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# Financial crisis, exchange rate and stock market integration

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## Abstract

After the financial crisis originating from the collapse of the US housing market in 2007, financial markets, including stock markets and foreign exchange markets, experienced drastic fluctuations during an adjustment stage. We examine how the financial crisis affected the linkage between foreign exchange markets and stock markets. First, by examining the daily stock market returns of both Japan and the United States for the sample covering the 2007-2008 financial crisis, we test whether there is a shift in correlation in a smooth-transition correlation GARCH model. We find strong evidence that there was an abrupt upward shift in correlation in June of 2001. There may have been another upward shift of correlation in June of 2008, although the evidence is statistically weak. Second, after adding the JPY/USD exchange rate into a model, we find little correlation between returns of exchange rate and the Japanese stock market, although evidence indicates that there is two-way causality effect between the exchange rate and the Japanese stock market in a VAR framework. This paper provides evidence that a large financial shock may bring financial markets around the world closer to one another.

**Keywords:** Exchange rate; Financial crisis; Japan and US; Smooth transition; Stock market integration.

**JEL classification codes**: F31; F36; G15.

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## 1. Introduction

The financial turmoil that originated from the US housing market caused the market value of listed firms to plummet in stock markets globally. The globalization of financial markets enabled risky lending, namely in the form of housing loans to subprime borrowers, to be shared world-wide via the securitization of loans as mortgage backed securities and consequently affected financial markets everywhere. The most severely affected markets in Europe were Iceland, the UK and Germany. Severe downturns in stock markets also occurred in Asian countries. For example, in Japan, the Nikkei255 index dropped from 18,269 yen at its recent peak in July of 2007 to 7,059 yen at the bottom in March of 2009. However, stock markets seemed to regain their confidence in the last half of 2009.

While we observe the first impact of financial shock to have abated, a negative impact on the real economy can still be observed internationally. The industrial production index for the Japanese manufacturing industry indicates that the production level is only at 80 percent of its level in the pre-crisis period; see Figure 1. The unemployment rate was still rising in the US as of October of 2009, as pictured in Figure 2. Thus, an important general question arises: does a severe financial crisis reverse the process of integration of financial markets across nations?

In this paper, we focus on the degree of integration of financial markets between the US and Japan, including the sample after the subprime loan crisis. In particular, we measure the integration of stock markets of both countries by estimating the time-varying correlation of innovations in two markets using a multivariate GARCH model.

In addition, we investigate causality between the exchange rate and stock markets in a multivariate GARCH model. Both the exchange rate and the stock price index are the discounted present value of expected future economic fundamentals. Expected improvements in future domestic economic conditions increase (appreciate) the price on the stock market (exchange rate). However, the appreciation of domestic currency may have a negative effect on export-oriented manufacturing industries, and this indirect effect of the exchange rate may dwarf the rise in stock market prices.

Our findings are two-fold. First, we found strong evidence that there was an abrupt upward shift in correlation between the US and Japanese stock market returns in June of 2001. There may have been another upward shift of correlation in June of 2008, although the evidence is statistically weak. Second, we found little correlation between exchange rate returns and the Japanese stock market, although evidence indicates that there is one-way causality effect from the exchange rate to the Japanese stock market. We provide preliminary evidence that the zero correlation between the exchange rate and the Japanese stock market may be spurious due to unsynchronized timing.

The reminder of the paper is organized as follows. Section 2 reviews previous empirical evidence on stock market integration between the US and Japan and discusses a subtle distinction between contagion and integration. Section 3 discusses the possible link between the exchange rate and stock market co-movement for Japan. Empirical model specifications and a test for constancy of correlation are provided in section 4. In section 5, we provide evidence for the strengthening of stock market integration between the US and Japan, possibly abruptly in 2001 and gradually again in 2008. The last section concludes with a brief summary and some suggestions for future research.

## 2. Financial Crisis and Stock market integration between the US and Japan

Empirical studies on stock market integration are extensive, covering the relationships between developed countries, between developing countries, and between developed and developing countries, as well as covering all regions. The literature provides evidence that the degree of market integration has increased for most countries.

For emerging countries in Eastern Europe, applying a smooth-transition model for monthly correlations, Chelley-Steely (2005) provides evidence that the stock market integration of Hungary, Poland and the Czech Republic with developed countries increased between 1994 and 1999. Wang and Moore (2008) provide further evidence for increasing correlation between these Eastern European countries and the EMU index for the period between 1994 and 2006. The existence of a high correlation between stock market returns for an emerging country and developed countries may only indicate that the fluctuations in the Dow Jones create ripple effects around the world. By examining 14 emerging countries, nevertheless, Cuadro-Saez et al. (2009) show that shocks in emerging markets have a significant impact on global equity markets.

Many of the existing studies investigating the degree of market integration between the US and Japan, however, have provided evidence that integration between these two countries remain relatively limited. Yang et al. (2003), using the recursive co-integration approach of Hansen and Johansen (1999), find that there is no long-run relationship between the US and Japan for the period between 1970 and 2001. It is noteworthy that a long-run relationship with the US is found for neither the UK or Germany, while the movements of stock market returns are found to be co-integrated between the US and developing countries. Additionally, from the perspective of long-term relationships, Li (2006) examines the stock price indexes of Australia, Japan, New Zealand, the UK, and the US via nonlinear co-integration approach with daily data from May 1992 to April 2001. Using a bivariate model, Japan does not co-integrate with any country, whether via a linear or a nonlinear relationship. Using a constant correlation test for a smoothly time-varying correlation model, Berben and Jansen (2005) provide the evidence that the correlation between the US and Japan remained the same for the period between 1980 and 2000.

However, the casual observation of market headlines suggests a closer co-movement of the US and Japanese stock markets. It is possible that the degree of integration increased after the turn of this century, which is the end of the period investigated in the previous studies. Therefore, our primary interest is to extend the sample coverage to include from 1984 to 2009.

## 3. Exchange rate and stock market

The exchange rate is only a denomination issue in stock market integration literature. For example, Masih and Masih (2001) use the US dollar denomination for all stock market price indexes to adopt the perspective of US investors.

The asset approach to the exchange rate puts forth the current exchange rate as the discounted present value functions of economic fundamentals such as the real income and money supply of two countries. On the other hand, stock price is also represented as the discounted present value of future dividends. Aggregating over all prices of stocks listed in a country, real income can approximately replace aggregated future dividends. For an open economy, the foreign exchange rate may influence the prices listed on the stock market, especially for manufacturing industries.

$$s_t = \sum_{i=0}^{\infty} \gamma_0^i \phi_{0,t+1+i}^e (Y^J, Y^U, M^J, M^U) \tag{1}$$

$$P_t^J = \sum_{i=0}^{\infty} \gamma_1^i \phi_{1,t+1+i}^e (Y^J, s) \tag{2}$$

An increase in the future real income of home country raises the price of stocks and appreciates the home currency. However, the appreciation of the currency may diminish the original impact on prices in the stock market.

## 4. Empirical Model

The return variable is calculated as $R_t^X = 100 * (\ln X_t - \ln X_{t-1})$ for the US stock price, the Japanese stock price, and the value of US dollars in terms of Japanese Yen in the place of X.

### 4-1. Stock market integration model

This subsection introduces a preliminary model that only integrates the stock price index of the US and Japan. This approach is useful in providing a benchmark for the later inclusion of the foreign exchange rate in the trivariate model. First, the returns in the US and Japanese markets are assumed to have a VAR structure, as in equation (3); (see King and Wadhwani, 1990; Hamao et al., 1990). Each return is assumed to possess an autoregressive component and cross-market effect component.

$$\begin{bmatrix} R_t^J \\ R_t^U \end{bmatrix} = \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} + \begin{bmatrix} \beta_{11} & \beta_{12} \\ \beta_{21} & \beta_{22} \end{bmatrix} \begin{bmatrix} R_{t-1}^J \\ R_{t-1}^U \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} \tag{3}$$

The variance of disturbance terms are modeled with a GARCH(1,1) structure in which the variance and covariance of disturbance terms follow ARMA structures;

Bollerslev (1988).

$$\begin{bmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \end{bmatrix} \sim \mathbf{N}(0, H_t), \qquad \text{where } H_t = \begin{bmatrix} h_{11,t} & h_{12,t} \\ h_{21,t} & h_{22,t} \end{bmatrix}. \tag{4}$$

We must discuss the importance of having a VAR-GARCH structure for the investigation of stock market integration. The cross-market coefficients in equation (3), $\beta_{12}$ and $\beta_{21}$, capture the extent of causality, while the correlation term calculated from the GARCH structure in variance-covariance terms captures the co-movement of two stock markets regardless of the direction of causality. We will elaborate on the alternative specifications for the correlation between two markets in the subsections that follow.

While a higher correlation between stock market returns is interpreted as an increase in the degree of market integration, the imperfect information model for two stock markets leads to a contagion effect in which the return of the other market affects the return of the domestic market; King and Wadhwani (1990) and Hamao et al. (1990)[1]. We will examine this effect by applying a vector autoregressive framework to stock market returns in the US and Japan.

**4-1-a. Constant correlation model and tests for constancy of correlation**

For a simple parameterization and for ease in imposing the conditions for the positive definiteness of the covariance matrix, Bollerslev (1990) introduces the constant correlation assumption in GARCH. Each variance term follows the GARCH process in equation (5). The covariance term is restricted using constant correlation parameters and variance terms in equation (6)

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1 A closely related body of literature is that which has to do with contagion in currency crisis studies. After observing the drastic devaluation of Thai bahts, Indonesian rupees and Korean won in the midst of the Asian currency crisis, the research investigated the ‘contagion’ or ‘Asian flu’ phenomenon associated with the currency crisis.

$$h_{11,t} = w_1 + a_1\varepsilon_{1,t-1}\varepsilon_{1,t-1} + b_1 h_{11,t-1}$$
$$h_{22,t} = w_2 + a_2\varepsilon_{2,t-1}\varepsilon_{2,t-1} + b_2 h_{22,t-1} \tag{5}$$

$$h_{21,t} = \rho(h_{11,t}h_{22,t})^{1/2} \tag{6}$$

For ease of presentation with regard to the test statistics, the scaled residuals are denoted as $\hat{\varepsilon}^*_{i,t} = \hat{\varepsilon}_{i,t} \big/ \sqrt{h_{i,t}}$ and pre-multiplied by the inverse of correlation matrix as

$$\left(\hat{v}^*_{1t},\quad \hat{v}^*_{2t}\right)' = \left(\frac{\hat{\varepsilon}^*_{1,t} - \hat{\rho}\hat{\varepsilon}^*_{2,t}}{\sqrt{1-\rho^2}},\quad \frac{\hat{\varepsilon}^*_{2,t} - \hat{\rho}\hat{\varepsilon}^*_{1,t}}{\sqrt{1-\rho^2}}\right). \tag{7}$$

Bera and Kim (2002) propose two versions of IM tests:

$$IM_e = \frac{\left[\sum_{t=1}^{T}(\eta_t)\right]^2}{4T\left(1+4\hat{\rho}^2+\hat{\rho}^4\right)}, \quad \text{where } \eta_t = \hat{v}^{*2}_{1t}\hat{v}^{*2}_{2t} - 1 - 2\hat{\rho}^2 \tag{8}$$

$$IM_s = \frac{\left[\sum_{t=1}^{T}\eta_t\right]^2}{\sum_{t=1}^{T}(\eta_t - \bar{\eta})^2} \tag{9}$$

Berben and Jansen (2005) apply Tse's (2000) Lagrange multiplier test for a smooth-transition GARCH model.

$$\frac{\partial l_t}{\partial \hat{\gamma}} = \left[\left(\hat{v}^*_{1t}\hat{v}^*_{2t} \big/ 1-\rho^2\right) + \left(\rho \big/ 1-\rho^2\right)\right] \tag{10}$$

$$LMC = \frac{\left(\sum_{t=1}^{T}\frac{\partial l_t}{\partial \hat{\gamma}}\right)\left(\sum_{t=1}^{T}\frac{\partial l_t}{\partial \hat{\gamma}}\right)}{\sum_{t=1}^{T}\left(\frac{\partial l_t}{\partial \hat{\gamma}}\frac{\partial l_t}{\partial \hat{\gamma}}\right)} \tag{11}$$

The limiting distributions of all three test statistics are $\chi^2(1)$. Small sample properties via Monte Carlo simulation indicate that the *IMs* test performs better than the *IMe* test in terms of power when the disturbance term follows the t-distribution; Bera and Kim

(2002). The power of the LMC test declines when the transition is linear and the location of the transition is closer to either end of the sample period; Berben and Jansen (2005).

**4-1-b. Smooth-transition correlation GARCH(1,1) model**

Following Lin and Terasvirta (1994), we can model the correlation between the US and Japanese stock markets to follow a smooth transition over the sample period. We follow the specification of Berben and Jansen (2005).

$$\begin{aligned} h_{11,t} &= w_1 + a_1 \varepsilon_{1,t-1} \varepsilon_{1,t-1} + b_1 h_{11,t-1} \\ h_{22,t} &= w_2 + a_2 \varepsilon_{2,t-1} \varepsilon_{2,t-1} + b_2 h_{22,t-1} \end{aligned} \tag{12}$$

$$h_{21,t} = \rho_t (h_{11,t} h_{22,t})^{1/2} \tag{13}$$

$$\rho_t = \rho_0 (1 - G(s_t;\gamma,c)) + \rho_1 G(s_t;\gamma,c) \tag{14}$$

$$G(s_t;\gamma,c) = \frac{1}{1+\exp(-\gamma(s_t - c))}, \qquad \gamma > 0 \tag{15}$$

Correlation in the first regime and second regime is denoted as $\rho_0$ and $\rho_1$, respectively. Time-varying correlation is therefore a weighted average of these two correlations, as in equation (14). The weighting function, $G(s_t;\gamma,c)$, follows the logistic specification, and $\gamma$ and $c$ denote the 'speed' of transition and the (mid) 'point' of transition, respectively. The transition variable, $s_t$, is specified as $t$ divided by the number of observations; therefore, $t \in (0,1]$. Then weight becomes a monotonic function of the transition variable.

We first obtain consistent estimators for $\hat{\alpha}_i$ and $\hat{\beta}_{ij}$ by estimating equation (3) using ordinary least squares. Then using these estimators as initial values, we obtain maximum likelihood estimators for the parameters in their totality, including the

GARCH component of stochastic variances. After we obtain estimated coefficients for $\rho_0, \rho_{1,}\gamma,$ and $c$, we can calculate the time-varying correlation between US and the Japanese stock market innovations.

### 4-1-c. general time-varying correlation model

Symmetric covariance can be simplified in vector representation (Vech) form, imposing restrictions on off-diagonal components of the matrix to be zero, following Bollerslev et al. (1988).

$$\begin{bmatrix} h_{11,t} \\ h_{21,t} \\ h_{22,t} \end{bmatrix} = \begin{bmatrix} w_{11} \\ w_{22} \\ w_{33} \end{bmatrix} + \begin{bmatrix} a_{11} & 0 & 0 \\ 0 & a_{22} & 0 \\ 0 & 0 & a_{33} \end{bmatrix} \begin{bmatrix} \varepsilon_{1t-1}\varepsilon_{1t-1} \\ \varepsilon_{2t-1}\varepsilon_{1t-1} \\ \varepsilon_{2t-1}\varepsilon_{2t-1} \end{bmatrix} + \begin{bmatrix} b_{11} & 0 & 0 \\ 0 & b_{22} & 0 \\ 0 & 0 & b_{33} \end{bmatrix} \begin{bmatrix} h_{11,t-1} \\ h_{21,t-1} \\ h_{22,t-1} \end{bmatrix} \quad (16)$$

We first obtain consistent estimators for $\hat{\alpha}_i$ and $\hat{\beta}_{ij}$ by estimating equation (3) via the ordinary least squares method. Then using these estimators as initial values, we obtain maximum likelihood estimators for the parameters as a whole, including the GARCH part of stochastic variances.

After we obtain the time-varying variances and covariance, we can calculate the time-varying correlation between US and the Japanese stock market innovations.

$$\rho_t = \frac{h_{12,t}}{\sqrt{h_{11,t}h_{22,t}}} \quad (17)$$

### 4-2. Trivariate GARCH(1,1) model

In this subsection, we incorporate the exchange rate series into the previous bivariate model to present the trivariate GARCH (1,1) model. We continue to assume that the returns in the stock markets and foreign exchange market follow a VAR structure as in equation (3).

$$
\begin{bmatrix} R_t^J \\ R_t^U \\ R_t^e \end{bmatrix} = \begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_3 \end{bmatrix} + \begin{bmatrix} \beta_{11} & \beta_{12} & \beta_{13} \\ \beta_{21} & \beta_{22} & \beta_{23} \\ \beta_{31} & \beta_{32} & \beta_{33} \end{bmatrix} \begin{bmatrix} R_{t-1}^J \\ R_{t-1}^U \\ R_{t-1}^e \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \varepsilon_{3t} \end{bmatrix} \tag{18}
$$

The full stochastic variance structure expands to include 78 parameters; however, we continue to restrict off-diagonal elements to zero, reducing the number of parameters to only 18.

$$
\begin{aligned}
\begin{bmatrix} h_{11,t} \\ h_{21,t} \\ h_{31,t} \\ h_{22,t} \\ h_{32,t} \\ h_{33,t} \end{bmatrix} &= \begin{bmatrix} w_{11} \\ w_{22} \\ w_{33} \\ w_{44} \\ w_{55} \\ w_{66} \end{bmatrix} + \begin{bmatrix} a_{11} & & & & & 0 \\ & a_{22} & & & & \\ & & a_{33} & & & \\ & & & a_{44} & & \\ & & & & a_{55} & \\ 0 & & & & & a_{66} \end{bmatrix} \begin{bmatrix} \varepsilon_{1t-1}\varepsilon_{1t-1} \\ \varepsilon_{2t-1}\varepsilon_{1t-1} \\ \varepsilon_{3t-1}\varepsilon_{1t-1} \\ \varepsilon_{2t-1}\varepsilon_{2t-1} \\ \varepsilon_{3t-1}\varepsilon_{2t-1} \\ \varepsilon_{3t-1}\varepsilon_{3t-1} \end{bmatrix} \\
&+ \begin{bmatrix} b_{11} & & & & & 0 \\ & b_{22} & & & & \\ & & b_{33} & & & \\ & & & b_{44} & & \\ & & & & b_{55} & \\ 0 & & & & & b_{66} \end{bmatrix} \begin{bmatrix} h_{11,t-1} \\ h_{21,t-1} \\ h_{31,t-1} \\ h_{22,t-1} \\ h_{32,t-1} \\ h_{33,t-1} \end{bmatrix}
\end{aligned} \tag{19}
$$

As in the previous section, time-varying correlations can be calculated from estimated stochastic variances and covariances. In addition to the correlation between the innovations of two stock markets, we obtain the correlations between the exchange rate and two stock markets.

$$
\rho_t^{JU} = \frac{h_{21,t}}{\sqrt{h_{11,t}h_{22,t}}} \tag{20}
$$

$$\rho_t^{Je} = \frac{h_{31,t}}{\sqrt{h_{11,t}h_{33,t}}}, \quad \rho_t^{Ue} = \frac{h_{32,t}}{\sqrt{h_{22,t}h_{33,t}}} \tag{21}$$

### 4-3. Data

Daily series of foreign exchange rate data for Japanese Yen in terms of the US dollar are obtained from the *Federal Reserve Statistical Release, H.10 Foreign Exchange Rate*, The Federal Reserve Board. The data detail noon buying rates in New York for cable transfers payable in the listed currencies. The stock market prices for the Nikkei Index and the Dow Jones Industrial Average are taken at the opening hour of the corresponding markets. The sample period covers January 1, 1984 to September 18, 2009. All data series are adjusted to have same number of observations via the deletion of data from all series if at least one series reports missing data on any date due to national holidays.

## 5. Empirical results

### 5-1. Financial crisis and stock market integration

As our preliminary analysis, we apply a vector autoregressive model in equation (3) and a variance-covariance matrix to follow equation (15). The result is shown in Table 1. Regarding lag variables in the two markets, we can say that the US market shows very low positive autocorrelation in returns, while the Japanese coefficient is not statistically significant. More interestingly, the linkage between two markets can be measured via the estimated cross-market coefficients. Conforming to casual observations of headline news at the opening hours of Japanese stock market, the US market outcome has a significant impact on the Japanese stock market. The coefficient estimates indicate that a five percent increase in the US market (open during the night in Japan) causes a one percent increase in the Nikkei225 index. On the other hand, no effect of the Japanese market outcome on its US counterpart exists. Our analysis indicates that causality runs only from the US to Japan and that this result is consistent with the findings in Hamao et al. (1990).

Turning to the GARCH coefficients, we note that all parameters in the stochastic variance specification are statistically significant. With the specification in equation (17), the time-varying correlation is plotted in Figure 3. Conforming to the findings of Yang et al. (2003) and Berben and Jansen (2005), the correlation fluctuates around the low level, around 0.2, until the middle of 2001. However, the correlation exhibits a hike and remains at the higher level up to 2005, then experiences a sudden drop in September of 2005 and finally another substantial increase that has extended up to now (2009, September).

Observing a clear change in the correlation after around 2001, we proposed to further analyze the possible existence of a transition of correlation regimes in the following two steps. First, we would formally test the constancy of correlation using the IMe test and IMs test of Bera and Kim (2002) in equations (8) and (9) and the LMC test proposed by Berben and Jansen (2005) in equation (11). Second, we would estimate a smooth-transition correlation model proposed by Berben and Jansen (2005) in equation (5') and (12) through (14) if the null hypothesis of the constancy of correlation was rejected.

The test results are 16.88, 1.86 and 2.91 for IMe, IMs, and LMC, respectively. The critical value at a ten-percent statistical significance level is 2.71. Therefore, we reject the null hypothesis of the constancy of correlation between the US and Japanese stock market. Now, we estimate a smooth-transition correlation model; the results are provided in Table 3. The estimated coefficients are all statistically significant. The initial correlation is 0.158, and the correlation in the next regime is 0.541. The speed of transition is extremely high, 129, and the transition point is 0.680 on June 11, 2001. The correlation plot is depicted in Figure 5. It is noteworthy that our findings are not contradictory to the low correlation found in the studies of Yang et al. (2003) and Berben and Jansen (2005), whose datasets span only up to 2001.

Regarding the effect of the financial crisis on market integration between the US and Japan, we split the sample at the estimated transition point (June 11, 2001) and estimate the latter sub-sample. The estimated coefficient for the transition point is 0.842, which indicates June,2008. Because a comparison of initial and final correlation indicates a (smooth) increase after the transition point, we find that the timing quite consistently matches with the fall of Lehman Brothers in September 2008. The evidence indicates that the financial crisis may have strengthened the integration between the US and Japanese stock markets. This result is consistent with the results of Andersen et al. (2007), who found a stronger contemporaneous effect among the US, UK, and European stock market indices during the contraction period.

**5-2. Financial crisis, exchange rate and stock market**

The impact of exchange rate change on stock markets can be measured in the second column and second row, $\beta_{iFX}$, in Table 4. The depreciation of the Japanese Yen has a positive impact on the Japanese market. This positive causality is consistent with the fact that many listed Japanese companies in the manufacturing industry are heavily involved in export activity. It is also noteworthy that the share of the US in total Japanese exports is substantial and that the use of the US dollar in invoicing currency, even for non-US markets, is exceptionally high. Regarding the reverse effect of a rise in the Japanese stock market on the foreign exchange rate, it can be said that a negative estimated coefficient of $\beta_{iJPN}$ implies the appreciation of the Japanese yen. Thus, these estimated signs are quite consistent with the idea, as discussed, in section 3, that the appreciation of the yen based on the expected future growth of the Japanese economy may mitigate initial rise of stock market prices.

The full trivariate specification including two stock market and exchange rate returns does not qualitatively modify the result in terms of the relationship between the US and Japanese stock market (as noted in the previous section) or between the exchange rate and the Japanese stock market. Puzzlingly, the appreciation of US dollar also has positive effect on the US stock market. This might due in part to the special role of the US currency in the world. The relationship between the foreign exchange rate and the stock market can be quite heterogeneous, reflecting the different economic and political roles of two economies. It would be interesting to pursue this FX-stock market relationship for other developed economies as well as developing economies.

### 5-3. Revisiting (low) negative correlation between the FX rate and Nikkei index

Our analysis used opening-time to opening-time daily return, as shown in Figure 6. It is suspected that this mismatched time-frame for calculating returns may bias the correlation between the exchange rate and the Japanese stock market returns[2].

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[2] Karolyl and Stulz (1996) allow further decompositions of daily returns into closing-time in previous day to opening-time return and opening-time to closing-time in the same day returns. This decomposition makes clearer the timing of innovations in each market and may mitigate this lag problem.

We revisit this problem by comparing matched hourly returns for the period from August 19, 2009 to September 18, 2009[3]. Figure 8 plots the hourly returns of the Nikkei index and the matched hourly exchange rate returns, while daily returns used in this study are plotted in Figure 7. It is surprising that the unconditional correlation for hourly returns is 0.23, while unconditional correlation for daily returns is -0.41. This is just one example for only a month-long period, but it is striking that the sign of the correlation can be reversed when a matched time-frame and higher frequency is used.

## 6. Conclusions and future tasks

We investigated the degree of integration between the Japanese and the US stock market over the sample covering both the pre- and the post-financial crisis in 2007. On the surface, both markets seem to affect each other in terms of daily returns, but the results procured using the vector autoregressive model indicate that causality runs only from the US to Japan. We also provide evidence that the depreciation of the Japanese yen raises the prices in the stock market in Tokyo, while an increase in the Nikkei index alternately appreciates the value of the Japanese yen. Moreover, the degree of integration, measured in terms of a time-varying correlation between innovations in two markets, undergoes a drastic change around 2001. There may have been another upward shift in correlation in June of 2008, although the evidence is statistically weak. This paper provides evidence that a large financial shock may bring financial markets around the world closer together.

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[3] Hourly series are obtained from using Reuter 3000Xtra of Thomson-Reuter.

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Table1. Bivariate GARCH(1) model, DJI and Nikkei

| | US | Japan |
|---|---|---|
| $\alpha_i$ | 0.066*** | 0.056*** |
| | (0.011) | (0.012) |
| $\beta_{iUS}$ | 0.026* | 0.205*** |
| | (0.014) | (0.013) |
| $\beta_{iJPN}$ | 0.000 | 0.008 |
| | (0.009) | (0.013) |
| | GARCH parameters | |
| $c_{11}$ | 0.014*** | |
| | (0.002) | |
| $c_{22}$ | 0.003*** | |
| | (0.001) | |
| $c_{33}$ | 0.018*** | |
| | (0.003) | |
| $a_{11}$ | 0.073*** | |
| | (0.005) | |
| $a_{22}$ | 0.030*** | |
| | (0.004) | |
| $a_{33}$ | 0.084*** | |
| | (0.006) | |
| $b_{11}$ | 0.917*** | |
| | (0.006) | |
| $b_{22}$ | 0.960*** | |
| | (0.004) | |
| $b_{33}$ | 0.908*** | |
| | (0.006) | |
| NOB | 6140 | |
| Log likelihood | -17707.8 | |

Note: Standard errors are computed from numeric second derivatives.

Table 2. Bivariate Smooth time-varying correlation GARCH(1) model

| | US | Japan |
|---|---|---|
| $\alpha_i$ | 0.063*** | 0.057*** |
| | (0.010) | (0.012) |
| $\beta_{iUS}$ | 0.025* | 0.201*** |
| | (0.014) | (0.013) |
| $\beta_{iJPN}$ | 0.000 | 0.012 |
| | (0.009) | (0.013) |
| | GARCH parameters | |
| $c_{11}$ | 0.017*** | |
| | (0.003) | |
| $c_{33}$ | 0.019*** | |
| | (0.003) | |
| $a_{11}$ | 0.081*** | |
| | (0.006) | |
| $a_{33}$ | 0.095*** | |
| | (0.007) | |
| $b_{11}$ | 0.906*** | |
| | (0.007) | |
| $b_{33}$ | 0.898*** | |
| | (0.007) | |
| $\gamma$ | 129.476** | |
| | (62.659) | |
| c | 0.680*** | |
| | (0.005) | |
| $p_0$ | 0.158*** | |
| | (0.015) | |
| $p_1$ | 0.541*** | |
| | (0.015) | |
| NOB | 6140 | |
| Log likelihood | -17679.4 | |

Note: Standard errors are computed from numeric second derivatives. Estimated coefficient for c indicates June 11, 2001.

Table 3. Bivariate Smooth time-varying correlation GARCH(1) model: Sub-sample

| | US | Japan |
|---|---|---|
| $\alpha_i$ | 0.039** | 0.026 |
| | (0.019) | (0.026) |
| $\beta_{iUS}$ | -0.069*** | 0.187*** |
| | (0.027) | (0.029) |
| $\beta_{iJPN}$ | 0.017 | -0.031 |
| | (0.019) | (0.026) |
| | GARCH parameters | |
| $c_{11}$ | 0.018*** | |
| | (0.004) | |
| $c_{33}$ | 0.065*** | |
| | (0.015) | |
| $a_{11}$ | 0.085*** | |
| | (0.010) | |
| $a_{33}$ | 0.115*** | |
| | (0.013) | |
| $b_{11}$ | 0.901*** | |
| | (0.010) | |
| $b_{33}$ | 0.854*** | |
| | (0.016) | |
| $\gamma$ | 0.873** | |
| | (0.422) | |
| c | 0.842 | |
| | (1.154) | |
| $p_0$ | 0.206 | |
| | (0.000) | |
| $p_1$ | 0.973** | |
| | (0.384) | |
| NOB | 1966 | |
| Log likelihood | -5816.35 | |

Note: Standard errors are computed from numeric second derivatives. Estimated coefficient for c indicates June 3, 2008.

Table 4. Bivariate GARCH(1) model, exchange rate and Nikkei

| | FX | Japan |
|---|---|---|
| $\alpha_i$ | -0.008 | 0.061*** |
| | (0.008) | (0.013) |
| $\beta_{iFX}$ | 0.019 | 0.082*** |
| | (0.014) | (0.021) |
| $\beta_{iJPN}$ | -0.022*** | 0.065*** |
| | (0.007) | (0.014) |
| | GARCH parameters | |
| $c_{11}$ | 0.011*** | |
| | (0.002) | |
| $c_{22}$ | -0.010 | |
| | (0.008) | |
| $c_{33}$ | 0.023*** | |
| | (0.004) | |
| $a_{11}$ | 0.042*** | |
| | (0.005) | |
| $a_{22}$ | -0.009 | |
| | (0.012) | |
| $a_{33}$ | 0.101*** | |
| | (0.008) | |
| $b_{11}$ | 0.936*** | |
| | (0.007) | |
| $b_{22}$ | -0.255 | |
| | (0.299) | |
| $b_{33}$ | 0.890*** | |
| | (0.008) | |
| NOB | 6096 | |
| Log likelihood | -15893.3 | |

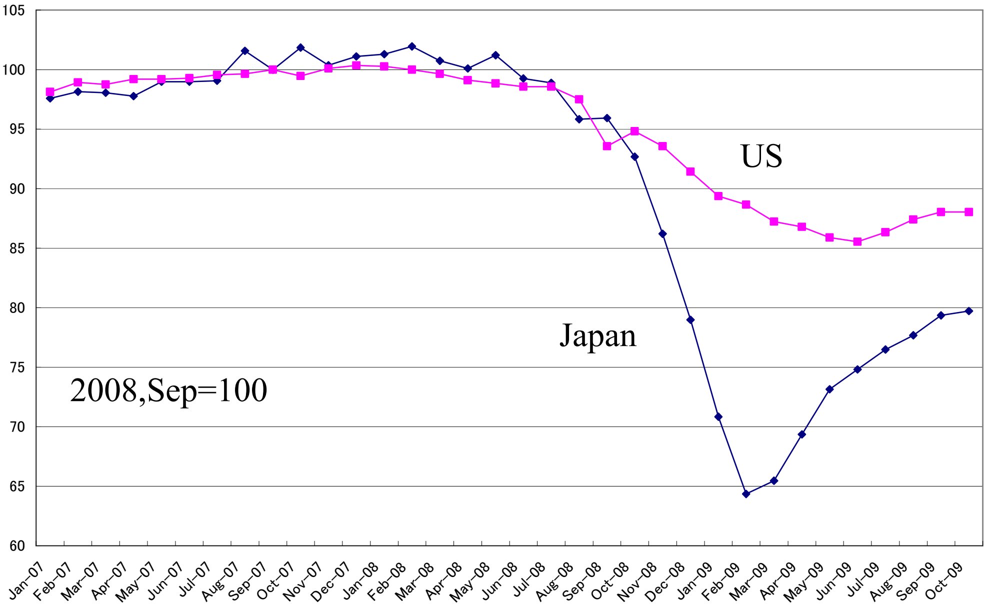
Note: Standard errors are computed from numeric second derivatives.

Table 5. Trivariate ARCH(1) model

| | US | Japan | FX |
|---|---|---|---|
| $\alpha_i$ | 0.058*** | 0.013 | -0.006 |
| | (0.012) | (0.016) | (0.009) |
| $\beta_{iUS}$ | -0.009 | 0.232*** | -0.016** |
| | (0.008) | (0.008) | (0.007) |
| $\beta_{iJ}$ | -0.043*** | -0.021** | -0.020*** |
| | (0.006) | (0.010) | (0.006) |
| $\beta_{iFX}$ | 0.126*** | 0.119*** | 0.019 |
| | (0.014) | (0.017) | (0.014) |
| | ARCH parameters | | |
| $c_{11}$ | 0.898*** | | |
| | (0.013) | | |
| $c_{21}$ | 0.383*** | | |
| | (0.015) | | |
| $c_{31}$ | 0.033*** | | |
| | (0.008) | | |
| $c_{22}$ | 1.359*** | | |
| | (0.020) | | |
| $c_{32}$ | -0.008 | | |
| | (0.010) | | |
| $c_{33}$ | 0.422*** | | |
| | (0.006) | | |
| $a_{11}$ | 0.305*** | | |
| | (0.012) | | |
| $a_{21}$ | 0.078*** | | |
| | (0.014) | | |
| $a_{31}$ | 0.007 | | |
| | (0.013) | | |
| $a_{22}$ | 0.243*** | | |
| | (0.013) | | |
| $a_{32}$ | -0.011 | | |
| | (0.014) | | |
| $a_{33}$ | 0.156*** | | |
| | (0.010) | | |
| NOB | 6095 | | |
| Log likelihood | -19797.1 | | |

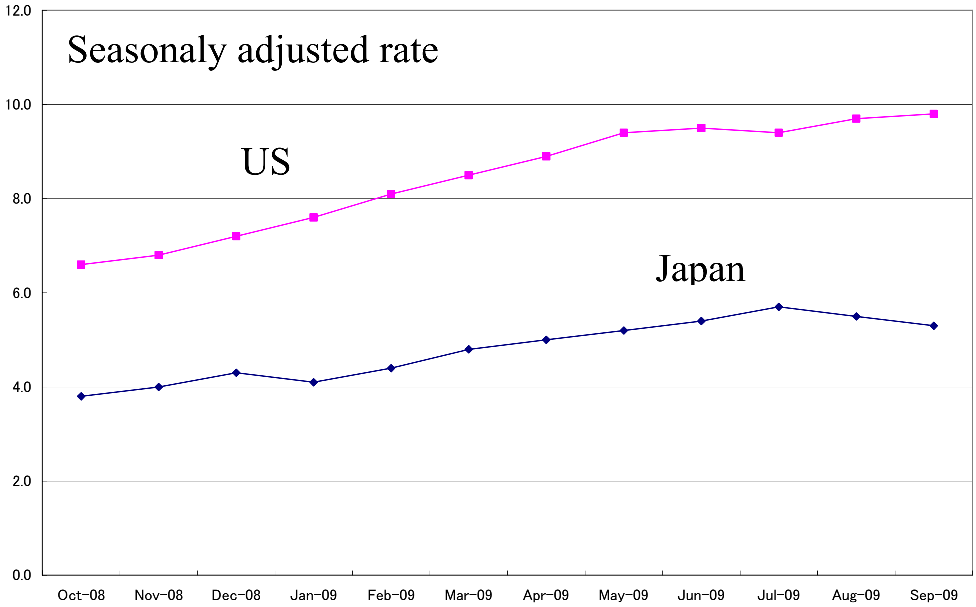
Note: Standard errors are computed from covariance of analytic first derivatives of BHHH.

**Figure 1. Industrial production index for US and Japan**




Source: Industrial production index for total industry form the US Federal Reserve Statistical Release G. 17 and index of industrial production for manufacturing from the Ministry of Economy, Trade and Industry.

**Figure 2. Unemployment rate for US and Japan**




Source: US Bureau of Labor Statistics and Labor Force Survey, Statistical Bureau.

**Figure 3. Time-varying correlation between US and Japanese stock innovations (2-variate GARCH)**

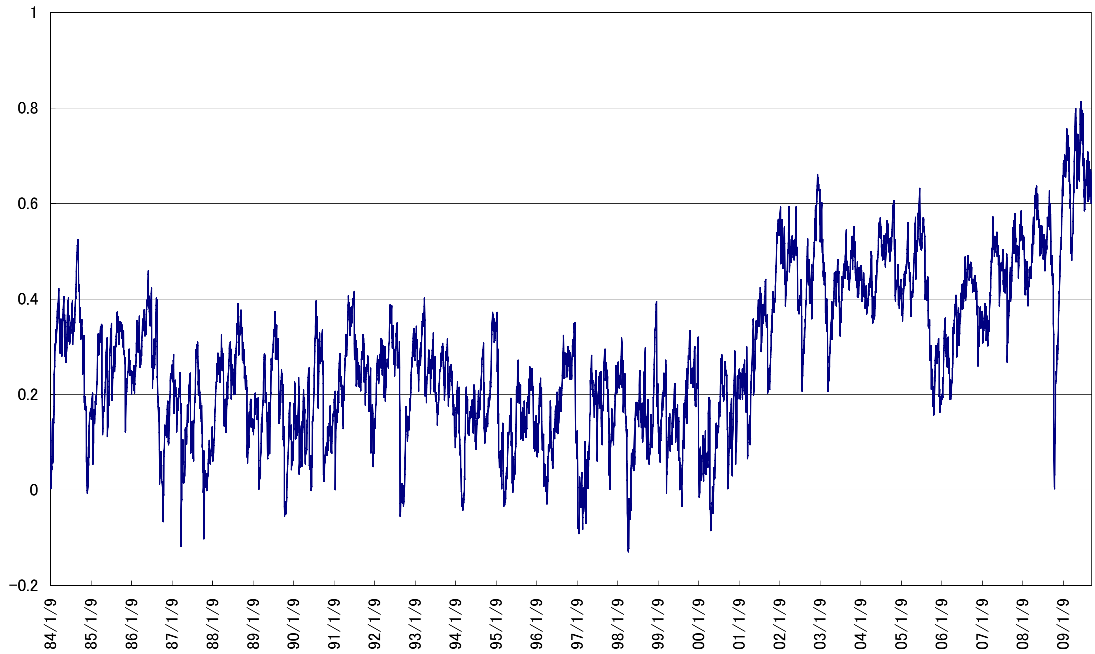



**Figure 4. Time-varying correlation between Japanese stock and JPY/USD innovations (2-varaite GARCH)**

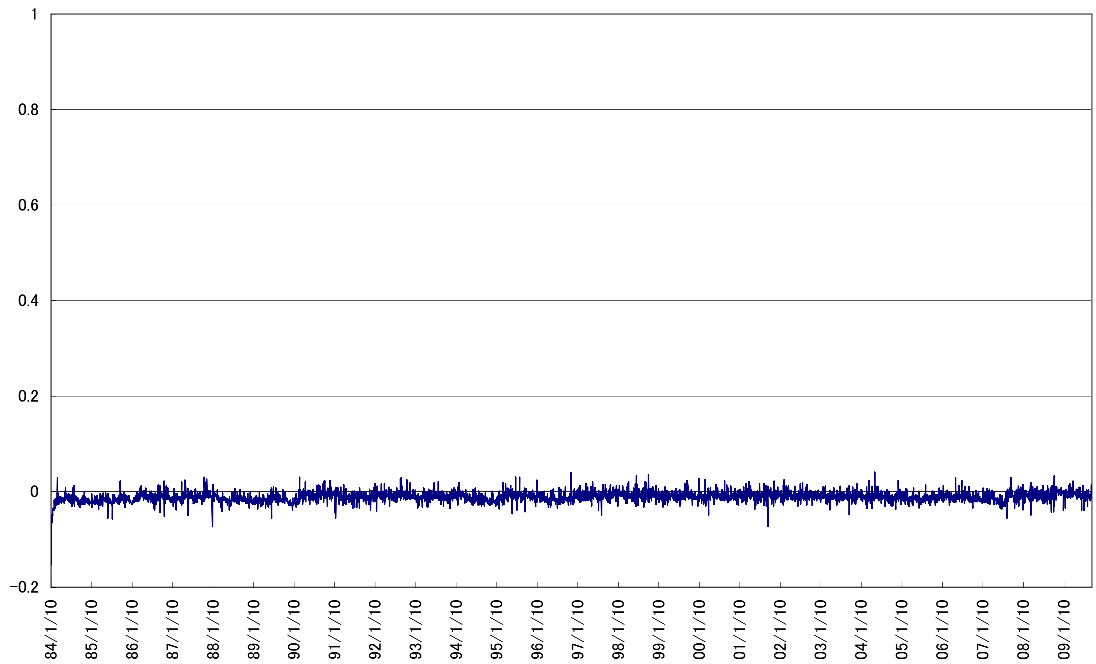


**Figure 5. Smooth transition correlation between the US and Japanese stock market returns**

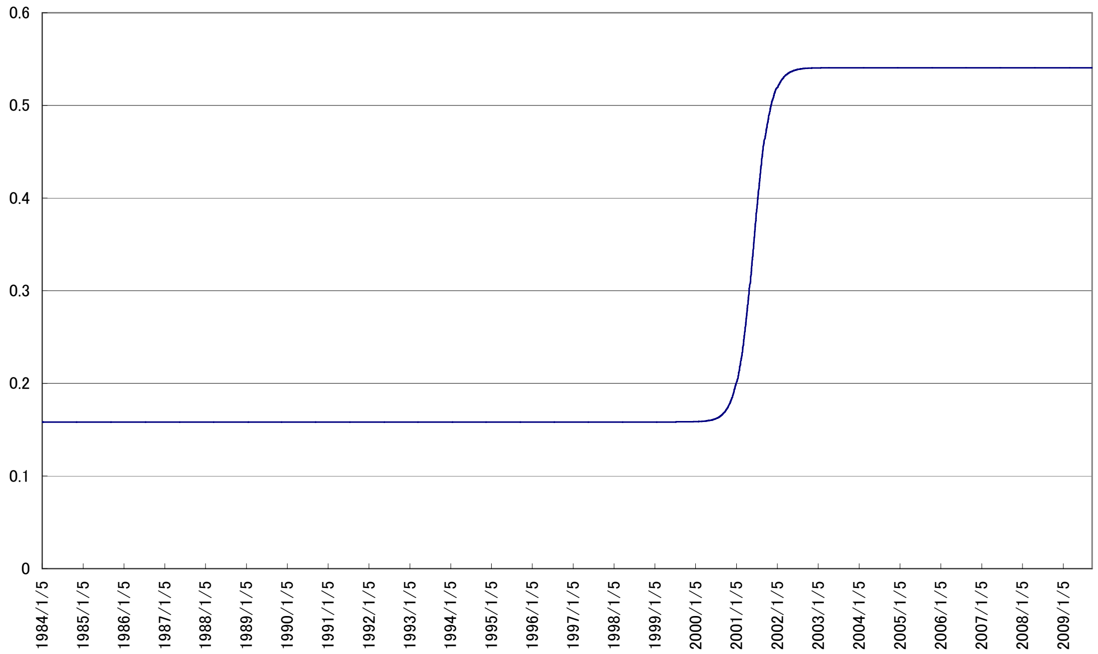



**Figure 6. Sequence of market returns**

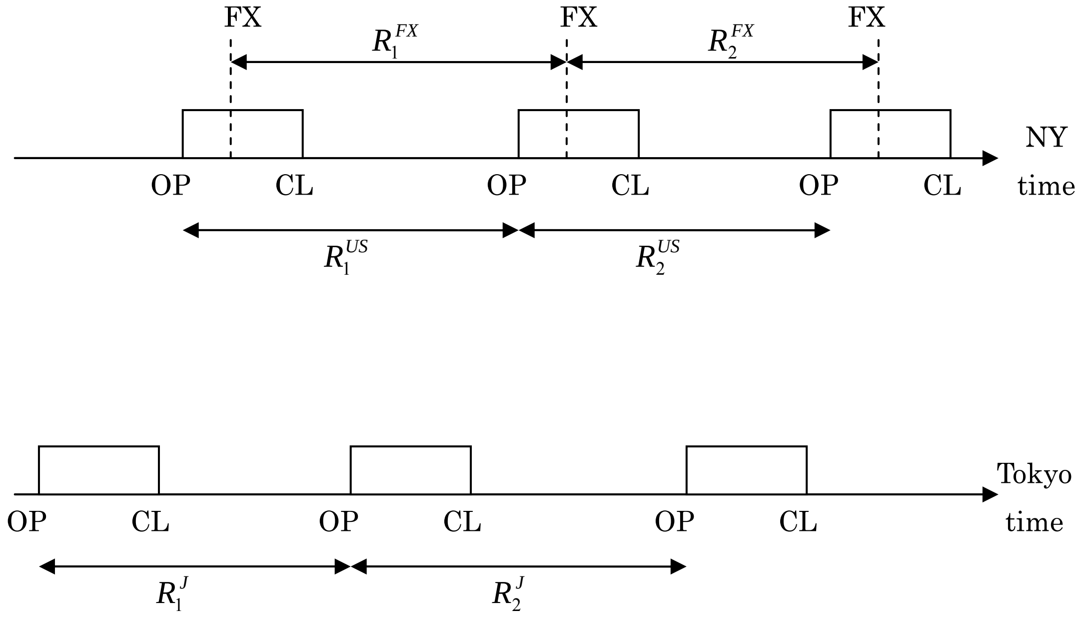


**Figure 7. Daily returns of Nikkei index and JPY/USD rate (Aug 19 – Sep 18)**

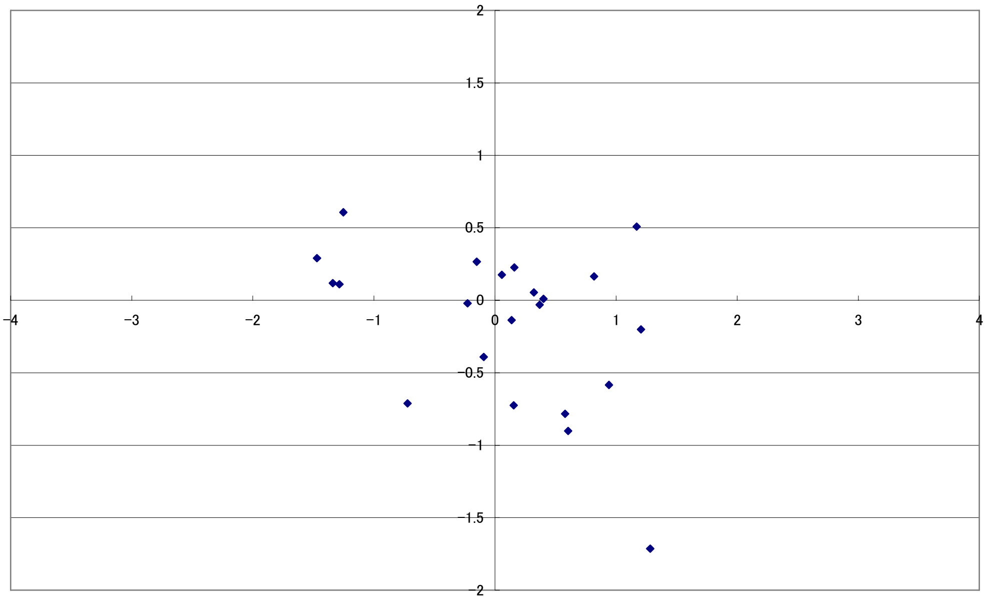



**Figure 8. Hourly returns of Nikkei index and JPY/USD rate**

**(Aug 19, GMT2:00 – Sep 18, GMT 7:00)**