E C O N O M I C S B U L L E T I N

Are International Interest Rate Differentials Driven by the Risk Premium? The Case of Asian Countries

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Abstract

This paper investigates the relationship between international interest rate differentials and the risk premium during the 1997-1998 Asian crisis. Variables standing for the accumulation of imbalances in the monetary sector are used as proxies for the risk premium. We show,using a Vector Error Correction Model (VECM) on monthly data from January 1994 to December 2002, that the international interest rate differentials are driven by the risk premium indicators. This result explains the temporary inability of high interest rates to support exchange rates. However, the risk premium considered in this paper would have been required regardless of the interest rate policy. Consequently, high interest rates helped to prevent exchange rates from depreciating more.

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1 Introduction

The Uncovered Interest Parity (UIP) condition defines foreign exchange market efficiency. It states that two assets which are strictly identical except for currency of denomination should have the same rate of return through the joint assumptions of rational expectations, risk neutrality, free capital mobility and the absence of taxes on capital transfers. As a result, the international interest rate differential (called the forward premium) should, on average, be equal to the expected exchange rate change if all riskless arbitrage opportunities have been exploited. This implies that *ex-post* changes in exchange rates should be positively related to international interest rate differentials with a unit coefficient.

However, many researchers have highlighted a forward premium puzzle which points out that the forward premium mispredicts the direction of the subsequent change in the spot rate (Taylor, 1995). Many estimations for a large variety of currencies and time periods show that *ex-post* changes in exchange rates are generally negatively related to international interest rate differentials. These results could support evidence of expectational errors, or of a time-varying risk premium (Froot and Thaler, 1990; Lewis, 1995; Taylor, 1995).

Over the last decade, the accumulation of new data and the emergence of new markets have provided empirical support for interest rate parity. Flood and Rose (2001) and Chinn (2006) find that UIP works better, on average, in emerging countries and particularly during the 1990s[1]. Francis et al. (2002) focus on the time-varying risk premium explanation of deviations from UIP in emerging countries. They show that a significant part of the excess currency return is due to a time-varying risk premium and that, except for during the 1997-1998 financial crisis, financial liberalization leads to a decline in this excess return in Asia.

These empirical tests of the UIP condition are important because this condition is widely used for monetary policy recommendations during currency crises. Interest rate defenses of fixed exchange rates during such crises suggest that raising the domestic interest rate appreciates the spot exchange rate, but only if the expected exchange rate and the risk premium are kept constant. These two assumptions are challenged during a currency crisis, and then theory becomes ambiguous about the relationship between interest rates and exchange rates. The traditional view claims that a tight monetary policy strengthens a currency since it raises the return obtained from investing in the country, reduces capital flight, and discourages speculation (Dekle et al., 2002). The revisionist view (Furman and Stiglitz, 1998; Radelet and Sachs, 1998), in contrast, argues that raising interest rates has a negative impact on exchange rates. The revisionist view emphasizes that the risk premium could be positively correlated with interest rates during a currency crisis because, high interest rates worsen the financial position of debtors, therefore raising default probabilities.

This ambiguous relationship between exchange rates and interest rates is also found

[1] Frankel and Poonawala (2002) find a smaller bias for emerging market currencies than for advanced country currencies. They conclude that a time-varying risk premium is not an appropriate explanation for the forward premium puzzle in advanced country currencies (which is consistent with Frankel (1988) and Engel (1996)), since emerging markets are riskier and UIP tests perform better on these markets. Nevertheless, their results do not mean that excess currency returns in emerging countries cannot be explained by a time-varying risk premium.

in empirical research. With a large set of cross-country data covering the period of 1980-1998, Goldfajn and Gupta (1999) show that tight monetary policy eases the reversal of currency undervaluation through nominal appreciation, but that this result is not robust when a country faces a twin financial crisis. Furthermore, Furman and Stiglitz (1998) and Kraay (2003) find no evidence in favor of the traditional view. In a sample of low-inflation emerging countries including East Asian countries, Furman and Stiglitz (1998) show that high interest rates lead to exchange rate depreciations. In a sample of currency crisis episodes from 1960 to 1997, Kraay (2003) finds no evidence that raising interest rates lowers the probability of success of a speculative attack. Time series analyses over the 1997-1998 Asian crisis also find mixed results. Baig and Goldfajn (2002) and Dekle et al. (2001, 2002) support the traditional view, even if a tight monetary policy leads to a limited exchange rate appreciation. In contrast, Ohno et al. (1999) show that correlation and causality relationships between exchange rates and interest rates change significantly during a crisis, and then high interest rates fail to support exchange rates. The role of the risk premium is particularly well documented by Gould and Kamin (2000). They use international credit spreads and domestic stock prices as proxies for the risk premium and they find that exchange rates were significantly affected by the risk premium indicators but not by interest rates during the 1997-1998 Asian crisis.

The empirical studies mentioned above focus on the behavior of the exchange rate. The focus of this paper is rather on the behavior of the interest rate differential as an explanation of the ambiguous results concerning the ability of high interest rates to strengthen exchange rates. The behavior of the interest rate differential in advanced countries has been notably investigated by Caramazza (1993), Bernhardsen (2000) and Lane et al. (2001). These empirical studies show that the interest rate differential depends on macroeconomic variables such as the inflation differential, the current account or the real income growth differential. The relationship between the interest rate differential and these variables underlines the fact that macroeconomic policy influences depreciation expectations and the risk premium. Moreover, Borensztein et al. (2001) show that the domestic interest rates in Hong Kong, Singapore, Argentina and Mexico react to risk premium shocks. Policymakers can therefore resist depreciation following risk premium shocks in emerging countries. In this paper, a Vector Error Correction Model (VECM) is used to estimate to what extent international interest rate differentials are driven by the risk premium in four Asian countries (Thailand, South Korea, the Philippines and Malaysia), using monthly data from January 1994 to December 2002[2]. Variables standing for the accumulation of imbalances in the monetary sector are used as proxies for the risk premium. The relationship between the interest rate differential and the risk premium is particularly relevant, since it could explain the temporary inability of high interest rates to support exchange rates during the 1997-1998 Asian crisis and thereby explain the ambiguous relationship between exchange rates and interest rates found in numerous empirical studies. We therefore suggest that the risk premium would have called for stronger exchange rate depreciations if Asian countries had not implemented tight monetary policies.

Section 2 presents the model and the variables used as proxies for the risk premium. Section 3 discusses the empirical results. Section 4 concludes.

[2]Gould and Kamin (2000) propose an alternative approach. They estimate to what extent the risk premium indicators affect exchange rates.

2 International interest rate differential determination

The UIP condition suggests that the international interest rate differential is a function of both the expected exchange rate change and the risk premium. The IUP condition is given by

$$i_t - i_t^* = \frac{\mathrm{E}_t\{S_{t+1}\} - S_t}{S_t} + \rho_t, \tag{1}$$

where i_t is the domestic interest rate, i_t^* the foreign interest rate, ρ_t the risk premium, S_t the nominal exchange rate and $\mathrm{E}_t\{\cdot\}$ the expectation operator. The risk premium ρ_t incorporates both the exchange rate risk premium and the default risk premium on domestic bonds.

We assume that efficient market purchasing power parity (EMPPP) holds. Under the efficient market assumption, *ex-ante* deviations from purchasing power parity are unpredictable. The EMPPP condition is given by

$$\frac{\mathrm{E}_t\{S_{t+1}\} - S_t}{S_t} = \mathrm{E}_t\{\pi_{t+1}\} - \mathrm{E}_t\{\pi_{t+1}^*\}, \tag{2}$$

where π_{t+1} is the domestic inflation rate over the period t to $t+1$ and π_{t+1}^* is the foreign inflation rate over the period t to $t+1$.

Using equations (1) and (2), one obtains

$$i_t - i_t^* = \mathrm{E}_t\{\pi_{t+1}\} - \mathrm{E}_t\{\pi_{t+1}^*\} + \rho_t. \tag{3}$$

Equation (3) shows that the real interest rate differential is simply the risk premium if the EMPPP condition holds.

Equation (3) is estimated to analyze to what extent international interest rate differentials are driven by the risk premium. Variables standing for the accumulation of imbalances in the monetary sector are used as proxies for the risk premium. Third-generation currency crisis models (Aghion et al., 2001) and early warning systems on twin crises (Kaminsky and Reinhart, 1999; Kaminsky, 2003) identify indicators embodying the risk premium. More precisely, currency crisis models emphasize balance sheet effects, and early warning systems focus on macroeconomic indicators that capture different varieties of crises.

Four indicators are considered to represent the risk premium. First, the domestic credit growth rate (variable $Credit_t$) allows one to characterize boom and bust episodes and exhibit domestic economic fragilities. A positive relationship is expected with the international interest rate differential. Second, the ratio of gross foreign liabilities of the banking sector to gross foreign assets (variable $Liab/Assets_t$) should exhibit a positive relationship with the international interest rate differential, since balance sheet effects are based on foreign liabilities. This variable could underestimate the exposure to balance sheet effects, since foreign liabilities are also held directly by firms. However, we do not have any monthly statistics to represent this aspect, and statistics on short-term external debts are, at best, quarterly. In addition, banks play a key role during the build-up of monetary imbalances – domestic banks are intermediaries between foreign lenders and domestic producers (Agénor and Aizenman, 1997) – and during financial crises (Mishkin, 1996). Third, the international reserves growth rate (variable GIR_t) should capture the reversal of international capital flows. Variable GIR_t is therefore related to sudden stop phenomena. A negative relationship is expected with the international interest rate differential. Finally, the ratio of M2 to international reserves (variable $M2_t/Reserves_t$), initially studied by Calvo and Mendoza (1996), allows one to capture monetary aggregate imbalances, since domestic credit expansions have to be consistent with the fixed exchange rate system. A positive relationship is therefore expected with the international interest rate differential.

The following empirical model is estimated

$$(i_t - i_t^*) = \beta_0 \left(\pi_{t+1} - \pi_{t+1}^* \right) + \sum_{i=1}^{4} \beta_i \, \rho_{it} + cst + \varepsilon_t, \tag{4}$$

where ρ_{1t} represents the variable $Credit_t$, ρ_{2t} the variable $Liab/Assets_t$, ρ_{3t} the variable GIR_t, ρ_{4t} the variable $M2_t/Reserves_t$, cst the intercept and ε_t the error term. We assume perfect anticipations (expected variables equal their *ex-post* value) to estimate equation (4) and the reference country is the United States.

3 Empirical results

We consider four Asian countries – South Korea, Malaysia, the Philippines and Thailand – from January 1994 to December 2002. We use monthly data from the International Financial Statistics database and from the Asia Regional Information Center database (see the data appendix). Figure 1 displays the three-month nominal interest rate differentials between Asian countries and the United Stated. Series on inflation, domestic credit and international reserves consist of monthly year-on-year growth rates.

We examine the stationarity of the individual series, using standard unit root tests (Augmented Dickey-Fuller (ADF), Elliott-Rothenberg-Stock (DF-GLS), and Kwiatkowski-Phillips-Schmidt-Shin (KPSS))[3]. The international interest rate differential, the inflation differential and variables $Credit_t$, $Liab/Assets_t$, and GIR_t appear to be non-stationary in level but stationary in first-difference for each country. The variable $M2_t/reserves_t$ is non-stationary in level but stationary in first-difference for South Korea, Malaysia and the Philippines. These variables are therefore considered as $I(1)$ variables. The variable $M2_t/reserves_t$ for Thailand are stationary in level. This variable is therefore considered as $I(0)$.

Equation (4) is estimated with a Vector Error Correction Model (VECM). The model has the form

$$\Delta Y_t = \mu_0 + \sum_{i=1}^{k} \phi_i \left(\Delta Y_{t-i} \right) + \alpha \left(\beta Y_{t-1} + \mu_1 \right) + u_t, \tag{5}$$

where $Y_t' = \left[\; (i_t - i_t^*) \quad \left(\pi_{t+1} - \pi_{t+1}^*\right) \quad Credit_t \quad Liab/Assets_t \quad GIR_t \quad M2_t/Reserves_t \; \right]$ is a (5×1) vector[4], μ_0 is a (5×1) constant vector and the error vector u_t is such that $E(u_t) = 0$, $E(u_t u_s) = 0$ if $t \neq s$ and $E(u_t u_s) = \Omega$ if $t = s$ with $\det(\Omega) \neq 0$.

The Johansen and Juselius (1990) cointegration method is used to estimate equation (5). This method allows one to test for the number of cointegrating vectors using a trace test. However, this test could lead to an over-rejection of the no cointegration

[3] More details of these tests are available upon request.

[4] Except for Thailand where $M2/Reserves$ is stationary in level.

hypothesis, due to the finite sample bias and the possible cointegration rank inconstancy. Consequently, forward recursive trace tests are implemented to investigate the cointegrating rank stability. Moreover, the trace test statistic is corrected for the finite sample bias as suggested by Reinsel and Ahn (1992) and Reimers (1991)[5].

The trace tests support the existence of one cointegrating vector at the 1% level over the whole sample for Thailand, South Korea, the Philippines (Table 1)[6]. The forward recursive trace tests with the finite sample bias correction also support the existence of one cointegrating vector for South Korea and the Philippines at the 10% level (Figures 3 and 4). We conclude that the forward recursive trace test with the finite sample bias correction also supports the existence of one cointegrating vector for Thailand even if the trace statistic is slightly below the 10% level during 7 months in 2000 and 2001 (Figures 2). The forward recursive trace test is not implemented for Malaysia. This country adopted an unorthodox policy response to the crisis, including strong capital controls and a pegged exchange rate (Nambiar, 2003; Tamirisa, 2004). As a result, we consider a smaller sub-sample before the implementation of this policy package. The trace test with the finite sample bias correction indicates one cointegrating vector at the 5% level over the sub-sample.

Estimated cointegrating vectors are reported in Table 1. The lag length is chosen for obtaining normality and independence of residuals. The LR(#1) statistic tests the homogeneity hypothesis between the international interest rate differential and the inflation differential. At the 10% level, this restriction is supported for each country. The LR(#2) statistic tests the weak exogeneity (nullity of adjusting coefficients) of the risk premium indicators and the inflation differential. At the 1% level, these restrictions (which include the homogeneity restriction) are supported in each country. The adjustment coefficient is therefore only significant on the international interest rate differential. Cointegrating vectors in Table 1 represent the estimated interest rate parity and more particularly the long term equation for the determination of the international interest rate differential. The risk premium indicators play a key role in this relation. The joint nullity of coefficients associated with these indicators is tested with a LR test, and the result is that this hypothesis is never supported. However, the four risk premium indicators are not significant in each country. The variable GIR_t is dropped for South Korea and Malaysia. Furthermore, the variables $Credit_t$ and $M2_t/Reserves_t$ are dropped for the Philippines. These results reflect the fact that the accumulation of imbalances in the monetary sector was different from one country to another.

The adjustment coefficient on the error correction term is negative and significant at the 1% level for each country, which is consistent with the error correction behavior. When the error correction term is positive, due, for example, to a decline in the inflation differential, the international interest rate differential decreases to reach its equilibrium level. The speed of adjustment of the international interest rate differential to its equilibrium value is equal to one minus the first-order autoregressive coefficient of the error-correction term (Phylaktis and Kassimatis, 1994). The speed of adjustment is high in Thailand, South Korea and Malaysia, respectively at 51.63%, 29.79% and 44.54% per month. One

[5] This correction does not consist in estimating new critical values but in multiplying the trace test statistic by the scale factor $(T - pk)/T$, where T is the number of observations, p the number of endogenous variables, and k the number of lags .

[6] The relation between the international interest rate differential and the risk premium indicators varies from one country to another. Insignificant indicators are therefore dropped from the cointegrating vector.

can deduce the number of months needed to reduce to 90% of its original amount a given deviation in the cointegrating relation[7] (Phylaktis and Kassimatis, 1994). As a result, 90% of the gap between the international interest rate differential and its equilibrium level is eliminated in 3.2 months in Thailand, 6.5 months in South Korea and 3.9 months in Malaysia. The level feedback is weaker in the Philippines, where the speed of adjustment is 18.71% per month. A given deviation in the cointegrating relation is reduced to 90% of its original amount in 11/1 months. Adjustment forces in each country are consistent with fixed and highly managed floating exchange rate systems applied in these countries (Hermandez and Montiel, 2002).

Figure 5 displays the estimated risk premiums. They increase during 1996 and 1997 in Thailand and South Korea and during 1997 in the Philippines. Accumulation of imbalances in the monetary sector in Thailand and South Korea, through credit expansion and the build-up of foreign liabilities, explains the risk premium rise. Domestic credit and monetary expansions peak in 1996 in the Philippines, which explains why the variables $Credit_t$ and $M2_t/Reserves_t$ are not significant (they are dropped in Table 1 for this country). The estimated risk premium for the Philippines is related more to the fall in international reserves (variable GIR_t) following speculative attacks against the peso. This effect is also significant in Thailand, but not in South Korea. International reserves increase sharply in 1998 in South Korea following both a current account surplus and inflows of IMF funds, but the Bank of Korea gradually decreases the overnight rate and gives more importance to exchange rate appreciation. Consequently, the variable GIR_t does not capture the effect of risk premium changes on the interest rate differentials in South Korea. Nevertheless, Figure 5 shows that the risk premium is a driving force of the international interest rate differential during the crisis in each country. From the middle of 1996 to the end of 1997, the estimated risk premium increases by more than 5 points in Thailand and by more than 4 points in South Korea and the Philippines. These rises reduce the ability of tightening monetary policies to stabilize exchange rates. The estimated risk premium in Malaysia displays a particular path, due to the weak international interest rate differential and the unorthodox crisis management plan implemented by this country.

4 Conclusion

This paper shows that international interest rate differentials are driven by the risk premium during the 1997-1998 Asian crisis. The risk premium variations therefore explain why the interest rate hikes implemented by Asian countries failed to stabilize their exchange rates. However, the risk premium considered in this paper depends on the accumulation of monetary imbalances. This risk premium would have been required regardless of the interest rate policy. As a result, high interest rates helped to prevent exchange rates from depreciating more.

[7] If s is the speed of adjustment and if f and g are respectively the initial and final percentage deviation from equilibrium, the number of intervals from f to g is given by $r = (\ln(g) - \ln(f))/\ln(1-s)$.

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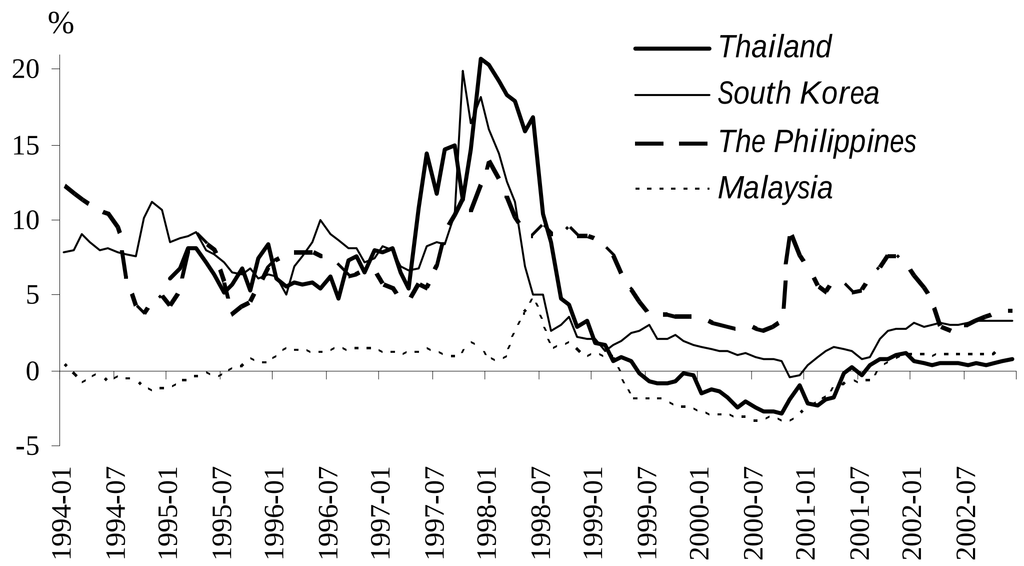
Appendix: Data source descriptions

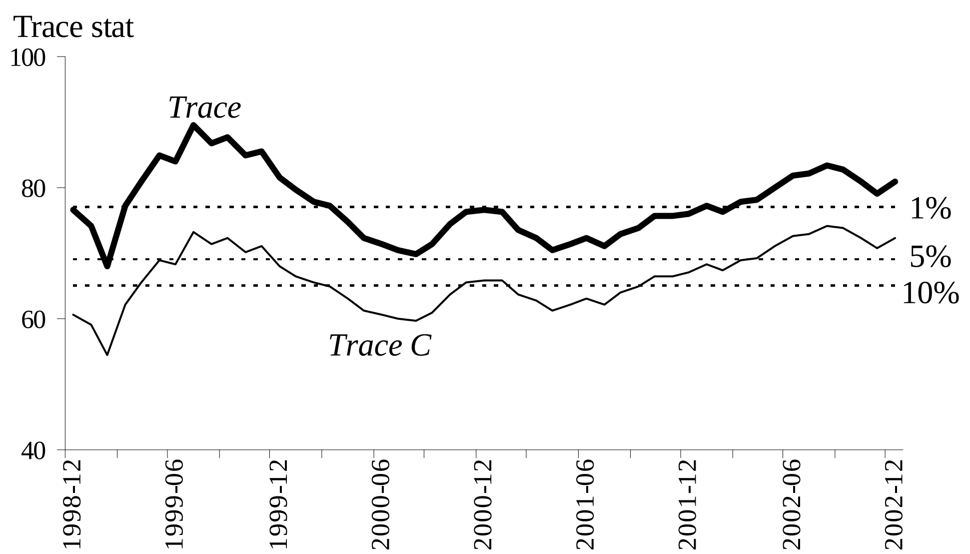
Sources: International Financial Statistics (IFS), International Monetary Fund (IMF), Asia Regional Information Centre (ARIC).
Three month interest rate: Three-Month Interbank Lending Rate for Thailand and South Korea (ARIC indicators / Macroeconomy / Monetary and Fiscal Sectors). Treasury Bill Rate for Philippines, Malaysia and the United States (IFS line 60.C). The IFS lines 60.C are not available for Thailand and South Korea whereas the lines Three-Month Interbank Lending Rate are not available on the whole sample for Philippines and Malaysia in the ARIC database.
Consumer prices: IFS line 64
Real effective exchange rate: ARIC indicators / Macroeconomy / External Sector.
Domestic credit: ARIC indicators / Financial and Corporate Sectors / Financial Soundness Indicators.
Ratio Liab/Assets: ARIC indicators / Financial and Corporate Sectors / Financial Soundness Indicators.
International Reserves: IFS line IL.d.
Ratio M2/Reserves: IFS lines 34 plus 35 converted into dollars (using IFS line ae) divided by IFS line IL.d.

Figure 1: International interest rate differentials



Note: The sample for Thailand starts in January 1995, the 3 month interest rate is not available in 1994 in the ARIC database.

Figure 2: Forward recursive trace test: Thailand



Note: *Trace* represents the trace statistic for 1 cointegrating relation. *Trace C* represents the corrected trace statistic for 1 cointegrating relation.

Figure 3: Forward recursive trace test: South Korea

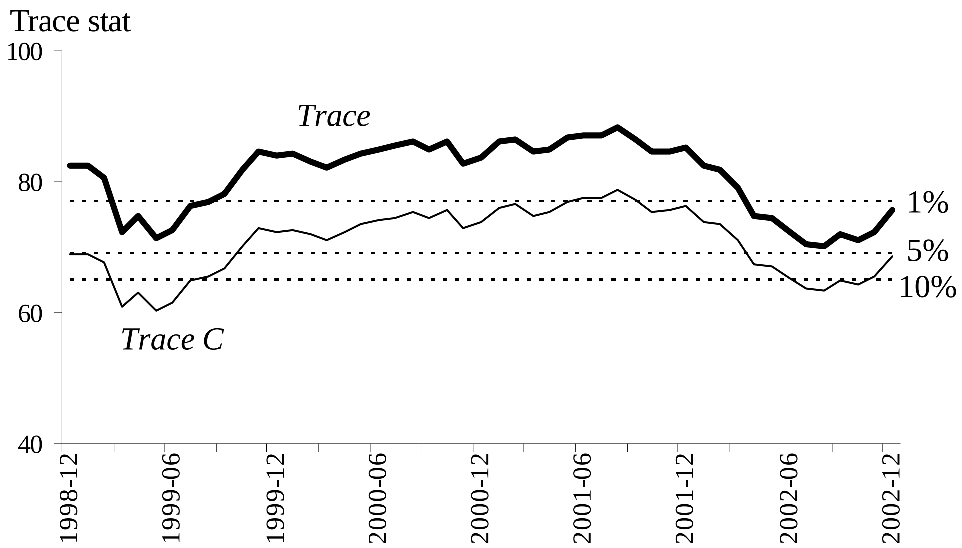
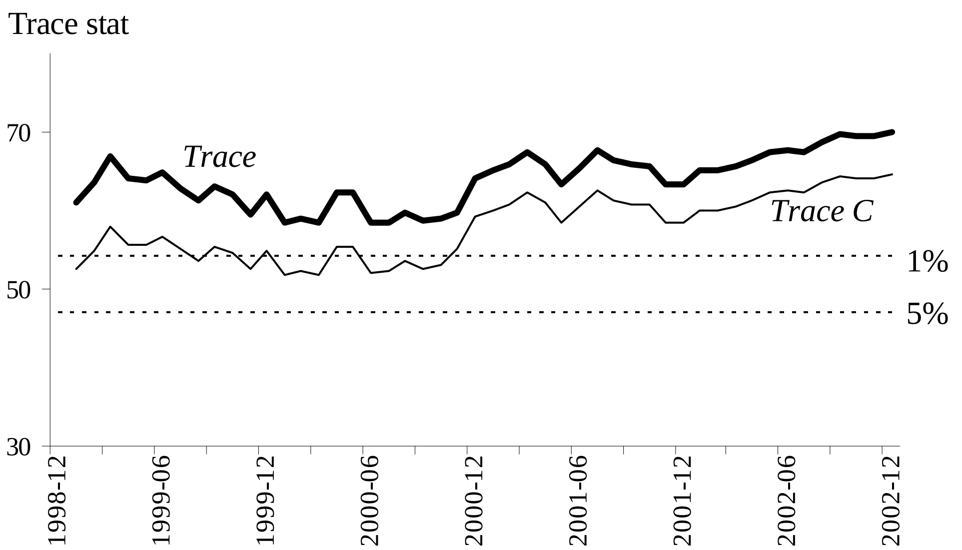


Figure 4: Forward recursive trace test: the Philippines



Note: *Trace* represents the trace statistic for 1 cointegrating relation. *Trace C* represents the corrected trace statistic for 1 cointegrating relation.

Figure 5: Estimated risk premiums

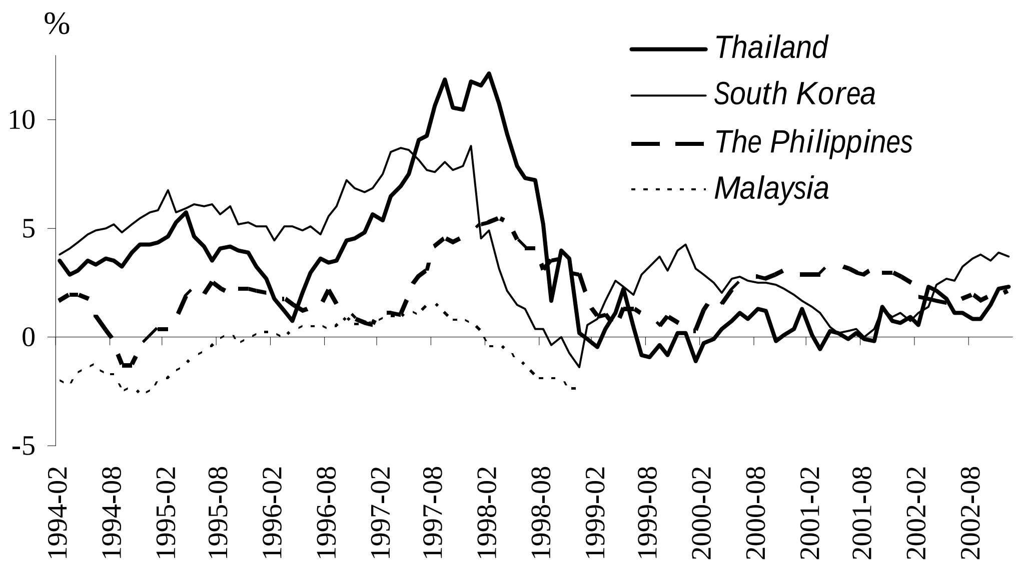


Table 1: Estimated cointegrating vectors

$$(i_t - i_t^*) = \beta_0 \left(\pi_{t+1} - \pi_{t+1}^*\right) + \beta_1 Credit_t + \beta_2 Liab/Assets_t + \beta_3 GIR_t + \beta_4 M2_t/Reserves_t + \mu_1$$

	Thailand	**South Korea**	**The Philippines**	**Malaysia**
$(\pi_t - \pi_t^*)$	1	1	1	1
$Credit_t$	0.1498	0.2210	-	0.0899
$Liab/Assets_t$	0.3678	4.8538	1.8103	0.6034
GIR_t	-0.1747	-	-0.0543	-
$M2_t/Reserves_t$	-	1.0410	-	1.0681
μ_1	2.5351	-8.8378	0.5187	-6.4791
Error-correction term (α) $[t-stat]$	-0.3475 $[-4.46]$	-0.3145 $[-5.40]$	-0.1924 $[-4.77]$	-0.6028 $[-4.72]$
Speed of adjustment	0.5163	0.2979	0.1871	0.4454
90% adjustment (months)	3.1703	6.5103	11.1156	3.9059
LR (#1) $(p-stat)$	Chi-square(1) 0.46 (0.49)	Chi-square(1) 2.12 (0.14)	Chi-square(1) 0.64 (0.42)	Chi-square(1) 1.27 (0.25)
LR (#2) $(p-stat)$	Chi-square(5) 6.53 (0.25)	Chi-square(5) 11.98 (0.04)	Chi-square(4) 7.48 (0.11)	Chi-square(5) 4.41 (0.49)
Trace-stat $(p-stat)$	80.84 (0.00)	75.78 (0.01)	70.05 (0.00)	77.01 (0.01)
Jarque-Bera	3.08	0.37	1.74	1.21
Lag number	2	2	2	2
Dummy	07/1998	10/1997 05/1998	11/2000	-
Obs	96	108	108	60

Note: The speed of adjustment of the international interest rate differential to its equilibrium value is equal to one minus the first-order autoregressive coefficient of the error-correction term. The 90% adjustment is equal to ln(0.10)/ln(1-s) where s is the speed of adjustment (Phylaktis and Kassimatis, 1994).