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

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MONETARY POLICY AND THE EXCHANGE RATE

DURING THE ASIAN CRISIS:

IDENTIFICATION THROUGH HETEROSCEDASTICITY

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Abstract

This paper examines whether a monetary policy tightening (i.e., an increase in the domestic interest rate) was successful in defending the exchange rate from speculative pressures during the Asian financial crisis. We estimate a bivariate VECM for four Asian countries, and improve upon existing studies in two important ways. First, by using a long data span we are able to compare the effects of an interest rate rise on the nominal 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 time series, following Rigobon (2002). We find that while tight monetary policy helped to defend the exchange rate during tranquil periods, it had the opposite effect during the Asian crisis.

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 a 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 Rigobon (2002)[1]. This method enables us to address the endogeneity of interest rates, a thorny econometric

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[1] See also Sentana and Fiorentini (2001) for a similar identification methodology.

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

2.1 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 to 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 macromodel 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. [2]

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. 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 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

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

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, 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 Kamin (1999) use Granger causality tests 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 tests of 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.

2.2 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.

Perhaps the most important empirical challenge is the identification of monetary policy exogenous shocks as distinct from monetary policy actions (see also Kraay, 1999[3]). 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 rate rise on the exchange rate by taking into account the heteroscedasticity property of the time series under investigation, following the method put forward by Rigobon (2002).

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[4]. 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 Kamin (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 empirical analysis, as we believe this 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 Kraay (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

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[3] 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.

exchange rates and interest rates in tranquil and turbulent periods. Therefore, we model policy shifts by defining appropriate dummies, which are fully described in Section 4.

3. Empirical Methodology: Identification through Heteroscedasticity

Given the 2×1 vector of endogenous variables z_t, consider the structural VAR of order p:

$$A_0 z_t = B(L) z_{t-1} + \boldsymbol{e}_t \tag{1}$$

where $B(L)$ is a polynomial in the lag operator, A_0 is the 2×2 matrix which captures the contemporaneous interaction between the variables included in z_t, and ε_t is the vector of structural innovations. It is usually assumed that the covariance matrix of the structural innovations Γ is diagonal, e.g. that the structural shocks are orthogonal to each other. Furthermore, a normalisation to unity of the elements of the main diagonal of A_0 is imposed. The corresponding reduced form of the model in (1) is:

$$z_t = C(L) z_{t-1} + \boldsymbol{n}_t \tag{2}$$

where $C(L) = A_0^{-1} B(L)$ and $v_t = A_0^{-1} \boldsymbol{e}_t$. The covariance matrices of the reduced form innovations v_t is Σ. If the residuals are homoscedastic, then the system:

$$\Sigma = A_0^{-1} \Gamma A_0^{-1} \tag{3}$$

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[4] 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.

provides three covariance equations (given the symmetry of the covariance matrix) and four unknowns, namely the off-diagonal elements of A_0 and the variances of the two structural innovations, and hence the system (1) is not identified. More generally, for an $n \times 1$ vector of endogenous variables, the set of restrictions described above is equal to $n(n+1)/2$ 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 (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 methodology of Rigobon (2002) (see also Sentana and Fiorentini, 2001 for a similar identification scheme)[5], which enables us to identify a bivariate (cointegrating) VAR model including interest rates and exchange rates only

by exploiting the heteroscedastic time series properties of the two financial series. The intuition behind this approach is that hederoskedasticity adds equations to the system, thereby allowing the number of unknowns to match the number of equations, which enables us to tackle the endogeneity bias:

$$\Sigma_s = A_0^{-1} \Gamma_s A_0^{-1} \tag{4}$$

In this case, due to heteroscedasticity, the covariance matrices of the structural and reduced form shocks are time variant (hence the subscript s in (4), denoting a different regime or time period). The time varying covariance matrix of the structural form shocks is also assumed to be diagonal (e.g. the structural form innovations are restricted to be orthogonal to each other). For a 2×1 vector of endogenous variables z_t with a shift in the variance across two regimes, (e.g. s = 1,2), the system in (4) provides six covariance equations (three in each period) and six unknowns (two are given by the off-diagonal elements of A_0 and the remaining four are given by the variances of the two structural shocks in each regime), and hence identification of (1) is achieved (through the order condition). [6]

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

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[5]For empirical applications of identification through heteroscedasticity, see King, Sentana, Wadhwani (1994), Normandin and Phaneuf (1997) and Rigobon (2002).

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.[7] In each case the conditional mean equations are specified as a VECM:

$$A_0 z_{0t} = A_1 z_{1t} + A_2 x_{2t} + a * ect_{t-1} + e_t \tag{5}$$

where:

$$A_0 = \begin{bmatrix} 1 & -a_{01} \\ -a_{02} & 1 \end{bmatrix}; z_{0t} = \begin{bmatrix} \Delta xr_t \\ \Delta i_t^d \end{bmatrix}$$

The two endogenous variables in the vector z_t are a proxy for the nominal exchange rate depreciation rate (in percent values), that is $100 \times \Delta xr_t$, where Δxr_t is the first-order difference of the log of the nominal bilateral exchange rate (with respect to the US dollar), and Δi_t^d the first-order difference of the domestic short-term interest rate.

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[6] The rank condition states that shifts in the variance of the structural shocks must not to be proportional (for details, see Rigobon, 2002).
[7] An augmented Dickey Fuller test was carried out. Results are available upon request.

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 D_1. 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.

hence in (3) we can define:

$$A_1 = \begin{bmatrix} 0 & a_{11} \\ a_{12} & 0 \end{bmatrix}; z_{1t} = \begin{bmatrix} D_{1t} * \Delta xr_t \\ D_{1t} * \Delta i_t^d \end{bmatrix}$$

Our main focus is on the estimation of the coefficients a_{01} and $(a_{01} + a_{11})$ which measure the contemporaneous effect of an increase in the domestic interest rate on the nominal exchange rate during a period of calm and of crisis, respectively. We are also interested in the estimation of the coefficients a_{02} and $(a_{02} + a_{12})$ which measure the contemporaneous response of the domestic interest rate to a nominal exchange rate increase (e.g. appreciation) during a period of calm and of crisis, respectively.

The US federal funds rate, i^f_t, is treated as a strictly exogenous regressor, and in (3) we can define:

$$A_2 = \begin{bmatrix} a_{21} \\ a_{22} \end{bmatrix}; \ x_{2t} = \Delta i^f_t$$

In the first stage of the empirical analysis, it was found that the interest parity condition holds in the long-run in Indonesia and the Philippines once we take into account an increase in the mean (captured by the intervention dummy D_1 described above) during the crisis period. 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 (downward) shift in the mean is taken into account. Thus, an additional intervention dummy (D_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 intercept shifts are usually interpretable as changes in the country risk premium (see Obstfeld and Rogoff, 1996). In light of the discussion above, in (3) $a = \begin{bmatrix} a_1 \\ a_2 \end{bmatrix}$ is a 2x1 speed of adjustment coefficient matrix and the estimated long-run equilibrium relationships defined by ect_{t-1} in (5) are (t-ratios in parentheses):

Thailand: $\quad (i^d_t - i^f_t - \underset{(6.10)}{2.62} - \underset{(9.27)}{11.81} * D_{1t} + \underset{(8.68)}{6.60} * D_{2t})$

Korea: $\quad (i^d_t - i^f_t - \underset{(10.34)}{7.39} - \underset{(4.06)}{12.15} * D_{1t} + \underset{(6.18)}{8.15} * D_{2t})$

Indonesia: $\quad (i^d_t - i^f_t - \underset{(5.32)}{7.08} - \underset{(12.02)}{37.10} * D_{1t})$

Philippines: $(i_t^d - i_t^f - \underset{(11.83)}{7.66} - \underset{(4.91)}{34.19} * D_{1t})$

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 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 in both these countries (measured as the sum of the coefficient on the intercept and the one associated with the Dummy D_{2t}) seem to have turned negative overall, around -4% and -1% respectively.

Since we allow a shift the structural form slope coefficients, the system given in (4) now becomes:

$$\Sigma_s = A_0^{-1}\Gamma_s A_0^{-1} + A_1^{-1}\Phi A_1^{-1} \qquad\qquad (4')$$

where Φ is the covariance matrix for the shocks to the variables $D_{1t} * \Delta xr_t$ and $D_{2t}\Delta i_t$. These shocks are orthogonal to each other and homoscedastic (given the fact that the dummy captures the dynamics of the nominal interest rate and of the nominal exchange rate only during the crisis period), with variances normalised to unity. Hence in the system given by (4') we have two extra unknowns, i.e. the coefficients a_{11} and a_{12}, which implies that heteroscedasticity alone is not sufficient to identify the simultaneous equation system.

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

$$L_t = -\frac{1}{2}\log|\Gamma_t| - \frac{1}{2}\boldsymbol{e}'_t(\Gamma_t)^{-1}\boldsymbol{e}_t \tag{6}$$

where $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t})'$ is the vector of structural innovations. To explicitly recognize the existence of heteroscedasticity we use the following GARCH(1,1) specification for the variance for the ith equation (with i = 1,2): [8]

$$\boldsymbol{s}^2_{i,t} = (1 - \boldsymbol{g}_{1i} - \boldsymbol{g}_{2i}) + \boldsymbol{g}_{1i}\boldsymbol{e}^2_{i,t-1} + \boldsymbol{g}_{2i}\boldsymbol{s}^2_{i,t-1} \tag{7}$$

where the constraints $\gamma_{ij} \geq 0$ and condition $\gamma_{i1} + \gamma_{i2} < 1$ ensure non-negative variances, and allow for covariance stationary variances, respectively. In (7) heteroscedasticity is modelled though shifts in the conditional variances. The normalisation to unity of both unconditional variances (see King, Sentana, and Wadhwani, 1994, and Normandin and Phaneuf, 1997 for an application) adds the two additional restrictions which solve the system given by (4') and, consequently, identify the system given by (5).

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 Bollerslev and Woodlbridge, 1992) estimator was used in

order to obtain robust standard errors, given the evidence of non-Gaussian standardized residuals.

5. Estimation and Empirical Results[9]

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[10]. 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[11]. 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.[12] 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 computed by implementing a Newey – West (1987) type of correction for the presence of residual autocorrelation.

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[8] 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).

[9] The empirical analysis has been carried using the RATS software.

[10] 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.

[11] 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.

[12] 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 conditional variance to unity. Furthermore, the results do not change if we adopt an IGARCH specification for the other countries as well.

The results reported in Table 1 confirm the presence of cointegration, as the error correction coefficient, α_2, is found to be negative and statistically significant in the estimated domestic interest rate equations. They also suggest that there was a significant monetary policy contraction during turbulent periods, since the sum of the coefficients a_{02} and a_{12} is negative in all cases, and a_{12} 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 provides clear evidence of a nominal exchange rate appreciation in response to a domestic interest rate increase during tranquil periods. This is indicated by the coefficient a_{01}, 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 shown by the sum of the coefficients a_{01} and a_{11}, which is clearly negative. Note that a_{11} is much larger in absolute terms than a_{01}, and that it is highly significant in all four cases, suggesting that the effects of a 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 econometric 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 Rigobon (2002), 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 between exchange rates and interest rates during tranquil periods with that during more turbulent periods.

References

Bagliano, F., Favero, C., 1999. Informations from financial markets and VAR measures of monetary policy. European Economic Review 43, 825-837.

Backus D., Driffill, J., 1985. Rational expectations and policy credibility following a change in regime. Review of Economic Studies 52 (2), 211-221.

Basurto, G., Ghosh, A., 2000. The interest rate-exchange rate nexus in the Asian crisis countries. IMF Working Paper. no. 00/19.

Bensaid, B., Jeanne, O., 1997. The instability of fixed exchange rate systems when raising the nominal interest rate is costly. European Economic Review 41 (8), 1461-78.

Bernanke, B., 1986. Alternative explanations of the money-income correlation. in K. Brunner, K., Metzler, A.H. (Eds.), Real Business Cycles, Real Exchange Rates, and Actual Policies. Carnegie-Rochester Series on Public Policy 25, North-Holland, Amsterdam, 49-99.

Bernanke, B., Gertler, M., 1999. Monetary policy and asset price volatility", in New Challenges for Monetary Policy. Proceedings for the symposium sponsored by the Federal Reserve Bank of Kansas City.

Bollerslev, T., Chou, R., Kroner, K. 1992. ARCH modelling in finance. Journal of Econometrics 52, 5-59.

Bollerslev, T., Wooldridge, T.,1992. Quasi Maximum Likelihood Estimation and Inference in Dynamic Models with time varying covariances. Econometric Reviews 11, 143-172.

Cho, D., West, K.W., 2001. Interest rates and exchange rates in the Korean, Philippine and Thai exchange rate crises. in Frenkel, J., Dooley. M. (Eds.), Management of Currency Crises, NBER, forthcoming.

Dekle, R., Hsiao, C., Wang, S., 1998. Do high interest rates appreciate exchange rates during crises? The Korean evidence. Mimeo, Department of Economics, University of Southern California.

Doornik, J.A., Hansen, H. 1994. An omnibus test for univariate and multivariate normality. Nuffield College Working Paper.

Drazen, A., Masson, P., 1994. Credibility of policies versus credibility of policymakers. Quarterly Journal of Economics 109 (3), 735-54.

Feldstein, M., 1998. Refocusing the IMF", Foreign Affairs 77, 20-33.

Furman, J., Stiglitz, J.E., 1998. Economic crises: evidence and insights from East Asia. Brookings Papers on Economic Activity 1-135.

Gertler, M., Gilchrist, S., Natalucci, F., 2000. External constraints on monetary policy and the financial accelerator. New York University Working Paper.

Goldfajn, I., Baig, T. 1998. Monetary policy in the aftermath of currency crises: the case of Asia. Mimeo, IMF.

Goldfajn, I., Gupta, P., 1999a. Overshooting and reversals: the role of monetary policy. Mimeo, IMF.

Goldfejn, I., Gupta, P., 1999b. Does monetary policy stabilize the exchange rate following a currency crisis. IMF Working Paper, no. 99/42.

Gould, D., Kamin, S.B., 1999. The impact of monetary policy on exchange rates during financial crisis. Paper presented at the 1999 Pacific Basin Conference, San Francisco, September.

Kim, S., Roubini, N., 2000. Exchange rate anomalies in the industrial countries: a solution with a structural VAR approach. Journal of Monetary Economics 45, 561-86.

King, M., Sentana, E., Wadhwani, S.B., 1994. Volatility and links between national stock markets. *Econometrica* 62, 901-933.

Kraay, A., 1999. Do high interest rates defend currencies during speculative attacks?. Mimeo. World Bank.

Krugman, P., 1998. Will Asia bounce back?. Internet Website http://.web.mit.edu.

Newey, W., West, K., 1987. A simple positive-definite heteroscedasticity and autocorrelation consistent covariance matrix. Econometrica 55, 703-708.

Normandin, M., Phaneuf, L. 1997. The liquidity effect: testing identification conditions under time varying conditional volatility. Centre for Research and Economic Fluctuations and Employment, Working Paper no. 41.

Obstfeld, M., 1994. The logic of currency crises. Cahier economiques et monetaires, 43, 189-214.

Obstfeld, M., Rogoff, K., 1996. Foundations of International Macroeconomics. Cambridge, MIT Press.

Ohno, K., Shirono, K., Sisli, E., 1999. Can high interest rates stop regional currency falls?. ADB Institute, Working Paper Series, no. 6.

Radelet, S., Sachs, J., 1998. The East Asian financial crisis: diagnosis, remedies, prospects. Brookings Papers on Economic Activity 1, 1-74.

Rigobon, R., 2002. Identification through heteroscedaticity. Review of Economics and Statistics, forthcoming.

Sentana, E., Fiorentini, G., 2001. Identification, estimation and testing of Conditionally Heteroskedastic Factor Models. Journal of Econometrics 102, 143-164.

Smets, F., 1996. Measuring monetary policy in the G7 countries: interest rates versus exchange rates. Paper presented at CEPR-Banca d'Italia workshop on "Model Specification, Identification and Estimation in Empirical Macroeconomics".

Smets, F., 1997. Measuring monetary policy shocks in France, Germany and Italy: the role of the exchange rate. BIS Working Paper, no. 42.

Stiglitz, J.E., 1999. Interest rates, risk, and imperfect markets: puzzles and policies. Oxford Review of Economic Policy 15 (2), 59-76.

Tanner, E., 1999. Exchange market pressure and monetary policy: Asia and Latin America in the 1990s. IMF Working Paper, no. 99/114.

Toda, H., Yamamoto, T., 1995. Statistical inference in vector-autoregressions with possibly integrated processes. Journal of Econometrics, 66, 225-250.

Table 1: Estimation results

	Thailand	Korea	Indonesia	Philippines
α_2	-0.38 (6.09)	-0.14 (13.24)	-0.41 (7.25)	-0.20 (2.45)
a_{01}	0.10 (1.85)	0.12 (4.74)	0.03 (3.25)	0.15 (3.09)
a_{11}	-4.10 (12.35)	-1.41 (14.06)	-0.34 (5.44)	-0.56 (2.41)
a_{02}	0.07 (2.73)	0.01 (2.64)	0.04 (1.42)	-0.08 (1.07)
a_{12}	-0.11 (2.94)	-0.18 (26.33)	-0.39 (3.48)	-0.34 (1.98)
γ_{11}	0.60 (6.48)	0.71 (6.00)	0.27 (5.35)	0.61 (15.18)
γ_{12}	0.36 (4.39)	0.28 (6.69)	0.63 (12.97)	0.36 (9.22)
γ_{21}	0.84 (20.15)	0.48 (14.87)	0.66 (7.43)	0.55 (9.22)
γ_{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 γ_{21} and γ_{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
$Ep\text{-}Stat_1$	0.00	0.00	0.00	0.00
$Ep\text{-}Stat_2$	0.14	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 same test for the squared standardised residuals. Ep-Stat is the p-value for the normality test on the residuals (see Doornik and Hansen, 1994). The subscript i (where, i = 1,2) denotes the ith equation of the VECM.