

US Equity Market Spill-Over and Contagion Effects On Selected Asian Markets Vis-à-vis September 11

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Abstract

Considering the global dominance of the US equity market, it is expected that the impact of September 11 on the US market would spill-over to other markets. Since this terrible event had created a global climate of fear and uncertainty, it is possible that the US spill-over effect could have been driven not only by fundamental factors but also by non-fundamental ones. Thus, contagion could have played a significant role in transmitting the effect of September 11 from the US to other markets. In this paper, we verify whether indeed this has happened. Based on leveraged bootstrap causality tests, we examine the causal effects of the US on Indonesia, China, Japan, Taiwan and Singapore in relation to September 11. We found that the causal effects of the US on Japan and Singapore intensified as a result of September 11; however, we did not find these to be associated with contagion effects. It appears that the September 11 impact on the US market was not transmitted to the other Asian markets as the US influence on these markets either remained the same or even diminished.

JEL Classification: G15; C32

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1. Introduction

Considering the global dominance of the US stock market, it is expected that the effect of such a horrible event as September 11 would spill over to other markets. Since this tragic event had created a global climate of fear and uncertainty, it is possible that these spill over effects could be driven not only by fundamental factors but also by non-fundamental ones and thus contagion could play a significant role in this transmission process. Since developed markets are more efficient in processing information and have fewer problems with information asymmetry, it is expected that they will be less susceptible to spill over effects and contagion. However, this possibility is negated by the fact they are also more open and integrated with the world markets. In the case of less developed markets, it is argued that they are more prone to spill-over effects and contagion as they are exposed more to information asymmetry effects and herding behaviour. But at the same time, they are also less open and less efficient in responding to new information and thus they may be less affected by foreign spill-over and contagion. Thus, it is unclear how the September 11 impact on the US market would have been transmitted to other markets. Unfortunately, at present there is close to nothing in terms of empirical research on this issue.

We therefore address this gap in the literature. We empirically investigate the extent of the US equity market spill-over to Asian markets arising from the September 11 shock and whether such spill-over also exhibit contagion characteristics. We examine the causal effects, vis-à-vis September 11, of the US equity market on the stock markets of Indonesia, China, Taiwan, Singapore and Japan based on bootstrap causality test with leveraged adjustments in accordance with the procedures

developed by Hacker and Hatemi-J (2005a). This method has been shown by Hacker and Hatemi-J (2005a) to be robust to non-normality and ARCH effects in the data.

The issue of financial market spill-over as a result of a shock such as September 11 and the role of contagion in this spill-over is one that is highly important to investors as well as policymakers. Spill-over and contagion effects impact on the linkages between markets and therefore have implications on the extent of diversification benefits that are available for investors. As a result of September 11, there will be more impetus for investors to diversify their portfolios across more markets in order to protect themselves against a shock like this that occurs in a particular market (Eichengreen, 2002). Thus, knowledge about spill-overs and contagion arising out of September 11 would be highly important to investors in their quest for international diversification.

As argued by Baig and Goldfajn (2000), knowing whether a crisis is due to a spill-over effect that is mainly coming from contagion or not will greatly assist in formulating the appropriate policies that are needed to respond to a crisis situation. If the crisis is mainly due to contagion, that is, due to temporary, non-fundamental factors, then short-run isolation strategies – such as capital controls – could be highly effective in reducing the effect of the crisis. On the other hand, if the crisis is transmitted mainly through permanent factors, i.e. fundamental factors, that are prevalent in all states of the world, then these short-run isolation strategies will only delay a country's adjustment to a shock (Forbes and Rigobon, 2001b; Goldstein, 1998; Lane and Phillips, 2001 and Dungey et al, 2003). In fact, inappropriate policies may result in exacerbating contagion in the sense that these policies create “moral hazard”

which has been shown to also act as channels of contagion effects (Wälti, 2003).

The investigation of contagion also has important theoretical implications. Studies of contagion in crises periods provide new evidence in relation to the efficient market hypothesis. As Blose et. al (1996) emphasized, catastrophic events can place extreme stress on the stock market, and under such circumstances, both experienced as well as unsophisticated analysts can be misled by breaking news items that may be inaccurate (or at least exaggerated), incomplete, or biased. Contagion effect caused by panic in which investors decide to liquidate their current holdings in the absence of analysis based on high quality information is evidence of market inefficiency.

In this study, we use a new approach to the study of financial market spill-over and contagion. We perform our investigation using causality tests, which we bootstrap with leveraged adjustments following the procedures developed by Hacker and Hatemi-J (2005a). This approach provides two important advantages over previous approaches used in the literature. First, testing financial market interaction and contagion within a causality context enables us to capture the dynamic nature of the spill-over and contagion since such effects are viewed as transmissions from an originator country to the recipient countries, which accords with the definition of causality. Secondly, the use of leveraged bootstrapping in performing the causality test provides results that are robust to non-normalities and ARCH effects in the data. The probability of extreme events in the financial markets is much higher than what normal distribution would suggest (Hatemi-J, 2002).

One of the most highly used methods in analysing financial market linkages and contagion is the use of correlation. This approach, however, suffers from problems arising from heteroscedasticity, (as documented in Boyer, Gibson and Loretan, 1999; Loretan and English, 2000; and Forbes and Rigobon, 2001a), endogeneity and omitted variables (as discussed in Forbes and Rigobon 2001a). A number of innovative papers proposing new methods of calculating correlation have been published in an attempt to overcome these problems; however, no approach has been able to handle all the three problems simultaneously. Forbes and Rigobon (2001a), for example, proposed a new method for estimating correlations that adjust for heteroscedasticity based on the assumption that the other two problems are controlled.

In our study, we make use of a methodology that addresses these three problems simultaneously in order to detect whether a spill-over arising from a financial turmoil in one country is characterised by contagion. As stated earlier, we employ the Hacker and Hatemi-J (2005a) causality test, which is based on bootstrap simulations with leveraged adjustments.

There is a voluminous study on financial market movements' spill-over. However, very few of these studies focus on the role of contagion in these spill-overs. With reference to the September 11, 2002 disaster, there is close to nothing in terms of research on contagion vis-à-vis spill-overs. As far as we know, there is only one study, that of Hon et. al. (2004) that examined contagion in selected markets after the September 11 attack. In this study, they utilised a traditional approach by evaluating whether the correlation coefficient amongst international stock returns has changed.

This study is therefore not able to address the three problems stated earlier. Furthermore, the said study does not include China in their analysis.

The remaining parts of this paper are organised as follows. Section two presents market characteristics and data properties. Section three describes the bias–correction procedure proposed by Hacker and Hatemi-J (2005a) and adopted in this paper. Section four presents the empirical results relative to the existence of contagion and explores the pattern of propagation. Section five embodies the conclusion and suggestions for further research.

2. Market characteristics and data properties

The countries in our sample provide a very heterogenous group in terms of capitalisation, liquidity, openness to international equity markets and experience in previous crises. Therefore, we expect these markets to react differently to the September 11 shock.

The Japanese market is the largest in terms of capitalisation, number of listed equities, and market turnover. After Japan, the Singaporean market is one of the leading regional markets in East Asia. The Indonesian, Taiwanese and Chinese markets tend to be at the lower end of the spectrum.

[INSERT TABLE 1]

In relation to contagion, developing markets have been found to usually exhibit more

contagion than developed markets (Calvo, 1999). Developed markets are often modeled as conduits of contagion between developing markets (Kodres and Pritsker, 2001; Kaminsky and Reinhart, 2002). Lower information asymmetries in developed markets may explain why such developed markets remain unscathed by contagion while emerging markets are usually hit hard. That suggests that in this case, contagion will most likely emerge in underdeveloped financial markets such as Indonesia, China and Taiwan.

Kaminsky and Reinhart (2000) stressed the possibility of cross-hedging as a channel for the transmission of financial tremor. When there is a negative shock to or within one country, investors may be tempted to sell that country's assets and buy assets of another country, thereby increasing their exposures to the idiosyncratic factor of the second country. In order to hedge this new position, investors sell the assets of a third country, and this process may form a channel for transmitting a financial crisis. However, this channel works in a circumstance where market liquidity is sufficient. The average amount of transaction and the number of shares traded provide indications of the liquidity in each market. Data for each market in relation to these items are presented in Table 2. These data show that Japan had the highest average amount of transaction, followed by Taiwan. The ratios for Singapore and Indonesia are much lower than those of Japan and Taiwan.

Sarr and Lybek (2002) also tested the relation between liquidity and financial crisis. They reported that the uncertainty about the equilibrium price can be an outcome of illiquidity in financial markets. The doubts about the equilibrium price may shift a country's equilibrium to a "bad" one with or without fundamentals analysis when a

financial crisis happens in other countries, so that in this situation, more interaction, even contagion can occur (Masson, 1998). Illiquidity can be indicated by small trading volumes and low turnover in terms of trading. This is because large numbers of trades constitutes a valuable information source to investors, and allows investors to adjust investment appropriately based on it. Trading volume and turnover ratio for each country in question is listed in Table 1. China, Taiwan and Japan had high turnover ratios in 2001. Taiwan's turnover ratio is the highest among the five Asian countries, followed by China and Japan with the ratio in China being only half of that in Japan.

It is therefore our expectation that Singapore and Indonesia with relatively low liquidity in 2001 would be more affected by external influence and even contagion in comparison to Japan, China and Taiwan, which had high liquidity.

[INSERT TABLE 2]

The efficient market hypothesis implies that markets react instantaneously and properly to unanticipated information relative to economic fundamentals. Contagion is an evidentiary factor of market inefficiency as it emerges in a situation where investors overreact to unanticipated external information without high quality fundamental analysis. Smiles (2003) argued in his paper that some underdeveloped Asian countries possibly do not react to unanticipated external news. This is not necessarily an evidence of market inefficiency, but could possibly be an outcome of "information inefficiency". China, Taiwan and Indonesia have high foreign investment restrictions. As can be seen in Table 3, these markets have relatively few

foreign companies. The lack of foreign participants may lead those markets to view that external information is irrelevant or may not realise the implications of such information. Thus, we expect that these markets will experience less spill-over effects from the US arising out of September 11 and would also be less susceptible to contagion.

[INSERT TABLE 3]

Empirical evidence regarding the effect of contagion upon recent financial crises varies in Asian equity markets. Developed markets such as Japan and Singapore were affected by the Asian crises in 1997, as were underdeveloped markets of Taiwan and Indonesia (Baig and Goldfajn 1999, Park and Song 2000). China, however, seemed to have been unscathed even when it suffered from many (if not more) of the same structural ills that brought the other Asian countries down in 1997. Thus, we expect China not to be significantly affected by September 11 in terms of spill-over and contagion effects from the US market.

As discussed above, Asian markets are heterogeneous. The way in which different Asian markets would be impacted by the shock that originated in the U.S. market as a result of September 11 may differ greatly. Developing markets are expected to be hit harder than developed markets if there is any contagion effect from the US market to Asian markets due to the high liquidity and less asymmetry of those markets. On the other hand, lower foreign participation and associated liquidity might prevent these developing markets from being affected by external shocks. Thus, it is unclear how the propagation pattern of shocks from the US will proceed.

We utilise daily MSCI price index data, expressed in US dollars, covering the period from 12 March 2001 to 12 March 2002. We centre our event data on 11 September 2001 and designate a six months period before the event and six months after the incident as the two sub-periods of study.

Descriptive statistics in relation to the returns generated in each market are presented in Table 4, which show the existence of non-normality in the data. The null hypothesis of normality is rejected for each variable in each period with exceptions of Taiwan and Indonesia in period 2. Negative value of skewness in period 2 that characterises every market is consistent with what we stated earlier in this paper that most of the equity markets downsized after the September 11 attack.

[INSERT TABLE 4]

We further test for multivariate normality using Doornik and Hansen (1994) and ARCH effects by using Hacker and Hatemi-J (2005b) test. The results of these tests are presented in Table 5. The null hypothesis of no multivariate ARCH(1) as well as the null hypothesis of multivariate normality in the VAR model is strongly rejected for each sub-period. This justifies the usage of the leveraged bootstrap causality test in order to draw valid inference. This method is explained in a more detailed manner in the next section.

[INSERT TABLE 5]

3. Methodology

We postulate that in normal periods, if markets are efficient, the co-movement between markets would be guided by fundamentals. In a crisis period, if other factors beyond fundamentals start to also impact on market relationships, then the co-movement among markets would increase above that dictated by fundamentals. Because of “contagion” and other factors (e.g. “monsoonal” effect, spill-over, highlighted in Masson, 1998) there can be excessive co-movements between markets. Baig and Goldfajn (1999) argued further that any one of those driving factors can lead to an increased cross-markets linkage between markets in a crisis period. Therefore, a high degree of co-movement or spill-over during a crisis period is not a sufficient proof of contagion. It could be a continuation of historical cross-correlation between markets – that is, the co-movement is driven by factors other than contagion that is associated with fundamentals. The scenario is quite different if the correlations change substantially subsequent to the onset of the crises – in which case, one can indeed make the case for contagion. Forbes and Rigobon (2001c) use *t*-tests to ascertain if there is a significant increase in any of correlation coefficients during the turmoil period. They considered any statistic that is greater than the critical value at the 5% level of significance as an indication of contagion, while any statistic less than or equal to this value as an indication of the absence of contagion.

In this paper, we apply this evaluation procedure in a causality test scenario. In order to meet the first objective of the paper, we determine the extent of the causal effects of the US on the other markets during normal (before September 11) and crisis (after September 11) periods. In order to determine whether these causal effects contain contagion, which is the second objective of the study, we do the following. Since

contagion is represented by an excessive increase in spill-over, we determine the extent of increase in the causal effects of the US during the crisis period and test whether this is excessive. If the increase is significant at the 1% level, we deem this to be excessive and therefore the spill-over contains contagion effects.

In the ensuing paragraphs, we describe the causality test that is employed in this study. As previously mentioned, we make use of a causality test, which is based on bootstrap with leveraged adjustments as introduced by Hacker and Hatemi-J (2005a).¹ By causality, we mean causality in the Granger (1969) sense. That is, we aim to find out whether one variable precedes another variable or not. For this purpose, the following vector autoregressive model of order p , VAR(p), is utilised:

$$y_t = \nu + A_1 y_{t-1} + \dots + A_p y_{t-p} + \varepsilon_t, \quad (1)$$

Here:

y_t = the number of variables in the VAR model, which is six in this particular case,

ν = a six-dimensional vector of intercepts,

ε_t = a six-dimensional vector of error terms without the assumption of normality, and

A_r = a 6×6 matrix of parameters for lag r ($r = 1, \dots, p$).

It is important to choose an optimal lag order in this case because all inference in the VAR is naturally based on the chosen lag order. We use a new information criterion introduced by Hatemi-J (2003), which performs well. This information criterion, denoted by HJC, is presented below

$$HJC = \ln(\det \hat{\Omega}_j) + j \left(\frac{n^2 \ln T + 2n^2 \ln(\ln T)}{2T} \right), \quad j = 0, \dots, p, \quad (2)$$

¹ For another application when the variables are non-stationary see Hatemi-J (2004).

Here:

\ln = the natural logarithm,

$\det \widehat{\Omega}_j$ = the determinant of the estimated maximum-likelihood variance-covariance matrix of ε_t for lag order j ,

n = the number of variables in the VAR model, and

T = the sample size used to estimate the VAR model.

The optimal lag order (p) is chosen based on the minimisation of HCI . The null hypothesis of non-Granger causality from variable i on variable k in the vector y_t is then defined as the following:

$$H_0: \text{the row } i, \text{ column } k \text{ element in } A_r \text{ equals zero for } r = 1, \dots, p. \quad (3)$$

Assuming initial values are given, the estimated VAR(p) model can be written compactly as:

$$Y = DZ + \delta. \quad (4)$$

where

$$Y := (y_1, \dots, y_T) \quad (n \times T) \text{ matrix,}$$

$$D := (v, A_1, \dots, A_p) \quad (n \times (1 + n \times p)) \text{ matrix,}$$

$$Z_t := \begin{bmatrix} 1 \\ y_t \\ y_{t-1} \\ \vdots \\ y_{t-p-d+1} \end{bmatrix} \quad ((1 + n \times p) \times 1) \text{ matrix, for } t = 1, \dots, T,$$

$$Z := (Z_0, \dots, Z_{T-1}) \quad ((1 + n \times p) \times T) \text{ matrix, and}$$

$$\delta := (\varepsilon_1, \dots, \varepsilon_T) \quad (n \times T) \text{ matrix.}$$

The multivariate WALD statistic for testing the null hypothesis of non-Granger causality is defined as

$$WALD = (C\hat{\beta})' [C((Z'Z)^{-1} \otimes S_U)C']^{-1} (C\hat{\beta}) \sim \chi_p^2, \quad (5)$$

Here:

\otimes = element by all element matrix multiplication operator (the Kronecker product).

C = a $p \times n(1+n \times p)$ selector matrix, which indicates parameters that should take value zero under the null hypothesis.

S_U = the estimated variance-covariance matrix of residuals in equation (4) when the null hypothesis of no-Granger causality is not imposed (unrestricted model).

$\hat{\beta} = \text{vec}(\hat{D})$, where *vec* signifies the column-stacking operator.

This WALD test statistic has asymptotically a χ^2 distribution with degrees of freedom equal to the number of restrictions under the null hypothesis. However, Hacker and Hatemi-J (2005a) show through Monte Carlo experiments that the multivariate WALD test statistic overrejects the null hypothesis if the error terms in the VAR model are non-normal with autoregressive conditional heteroscedasticity (ARCH). To improve on the inference based on tests for causality in such situations, the authors suggest an alternative approach based on leveraged bootstrap simulations. The bootstrap method resamples the underlying data to estimate the distribution of a test statistic. Using this distribution can decrease bias in inference by providing more precise critical values.

The bootstrap simulations are performed in the following way. We first estimate equation (4). Then, we simulate the bootstrap data, Y^* , based on the estimated

coefficient matrix from this regression, \hat{D} ; the original Z , data; and δ^* (the bootstrapped residuals). That is the bootstrap data is obtained by the following equation:

$$Y^* = \hat{D}Z + \delta^* \quad (6)$$

It should be mentioned that the bootstrap residuals are drawn with replacement from the regression's modified residuals, each with equal probability of $1/T$. These residuals are mean adjusted to have the expected value that is equal to zero for each bootstrap sample. The residuals are also weighted by the leverages to make sure that the variance is constant.² Note that \hat{D} is estimated as the following:

$$\hat{D} = YZ'(ZZ')^{-1}. \quad (7)$$

The bootstrap simulation is repeated 10000 to calculate the WALD test statistic each time.³ These values give us the possibility to generate the empirical distribution for the multivariate WALD test. Then, we take the $(\alpha)th$ upper quantile of this distribution, which is the α -level “bootstrap critical values” (c_α^*). In this study we produce the bootstrap critical values for 1%, 5% and 10% significance levels. The next step is to calculate the WALD statistic using the original data (not the bootstrapped simulated data) and compare it to the bootstrap critical value at a given significance level. The null hypothesis of no Granger causality is rejected based on bootstrapping if the actual WALD is greater than the critical value. The simulations are performed by making use of a programme procedure written in GAUSS.⁴

² For more details on leverage adjustment, see Davison, and Hinkley (1999) and Hacker and Hatemi-J (2005a). The latter authors introduce this adjustment for multivariate equation cases.

³ However, the results are not sensitive to the number of simulations.

⁴ A program procedure written in GAUSS to conduct leveraged bootstrap simulations as introduced by Hacker and Hatemi-J (2005a) is available on request from the first author.

In order to determine whether the US spill-over is driven by contagion, we add a dummy variable that is equal to zero for the low volatility period and one for the high volatility period. The addition of the dummy variable into the VAR model is meant to capture the increase in causal links and therefore allows to determine the existence of contagion. We test the hypothesis that the coefficient of the dummy variable is equal to zero. If the null hypothesis is rejected at the 1% level of significance, then the increase in causality is deemed to contain contagion effects.

Before conducting the causality test, it is well accepted in the literature that it is important to check the time series properties of the underlying data in order to avoid false and spurious inference. It is also well known that standard tests for unit roots have low power if structural breaks have occurred during the period of study. In order to take into account the effect of a potential structural break arising from the September 11 terrorist attack, we make use of a test suggested by Perron (1989) to test for the integration order of the variables. This test permits for a structural break in both the mean value and the deterministic trend of the variable under investigation for unit root. The Perron (1989) test for unit roots of variable z is based on the following regression:

$$z_t = c_1 + c_2 D_t + d_1 t + d_2 D_t t + g J_t + \gamma z_{t-1} + \sum_{i=1}^m b_i \Delta z_{t-i} + u_t \quad (8)$$

where

t = the time period (the linear trend term),

D_t = a (binary) dummy variable that takes value zero for the time period before break and one for the rest of the period,

J_t = a binary variable that is equal to one if the time period t is the first period after that of the structural break, and is zero otherwise,

Δ = the first difference operator and

u_t = a white noise error term.

The null hypothesis of a unit root is $\gamma = 1$, and the alternative hypothesis of stationarity is $\gamma < 1$. If necessary, we will include lag values of Δz in equation (1) to make sure that the error term is white noise. The optimal number of lagged differences (m) is determined by including more lags until the null hypothesis of independence for u_t is not rejected by the LM test at the 5% significance level.⁵

4. Empirical Results

The Perron (1989) test was conducted to test for the stationarity of the series for all six countries. For the null hypotheses of I(1), i.e. integration of the first order, the estimated test values in absolute terms are found to be greater than the critical values at the conventional significance levels. Hence, the null hypothesis that each variable is I(1) is rejected and therefore the variables are stationary.

[INSERT TABLE 6]

The results of the bootstrapped causality tests are presented in Table 7. During the period before September 11, it can be seen from this Table that, with the exception of Indonesia, the US significantly influenced the Asian markets. The US was most influential on China, Singapore and Taiwan – Granger causing these markets at the 1% level of significance. It was less influential on Japan – Granger causing Japan at the 5% level of significance. After September 11, the US ceased to be influential on China and became less influential on Taiwan (the US Granger caused this market at

⁵ It is shown by Hatemi-J (2004b) that the LM test has better size properties compared to alternative tests for autocorrelation.

only the 5% level as compared to 1% before). However, it became more influential on Japan and Singapore as shown by the bigger causal parameters associated with these markets, which were all significant at the 1% level. Thus, September 11 resulted in increased US spill-over effects on Japan and Singapore but had the opposite effects on China, Indonesia and Taiwan.

[INSERT TABLE 7]

The assessment issued by the World Bank in “Asian Development Outlook 2002” supports the above picture to a large degree. Increased causal links from the US to Japan and Singapore could be due to the fact that Singapore and Japan were the most international markets during this period. In addition, Japan was experiencing a recession during that period, with the loss of market confidence in growth prospects translating into significant additional wealth destruction. This resulted in a sharp fall in investment as equity prices dived, with investors becoming more likely to panic as a consequence of external shocks. The deterioration of the economic condition worldwide and a severe downward in electronic demand had dragged down the Singapore economy and its financial market since Singapore is heavily dependent on foreign demand, especially for electronic products.

On the other hand, the result in which China showed decreasing interdependence with the US and therefore was not affected significantly by the financial turmoil caused by the attack is not surprising in the context of the assessment that:

“In spite of the global economic slowdown, the People’s Republic of China maintained its robust economic expansion in 2001, though some signs of slowing appeared in the second half of the year, due partly to a sharp deceleration in growth of external trade...” (Asian Development Outlook 2002).

Indonesia and Taiwan were not influenced so much by the September 11 shock as revealed by the causality test, despite the fact that their economy growth also decelerated rapidly in 2001 in the same manner as Japan and Singapore. This result could be therefore due to factors other than economic conditions, such as market regulations, information transmission, etc. which were not discussed in this study.

In order to determine whether the increased US causal effects on Japan and Singapore contain contagion characteristics, we tested whether the size of the increase in the causal parameters were significant at the 1% level using the *t*-test. The results, presented in Table 8, show that the increases in the size of the parameters were only significant at the 10% level. Hence, the increase in spill-over from the US market to Japan and Singapore did not carry contagion effects and were mainly due to fundamental factors.

[INSERT TABLE 8]

The observation of the absence of contagion with respect to the September 11 spill-over effect is inconsistent with the findings of Hon et. al. (2004). In that study, it was found that Japan experienced significant increases in correlation with the US market, and the correlation remained significant six months after the attack which the authors cite as evidence of contagion. The observations generated by our study are also inconsistent with the results of studies on the spill-over effects of the Asian crisis. In those studies, Asian markets were found to be generally vulnerable to external shocks (see Corsetti, Pericoli and Sbracia, 2000; Baig and Goldfajn, 1999; Kaminsky and Schmukler, 2002; Cerra and Saxena, 2000; among others).

The divergence between our study and that of Hon et. al. (2004) could be due to the different methods and data used. Hon et. al. (2004) utilised the conventional correlation coefficient analysis, whereas our study applies causality tests based on leveraged bootstrap procedures. No study to date has used this method to test contagion between equity markets. The discrepancy of the results between our study and those studies focusing on the Asian crises episodes could also be a consequence of the utilisation of different methods. Corsetti, Pericoio and Sbracia (2000) and Baig and Goldfajn(1999) relied on standard correlation coefficient analysis, Kaminsky and Schmuler (2002) adopted regression and event studies, Cerra and Saxena (2000) used the time-varying transition probability model and Markov-Switching model. The fact that Rigobon (2000) used a purified correlation coefficient test and did not observe any contagion could be evidence that the finding of contagion can be attributed to the limitations of statistical models. In addition to the reasons discussed above, the divergence of our results to those of others also could have emerged from the different markets and time periods used.

Our findings, however, draw support from the report of the World Bank (Regional Economic Developments and Prospective, 2000) which stated that the external position (demonstrated by high international reserves, low short-term debt and positive current accounts, etc.) of all countries was significantly better than in 1997, with the capacity to absorb external shocks increasing ever since, prevent it from reach high co-movement level, contagion.

5. Conclusions

We investigate whether September 11 significantly changed the links between the US equity market and those of China, Taiwan, Indonesia, Singapore and Japan using leveraged bootstrap causality tests developed by Hacker and Hatemi-J (2005a). We also investigated whether September 11 spill-over effects from the US to the latter markets exhibit contagion. Our results reveal that September 11 significantly changed the interaction of the US market with the Asian markets covered in this study. The US became more highly linked with Japan and Singapore but less with Taiwan, ceased to be at all with China and continued not to be so with Indonesia. Thus, September 11 diverted the US influence to the developed markets of Japan and Singapore away from the less developed markets of China and Taiwan. We did not find, however, any contagion effects in the increased interaction of the US with Japan and Singapore. Hence, in spite of the climate of uncertainty created by September 11, the US market continued to deal with other markets based on fundamental factors.

Future research into the issue of contagion and interdependence has scope in at least three areas. First, the findings in this study are limited to providing evidence of existence/absence of contagion and spill-over. The spill-over was broadly categorised into either due to contagion or fundamental factors. The findings, however, offer no explanation for which of those possible transmission channels of contagion or spill-over was effective in the wake of equity markets collapse after the September 11 attack. A future study can develop a new model to quantify transmission channels and the results produced from such a study will be of assistance to policymakers in developing appropriate strategies to contain negative external spill-overs.

Second, considering that only a relatively small sample of countries was used in this study, there is the risk of omission of some important financial links amongst equity markets. Future studies may attempt to cast the net as wide as possible using a sample that will cover more equity markets.

Third, it is possible that our results may have been distorted by other financial and policy events which occurred in the examine window. A future study may look for a reduction of these anomalies and purify the results. Two methods can be used to achieve this. The first is to clear the examined window and control the effect from other events. The second is to develop a method that traces the movement of asset prices and cross-market interactions continuously (say, on a day-to-day basis). Price changes and cross-market interaction measured in this way can be easily related to the effect of every event.

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Tables

Table 1. Some Important Features of the Equity Markets of Selected Asian Markets

	Market Capitalisation		Value Traded as a	Turnover Ratio as a	Number of Listed
	In millions of dollars	Percentage of GDP	Percentage of GDP	Percentage of GDP	Domestic Companies
Japan	4,546,937	104.6	42.5	52.5	2470
Singapore	198,407	236.8	115.4	49.4	355
China	330,703	33.2	38.1	114	950
Indonesia	26,834	45	14	32.9	290
Taiwan	375,991	130.6	316.1	242	462

Source: World Development Indicators (2001) and Emerging Markets Factbook (2000)

Table 2. Trading Statistics of the Equity Markets of Selected Asian Markets

	Number of Transaction	Average Amount of Transaction	Number of Shares Traded
Japan	NA	6750.1	204194
Singapore	NA	284.4	85138.7
China	NA	NA	NA
Indonesia	3621.6	38.7	148381.3
Taiwan	144280	2231.9	606420

Source: International Federation of Stock Exchange, 2001

Table 3. Foreign Companies with Shares Listed on Major Securities Exchanges

	Number of Foreign Companies	Foreign Companies as a % of all companies	Share traded of Foreign companies (US\$B)	Foreign Companies Trading as a % of all trading value
Japan	41	2	1	0.3
Singapore	63	13.1	0	0
China	0	0	0	0
Indonesia	0	0	0	0
Taiwan	0	0	0	0

Source: International Federation of Stock Exchange, 2000

Table 4. Descriptive Statistics Associated with Returns in Each Market

Market	Mean	S.D.	Min	Max	Skewness	Kurtosis	JB
<i>Sub-period 1 (12 Mar, 2001-11 Sep, 2001)</i>							
China	-12.53482	107.2382	-324.8222	278.8610	0.000894	3.611449	2.040723***
Indonesia	0.087687	1.120811	-3.420764	4.611866	0.407467	6.060037	54.73587***
Japan	-0.037559	0.715883	-2.106222	2.494614	0.375237	3.937147	7.867940***
Singapore	-6.910116	56.61427	-238.3766	140.1325	-0.292748	4.576393	15.43518***
Taiwan	-0.102534	0.745426	-1.697202	1.922544	0.259938	2.692032	1.992916***
US	-0.023871	0.569705	-1.573560	1.897505	0.308889	3.652766	4.408984***
<i>Sub-period 2 (12 Sep 2001- 12 Mar, 2002)</i>							
China	8.386382	154.7698	-431.1492	283.3425	-0.578482	3.728927	2.415290***
Indonesia	-0.421235	1.249715	-3.269206	2.313089	-0.229657	3.160931	0.305956
Japan	-0.039814	1.003178	-3.108761	1.737715	-0.986061	4.325131	7.291768***
Singapore	-19.43890	109.7463	-298.6114	165.3724	-0.644900	3.166693	2.184687***
Taiwan	-0.068919	1.185662	-2.589130	2.140111	-0.197806	2.524300	0.494449
US	-0.009705	0.741436	-2.268921	1.725290	-0.571456	4.850327	6.109531***

(a) JB denotes Jarque-Bera test for normality.

(b) The notation *** means that the null hypothesis of normality is rejected at the 1% significance level.

Table 5. P-Values of the Tests for Multivariate ARCH and Multivariate Normality in the VAR model

Test	P-value
Period 1	
ARCH	0.007
Normality	0.000
Period 2	
ARCH	0.001
Normality	0.000

Notes: Doornik and Hansen (1994) test was used to test for multivariate normality. The null of no multivariate ARCH(1) effect was tested by using a test method suggested by Hacker and Hatemi-J (2004b).

Table 6. Test for Unit Roots Using the Perron Test

$H_0: I(1), H_1: I(0)$	TEST VALUE	$H_0: I(2), H_1: I(1)$	TEST VALUE
Japan	-17.17***	China	-16.76 ***
US	-15.70***	Indonesia	-14.47 ***
Singapore	-15.61***	Taiwan	-14.66 ***

Notes:

- (a) The critical value is -4.78 and -4.24 at the 1% and 5% significance level, respectively.
 (b) The notation *** implies significance at the 1% significance level.

Table 7. Results of Causality Test Based on Hacker-Hatemi-J Test

THE NULL HYPOTHESIS	THE ESTIMATED TEST VALUE (MWALD)	1% BOOTSTRAP CRITICAL VALUE	5% BOOTSTRAP CRITICAL VALUE	10% BOOTSTRAP CRITICAL VALUE	CAUSAL PARAMETERS
<i>First Sub-Period (Day ending on 12 Mar, 2001 to day ending on 11 Sep, 2001)</i>					
$SP_{US} \neq SP_{CHI}$	45.386***	5.373	3.284	2.324	0.967
$SP_{US} \neq SP_{INDO}$	0.290	6.598	3.947	2.694	0.001
$SP_{US} \neq SP_{JAP}$	5.243**	5.917	3.941	2.797	0.025
$SP_{US} \neq SP_{SING}$	13.664***	6.295	4.237	2.840	0.296
$SP_{US} \neq SP_{TAI}$	6.721***	6.242	3.804	2.746	0.028
<i>Second Sub-Period (Day ending on 12 Sep, 2001 to day ending on 12 Mar, 2002)</i>					
$SP_{US} \neq SP_{CHI}$	2.696	6.241	3.800	2.664	0.325
$SP_{US} \neq SP_{INDO}$	1.128	6.727	3.992	2.714	0.002
$SP_{US} \neq SP_{JAP}$	13.893***	6.929	3.815	2.670	0.053
$SP_{US} \neq SP_{SING}$	16.452***	6.724	3.564	2.385	0.546
$SP_{US} \neq SP_{TAI}$	4.971**	6.737	3.722	2.609	0.012

Notes: (a) The notation \neq implies non-Granger causality; (b) MWALD represents the modified Wald test statistic as described in Eq. (5); and (c) the lag order of the VAR model, p , was set to one for each sub-period. Also the augmentation lag, d , was to one since each variable contains one unit root, (d). The notations *, **, and *** imply significance at the 1%, 5% and 10% significance level, respectively.

Table 8. Change in US Causal Effects on Selected Asian Markets After September 11

Market	Change in Causal Effect
China	-0.642***
Japan	0.028*
Indonesia	0.010
Singapore	0.250*
Taiwan	0.008

Notes:

Null hypothesis is that the change in the causal parameter from period 1 to period 2 is equal to zero.

* indicates null hypothesis is rejected at the 10% significance level

** indicates null hypothesis is rejected at the 5% significance level

*** indicates null hypothesis is rejected at the 1% significance level.

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