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Monetary Policy and the Exchange Rate During the Asian Crisis: Identification Through Heteroscedasticy

MONETARY POLICY AND THE EXCHANGE RATE DURING THE ASIAN CRISIS: IDENTIFICATION THROUGH HETEROSCEDASTICITY

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Abstract

This paper evaluates whether a tight monetary policy (i.e., an increase in the domestic interest rate) was successful in defending the exchange rate from speculative pressures during the Asian financial crisis. The empirical analysis applied to five Asian countries utilizes a bivariate VAR model, which is identified by taking into account the heteroscedasticity properties of the time-series of interest, following Sentana and Fiorentini (1999). The empirical evidence shows that tight monetary policy did not help to stabilize the currencies under investigation.

Keywords: *Monetary Policy, Financial Crisis, Identification*

JEL classification*: E52, C32*

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1. Introduction

The nature of the relationship between exchange rates and interest rates during the Asian financial crisis has been at the centre of a hot policy debate between the World Bank, the IMF and the US Treasury. While the IMF has argued that steep rises in interest rates were vital in stabilizing Asian exchange rates, the World Bank position, partly reflecting the views of its former chief economist Joseph Stiglitz, has been that interest rate hikes destabilized the currencies further by increasing the risk of bankruptcy, which led to further loss of confidence in these economies. Even though the Asian financial crisis is no longer a headline grabbing issue, drawing out policy lessons from such episodes remains vital for safeguarding international financial stability in the future. However, and perhaps not surprisingly, much of recent literature on the issue has been produced by the key players in the debate. Independent academic studies can make a useful contribution to gaining further insights into this key policy issue.

This paper provides new empirical evidence on whether higher interest rates were successful in defending Asian exchange rates from speculative pressures during the crisis period. The empirical analysis applied to five Asian countries is based upon a bivariate VAR model, which attempts to capture the relationship between the exchange rate and the interest rate. The identification of the system is achieved by taking into account the heteroscedasticity properties of the time-series under investigation, following Sentana and Fiorentini (1999). This method enables us to address the endogeneity of interest rates, a thorny econometric problem under any circumstances but especially acute during periods of speculative attacks. In this respect, the paper represents an important step forward over existing empirical studies, most of which either do not recognize or are have been unable to address this serious econometric problem, which can be a source of biased estimates.

The rest of the paper is organized as follows. Section 2 briefly reviews the existing theoretical and empirical literature, and outlines the methodological issues which a rigorous empirical analysis needs to address. Section 3 explains the empirical methodology used to identify the model, which utilizes the heteroscedasticity properties of the series. Section 4 gives details of the data and the equations to be estimated. Section 5 presents and discusses the empirical results. Finally, section 6 offers some concluding remarks.

2. Literature Review and Methodological Issues

The traditional view on the relationship between monetary policy and the exchange rate, on which the IMF position is based, is that tight monetary policy strengthens the exchange rate by sending a signal that the authorities are committed to maintaining a fixed rate, thereby increasing capital inflows (Backus and Driffill, 1985). A number of authors, however, have argued against the signaling value of a monetary tightening. Obstfeld (1994), Drazen and Masson (1994), and Bensaid and Jeanne (1997) provide a theoretical framework where the policymakers face a trade-off when pegging the exchange rate. The nature of the trade-off varies across models but they all have a common flavor. On one hand, letting the exchange rate float implies a fixed cost arising from the loss of credibility. This cost reflects the fact that policymakers have to abandon their disinflation goal linked to an exchange rate anchor. On the other hand, the cost of maintaining the peg is associated with either the output costs of an overvalued currency, or the excess current deficit of an overvalued currency, or the budgetary consequences of the higher interest rates needed to defend the currency. This framework has been associated with self-fulfilling currency crises because the relative cost of defending the currency increases substantially during a speculative attack and policymakers may choose to abandon the peg once the attack occurs.

In the Asian case, a number of economists (see, e.g., Radelet and Sachs, 1998, Feldstein, 1998; Furman and Stiglitz, 1998; Stiglitz, 1999) argued against the signaling value of monetary policy by considering the positive effect of interest rates on the likelihood of bankruptcy for highly leveraged borrowers. This will result in a larger country risk premium, a lower expected return to investors, and capital flight, which generates more downward pressure on the exchange rate. The foundations of this "revisionist" view, which predicts the "foreign exchange-interest rate Laffer curve" have been criticized by Krugman (1998).[1] He argues that even very high

[1] The focus of this paper is on the interest rate-exchange rate relationship. Other important monetary policy issues in the aftermath of currency crises are evaluating whether the real exchange rate (RER) has overshot and has become undervalued with respect to its long-run equilibrium; whether nominal currency appreciation or higher domestic inflation should be used as a means to bring it back to

interest rates might be preferable to a free fall in the exchange rate in countries with a large external debt denominated in foreign currency.

The empirical evidence on the issue is mixed. Some empirical studies (based on VAR model specification) support the traditional view of monetary policy. Specifically, Dekle et al (1998), using high-frequency (weekly) data, find that in the case of Korea the increase in the interest rate differential helped to appreciate the Korean Won. The analysis of Basurto and Gosh (2000), based on monthly data for Indonesia, Korea and Thailand, provides little evidence that higher real interest rates resulted in a higher risk premium, whilst they appear to be associated with an appreciation of the currency. Tanner's (1999) empirical study, which uses monthly data, focuses on an index of the exchange market pressure, which is measured by the sum of exchange rate depreciation and reserve outflows. Examining individual and pooled estimates of a VAR model estimated for Brazil, Chile, Mexico, Indonesia, Korea and Thailand, the author (op. cit.) finds that a contractionary monetary policy helps to reduce exchange rate market pressure.

On the other hand, a number of empirical studies support the revisionist view. The VAR model estimation and impulse response analysis of Goldfajn and Baig (1998), based on daily data, provide evidence of a perverse effect of a tight monetary policy on the exchange rate in Thailand, Indonesia, Philippine, Korea and Malaysia. Gould and Kamin (1999) use Granger causality test on weekly observations on the interest and exchange rates for six countries: Indonesia, Korea, Malaysia, Philippines, Thailand and Mexico. They find that during financial crises, exchange rates are not significantly affected in any of the countries examined by changes in the interest rates. A similar finding is reported in Ohno, Shirono and Sisly (1999), who apply the Toda-Yamamoto (1995) methodology (which allows to test Granger causality regardless of the order of integration of the time series) to daily observations on interest rates and exchange rates in Thailand, Korea, Indonesia, Philippines, Malaysia, Taiwan and Singapore.

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equilibrium; and finally, the costs of raising interest rates in terms of output losses and financial system fragility (see Goldfajn and Baig, 1998).

Empirical studies based on panel data analysis support the revisionist view. Goldfajn and Gupta (1999a,b), using monthly data from 80 countries covering the period of 1980-98, find that the probability of currency appreciation conditional on tight monetary policy is much lower in countries (such as the those in East Asia) with a weak banking sector. Kray (1999) examines factors determining whether or not defenses of a fixed peg against speculative attack succeed. The author (op. cit.), using monthly observations, instruments for the policy endogeneity of interest rates, and finds, in a sample of 75 developed and developing countries, that a tight monetary policy does not increase the likelihood of a successful defense. Finally, Furman and Stiglitz (1998) examine nine emerging markets with episodes of temporarily high interest rates. Using simple regression analysis, they find that both the magnitude and duration of such interest rates hikes are associated with exchange rate depreciation.

Methodological Issues

Perhaps the most important challenge in examining this issue empirically is the identification of monetary policy exogenous shocks as distinct from monetary policy actions (see also Kray, 1999[2]). Policy makers' actions to some extent respond to current developments in the economy, such as a speculative attack on the currency. This response may be captured by a policy reaction function and is distinct from exogenous policy shocks, which are defined as deviations of the authorities' behavior from their rule. In other words, an identification scheme is needed to solve the simultaneity problem between policy instruments and other endogenous variables, such as exchange rates to which monetary policy systematically reacts. Past empirical studies of the Asian crisis based on VAR analysis do not explicitly recognize the simultaneous feedback between exchange rates and interest rates. In the present paper we are able to identify, the effect of an interest rates rise on the exchange rate by taking into account the heteroscedasticity property of the time series under investigation, following the method put forward by Sentana and Fiorentini (1999).

A second empirical concern relates to the measurement of the monetary policy stance. Monetary authorities defend a currency from speculative attacks, not only by raising

[2] The author (op. cit.) uses an instrumental variable technique by employing changes in foreign reserves and changes in the country borrowing from the IMF as instruments. However, the latter are likely to be imperfect instruments: they are both endogenous during speculative attacks episodes.

domestic interest rates, but also by selling foreign reserves and contracting credit and monetary aggregates. A number of authors have therefore suggested that by considering only interest rates, it is not possible to obtain an accurate indicator of the stance of monetary policy (Tanner, 1999; Basurto and Ghosh, 2000). We believe that, since much of the economically interesting variation during speculative attack episodes is likely to occur at high frequencies, one needs to utilize daily data (in a way similar to Goldfajn and Baig, 1998). The use of low frequency data precludes modeling the likely path-dependence in the effects of interest rates on speculative pressures, a point emphasized by Drazen (1999). Since only the observations of financial assets like nominal exchange rates and interest rates are available at such high frequency, then the empirical analysis in the present paper is based only on these series.

Furthermore, it has been argued that the ex ante real interest rate is the most appropriate measure of the tightness or looseness of monetary policy (see Goldfain and Baig, 1998), but as inflation expectations generally are not observed directly, frequently the ex post real interest rate must be used instead. As Gould and Kamin (1999) pointed out, unfortunately, while actual inflation may be an adequate proxy for inflation expectations during normal periods, actual inflation may diverge considerably from inflation expectations during financial crises involving sharp depreciations of the exchange rate. Such depreciations may cause short bursts of inflation that cause real ex post interest rates to fall or even become negative temporarily, even though nominal interest rates may have been raised substantially. Therefore, the results of studies that rely on the real interest rate as a measure of monetary stance may be misleading.

Finally, the last issue that needs to be addressed (as Kray, 1999, also points out) concerns the possibility of asymmetries when considering long time periods. However, a non-linear specification (for instance, regime switching models) to capture the different system dynamics between normal and turbulent periods is beyond the scope of this paper. In the light of the previous considerations, we focus only on the period of financial turmoil in the East Asian countries.

3. Empirical Methodology: Identification through Heteroscedasticity

Given the $n\times1$ vector of endogenous variables Z_t, consider the structural VAR of order p:

$$AZ_t = B(L)Z_{t-1} + e_t \quad (1)$$

where B(L) is a polynomial in the lag operator, A is the $n\times n$ matrix which captures the contemporaneous interaction between the variables included in Z_t, and ε_t is the vector of structural innovations. Commonly, it is assumed that the structural innovations are unconditionally and conditionally orthogonal. Therefore both the unconditional covariance matrix $E(\varepsilon_t\varepsilon'_t) = \Gamma$ and the conditional covariance matrix $E_{t-1}(\varepsilon_t\varepsilon'_t) = \Gamma_t$ are diagonal. Furthermore, the unconditional variance of the structural innovations is normalised to unity, and, consequently, $E(\varepsilon_t\varepsilon'_t) = I$, where I is an identity matrix. The corresponding reduced form of the model in (1) is:

$$Z_t = C(L)Z_{t-1} + n_t \quad (2)$$

The unconditional and conditional covariance matrices of the reduced form innovations ν_t are $E(\nu_t\nu'_t) = \Sigma$, and $E_{t-1}(\nu_t\nu'_t) = \Sigma_t$, respectively. If the residuals are homoscedastic, then $\Sigma_t = \Sigma$, whereas, under conditional heteroscedasticity, $\Sigma_t \neq \Sigma$. Under homoscedasticity, there exists a time invariant orthogonal transformations such that $A\nu_t=\varepsilon_t$ is observationally equivalent to $A^*\nu_t = \varepsilon^*_t$, where $A^* = (Q^{-1})'A$, $\varepsilon_t^* = Q\varepsilon_t$, and Q is an arbitrary $n\times n$ orthogonal matrix. The observational equivalence occurs since both ε_t and the corresponding orthogonal rotated innovations ε^*_t imply that $\Sigma = A^{-1}(A^{-1})' = (A^*)^{-1}(A^*)^{-1\prime}$. In order to (exactly) identify the matrix A, we need n^2 restrictions. It is customary to impose, as a set of identifying restrictions, the normalisation to unity of the elements of main diagonal of Γ, and to assume orthogonal structural innovations. This provides a set of $n(n+1)/2$ restrictions and, therefore, it is not sufficient to identify the parameters.

Traditional VAR models of monetary policy are based on Bernanke’s (1986) methodology which provides the remaining n(n+1)/2 identifying restrictions by imposing a recursive structure on the impact multiplier matrix A. However, this identifying scheme, which is rationalized in terms of informational delays in the monetary authorities feedback rules, is hard to justify in open economies, where mutual contemporaneous feedback between interest rates and exchange rates may be more plausible. Smets (1996, 1997) and Kim and Roubini (2000) propose a non-recursive identifying scheme for a VAR including a few other variables in addition to interest and exchange rates. Bagliano and Favero (1998) use a non-VAR measure of monetary policy shocks to explicitly address the identification problem arising from the simultaneity of interest rates and exchange rates. They consider the US-Germany case, and derive a direct measure of German monetary policy shocks by using information extracted from financial markets.

In this paper we follow the suggestions of Sentana and Fiorentini (1999)[3] (see King, Sentana, Whadwani, 1994 and Normandin and Phaneuf, 1997 for empirical applications) and we are able to identify a bivariate VAR including interest rates and exchange rates only by considering the time varying conditional variances properties of the two financial series. Under conditional heteroscedasticity, $A\nu_t=\varepsilon_t$ is not observationally equivalent to $A_t^*\nu_t = \varepsilon^*_t$, where $A_t^* = (Q^{-1})'\Gamma_t^{-1/2}A$ and $\varepsilon^*_t = Q\Gamma_t^{-1/2}\varepsilon_t$. Observational equivalence does not occur since, unlike the homoscedastic case, the orthogonal rotations of the vector of structural innovations ε_t are different in each time period, given that the conditional covariance matrix Γ_t, and, consequently, A_t^* are time varying (for details, see Sentana and Fiorentini, 1999).

4. Data and Empirical Model

The analysis was carried out using daily data for the five countries most heavily affected by the Asian financial crisis: Thailand, South Korea, Malaysia, Indonesia and Philippines. A more detailed description of the data and their sources is provided in the Data Appendix. The focus of the analysis is on the behaviour of short-term interest rates during a period of speculative pressure and of floating of the exchange rates. However, we do not concentrate exclusively on the interest rates hikes which

[3] Rigobon (2000) provides an alternative identification scheme based on heteroscedasticity.

most of the countries experienced only for a very short time period. In Thailand, interest rates tended to be high even before July 2 (when the Thai baht was floated). Therefore, the estimation periods begin at different points in time: on May 1, 1997 for Thailand and on June 1, 1997 for the other countries. All estimation samples extend through to July 31, 1998, with the exception of Indonesia[4]. Following the suggestions of Gould and Kamin (1999), this end date was chosen for two reasons. First, the focus of our analysis is the behaviour of floating exchange rates during financial crises; by the end of July 1998, financial conditions had stabilised to a substantial degree in most of the Asian economies. Second, following the Russian devaluation and default in August 1998, credit spreads on emerging market bonds rose sharply throughout the world. The universal nature of the rise, and the lack of precipitating events in many of the countries experiencing the rise, suggest that after July 1998 spreads may have been influenced by more generalised trends in investor liquidity and risk aversion rather than by country-specific risk factors.

For each of the following countries we specify the conditional mean equations as a structural VAR with a lag order of 1:

$$\Delta e_t = a_{12}\Delta i_t + b_{11} + b_{12}\Delta e_{t-1} + b_{13}\Delta i_{t-1} + b_{14}\Delta i^*{}_{t-1} + e_t$$

$$\Delta i_t = a_{21}\Delta e_t + b_{21} + b_{22}\Delta e_{t-1} + b_{23}\Delta i_{t-1} + b_{24}\Delta i^*{}_{t-1} + e_t$$

Some unit root pre-testing analysis was carried out, showing evidence of a unit root in each series[5]. The two endogenous variables are a proxy for the nominal exchange rate depreciation rate (in percent values), that is $100\times\Delta e_t$, where Δe_t is the first order difference of the log of the nominal bilateral exchange rate (with respect to the US dollar), and Δi_t, the first order difference of the domestic short term interest rate. We also include the first order difference of the US federal funds rate Δi^*_t as a strictly exogenous regressor. Assuming that the structural innovations are Gaussian, the conditional log-likelihood (ignoring a constant term) is:

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[4] Given the financial and political turbulence the estimation sample in Indonesia extends through to December 31, 1998.

[5] An augmented Dickey Fuller test was carried out. Results are available upon request.

$$L_t = -1/2 \log\left|\Gamma_t\right| - 1/2 e'_t \, (\Gamma_t)^{-1} e_t \qquad (3)$$

where the n-dimensional vector of structural innovations ε_t = AZ_t - $B(L)Z_{t-}$ To explicitly recognize the existence of conditional heteroscedasticity we use the following GARCH(1,1) specification[6] for the conditional variance for the ith equation (with i = 1,2):

$$s^2_{i,t} = (1 - d_{1i} - d_{2i}) + d_{1i} e^2_{t-1} + d_{2i} s^2_{i,t-1}$$

where the constraints $\delta_{ij} \geq 0$ ensure non-negative conditional variances, and the condition $\delta_{i1} + \delta_{i2} < 1$ allows for covariance stationary conditional variances. The constrained intercept terms (see King, Sentana, Whadwani, 1994 and Normandin, 1997) ensure that the unconditional variance of each innovation is unity, and, consequently, the ith structural disturbance is homoscedastic if $\delta_{i1} = \delta_{i2} = 0$.

5. Estimation and Empirical Results

We maximize the joint log-likelihood $\Sigma_t L_t$ over the parameters of the conditional mean and variance equations (A, B(L), δ_{ij}, where i,j = 1,2) by using the simplex algorithm in the first few iterations and then the BFGS algorithm. The Quasi Maximum Likelihood (see Bollersev and Woodlbridge, 1992) estimator was used in order to obtain robust standard errors, given the evidence of non Gaussian standardized residuals. The sum of the estimated parameters in the conditional variance equations gave evidence of persistence in the conditional variance of the exchange rate depreciation rate in each country considered, and in the conditional variance of the first order difference of the domestic policy rate in South Korea and Thailand. As argued in Bollersev, Chou and Kroner (1992) this is a common finding in much of the empirical literature using high frequency financial data. Therefore, we specified an Integrated GARCH (IGARCH) model, by imposing $\delta_{i1} + \delta_{i2} = 1$ (where i = 1,2), for the corresponding conditional variance equations[7].

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[6] This model specification has been useful to describe the time varying conditional volatility of many macroeconomic and financial time series. See Bollersev, Chou and Kroner (1992) for a survey of the literature.

[7] As shown in Nelson (1990) and in Bougerol and Picard (1992), the IGARCH model is strictly stationary and ergodic, though not covariance stationary. Furthermore, Lumsdaine (1996) shows that in

The estimates of the conditional mean and variance equations parameters are reported in Table 1. The contemporaneous response of the domestic interest rate to an exchange rate depreciation (measured by the coefficient a_{21}) is positive in each country, suggesting a tight monetary policy in response to an exchange rate depreciation. The contemporaneous response of the nominal exchange rate to an increase in the domestic interest rate (measured by the coefficient a_{12}) varies across countries. Specifically, there is evidence of a statistically significant perverse impact effect on the exchange rate in Thailand (a one point increase in the domestic policy rate leads to a seven percent contemporaneous currency depreciation). A similar perverse effect (even though it is statically insignificant) occurs in Indonesia. There is an exchange rate appreciation effect in Korea and the Philippines, but it is very marginal and statistically insignificant. In Malaysia also there is a contemporaneous appreciating effect (of modest magnitude: around 0.08%, although it is statistically significant) due to an interest rate rise. The presence of such "traditional" monetary policy effect in Malaysia (even though it is very marginal) can be reconciled with the "revisionist view". Given the modest rise (relative to the other countries) in the Malaysian interest rates during the crisis period, the financial burden of the leveraged corporations was not significantly affected.

The impulse responses in Figure 1 show an immediate response of the nominal exchange rate level to a unit shock to the interest rate. Specifically, after an impact overshooting, or undershooting (see Dornbusch, 1976), depending on the depreciating or appreciating effect on the exchange rate, the series converges to its steady state value quickly (within four days).

The diagnostic tests described in Table 2 show that the model is not mis-specified. According to the Ljung-Box test on the standardized residuals, the latter do not show any evidence of serial correlation. There is no evidence of remaining heteroscedasticity, according to the Ljung-Box on the squared standardized residuals. To be more specific, an IGARCH(1,1) specification does not provide any evidence of time varying variances in the standardized residuals of both equations in Thailand,

contrast to the case of a unit root in the conditional mean, the presence of a unit root in the conditional variance does not affect the limiting distribution of the estimators.

Malaysia, and Indonesia. In order to obtain heteroscedasticity-free standardized residuals, we included only a moving average error term of order one and the squared residuals lagged once and three times in the conditional variance specification for the first order difference in the Korean interest rate (see the coefficients δ_{21}, δ_{25} and δ_{26} respectively in Table 1). For the same reason, we included a moving average error term of order two in the conditional variance equation for the rate of depreciation of the Philippines exchange rate (see the coefficients δ_{13}, δ_{14} in Table 1).

6. Conclusions

This paper has focused on the effect of monetary policy tightening on the exchange rates during the recent Asian crisis. Advocates of the "revisionist view" (see Stiglitz, 1999) emphasize the perverse effect of an increase in the domestic interest rates on the domestic currency, due to a higher probability of bankruptcy of highly leveraged corporations. The simultaneous feedback between exchange rates and interest rates implies an empirical identification problem which has not been addressed satisfactorily by previous empirical studies of the Asian crisis. In this paper we specify a VAR model and identify the system using the procedure suggested by Sentana and Fiorentini (1999). The latter is based on the presence of heteroscedasticity in the high frequency financial time series under investigation. In line with a number of empirical studies (see above), the analysis carried in this paper suggests that monetary policy tightening did not help to defend the exchange rates that came under speculative attack.

Table 1: Estimates of the conditional mean and variance equations

	Thailand	Korea	Malaysia	Philippines	Indonesia
b_{11}	0.142 (1.75)	0.024 (1.30)	0.178 (1.63)	0.119 (1.17)	0.086 (0.52)
b_{12}	0.038 (0.78)	0.225 (6.02)	-0.005 (0.11)	0.142 (2.71)	0.137 (2.23)
b_{13}	-0.303 (3.58)	-0.013 (0.20)	-0.026 (0.70)	0.082 (1.04)	-0.133 (1.60)
b_{14}	0.681 (2.88)	(-)	-0.171 (0.54)	-0.503 (1.14)	-0.622 (1.63)
b_{21}	-0.005 (0.75)	-0.040 (3.29)	-0.171 (6.31)	-0.084 (9.59)	0.048 (0.47)
b_{22}	0.034 (2.48)	0.063 (14.18)	0.030 (0.90)	0.012 (0.90)	-0.005 (0.71)
b_{23}	0.242 (5.88)	0.217 (4.37)	0.087 (3.49)	0.039 (0.32)	-0.022 (0.84)
b_{24}	0.944 (25.02)	(-)	0.204 (4.99)	0.174 (1.80)	-0.229 (1.31)
a_{12}	0.722 (16.50)	-0.042 (0.66)	-0.083 (1.90)	-0.100 (1.24)	0.183 (1.25)
a_{21}	0.087 (7.04)	0.007 (1.54)	0.017 (1.51)	0.048 (13.14)	0.012 (0.50)
δ_{11}	0.069 (-)	0.272 (-)	0.066 (-)	0.000 (-)	0.266 (-)
δ_{12}	0.931 (130.55)	0.728 (73.77)	0.934 (161.43)	0.029 (4.15)	0.733 (31.28)
δ_{21}	0.086 (-)	0.273 (10.82)	0.883 (35.82)	0.063 (-)	0.000
δ_{22}	0.904 (149.94)	0.000 (-)	0.000 (-)	0.000 (-)	0.000
δ_{13}	0.000	0.000	0.000	0.030 (-)	0.000
δ_{14}	0.000	0.000	0.000	0.906 (105.18)	0.000
δ_{25}	0.000	0.172 (-)	0.000	0.000	0.000
δ_{26}	0.000	0.555 (36.25)	0.000	0.000	0.000

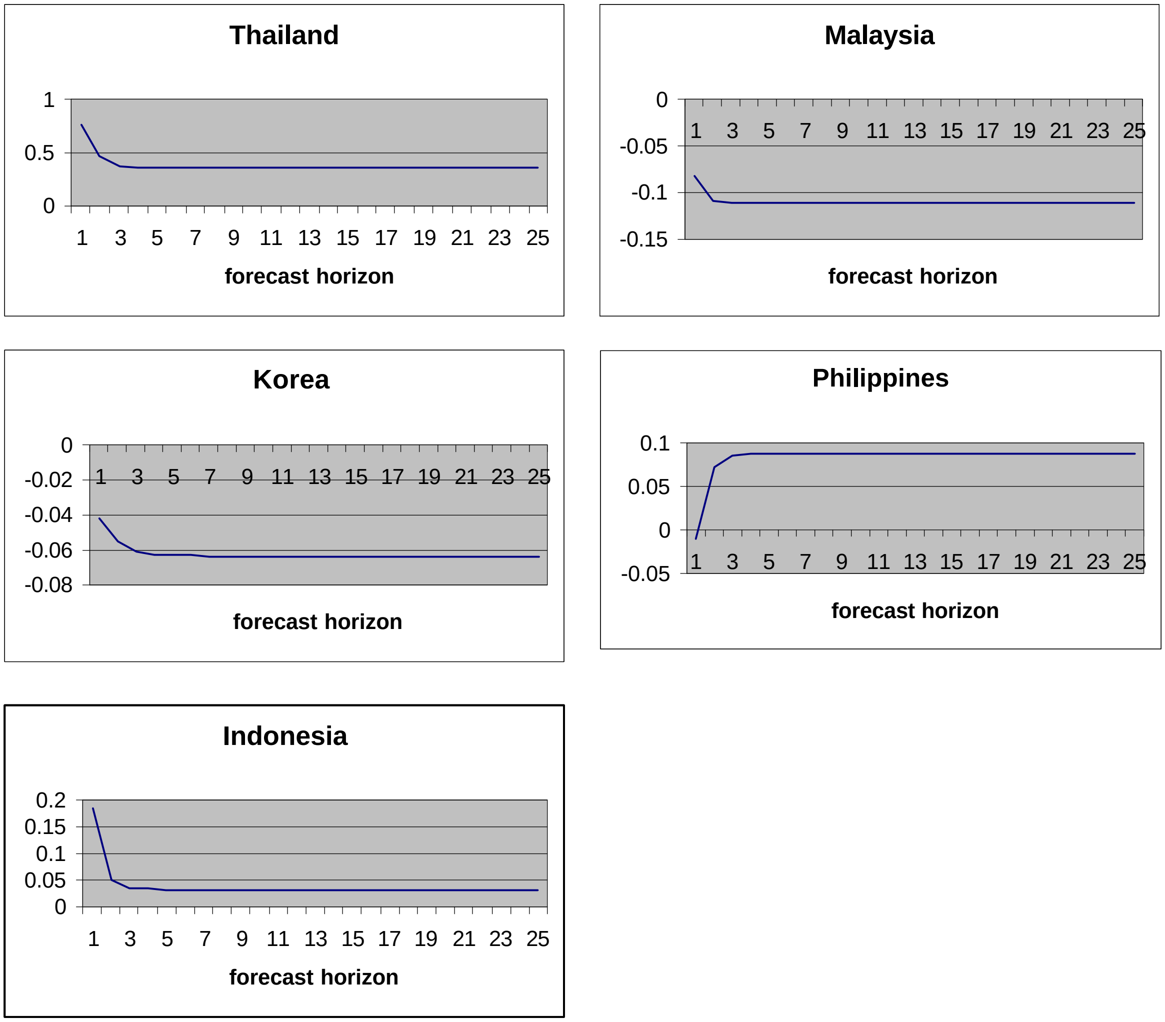
Note: The (-) symbol is for the estimated restricted coefficients when imposing non-negativity constraints and an IGARCH behavior on the conditional variance equations. The numbers in parentheses are the t-ratios.

Table 2: Diagnostic tests on the residuals

	Thailand	Korea	Malaysia	Philippines	Indonesia
$LB_1(20)$	0.73	0.20	0.58	0.67	0.76
$LB_2(20)$	0.67	0.19	0.24	0.72	0.99
$LB^2_1(20)$	0.68	0.58	0.88	0.29	0.35
$LB^2_2(20)$	0.98	0.23	1.00	0.90	1.00
Ep-$Stat_1$	0.00	0.00	0.00	0.00	0.00
Ep-$Stat_2$	0.00	0.00	0.00	0.00	0.00

Note: The diagnostics are computed for the standardized residuals ε_i (where, i = 1,2). LB is the p-value of the Ljung-Box test for the null of no autocorrelation against the alternative of autocorrelation up to order 20 for the standardized residuals. LB^2 is the p-value of the Ljung-Box test for the null of no autocorrelation against the alternative of autocorrelation up to order 20 for the squared standardised residuals. Ep-Stat is the p-value for the normality test on the residuals (see Doornik, Hansen, 1994). The subscript i (where, i = 1,2) denotes i^{th} equation of the VAR model.

Figure 1: Impulse response of the level of the nominal exchange rate to a unit shock to the level of the domestic nominal interest rate



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Data Appendix

The Thai repurchase rate, the Malaysian overnight call rate, the Korean call rate were obtained from the corresponding Central Bank websites. The Philippines prime rate and the Indonesian one month deposit rate were obtained from DATASTREAM. The Indonesian rupiah and the Philippines peso (in terms of US dollars) were also obtained from DATASTREAM, whereas the Thai baht, the Korean won and the Malaysian ringitt were obtained from the Federal Reserve Bank website.