Structural Breaks and Diversification: The Impact of the 1997 Asian Financial Crisis on the Integration of Asia-Pacific Real Estate Markets

Patrick Wilson
Ralf Zurbruegg
Richard Gerlach

ISSN: 1036-7373
Abstract

Currently, there exists relatively little research on the influence that the 1997 Asian financial crisis has had upon capital flows within the securitized property market and the associated long run implications of it. This paper examines the impact that the crisis has had upon the integration and dynamic links between a number Asia-Pacific real estate markets. This is achieved through the use of multivariate cointegration analysis that determines and accounts for structural breaks endogenously. The procedures used include those developed by Inoue (1999) and Johansen, Mosconi and Nielsen (2000). The results show that the integration of Asia-Pacific property markets is prevalent despite a structural shift, and that the benefits to securitized real estate diversification maybe far less than originally perceived. These results are a particularly important finding for fund managers concerned with the impact of globalization on the performance of their real estate portfolios.

JEL codes: G15, F30, C22, C52

Key Words: cointegration, structural breaks, Asian financial crisis, real estate markets

**Corresponding Author:**
Patrick J. Wilson
Associate Professor of Finance
School of Finance and Economics
University of Technology, Sydney
PO Box 123
Broadway NSW 2007
Australia
email: Patrick.Wilson@uts.edu.au
fax: 61(0)295147711
phone: 61(0)295147777

*This research has received financial support from the Faculty of Business*
Structural Breaks and Diversification: The Impact of the 1997 Asian Financial Crisis on the Integration of Asia-Pacific Real Estate Markets

Abstract

Currently, there exists relatively little research on the influence that the 1997 Asian financial crisis has had upon capital flows within the securitized property market and the associated long run implications of it. This paper examines the impact that the crisis has had upon the integration and dynamic links between a number Asia-Pacific real estate markets. This is achieved through the use of multivariate cointegration analysis that determines and accounts for structural breaks endogenously. The procedures used include those developed by Inoue (1999) and Johansen, Mosconi and Nielsen (2000). The results show that the integration of Asia-Pacific property markets is prevalent despite a structural shift, and that the benefits to securitized real estate diversification maybe far less than originally perceived. These results are a particularly important finding for fund managers concerned with the impact of globalization on the performance of their real estate portfolios.

JEL codes: G15, F30, C22, C52

Key Words: cointegration, structural breaks, Asian financial crisis, real estate markets
1. Introduction

International diversification in real estate has recently become a more important issue among academics because the evidence is not entirely clear on the benefits from diversification across property markets. Portfolio managers are faced with the persistent problem of maintaining investment returns while simultaneously reducing risk. This outcome is achieved through portfolio diversification - allocating resources across a number of asset classes and sub-classes as well as across countries. Ideally such managers seek investments in markets that are insulated from each other so that, in particular, the effects of a collapse in one market, or one segment of the market, are not transmitted to investment holdings in other areas. Hence the notion of market segmentation is of prime importance in property portfolio management.

The outcomes from academic research on property markets over the last few years, however, have been unable to reach a firm conclusion on whether international diversification in real estate is beneficial. Much research has centered on the issue of market segmentation/integration. For instance, if international property markets are well integrated then little gain in risk reduction may be achieved through holding internationally diversified investments. Well integrated property markets can also imply that such markets may respond to the same economic stimuli, thereby providing little gain in diversifying across these markets. On the other hand, if markets are clearly segmented and respond to different economic stimuli, then it is important for portfolio managers not only to diversify, but to be able to allocate resources in a dynamic fashion so as to take advantage of changing conditions in each market.
Moreover, even if markets are integrated, the impact that a shock transmission may have upon these markets may differ. This can be an important issue if there are structural shifts in one or more of the asset series making up the portfolio since the existence of a structural break may disguise the true nature of any potential relationships between assets within the portfolio. This may be particularly crucial to those portfolio managers concerned with strategic asset allocation – i.e. the diversification strategies to be pursued over the long term.

This paper aims to analyze the effect that the 1997 Asian financial crisis had on the interdependence among several Asia-Pacific real estate markets. In particular, the paper is concerned with whether the benefits of diversifying across such real estate markets may have altered because of the crisis. Not only will this contribute to the research on whether property markets are integrated, but also provide a very useful look at how the real estate markets adapted to the Asian crisis, which some literature argues was the original catalyst for the crisis itself\textsuperscript{1}. The finding is that failure to take into account the events of 1997 disguises the true nature of the long run inter-linkages between these property markets. Specifically, the real estate markets in the study group of countries are found to be cointegrated when allowance is made for the 1997 crisis. This finding of cointegration has important implications for property portfolio managers within any country making up the cointegrated system since such managers not only have to be aware of influencing events within their own market, but also events in the other cointegrated markets.

The structure of the rest of the paper is as follows: section two briefly considers some recent literature on diversification in real estate markets; section three considers
how the methodology of cointegration can be used to indicate the likely diversification benefits that exist between markets; section four describes the data and results; while section five offers some conclusions.

2. A Literature Review on Real Estate Diversification

Since the jury is still out on the benefits of international diversification in real estate there is, accordingly, a body of literature suggesting that property markets are segmented and, consequently, research showing the opposite.

Evidence illustrating that markets are integrated include research by Myer, Chaudhry and Webb (1997). Using the Johansen cointegration methodology on appraisal based property series across three countries (US, Canada and the UK) they found that these series were highly cointegrated. More recently, Case, Goetzman and Rouwenhorst (2000), using appraisal based property data over 22 countries, presented strong evidence to support the notion of globalisation of property markets. These authors argued that, since property is location specific there would, on an intuitive level, be no reason to suppose that such markets should be linked. Quite significantly their research, in fact, suggested that world real estate markets are correlated and that this correlation was due, in part, to common exposure to fluctuations in the global economy, as measured by an equal weighted index of international GDP changes.

Research showing international property markets are segmented include, among others, Ziobrowski and Curcio (1991) and Sweeney (1993). Eichholtz, Huisman, Koedijk and Schuin (1998) also found segmentation generally between continents but integration within continents. This was particularly so for Europe and true to a lesser
extent for North America. Eichholtz et al. found, for example, that European investors would need to look outside Europe for diversification benefits. Interestingly, these authors did not find a continental factor for the Asia-Pacific region. Liu and Mei (1998) point out that international property markets are segmented and that there are benefits to international diversification in real estate. More recent research by Eichholtz, Koedijk and Schweitzer (2001) suggests that, while there are benefits to international diversification, there is a trade-off between the benefits and costs of such diversification. Their findings suggest that property investors can gain substantially in terms of reduced costs by investing in securitised property companies that concentrate on their local, domestic market.

Clarifying this issue on segmentation vs integration is important because market integration implies reduced, or no diversification benefits and portfolio managers need such information so appropriate diversification strategies can be implemented. In addition it is important to know whether the degree of integration across assets and countries vary according to different economic climates (different regimes). That is, does the existence of a structural break impact on the benefits from diversifying internationally? Here the temporal instability of the correlation structure is an important issue. Tarbert (1998) has raised concern over the dangers of using conventional correlation techniques in preliminary portfolio construction due to the temporal instability of such correlations, pointing to earlier work on this by Baum and Schofield (1991). The main difficulty revolves around the idea that, since correlation coefficients are temporally unstable, a well diversified portfolio initially selected through correlation analysis in one period may not hold in subsequent periods. In a move away from correlation analysis
Tarbert (1998) applied cointegration techniques for initial property portfolio selection and found that the potential risk reduction benefits of property diversification by region and sector within the UK were more limited than previously thought.

Temporal instability of correlation coefficients may be caused, in part, by structural breaks in one or more of the assets under consideration. From a diversification viewpoint, such instability is undesirable since an underlying tenet of diversification is that shocks to assets in the diversified portfolio are asset specific. The direction of movement in the coefficient caused by a shock can have important ramifications for asset allocation since this will not only influence the assets entering the initial portfolio, but also the weightings of assets in the portfolio.

There has been some interest among researchers on the impact of the Asian crisis on property markets, but such researchers have focused on the interdependence between real estate and other asset classes. For example, Renaud (2000) noted that the interdependent roles of real estate and banking in the Asian crisis has highlighted the conspicuous need for much better price and quantity monitoring of real estate cycles. Research by Renaud, Zhang and Koeberle (2001) suggests a real estate crisis in 1996/97 in Thailand precipitated a domestic financial crisis whose large cost was further amplified by a currency crisis in 1997, and it was from this point that the crisis spread quickly to financial and property assets held in other economies. In contrast, a study by Kim (2000) on the Korean real estate market presented strong evidence to suggest that the real estate sector could not have been a major cause of the economic crisis in that country. From an international diversification perspective, and the focus of this paper, the important question is whether there are common linkages between Asia-Pacific real estate markets
and whether these linkages may have altered due to the crisis. Answers to these questions will determine the potential existence of either long term or short-run diversification benefits in holding international real estate assets and the impact foreign markets can have upon domestic securitized property prices.

3. Methodological Issues

A conventional approach to initial asset selection in portfolio construction is through the use of cross correlation analysis. However, as noted earlier, Tarbert (1998) points to the dangers of solely relying on correlation analysis because of temporal instability of correlation coefficients. Also Forbes and Rigobon (2002) show that conventional cross-correlation coefficients are biased upwards during a period of increased volatility in just one of the relevant variables (markets). This may occur during a turbulent period such as the 1997 Asian crisis.

In cointegration analysis, as with correlation analysis, failure to consider possible structural breaks in one or more of the series can impact on the results. In the event that the structural break is known then one approach may be to sub-sample either side of the break. For example, Sheng and Tu (2000) used this approach for examining stock market data sampled before and during the Asian financial crisis. Their research suggested that stock markets were not cointegrated before the crisis of 1997 but that there was some degree of cointegration during the crisis. Sheng and Tu pre-judged the sampling break – a procedure which may or may not have impacted on their results.

This study will similarly use a cointegration framework, but will depart from the practice of pre-selecting sub-sampling periods. Instead we will adopt the procedure
developed by Inoue (1999) for determining a potential structural break endogenously within a multivariate cointegrated system. The Inoue (1999) procedure allows for a test of cointegrating rank within the presence of a mean- or trend-break. A significant advantage from an analyst’s viewpoint here is the fact that this is a Johansen (1988, 1991) type test and does not require prior specification of the structure of a cointegrating system. That is, a whole portfolio can be analysed in one pass to examine the number of common linkages that may exist among assets (real estate markets of different countries) given the presence of an unknown structural break.

Distinctively, the Inoue (1999) procedure has an advantage over alternative tests for cointegration in the presence of structural breaks when one is trying to determine the cointegrating rank. This is particularly useful when a multivariate system is under analysis. In the presence of breaks, standard Johansen based tests may incorrectly infer that no, or only a limited number of, cointegrating vectors exist, when the system may in fact be highly cointegrated. As the rank of a system is intrinsically linked to the number of common stochastic processes present within a system (see Stock and Watson, 1988) the rank can reveal useful information relating to how integrated the system is over the long-run. For example, since cointegrated variables share common stochastic trends, if the cointegrating rank, \( r \), of a system is, say, \( r = n-1 \), then there is a single common trend (i.e. \( n - r = 1 \)) driving all \( n \) series. In economic terms, such a scenario would lead to no diversification benefits in the long-term as all the markets will follow the same long-run trend. That is, once the system is pushed out of equilibrium by some shock there is a single, common force pulling all variables (countries) back towards equilibrium.
Therefore, knowledge of the cointegrating rank of a system can help determine the degree of integration prevalent within the markets under analysis.

The Inoue (1999) methodology follows closely that of the Johansen type tests. Three models are examined (A, B and C) that allow for possible mean and trend breaks, with model B being recommended when the form of a possible break is unknown. Intuitively, Inoue’s procedure involves sequentially estimating a model that incorporates a step dummy to account for a possible break. From this it also provides an indication as to where the break most likely will have occurred, based upon the relative size of the eigenvalues obtained from the method described below.

As Inoue (1999) outlines in his paper, the models can be written as $n$-dimensional vector autoregressions (VAR) such that:

$$ Y_i^i (\xi_0) = \sum_{j=1}^{p} \Phi_j Y_{i-j}^i (\xi_0) + u_i, i = A, B, $$

$$ Y_t^A (\xi) = X_t - \mu - \tilde{\mu} U_t (\xi), $$

$$ Y_t^B (\xi) = X_t - \mu - \tilde{\mu} U_t (\xi) - \delta t - \tilde{\delta} T_t (\xi), $$

$$ Y_t^C = c + \tilde{\mu} U_t (\xi_0) + \sum_{j=1}^{p} \Phi_j Y_{i-j}^C + u_i, Y_i^C = X_t, $$

where $\{\phi_j\}_{j=1}^{p}$ are $n \times n$ matrices, $u_t \sim NID(0, \Omega), c, \mu, \tilde{\mu}, \delta$ and $\tilde{\delta}$ are $n$-dimensional vectors, $DU_t (\xi) = I(t > [T\xi])$ and $DT_t (\xi) = (t - [T\xi])I(t > [T\xi])$ where $I(\xi)$ denotes an indicator function with $[x]$ being the integer component of $x$. $\xi_0$ is the break fraction.

The above equations can also be written in an error-correction form (see Inoue, 1999), such that:
\[
\Delta Y_t^i(\xi_0) = \Pi Y_{t-1}^i(\xi_0) + \sum_{j=1}^{q-1} \Gamma_j \Delta Y_{t-j}^i(\xi_0) + u_i, i = A, B, \tag{3}
\]

\[
\Delta Y_t^c = c + \tilde{\mu} D U_t(\xi_0) + \Pi Y_{t-1}^c + \sum_{j=1}^{p-1} \Gamma_j \Delta Y_{t-j}^c + u_t, \tag{4}
\]

where \( q \) and \( r \) are integers for \( 1 \leq q \leq n \) and \( 0 \leq r \leq q \), and \( \alpha \) and \( \beta \) are \( n \times q \) matrices such that \( \alpha \beta' = \Pi \) where \( \{\Gamma_j\}_{j=1}^{p-1} \) and \( \Pi \) are \( n \times n \) matrices. From this Inoue (1999) develops trace and maximum eigenvalue statistics that are similar in taxonomy to Johansen (1988, 1991) with the null hypothesis of

\[
H_0: \text{rank}(\alpha) = \text{rank}(\beta) \leq r, \quad \tilde{\mu} = \delta = 0_{n \times 1},
\]

being tested against either the alternative:

\[
H_1: \text{rank}(\alpha) = \text{rank}(\beta) > r.
\]

using the trace statistic:

\[
\sup_{\xi \in \Xi} \{-T \sum_{j=r+1}^{T} \ln (1 - \hat{\lambda}_j(\xi))\}; \tag{5}
\]

or, by applying the maximum eigenvalue statistic:

\[
\sup_{\xi \in \Xi} \{-T \ln (1 - \hat{\lambda}_{r+1}(\xi))\} \tag{6}
\]

one can test against the alternative:

\[
H_2: \text{rank}(\alpha) = \text{rank}(\beta) = r + 1
\]

Inoue (1999) provides asymptotic critical values for these test statistics as well as evidence that these tests perform as well, if not better, than the Gregory and Hansen (1996) residual-based tests for cointegration with breaks, plus are more appropriate than the standard Johansen methodology where breaks are present.
The above test statistics are therefore calculated in the empirical section to provide a direct comparison with the standard Johansen tests in order to determine the rank, as well as to determine the potential benefits to diversification, once the 1997 crisis is explicitly taken into account.

Finally, to examine in more depth the cointegrative relationship between the various real estate markets, we construct a cointegrative model that includes intervention dummies to account for possible breaks determined by the Inoue(1999) procedure. Following Johansen, Mosconi and Nielsen (2000), we, again, determine the rank of the system to check of consistency, before conducting exogeneity and exclusion tests within the model and the cointegrating space.

What will these tests tell us? Applying a combination of various cointegration tests will allow us to determine how any given structural break has affected the flow of potential benefits from diversifying internationally, albeit within a similar geographic region. In addition, by also conducting tests on exogeneity and exclusion a number of very useful questions can be addressed, including the extent to which the Asian crisis may have affected international linkages in real estate markets; whether managers may need to adjust the composition of their property portfolios; and on the extent to which portfolio managers need to be aware of foreign country events as well as their own, in making portfolio decisions.

4. Data Description and Results

4.1. Data

As it is the intent of this paper to show how possible inter-relationships may have altered due to the crisis, four countries are chosen that should, a priori, be fundamentally
linked within the East Asian Region. These are: Singapore, Malaysia, Hong Kong and Japan. Japan enters as the premier economy in the Asia region, and therefore a potential driver for the regional economy, while the remaining countries are members of ASEAN, an important trading alliance in the area. Moreover, Hong Kong, Singapore and Malaysia are all major trading partners with one another, which would lead to the expectation that these markets maybe strongly integrated. Data limitations prohibited the inclusion of other important regional economies such as Thailand, Indonesia and Korea. Inability to include Thailand due to the unavailability of sufficient real estate data is particularly disappointing since it has been suggested this country was an important player in the Asian financial crisis (cf. Renaud, Zhang and Koeberle (2001)).

Other countries, either individually or as groups, could have been incorporated into the analysis. Both the US and the EU could have entered as major economic forces, with direct and indirect trading links to most countries. However, the primary objective of this paper is to examine inter-linkages amongst Asian real estate markets and the possible impact of the 1997 crisis. While the 1997 Asian crisis may have had effects reaching beyond the Asian region, that issue can be an emphasis for future research.

Weekly price return indices from the first week in January, 1993 to the end of 2001 was obtained from Datastream International for each of the study countries. These were exchange and CPI adjusted for US dollars and the base date was set to January, 1993. The start date was chosen to avoid problems with changes to market regulations that were far more restrictive prior to 1993 in some of the markets. It is also worthwhile drawing attention to the fact that our analysis uses an aggregate index of a number of companies for each country. Although these companies invest mostly in their own
domestic markets, some may have property investments abroad including the other countries under analysis. This implies that a certain degree of integration is at least expected. However, this integration may still not show up if the 1997 crisis has severely distorted any standard cointegration analysis.

While all analyses were undertaken on the natural logarithm of the data figure 1 plots the series in unlogged form to obtain a clearer visual impression prior to our analysis. A visual inspection will also show how all the series, with the possible exception of Japan, roughly follow a positive drift over time until the 1997 crisis. The Japanese economy has had difficulty recovering from the general recession of 1990/91 and this has been reflected in its property sector. To illustrate the impact of the 1997 crisis on the Japanese property market we have scaled Japan on the secondary (RHS) axis in figure 1. We note that from the time of the crisis, and particularly between mid 1997 to the end of 1998, a significant change occurs within all the price series before resuming something close to the previous pattern prior to the crisis.

Table 1 provides a brief description of the returns series for each country. From the table we see that none of the returns series are normally distributed. There is some positive skewness in both Japan and Malaysia and negative skewness in both Hong Kong and Singapore, although the skew for Singapore is only very slight. We would expect the thickness in the tails as indicated by the excess kurtosis since this has been likely generated by the abnormal returns due to the turmoil of 1997.

4.2. Unit Root Tests

Prior to undertaking any cointegration analysis it is necessary to ascertain the degree of integration for each series. However, research by Perron (1989) has shown that
the existence of a structural break in a series can affect its stationarity properties. Unit root testing procedures developed by Zivot and Andrews (1992) allow for the testing of a unit root in the presence of a possible structural break in the series (a break in the intercept, slope or both – their models A, B and C). Table 2 presents the outcomes from both conventional Augmented Dickey-Fuller (ADF) unit root tests along with Zivot and Andrews (ZA) unit root tests. Both ADF and ZA tests indicate that the series are I(1). Also, all of the significant ZA results indicate a break date about mid to late 1997, which coincides with the Asian financial crisis. While the ZA tests are univariate tests it is quite clear that breaks in the univariate series could well have implications for possible breaks in any cointegrating relationships.

4.3. Johansen Rank Tests

As both figure 1 and the Zivot and Andrews tests in table 2 indicate a break around the time of the 1997 financial crisis, the standard cointegration tests presented are not only conducted for the full sample but also for a pre-crisis period and post-crisis period. Doing so may help determine whether there has been a change in the cointegrating rank before and after the crisis. Moreover, a comparison of the sub-sample results with the full data set may reveal the impact of the crisis upon the cointegration process.

To choose a lag order for the VAR several criteria were initially examined. The standard Akaike and Schwartz information criteria were applied but returned an optimal lag order of 2. Unfortunately, the residuals from taking such a short lag structure were highly non-Gaussian and therefore a lag of 26 was chosen as this was the minimum lag
providing diagnostic results that returned uncorrelated residuals. It would also seem that the lag length chosen is not dissimilar to other studies utilizing weekly data. Manning (2002) discusses this issue in relation to using stock index prices. Table 3 provides residual based tests demonstrating that none of the series show evidence of autocorrelation and only Hong Kong and Singapore display ARCH effects.

To split the data into two sub-samples the earliest breakpoint identified by the Zivot and Andrews test is used (29th July 1997) as an indicator for the end-point of the first sample. Similarly, we chose the start of the second sub-sample to begin at the last significant breakpoint established for any of the series by the Zivot and Andrews tests. For our data this is the 21st October 1997.

The results for the Johansen rank tests are tabulated in table 4. The procedure allows for a constant to lie within the cointegrating space but not outside it (i.e. no linear deterministic trend) as table 1 shows little evidence that E[Δy_t] ≠ 0.

The results do indeed show differing outcomes for the sub-samples and full sample. For the full sample the results show no presence of any cointegrating equations, suggesting these real estate markets are not integrated and may offer substantial diversification benefits over the long run. However, as previously discussed, the result for the full sample may be erroneous as the Johansen rank tests have not taken into account the possibility of a break within the data, consequently the extent of diversification benefits to a portfolio manager may be exaggerated.

Also, in table 4 we see that both sub-samples show evidence of cointegration. In fact, the results strongly support the presence of at least two cointegrating equations at the 1% significance level. With evidence of a long run equilibrium condition, portfolio
managers must be aware of not only events in their own countries that may impact on the investment performance of their property portfolios, but must also consider events in other countries since these may also impact on the long run performance of their portfolios.

4.4. Parameter Constancy Tests

One possible reason for the results to differ between the sub-periods and full sample is from the model parameters not remaining constant for the full data set. This may well be due to the influence of the crisis. As discussed by Hansen and Johansen (1992, 1999) a simple test for parameter constancy can be applied under the estimated rank of the system. Assuming a rank of two for the full sample, the Johansen procedure is re-run and an analysis of the recursive eigenvalues is then conducted to test whether the cointegrating equations are stable, under the null hypothesis of sample independence for recursive sub-samples. The test statistic is a likelihood ratio given by:

\[
LR = T \sum_{j=1}^{r} \left\{ \ln[1 - \hat{\rho}_j(\tau)] - \ln[1 - \hat{\lambda}_j(\tau)] \right\} \quad \text{for} \quad \tau = T_0, \ldots, T
\]

where \( T_0 \) is the sub-sample size, \( T \) is the full sample size, \( r \) is the hypothesized number of cointegrating vectors, and \( \hat{\rho}_j(\tau) \) and \( \hat{\lambda}_j(\tau) \) are the restricted and unrestricted solutions to the eigenvalue problems. This likelihood ratio is distributed as a Chi-square and figure 2 shows the Chi-square test statistic computed over the sample period. It is noticeable that in June 1997 the null of parameter constancy is rejected, occurring within the same timeframe as the Zivot and Andrews breakpoints. This essentially indicates that even if
two cointegrating equations are present within the full sample, there is a break in the data that leads these relationships to be unstable, even if it is for a short period.

4.5. Rank Tests Accounting for the Crisis

To correct for what seems to be the presence of a break within the integrated real estate markets, the Inoue results for all three models are presented in table 5 for the full sample period. The results show the presence of at least two cointegrating equations at the 5% level for all three model types and at least one cointegrating equation at the 1% level.

These results also compliment the Johansen sub-sample test results shown in table 4. As the rank has remained the same, it may be that the crisis has had little impact upon the integration of the markets, resulting in nothing more than a simple structural shift in the cointegrating relations. Interestingly, in their study on stock market co-movements, Forbes and Rigobon (2002) developed a simple adjustment procedure for examining correlation coefficients and then analyzed a number of stock markets associated with various crisis events such as the Asian crisis of 1997, the 1994 Mexican peso collapse and the 1987 US stock market crash. Their finding was strongly supportive of stock market interdependence. That is, cross market linkages did not change significantly either pre- or post- the particular event. Such a result supports our own finding as to the impact the 1997 crisis has had on the degree of integration upon the property markets. Moreover, the likely breakpoints given by the Inoue process are either in June or July 1997, coinciding closely with both the date that the recursive parameter constancy tests are first rejected, as well as with the Zivot and Andrews breakpoints.
In table 6 we also present the results of the Johansen procedure for the full sample, only this time we have included intervention dummies to account for the break as identified by the Inoue process. A step dummy (*permdum*) was added into the cointegration space to account for any permanent shift. The dummy was implemented such that it took the value of zero until 15\textsuperscript{th} July 1997 and unity afterwards. The break date being the earliest breakpoint identified by the Inoue procedure\textsuperscript{7}. Also, to account for any transient effect, an unrestricted impulse dummy (*transdum*) was also incorporated in the model, taking the value of 1 at time *t*, -1 at time *t+1* and zero otherwise.

It is evident from this table that, once the turmoil of the Asian crisis has been taken into consideration, there are, again, at least two cointegrating equations. Critical values at the 95% percentile were derived from Johansen, Mosconi and Nielsen’s (2000) distribution where we allow for one break date within the sample. The results show there are two common stochastic trends pushing this system of real estate markets towards equilibrium.

\textit{4.6. Exogeneity and Exclusion Tests}

The analysis so far has clearly indicated that these Asian securitized real estate markets are interlinked, with a structural shift during the 1997 crisis. The interesting question now is whether one or more of these property markets are influencing the long run development of the other markets, but is not influenced by them (i.e. there is no ‘levels feedback’ effect). In other words, can one or more of these real estate markets be considered primary drivers within the system? This is a particularly important issue for a portfolio manager because identifying such a primary driver (or drivers) would permit the
manager the option of focusing on what is happening in the driver market. Several tests for this purpose can be performed and we begin by testing for weak exogeneity as presented in table 7.

To read table 7 we view each country sequentially as a LHS variable in the error correction model. To examine long-run weak exogeneity, we check to see if the alpha coefficients for each country are significantly different from zero. The results of this Chi-square test show that the error-correction terms are indeed highly significant, at the 1% significance level, for determining market returns in each country. This indicates that each of these countries are influenced by long-run equilibrium relationships prevalent within the system. Therefore, if there is a disturbance to the long run equilibrium of this system, then the forces pushing the system back towards equilibrium will, either directly or indirectly, have some impact on all the property markets.

For the short-run dynamics, block exogeneity tests on the joint significance of each market’s lagged returns show some interesting results. First, in the short-run Hong Kong returns have no significant impact upon any market. However, Malaysian lagged returns have a significant impact on all other markets. Japanese returns affect Malaysia and Hong Kong, whereas Singapore returns influence Japanese and Malaysian markets. The impulse dummy is significant for Japan and Hong Kong, where it is also negative as would be expected. However, as the crisis would have affected each market at a slightly different interval, the transitory dummy is probably not that useful to explain all of the markets downturns at a particular point in time.

To place a somewhat intuitive interpretation on the above results we might briefly return to figure 1. In table 7 we found that Malaysian lagged returns affected all other
markets. In figure 1 we see that, of the four study countries, the Malaysian property market appeared to be the first to turn down, with Japan and Singapore following shortly thereafter. It would seem that the last market to fall into the crisis was Hong Kong. This casual view of the timing certainly appears to support the outcome from the statistical analysis and may in part explain the behaviour of these markets for the whole sample. Also, economic and geographic ties can explain some of the short-run relationships. Singapore and Malaysia, for example, are neighbors sharing many economic conditions and both may very easily influence each other’s property market in the short-term.

However, to ascertain any real long-run influences that one market may have on another, an examination of whether any or all of the countries form part of the cointegrating space needs to be conducted. If it turns out that not all the countries help drive the common trends, then it can be argued their long-run impact upon the other Asia-Pacific property markets is not permanent.

Table 8 presents tests to determine whether all or only some of the markets form part of the cointegrating space. Over-identifying zero exclusion tests were performed on each country respectively and the $p$-values show that apart from Japan and the permanent shift dummy, none of the variables are significant at the 1% level. This means that Japan will always form part of the cointegrating space within this system of property markets, at the 99% significance level.

This outcome is quite reasonable in view of the prominence of the Japanese economy within the region. However, and as an explanation for there being at least two cointegrating equations, at the 5% level Singapore is significant and at 10% so is Hong
Kong. While Japan is part of the high income OECD countries both Singapore and Hong Kong are classified by the World Bank as high income, dynamic economies. Both of which rely heavily on trade, within the South East Asian region and in the broader world economy. Therefore, even if the individual property markets themselves do not directly influence each other, as securitized real estate their returns will in part be related to general economic conditions and forecasts. Moreover, as Barberis, Shleifer and Wurgler (2002) indicate that general comovements between indices can also be explained by ‘category-based’ comovements in particular sectors and securities, a downturn in a major property market, in say, Japan, may very well lead to a drop in property prices in Malaysia as investors shift money away from securitized real estate. However, it is unlikely that the reverse would also be true, given the relatively small size of the Malaysian market.

5. Summary and Conclusions

This paper set out to examine the question of whether a select group of real estate markets in the Asia Pacific region were interlinked, whether such interlinkages might be influenced by the Asian crisis, and what the implications of the finding would mean for diversification into Pacific-Rim real estate markets. The tool chosen to examine this was cointegration analysis. Specifically we queried whether Asian real estate markets were integrated over the long term when accounting for (unknown) structural breaks.

The results of this paper illustrated that if the possibility of a structural break within the data is ignored, then the conventional Johansen procedure may yield incorrect results and lead a portfolio manager to erringly diversify across assets that are not likely to yield as good a risk reduction benefit as anticipated. We showed this in three ways
viz. (i) through the use of the Johansen methodology on a simple data split pre- and post-crisis in which we demonstrated that the number of stochastic processes were higher than was the case when the crisis was ignored; (ii) through the use of a parameter constancy test on the supposition that the pre- and post- number of cointegrating equations held for the full sample; and (iii) through the use of the Inoue technique which allows for structural breaks. The impact of the crisis seemed to primarily create a structural shift in the property markets which are cointegrated with two stochastic processes both before the crisis and after it.

In addition we used exogeneity and exclusion tests to determine if there were important long run drivers in the system, and it appeared that Japan, at the 1% level, and Singapore, at the 5% level, could be considered important factors in influencing the long run equilibrium within these markets. From a portfolio manager’s perspective, the influence from these markets, particularly Japan, cannot be ignored when examining local property market performance. Moreover, the above results not only show that there are both long and short-run linkages between the property markets within the Asia-Pacific Region, but that the impact of the crisis has had little impact upon the degree of integration within these markets.
6. **Bibliography**


Figure 1. Asian Property Market Indices.

Figure 2. Recursive Parameter Constancy Test.

The likelihood ratio test statistics on the vertical axis have been scaled by the 1% critical value. Values greater than unity indicate rejection of the null of parameter constancy.
Table 1. Descriptive Statistics.

<table>
<thead>
<tr>
<th>Country</th>
<th>Mean</th>
<th>Std Dev</th>
<th>Skewness</th>
<th>Excess Kurtosis</th>
<th>Normality Test</th>
</tr>
</thead>
<tbody>
<tr>
<td>Japan</td>
<td>-6.482e-04</td>
<td>0.0424</td>
<td>0.2789</td>
<td>1.3064</td>
<td>25.342(^a)</td>
</tr>
<tr>
<td>Malaysia</td>
<td>-9.871e-04</td>
<td>0.0659</td>
<td>0.9296</td>
<td>10.7080</td>
<td>380.42(^a)</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>1.064e-03</td>
<td>0.0567</td>
<td>-0.5598</td>
<td>5.4482</td>
<td>185.63(^a)</td>
</tr>
<tr>
<td>Singapore</td>
<td>-8.273e-05</td>
<td>0.0569</td>
<td>-0.0679</td>
<td>4.9463</td>
<td>204.36(^a)</td>
</tr>
</tbody>
</table>

All statistics are for logarithmic returns. The normality test is derived from Doornik and Hansen (1994) and is \(\chi^2\) distributed. \(^a\) indicates rejection of the null of normality at the 1% level.

Table 2. Unit Root Tests.

<table>
<thead>
<tr>
<th>Country</th>
<th>ADF Tests (Levels)</th>
<th>ADF Tests (Returns)</th>
<th>Zivot and Andrews Tests</th>
</tr>
</thead>
<tbody>
<tr>
<td>Japan</td>
<td>-3.71</td>
<td>-9.66(^a)</td>
<td>-6.04(^a)</td>
</tr>
<tr>
<td></td>
<td>(21-10-97)</td>
<td>(21-10-97)</td>
<td>(21-10-97)</td>
</tr>
<tr>
<td>Malaysia</td>
<td>-2.50</td>
<td>-8.61(^a)</td>
<td>-4.68(^c)</td>
</tr>
<tr>
<td></td>
<td>(29-07-97)</td>
<td>(05-08-97)</td>
<td>(05-08-97)</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>-2.95</td>
<td>-8.80(^a)</td>
<td>-5.24(^b)</td>
</tr>
<tr>
<td></td>
<td>(7-10-97)</td>
<td>(14-10-97)</td>
<td>(14-10-97)</td>
</tr>
<tr>
<td>Singapore</td>
<td>-2.36</td>
<td>-8.72(^a)</td>
<td>-4.06</td>
</tr>
<tr>
<td></td>
<td>(5-08-97)</td>
<td>(5-08-97)</td>
<td>(5-08-97)</td>
</tr>
</tbody>
</table>

Augmented Dickey-Fuller (ADF) tests were performed on logarithmic values (levels) and their first differences (returns). A, B and C denote the three different model types presented in Zivot and Andrews (1992). Break dates are in brackets. The number of lags for the tests were determined by following a sequential, downward t-test on all lags.

\(^a\) Indicates significance at the 1% level. \(^b\) indicates significance at the 5% level and \(^c\) at the 10% level.

Table 3. Residual Diagnostic Tests.

<table>
<thead>
<tr>
<th>Country</th>
<th>LM autocorrelation test</th>
<th>LM ARCH test</th>
</tr>
</thead>
<tbody>
<tr>
<td>Japan</td>
<td>1.4084</td>
<td>0.5569</td>
</tr>
<tr>
<td>Malaysia</td>
<td>0.9649</td>
<td>1.5660</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>0.8232</td>
<td>3.7052(^a)</td>
</tr>
<tr>
<td>Singapore</td>
<td>0.9836</td>
<td>2.1277(^a)</td>
</tr>
</tbody>
</table>

Residual tests are for 12 lags and are based on a VAR with 26 lags. Both tests are \(\chi^2\) distributed.
Table 4. Johansen Rank Tests.

<table>
<thead>
<tr>
<th>$H_0$: $r$</th>
<th>$\text{Full Sample (5/1/93 – 11/12/01)}$</th>
<th>$\text{Pre- Crisis Sample (5/1/93 – 29/7/97)}$</th>
<th>$\text{Post- Crisis Sample (21/10/97 – 11/12/01)}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$r=0$</td>
<td>$\lambda_{\text{Trace}}$ $91.49^a$ $\lambda_{\text{Max}}$ $61.738^a$</td>
<td>$\lambda_{\text{Trace}}$ $123.321^a$ $\lambda_{\text{Max}}$ $69.302^a$</td>
<td>$\lambda_{\text{Trace}}$ $94.186^a$ $\lambda_{\text{Max}}$ $63.358^a$</td>
</tr>
<tr>
<td>$r\leq1$</td>
<td>$\lambda_{\text{Trace}}$ $50.619^b$ $\lambda_{\text{Max}}$ $36.046^b$</td>
<td>$\lambda_{\text{Trace}}$ $62.302^b$ $\lambda_{\text{Max}}$ $40.577^b$</td>
<td>$\lambda_{\text{Trace}}$ $51.832^a$ $\lambda_{\text{Max}}$ $36.788^b$</td>
</tr>
<tr>
<td>$r\leq2$</td>
<td>$\lambda_{\text{Trace}}$ $17.939$ $\lambda_{\text{Max}}$ $14.868$</td>
<td>$\lambda_{\text{Trace}}$ $34.153$ $\lambda_{\text{Max}}$ $24.171$</td>
<td>$\lambda_{\text{Trace}}$ $17.711$ $\lambda_{\text{Max}}$ $14.595$</td>
</tr>
<tr>
<td>$r\leq3$</td>
<td>$\lambda_{\text{Trace}}$ $6.291$ $\lambda_{\text{Max}}$ $6.291$</td>
<td>$\lambda_{\text{Trace}}$ $14.021$ $\lambda_{\text{Max}}$ $14.024$</td>
<td>$\lambda_{\text{Trace}}$ $5.889$ $\lambda_{\text{Max}}$ $5.888$</td>
</tr>
</tbody>
</table>

The results presented are the Johansen trace and maximum eigenvalue statistics. Critical values are taken from Osterwald-Lenum (1992). $^a$ Indicates rejection of the null at the 1% significance level.

Table 5. Inoue Rank Tests.

<table>
<thead>
<tr>
<th>$H_0$: $r$</th>
<th>$\text{Model A}$</th>
<th>$\text{Model B}$</th>
<th>$\text{Model C}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$r=0$</td>
<td>$\lambda_{\text{Trace}}$ $95.491^a$ $\lambda_{\text{Max}}$ $61.738^a$</td>
<td>$\lambda_{\text{Trace}}$ $123.321^a$ $\lambda_{\text{Max}}$ $69.302^a$</td>
<td>$\lambda_{\text{Trace}}$ $94.186^a$ $\lambda_{\text{Max}}$ $63.358^a$</td>
</tr>
<tr>
<td>$r\leq1$</td>
<td>$\lambda_{\text{Trace}}$ $50.619^b$ $\lambda_{\text{Max}}$ $36.046^b$</td>
<td>$\lambda_{\text{Trace}}$ $62.302$ $\lambda_{\text{Max}}$ $40.577^b$</td>
<td>$\lambda_{\text{Trace}}$ $51.832^a$ $\lambda_{\text{Max}}$ $36.788^b$</td>
</tr>
<tr>
<td>$r\leq2$</td>
<td>$\lambda_{\text{Trace}}$ $17.939$ $\lambda_{\text{Max}}$ $14.868$</td>
<td>$\lambda_{\text{Trace}}$ $34.153$ $\lambda_{\text{Max}}$ $24.171$</td>
<td>$\lambda_{\text{Trace}}$ $17.711$ $\lambda_{\text{Max}}$ $14.595$</td>
</tr>
<tr>
<td>$r\leq3$</td>
<td>$\lambda_{\text{Trace}}$ $6.291$ $\lambda_{\text{Max}}$ $6.291$</td>
<td>$\lambda_{\text{Trace}}$ $14.021$ $\lambda_{\text{Max}}$ $14.024$</td>
<td>$\lambda_{\text{Trace}}$ $5.889$ $\lambda_{\text{Max}}$ $5.888$</td>
</tr>
</tbody>
</table>

Critical values for the trace and maximum eigenvalue statistics are taken from Inoue (1999). The lag order was determined by sequential LR tests on the lags as followed by Inoue (1999). Breakpoint dates are presented in brackets under the last significant test statistic. $^a$ Indicates rejection of the null at the 1% level and $^b$ indicates rejection at the 5% level.

Table 6. Johansen Rank Tests for Break Model.

<table>
<thead>
<tr>
<th>$H_0$: $r$</th>
<th>$\text{Full Sample (5/1/93 – 11/12/01)}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$r=0$</td>
<td>$\lambda_{\text{Trace}}$ $91.42$ $\lambda_{\text{Max}}$ $68.952$</td>
</tr>
<tr>
<td>$r\leq1$</td>
<td>$\lambda_{\text{Trace}}$ $48.17$ $\lambda_{\text{Max}}$ $47.169$</td>
</tr>
<tr>
<td>$r\leq2$</td>
<td>$\lambda_{\text{Trace}}$ $19.55$ $\lambda_{\text{Max}}$ $29.072$</td>
</tr>
<tr>
<td>$r\leq3$</td>
<td>$\lambda_{\text{Trace}}$ $2.975$ $\lambda_{\text{Max}}$ $14.354$</td>
</tr>
</tbody>
</table>

The 95% critical values are derived from the estimated distribution for the model $H_c(r)$ presented in Johansen, Mosconi and Nielsen (2000).
Table 7. Exogeneity Tests

$\Delta X_t = \alpha_1 \text{ECM}_1 + \alpha_2 \text{ECM}_2 + \sum_{i=1}^{n} a_i \Delta \text{Japan}_{t-i} + \sum_{i=1}^{n} b_i \Delta \text{Malaysia}_{t-i} + \sum_{i=1}^{n} c_i \Delta \text{Hong Kong}_{t-i} + \sum_{i=1}^{n} d_i \Delta \text{Singapore}_{t-i} + \text{dum}_1 \text{ transdum}_t + e_t$

<table>
<thead>
<tr>
<th>Dependent Variable ((\Delta X_t))</th>
<th>(H_0: \alpha_1 = \alpha_2 = 0)</th>
<th>(\Sigma a_i)</th>
<th>(\Sigma b_i)</th>
<th>(\Sigma c_i)</th>
<th>(\Sigma d_i)</th>
<th>(\text{dum}_1)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Japan</td>
<td>24.2991(^a)</td>
<td>NA</td>
<td>43.0452(^b)</td>
<td>31.2425</td>
<td>40.0614(^c)</td>
<td>-0.0299(^b) (0.0120)</td>
</tr>
<tr>
<td>Malaysia</td>
<td>12.2445(^a)</td>
<td>61.7909(^a)</td>
<td>NA</td>
<td>37.0125</td>
<td>39.3623(^c)</td>
<td>-0.0111(^b) (0.022)</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>9.5538(^a)</td>
<td>40.2887(^c)</td>
<td>43.0512(^b)</td>
<td>NA</td>
<td>36.6753</td>
<td>-0.0194(^c) (0.0107)</td>
</tr>
<tr>
<td>Singapore</td>
<td>9.4517(^a)</td>
<td>29.4880</td>
<td>59.1953(^a)</td>
<td>36.0626</td>
<td>NA</td>
<td>0.0088(^a) (0.0429)</td>
</tr>
</tbody>
</table>

Joint tests on the alpha coefficients are \(\chi^2\) tests for weak long-run exogeneity, while the block exogeneity tests are F-distributed. Figures in brackets are standard errors. \(^a\) Indicates rejection of the null at the 1% level, \(^b\) at the 5% level and \(^c\) at the 10% significance level.

Table 8. Exclusion Tests.

ECM\(_1\): Japan – 0.010 Malaysia + 0.032 Hong Kong – 0.098 Singapore – 4.572 + 0.361 permdum
ECM\(_2\): -0.741 Japan – 1.503 Malaysia + 1.805 Hong Kong + Singapore – 3.611 – 0.511 permdum

<table>
<thead>
<tr>
<th>Country</th>
<th>(\chi^2(2)) statistic</th>
<th>(P)-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Japan</td>
<td>25.011</td>
<td>0.0000</td>
</tr>
<tr>
<td>Malaysia</td>
<td>4.029</td>
<td>0.1334</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>4.962</td>
<td>0.0837</td>
</tr>
<tr>
<td>Singapore</td>
<td>7.472</td>
<td>0.0239</td>
</tr>
<tr>
<td>permdum</td>
<td>23.657</td>
<td>0.0000</td>
</tr>
</tbody>
</table>

The exclusion tests were based on there being 2 cointegrating equations and are \(\chi^2\) distributed.
The authors wish to thank Walter Enders and Gregory Schwann for several helpful comments on earlier drafts.

1 The role of the real estate market upon the Asian crisis has been documented frequently. See Renaud (2000) for an overview.

2 But Ziobrowski and Curcio also found that diversification benefits were dissipated through exchange rate risk.

3 Monthly data from Global Property Research and Datastream International were also examined to provide a comparison using different property indices. The results from using the alternate series were relatively the same, although the low frequency of observations and small sample size did impact upon some of the cointegration analysis in relation to the weekly data.

4 If the sample excludes the crisis period excess kurtosis is 0.7324, 5.4453, 2.2185 and 3.2553 for Japan, Malaysia, Hong Kong and Singapore, respectively.

5 Choosing higher lag orders to eliminate all ARCH effects (41 lags) within the residuals did not lead to any significant changes to the outcomes presented in the paper and was at a too high a cost of degrees of freedom.

6 However, not allowing for a linear deterministic trend does not preclude deterministic components within the cointegrating relations.

7 Choosing the other breakpoints identified makes little difference to the results as they are no more than two weeks apart.

8 The two cointegrating equations were normalized on Japan and Singapore, based on having the largest significant negative alpha coefficients within each of the cointegrating equations.

9 cf. World Bank World Tables.