The Degree of Capital Mobility in Thailand: Some Estimates Using a Cointegration Approach

by

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Abstract

This paper provides an empirical estimate of the degree of international capital mobility in Thailand. A model of interest rate determination which allows for imperfectly mobile capital and the impact of both domestic and international influences on the domestic rate is developed and estimated using cointegration techniques. The results indicate that domestic interest rates have been influenced significantly by foreign interest rates. However, despite several reductions in capital control in the late 1980s and early 1990s, the degree of capital mobility appears not to be as high as previously reported nor to have increased significantly. The finding of a moderately high degree of capital mobility coupled with a reasonable EC-type, short-run dynamic adjustment equation for the domestic interest rate implies that there is still some scope for the Thai authorities to conduct an independent monetary policy.

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THE DEGREE OF CAPITAL MOBILITY IN THAILAND: SOME ESTIMATES USING A COINTEGRATION APPROACH

I INTRODUCTION

The impact of international capital mobility on the conduct of monetary policy in developing countries has received increased attention from policy makers who have realised, perhaps belatedly, that large capital inflows may cause severe economic problems.[1] A dramatic increase in capital inflow was claimed to be the main cause of a rapid rise in the growth of monetary aggregates in Thailand between 1987 and 1992, a major cause of concern for the Thai monetary authorities.[2] In general, the more open is the capital account, the more difficult it will be to conduct independent domestic monetary policy.

Edwards and Khan (1985) were probably the first to attempt to estimate empirically the extent of international capital mobility in developing countries.[3] Using a similar approach, Robinson *et al* (1991) estimated the degree of capital mobility for Thailand for the period 1980-1990 and found it to be so high that it implied little scope for the Bank of Thailand (BOT) to conduct an independent monetary policy. However, as the authors note, there remain some deficiencies in their estimation.[4] The present study avoids some of these problems. In particular, we test for nonstationarity of the data employed and subsequently use appropriate estimation techniques. Furthermore, our sample period encompasses the late 1980s and early 1990s when substantial changes were made to Thai capital controls.

This paper employs a modified version of Edwards and Khan (1985) model to provide empirical estimates of the degree of international capital mobility in Thailand. The

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[1] See, for instance, Corbo and Hernandez (1993).

[2] Discussion of this may be found, for example, in Nijathaworn (1993).

[3] Their estimates indicated that capital mobility and the 'degree of openness of the capital account' were high for both Singapore and Columbia. Haque and Montiel (1990) modified Edwards and Khan's model and estimated the degree of capital mobility for some 15 developing countries and found evidence of relatively high capital mobility in these countries.

[4] Robinson *et al* (1991) do not provide formal tests for stationarity of the data series used, although they acknowledge the importance of this in model estimation. There is no explicit reference to the sort of econometric techniques employed in their study. Furthermore, estimated coefficients on variables describing monetary disequilibrium in the model, such as real income, lagged money balances and inflation, are wrongly signed and not statistically significant.

model allows for imperfectly mobile capital, allowing the relative importance of domestic and foreign influences on domestic interest rates to be estimated. This is done using cointegration and error-correction models.

II THE MODEL: LINKING DOMESTIC AND FOREIGN INTEREST RATES

Following Edwards and Khan (1985), the domestic interest rate is hypothesised to depend on a weighted average of domestic and foreign factors. The estimated value of this weight is viewed as a measure of the degree of financial openness of the capital account. Thus, the domestic interest rate is expressed as:

$$i_t = \varphi(i_t^f + \dot{e}_t) + (1-\varphi)(rr_t + \pi_t^e); \qquad 0 < \varphi < 1 \qquad (1)$$

where i_t is the domestic nominal rate of interest, i^f is the foreign nominal interest rate, $\dot{e}_t$ is the expected rate of depreciation of the domestic currency, rr_t is the real domestic rate of interest, π_t^e represents the expected domestic inflation rate and φ is a measure of the degree of openness of the capital account. In the case of a fully open economy where there are no impediments to capital mobility, $\varphi = 1$ and equation (1) reduces to the *uncovered interest rate parity* (UIP) relationship, reflecting an overwhelming influence of foreign factors on the domestic rate. For a fully closed economy where capital flows are prohibited, $\varphi = 0$ and equation (1) collapses to the so-called *Fisher* equation, where external factors have no direct role to play in the determination of the domestic interest rate.

Following Edwards and Khan (1985), we assert that the real interest rate (rr) is influenced, in the short run, by the excess supply of real money balances (ESM). Thus:

$$rr_t = \rho - \lambda ESM_t + \omega_t \qquad (2)$$

where ρ may be thought of as the constant, long-run equilibrium real interest rate and ω is a random error term. ESM is defined as the difference between the log of actual real money (M1) balances (m) and the log of desired real money balances (m^d). The equilibrium demand for money (m^d) in real terms is hypothesised to take the following form:

$$m_t^d = \alpha_0 + \alpha_1 y_t - \alpha_2 i_t - \alpha_3 fi_t + u_t \qquad (3)$$

where y is the log of real income. The specification of the demand for money in (3) differs from that in Robinson *et al* (1991) by explicitly taking into account the impact of rapid

financial innovation and liberalization in Thailand on the demand for money.[5] Such an impact is modelled here using a proxy variable fi, measured as the log of the ratio M2 to M1.[6] The justification for the use of this proxy is that the increased scope for substitution into M2 provided by financial innovation lowers the demand for M1.

To close the model, the stock of real money balances (m) is assumed to adjust according to

$$\Delta m_t = \beta[m_t^d - m_{t-1}]; \qquad 0 < \beta < 1 \qquad (4)$$

Equation (4) characterises a process by which the nominal interest rate moves to its equilibrium level. Using equations (1), (2), (3) and (4), the reduced-form equation for the domestic nominal rate of interest is:

$$i_t = \gamma_0 + \gamma_1 i_t^* + \gamma_2 y_t + \gamma_3 fi_t + \gamma_4 \pi_t^e + \gamma_5 m_{t-1} + z_t \qquad (5)$$

where i^* is the 'foreign rate of return' defined as $i^f + \dot{e}$. It is expected that $\gamma_1, \gamma_2, \gamma_4 > 0$ and $\gamma_3, \gamma_5 < 0$. Note that $\gamma_1 = \varphi / [1 + \lambda(1-\beta)(1-\varphi)\alpha_2] \approx \varphi$ since α_2 is estimated to be small. In the empirical estimation of (5), the unobservable variables π^e and $\dot{e}$ are proxied by the actual rate of inflation and the actual rate of change in the nominal spot exchange rate of the Thai baht against the U.S. dollar respectively.[7]

The use of cointegration and error-correction methodologies to estimate (5) allows for a distinction between the long-run and short-run relationships among the variables involved. Testing for the presence of a long-run relationship among the variables involved in equation (5) is tantamount to testing whether they are cointegrated. This is done using the maximum likelihood method of Johansen (1988) and Johansen and Juselius (1990).[8]

III DATA AND EMPIRICAL RESULTS

The data employed in this study are for the sample period 1980Q.1 to 1992Q.4.[9] Before proceeding to cointegration tests, all data are tested for the presence of unit roots by using the Dickey and Fuller (1979 and 1981) methodology. The Dickey and Fuller (DF)

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[5] See, *inter alia*, Hataiseree (1991).

[6] The use of this proxy variable is similar to that used in Arrau *et al.* (1991).

[7] Forward rates could be used as a proxy for the expected exchange rate but we do not do so because the nature and size of the Thai forward market are far from fully developed, and the forward rate does not well represent the underlying economic forces in the market.

[8] It has been found to perform generally better than single equation methods and alternative multivariate methods (Gonzalo, 1994).

[9] More detailed descriptions of the data and sources are provided in Appendix A.

and Augmented Dickey and Fuller (ADF) tests for stationarity are reported in Table 1. Since all of the data concerned, except that of the inflation rate[10], are found to be non-stationary, the application of the maximum likelihood cointegration technique of Johansen (1988) to the estimation of equation (5) is appropriate.

Table 1 Dickey and Fuller Tests for Unit Roots

Variables	Levels τ_τ	Levels τ_μ	First-Differences τ_μ
i	-2.25 [0]	-1.77 [0]	-6.49 [0]
	(4.72)	(3.98)	(5.49)
i^*	-2.55 [0]	-1.59 [0]	-7.00 [0]
	(5.54)	(5.85)	(3.90)
y	-2.39 [0]	1.42 [0]	-9.00 [0]
	(6.88)	(7.07)	(6.62)
fi	-1.19 [0]	-1.90 [0]	-5.64 [0]
	(3.71)	(2.30)	(4.20)
π	-3.50 [0]	-3.79 [0]	-5.21 [0]
	(8.91)	(5.87)	(5.25)
$m1$	-2.36 [0]	1.04 [0]	-7.61 [0]
	(6.89)	(6.20)	(7.26)

Notes: Figures in square brackets represent the number of lagged dependent variables used in the autoregression to ensure the residual terms are white noise. The choice between zero and non-zero lags was based on the Lagrange multiplier (LM) test for fourth-order serial correlation of the residuals. Figures in parentheses refer to the values of the LM(4) statistic. The rest of the entries are the reported DF and ADF statistics (τ). The respective critical values for τ_τ and τ_μ at the 5% significance level are -3.50 and -2.93 for N = 50. The tests were conducted using MICROFIT 3.0.

[10] The results for this variable were inconclusive.

Table 2 Cointegration Tests for Variables in Equation (5)

A. Cointegrating LR test Based on **Maximal Eigenvalue** of the Stochastic Matrix

Null	Alternative	Statistic	95% Cr.Value	90% Cr.Value
r <= 3	r = 4	23.5849	22.0020	19.7660
r <= 4	r = 5	17.1109	15.6720	13.7520

B. Cointegrating LR test Based on **Trace** of the Stochastic Matrix

Null	Alternative	Statistic	95% Cr.Value	90% Cr.Value
r <= 4	r = 5	26.1004	19.9640	17.8520
r <= 5	r = 6	8.9895	9.2430	7.5250

C. Estimated cointegrating vectors, coefficients normalized on i in parentheses.

Vector	i	i^*	y	m	π	fi
1	0.21	-0.18	-5.92	5.09	-0.01	0.18
	(-1.00)	(0.85)	(28.21)	(-24.26)	(0.04)	(-0.87)

LR test of restriction that coefficient on $i^* = 1$ $\chi^2(4) = 16.6$ [0.002]

Note: r denotes the number of cointegrating vectors.

The results of the tests for cointegration among the variables appearing in equation (5),[11] presented in Table 2, indicate that there are possibly five cointegrating vectors[12] between the domestic interest rate and foreign interest rate. As indicated by the trace test, the null hypothesis of at least five cointegrating vectors cannot be rejected at the 95% critical value. This result is reinforced by the maximum eigenvalue test. However, only one of these five estimated cointegrating vectors - the first - contains a sensible and meaningful economic interpretation consistent with the model employed here, and this is shown in panel C of Table 2. In this cointegrating vector, all of the estimated coefficients

[11] Four lags were selected for use in the VAR models. The choice for this is dictated by the nature of the quarterly data and the number of observations available.

[12] The existence of a cointegrating vector for the domestic interest rate equation suggests that there is a stable and predictable relationships between domestic and foreign interest rates over the period 1980.Q1 to 1992.Q4.

have the expected signs and are of reasonable magnitudes. The long-run influence of the foreign interest rate on the domestic interest rate is positive and has a plausible magnitude of 0.85, implying that an increase in the foreign interest rate of 1% induced, on average, an increase in the domestic interest rate of 0.85%. The data rejects strongly the restriction that this coefficient is unity. The coefficients of other variables accounting for the influences of domestic monetary conditions are also consistent with the model. The coefficient on income has a positive sign as expected. The coefficient on the real money stock has a negative sign which implies that higher liquidity lowers interest rates. The inflation rate is found to have a positive influence on the nominal domestic interest rate, although its influence seems to be negligible.

It is important to note that the estimated coefficient on the foreign interest rate (a measure of the degree of openness of the Thai capital account) is smaller than the 1.01 obtained by Robinson *et al* (1991). The relatively lower value estimated here suggests that there was some scope, albeit small, for the conduct of independent monetary policy in Thailand during the period under review. The difference between the two estimates may be due to any number of the following factors: First, the money demand function specified in the study of Robinson *et al* is assumed to have a conventional form, while that used here is formulated to take account additionally of the impact of financial innovation. Second, the influence of the rate of depreciation of the expected exchange rate (e) was assumed to be zero in the previous study, while here it is proxied by the actual change in the nominal spot exchange rate. Third, the estimation method employed in the study by Robinson *et al*, while not stated explicitly, appears to be based on the Engle and Granger (1987) two-step approach, while the estimation method used here is based on the Johansen and Juselius technique.

The finding of one sensible cointegrating vector among the variables in equation (5) allows us to formulate the short-run adjustment process in terms of an error-correction (EC) model (Granger, 1986). An EC model describing adjustment towards the long-run domestic interest rate can be formulated as:

$$\Delta i_t = \sum_{i=1}^{4} \beta_i \Delta i_{t-i} + \sum_{i=0}^{4} [\phi_i \Delta i^*_{t-i} + \gamma_i \Delta y_{t-i} + \mu_i \Delta fi_{t-i} + \delta_i \Delta \pi^e_{t-i} + \vartheta_i \Delta m_{t-i}] + \lambda (i - i^d)_{t-1} + \alpha + \varepsilon_t \qquad (6)$$

where the term $(i - i^d)$ is the residual from the first cointegrating vector, representing

deviations from long-run equilibrium. The parameter λ should be *negative* indicating that, if i (the interest rate in the short run) is above i^d (the desired long-run interest rate), then i tends to decrease.

Table 3 Empirical Estimates of the EC Model for the Domestic Interest Rate

$$\Delta i_t = \underset{(2.58)}{0.38}\Delta i_{t-1} + \underset{(2.48)}{0.33}\Delta i_{t-3} + \underset{(2.47)}{0.69}\Delta i^*_t - \underset{(-2.33)}{19.59}\Delta y_{t-2} - \underset{(-3.07)}{28.02}\Delta y_{t-3} - \underset{(-3.94)}{17.61}\Delta m_t + \underset{(4.04)}{44.36}\Delta m_{t-3} + \underset{(14.39)}{59.08}\Delta fi_{t-3} - \underset{(-4.27)}{0.71}\, EC_{t-1}$$

$R^2 = 0.48$ $DW = 2.01$ F(8,24) = 3.93

$\chi^2_{sc}(4) = 2.15$ [0.71]; $\chi^2_{ff}(1) = 0.87$ [0.35];

$\chi^2_{nm}(2) = 0.44$ [0.80]; $\chi^2_h(1) = 0.50$ [0.22] $\chi^2_{ch}(9) = 0.50$ [0.22]

Notes: *EC* is the error-correction term from the most significant cointegrating vector reported in Table2. The equation contains a constant and seasonal dummy variables. The figures in round parentheses are *t*-values while those in square brackets are *p-values*. The diagnostic-test statistics are - $\chi^2_{sc}(4)$ for residual serial correlation; $\chi^2_{ff}(1)$ for functional form; $\chi^2_{nm}(2)$ for normality of residuals; $\chi^2_h(1)$ for heteroscedasticity; $\chi^2_{ch}(9)$ Chow test for structural stability [with sample period broken at 1989Q(4)] - as embodied in Microfit 3.0.

Table 3 reports the final parsimonious error-correction equation obtained by successive exclusion of insignificant variables from the original EC model.[13] The EC equation for the domestic interest rate performs satisfactorily with well-determinded coefficients. In particular, the equation yields a highly-significant, correctly-signed coefficient on the foreign interest rate. The EC term was found to be negative and highly significant. The estimated value of 0.74 implies that approximately three-quarters of the previous discrepancy between the actual and long-run domestic interest rate is corrected in

[13] Exclusion in this way is similar in spirit to the familiar general-to-specific approach of Hendry *et al* (1984).

each quarter.[14] Importantly, a Chow test indicates parameter stablity for a sample break at 1989Q(4). Loosening of capital restrictions in 1990 had little impact on international capital mobility in Thailand.

IV CONCLUSIONS AND POLICY IMPLICATIONS

This paper has examined the degree of international capital mobility in Thailand during the period 1980 to 1992. models. The application of cointegration and error-correction techniques to the Thai data has provided a number of interesting results. First, the Johansen and Juselius estimation technique yields one cointegrating relationship between the domestic and foreign interest rate with a sensible economic interpretation and a well-defined coefficient on the foreign interest rate. Second, the degree of openness of the capital account, measured by this coefficient was found to be 0.85, indicating that both the fully-closed and fully-open hypotheses were rejected by the Thai data. The estimated long-run coefficient on the foreign interest rate, albeit large, was smaller than previous findings. Finally, an EC-type of short-run dynamic adjustment equation for the domestic interest rate was found to fit the data well and passed a set of diagnostic tests.

Important implications for the conduct of stabilisation policy in Thailand are: First, it indicates that the domestic interest rate was influenced by both internal and external factors. Second, it suggests that monetary policy still has a role, albeit limited, to play in stabilisation policy in the short run. Policy actions, such as variations in domestic credit and/or domestic rates of interest, may have some success in offsetting the impacts of foreign capital inflows on domestic money market conditions, but they are severely circumscribed by the degree of openness of the capital account.

[14] Additionally, the estimated equation passes a battery of diagnostic tests; there is no evidence of problems with serial correlation, functional form, non-normality, heteroscedasticity and parameter stability. A CUSUM (not shown here) supports parameter stability, in spite of changes to the financial environment in the sample period

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APPENDIX A

DATA DEFINITIONS AND SOURCES

The data employed are seasonally unadjusted for the period 1980Q.1-1992Q.4 and are taken from various issues of: (i) the Bank of Thailand Monthly Statistical Bulletin and /or the Bank of Thailand's in-house data base (hereafter BOT) and (ii) International Financial Statistics (hereafter IFS) published by the International Monetary Fund, (Washington D.C.). They are:

i Three-month interbank rate of Thailand (percentage). Source: IFS

i^f Three-month Eurodollar rate (percentage). Source: IFS

$M1$ Narrow definition of money, definded as currency in circulation plus demand deposits (billions of Baht). Source: IFS

pc Thailand's consumer price index (log). Source: IFS

m Real money balances (log), definded as logM1- pc. Source: IFS

Y Gross Domestic Product (GDP) of Thailand (billions of Baht). Since the data for GDP are not available on a quarterly basis, they are interpolated as described in Bank of Thailand (1991). Source: BOT

y real GDP (log), definded as logY- pc.

π^e Expected inflation rate for Thailand (percentage), measured as $pc_t - pc_{t-4}$

e Spot exchange rate of the Thai baht against the U.S. dollar (log). Source: BOT

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