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Do equity market correlations really change over time? The case of the US and Asia-Pacific markets

Existing literature suggests that conditional correlations between equity markets vary over time, and increase over periods of financial crises. The author tests this hypothesis on a set of eight national equity indices from the Asia-Pacific region on the one hand, and the US market on the other. Tse's (2000) constant conditional correlation test suggests that three out of eight market pairs exhibit constant conditional correlations. The remaining five correlations are time varying, but can be further subdivided into those that are characterized by high persistence in the dynamics, and those which display strong mean reversion graphs. Global asset managers should take these features into account when allocating funds across the Asia-Pacific as the expected diversification gains are likely to differ across the region, and depend on the characteristics of the correlation coefficients.

Keywords: stock market interdependencies, Asia-Pacific region, constant conditional correlation, dynamic conditional correlation, international diversification.

Introduction

Early studies on international equity markets such as Koch and Koch (1991), and Von Furstenberg and Jeon (1989) reveal growing interdependence across national share markets. It has since been argued that globalization, financial deregulation and contagion strengthen such dependencies and reduce diversification benefits (e.g., Longin and Solnik, 1995; Koutmos and Booth, 1995; Billio and Pelizzon, 2003). In particular, correlation coefficients are found to change over time and increase during bear markets, thus reducing the appeal of investing internationally.

The aim of this paper is twofold. First, we test the hypothesis of time varying conditional correlations between eight Asia-Pacific share markets and the US market with the aim of identifying which markets continue to exhibit stable correlation coefficients over the 2005-2011 time interval. This sample period covers the recent global financial crisis (GFC), and hence will allow us to provide some new evidence on the international diversification issue under the conditions of extreme market turbulence. Second, we estimate conditional correlations for the eight market pairs using two econometric models: (a) constant conditional correlation (CCC) model of Bollerslev (1990); (b) dynamic conditional correlation (DCC) model of Engle (2002). The choice between the two models is decided based on Tse (2000) constant conditional correlation test performed in step one of the study. In instances where the null of constancy is not rejected we estimate the CCC; where the null is rejected we proceed to estimate the dynamic conditional correlation model. In both cases, however, we account for the time varying nature of the volatilities using the GARCH model (Bollerslev, 1986). The set of Asia-Pacific markets used in this study includes: Japan, Hong Kong, Singapore, Korea, Taiwan, Thailand, Australia and New Zealand.

Interestingly, out of the eight market pairs (each of the Asia-Pacific markets is paired with the US market) there are three Asian markets that exhibit constant conditional correlations with the US. These markets are Singapore, Korea and Taiwan, and although they are highly correlated with the US, the correlation coefficients do not appear to be time varying. Since diversification gains depend both on correlations and volatilities of constituent portfolio securities, this finding implies lower risk of unexpected changes in the correlations between the three markets and the US. The markets of Australia and New Zealand are characterized by most persistence in the correlation dynamics with the US. Their correlations assume changing trends over prolonged periods of time and increase considerably during the period of the GFC, although the New Zealand-US correlation has decreased since. Lastly, the correlations between the markets of Japan, Hong Kong and Thailand and the US are time varying but with less persistence, as judged by the sum of the parameters on the lagged regressors in the DCC equation. The correlations between these three Asian markets and the US display strong mean reverting behavior illustrated in Figure 3, which would be regarded more favourably in a portfolio diversification context.

The rest of the paper is organized as follows. The econometric methodology is discussed in section 1, while section 2 describes the dataset and presents summary statistics for the data. Model estimates are presented in section 3 and the last section contains a conclusion.

1. Econometric specification

In order to formulate an econometric model for conditional correlations and test the hypothesis of their constancy we first remove any autocorrelation that may exist in the data. Specifically, we fit a vector autoregression (VAR) to weekly log returns of the nine national share markets as follows:

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$$r_t = c + \Pi_1 r_{t-1} + \ldots + \Pi_p r_{t-p} + u_t \,, \quad (1)$$

where $r_t = [r_{1t}, \ldots, r_{9t}]'$ is a vector of log-returns constructed from share market index levels P_{it} as r_{it} = 100 × log(P_{it} / P_{it-1}) for each country i. In the above model c is a (9 × 1) vector of constants, while Π_t are (9 × 9) autoregressive coefficient matrices. The correct lag length of one (i.e., p =1) is selected using the Schwarz (SBIC) and Hannan-Quinn (HQ) information criteria. The vector of innovations is assumed to be conditionally normally distributed $u_t | \Omega_{t-1} \sim N(0, H_t)$, where Ω_t is the information set available at time t.

Time varying nature of the conditional covariances (i.e., the elements of H_t) has been well documented. The advent of multivariate GARCH models (Bollerslev, Engle and Wooldridge, 1988) has enabled researchers to explicitly specify conditional covariance and correlation equations and thus study their estimates over time. For example, Longin and Solnik (1995) apply a multivariate GARCH model to seven major markets and conclude that covariances and correlation matrices change over time. They also find that correlations increase during periods of high volatility. In this paper, the benchmark model for the covariance matrix is:

$$H_t = D_t R D_t = \left[\rho \sqrt{h_{ii,t}} \sqrt{h_{jj,t}}\right]. \quad (2)$$

This is the constant conditional correlation model of Bollerslev (1990), which is characterized by constant conditional correlations so that conditional covariances are proportional to the product of the corresponding conditional standard deviations. D_t is specified as a diagonal matrix of time varying standard deviations of dimension (n × n), while R is defined as a symmetric matrix of conditional correlation coefficients which elements are pair-wise correlations ρ_{ij}. This specification provides a parsimonious model for the covariance matrix and significantly simplifies estimation. The diagonal elements of H_t, the conditional variances, which square roots are elements of D_t are specified in the form of Bollerslev (1986) GARCH (1, 1) models:

$$h_{ii,t} = \omega_i + \alpha_i u_{i,t-1}^2 + \beta_i h_{ii,t-1}. \quad (3)$$

Thus, the variance for each national market i = 1,...9 depends on a constant, a volatility shock $u_{i,t-1}^2$, and its lagged volatility $h_{ii,t-1}$.

1.1. Testing for constant conditional correlations. Given the above specification in equation (1) we wish to test for the constancy of the correlation coefficients ρ_{ij}. This test is implemented using the approach of Tse (2000), which formulates the following hypotheses:

$$H_0 : h_t^{US,i} = \rho^{US,i} \sqrt{h_{US,t}^2 h_{i,t}^2},$$
$$H_1 : h_t^{US,i} = \rho_t^{US,i} \sqrt{h_{US,t}^2 h_{i,t}^2} \quad (4)$$

for each i ∈ {Japan, Hong Kong, Singapore, Korea, Taiwan, Thailand, Australia and New Zealand}. The difference between the hypotheses being the subscript t on the correlation coefficient in the alternative hypothesis. Tse (2000) derives a Lagrange Multiplier (LM) version of the test which under the null hypothesis is distributed asymptotically as a χ^2 variable with (N (N − 1) / 2) degrees of freedom. If the null hypothesis in not rejected for a market pair we estimate the CCC model described above. On the other hand, in those instances where we reject the null of constancy in the conditional correlations we proceed to model correlation dynamics using the DCC model of Engle (2002).

1.2. DCC. The DCC model replaces the constant correlation matrix R with a time varying matrix R_t so that the covariance matrix may be re-written as:

$$H_t = D_t R_t D_t = \left[\rho_{ij,} \sqrt{h_{ii,t}} \sqrt{h_{jj,t}}\right]. \quad (5)$$

Next, a specification is formulated for R_t under the assumption that the standardized innovations ε_t obtained from $\varepsilon_t = D_t^{-1} u_t$ are conditionally normally distributed, i.e., $\varepsilon_t | \Omega_{t-1} \sim N(0, R_t)$. The last step is to specify the form for R_t:

$$R_t = diag\,\{Q_t\}^{-1} Q_t diag\,\{Q_t\}^{-1}, \quad (6)$$

where

$$Q_t = c + a\varepsilon_{t-1}\varepsilon'_{t-1} + bQ_{t-1}. \quad (7)$$

The above models are estimated using the quasi maximum likelihood approach in Ox 6.01 software.

2. Data summary and preliminary statistical analysis

We use weekly series for eight Asia-Pacific countries, plus the US market index, over the sample period of January 3, 2005-October 24, 2011. The indices included in the study are the Standard and Poor's 500 (the US), Nikkei Stock Average (Japan), Hang Seng (Hong Kong), Straits Times (Singapore), Kospi (Korea), Taiwan SE (Taiwan), Bangkok SET (Thailand), S&P/ASX 200 (Australia) and NZ50 (New Zealand). All data is obtained from DataStream® and is denominated in local currency. Weekly rather than daily frequency was chosen to circumvent the problems associated with non-synchronous data as described in Burns, Engle and Mezrich (1998) and Martens and Poon (2001)[1].

[1] Estimating second conditional moments on non-synchronous data leads to underestimation of conditional correlations/covariance and inability to distinguish between contemporaneous correlations and lagged spillover effects.

Figure 1 depicts the national share indices in levels. A number of interesting observations can be made regarding the movement of the markets. First, although there is significant variation across the individual markets, the nine series follow similar long-term trends.

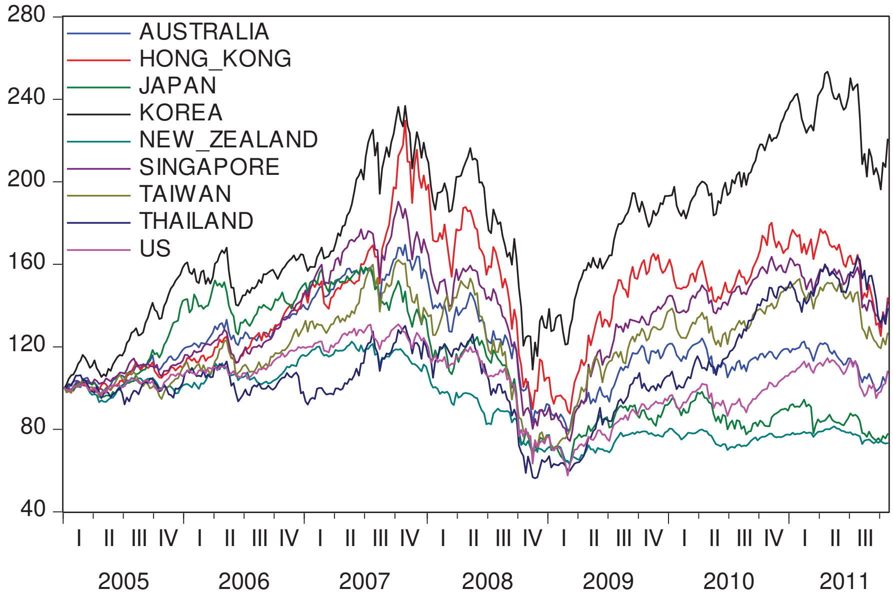

Fig. 1. National stock market indices (rebased to 100 in January 2001)

In particular, there is an upward trend across all nine markets over the period of 2005-mid 2007, when the markets peak with the onset of the global financial crisis. The beginning of the GFC is frequently dated to the summer of 2007, see, e.g., Reinhart and Rogoff (2008), Krugman (2009) and Swagel (2009), and this is clearly evident in the above graph. Following the start of the second quarter of 2009 most of the nine markets resume their upward trends. Next we present log returns for the above indices in Figure 2.

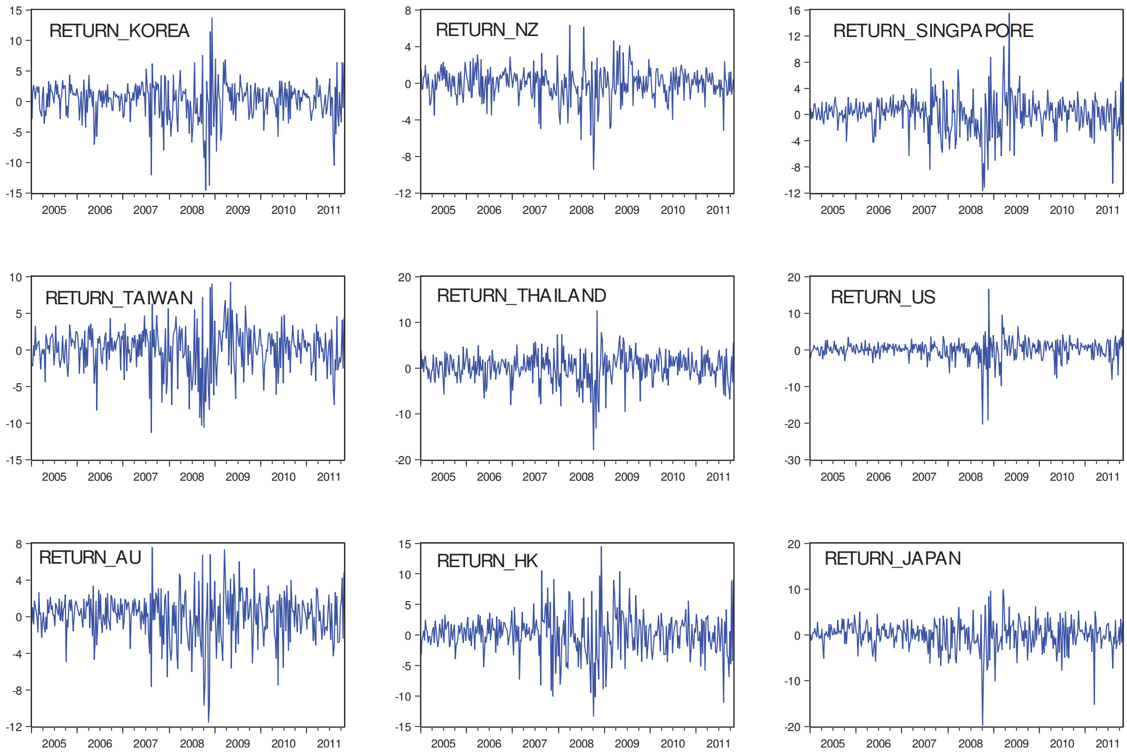

Fig. 2. Log returns for national stock market indices

As illustrated above the return series are consistent with the frequently cited stylized facts about financial time series. In particular, we can observe the existence of large outliers, synchronicity of extreme observations and volatility clustering in the nine return series. Further, the GFC period is clearly characterized by a significant increase in volatility and the frequency of extreme observations. In order to gain further information about the return series we provide some summary statistics in Table 1.

The highest return, over the January 2005-October 2011 period, was recorded in Korea of 11.15% p.a., followed by Thailand with 5.21% p.a and 3.14% p.a. for Taiwan. The US market returned only 0.83% p.a. while the lowest returns were in New Zealand and Japan with -4.30% p.a. and -3.69% p.a., respectively. Over the same time period Thailand had the highest levels of risk with a standard deviation of 23.71% p.a., followed by Korea and Japan. New Zealand had the least volatile market. All of the markets exhibit some negative skewness and excess kurtosis. The Jarque-Bera tests reject the null hypothesis of normality in all nine instances.

Table 1. Descriptive statistics

	US	Japan	Hong Kong	Singapore	Korea	Taiwan	Thailand	Australia	New Zealand
Mean (% p.a.)	0,83	-3.69	4.81	5.21	11.15	3.14	5.30	1.02	-4.30
St. dev. (% p.a.)	20.90	22.90	25.09	20.55	22.73	21.99	23.71	18.47	12.91
Skewness	-1.58	-1.02	-0.10	-0.16	-0.73	-0.61	-0.79	-0.72	-0.70
Kurtosis	17.46	8.31	5.10	7.29	6.83	4.46	6.28	5.24	6.30
Jarque-Bera	3251.23	480.43	65.93	275.01	249.39	53.35	196.23	104.79	190.72
JB p-value	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

2.1. Unit root tests. The index series appear non-stationary in levels but stationary in returns, as illustrated in Figures 1 and 2, respectively. In order to formally test these hypotheses we perform the Augmented Dickey-Fuller (ADF, 1979) and Phillips-Perron (PP, 1988) tests.

Table 2. Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests

	Levels		Log differences	
	ADF	PP	ADF	PP
US	-1.58	-1.56	-12.22**	-21.79**
Japan	-0.84	-0.94	-18.25**	-18.25**
Hong Kong	-1.93	-2.00	-18.64**	-18.66**
Singapore	-1.71	-1.89	-17.60**	-17.93**
Korea	-1.82	-1.81	-11.79**	-19.36**
Taiwan	-1.77	-1.99	-8.80**	-20.02**
Thailand	-1.12	-1.26	-19.64**	-19.86**
Australia	-1.59	-1.62	-18.92*	-18.92**
New Zealand	-0.63	-0.79	-17.34**	-17.44**

Note: Log difference are approximately percentage returns. The critical values for the two tests are -3.44 at the 1% and -2.87 at the 5% significance. ** denotes rejection of the null hypothesis at the 1% significance.

Both of these procedures test the null hypothesis of unit root, but differ in the way they adjust for autocorrelation in the residuals. The Phillips-Perron test accounts for autocorrelation non-parametrically, while the ADF includes lagged regressors in the test equation.

As illustrated in Table 2 we are unable to reject the null hypothesis of unit-root for any of the series in levels, i.e., prices. On the other hand, the same null hypothesis is rejected for all return series at the 1 percent significance level according to both ADF and PP tests. These findings give us confidence that the return series are stationary, i.e., $I(0)$, and hence can be adequately modelled using the methods outlined in section 1.

3. Empirical findings

In the first step of the estimation procedure we need to decide whether to use the CCC or the DCC model. This decision is based on Tse (2000) test for constant conditional correlation described in section 1.1. Below we present the results of the test.

Table 3. Tests for constant conditional correlations (p-values)

	Constant conditional correlation test Tse (2000)
Japan-US	0.00
Hong Kong-US	0.04
Singapore-US	0.19
Korea-US	0.31
Taiwan-US	0.33
Thailand-US	0.04
Australia-US	0.01
New Zealand-US	0.01

There are three instances where the test is unable to reject the null hypothesis of constant conditional correlation. In particular, the market pairs Korea-US, Taiwan-US and Singapore-US appear to characterized by constant conditional correlations at any conventional level of significance (indicated by the p-values presented above). On the other hand, the rest of the correlations are better described by a time varying conditional correlation model, and for this purpose we use the DCC specification given in equations (5)-(7). Estimates of the volatility and correlation equations is given in Table 4 below. The conditional variance estimates are presented on the left side of the table, while the conditional correlation equation parameters are given on the right side.

Table 4. Estimates of the conditional variance and correlation equations

	Variance equation				Correlation equation		
	Intercept	ARCH	GARCH		*c*	*a*	*b*
Japan	0.94 [0.08]	0.20 [0.04]	0.71 [0.00]	Japan-US	0.54 [0.00]	0.11 [0.08]	0.30 [0.39]
Hong Kong	0.37 [0.07]	0.15 [0.00]	0.82 [0.00]	Hong Kong-US	0.58 [0.00]	0.07 [0.06]	0.79 [0.00]
Singapore	0.28 [0.08]	0.19 [0.00]	0.79 [0.00]	Singapore-US	0.60 [0.00]		
Korea	0.50 [0.02]	0.21 [0.00]	0.75 [0.00]	Korea-US	0.57 [0.00]		
Taiwan	0.37 [0.06]	0.15 [0.00]	0.81 [0.00]	Taiwan-US	0.47 [0.00]		
Thailand	1.17 [0.00]	0.19 [0.00]	0.69 [0.00]	Thailand-US	0.34 [0.00]	0.01 [0.07]	0.80 [0.00]
Australia	0.26 [0.07]	0.26 [0.00]	0.72 [0.00]	Australia-US	0.64 [0.00]	0.04 [0.03]	0.93 [0.00]
New Zealand	0.06 [0.015]	0.08 [0.00]	0.90 [0.00]	New Zealand-US	0.44 [0.00]	0.03 [0.04]	0.95 [0.00]
US	0.26 [0.09]	0.21 [0.03]	0.77 [0.00]				

Note: Variance specification is given in equation (3), while the correlation part is in equation (6); p-values condtructed using Bollerslev-Wooldridge (1992) robust standard errors are presented in square brackets.

Judging by the above estimates, each of the nine variances exhibits strong volatility persistence, with the sum of the ARCH and GARCH parameters close to unity. A significant amount of persistence is also found in the correlations between the following market pairs: Australia-US and New Zealand-US, and to a lesser extent Hong Kong-US and Thailand-US. In contrast, the Japan-US correlation has less memory with the correlation parameter *b* being statistically indifferent from zero. Lastly, the correlations between the markets of Singapore, Korea and Taiwan on one the hand, and the US on the other hand, exhibit no persistence or time variation. The fitted CCC model for those three pairs suggests that the highest level of correlation is between the US and Singapore followed by Korea-US and Taiwan-US correlations. Next we present the estimated weekly correlations graphically in Figure 3.

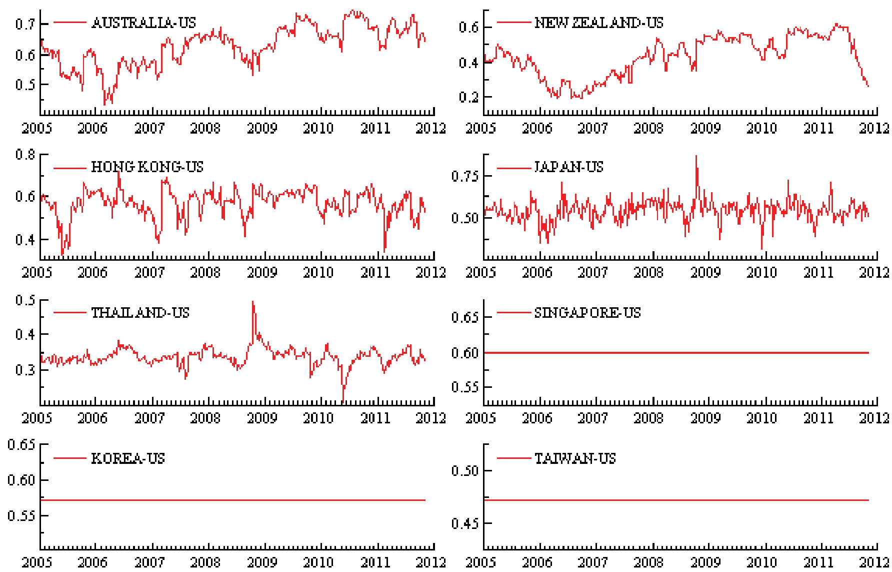

Fig. 3. Estimated weekly correlations

It appears that there are three distinctly different patterns in the correlations. First, Australia-US and New Zealand-US correlations increase steadily from the middle of 2006 and continue their upward trend up until the beginning of 2011. This is when the Australia-US correlation levels off, while the New Zealand-US correlation commences a strong downward trend. Second, Hong Kong-US, Japan-US and Thailand-US market correlations clearly exhibit time varying behavior but also display a strong mean reversion. Each of the three correlation coefficients oscillate around a stable mean with no tendency to follow prolonged periods of upward or downward trend, with a possible exception of the Hong Kong-US correlation. Nevertheless, the Japan-US and Thailand-US market pairs show transitory spikes in the estimated correlations coinciding with the period of the GFC. Lastly, the correlation for Singapore-US, Korea-US and Taiwan-US market pairs are depicted with straight lines generated from the CCC model.

Time varying volatilities corresponding to the parameters estimates presented on the left hand side of the Table 4 above are depicted in Figure 4.

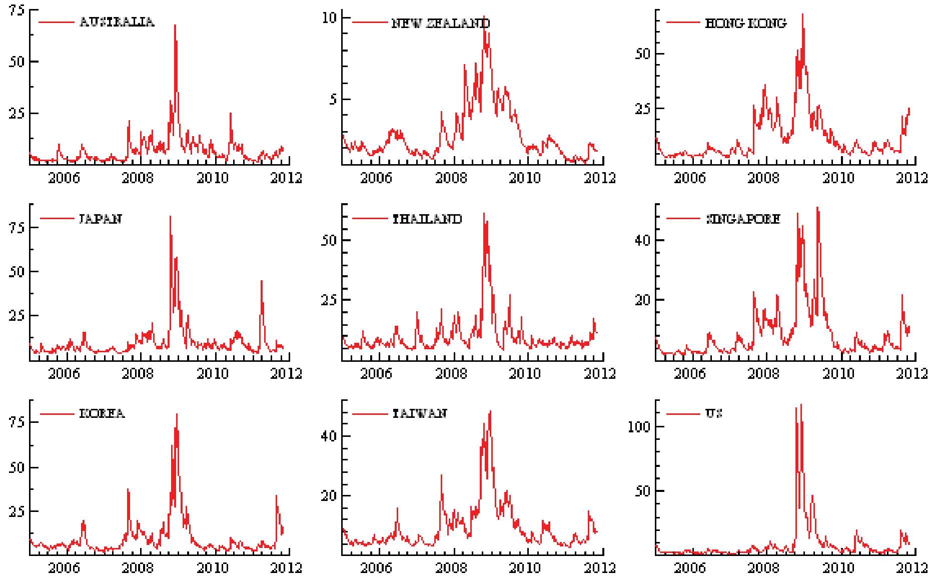

Fig. 4. Estimated weekly variances

The conditional variances illustrate the impact of the GFC over the mid 2007-early 2009 time period. In most cases the variances increase by more than fourfold, while in some instances the change is even more significant. Another interesting observation is the increase in the variances commencing in mid-2011, which may be associated with the European debt problems. This illustrates the phenomenon of financial contagion whereby financial turbulence spreads quickly across different geographical regions.

Lastly, we evaluate how well the estimated models fit the data by providing some residual diagnostic tests in Table 5. Overall, the models fit the data well, with no residual autocorrelation or heteroskedasticity left in the estimated innovations. Specifically, the VAR model fitted in the first step of estimation appears to filter out all of the autocorrelation present in the returns data, as implied by the large p-values of the Q-statistics in the residuals. Similarly the Q-statistics for autocorrelations in squared standardized residuals, which measure persistence in the volatility, indicate that the fitted DCC (CCC) specifications account for most of time varying variances.

Table 5. Residual diagnostic tests (p-values)

	q-statistic (5) (residuals)	q-statistic (10) (residuals)	q-statistic (5) (squared residuals)	q-statistic (10) (squared residuals)
US	0.65	0.77	0.71	0.45
Japan	0.56	0.12	0.78	0.97
Hong Kong	0.99	0.99	0.97	0.87
Singapore	0.59	0.82	0.78	0.92
Korea	0.59	0.44	0.73	0.91
Taiwan	0.53	0.84	0.16	0.39
Thailand	0.21	0.11	0.58	0.42
Australia	0.98	0.92	0.20	0.31
New Zealand	0.91	0.71	0.15	0.29

Conclusion

We investigate the degree of time variation in conditional correlations between eight Asia-Pacific markets and the US equity market. Based on a constant conditional correlation test of Tse (2000) we conclude that three of the eight Asia-Pacific markets exhibit constant conditional correlations with the US. These markets are Singapore, Korea and Taiwan. On the other hand, the markets of Japan, Hong Kong, Thailand, Australia and New Zealand are characterized by dynamic conditional correlations with the US.

Based on the above findings, we fit two types of conditional correlation model to our market pairs; the constant conditional correlation model is fitted to the correlations between Singapore, Korea and Taiwan with the US, while the correlations between the remaining markets of the Asia-Pacific region and the US are modeled using the dynamic conditional correlation model. Conditional variances are estimated with GARCH specifications in for all nine markets.

A practical implication of our study is that we may classify the risk of changing correlations in the context of international portfolio management in three levels. There is a low level of risk of unexpected changes in the correlations between the markets of Singapore, Korea and Taiwan on one hand, and the US on the other. Medium levels of risk of sustained changes in the correlations are present for the markets where correlations are time varying but exhibiting low persistence and high mean reversion. These are the correlations between the US and the equity markets of Japan, Hong Kong and Thailand. Lastly, high risk is present in the correlations for the Australia-US and New Zealand-US market pairs, which pursue prolonged trends in the conditional correlations as displayed in Figure 3. This is due to significant persistence in the estimated correlations implied by the large values of the estimated parameters for these two correlation equations, which sum close to unity.

Lastly, we observe a significant impact of the GFC on the estimates of the GARCH volatilities, which increase simultaneously for the nine markets over the crisis period. The correlations for the Australia-US and New Zealand-US pairs follow an upward trend that covers the period of the GFC so it is difficult to judge the exact impact of the crisis. In contrast, the Japan-US and Thailand-US pairs show clear transitory spikes in the estimated correlations over the period of the crisis. The remaining correlations appear unaffected by the GFC.

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