International linkage of real interest rates: the case of East Asian countries

Philip I. Ji
Discipline of Economics
University of Sydney
Email: P.Ji@econ.usyd.edu.au

Jae H. Kim
Department of Econometrics and Business Statistics
Monash University, Australia
Email: Jae.Kim@buseco.monash.edu.au

Abstract

This paper examines linkage of real interest rates for a group of selected countries in East Asia. The countries under study include Japan, Korea, Singapore, Malaysia and Thailand. The long run relationship is tested and estimated using the conitegration analysis. We also have conducted the impulse response analysis based on unrestricted vector autoregression, using the bias-corrected wild bootstrap for statistical inference. Our results show that (1) there exists a long run equilibrium relationship, (2) there are interesting short run dynamic interactions, in which Singapore, Malaysia and Thailand play the role of equilibrating factors.

Key Words: Financial linkage; Real interest rate parity; Cointegration analysis; Wild bootstrap

JEL classifications: F36; E44

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 for this hypothesis used the conventional regression technique, but the results were overwhelmingly against the real interest rate equalization (see, for example, 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 the real interest rate equalization rejected in the regression context, even when the capital markets are efficient and fully integrated. Moreover, as Goodwin and Grennes (1994) pointed out, statistical inference based on the conventional regression technique may not be valid, when the 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 the long run equilibrium among the real interest rates should have strong implications to the interest parity and efficiently integrated markets. They suggested the use of the cointegration analysis (Engle and Granger; 1987; and Johansen; 1988) as an alternative, since it provides a suitable framework to test and estimate long run equilibrium relationships. Their cointegration analysis revealed strong evidence of the interest parity and market integration among a number of countries. Subsequent studies in this area by Hutchison and Singh (1997), Phylaktis (1999) and Yamada (2002a, 2002b) have also used the cointegration analysis and identified strong linkages among the real interest rates. These studies have found high degrees of

capital market integration, although the condition for the real interest rate equalization hypothesis does not hold in general.

In this paper, we pay attention to the case of East Asian countries. Most of the countries in the region are rapidly growing economies whose capital markets are at immature stages. It should be noted that the degrees of efficiency and integration of their capital markets have not been extensively investigated, as the above-mentioned past studies largely neglected the case of East Asian countries. To the best of our knowledge, Phylaktis (1999) is the only study where a number of East Asian countries are examined as a part of the Pacific Basin countries. Note that Phylaktis (1999) used the data from early seventies to early nineties. In this paper, we follow the recent studies in adopting the cointegration analysis using the data from 1980 to 2002. In addition, we conduct the impulse response analysis based on the vector autoregression (VAR) to examine the dynamic interactions among the real interest rates.

For statistical inference for impulse response analysis, we resort to bias-corrected confidence intervals based on the bootstrap method. Given the strong evidence of unit-root non-stationarity of the real interest rates, the conventional Granger non-causality test or statistical inference for impulse response analysis may show deficient small sample properties. Since this small sample deficiency is caused in large part by biases of VAR parameter estimators, we use the bias-corrected bootstrap confidence interval of Kilian (1998), which has been found to exhibit much better small sample properties than the conventional confidence intervals. The bootstrap (Efron, 1979) is a computer-intensive method of approximating the sampling distribution of a statistic. It has been applied widely in statistics and econometrics, and often found to provide a

2

superior alternative to the conventional methods in small samples (see, for example, Li and Maddala, 1996; and Berkowitz and Kilian, 2000). The conventional bootstrap, however, is applicable to the data generated from an iid distribution. Similarly, Kilian's (1998) bias-corrected bootstrap is applicable to the VAR model whose innovations follow an iid distribution. This may not work properly when the VAR model shows conditionally heteroskedastic error terms, which is the case for our VAR model (see Section 4). Recently, a bootstrap procedure called the wild bootstrap (see, for example, Mammen, 1993) has been developed, which is applicable to time series with conditional or unconditional heteroskedasticity of unknown form. In this paper, we use the wild bootstrap to construct bias-corrected bootstrap confidence interval of Kilian (1998). The theoretical underpinning of the wild bootstrap in the context of univariate AR model can be found in Gonclaves and Kilian (2004).

The main finding of the paper is that the real interest rates of the East Asian countries are closely related in the long run and short run. We have found one long run relationship, which implies a weak degree of market integration in this region. We also have found interesting dynamic short run interactions among the real interest rates, whereby Singaporean, Malaysian and Thai rates playing the role of equilibrating factors towards the long run equilibrium. Further interesting causal relationships are also observed using the impulse response analysis. In the next section, 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 used monthly time series from 1980:1 to 2003:3, which includes 267 observations. The starting date reflects the timing of deregulation where the most of Asian countries started to open their financial markets. We have selected the real interest rates of five countries, which include Japan, Korea, Singapore, Malaysia and Thailand. This choice is based on the consideration that VAR dimension should be kept manageable for parsimonious parameterization. We also feel that these countries represent a good mixture of developed and developing countries in the region, with diverse characteristics and different degrees in the maturity of capital markets.

We use short-term interest rates for these countries. Monthly money market rate is used for Korea, the Philippines and Thailand, and T-bill rate for Malaysia. 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 (CPI) is used. The CPI is seasonally adjusted using the X-12 method. All nominal interest rates are then deflated by inflation in order to generate 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 trend. On this basis, we decide not to include linear time trend in our testing and estimation below, because there is little evidence for the presence of linear trend. As argued by Yamada (2002a; p.280), this can provide more reliable empirical results. To determine whether the real interest rates series possess

unit-root, we conducted the augmented Dickey-Fuller tests[1]. As reported in Table 1, all real interest rates appear to be integrated of order 1 process at 5% significance level. This casts the possibility of cointegration among these rates and the existence of long-run relationships.

We have conducted pairwise and multiple cointegration testing using the Engel-Granger (1987) methodology. Although the cointegrating relationship has been found in many cases, it is found that the condition of the 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 results. 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

VAR Model and Cointegration

We consider the vector autoregressive (VAR) model of the form

$$Y_t = v + B_1 Y_{t-1} + \ldots + B_p Y_{t-p} + u_t, \tag{1}$$

where Y_t is the $K \times 1$ vector of variables at time t, v is the $K \times 1$ vector of intercepts, and B_i's are the $K \times K$ matrices of coefficients. Note that u_t is the $K \times 1$ vector of innovations with $E(u_t) = 0$ and $E(u_t u_t') = \Sigma_u = PP'$. The above VAR system can be written in the vector error correction (VEC) form as

$$\Delta Y_t = v + \Gamma_1 \Delta Y_{t-1} + \ldots + \Gamma_{p-1} \Delta Y_{t-p+1} + \Pi Y_{t-1} + u_t, \tag{2}$$

---

[1] The presence of a unit root in real interest rate may be arguable. However, as Goodwin and Grennes (1994; footnote 5) pointed out, its justification can be found from past empirical evidence and practical considerations. See also Shea (1992) for a similar argument. In addition, there is growing evidence that the real interest rate can be modelled with non-linear time series model. However, this possibility is not considered here and left as a future research.

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

The unknown VAR order p in (1) is estimated to ensure that the residuals of each equation in VAR mimic a white noise. We employ a simple to general approach for parsimonious parameterisation. To this end, visual inspection of residual autocorrelation function is conducted, in addition to the use of 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 methods are not presented in this paper, because they are well documented elsewhere (see, for example, Lütkepohl, 1991; Chapter 11; Hamilton; 1994; Chapter 20).

Impulse response analysis

We examine how the real interest rates are inter-related over time by resorting to the orthogonalized impulse response analysis (see, for details, Lütkepohl, 1991). It is a dynamic multiplier analysis among the variables in the VAR system, measuring how a standard deviation shock to a variable in the system is transmitted to others over time. It is particularly useful as a means of examining the short-run dynamic interactions among the variables in the VAR system, and has been applied widely to empirical macroeconomics and international finance (see, for example, Eichenbaum

and Evans, 1995). It is also closely related to the causality analysis, as zero impulse responses between two variables mean no causality (Lütkepohl, 1991; p.45).

The orthogonalized impulse responses are calculated from the coefficients of the MA(∞) representation of the VAR model and residual covariance matrix. In conducting the impulse response analysis, it is important to test whether impulse response estimates are statistically different from 0. The conventional method of statistical inference for impulse responses is based on asymptotic methods (see, for example, Lütkepohl; 1991). These asymptotic methods can perform poorly when the sample size is small, or when the VAR model is subject to a non-normal innovation process (see, for details, Kilian, 1998). As a result, the asymptotic method can lead to misleading inferential outcomes for the statistical significance of impulse response estimates. In addition, impulse response estimates are necessarily biased in small samples, due to small sample biases present in 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.

Given n realizations (Y_1, …, Y_n) of (1), the unknown coefficients are estimated using the least-squares (LS) method. The LS estimators for $B = (v, B_1, …, B_p)$ and Σ_u are denoted as $\hat{B} = (\hat{v}, \hat{B}_1, …, \hat{B}_p)$ and $\hat{\Sigma}_u$, and the associated 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 period ago. Using $\hat{B}$ and $\hat{\Sigma}_u$, the estimator for impulse response $\hat{\theta}_{kl,i}$ for $\theta_{kl,i}$, can be calculated. For bias-correction of VAR parameter estimates, we follow Kilian (1998) to use Pope's (1990) bias formula to obtain bias-corrected parameter estimators, and they are used for the bias-corrected bootstrap. Note that Pope's (1990) formula estimates bias to the order of n^{-1}, and is applicable to VAR model with martingale difference innovations with a fixed covariance matrix, which includes conditionally heteroskedastic errors as special cases. Let the bias-corrected estimators for B be $\hat{B}^c = (\hat{v}^c, B_1^c, ..., \hat{B}_p^c)$. Further details of bias-correction can be found in Kilian (1998).

The bias-corrected confidence interval based on the wild bootstrap for $\theta_{kl,i}$ can be outlined as below:

In Stage 1, generate pseudo data set following the recursion

$$Y_t^* = \hat{v}^c + B_1^c Y_{t-1}^* + ... + B_p^c Y_{t-p}^* + u_t^*, \tag{3}$$

using the first p values of the original data as starting values. The wild bootstrap 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. Note that the procedure we adopt here is different from Kilian's (1998) on this point. With Kilian's (1998) procedure where iid innovations are assumed, u_t^*'s are generated as random resampling of $\hat{u}_t$'s with replacement. The distinct feature of the wild bootstrap is that u_t^*'s are generated as a random weighting of $\hat{u}_t$'s, so that $E(u_t^* | \hat{u}_t) = 0$ and $E(u_t^* u_t^{*\prime} | \hat{u}_t) = u_t \hat{u}_t'$.

In Stage 2, using $\{Y_t^*\}_{t=1}^n$, the VAR coefficient matrices are re-estimated and denoted as $\hat{B}^* = (\hat{v}^*, \hat{B}_1^*, ..., \hat{B}_p^*)$. Pope's (1990) bias formula is then applied to $\hat{B}^*$ in order to obtain a bias-corrected version $\hat{B}^{*c} = (\hat{v}^{*c}, B_1^{*c}, ..., \hat{B}_p^{*c})$ of $\hat{B}^*$. Repeat Stages 1 and 2 sufficiently many times, say m, to generate bootstrap replicates of $\{\hat{B}^{*c}(j)\}_{j=1}^m$, from which B bootstrap replicates $\hat{\theta}_{kl,j}^*$ of impulse responses are obtained.

The $100(1-2\alpha)\%$ bias-corrected bootstrap confidence intervals for $\theta_{kl,i}$ can be constructed as the interval $[\hat{\theta}_{kl,j}^*(\alpha), \hat{\theta}_{kl,j}^*(1-\alpha)]$, where $\hat{\theta}_{kl,j}^*(q)$ is the qth percentile from the distribution of m bootstrap replicates $\{\hat{\theta}_{kl,j}^*\}_{j=1}^m$, based on the percentile method of Efron and Tibshirani (1993, p.160). 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.

4. Empirical Results

In conducting the structural analysis based on the VAR, the ordering of the variables in the VAR system is important. This is closely related to identifying restrictions to the shocks in the VAR system. In this paper, we follow the ordering according to the Wold-causality (see, Lütkepohl, 1990; p.52). That is, we place Japanese real interest rate first, followed by Korean, Singaporean, Malaysian, and Thai real interest rates. In the context of orthogonalized impulse response analysis, this amounts to assuming the instantaneous causality runs one way from Japanese to Thai rates. This seems reasonable, considering the relative power and scale of the economies of these countries in East Asia.

Table 2 reports the estimation results for the VAR model. We have chosen the VAR order 2, and this choice is found to be statistically adequate with the residuals from all equations mimicking white noise with an exception of the equation for Thailand. This choice is further justified statistically, as it is the order also preferred by AIC. A notable feature of the estimated VAR(2) model is that the residuals show strong evidence of non-normality, according to the Jarque-Bera test for normality. In addition, the ARCH LM test reveals strong evidence of conditionally hetroskedastic errors.

Table 3 reports the tests for cointegration rank. Both trace and maximal eignevalue tests indicate that the VAR model is cointegrated with cointegrating rank of 1, at the level of significance 5%. That is, we have found one long run equilibrium relationship, which implies the existence of four common trends. This indicates a low degree of market integration, which seems reasonable given the immaturity and diversity of the markets in this region. The coefficients of the VEC model (2) are estimated, and the estimates of the cointegrating vector are presented in Table 4. The estimated cointegration vector reported in Table 3 indicates that, in the long run, the Japanese rate is negatively related to the Korean and Thai rates, but positively with the Malaysian and Singaporean rates. According to the LR test conducted on the coefficients, all long run coefficients are different from zero statistically. Moreover, we cannot reject the hypothesis that the long run coefficients associated with Japanese and Korean rates are equal; and the hypothesis that those of Singaporean and Malaysian rates are equal. This means that the real interest rates are moving together in the long run, with some of the rates exhibiting close relationships. However, we do

not find any evidence in support of the real interest rate equalisation hypothesis among these countries.

Table 5 reports the results of short run coefficients of the VEC model. It appears that the real interest rates of all countries are closely related, with rich dynamic interactions in the short run. The speed-of-adjustment coefficients are statistically different from zero only in the equations of Singaporean, Malaysian and Thai relates. This means that these rates act as equilibrating factors in the system, while Japanese and Korean rates do not respond to the past equilibrium error. One may conduct the Granger non-causality tests based on the VEC model. However, given the strong evidence of non-normal and conditionally heteroskedastic errors, the conventional Wald-type test may show deficient properties. On this basis, we resort to the impulse response analysis based on the wild bootstrap.

Figure 1 presents dynamic responses to a positive one-deviation shock to a real interest rate, along with the 90% and 95% confidence intervals. If the confidence interval contains zero, the null hypothesis that the true value of response is zero cannot be rejected at the specified level of significance. For all countries, it is plausible to observe that the real interest depends on its own past, with positive impact lasting more than 12 months. To a shock to Japanese rate, only the Singaporean rate responds positively over a period of 3 months to 11 months, at the level of significance 10%. There is no evidence that the other rates respond to a change in the Japanese rate. To a shock to Korean rate, only the Malaysian and Thai rates react positively. The Malaysian rate shows longer-term responses than the Thai rates. To a shock to Singaporean rate, only the Thai rate responds positively at 10% level of

significance, with a short-term effect. There is also evidence that Malaysian rate affect Singaporean rate negatively for a period from 3 to 12 months. When there is a shock to the Thai rate, only the Singaporean and Malaysian rates react positively with similar dynamic pattern.

The overall summary of causal relationships is illustrated in Figure 2. It can be seen that the Japanese and Korean rates are not related over time in the short run. The rates of these countries affect the rates of the other countries dynamically, but there is no causality running from the rates of the other countries to those of Korea and Japan. The Singaporean, Malaysian and Thai rates are inter-related dynamically. There are one-way causal relationships running from Thai to Malaysian rate and from Malaysian to Singaporean rate, while a feedback system is present between Singaporean and Thai rates.

5. Concluding remarks

This paper examines the long run and short run relationships of the real interest rates of several East Asian countries. We have used monthly data for Japan, Korea, Singapore, Malaysia and Thailand from 1980 to 2003. From the cointegration analysis, we have found that the real interest rates of these countries are closely related in the long run. We have found one cointegration vector, which indicates a low degree of market integration.

Rich dynamic short run interactions have been found from the error correction model estimation, with Singaporean, Malaysian and Thailand playing the roles of equilibrating factors towards the long run equilibrium. Further Interesting dynamic

relationships have been revealed using the impulse response analysis. Korean and Japanese rates affect the others over time, but the others show no dynamic influence to Korean and Japanese rates over time. In addition, there is no evidence of short run dynamic interactions between the Korean and Japanese rates.

Although we have not found any direct evidence of the real interest rate equalisation, this study has found that the real interest rates of East Asian countries are closely related in the long run and short run. These features should have strong implications to the interest parity and efficiently integrated capital markets in this region.

Table 1. ADF test results

	No intercept and no trend		Intercept with no trend	
	Level	First Difference	Level	First Difference
JAP	-1.65 (1)	-11.86* (1)	-1.62 (2)	-11.95* (1)
KOR	-1.76 (7)	-7.69* (6)	-2.44 (7)	-7.77* (6)
SIN	-1.83 (0)	-5.51* (12)	-2.42 (13)	-5.70* (12)
MAL	-0.72 (3)	-11.13* (2)	-2.44 (1)	-11.11* (2)
THA	-1.25 (6)	-8.31* (5)	-2.31 (9)	-8.32* (5)

"*" indicates the rejection of the null hypothesis at 5% level.
The numbers in the bracket are the orders of the ADF regression to ensure no serial correlation in residuals

Table 2. VAR Model Estimation Results

	JAP	KOR	SIN	MAL	THA
JAP(-1)	0.802*	0.138	0.093	-0.052	-0.012
JAP(-2)	0.173*	-0.106	-0.016	0.057	-0.019
KOR(-1)	0.041*	0.935*	-0.015	0.042	0.388*
KOR(-2)	-0.039	-0.035	0.032	-0.001	-0.308*
SIN(-1)	-0.006	0.206*	0.835*	-0.047	0.323*
SIN(-2)	0.013	-0.199*	-0.069	-0.043	-0.149
MAL(-1)	0.091	-0.053	-0.001	0.679*	-0.175
MAL(-2)	-0.119*	0.055	-0.174*	0.161*	0.188
THA(-1)	0.000	0.004	0.023	0.032*	0.645*
THA(-2)	0.007	0.043	0.051*	0.009	0.179*
Intercept	0.059	0.490	0.727	0.288	-0.069
Diagnostic Tests					
Normality	118.73*	287.74*	191.68*	190.8*	192.5*
ARCH	16.16*	20.72*	15.54*	13.21*	44.03*
Auto	2.06	3.91	5.81	2.34	9.00*
Adjusted R^2	0.97	0.91	0.91	0.88	0.85

"*" indicates the significance of the coefficients (or rejection of the null hypothesis) at 10% level.
VAR order 2 is chosen using AIC.
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

Table 3. Johansen's cointegration test results

Null hypothesis	Maximal Eigenvalue	Null hypothesis	Trace
r = 0	55.16*	r = 0	101.24*
r ≤ 1	22.65	r ≤ 1	46.07
r ≤ 2	13.29	r ≤ 2	23.43
r ≤ 3	7.11	r ≤ 3	10.14
r ≤ 4	3.02	r ≤ 4	3.02

"*" indicates the rejection of the null hypothesis at 5% level.
The results are based on VAR(2) model, assuming restricted intercept and no trends in VAR

Table 4. Cointegration vector estimate and test results

Cointegrating Vector Estimate	
(a1, a2, a3, a4, a5, a6) = (1, 0.86, -4.93, -5.02, 2.10, 13.28)	

Testing the restriction on the cointegrating vector

Null hypothesis	Test Statistic
a1 = 0	4.08*
a2 = 0	3.65*
a3 = 0	31.33*
a4 = 0	24.71*
a5 = 0	22.64*
a1 = a2	0.04
a3 = a4	0.01

"*" indicates the rejection of the null hypothesis at 5% level.
(a1, a2, a3, a4, a5, a6) represent the coefficients of cointegrating vectors associated with Japanese, Korean, Singaporean, Malaysian, Thai real interests, plus intercept.
The likelihood ratio test results are given, which asymptotically follows the chi-squared distribution with the degree of freedom one for all cases.

Table 5. Parameter estimates of the error correction models

	ΔJAP	ΔKOR	ΔSIN	ΔMAL	ΔTHA
ΔJAP(-1)	-0.18*	0.12	0.03	-0.059	0.032
ΔKOR(-1)	0.04*	-0.02	-0.05	0.018	0.388*
ΔSIN(-1)	0.00	0.23*	0.05	0.053	0.137
ΔMAL(-1)	0.11*	-0.08	0.16*	-0.170*	-0.295
ΔTHA(-1)	0.00	-0.03	-0.06*	-0.010	-0.214*
ECM(-1)	0.24	0.29	4.25*	2.612*	-4.128*

"*" indicates the significance of the coefficients at 10% level.
$\Delta \equiv 1-B$, where B is the lag operator.
ECM is the error correction term calculated from the cointegrating vector.

References

ABADIR, K. M., K. HADRI, and E. TZAVALIS (1999): "The Influence of VAR Dimensions on Estimator Biases," *Econometrica*, 67, 163-181.

BERKOWITZ, J., and L. KILIAN (2000): "Recent Developments in Bootstrapping Time Series," *Econometric Reviews*, 19, 1-48.

CUMBY, R. E. and F. MISHKIN (1986) The international linkage of real interest rates: The European±US connection, *Journal of International Money and Finance*, 5, 5-23.

EFRON, B. (1979): "Bootstrap Methods: Another Look at the Jackknife," *Annals of Statistics*, 7, 1-26.

EFRON, B., and R. J. TIBSHIRANI (1993): *An Introduction to the Bootstrap*. New York: Chapman & Hall.

ENGLE, R. E. and Granger, C. W. J. (1987) Cointegration and Error-Correction: Representation, Estimation, and Testing. *Econometrica* 55, 251-76.

EICHENBAUM, M., and C. L. EVANS (1995): "Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rates," *Quarterly Journal of Economics*, 110, 975-1009.

GONCALVES, S., and L. KILIAN (2004): "Bootstrapping Autoregressive with Conditional Heteroskedasticity of Unknown Form," *Journal of Econometrics*.

GOODWIN, B. K. and GRENNES, T. J. (1994) Real interest rate equalization and the integration of international Financial markets, *Journal of International Money and Finance*, 13, 107-24.

HAMILTON, J. D. (1994): *Time Series Analysis*. Princeton: Princeton University Press.

HUTCHISON, M. M. and SINGH, N. (1997) Equilibrium real interest rate linkages: The United States and Japan, *Journal of the Japanese and International Economies*, 11, 208-27.

JOHANSEN, S. (1988): "Statistical Analysis of Cointegrating Vectors," *Journal of Economic Dynamics and Control*, 12, 231-254.

JOHANSEN, S. (1991) Estimation and hypothesis testing of cointegration vectors in Gaussian autoregressive models, *Econometrica*, 59, 1551-80.

JOHANSEN, S. (1995) *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford University Press, Oxford.

KILIAN, L. (1998): "Confidence Intervals for Impulse Responses under Departure from Normality," *Econometric Reviews*, 17, 1-29.

LI, H., and G. S. MADDALA (1996): "Bootstraping Time Series Models," *Econometric Reviews*, 15, 115-158.

LÜTKEPOHL, H. (1991): *Introduction to Multiple Time Series Analysis*. Berlin: Springer-Verlag.

MAMMEN, E. (1993): "Bootstrap and Wild Bootstrap for High Dimensional Linear Models," *The Annals of Statistics*, 21, 255-285.

MARK, N. N. (1985) Some evidence on international inequality of real interest rates, *Journal of International Money and Finance*, 4, 189-208.

MERRICK, J. J. and Saunders, A. (1986) International expected real interest rates: New tests of the parity hypothesis and US Fiscal-policy effects, *Journal of Monetary Economics*, 18, 313-22.

MISHKIN, F. (1984) Are real interest rates equal across countries? An empirical investigation of international parity conditions, *Journal of Finance*, 39, 1345-57.

NICHOLLS, D. F., and A. L. POPE (1988): "Bias in Estimation of Multivariate Autoregression," *Australian Journal of Statistics*, 30A, 296-309.

POPE, A. L. (1990): "Biases of Estimators in Multivariate Non-Gaussian Autoregressions," *Journal of Time Series Analysis*, 11, 249-258.

STOCK, J. (1987) Asymptotic Properties of Least-Squares Estimators of Cointegrating Vectors, *Econometrica*, 55, 1035-56

TJOSTHEIM, D., and J. PAULSEN (1983): "Bias of Some Commonly-Used Time Series Estimates," *Biometrika*, 70, 389-399.

YAMADA, H. (2002a): "On the Linkage of Real Interest Rates between the US and Canada: Some Additional Empirical Evidence," *Journal of International Financial Markets, Institutions and Money*, 12, 279-289.

YAMADA, H. (2002b): "Real Interest Rate Equalization: Some Empirical Evidence from the Three Major World Financial Markets," *Applied Economics*, 34, 2069-2073.

Figure 1. Impulse respons estimates and confidence bands
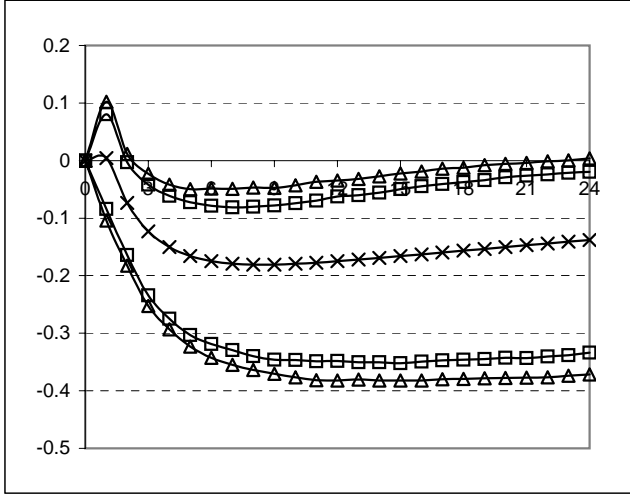
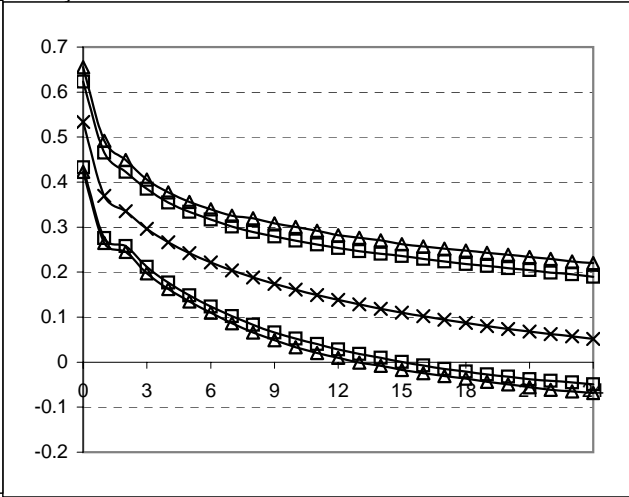Shock to Japan

Japan
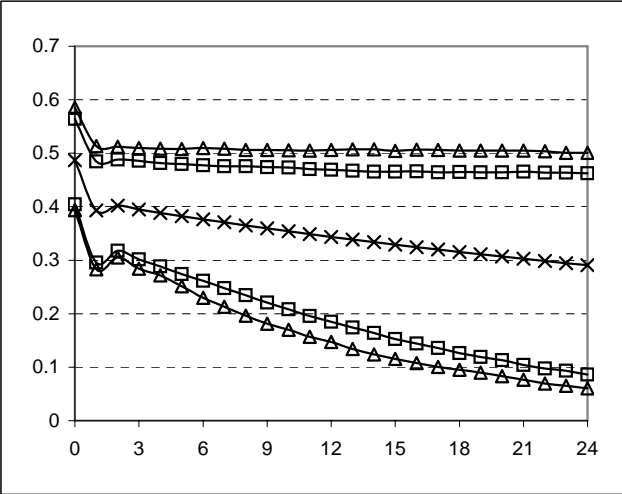
Korea
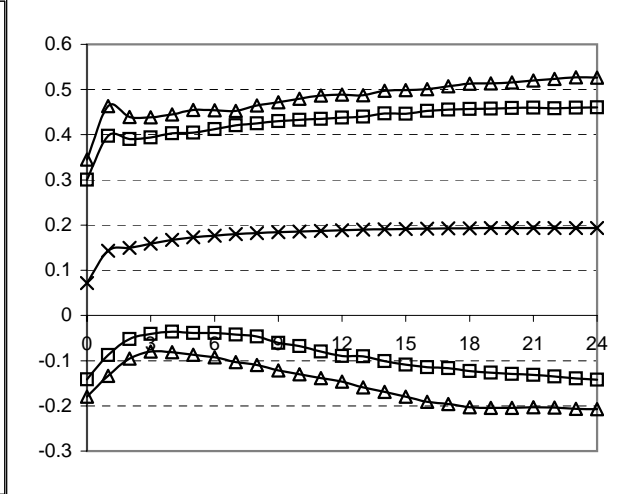
Singapore
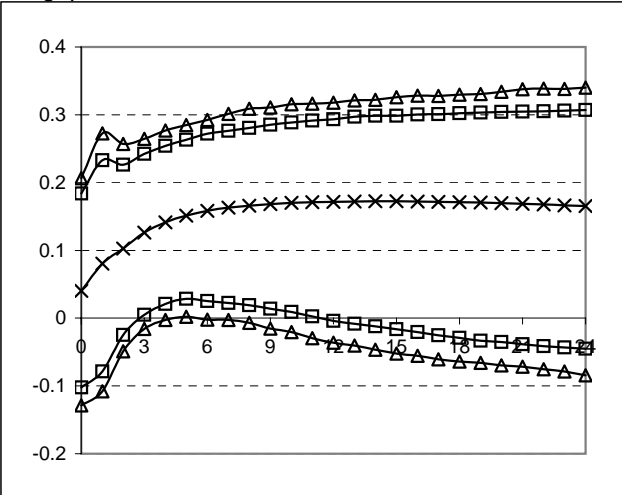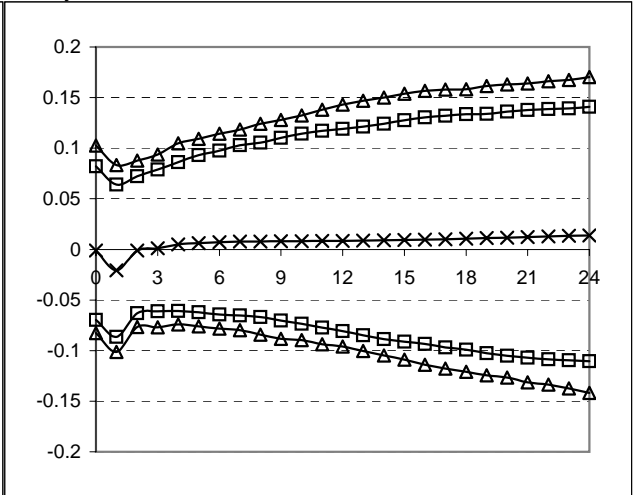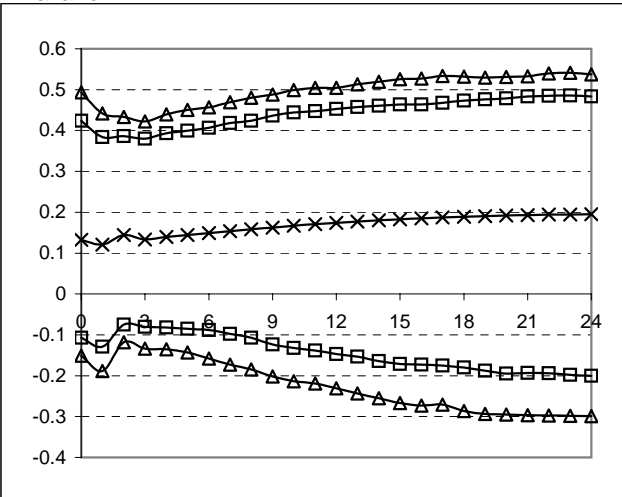
Malaysia

Thailand

Cross: Impulse response estimate; Square: 90% confidence band; Triangle: 95% confidence band

Figure 1. Continued.
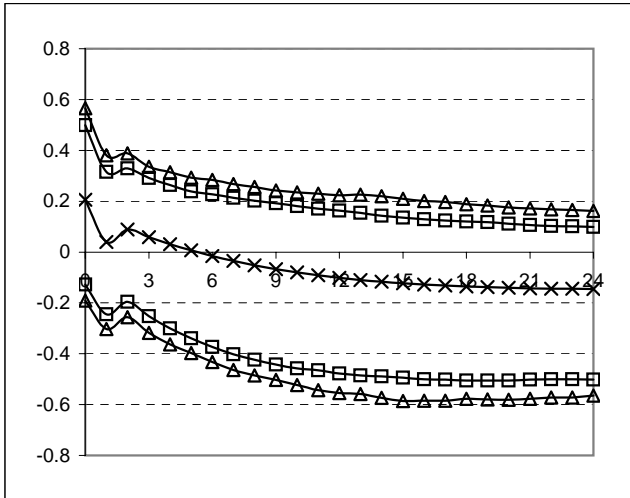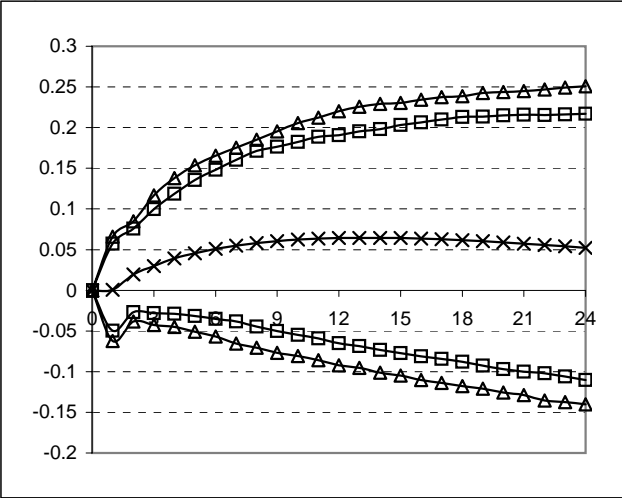
Shock to Korea

Japan
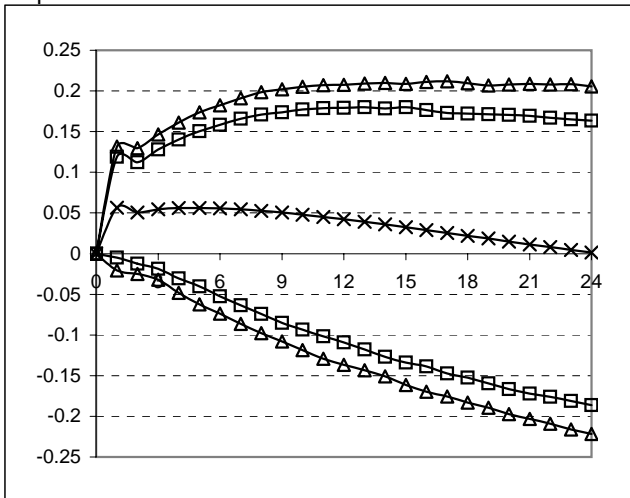
Korea
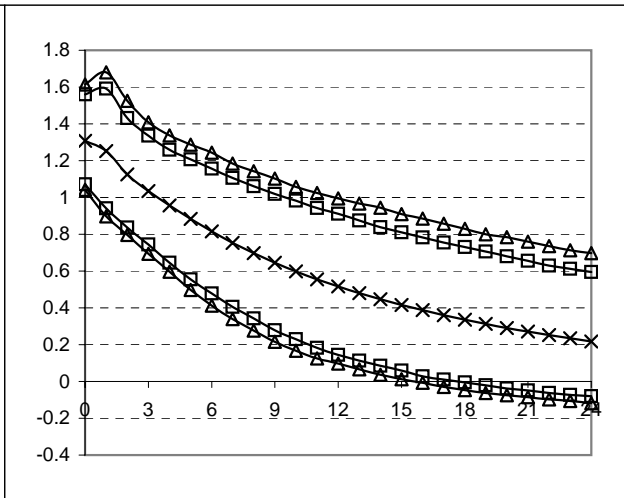
Singapore
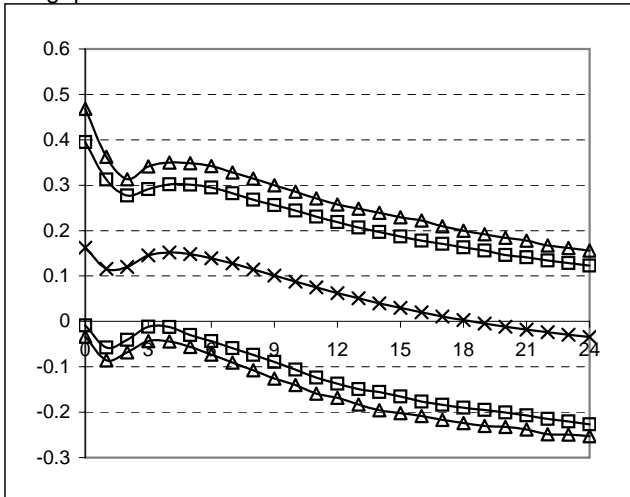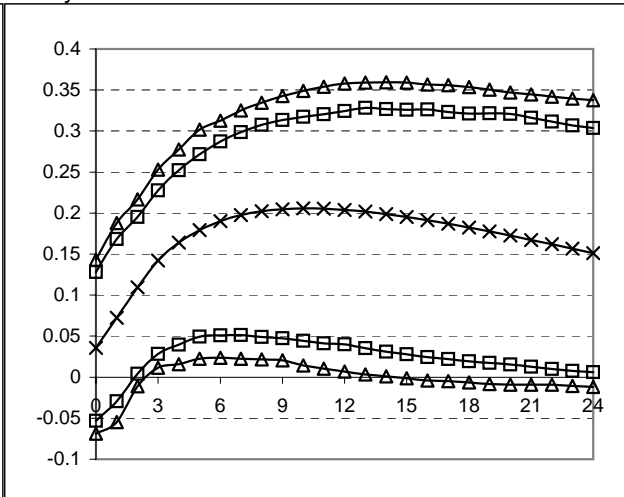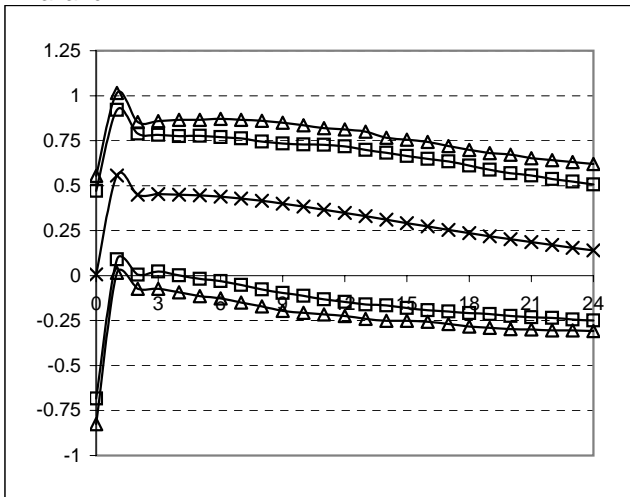
Malaysia

Thailand

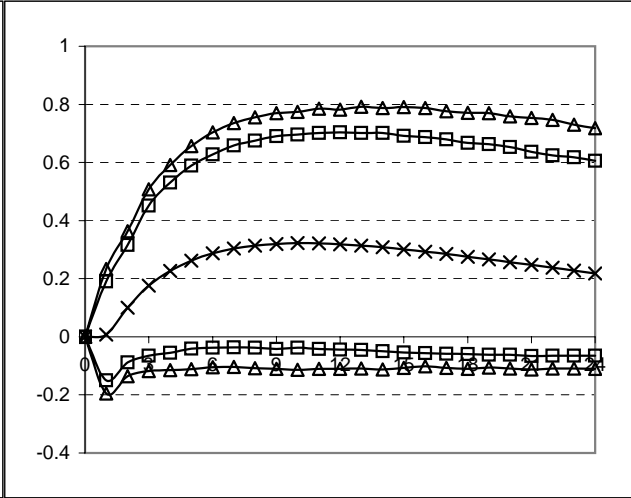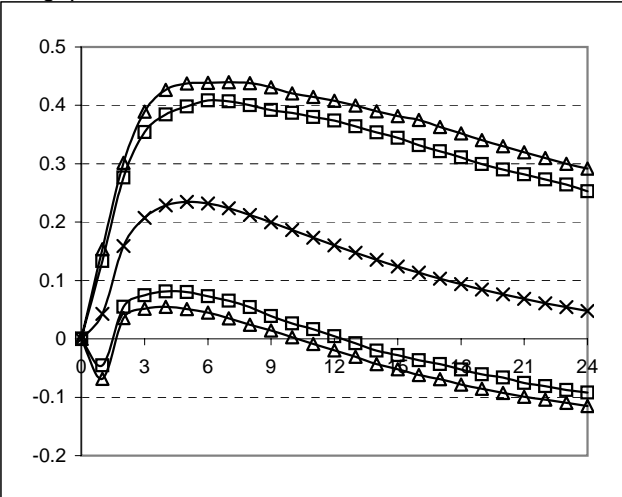Cross: Impulse response estimate; Square: 90% confidence band; Triangle: 95% confidence band

Figure 1. Continued

Shock to Singapore

Japan
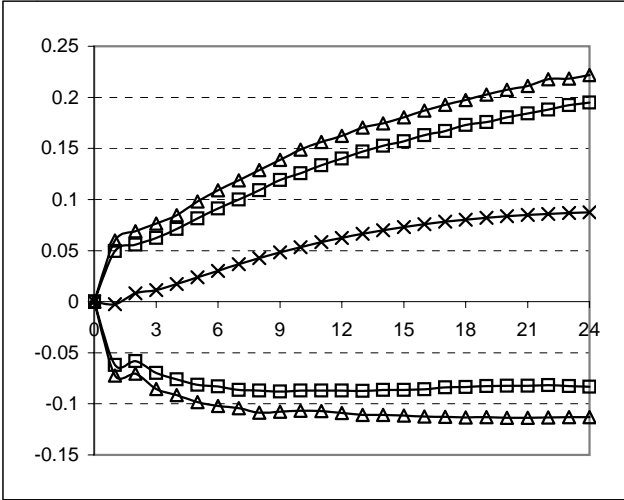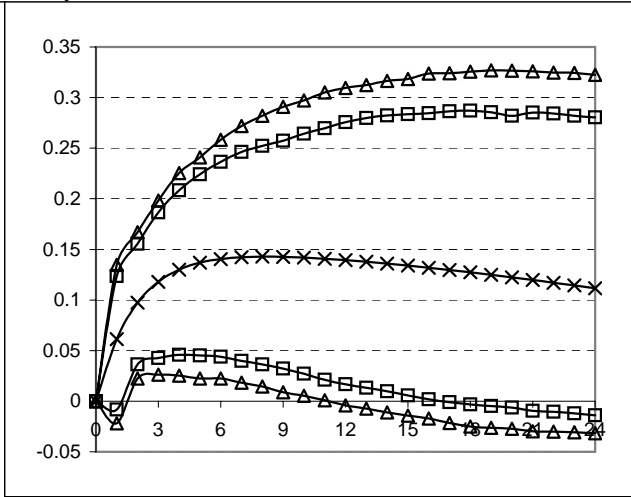

Korea
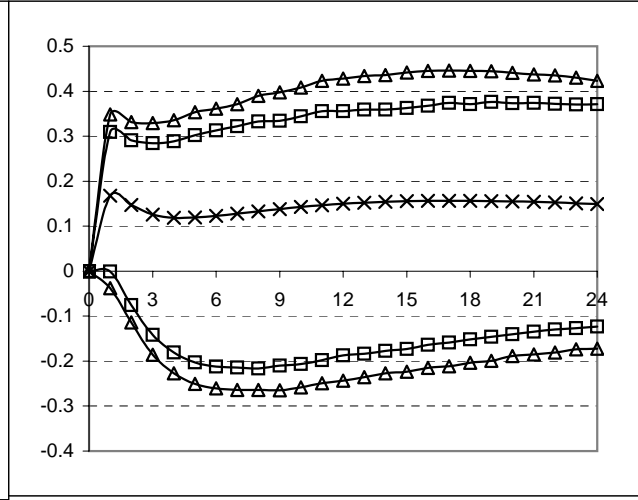
Singapore
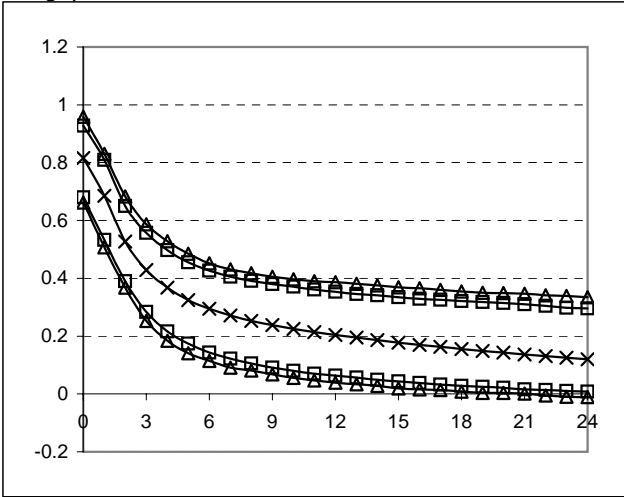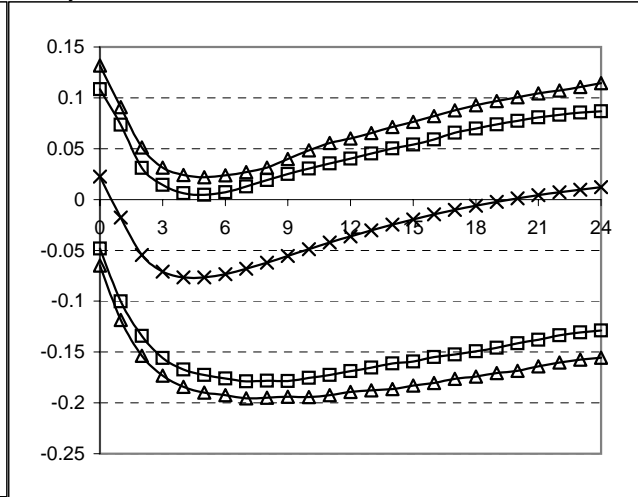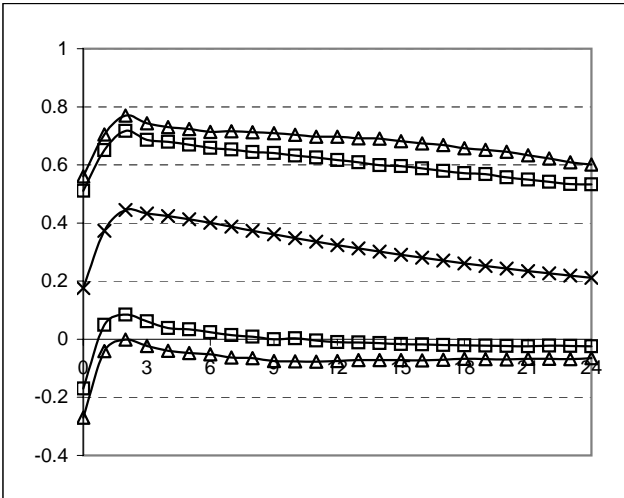
Malaysia

Thailand

Cross: Impulse response estimate; Square: 90% confidence band; Triangle: 95% confidence band

Figure 1. Continued

Shock to Malaysia

Japan
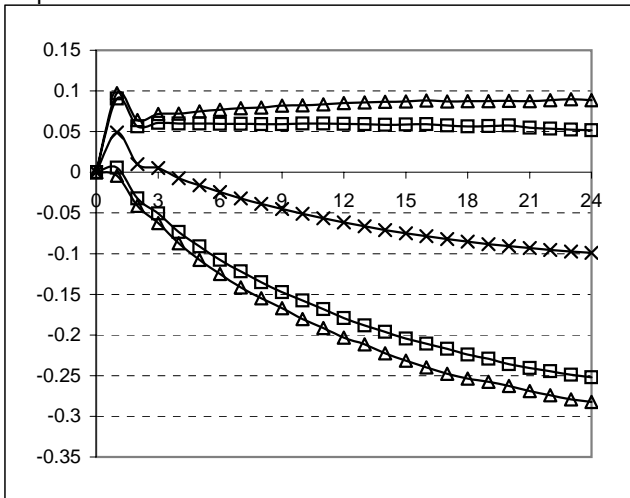
Korea
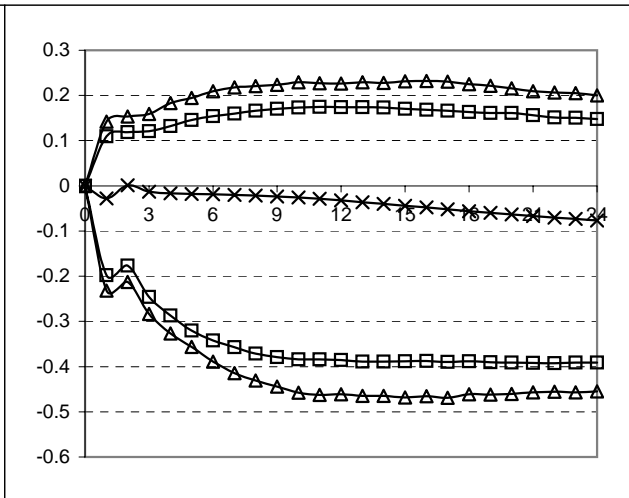
Singapore

Malaysia

Thailand

Cross: Impulse response estimate; Square: 90% confidence band; Triangle: 95% confidence band

Figure 1. Continued

Shock to Thailand
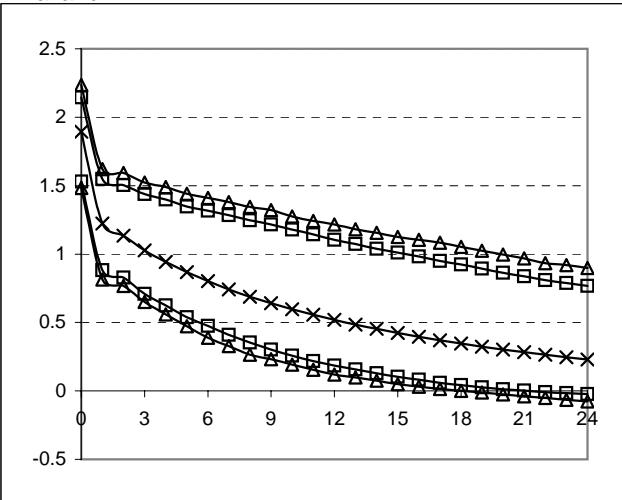
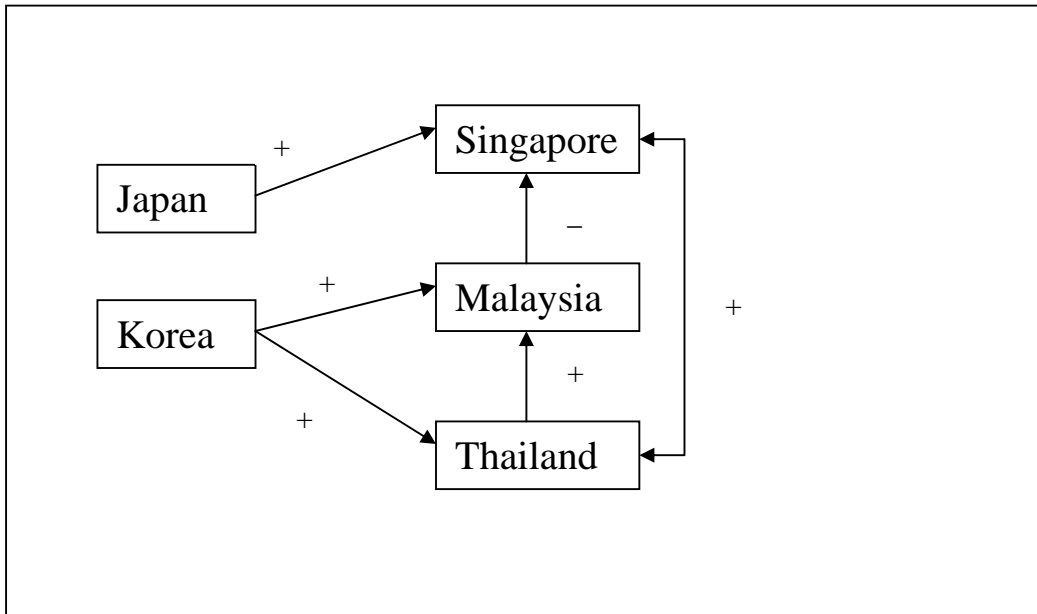Japan

Korea

Singapore

Malaysia

Thailand

Cross: Impulse response estimate; Square: 90% confidence band; Triangle: 95% confidence band

Figure 2. Directions and signs of causal relationships

The positive and negative signs respectively indicate positive and negative effects.