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The Interest Rate–Exchange Rate Nexus in Currency Crises

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Sharp exchange rate depreciations in the East Asian crisis countries (Indonesia, Korea, and Thailand) raised doubts about the efficacy of increasing interest rates to defend the currency. Using a standard monetary model of exchange rate determination, this paper shows that tighter monetary policy was in fact associated with an appreciation of the exchange rate in these countries and during the Mexican currency crisis. Moreover, there is little evidence of higher real interest rates contributing to a widening of the risk premium. [JEL F31, G15, E40]

One of the more controversial elements of the stabilization programs in the East Asian crisis countries (Indonesia, Korea, and Thailand) was the stance of monetary policy. With the sharp exchange rate depreciations experienced at the onset of the crisis, standard policy prescriptions called for an immediate tightening of monetary policy.

But continued depreciation of the exchange rates—well into the stabilization programs—began to raise doubts about the efficacy of raising interest rates to defend the currency.[1] Some commentators, indeed, started suggesting that raising

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[1]IMF-supported programs began in August 1997 in Thailand, November 1997 in Indonesia, and December 1997 in Korea, while the most depreciated exchange rates were in January 1998 in Korea and Thailand and in July 1998 in Indonesia. Lane and others (1999) provides a useful summary.

interest rates, far from stabilizing the exchange rate, could actually prove counterproductive: further depreciating the exchange rate instead of appreciating it. The mechanism of this "perverse" effect is straightforward. High (presumably, real) interest rates, by causing widespread bankruptcies (or the expectation thereof), result in larger country risk premiums—so much so that the expected return to investors actually declines as interest rates increase, thus prompting even more capital flight and generating greater downward pressure on the exchange rate.[2]

Establishing whether tighter monetary policy—often taken to mean an increase in nominal interest rates—appreciates or depreciates the currency turns out to be a surprisingly difficult task. Such studies as do exist typically use regressions or vector autoregressions to correlate exchange rate movements to changes in nominal interest rates. This approach, however, runs into two main problems.[3]

First, the level of nominal interest rate is simply not a good measure of the monetary stance. To give but the starkest example, in January 1998 interest rates in Indonesia reached almost 60 percent per year (far higher than the interest rates witnessed in the other Asian crisis countries) at a time when the money supply was expanding at a *monthly* rate of 30 percent—scarcely a tight monetary stance. Second, simple time series correlations or vector autoregressions provide very little structure on the model, and their empirical performance in explaining exchange rate movements—even in the absence of a crisis—is, at best, limited. It is difficult to know what to make of a statement such as "higher interest rates are not correlated with exchange rate appreciations during the East Asian crisis" when the model is mute on what *is* driving the exchange rate.

In this paper, we propose an alternative approach to examining whether high real interest rates resulted in exchange rate depreciations. We start from the simple proposition that, as the relative price of two monies, the exchange rate should appreciate in response to a contraction of the domestic money supply. This, together with the empirical observation that in the Asian crisis countries there is a somewhat better correspondence between the exchange rate and the money supply (than between the exchange rate and interest rates), suggests that a standard monetary model may be useful for explaining the bulk of the exchange rate dynamics. This allows us to isolate the risk premium, controlling for changes in monetary policy, and permits a direct test of whether higher real interest rates are associated with a larger risk premium—and thus, *ceteris paribus*, downward pressure on the exchange rate.

[2]An important proponent of this school of thought is Joseph Stiglitz; see, for example, Furman and Stiglitz (1998).

[3]A third problem is that of policy endogeneity and causality. Interest rates were raised in East Asia precisely *because* the exchange rate was depreciating. The issue goes beyond finding appropriate instruments for interest rate policy (itself no easy task): In an environment in which policies are being set in anticipation of reactions of the exchange market, and the market-determined exchange rate embodies expectations of future policy, it becomes virtually impossible to disentangle cause from effect. Kraay (1999) reports results using an instrumental variable technique.

By measuring the monetary stance by the money supply, and by using an explicit model of exchange rate determination, our approach goes at least part of the way in addressing the methodological problems identified above. Of course, even if higher real interest rates are correlated with a larger risk premium, it does not necessarily follow that tightening monetary policy is counterproductive for stabilizing the exchange rate. The magnitude of the effect on the risk premium may be small. And, of course, there may be third factors (such as adverse political news) affecting both the real interest rate and the risk premium on the exchange rate. Nonetheless, if the findings suggest no correlation between real interest rates and the risk premium, then the possibility of the perverse effect (of tight monetary policy causing an exchange rate depreciation) can be ruled out.

We apply our methodology to the 1997 currency crises in the three Asian countries and, by way of comparison, to the 1994 Mexican crisis. Our results may be summarized briefly. We find that the pure monetary model does credibly well in explaining much of the observed exchange rate movements (though the stringent cross-equation constraints are rejected). Augmenting this framework to allow for a time-varying risk premium, we find little evidence that high real interest rates are correlated with a larger risk premium in any of the countries except Korea. Once a simple contagion variable is added to the explanatory variables of the risk premium, moreover, the significance of the real interest rate diminishes even in the case of Korea. We conclude that there is little evidence of a "perverse" effect of a monetary tightening on the exchange rate.

The remainder of the paper is organized as follows. Section I provides a brief review of the literature and an overview of exchange rate developments during the crisis. Section II lays out the methodology. Section III reports the parameter estimates of the monetary model. Section IV turns to the behavior of the risk premium. Section V concludes.

I. Background

Perhaps the most dramatic aspect of the East Asian crisis was the sharp depreciation of exchange rates. Before the crisis, the nominal exchange rates in these countries had, to varying degrees, been de facto pegged against the U.S. dollar. In July 1997, the Thai baht depreciated by 18 percent, eventually going from baht 25 to baht 54 per U.S. dollar (at its most depreciated rate, in January 1998). The initial (sharp) depreciations in Indonesia and Korea were somewhat smaller, around 12 percent (in August 1997 and November 1997, respectively), though the maximum depreciations—from, 2,400 rupiah to 15,000 rupiah per U.S. dollar; and 850 won to 1,700 won per U.S. dollar (January 1998)—were, if anything, even more spectacular.[4]

Confronted by sharply depreciating exchange rates, monetary policy had to tread warily between two objectives. Under the assumption that tighter monetary

[4]Exchange rates in Indonesia and Korea were not, in fact, pegged, and the exchange rate had already started depreciating, so the onset of the crisis in each country is not precisely defined. Below, we use August 1997 and November 1997, respectively, as the start dates of the "floating period" in Indonesia and Korea.

policy would stabilize the exchange rate, there was an obvious need to limit the currency depreciation, not least because of the large foreign currency debt exposures of the banking and corporate sectors (particularly in Thailand and Indonesia). Against this was the danger of an excessive contraction that could severely weaken economic activity. In the event, this dilemma resulted in stop-go policies, with significant declines in money growth rates occurring only in early 1998 in Korea and Thailand. In Indonesia, a deepening banking sector crisis necessitated massive liquidity injections, and the money supply grew rapidly until mid-1998.

The continued exchange rate depreciations despite (generally) rising interest rates began to raise doubts about the efficacy of raising interest rates to defend the currency. On the other hand, at least to date, direct evidence that higher interest rates—brought about by raising risk premiums—resulted in further depreciation of the exchange rate (whether in East Asia or elsewhere) has been scant.

Furman and Stiglitz (1998) identify a set of 13 episodes, in nine emerging markets, of "temporarily high" interest rates. Using a simple regression analysis, they find that both the magnitude and the duration of such interest rate hikes are associated with exchange rate *depreciation.* Though they caution that this evidence is not definitive and that its interpretation is fraught with difficulties concerning endogeneity, they conclude that it at least questions the usefulness of raising interest rates to defend the exchange rate. Conversely, Goldfajn and Baig (1999), using daily data to analyze the relationship between nominal interest rates and nominal exchange rates during the Asian crisis, find no evidence that higher interest rates resulted in weaker exchange rates—if anything, they find support for the "orthodox" relationship. Finally, Kraay (1999) uses a large panel data set to examine whether higher interest rates helped stave off speculative attacks. Importantly, he instruments for the policy endogeneity of interest rates, though he notes the difficulties in finding adequate instruments. He finds very little association (positive or negative) between raising interest rates and the outcome of the speculative attack.[5]

Overall, perhaps the most robust finding of these papers is that the interest rate–exchange rate nexus does not lend itself very easily to econometric analysis, particularly in the East Asian crisis context. Figures 1–4 show why.

Take the case of Thailand (Figure 1). Until May 1997, interest rates hovered between 8 and 15 percent, while the exchange rate remained virtually constant (the currency was de facto pegged against the U.S. dollar). From May 1997 to September 1997, higher interest rates were generally associated with continual exchange rate depreciation (the "perverse" effect), but from September 1997 to December 1997, interest rates *fell* and the exchange rate depreciated (the "orthodox" relationship). Interest rates then rose (with continued exchange rate

[5]Goldfajn and Gupta (1998) take another tack and study the behavior of nominal exchange rates in the *aftermath* of a speculative attack and, in particular, whether higher interest rates are associated with the reversal of the overshooting of the real exchange rate taking place through a nominal exchange rate appreciation rather than through higher inflation. They find that higher interest rates are indeed associated with the real appreciation taking place through the nominal exchange rate, with the important caveat that this result does not apply in countries that also suffered a banking crisis.

Figure 1. Thailand: Interest Rate and Exchange Rate

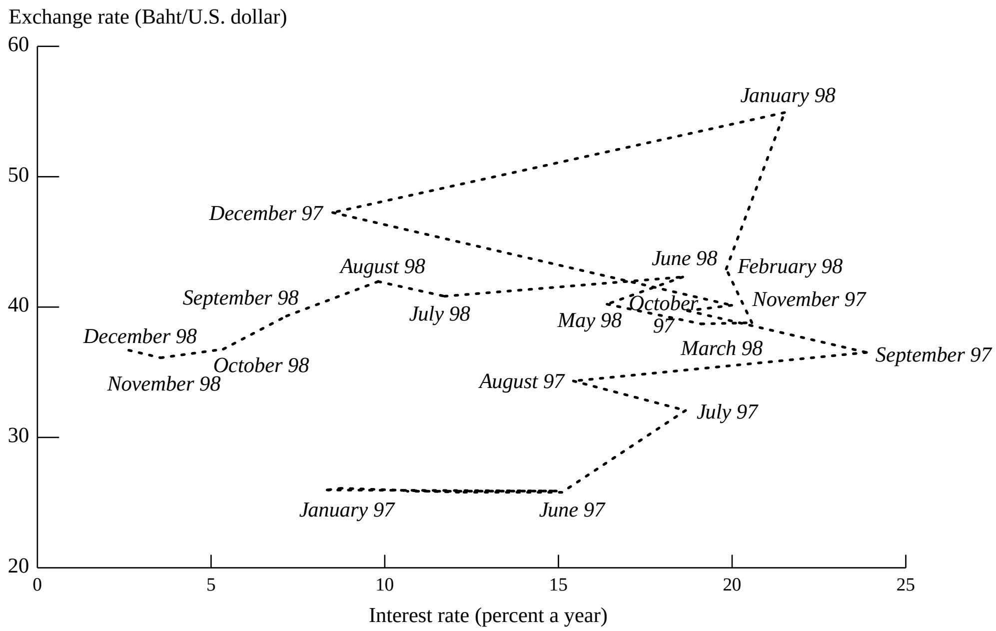


Figure 1. Thailand: Interest Rate and Exchange Rate

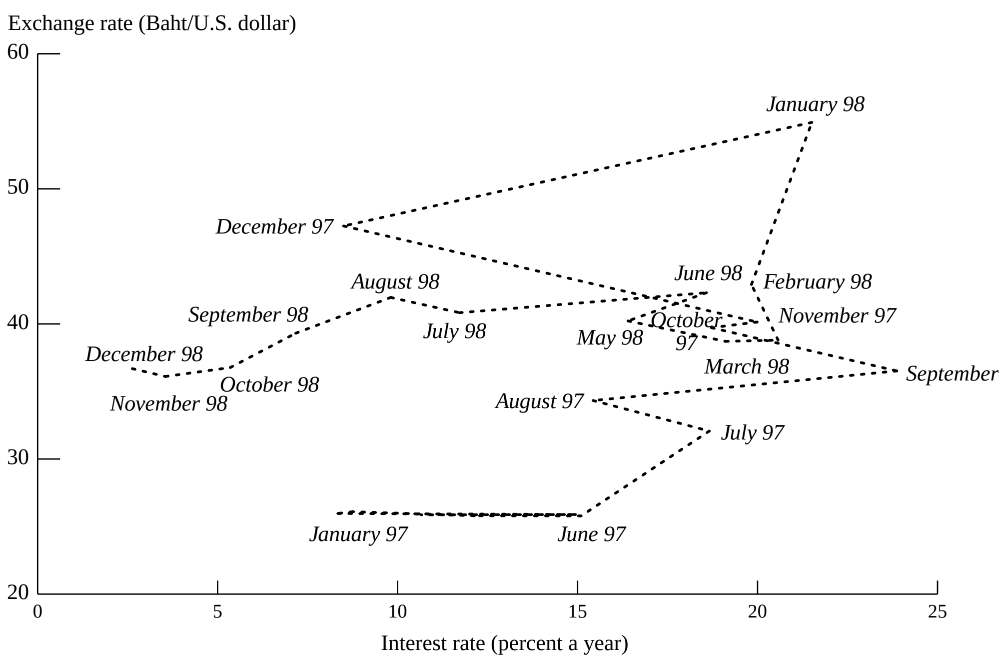


Figure 3. Korea: Interest Rate and Exchange Rate

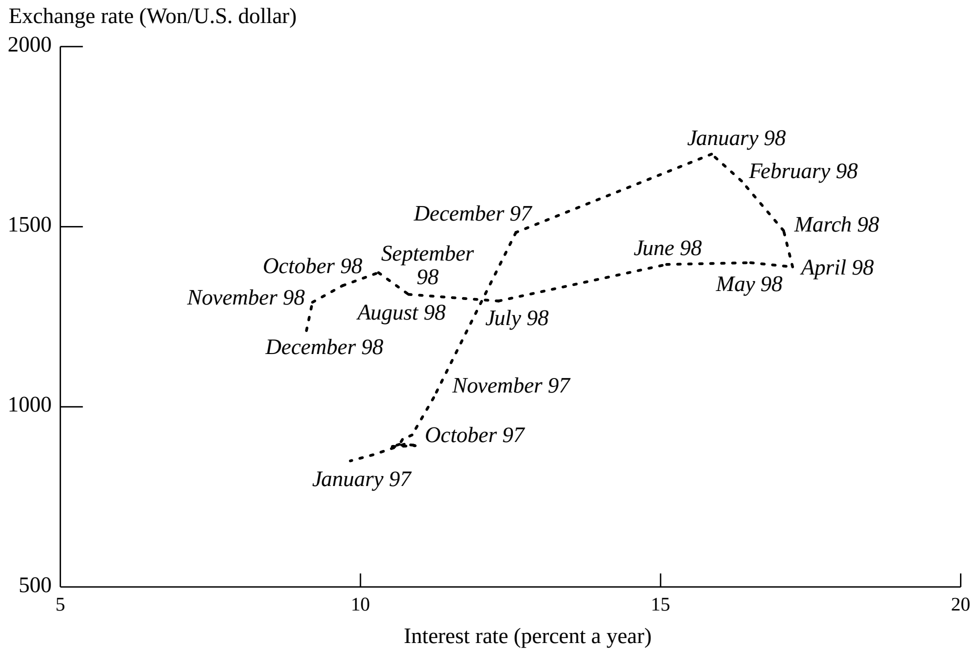
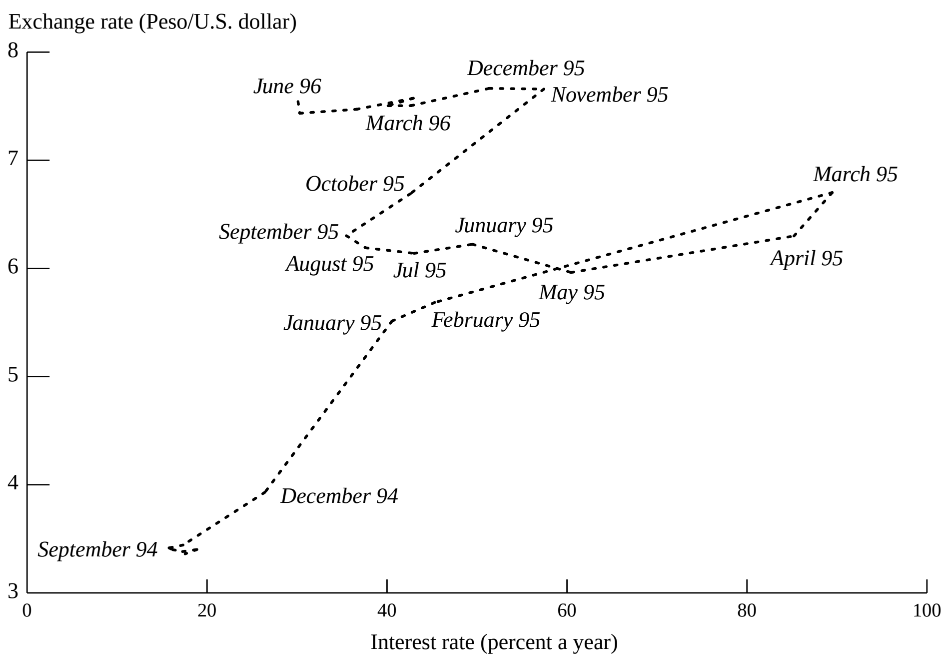


Figure 4. Mexico: Interest Rate and Exchange Rate



depreciation) until January 1998, and from January to March 1998, higher interest rates were associated with an exchange rate appreciation (the orthodox effect). Finally, since June 1998, interest rates have fallen steadily—with few ill effects on the exchange rate.

It is hard to know what to make of all this, with neither the orthodox school ("high interest rates appreciate the exchange rate") nor the "perverse" school ("high interest rates depreciate the exchange rate") receiving unequivocal support. A quick check of the other countries likewise shows periods during which interest rate and exchange rate movements were positively correlated, but also periods when higher interest rates were associated with exchange rate appreciations.

Of course, it is always difficult to know the counterfactual, and many factors other than interest rates—the availability of official external financing, debt deals with creditors, and political uncertainty—must have been impinging on the exchange rate as well. A further difficulty, alluded to above, is that interest rates often reflect risk premiums and expectations of inflation and/or depreciation and, as such, do not provide a very clear indication of the monetary stance of the country.[6]

Do monetary aggregates tell a better story? Figures 5–8 show corresponding time plots for the exchange rate and broad money supplies in these countries. For Thailand and Indonesia, the orthodox relationship—greater monetary expansion is associated with an exchange rate depreciation—comes through reasonably clearly. For Korea, the time plot is more difficult to interpret: the exchange rate obviously overshot in late 1997 and then appreciated back, but taking the period as a whole, looser monetary policy is associated with an exchange rate depreciation. Finally, for Mexico, there is again a relatively clear positive relationship between the expansion of the money supply and the depreciation of the exchange rate.

This (comparatively) stronger relationship between monetary aggregates and the exchange rate suggests an alternative approach to studying whether higher interest rates contributed to an exchange rate weakening via the risk premium, using an explicit monetary model of exchange rate determination.

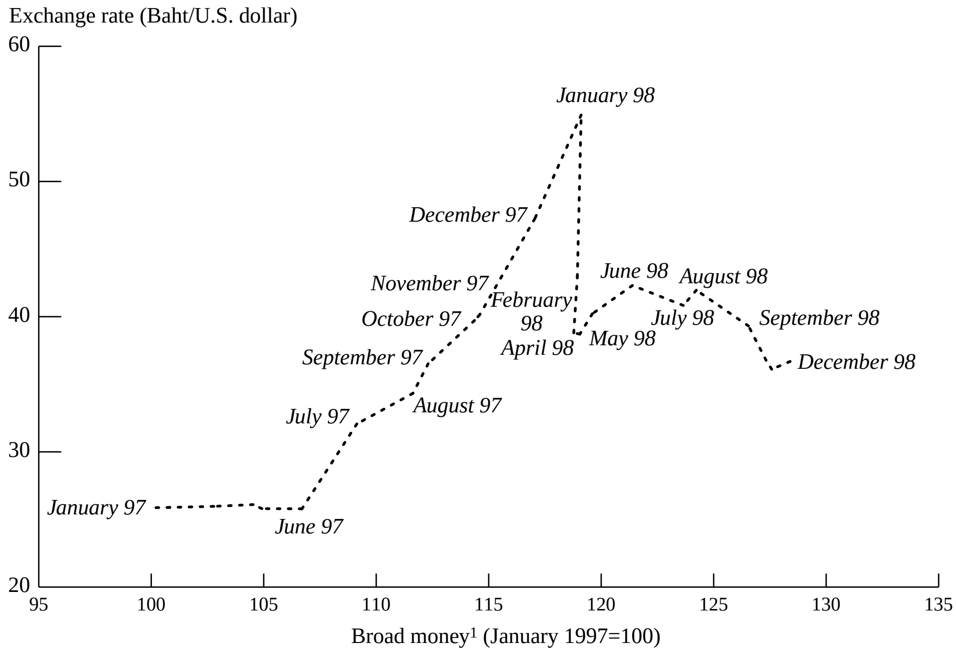
II. Methodology

The basic idea, which follows Ghosh (1992) in a similar context, is to calculate a "benchmark" exchange rate, based on a pure monetary model that abstracts from any (non-constant) risk premium.[7] The difference between the actual exchange rate and this benchmark exchange rate therefore captures the risk premium. The risk premium thus identified can then be correlated to explanatory variables, such as those capturing political events, contagion from other countries, and, in

[6]In fact, finding pure "policy" interest rates in these countries is not easy. In Korea, for instance, the so-called Bank of Korea discount rate barely moved during the crisis, and actually fell from 5.0 percent to 3.0 percent. In Indonesia, the market-determined interest rate rose to 60 percent even as broad money was expanding at a monthly rate of 30 percent, while Bank Indonesia's discount rate remained constant at 20 percent per year.

[7]The essential econometric methodology was developed by Campbell and Shiller (1987) in a somewhat different context.

Figure 5. Thailand: Broad Money and Exchange Rate



[1]Excludes valuation effects of foreign currency deposits.

Figure 6. Indonesia: Broad Money and Exchange Rate

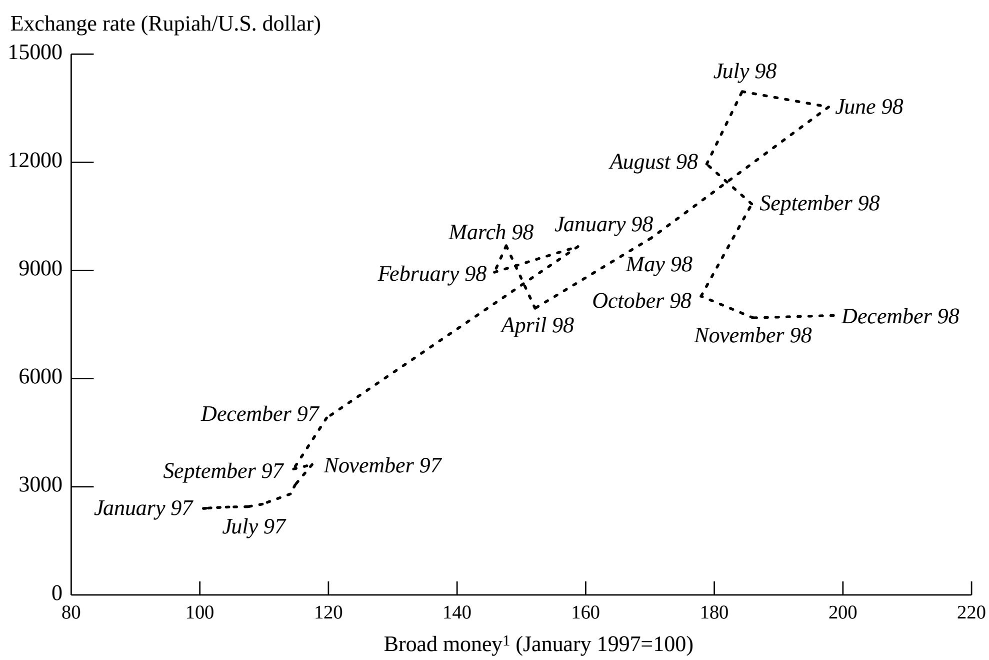
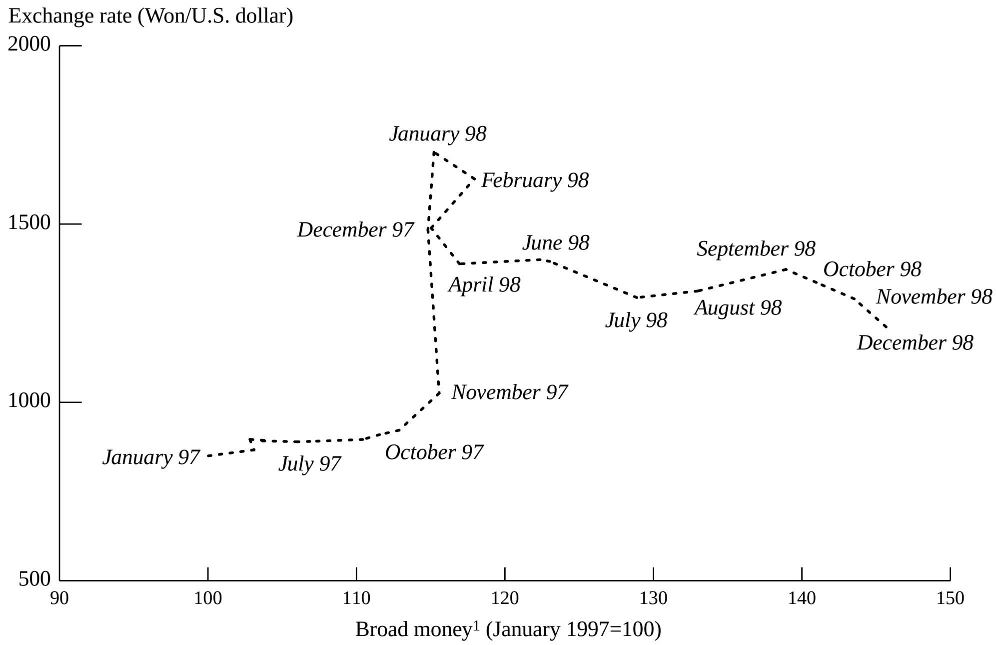
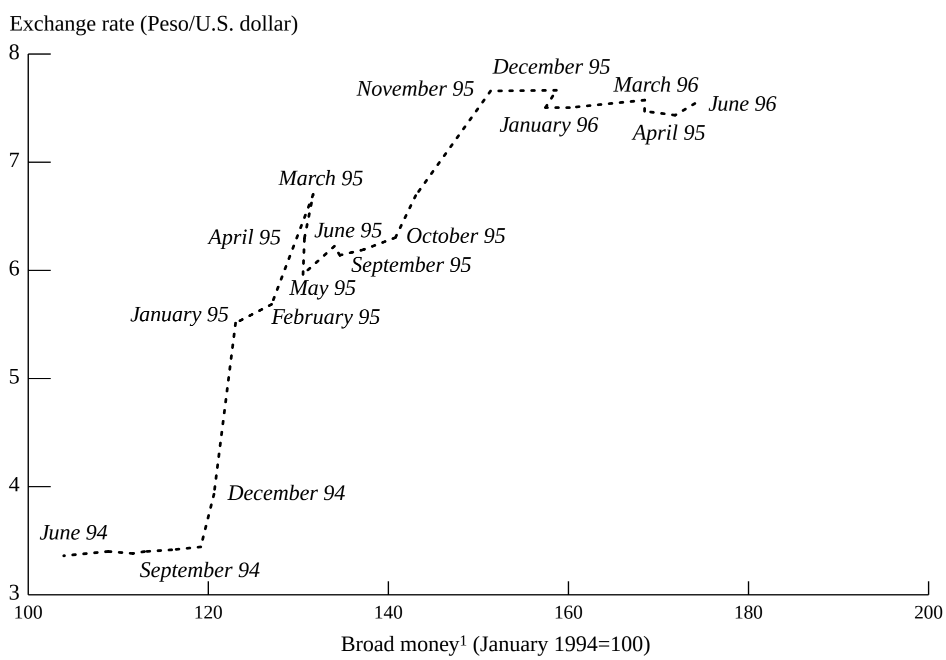


Figure 7. Korea: Broad Money and Exchange Rate



[1]Excludes valuation effects of foreign currency deposits.

Figure 8. Mexico: Broad Money and Exchange Rate



[1]Excludes valuation effects of foreign currency deposits.

particular, the level of real interest rates. If indeed high real interest rates are expected to cause widespread bankruptcies—and through this mechanism to exert downward pressure on the exchange rate—then they should be positively correlated with the risk premium.

To fix ideas, consider the case of Thailand (Figure 5). Between January 1997 and January 1998, broad money (excluding valuation effects) expanded by roughly 20 percent. The simplest "monetary" model (abstracting from changes in money demand, the foreign money supply, or expectations) would suggest a roughly commensurate depreciation of the exchange rate—say, 20 percent as well. This, then, would be the "benchmark" exchange rate. Meanwhile, the actual depreciation (change in the log) of the exchange rate was on the order of 75 percent—suggesting an "excess" depreciation of about 55 percent. This excess depreciation may be attributed (definitionally) to a widening risk premium. Our methodology thus consists of first calculating a benchmark exchange rate (using a somewhat more sophisticated monetary model) and then trying to correlate the excess depreciation to the rise in real interest rates (as suggested by Furman and Stiglitz).

Two points bear noting. First, the risk premium that this methodology identifies probably comes relatively close to the "credit" risk premium emphasized by Furman and Stiglitz. In particular, and as explained below, under the null hypothesis that the model is correct, the risk premium identified here controls for expectations of future monetary growth based on agents' *entire* information set. For instance, if there is an expectation of looser monetary policy (perhaps because of adverse political developments or the onset of a banking crisis), this is controlled for in identifying the risk premium.[8]

Second, the underlying framework is in the spirit of the monetary model of exchange rate determination. This model, a workhorse of exchange rate economics, fell into disuse after its relatively poor predictive performance in the 1970s and 1980s (Meese and Rogoff, 1983). In fact, however, the model has generally performed well in times of high nominal volatility (indeed, much of the early work on this model is based on the high inflation experience of the 1920s and 1930s) and, in its modern incarnation, actually performs rather well even for low-inflation, industrialized countries (as documented in Woo, 1985, and Ghosh, 1992, among others). Ultimately, of course, the proof of the model lies in its empirical fit and, as shown in Section III below, for the crisis countries considered here, the simplicity of the model notwithstanding, its fit is remarkably good. However, we are less interested in the fit of the monetary model than in using it as a *filter*—to take out the influence of the expansion of the money supply on the exchange rate and to see whether, once this has been controlled for, higher real interest rates are correlated with the "excess" depreciation of the currency.[9]

[8]However, other shocks, such as negative shocks to money demand, will be included in the risk premium. Thus, the test proposed here is probably conservative in the sense of being more likely to find a "perverse" effect of higher interest rates on the exchange rate.

[9]In this sense, the test proposed here is similar in spirit to variance bounds tests where the precise model is of less interest than the excess movement of the variable relative to the benchmark bound.

The model consists of three basic building blocks. Real money demand is assumed to depend positively on income and negatively on the nominal interest rate:[10]

$$m - p = \alpha y - \beta i,$$

where m is the log of money, p, the log of the domestic price index, y, the log of output, and i, the domestic interest rate. Domestic and foreign interest rates are linked by an interest parity condition:

$$i_t = i_t^* + s_{t+1}^e - s_t + \pi_t,$$

where s is the log of the exchange rate (an increase in s is a depreciation), and π is the risk premium. Finally, the real exchange rate is given by[11]

$$v_t = p_t + p_t^* - s_t.$$

Solving forward for the (first difference of the) nominal exchange rate yields[12]

$$\Delta s_t = \frac{1}{1+\beta}\sum_{j=0}^{\infty}\left(\frac{\beta}{1+\beta}\right)^j E_t\left\{\Delta x_{t+j} + \Delta\pi_{t+j} \middle| \Omega_t\right\}, \tag{1}$$

where $x_{t+j} = m_{t+j} - m_{t+j}^* - v_{t+j} - \alpha(y_{t+j} - y_{t+j}^*)$.

Equation (1) is merely a statement of the monetary model of exchange rate determination. According to equation (1), faster money growth in the home country (relative to the rest of the world) leads to a depreciation of the exchange rate, while faster output growth, achieved by raising money demand, results in an appreciation. (A widening risk premium, of course, depreciates the exchange rate.) Current movements of (any component of) x_t affect the exchange rate directly, while expected future movements are discounted at the rate $[\beta/(1+\beta)]^j$.

Notice that, in this model, a monetary contraction in the home country necessarily appreciates the exchange rate (via the first term in x_t)—unless higher real interest rates happen to cause a sufficiently large increase in the (present-value) risk premium, π.

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[10]Correspondingly, for the foreign country (the United States) $m^* - p^* = \alpha y^* - \beta i^*$.

[11]There are some subtle issues concerning the treatment of the real exchange rate. Clearly, the real exchange rate was not constant following the currency crises in these countries, so purchasing power parity (PPP) cannot be imposed. On the other hand, to the extent that the real exchange rate is driven entirely by movements of the nominal exchange rate, the "fundamentals" Δx will spuriously be correlated with the nominal exchange rate movement. In both Mexico and the Asian countries, however, real exchange rate changes were large and persistent, without a return to the pre-crisis level either through nominal appreciation (once the float began) or inflation—suggesting that real factors were also at play. Because our intention here is to create a benchmark model to filter out fundamentals, we include the real exchange rate in x. As a robustness check, we report results instrumenting for Δx with its lagged value.

[12]Ghosh (1992) works with a lagged adjustment money demand function and shows that the quasi-first difference, $s_t - \lambda s_{t-1}$, should be stationary (where $0 < \lambda < 1$ is a quadratic root that depends on the money demand parameters). The estimated value of λ is typically very close to unity, however; as a simplifying approximation, therefore, we use the first difference directly.

It is useful to define the benchmark exchange rate (excluding the risk premium) by

$$\Delta s_t^* = \frac{1}{1+\beta}\sum_{j=0}^{\infty}\left(\frac{\beta}{1+\beta}\right)^j E_t\left\{\Delta x_{t+j}\middle|\Theta_t\right\}. \tag{2}$$

Then, conceptually, our test consists of correlating the difference between $\Delta s - \Delta s^*$ to the variable of interest, r_t, such as the level of real interest rates (as suggested by Furman and Stiglitz).

The actual test is somewhat different and follows Campbell and Shiller (1987), who study such present value relations extensively in a somewhat different context, and Ghosh (1992), who studies the risk premium in a monetary model of the exchange rate.

The first step in estimating the pure monetary model is to create the projection of the expected future discounted monetary policy, Δx_t. The simplest approach would be to use a univariate autoregression. However, in general, agents have much more information about the evolution of future monetary policy than would be contained in past values of Δx_t. For instance, agents may be expecting looser monetary policy based on news about political events or adverse developments in the banking sector.

In general, it is difficult to identify and capture the additional information that is being used by agents to determine the exchange rate, and the econometrician's information set, Θ, will be only a small subset of the agent's information set, Ω. As shown by Campbell and Shiller (1987), however, it is possible to include *all* the relevant information in the econometrician's information set because, under the null, the exchange rate itself embodies this additional information.[13] (As discussed by Campbell and Shiller, one implication of this is that Δs should Granger-cause Δx.) Therefore, rather than use a univariate autoregression in Δx_t a vector autoregression (VAR) is estimated in $z = \{\Delta x_t, \Delta s_t\}$: $z_t = \Phi z_{t-1} + \varepsilon_t$. This turns out to be particularly convenient for computing the infinite sum on the right-hand side (RHS) of (2) because $E_t(z_{t+k}) = \Phi^k z_t$ so that the cross-equation constraint (2) becomes

$$g' z_t = \frac{1}{1+\beta}\sum_{j=0}^{\infty}\left(\frac{\beta}{1+\beta}\right)^j h'\Phi^j z,$$

where $g = [0 \;\; 1]$ and $h = [1 \; 0]$ so that the "benchmark" exchange rate (excluding the effects of higher real interest rates on the risk premium) may be computed as[14]

$$\Delta s_t^* = [1 \quad 0]' \frac{1}{1+\beta}\left(I - \frac{\beta}{1+\beta}\Phi\right)^{-1}\begin{pmatrix}\Delta x \\ \Delta s\end{pmatrix}. \tag{3}$$

[13]The issue is important because otherwise expectations of looser monetary policy are shifted to the risk premium term (since it is the residual), and the risk premium would be capturing not only credit risk, but also the risk associated with looser monetary policy.

[14]To see this, note that $E_t \Delta x_{t+j} = [1 \; 0]'\Phi^j z_t$. Therefore, $[1/(1+\beta)] \; \Sigma \; E_t \Delta x_{t+j}(\beta/(1+\beta))^j = [1/(1+\beta)]$ $[1 \; 0]'[I-\beta/(1+\beta)\Phi]^{-1}z_t$.

Writing out equation (3) explicitly yields $\Delta s_t^* = \Gamma_1 \Delta x_t + \Gamma_2 \Delta s_t$. This "benchmark" exchange rate can be compared to the actual exchange rate. Within the framework adopted here, to the extent that the actual depreciation, Δs, exceeds the benchmark exchange rate, Δs^*, it reflects a widening risk premium.

Next, in order to examine whether the risk premium depends on some variable r (such as real interest rates), the VAR is augmented to include it: $z = \{\Delta x_t, \Delta s_t, \Delta r_t\}$. Proceeding exactly as above yields an implied exchange rate (including the effects, if any, of higher real interest rates on the risk premium):

$$\Delta s_t^{**} = [1 \quad 0 \quad 0]' \frac{1}{1+\beta}\left(I - \frac{\beta}{1+\beta}\Phi\right)^{-1}\begin{pmatrix}\Delta x \\ \Delta s \\ \Delta r\end{pmatrix}. \tag{4}$$

Again, writing out equation (4) explicitly yields $\Delta s_t^{**} = \Gamma_1 \Delta x_t + \Gamma_2 \Delta s_t - \Gamma_3 \Delta s_t$. Under the null hypothesis that higher real interest rates have no effect on the risk premium, $\Gamma_3 = 0$, which can be tested using the appropriate Wald statistic. Under the alternative, that r_t is (positively) correlated with a currency depreciation, $\Gamma_3 > 0$.[15]

III. Empirical Results

Applying the methodology outlined above raises some practical issues. Because we use broad money as our monetary aggregate, part of the expansion of the money supply may be endogenous to the exchange rate because of foreign currency deposits. To address this, we remove the valuation effect of the stock of foreign exchange deposits on the money supply (numerically, it turns out to be important only in Indonesia, where foreign exchange deposits are substantial and the exchange rate depreciation was large).

A trickier issue concerns the choice of sample period. A natural choice would be the post-float period (i.e., once the fixed exchange rate regimes were abandoned). There are two drawbacks to this choice, however. First, except in the case of Thailand and Mexico, formal pegged exchange rate regimes were not in place. In Indonesia and Korea, for example, there was a lengthy period of successive depreciations before the very sharp depreciation at end-1997 and early 1998. Hence, the post-float period is not always clearly defined. Second, the post-float period may yield to few observations for reliably estimating the VARs. Without the passage of time, there is essentially no way around this problem. We proceed by reporting results both for the period 1990:Q1–2000:Q6 and for the post-float period.[16]

We begin by making some preliminary parameter estimates for the money demand function. With high rates of monetization and substantial financial innovation

[15]The pure monetary model also has implications for the other parameters; to wit, $\Gamma_1 = 0$, $\Gamma_2 = 1$.

[16]Thailand, July 1997 onward; Indonesia, August 1997 onward, Korea, November 1997 onward. For Mexico there are enough observations to use only the post-float period (December 1994 onward). Monthly data are taken from *International Financial Statistics*: exchange rate (line af); money plus quasi-money (lines 34+35); consumer price index (line 64); lending interest rate (line 60p); and industrial production (line 66). For Thailand, industrial production was taken from the Bank of Thailand *Monthly Bulletin*, and for Indonesia, quarterly data from *Biropustat Statitistik* are interpolated. Data on foreign exchange deposits to adjust the broad money figures are taken from the central bank bulletins or websites.

in the years preceding the crises, it is often quite difficult to obtain stable parameter estimates for the money demand functions. The estimates given in Table 1, however, are of plausible magnitude and statistically significant of the expected sign (a positive income elasticity and negative interest elasticity of money demand). Moreover, as discussed in the robustness section below, the main findings turn out not to be terribly sensitive to the exact parameter values of the money demand function. For instance, the interest elasticity (usually the most difficult parameter to estimate) only enters the expression for the exchange rate as the discount factor.

We therefore proceed on the basis of the estimates given in Table 1 and, in the robustness section, test the sensitivity of our main results to variations in the money demand parameter values.

Next, we check the order of integration of *s* and *x*. As indicated in the bottom panel of Table 1, the augmented Dickey-Fuller test cannot reject the null hypothesis of a unit root for the levels of *s* and *x*, but readily do so for their first differences; therefore it is appropriate to work with Δs_t and Δx_t.

With these preliminary transformations, we estimate the vector autoregression, $z_t = \Phi z_{t-1} + \varepsilon_t$. Table 2 reports the VAR parameters for a first-order system.[17] As the model would suggest, Δs_t Granger-causes subsequent movements in Δx_t in each case except Indonesia, where the *t*-statistic on Δs_{t-1} is marginal (about 1.20).

Before turning to the formal test of whether higher real interest rates are associated with a widening of the risk premium, we can gauge the usefulness of the monetary model as a "filter" by comparing the benchmark exchange rate (3) to the actual exchange rate. Here, the model performs credibly well, with the correlation coefficient during the float period ranging from 0.67 to 0.97, and the simple time series plots given in Figures 9–12 show that the model correctly captures much of the movement in the exchange rate.

These figures are also useful for identifying periods for which the pure monetary model does *not* work—which, in the framework adopted, means periods during which there were changes in the risk premium. In Indonesia, there seems to be little left to explain. Essentially, the massive liquidity injection in December 1997/January 1998 so swamps any other developments that the pure monetary model can account for nearly all of the exchange rate depreciation. In February 1998, however, the small re-appreciation of the actual exchange rate falls short of what the pure monetary model would predict—suggesting that the risk premium widened.

In Thailand, from July 1997 to January 1998, the actual exchange rate depreciated more than the monetary model would predict, suggesting a widening risk premium, with a decrease in the risk premium starting in March 1998. In Korea the story is much the same: a very large increase in the risk premium in December 1997, which starts reversing around April 1998. Finally, in Mexico, the risk premium widened in January 1995 and again significantly in March 1995, before narrowing in May 1995.

To summarize, the pure monetary model seems to characterize movements of the exchange rates reasonably well, and it provides a credible framework to control for the direct impact of monetary aggregates on the exchange rate.

[17]The order of the VARs was chosen using the Schwartz-Bayes criterion.

Table 1. Money Demand Parameter Estimates and Unit Root Tests

	Indonesia	Korea	Thailand	Mexico
Parameter estimates[a]				
α	0.26**	0.32**	1.33**	1.78**
t-statistic	2.59	4.87	20.78	8.91
β	0.04**	0.15**	0.13**	0.12**
t-statistic	2.96	5.28	2.38	2.13
R^2	0.87	0.99	0.82	0.49
Unit root tests[b]				
x	–0.66	–0.39	–1.28	–2.29
s	–0.98	–1.11	–1.03	–0.52
Δx	–3.63**	–4.70**	–5.06**	–3.92**
Δs	–3.38**	–4.73**	–4.16**	–4.17**
M^d residual	–3.54**	–3.66**	–3.93**	–3.50**

[a]OLS estimates; asterisks denote 10 (*) and 5 (**) percent significance levels, respectively.
[b]Augmented Dickey-Fuller tests with six lags.

Table 2. VAR Parameters

	Indonesia		Korea		Thailand		Mexico	
	Coefficient	*t*-statistic	Coefficient	*t*-statistic	Coefficient	*t*-statistic	Coefficient	*t*-statistic
Dependent variable: Δx								
$\Delta x(-1)$	–0.287	–0.99	0.088	0.55	–0.204	–2.04**	–0.340	–2.31**
$\Delta s(-1)$[a]	0.485	1.21	0.300	1.80*	0.515	2.44**	0.929	4.09**
$\Delta r(-1)$	–0.183	–0.21	–1.074	–1.53	1.999	1.61	0.064	0.04
Constant	0.019	1.71*	0.010	2.63**	0.006	0.74	–0.011	–1.13
Dependent variable: Δs								
$\Delta x(-1)$	–0.178	–0.85	0.252	1.80*	0.020	0.43	0.205	2.18**
$\Delta s(-1)$	0.460	1.59	0.235	1.51	0.174	1.73*	0.069	0.47
$\Delta r(-1)$	–0.548	–0.86	–0.418	–0.68	1.103	1.86*	2.323	2.27**
Constant	0.010	1.29	0.000	–0.03	0.003	0.72	0.012	1.82*
Dependent variable: Δr								
$\Delta x(-1)$	–0.144	–5.81**	–0.067	–3.18**	–0.007	–1.08	0.012	1.12
$\Delta s(-1)$	0.108	3.16**	0.049	2.10**	–0.012	–0.85	–0.056	–3.23**
$\Delta r(-1)$	–0.132	–1.76*	–0.155	–1.67	–0.334	–3.92**	–0.177	–1.46
Constant	0.001	1.47	0.001	1.28	0.000	0.25	0.001	0.95

Note: Asterisks denote 10 (*) and 5 (**) percent significance levels, respectively.
[a]Model implies that Δs should Granger-cause Δx.

However, it is also clear that the risk premium was not constant. In the next section, therefore, we relax this assumption and, in particular, allow the risk premium to depend upon real interest rates.

IV. Determinants of the Risk Premium

Many factors account for the widening risk premiums as the crisis deepened in each country—political uncertainties, contagion effects, corporate bankruptcies, banking system problems, prospects of possible capital controls, and indeed a seemingly never-ending stream of bad news. Most of these factors are difficult to

Figure 9. Thailand: Benchmark and Actual Exchange Rate
(log first difference)

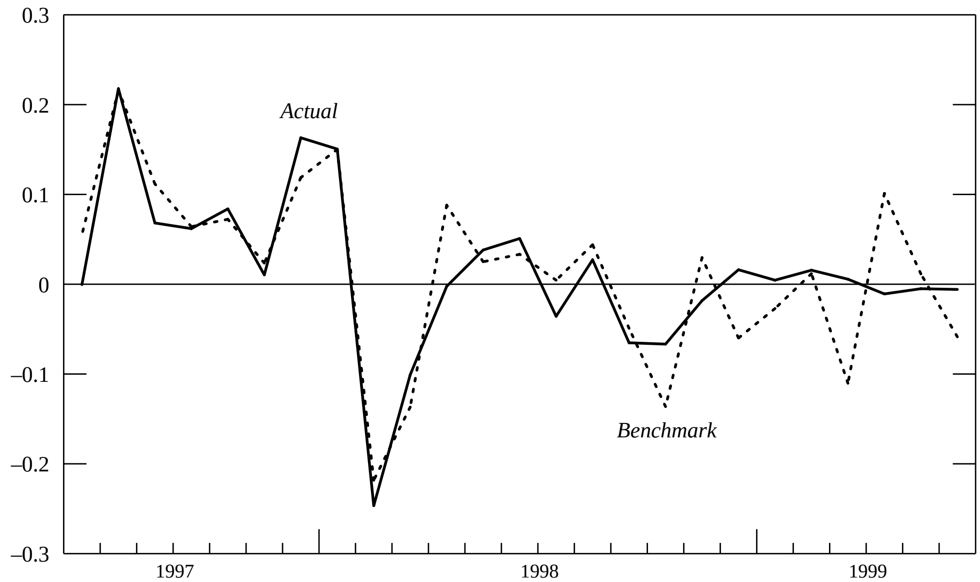


Figure 10. Indonesia: Benchmark and Actual Exchange Rate
(log first difference)

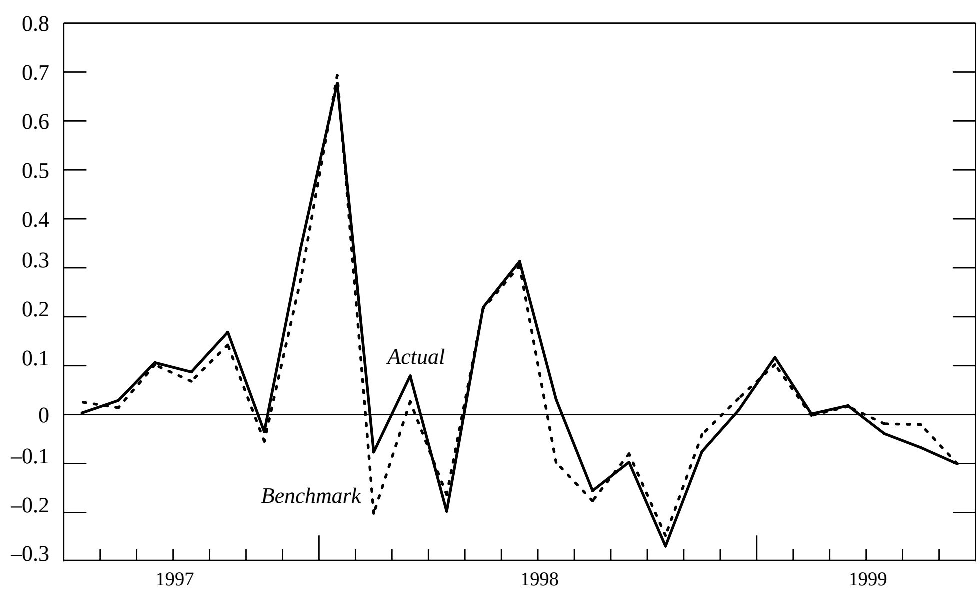


Figure 11. Korea: Benchmark and Actual Exchange Rate
(log first difference)

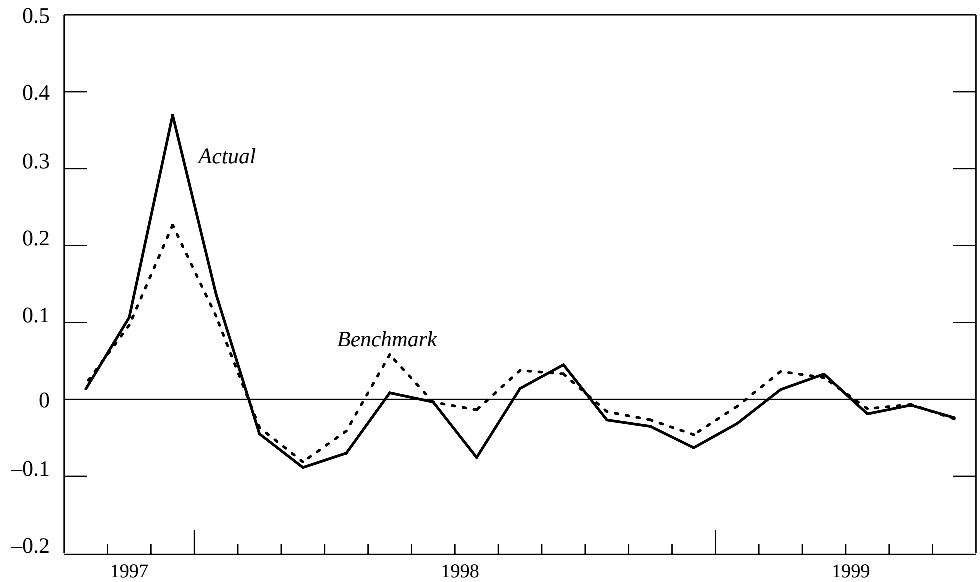
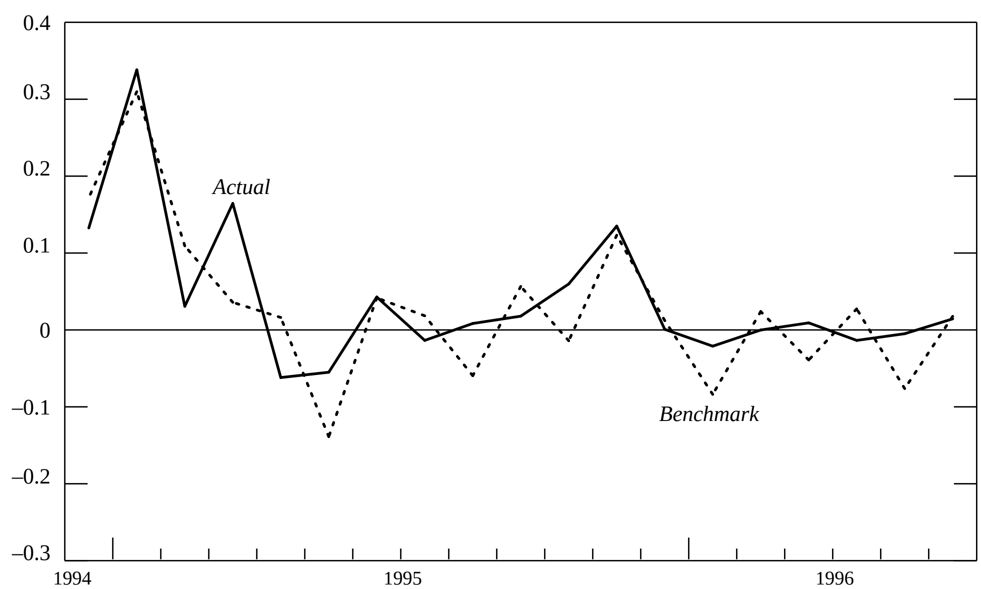


Figure 12. Mexico: Benchmark and Actual Exchange Rate
(log first difference)



capture econometrically, but to the extent that rising real interest rates contributed to widespread bankruptcies, part of the widening risk premiums may be correlated to higher real interest rates.

As discussed above, conceptually our test simply consists of regressing the difference between the actual and theoretical exchange rates on the real interest rate. Econometrically, however, it is preferable to do the estimation in a single step by augmenting the VAR to include the (change in) the real interest rate and then testing whether $\Gamma_3 = 0$ (the null), or $\Gamma_3 > 0$ (the alternate hypothesis, that higher real interest rates are associated with a widening premium).

Table 3 reports the empirical results. First, the top panel gives Wald test statistics on the pure "monetary" component of the model, based on the implied Γ coefficients from the estimated VAR parameters.[18] Although the estimates of Γ_2 are significantly different from zero (except for Indonesia), they are also significantly different from unity. Moreover, the Γ_1 coefficients are also significantly different from zero.[19]

Turning to the correlation between real interest rates and the risk premium, panel 2 of Table 3 reports the estimates of Γ_3. For Indonesia, there is a positive relation between real interest rates and the risk premium, though the coefficient is not significantly different from zero. For Thailand and Mexico, the coefficients are actually negative (suggesting, especially in the case of Mexico, that the risk premium went *down* only when investors saw higher real interest rates). Only for Korea do we find a positive and statistically significant relationship (*t*-statistic: 1.53), essentially because real interest rates started increasing around end-1997, when the risk premium also widened significantly.

But of course, this correlation between real interest rates and the risk premium does not prove that tighter monetary policy caused a widening of the risk premium. One possibility is that some other variable affected the risk premium. An obvious candidate is the contagion effect from other crisis countries in the region. To capture this, for each Asian crisis country, we simply use the unweighted average of the contemporaneous exchange rate movements in the *other* two countries. Once this variable is added to the explanatory variables of the risk premium, the real interest rate loses its significance even in the case of Korea, with the *t*-statistic falling to 1.33 (Table 3, panel 3).

Beyond their purely statistical significance, the effects are relatively small in economic terms as well. From Table 3, a 1 percentage point increase in real interest rates would be associated with a 0.1 percent depreciation of the currency. At their peak, real lending rates rose by about 10 percentage points in Korea (relative to the pre-crisis levels). Based on these estimates, the rise in real interest rates could account for less than a 1 percent depreciation of the won—a paltry effect relative to the observed depreciation.

[18]That is, we compute $\begin{bmatrix}1 & 0 & 0\end{bmatrix}' \frac{1}{1+\beta}\left(I-\frac{\beta}{1+\beta}\Phi\right)^{-1}\begin{pmatrix}\Delta x\\ \Delta s\\ \Delta r\end{pmatrix}$, then Γ_1 is the resulting coefficient on Δx, Γ_2 is the coefficient on Δs, and Γ_3 is the coefficient on Δr.

[19]Recall that the model implies $\Gamma 1 = 0$ and $\Gamma 2 = 1$. Standard errors were computed numerically as $\nabla Q' \Sigma \nabla Q$, where ∇Q is the gradient of Γ with respect to the VAR parameters, and Σ are the White-consistent standard errors.

Table 3. Cross-Equation Constraints

	Indonesia	Korea	Thailand	Mexico
Monetary model[a]				
Γ1	0.91**	0.89**	0.88**	0.87**
Standard error	0.02	0.02	0.01	0.01
Γ2	0.03	0.03*	0.05**	0.08**
Standard error	0.03	0.02	0.02	0.01
Real interest rate effect on risk premium[b]				
Γ3	0.01	0.11*	–0.18	–0.04
t-statistic	0.23	1.53	–1.66	–3.28
Real interest rate and contagion effect on risk premium[c]				
Γ3	0.01	0.10*	–0.18	. . .
t-statistic	0.26	1.33	–1.62	
Γ4	–0.16**	0.01	0.02	. . .
t-statistic	–7.27	1.53	1.48	
Correlation (Δs,ΔS^*)				
Full sample	0.86	0.95	0.50	. . .
Float period only	0.96	0.97	0.74	0.67

Note: Asterisks denote coefficients that are different from zero at the 10 (*) and 5 (**) percent significance levels, respectively.

[a]Pure monetary model, null hypothesis Γ1 =0, Γ2 = 1.

[b]If higher real interest rates result in larger risk premium, Γ3 > 0.

[c]If contagion results in larger risk premium, Γ4 > 0.

Robustness

To check for robustness regarding our main finding—that higher real interest rates are not particularly associated with a widening risk premium—we consider a number of alternative specifications.

As noted above, one issue concerns the choice of sample period. Specifications 1 and 2 of Table 4 repeat the analysis, but restrict the sample to the post-float period (for the Asian countries). This makes little difference to the results. One possibility is that the perverse effect of tighter monetary policy was purely a crisis phenomenon. In specification 3, therefore, we restrict the sample to the one-year period following the collapse of the fixed regime and the adoption of the float; again, the results differ little.

Specifications 4–7 vary the money demand parameters to within one standard error of the point estimates given in Table 1 above, while specifications 8 and 9 use an alternative interest rate (usually the deposit or money market rate rather than the lending rate) or an alternative deflator (the WPI instead of the CPI); specification 10 instruments for Δx using its lagged value; specification 11 includes the estimated residuals from the money demand functions in the forcing variable Δx. The estimated Γ_3 coefficient is generally not significantly positive (except for some of the specifications for Korea, which are borderline significant with *t*-statistics around 1.40).

Finally, beyond the risk premium effect emphasized by Furman and Stiglitz (1998), higher real interest rates could also affect the exchange rate by depressing

Table 4. Robustness Checks[a]

	Indonesia	Korea	Thailand	Mexico
[1] Float period, excl. contagion				
Γ3	0.03	0.19	–0.54	
t-statistic	0.22	0.94	–2.20	. . .
[2] Float period, incl. contagion				
Γ3	0.03	0.10	–0.52	. . .
t-statistic	0.27	0.49	–2.13	. . .
[3] Crisis year only				
Γ3	–0.01	0.04	–0.24	–0.71
t-statistic	–0.08	0.11	–0.88	–28.71
[4] Low α				
Γ3	0.01	0.09	–0.18	–0.04
t-statistic	0.19	1.26	–1.66	–3.28
[5] High α				
Γ3	0.01	0.10*	–0.18	–0.07
t-statistic	0.26	1.39	–1.58	–5.63
[6] Low β				
Γ3	0.01	0.09*	–0.15	–0.03
t-statistic	0.27	1.34	–1.62	–3.14
[7] High β				
Γ3	0.01	0.10*	–0.21	–0.05
t-statistic	0.26	1.33	–1.62	–3.63
[8] Deposit or money market interest rate				
Γ3	0.02	0.10*	–0.18	–0.01
t-statistic	0.37	1.40	–1.65	–0.92
[9] Real interest rate deflated by wholesale price index				
Γ3	0.00	0.06	–0.18	–0.19
t-statistic	0.71	1.12	–2.86	–11.64
[10] Lagged Δ*x* instrument				
Γ3	0.01	–0.01	0.03	0.00
t-statistic	0.58	–0.25	0.31	–0.49
[11] Including money demand residuals				
Γ3	0.07	0.18	–0.08	–0.14
t-statistic	1.30	1.28	–1.47	–8.45
[12] Not controlling for Δ*y*				
Γ3	0.00	0.06	–0.14	–0.14
t-statistic	0.10	0.87	–2.38	–9.20
Memorandum item				
Baseline model				
Γ3	0.01	0.10*	–0.18	–0.04
t-statistic	0.26	1.33	–1.62	–3.28

Note: Asterisk denotes significance at the 10 percent level (for a one-sided test for Γ3 > 0).
[a] Coefficient on real interest rate, Γ3. Contagion variable included (except in [1]), but not reported.

output. In the methodology adopted here, this would not be captured because the benchmark exchange rate controls for (actual and expected) changes in output. The issue is addressed readily enough, however, by simply dropping Δy from the definition of Δx in equation (4). As Table 4, specification 12, suggests, this makes little difference to the results.

V. Conclusions

One of the most controversial elements of the East Asian crisis was the stance of monetary policy. With continued depreciation of the exchange rate, some commentators suggested that higher interest rates, far from defending the currency, were having a perverse effect. In this view, higher interest rates, by creating the expectation of widespread bankruptcies, were widening the risk premium and resulting in downward pressure on the exchange rate.

In this paper, we argue that nominal interest rates are not a good gauge of the monetary stance—particularly in a crisis environment, where the nominal interest more likely reflects fears of inflation and of currency depreciation—and propose a method of testing whether higher real interest rates indeed contributed to a weakening of the currency.

We use a simple monetary model to filter out the effects of the monetary expansion on the exchange rate and to identify the risk premium. Turning to the determinants of this risk premium, we find little evidence that higher real interest rates contributed to a widening premium and hence, ceteris paribus, to a weakening of the exchange rate. Only for Korea is the coefficient even occasionally positive, and even there, it is generally statistically and economically insignificant once contagion effects are accounted for. We conclude that the perverse effect of higher interest rates on the exchange rate remains largely a theoretical curiosus.

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