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Real Interest Rate Linkages in the Pacific Basin Region

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Abstract

This paper examines the linkage of real interest rates of a group of Pacific-Basin countries with a focus on East Asia. We consider monthly real interest rates of the US, Japan, Korea, Singapore, and Thailand from 1980 to 2004. The impulse response analysis and half-life estimation are conducted in a multivariate setting, adopting the bias-corrected bootstrap as a means of statistical inference. It is found that the degree of capital market integration has increased after the Asian financial crisis in 1997. The evidence suggests that the crisis has substantially changed the nature of the short run interactions among the real interest rates. Before the crisis, both the US and Japanese capital markets dominated the region. However, after the crisis, the dominance of the Japanese market has completely disappeared, while the US remains as a sole dominant player.

JEL classifications: F36; E44

Key Words: Asian financial crisis, Bias-correction, Bootstrapping, Capital market Integration, Half-life, Impulse response analysis, Vector autoregression.

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1. Introduction

There is growing evidence to suggest that international capital markets have become increasingly integrated. Central to this issue is the real interest rate equalization hypothesis, and testing its empirical validity has been a subject of particular interest. Earlier attempts to test this hypothesis used conventional regression techniques, but the results were overwhelmingly against real interest rate equalization (see, for example, Mishkin, 1984; Mark, 1985; Cumby and Mishkin, 1986; and Merrick and Saunders; 1986). However, as Goodwin and Grennes (1994; p.109) demonstrated, the existence of non-traded goods and transaction costs can render the conditions for real interest rate equalization invalid in the regression context, even when capital markets are efficient and fully integrated. Moreover, Goodwin and Grennes (1994) pointed out that statistical inference based on the conventional regression technique might not be valid when real interest rates exhibit unit-root non-stationarity (see Stock, 1987).

In view of the points listed above, Goodwin and Grennes (1994) argued that the existence of a long run equilibrium among real interest rates should have strong implications for interest parity and efficiently integrated markets. They suggested the use of cointegration analysis (Engle and Granger; 1987; and Johansen; 1988), since it provides a suitable framework to test and estimate long run equilibrium relationships. Their cointegration analysis revealed strong evidence of interest parity and market integration among a number of countries. Subsequent studies by Chinn and Frankel (1995), Hutchison and Singh (1997), Phylaktis (1997, 1999), and Yamada (2002a, 2002b) have adopted cointegration analysis and identified long run relationships among real interest rates.

Given the existence of long run relationships among real interest rates, past studies have examined their short run interactions and attempted to determine which rate is leading others as a dominant force. Chinn and Frankel (1995) investigated the relative influence of US and Japanese real interest rates in the Pacific Rim region, where they presented evidence that most East Asian countries are linked with the US and Japan, forming another piece of the consensus in the literature that market integration has been increasing. They also found that the Japanese rate gained significant influence. Another notable example is Phylatkis (1999), who used impulse response analysis to examine the short run dynamics of the real interest rates of Pacific-Basin countries. She found an increasing degree of capital market integration after the financial market deregulation in the eighties. In addition, the US and Japanese capital markets were found to dominate the others, with the latter becoming increasingly more important. In this paper, we also examine the case of Pacific-Basin countries, with a focus on East Asian countries, using updated data and more sophisticated econometric methods. We consider three representative East Asian countries (Korea, Singapore and Thailand) along with the US and Japan, paying attention to the impact of the Asian financial crisis in 1997. We examine the existence of long run equilibrium relationships, short run dynamics and the issue of dominance among the real interest rates, before and after the crisis.

As in Phylaktis (1999), we employ impulse response analysis based on the vector autoregressive (VAR) model. We attempt to estimate half-lives of the real interest rates in the VAR context, as a measure of persistence. The half-life is defined as the number of periods required for the response of a time series, to its own shock, to be halved, and can be readily estimated from an impulse response function. If a time series in a cointegrated VAR model is found to be mean-averting in the sense that it shows a permanent response to its own shock, it can be argued that it represents a common trend of the system. It can also be regarded as a dominant force, since the others with smaller values of half-lives can be thought of as equilibrating factors with mean-reverting behaviour.

In order to conduct improved statistical inference for the impulse response analysis and half-life estimation, we resort to confidence intervals based on the bootstrap method (Efron, 1979). Bootstrap inference is useful in small samples, especially when the data is non-normal or heteroskedastic, where conventional asymptotic inference based on a normal approximation may perform poorly. In addition, small sample biases of VAR parameter estimators (see, for example, Abadir et al; 1999) can further undermine the reliability of the asymptotic method. In this paper, we use the bias-corrected bootstrap confidence interval of Kilian (1998a, 1998b)[1]. It has been found to exhibit much better small sample properties than conventional confidence intervals, especially for VAR models whose characteristic roots are close or equal to one. It can be made applicable to VAR models with non-normal or heteroskedastic innovations using the wild bootstrap of Mammen (1993).

The main finding of the paper is that the degree of capital market integration of Pacific-Basin countries has increased after the Asian financial crisis. The crisis also has changed the nature of short run dynamics among real interest rates. In particular, the dominance of Japan in this region appears to be a purely pre-crisis phenomenon, while the US maintains a strong dominance even after the crisis. In the next section,

[1] The importance of bias-correction in econometric analysis is well documented. See, for example, Andrews and Chen (1994).

we discuss the data details and the results of the preliminary analysis. Section 3 provides a summary of the methodologies used in the paper. Section 4 presents the empirical results, and Section 5 concludes the paper.

2. Data details and preliminary analysis

We have selected real interest rates of five countries; Japan, Korea, Singapore, Thailand and the US. This choice is based on the consideration that VAR dimension should be kept manageable for parsimonious parameterization. These countries also represent a good mixture of developed and developing countries in the Pacific-Basin region, with diverse characteristics and different degrees of maturity of capital markets. The Asian financial crisis in 1997 affected currencies, stock prices, and other asset prices of several East Asian countries. Korea and Thailand were the two severely hit, while Singapore was relatively unaffected. Japan was going through a long-term recession in the nineties, which was further exacerbated by the crisis.

We have used 204 monthly observations of real interest rates from 1980:1 to 1996:12 (pre-crisis period; Period I), and 70 for the period of 1999:1 to 2004:10 (post-crisis period; Period II). The starting date reflects the timing of deregulation where most Asian countries started to liberate their financial markets. All observations from 1997 and 1998 were excluded to eliminate noisy and unstable observations. Most of the East Asian countries hit by the crisis were already in financial distress in the fist half of 1997 and they only started to display a sign of recovery in September 1998.

We use short-term interest rates for these countries. The monthly money market rate is used for Korea and Thailand, and the T-bill rate for the US. For Japan, we use the call rate, while the interbank rate has been used for Singapore. To calculate the rate of inflation, the consumer price index is seasonally adjusted with the X-12 method using geometric weights. All nominal interest rates are then deflated by the ex post inflation rate in order to generate the ex post real interest rate series. All data are obtained from *International Financial Statistic* Database.

Visual inspection of the time plots indicates that the real interest rates show local trends with highly volatile fluctuations, although Japanese and Singaporean rates show fairly weak downward global trends. On this basis, we decided not to include a linear time trend in our testing and estimation below. As argued by Yamada (2002a; p.280), this can provide more reliable empirical results. To determine whether the real interest rates series possess unit-roots, we conducted the augmented Dickey-Fuller (ADF) test. All real interest rates appear to be integrated of order 1 at the 5% level of significance, both before and after the crisis, with the exception of the Japanese rate after the crisis. The details of the ADF test are available upon request.

The presence of a unit root in the real interest rate may be arguable. Apart from economic reasons, there are the well-known statistical issues that the ADF test is an asymptotic test that can have size distortion and low power in small samples. However, as Goodwin and Grennes (1994; footnote 5) pointed out, the justification for the presence of a unit root can be found from past empirical evidence and practical considerations. Based on this, we assume that all real interest rates are integrated of order 1 for the purpose of the cointegration analysis. It should be noted, however, that our analysis based on VAR impulse response functions and the associated bias-corrected bootstrap does not require the real interest rate to possess unit roots.

As a preliminary analysis, we have conducted pairwise and multiple cointegration testing using the Engel-Granger (1987) methodology. Although the cointegrating relationship has been identified in many cases, the condition of real interest rate equalization fails to hold for all cases. The use of fully-modified OLS estimation of Phillips and Hansen (1990) has led to similar findings. This may not be surprising in view of the argument put forward by Goodwin and Grennes (1994), in relation to the existence of non-traded goods and transaction costs.

3. Methodology

3.1 VAR Model and Cointegration

We consider the K-dimensional vector autoregressive (VAR) model of the form

$$Y_t = \nu + B_1 Y_{t-1} + \dots + B_p Y_{t-p} + u_t, \qquad (1)$$

where Y_t is the $K\times1$ vector of variables at time t, ν is the $K\times1$ vector of intercepts, and B_i's are the $K\times K$ matrices of coefficients. Note that u_t is the $K\times1$ vector of innovations with $\mathrm{E}(u_t) = 0$ and $\mathrm{E}(u_t u_t') = \Sigma_u = PP'$[2]. The above VAR system can be written in the vector error correction (VEC) form as

$$\Delta Y_t = \nu + \Gamma_1 \Delta Y_{t-1} + \dots + \Gamma_{p-1} \Delta Y_{t-p+1} + \Pi Y_{t-1} + u_t, \qquad (2)$$

where $\Pi = B_1 + \dots + B_p - I_K$ and $\Gamma_i = -(B_{i+1} + \dots + B_p)$. When Y_t is cointegrated with cointegration rank r, Rank(Π) = $r < K$ and $\Pi = \alpha\beta'$ where α and β are respectively $K\times r$ matrices.

[2] We assume homoskedastic innovations to begin with, but this assumption may be relaxed later to accommodate heteroskedastic innovations.

The unknown VAR order p in (1) is estimated to ensure that the residuals of each equation in the VAR mimic a white noise process. We employ a simple to general approach to model selection for parsimonious parameterisation. Visual inspection of residual autocorrelation function is conducted, in addition to the Ljung-Box test and Akaike's information criterion (AIC). To determine the cointegration rank and estimate the unknown parameters in the VEC model (2), we follow Johansen's (1988) method based on the maximum likelihood principle. The trace and maximal eigenvalue tests of Johansen (1988) are used to determine the cointegration rank. The details of this testing and estimation method are not presented in this paper, because they are well documented elsewhere (see, for example, Lütkepohl, 1991; Chapter 11; Hamilton; 1994; Chapter 20).

3.2 Impulse response analysis and half-life estimation

The VAR model given in (1) can be used for the (orthogonalized) impulse response analysis. It is a dynamic multiplier analysis among the variables in the VAR system, measuring how a one-standard deviation shock to a variable is transmitted to others over time (see, for details, Lütkepohl, 1991). It has been applied widely in empirical macroeconomics and international finance (see, for example, Eichenbaum and Evans, 1995). It is also closely related to testing for non-causality, as zero impulse responses between two variables imply no causality (Lütkepohl, 1991; p.45).

The orthogonalized impulse responses are calculated from the coefficients of the MA(∞) representation of the VAR model and the residual covariance matrix. Given n realizations (Y_1, …, Y_n) of (1), the unknown coefficients are estimated using the least-squares (LS) method. The LS estimators for $B = (\nu, B_1, \ldots, B_p)$ and Σ_u are denoted as $\hat{B} = (\hat{\nu}, \hat{B}_1, ..., \hat{B}_p)$ and $\hat{\Sigma}_u$, and the associated vector of residuals as $\{\hat{u}_t\}_{t=p+1}^{n}$. The orthogonalized impulse responses are defined as $\Theta_i = \Phi_i P$ where $\Sigma_u = PP'$ and Φ_i's are the coefficients of the MA(∞) representation of (1). A typical element of Θ_i is denoted as $\theta_{kl,i}$, and it is interpreted as the response of the variable k to a one-time impulse in variable l, i periods ago. The plot of $\theta_{kl,i}$ against i is called the impulse response function of the variable k to a one-time impulse in variable l. Using $\hat{B}$ and $\hat{\Sigma}_u$, the estimator for impulse response $\hat{\theta}_{kl,i}$ for $\theta_{kl,i}$, can be calculated.

As mentioned earlier, the half-life of a time series is defined as the number of periods required for the response of a time series, to its own shock, to be halved. As such, it can readily be obtained from the impulse response function of a time series. It has been used as a popular measure of persistence for key time series in international finance; see, for example, Rapach and Wohar (2004) for real interest rates and Murray and Papell (2002) for real exchange rates. For the univariate AR(1) model with the coefficient α, the analytical expression for the half-life is $h = \ln(0.5)/\ln(\alpha)$. As the value α of approaches 1, the value of h approaches infinity, indicating that the response of the time series to its own shock becomes permanent. For an AR(p) model with $p > 1$, the value of h can be calculated from the impulse response function. In the VAR case, the half-life of the k^{th} time series in the system, denoted as h_k can be calculated from the impulse response function to its own shock, namely $\theta_{kk,i}$, where $k = 1, \ldots, K$. The half-life estimator for h_k, $\hat{h}_k$, can be obtained from $\hat{\theta}_{kk,i}$.

3.3. Bias-corrected bootstrap

The bootstrap is a computer-intensive method of approximating the sampling distribution of a statistic. It has been applied widely in econometrics and is often found to provide a superior alternative to the conventional methods in small samples (see, Li and Maddala, 1996; Berkowitz and Kilian, 2000; and MacKinnon, 2002). The conventional bootstrap, however, is applicable to data generated from an identical and independently distributed (i.i.d.) random variable. Similarly, Kilian's (1998a, 1998b) bias-corrected bootstrap is applicable to the VAR model whose innovations follow an i.i.d. distribution. This conventional bootstrap may not work properly when the VAR model shows conditionally heteroskedastic error terms, which is the case for the VAR model fitted in this paper (see Section 4). Recently, a bootstrap procedure called the wild bootstrap (Mammen, 1993) has been developed, which is applicable to a time series with conditional or unconditional heteroskedasticity of unknown form. The theoretical underpinning of the wild bootstrap in the context of univariate AR model can be found in Gonclaves and Kilian (2004).

In conducting the impulse response analysis, it is important to test whether impulse response estimates are statistically different from 0. This is closely related to testing for causality among the variables in the VAR system. We employ confidence intervals for the impulse response for this purpose. Similarly, it is informative to report a confidence interval for the half-life, as it provides a range that contains the true value with a certain degree of confidence. Note that impulse response estimates and half-life estimates are necessarily biased in small samples, due to small sample biases present in the VAR parameter estimators (see Tjostheim and Paulsen, 1983; Nicholls and Pope, 1988; Pope, 1990; and Abadir et al., 1999). The biases are particularly severe when the VAR model has unit roots or near unit roots; when the VAR dimension K is larger; or when the sample size is smaller. It is highly likely that these biases adversely affect the small sample properties of the confidence intervals.

To obtain confidence intervals with improved small sample properties, Kilian (1998a, 1998b) proposed the use of the bias-corrected bootstrap (or bootstrap-after-bootstrap). It is a bootstrap method of constructing confidence intervals, in which the biases associated with parameter estimators are adjusted in two stages of the bootstrap. Kilian (1998a, 1998b) found that the bias-corrected bootstrap confidence interval has small sample properties far superior to its conventional alternatives, including those based on the asymptotic method detailed in Lütkepohl (1991). Although it was originally proposed for statistical inference of impulse response, the bias-corrected bootstrap can easily be adapted to half-life estimation. In the univariate case, Murray and Papell (2002) made a similar attempt to construct bias-corrected confidence intervals for the half-life of the deviation from purchasing power parity[3].

The bias-corrected bootstrap of Kilian (1998a, 1998b) involves two stages of bias-correction for VAR parameter estimates. Here we follow Kilian (1998b) in using Pope's (1990; p.253) asymptotic bias formula to obtain bias-corrected parameter estimators. Note that Pope's (1990) formula estimates bias to the order of n^{-1}, and is applicable to the VAR model with martingale difference innovations with a fixed

[3] Murray and Papell (2002) used the Andrews-Chen (1994) median-unbiased estimators for bias-correction. The bias-correction in the second stage of the bootstrap, however, was not conducted. The method here is more general, since it calculates the impulse response functions in a multivariate setting, with an additional stage of the bias-correction.

covariance matrix, which includes non-normal or conditionally heteroskedastic errors as special cases.

The bias-corrected confidence interval for $\theta_{kl,i}$ can be obtained as below:

In Stage 1, Pope's (1990) formula is applied to $\hat{B} = (\hat{\nu}, \hat{B}_1, ..., \hat{B}_p)$ to obtain the bias-corrected estimator $\hat{B}^c = (\hat{\nu}^c, \hat{B}_1^c, ..., \hat{B}_p^c)$ for B. It is possible that $\hat{B}$ satisfies the condition of stationarity, while $\hat{B}^c$ does not. In this case, Kilian (1998a, 1998b) suggested an adjustment to $\hat{B}^c$ so that it implies stationarity. This adjustment is called the stationarity correction[4], and its details can be found in Kilian (1998a, 1998b).

In Stage 2, generate a pseudo data set following the recursion

$$Y_t^* = \hat{\nu}^c + \hat{B}_1^c Y_{t-1}^* + ... + \hat{B}_p^c Y_{t-p}^* + u_t^*, \tag{3}$$

using the first p values of the original data as starting values. When the innovations are heteroskedastic, we adopt the wild bootstrap that involves generating $u_t^* = \eta_t \hat{u}_t$, where η_t is any scalar random variable whose mean is zero and variance is one. When the innovations are homoskedastic, u_t^*'s are generated as random resampling of $\hat{u}_t$'s with replacement following Kilian (1998b).

In Stage 3, using $\{Y_t^*\}_{t=1}^n$, the VAR coefficient matrices are re-estimated and denoted as $\hat{B}^* = (\hat{\nu}^*, \hat{B}_1^*, ..., \hat{B}_p^*)$. Pope's (1990) bias formula is again applied to $\hat{B}^*$ in order to

[4] This stationarity correction is also important in establishing the asymptotic validity of this bias-corrected bootstrap procedure. See, for details, Kilian (1998a, 1998b)

obtain a bias-corrected version $\hat{B}^{*c} = (\hat{v}^{*c}, \hat{B}_1^{*c}, ..., \hat{B}_p^{*c})$ of $\hat{B}^*$. The stationarity correction is again applied to $\hat{B}^{*c}$ if necessary.

Repeat Stages 2 and 3 sufficiently many times, say m, to generate bootstrap replicates of $\left\{\hat{B}^{*c}(j)\right\}_{j=1}^{m}$, from which m bootstrap replicates $\left\{\hat{\theta}^*_{kl,i}(j)\right\}_{j=1}^{m}$ of impulse responses are obtained. In this paper, m is set to 2000, which is sufficiently large to obtain reliable bootstrap confidence intervals (see Efron and Tibshirani, 1993). The 100(1-2α)% bias-corrected bootstrap confidence intervals for $\theta_{kl,i}$ can be obtained as the interval [$\hat{\theta}^*_{kl,i}(\alpha)$, $\hat{\theta}^*_{kl,i}(1\text{-}\alpha)$], where $\hat{\theta}^*_{kl,i}(q)$ is the qth percentile from the distribution of m bootstrap replicates $\left\{\hat{\theta}^*_{kl,i}(j)\right\}_{j=1}^{m}$, based on the percentile method of Efron and Tibshirani (1993, p.160).

To construct the bias-corrected bootstrap confidence interval for h_k, the bootstrap replicates of half-life $\left\{\hat{h}^*_k(j)\right\}_{j=1}^{m}$ are obtained from $\left\{\hat{\theta}^*_{kk,i}(j)\right\}_{j=1}^{m}$ in Stage 3. The 100(1-2α)% bias-corrected bootstrap confidence intervals for h_k can be constructed as the interval [$\hat{h}^*_k(\alpha), \hat{h}^*_k(1\text{-}\alpha)$], where $\hat{h}^*_k(q)$ is the qth percentile from the distribution of m bootstrap replicates $\left\{\hat{h}^*_k(j)\right\}_{j=1}^{m}$. The half-lives are calculated from the impulse responses over 240 months, i.e., $\hat{\theta}^*_{kk,i}$ with i =1, …, 240. If the impulse response does not halve in 240 months, $\hat{h}^*_k$ is set to infinity indicating that the response is practically permanent[5].

[5] Due to the stationarity correction implemented in the bootstrap procedure, a bootstrap replicate of half-life cannot take the value of infinity. This is different from Murray and Papell (2002) and Rapach

Note that the wild bootstrap described here is referred to as the recursive-design wild bootstrap, which is preferred by Gonclaves and Kilian (2004) to the other types of the wild bootstrap on the basis of superior small sample performance. The distinctive feature of the wild bootstrap is that u_t^*'s are generated as a random weighting of $\hat{u}_t$'s, so that $E(u_t^* \mid \hat{u}_t) = 0$ and $E(u_t^* u_t^*{}' \mid \hat{u}_t) = \hat{u}_t \hat{u}_t'$. Throughout the paper, we report the results associated with the case where η_t follows the standard normal distribution, since the results are not sensitive to the other choices.

4. Empirical Results

In conducting the orthogonalized impulse response analysis, the ordering of the variables in the VAR system is important. In this paper, we specify the ordering on the basis of the Wold-causality (see, Lütkepohl, 1991; p.52). We place the US real interest rate first, followed by the Japanese, Korean, Singaporean, and Thai real interest rates. In the context of orthogonalized impulse response analysis, this amounts to assuming the instantaneous causality running one way from the US rate to Thai rate. This is reasonable considering the relative power and scale of the economies of these countries. We also have conducted the generalized impulse response analysis of Pesaran and Shin (1998), which is invariant to the ordering of the variables. The results (point estimates and bootstrap confidence intervals as detailed in the previous section) are found to be qualitatively no different.

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and Wohar (2004), where the half-life estimate is allowed to take the value of infinity. This is because these authors have used estimation methods that allow parameter estimators to take non-stationary values.

4.1. Cointegration and error-correction models

On the basis of a number of statistical measures including AIC and residual portmanteau statistics, we have found the VAR order to be 2 for both Period I and II. The associated residual diagnostics are reported in Table 1. According to the Box-Ljung test, the residuals of all equations show no serial correlation, except for the US equation in Period I. The likelihood ratio test for multivariate white noise (not reported) indicates no serial correlation in the VAR innovations in Period I. All equations show good fit, except for the equation for the Japanese rate in Period II. There is strong evidence of non-normality and heteroskedasticity in the residuals in Period I, while the residuals in Period II show no evidence of departure from i.i.d. normality. This indicates that the Asian crisis has greatly altered the nature of the innovations in the data generation processes of the real interest rates.

Table 2 reports the Johansen cointegration test results. It is evident that there are two cointegrating vectors in Period I, while four cointegrating vectors are present in Period II. In other words, the number of common trends was three before the Asian crisis, but one common trend is driving the system afterwards. Hence, the degree of market integration has increased after the crisis[6]. From the error correction model estimation, rich short run dynamic interactions are found among the rates in the system. However, given the strong non-normality and heteoskedasticty observed in Period I and the small sample size of Period II, we prefer the bias-corrected bootstrap inference as a means of examining short run dynamics. The magnitudes of the estimated cointegrating vectors and tests using appropriate restrictions do not appear

[6] We have tested various restrictions on the cointegrating vectors in relation to the real interest rate equalization hypothesis, but none was supported by the data, both before and after the crisis.

to be supportive of the real interest equalisation proposition. Details of the estimated cointegrating vectors and error correction model parameters are available upon request.

4.2. Impulse response analysis and half-life estimation

Figure 1 presents impulse response functions and their 95% confidence bands for Period I. We also have calculated 90% and 99% confidence bands, but they are not reported for brevity. However, they will be discussed when necessary. There are five panels in Figure 1, each exhibiting dynamic responses of all real interest rates when a shock is given to a particular rate. If a confidence interval contains zero, the null hypothesis that the true response is zero cannot be rejected at the specified level of significance. The confidence intervals are calculated using the bias-corrected wild bootstrap for Period I, as it is evident that the VAR innovations in Period I are heteroskedastic.

From the first panel, the Singaporean and Thai rates show positive responses to a shock in the US rate for more than 12 months. The Japanese and Korean rates do not show any statistically significant non-zero responses. To a shock in the Japanese rate, there is weak evidence that only the Korean rate shows a positive response in month 1. The lower limit of the 95% interval is slightly smaller than zero, while that of the 90% interval (not reported) is positive. This indicates that the Korean rate shows a positive response to the Japanese rate in month 1, at the 10% level of significance. It is also evident that the Korean, Singaporean and Thai rates do not affect the other rates over time. Hence, we have identified one-way causality from the US rate to the

Singaporean rate, one-way causality from the US rate to the Thai rate, and one-way causality from the Japanese rate to the Korean rate.

The impulse response functions of the US and Japanese rates to their respective own shocks, given in Figure 1, are much flatter than the others, indicating high degrees of persistence. From Table 3, the half-life estimates of the US and Japanese rates are about 12 and 20 months respectively, much higher than those of the other rates. The upper limits of the 95% confidence intervals for the half-lives of the US and Japanese rates appear to be infinite. In this case, one cannot reject the null hypothesis that the response of a time series to its own shock is permanent at 5% level of significance. This means that these two rates have the degree of persistence equivalent to a time series with a unit root and represent common stochastic trends in the cointegrated VAR. Other rates have finite upper confidence limits, which mean that they show quick adjustment with mean-reversion. Thus, we found that the US and Japanese rates are the two dominant rates in the East Asia region, as also found by Phylaktis (1999).

Figure 2 presents the impulse response functions and their 95% confidence bands for Period II. Note that the confidence intervals are calculated using the bias-corrected bootstrap based on i.i.d. resampling, as there is no evidence of heteroskedasticity in the innovations. From the first panel, it is evident that the Korean and Singaporean rates are affected by the US rate, as they show positive responses for more than 12 months to a shock in the US rate. This is different from the case of Period I, where the US rate exerts direct influence only on the Singaporean and Thai rates. The Korean rate is directly affected by the US rate after the crisis, which was not the case before. As for the Thai rate, the reverse is evident, as it is no longer affected by the US rate after the crisis. This is interesting since these two countries were heavily affected by the crisis. From Figure 2, no further dynamic relationship is observed, except that the Singaporean rate affects the Thai rate positively for 2 months. Note that the Japanese rate does not affect other rates in Period II. Hence, there exist one-way causality from the US rate to the Korean rate, one-way causality from the US rate to the Singaporean rate, and one-way causality from the Singaporean rate to the Thai rate.

From Figure 2, it can be observed that the Japanese rate shows positive responses after period 6 in response to the shock in the US rate, and that the US rate responds negatively at period 3 to a shock in the Japanese rate. This may suggest a feedback between the two rates, at the 5% level of significance. However, if 99% confidence bands (not reported) are used, the confidence bands contain zeros for all lags, for both cases. Hence, we conclude that there is no dynamic causality between the two rates at the 1% level of significance.

Paying attention to the impulse response functions to the own shocks, it can be seen that the US function is flat, relative to the others that decline to zero quickly. In this respect, the Japanese rate shows markedly different behaviours before and after the crisis. From Table 3, the half-life of the US rate is 49 months, while that of the Japanese rate is 0.52 month. The latter implies a dramatic decline of persistence after the crisis. The 95% confidence interval for the US half-life has the upper limit of infinity, while those of the others are finite and fairly small. This means that the US rate represents the common stochastic trend, and is the dominant player in this region after the crisis. The dominance and persistence of the Japanese rate were confined to Period I.

The impulse response analysis and half-life estimation indicate that the main driving forces in Period I are the US and Japanese rates. This suggests two common trends in Period I, in conflict with the outcome of the cointegration test given in Table 2. Since the VAR innovations in Period I show strong non-normality and heteroskedasticiy, it seems reasonable to assume that the results associated with the bias-corrected wild bootstrap are more robust. This is because Johansen's (1988) tests depend heavily on the assumption of i.i.d. normality. In Period II, one common trend identified from the half-life estimation is in agreement with the outcome of the cointegration tests.

4.3. Further Discussion

A consensus has, as mentioned earlier, emerged in the literature that the real interest rate linkages in the Pacific Basin region have changed over time. Our evidence for the dichotomous phenomenon over the Asian crisis with regard to Japan provides the existing literature with another piece of a puzzle. Chinn and Frankel (1995) presented evidence that Japan possessed considerable market influence in the 1980's in the Pacific Basin region, although they cautiously pointed out that some of the evidence is tentative, indicating that Japan's influence may have been overstated. Phylaktis (1999) also found evidence that Japanese financial influence increased in the region. Our findings suggest that Japan has lost a significant level of dominance over the course of the financial crisis.

Although there are differing views as to the causes and effects of the financial crisis, it is an undoubted fact that a recession in Japan, coupled with the fragile condition of Japanese financial institutions in early 1997, exacerbated poor economic fundamentals in East Asia and worsened the crisis. Japan has not yet shown any sign of notable recovery as of early 2005. Japan was in a subtly promising shape showing a positive growth rate in early 1996, but this was halted by a recession after an increase in the consumption tax, which failed to generate sufficient import demand in the region. Moreover, lured by larger but riskier opportunities outside Japan, Japanese banks started to lend heavily in East Asia, which resulted in a hard hit-back with huge capital losses during the crisis. As Corsetti *et al.* (1999) pointed out, this naturally creates a distinctive comparison to the role of the US in the Mexican crisis[7]. In this respect, our results may be an indication that, after the crisis, the weakness of Japan was reflected in financial markets in the region where Japan is considered even less attractive and reliable than it was, and international investors diversified their portfolios more actively than before the crisis by moving away from Japan. As a result, the role of the Japanese Yen may have diminished as an alternative debt-denominating currency to the US dollar.

5. Concluding remarks

This paper examines the short run and long run relationships among the real interest rates of several Pacific-Basin countries with a focus on East Asia, paying attention to the impact of the Asian financial crisis. We have used monthly data for the US, Japan, Korea, Singapore, and Thailand from 1980 to 2004. We are concerned with the degree of capital market integration and the nature of short run dynamics. To investigate these issues, we adopt Johansen's (1988) cointegration test and impulse response analysis based on the unrestricted VAR model. We also have estimated the half-lives of the real interest rates to measure the persistence of real interest rates. For statistical

[7] A more comprehensive discussion on the causes and effects of the financial crisis can be found in Corsetti et al. (1999).

inference on impulse response analysis and half-life estimation, we use the bias-corrected (wild) bootstrap.

It is found that a selected group of capital markets in the Pacific-Basin region are highly integrated. The degree of integration has become stronger after the crisis in 1997. Rich dynamic interactions are observed from the impulse response analysis. Before the crisis, the US rate affects the Singaporean and Thai rates, while the Japanese rate affects the Korean rate. After the crisis, the US rate affects the Korean and Singaporean rates, while the Singaporean rate affects the Thai rate. The half-life estimation also reveals interesting features in relation to persistence in the real interest rates. Before the crisis, the US and Japanese rates show the degree of persistence equivalent to a unit root time series, while the others show strong mean-reversion. After the crisis, however, the US rate is the only time series that shows the degree of persistence of a unit root time series, while the others including the Japanese rate are highly mean-reverting. This indicates that the US and Japan were the two dominant capital markets in this region before the crisis, while the US capital market dominates the region after the crisis.

On a methodological note, this paper is distinct from past studies in the following aspects. First, our analysis is based on both point and interval estimates of impulse responses, in contrast with Phylaktis (1999) where only point estimates are analyzed. Second, we have estimated half-lives of real interest rates in a multivariate setting, while the previous studies, such as Rapach and Wohar (2004), are based exclusively on univariate methods. That is, our half-life estimates are obtained in a more general setting and are possibly more accurate. In combination with the cointegration test results, this also enables us to identify which real interest rates represent the common trends and thus dominate the others.

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Table 1. Residual Diagnostics from VAR estimation

	US	JAP	KR	SN	TH
Period I					
Normality	727.12*	29.58*	7.21*	47.45*	6.08*
ARCH	45.08*	7.64	7.84	17.84*	3.92
AUTO	16.27*	6.71	2.42	9.93	2.91
HETERO	72.65*	59.20*	25.59	47.37*	26.68
Adjusted R^2	0.94	0.96	0.87	0.91	0.82
Period II					
Normality	1.67	1.01	1.55	1.04	3.73
ARCH	3.20	2.02	3.82	3.33	6.45
AUTO	4.16	1.52	9.19	6.22	3.21
HETERO	11.72	16.51	22.16	23.47	12.13
Adjusted R^2	0.97	0.20	0.70	0.82	0.59

"*" indicates rejection of the null hypothesis at 10% level.
VAR order 2 is chosen for both periods
Normality is the Jarque-Bera test for the normality of residuals
ARCH is the Lagrange multiplier test for ARCH(6) model applied to residuals
AUTO is the Ljung-Box test for no serial correlation applied to the residuals with lag 6
HETERO indicates the White's heteroskedasticity test (no cross product terms) in each VAR equation. The squared residuals in each equation are regressed against the right-hand variables and their squares. The test statistic follows chi-squared with 20 degrees of freedom.

Table 2. Johansen's cointegration test results

	Period I		Period II	
H_0	λ_{max}	λ_{trace}	λ_{max}	λ_{trace}
$r = 0$	46.74*	99.30*	48.21*	133.58*
$r \leq 1$	30.82*	52.56*	42.52*	85.37*
$r \leq 2$	11.61	21.73	25.31*	42.84*
$r \leq 3$	9.37	10.12	16.96*	17.53*
$r \leq 4$	0.74	0.74	0.57	0.57

"*" indicates the rejection of the null hypothesis at 5% level.
The results are based on VAR(2) model in level or VEC(1) model, assuming restricted intercept and no trends in VAR
λ_{max}: Johansen's maximum eigenvalue statistic; λ_{trace}: Johansen's trace statistic

Table 3. Half-life estimates

	Period I		Period II	
	Point Estimate	Interval Estimate	Point Estimate	Interval Estimate
US	12.49	(3.62, ∞)	49.01	(4.31, ∞)
JP	20.23	(4.60, ∞)	0.52	(0.40, 0.76)
KR	6.26	(2.27, 15.17)	0.60	(0.43, 0.87)
SN	1.87	(0.89, 6.82)	0.69	(0.50, 1.11)
TH	2.40	(1.63, 3.81)	0.86	(0.57, 2.30)

The entries are the number of months. Interval estimates are 95% confidence intervals.

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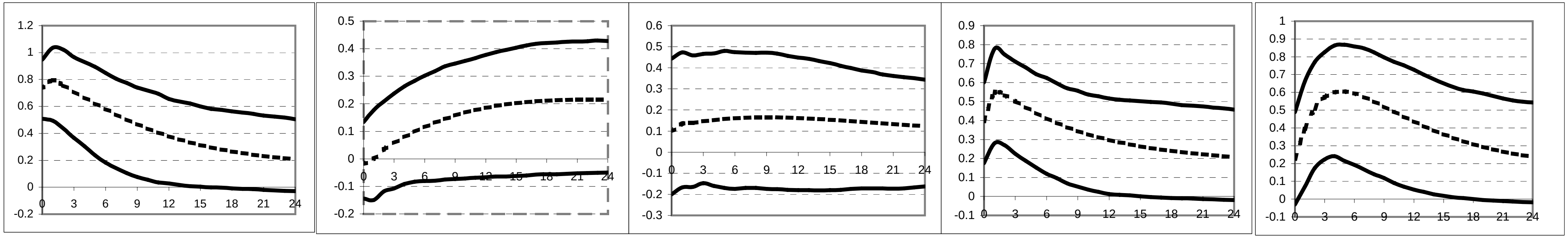
Figure 1. Impulse response functions and 95% confidence bands (Period I)

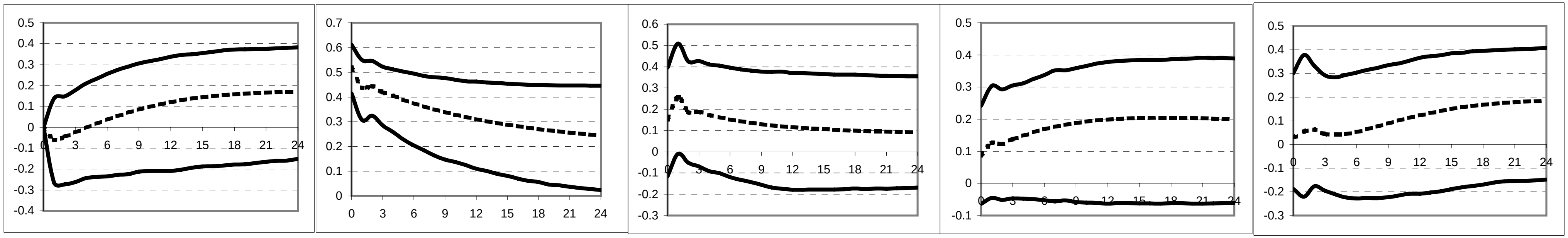
Shock to US

US JP KR SN TH



Shock to JP

US JP KR SN TH



Shock to KR

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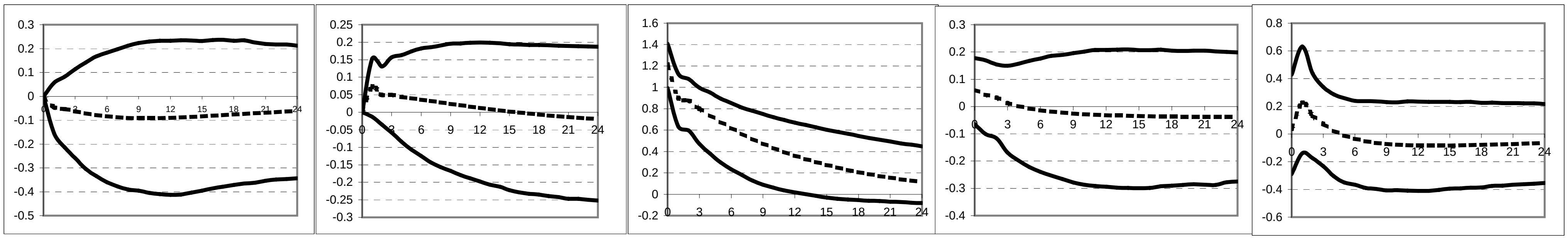

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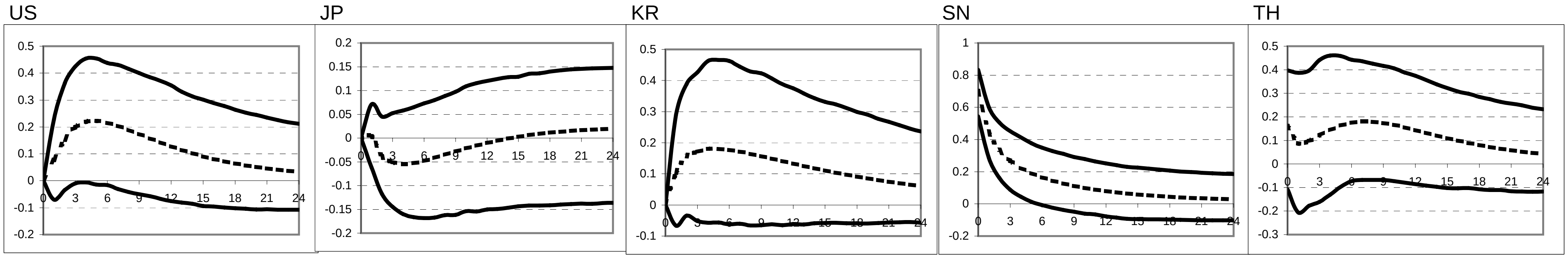

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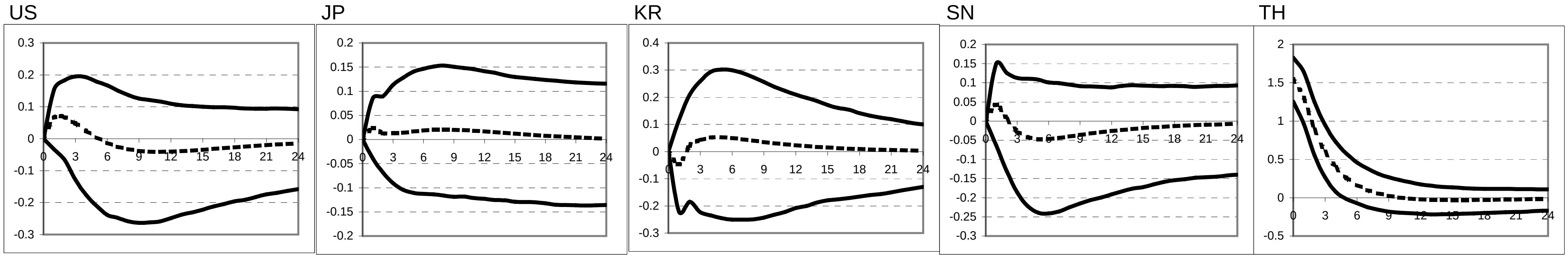
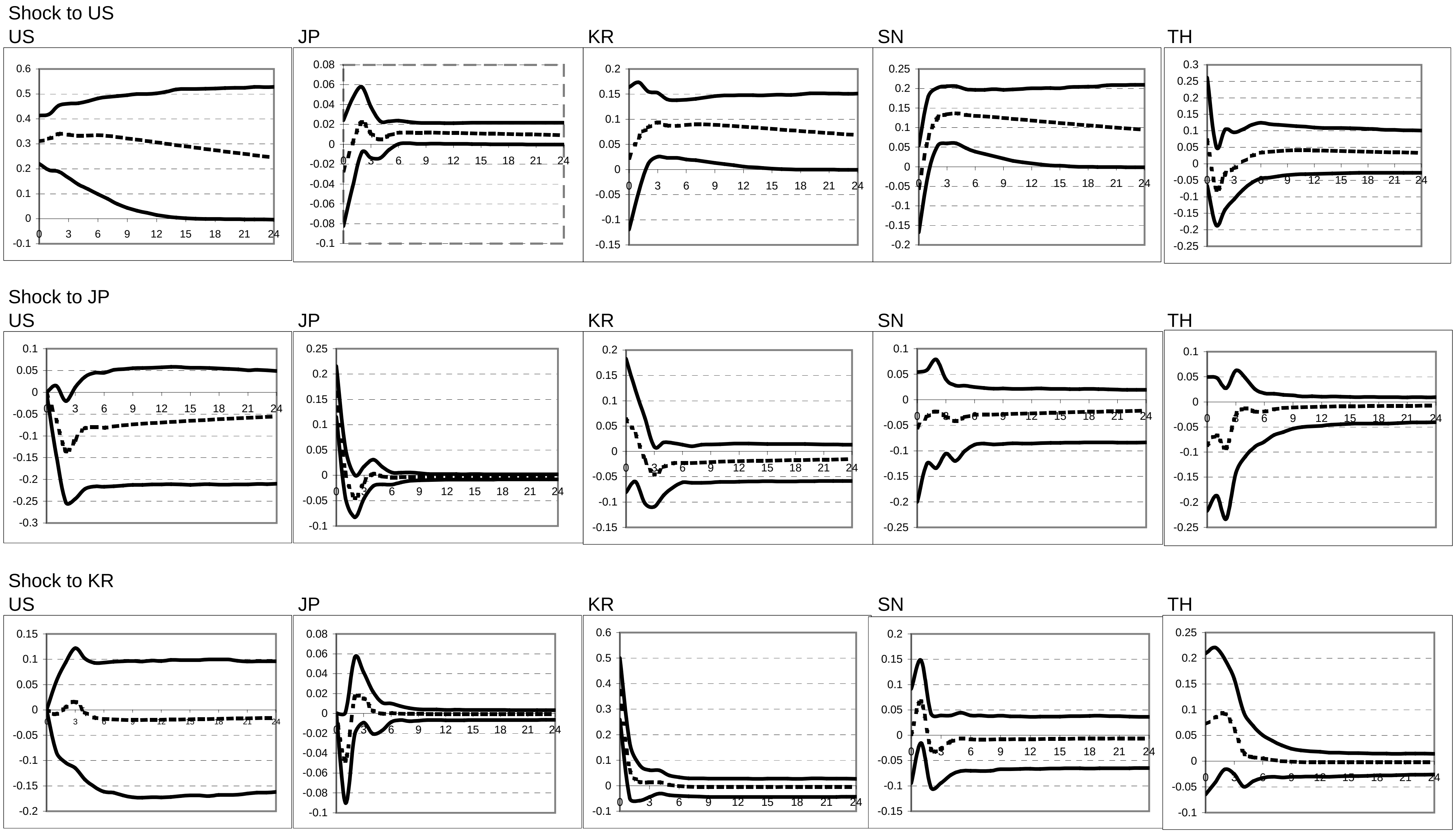


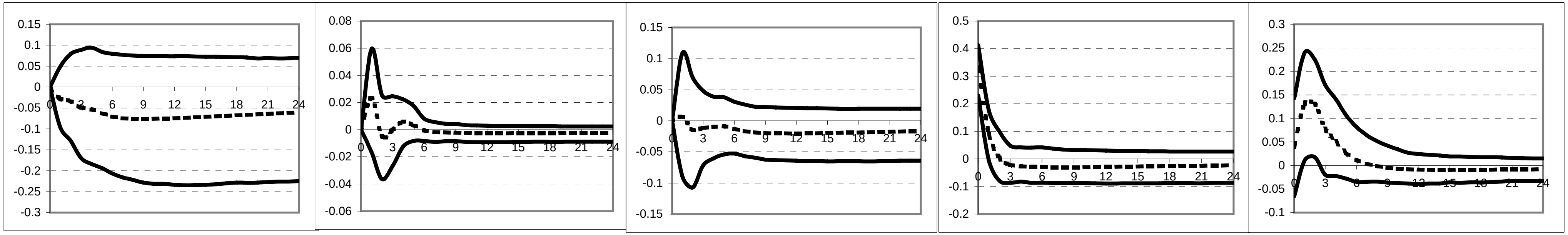
Each graph plots the responses over period 0 to 24.
Confidence intervals are calculated using the bias-corrected wild bootstrap

Figure 2. Impulse response functions and 95% confidence bands (Period II)



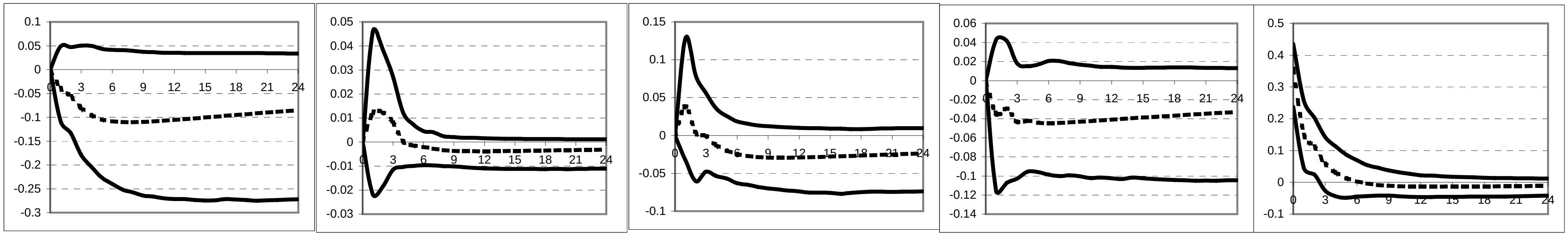
Shock to SN

US JP KR SN TH



Shock to TH

US JP KR SN TH



Each graph plots the responses over period 0 to 24.
Confidence intervals are calculated using the bias-corrected wild bootstrap