

M ONETARY POLICY AND THE EXCHANGE RATE  
DURING THE ASIAN CRISIS:  
IDENTIFICATION THROUGH HETERO SCEDASTIC ITY

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A bstract

This paper exam ines whether a m onetary policy tightening (i.e., an increase in the dom estic interest rate) was successful in defending the exchange rate from speculative pressures during the Asian financial crisis. W e estimate a bivariate VECM for four Asian countries, and improve upon existing studies in two important ways. F irst, by using a long data span we are able to compare the effects of an interest rate rise on the nom inal exchange rate during tranquil and turbulent periods. Second, we take into account the endogeneity of interest rates and identify the system by exploiting the heteroscedasticity properties of the relevant tim e series, as suggested by Sentana and F iorentini (2001). W e find that while tight m onetary policy helped to defend the exchange rate during tranquil periods, it had the opposite effect during the Asian crisis.

**Keywords:** M onetary Policy, F inancial C risis, 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 much debated by the World Bank, the IMF and the US Treasury, and has important policy implications. 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 (see Stiglitz, 1999). Drawing out policy lessons from episodes such as the Asian crisis is clearly vital for safeguarding international financial stability in the future.

This paper aims at contributing to this key policy debate by providing new empirical evidence on whether higher interest rates were in fact successful in defending Asian exchange rates from speculative pressures during the crisis period. We improve upon existing studies in two important ways. First, by using a long data span we are able to examine the effects of an interest rate rise on the nominal exchange rate during tranquil periods and to compare them with those during more turbulent periods. Second, we adopt an appropriate identification scheme. Specifically, we estimate a bivariate Vector Error Correction (VECM) model for four Asian countries in order 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 (2001). 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. This is in marked contrast to earlier empirical studies, most of which either did not recognize or were 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 exploits 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

### Existing Literature

The traditional view on the relationship between monetary policy and the exchange rate, on which the IMF position is based, is that a 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 the 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 resulting from it, 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 case of the Asian financial crisis, a number of economists, including Radelet and Sachs (1998), Feldstein (1998), Furman and Stiglitz (1998) and Stiglitz (1999), argued against the signaling value of tighter monetary policy by pointing out the effects of higher interest rates on the probability of bankruptcy of highly leveraged borrowers. These manifest themselves in the form of a larger country risk premium, a lower, possibly negative, expected return to investors, and capital flight, all of which generate downward pressure on the exchange rate. The role played by banks' and firms' balance sheets has been analysed by Stiglitz (1999) in a partial equilibrium model. More recently, Gertler et al. (2000) have stressed the perverse effect of a tight monetary policy occurring through the balance sheet channel in the context of a

general equilibrium model. This is essentially a small open economy macro model incorporating "a financial accelerator" mechanism (see also Bernanke and Gertler, 1999). Thus, the "revisionist" view predicts a "foreign exchange-interest rate Laffer curve". The foundations of this view have, however, been criticized by Krugman (1998), who argues that even very high interest rates might be preferable to a free fall in the exchange rate in countries with a large external debt denominated in foreign currency.<sup>1</sup>

The available empirical evidence is mixed. Empirical studies based on panel data analysis tend to support the revisionist view, while studies based on VAR model specifications provide conflicting results. 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 a tight monetary policy is much lower in countries (such as the those in East Asia) with a weak banking sector. Kraay (1999) examines factors determining whether or not defences of a fixed peg against speculative attack succeed. Using monthly observations, he 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. Fumman 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 rate hikes are associated with exchange rate depreciation.

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

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<sup>1</sup> 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 overshoot 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

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, 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, Philippines, Korea and Malaysia. Gould and Kamini (1999) use Granger causality test on weekly observations on 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 interest rates. A similar finding is reported in Ohno, Shirono and Sisly (1999), who apply the Toda and 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. Finally, the evidence in Cho and West (2001), who solve the identification problem by proposing a methodology mapping second moments to the point estimates of the relevant coefficients using structural restrictions, is mixed. Specifically, their main finding is that an exogenous increase in interest rates led to exchange rate appreciation in Korea and the Philippines, and to a depreciation in Thailand. However, as the authors recognize, the confidence intervals around the point estimates are very large, suggesting a cautious interpretation of their findings.

#### Methodological Issues

Three important issues need to be addressed by an empirical study of the effects of monetary policy on exchange rates. The first one relates to the likely endogeneity of monetary policy, the second to the measurement of the monetary policy stance, and the third to the possibility of regime switches. We discuss each of them in the remainder of this section.

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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).

Perhaps the most important empirical challenge is the identification of monetary policy exogenous shocks as distinct from monetary policy actions (see also Kraay, 1999<sup>2</sup>). 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. By contrast, in the present paper we are able to identify the effects 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 (2001).

According to Goldfain and Baig (1998), the ex-ante real interest rate is the most appropriate measure of the tightness or looseness of monetary policy<sup>3</sup>. However, while this may in principle be a valid economic argument, in practice there are thorny measurement issues associated with obtaining accurate measures of the real rate of interest. Because inflation expectations generally are not observed directly, this frequently leads to using ex-post measures of the real interest rate, by using realised rather than expected inflation. Unfortunately, as Gould and Kanin (1999) point out, while actual inflation may be an adequate proxy for inflation expectations during tranquil periods, it may diverge considerably from inflation expectations during financial crises that involve sharp depreciations of the exchange rate. Such depreciations may cause short bursts of inflation that lead to ex-post real interest rates falling or even temporarily becoming negative, even though nominal interest rates may have been raised substantially. Thus, the results of studies that rely on ex-post measures of the real interest rate as indicators of the monetary stance may be misleading. We, therefore, utilise the nominal interest rate in our estimations, which

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<sup>2</sup> This author uses an instrumental variable technique by employing changes in foreign reserves and changes in the country borrowing from the IMF as instruments. However, these are likely to be imperfect instruments, since they are unlikely to be exogenous during speculative attacks.

<sup>3</sup> Some authors (e.g. Tanner, 1999; Basurto and Ghosh, 2000), also suggest using other monetary indicators to capture the stance of monetary policy, such as foreign reserves and credit aggregates.

we believe is the most accurate and widely available indicator of the stance of monetary policy.

Finally, the possibility of asymmetries also needs to be taken into account, as Kearney (1999) points out. In particular, regime switches are likely to occur over longer time periods. Whilst the empirical studies reviewed above focus on the crisis period only, we use a long data span, which enables us to compare the relationship between exchange rates and interest rates in tranquil and turbulent periods. We, therefore, model policy shifts by defining appropriate dummies, which are fully described in Section 4.

### 3. Empirical Methodology: Identification through Heteroscedasticity

Given the  $n \times 1$  vector of endogenous variables  $Z_t$ , consider the structural VAR of order  $p$ :

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

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  $\varepsilon_t$  is the vector of structural innovations. It is usually 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:

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

The unconditional and conditional covariance matrices of the reduced form innovations  $v_t$  are  $E(v_t v_t') = \Sigma$ , and  $E_{t-1}(v_t v_t') = \Sigma_t$ , 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  $AV_t = \varepsilon_t$  is observationally equivalent to  $A^* v_t = \varepsilon_t^*$ , 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  $\varepsilon_t$  and the corresponding orthogonal rotated innovations  $\varepsilon_t^*$  imply that  $\Sigma = A^{-1}(A^{-1})' = (A^*)^{-1}(A^*)^{-1}$ . 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  $\Gamma$ , 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 King 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 (1999) 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 (2001)<sup>4</sup>, which enable us to identify a bivariate VECM model including interest rates and exchange rates only by considering the time-varying conditional variances properties of the two financial series. Under conditional heteroscedasticity,  $Av_t = \varepsilon_t$  is not observationally equivalent to  $A_t^*v_t = \varepsilon_t^*$ , 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 in the homoscedastic case, the orthogonal rotations of the vector of structural innovations  $\varepsilon_t$  are different in each time period, given that the conditional covariance matrix  $\Gamma_t$ , and, consequently,  $A_t^*$ , are time-varying (for details, see Sentana and Fiorentini, 2001).

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<sup>4</sup> Rigobon (2000) provides an alternative identification scheme based on heteroscedasticity. See King, Sentana, Wadhvani (1994) and Normandin and Phaneuf (1997) for empirical applications of the Sentana-Fiorentini approach.



#### 4. Data and Empirical Model

The analysis was carried out using monthly data for the period 1991:2-2001:10. The countries under investigation are those which experienced a temporary and significant monetary policy tightening during the Asian financial crisis: Thailand, South Korea, Indonesia and the Philippines. The bilateral nominal exchange rate series are defined as units of domestic currency per US dollar. The domestic interest rate series used are the Korean call overnight rate, the Indonesian interbank call rate, the Philippines interbank call loan rate, and the Thai repo rate, which are the relevant policy interest rates in each case. The US federal funds rate is used as the foreign interest rate. All series were obtained from Datastream.

Some unit root pre-testing analysis was carried out, showing evidence of a unit root in each series.<sup>5</sup> In each case the conditional mean equations are specified as a VECM:

$$\begin{aligned} \Delta x_{t} &= b_{12} * \Delta i_{t}^{d} + b_{12,D} * \Delta i_{t}^{d} * Dum1 + \\ &a_1 * (i_{t-1}^{f} - i_{t-1}^{f} - b_0 - b_{0,D} * Dum1) + d_1 \Delta i_{t}^{f} + e_{1t} \end{aligned} \quad (3)$$

$$\begin{aligned} \Delta i_{t}^{d} &= b_{21} * \Delta x_{t} + b_{21,D} * \Delta x_{t} * Dum1 + \\ &a_2 * (i_{t-1}^{f} - i_{t-1}^{f} - b_0 - b_{0,D} * Dum1) + d_2 \Delta i_{t}^{f} + e_{2t} \end{aligned}$$

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  $\Delta e_t$  is the first-order difference of the log of the nominal bilateral exchange rate (with respect to the US dollar), and  $\Delta i_t^d$  the first-order difference of the domestic short-term interest rate. The US federal funds rate,  $i_t^f$ , is treated as a strictly exogenous regressor.

To take into account policy shifts, particularly the tightening of monetary policy that took place in the four countries under investigation in an attempt to halt the slide of

the exchange rate during the crisis, we introduce an intervention dummy Dum1. Inspection of the data presented in Figures 1-4 suggests defining it in the following way in each case:

- Thailand: 1 during the Aug97-Dec97 and Feb98-Jul98 period, and 0 elsewhere.  
 Korea: 1 during the Dec97-May98 period, and 0 elsewhere.  
 Indonesia: 1 during the Aug97-Jul99 period, and 0 elsewhere.  
 Philippines: 1 during the Jul97-Oct97 period, and 0 elsewhere.

In the first stage of the empirical analysis, it was found that the interest parity condition holds in the long-run in both Indonesia and the Philippines. In the case of Thailand and Korea, however, we detected that the interest parity condition holds for the post-crisis period only if an additional shift in the mean is taken into account. These intercept shifts are usually interpretable as changes in the country risk premium (see Obstfeld and Rogoff, 1996), which mainly reflect the monetary policy tightening in the presence of speculative attacks. Thus, an additional intervention dummy (Dum 2) is defined that takes the value 1 from Nov98 onwards, and 0 elsewhere in the case of Thailand, and value 1 from Feb99 onwards and 0 elsewhere for Korea.

The estimated equilibrium relationships included in the VECMs are:

Thailand:  $(\hat{i}_t - \hat{i}_t^f - 2.62 - 11.81 * DUM 1 + 6.60 * DUM 2)$

Korea:  $(\hat{i}_t - \hat{i}_t^f - 7.39 - 12.15 * DUM 1 + 8.15 * DUM 2)$

Indonesia:  $(\hat{i}_t - \hat{i}_t^f - 7.08 - 37.10 * DUM 1)$

Philippines:  $(\hat{i}_t - \hat{i}_t^f - 7.66 - 34.19 * DUM 1)$

These relationships are plausible, suggesting sharp increases in the country risk premia during the crisis period, with Indonesia and the Philippines experiencing the largest increases, amounting to 37% and 34% respectively, compared with around

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<sup>5</sup> An augmented Dickey Fuller test was carried out. Results are available upon request.

12% in the case of both Korea and Thailand. Interestingly, the estimates also suggest that Thailand and Korea enjoy substantially lower risk premia in the post-crisis period than they experienced prior to the crisis. In fact the risk premia seem to have turned negative overall in both Thailand and Korea, approximately around -4% and -1% respectively.

Assuming that the structural innovations are Gaussian, the conditional log-likelihood (ignoring a constant term) is:

$$L_t = -1/2 \log |\Gamma_t| - 1/2 e_t' (\Gamma_t)^{-1} e_t \quad (3)$$

where  $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t})'$  is the vector of structural innovations. To explicitly recognize the existence of conditional heteroscedasticity we use the following GARCH (1,1) specification for the conditional variance for the  $i^{\text{th}}$  equation (with  $i = 1, 2$ ):<sup>6</sup>

$$s_{it}^2 = (1 - g_{1i} - g_{2i}) + g_{1i} e_{i,t-1}^2 + g_{2i} s_{i,t-1}^2$$

where the constraints  $\gamma_{ij} \geq 0$  ensure non-negative conditional variances, and the condition  $\gamma_{11} + \gamma_{22} < 1$  allows for covariance stationary conditional variances. The constrained intercept terms (see King, Sentana, and Waldhansl, 1994, and Normandin and Phaneuf, 1997) ensure that the unconditional variance of each innovation is unity, and, consequently, that the  $i^{\text{th}}$  structural disturbance is homoscedastic if  $\gamma_{11} = \gamma_{22} = 0$ . The normalisation to unity of the unconditional variances provides the two additional restrictions necessary to identify the shifts in the slope coefficients.

We maximize the joint log-likelihood  $\sum_t L_t$  over the parameters of the conditional mean and variance equations ( $A, B(L), \delta_{ij}$ , 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 Bollerslev and Woodbridge, 1992) estimator was used in order to obtain robust standard errors, given the evidence of non-Gaussian standardized residuals.

<sup>6</sup> This model specification has been found to be useful to describe the time-varying conditional volatility of many macroeconomic and financial time series (see Bollerslev, Chou and Kroner, 1992).

## 5. Estimation and Empirical Results<sup>7</sup>

The AIC and Schwarz criteria information suggest a VECM (0) for all countries except for Indonesia, for which a VECM (2) was selected. The estimates of the conditional mean and variance equations parameters are presented in Table 1, which only reports the parameters of interest<sup>8</sup>. As can be seen from Table 1, there is clear evidence of GARCH effects, with the estimated parameters of the conditional variance being significant, which supports the identification scheme proposed in this paper. Furthermore, the sum of the estimated parameters in the conditional variance of the domestic interest rate is less than unity, with the exception of the Philippines<sup>9</sup>. Consequently, for this country we specify an Integrated GARCH (IGARCH) model, by imposing  $\gamma_{i1} + \gamma_{i2} = 1$  (where  $i = 1, 2$ ), for the corresponding conditional variance equations.<sup>10</sup> Finally, the diagnostics presented in Table 2 are satisfactory, with the exception of Thailand and Korea, for which some autocorrelation in the standardized residuals is detected by the Ljung-Box statistic. The standard errors (and corresponding t-ratios) presented in Table 1 for Thailand and Korea are therefore corrected for the presence of residual autocorrelation.

The results reported in Table 1 confirm the presence of cointegration, as the error correction coefficient,  $\alpha_2$ , is found to be negative and statistically significant in the estimated interest rate equations. They also suggest that there was a significant monetary policy contraction during turbulent periods, since the sum of the coefficients  $\beta_{21}$  and  $\beta_{21D}$  is negative in all cases, and  $\beta_{21D}$  is highly significant. This confirms that in each of the four countries there was a contemporaneous increase of the domestic interest rate in response to exchange rate depreciation during the crisis period.

Table 1 also contains clear evidence of a nominal exchange rate appreciation in response to a domestic interest rate increase during tranquil periods. This is indicated

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<sup>7</sup> The empirical analysis has been carried using the RATS software.

<sup>8</sup> We do not report the coefficients on the speed of adjustment coefficient in the exchange rate equation, on the lags and the US federal funds rate.

<sup>9</sup> As pointed out by Bollerslev, Chou and Kroner (1992), the evidence of persistence in the conditional variance is a common finding in much of the empirical literature using financial data.

<sup>10</sup> The presence of IGARCH, as shown by Sentana and Fiorentini (2001), does not affect the identification of the system. In this case, the authors (op. cit.) suggest to restrict the constant part of the

by the coefficient  $\beta_{12}$ , which is found to be positive in all four countries. In contrast, during turbulent periods the nominal exchange rate appears to depreciate sharply in response to rises in the domestic policy rate. This is evidenced by the sum of the coefficients  $\beta_{12}$  and  $\beta_{12D}$ , which is clearly negative. Note that  $\beta_{12D}$  is much larger in absolute terms than  $\beta_{12}$  and that it is highly significant in all four cases, suggesting that the effects of monetary policy tightening on the exchange rate during turbulent periods were not only opposite to those during tranquil periods, but also substantially larger.

## 6. Conclusions

This paper has examined the effects of a monetary policy tightening on the exchange rates during the recent Asian crisis. Advocates of the "revisionist view", such as Stiglitz (1999), in a partial equilibrium model, and Gertler et al. (2000) in a general equilibrium framework, emphasize the perverse effect of an increase in the domestic interest rates on the domestic currency, owing to a higher probability of bankruptcy of highly leveraged corporations. Our empirical results are consistent with the conventional view in the sense that we find that monetary policy tightening leads to a nominal exchange rate appreciation during tranquil periods. However, they also provide support to the "revisionist" view in that they very clearly show that the tightening of monetary policy that occurred during the Asian financial crisis was excessive. By going beyond what was required to offset increasing risk premia, tighter monetary policy appears to have contributed to the collapse of the exchange rates when they came under speculative attack.

Our empirical findings are robust in the sense that we have taken care to address two fundamental economic problems that have plagued the empirical literature on this important policy issue. First, we have taken into account the simultaneous feedback between exchange rates and interest rates by specifying a VECM model and by utilising an appropriate identification procedure due to Sentana and Fiorentini (1999), which exploits the presence of heteroscedasticity in the time series under investigation. Second, we have considered a longer time period than other studies, which focus on the crisis period only. This has enabled us to compare the relationship

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conditional variance to unity. Furthermore, the results do not change if we adopt an IGARCH

between exchange rates and interest rates during tranquil periods with that during more turbulent periods.

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specification for the other countries as well.

Figure 1: Thailand int. rate and exch. rate

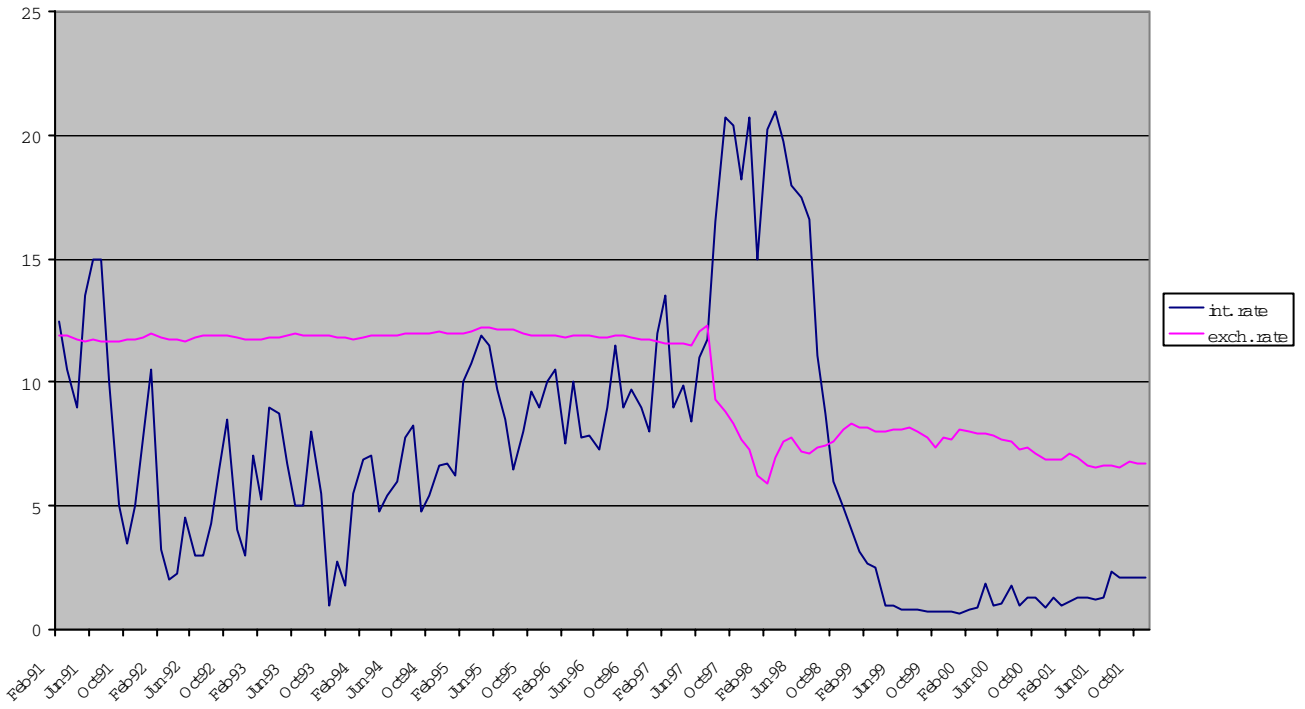


Figure 2: Korea int. rate and exch. rate

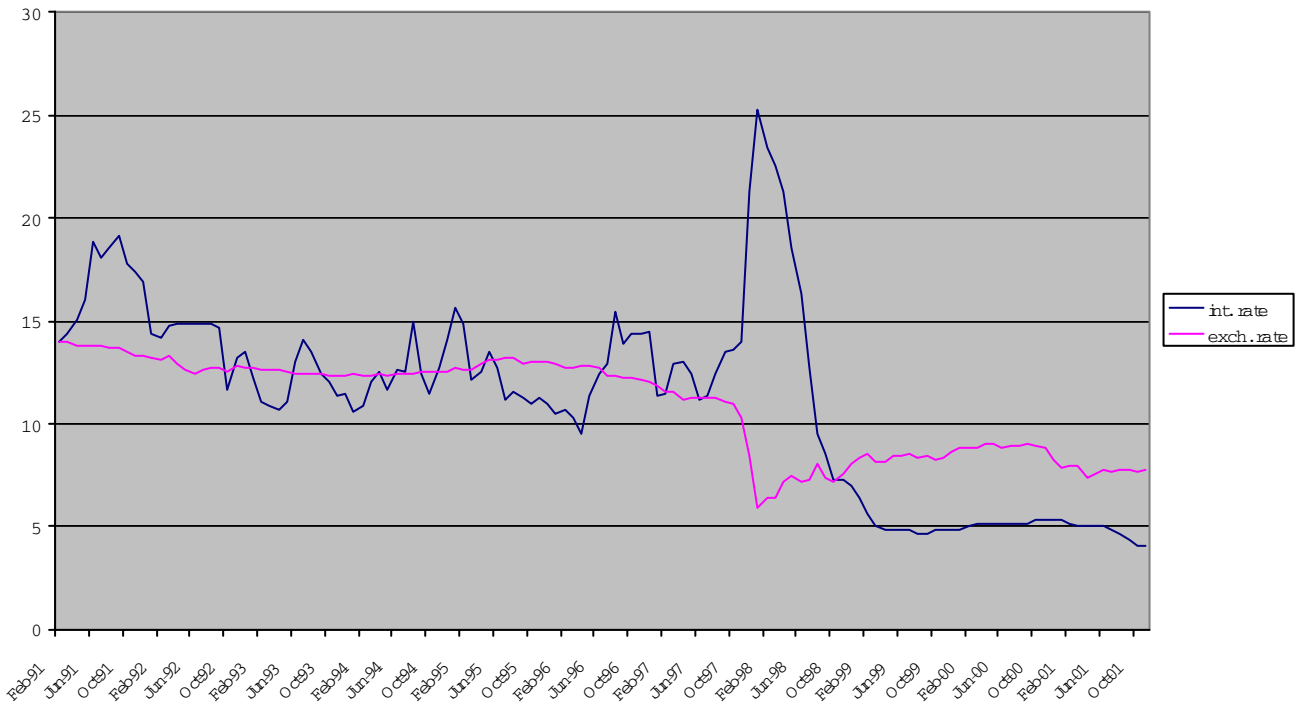


Figure 3: Indonesia int. rate and exch. rate

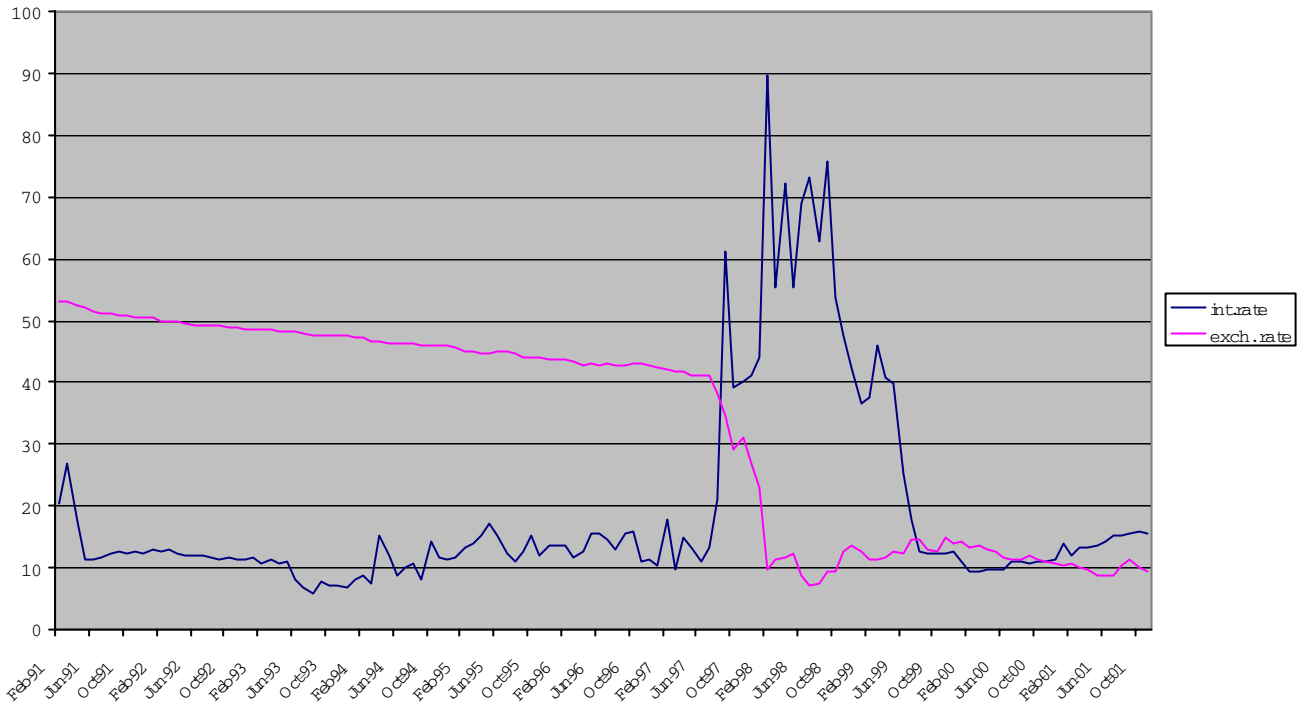


Figure 4: Philippine int. rate and exch. rate

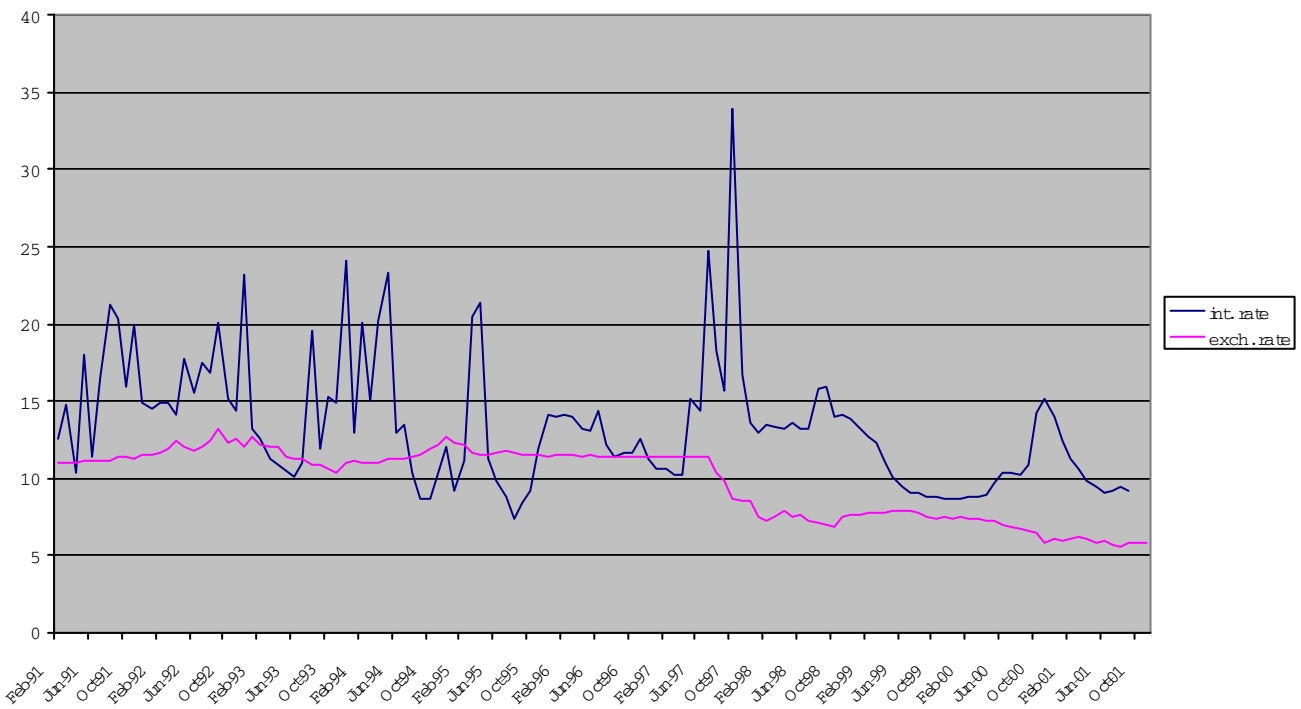




Table 1: Estimation results

	Thailand	Korea	Indonesia	Philippines
$\alpha_2$	-0.38 (6.09)	-0.14 (13.24)	-0.41 (7.25)	-0.20 (2.45)
$\beta_{12}$	0.10 (1.85)	0.12 (4.74)	0.03 (3.25)	0.15 (3.09)
$\beta_{12D}$	-4.10 (12.35)	-1.41 (14.06)	-0.34 (5.44)	-0.56 (2.41)
$\beta_{21}$	0.07 (2.73)	0.01 (2.64)	0.04 (1.42)	-0.08 (1.07)
$\beta_{21D}$	-0.11 (2.94)	-0.18 (26.33)	-0.39 (3.48)	-0.34 (1.98)
$\gamma_{11}$	0.60 (6.48)	0.71 (6.00)	0.27 (5.35)	0.61 (15.18)
$\gamma_{12}$	0.36 (4.39)	0.28 (6.69)	0.63 (12.97)	0.36 (9.22)
$\gamma_{21}$	0.84 (20.15)	0.48 (14.87)	0.66 (7.43)	0.55 (9.22)
$\gamma_{22}$	0.13 (3.08)	0.43 (6.06)	0.32 (5.60)	-

Note: T-ratios (adjusted for the presence of residual correlation) are in parentheses. In the Philippines, the conditional variance for the interest rate equation has been modelled as an IGARCH, as the sum of  $\gamma_{21}$  and  $\gamma_{22}$  had previously been found to exceed unity.

Table 2: Diagnostic tests on the residuals

	Thailand	Korea	Indonesia	Philippines
$LB_1(10)$	0.06	0.01	0.58	0.34
$LB_2(10)$	0.36	0.33	0.09	0.39
$LB^2_1(10)$	0.61	0.31	0.10	0.18
$LB^2_2(10)$	0.80	0.78	0.08	0.77
Ep-Stat <sub>i</sub>	0.00	0.00	0.00	0.00
Ep-Stat <sub>2</sub>	0.14	0.00	0.00	0.00

Note: The diagnostics are computed for the standardized residuals  $\varepsilon_i$  (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 same test 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 the  $i^{\text{th}}$  equation of the VECM.

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