.0009	0.0799	80.6722	1.8420	0.6840	0.0146	10.8299	0.0179	3.2718	2.4958	0.0910
	15	5.5999	0.0009	0.0799	80.6722	1.8420	0.6840	0.0146	10.8299	0.0179	3.2718	2.4958	0.0910
IND	2	2.2583	0.0075	0.0023	0.0011	98.9235	0.0072	0.0461	0.7331	0.0184	0.0364	0.2222	0.0022
	5	2.2587	0.0085	0.0024	0.0011	98.9166	0.0072	0.0466	0.7361	0.0191	0.0376	0.2225	0.0022
	10	2.2587	0.0085	0.0024	0.0011	98.9166	0.0072	0.0466	0.7361	0.0191	0.0376	0.2225	0.0022
	15	2.2587	0.0085	0.0024	0.0011	98.9166	0.0072	0.0466	0.7361	0.0191	0.0376	0.2225	0.0022
JPN	2	35.5732	0.0030	0.0866	0.0303	1.9242	93.7494	0.0713	3.0364	0.0025	0.3301	0.7546	0.0116
	5	35.5749	0.0039	0.0866	0.0318	1.9280	93.7407	0.0713	3.0362	0.0025	0.3324	0.7547	0.0119
	10	35.5749	0.0039	0.0866	0.0318	1.9280	93.7407	0.0713	3.0362	0.0025	0.3324	0.7547	0.0119
	15	35.5749	0.0039	0.0866	0.0318	1.9280	93.7407	0.0713	3.0362	0.0025	0.3324	0.7547	0.0119
KOR	2	2.7783	0.0612	1.0521	0.1154	5.7162	5.1448	78.7348	1.6538	0.0003	3.7018	1.5282	2.2913
	5	2.7788	0.0616	1.0541	0.1179	5.7313	5.1428	78.7044	1.6580	0.0004	3.7097	1.5289	2.2909
	10	2.7788	0.0616	1.0541	0.1179	5.7313	5.1428	78.7044	1.6580	0.0004	3.7097	1.5289	2.2909
	15	2.7788	0.0616	1.0541	0.1179	5.7313	5.1428	78.7044	1.6580	0.0004	3.7097	1.5289	2.2909
MLY	2	3.6642	0.0003	0.0058	0.1532	0.0692	0.0169	0.0113	99.5101	0.0197	0.0729	0.1126	0.0281
	5	3.6670	0.0004	0.0066	0.1748	0.0762	0.0180	0.0114	99.4541	0.0218	0.0784	0.1295	0.0291
	10	3.6670	0.0004	0.0066	0.1748	0.0762	0.0180	0.0114	99.4541	0.0218	0.0784	0.1295	0.0291
	15	3.6670	0.0004	0.0066	0.1748	0.0762	0.0180	0.0114	99.4541	0.0218	0.0784	0.1295	0.0291
PHL	2	5.2788	0.2694	0.4057	4.2533	0.1253	0.3540	0.0011	8.5444	79.5293	2.6716	3.6693	0.1766
	5	5.3032	0.2847	0.4096	4.3738	0.1361	0.3516	0.0017	8.9548	78.8228	2.6940	3.7905	0.1802
	10	5.3033	0.2847	0.4096	4.3739	0.1362	0.3516	0.0017	8.9551	78.8224	2.6940	3.7906	0.1802
	15	5.3033	0.2847	0.4096	4.3739	0.1362	0.3516	0.0017	8.9551	78.8224	2.6940	3.7906	0.1802
SNG	2	25.1073	0.1220	0.0611	0.0086	2.7401	5.2849	0.0065	19.3444	0.0147	68.0794	4.0941	0.2442
	5	25.1179	0.1255	0.0614	0.0159	2.7383	5.2841	0.0065	19.3735	0.0173	68.0273	4.1041	0.2462
	10	25.1179	0.1255	0.0614	0.0159	2.7383	5.2841	0.0065	19.3735	0.0173	68.0272	4.1041	0.2462
	15	25.1179	0.1255	0.0614	0.0159	2.7383	5.2841	0.0065	19.3735	0.0173	68.0272	4.1041	0.2462
THA	2	4.6930	0.0080	0.2792	0.0464	0.1048	0.0081	0.0001	12.1561	0.0606	0.3211	87.0059	0.0097
	5	4.6980	0.0082	0.2837	0.0597	0.1132	0.0102	0.0010	12.2529	0.0643	0.3378	86.8573	0.0115
	10	4.6980	0.0082	0.2837	0.0597	0.1132	0.0102	0.0010	12.2529	0.0643	0.3378	86.8573	0.0115
	15	4.6980	0.0082	0.2837	0.0597	0.1132	0.0102	0.0010	12.2529	0.0643	0.3378	86.8573	0.0115
TWN	2	4.4228	0.0052	0.1161	0.0048	1.8886	2.1371	0.1221	3.1234	0.0323	2.8602	0.5475	89.1627
	5	4.4242	0.0058	0.1182	0.0050	1.9032	2.1371	0.1227	3.1434	0.0324	2.8700	0.5532	89.1092
	10	4.4242	0.0058	0.1182	0.0050	1.9032	2.1371	0.1227	3.1434	0.0324	2.8700	0.5532	89.1092
	15	4.4242	0.0058	0.1182	0.0050	1.9032	2.1371	0.1227	3.1434	0.0324	2.8700	0.5532	89.1092

Notes: The ordering of the variance decomposition is based on the number of 'causes' in Table 3; the four rows shown for each market are for forecast periods of 2, 5,10 and 15 days, respectively