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Testing Monetary Model of Exchange Rates in Emerging Economies: New Evidence from ASEAN 5 + 2 Countries

MARIAL A.YOL*
Department of Economics, University of Juba, Khartoum, Sudan & Department of Economics, Universiti Putra Malaysia, Serdang Malaysia.

ABSTRACT

The objective of this paper is two-fold; first, to test whether exchange rates are cointegrated with macroeconomic fundamentals as the theory predicts and secondly, to examine whether flexible-price monetary models can be used to predict and forecast future exchange rates successfully. The panel cointegration tests find the series to be cointegrated. The panel group FMOLS estimation results indicate that the estimated coefficients of money supply and real output levels relative to the US, in addition to the interest rate differentials, are statistically significant and carry correct positive, negative and positive signs respectively. The findings show that monetary models could be an important tool for explaining and forecasting the exchange rates of these ASEAN 5+2 countries in the long run.

Keywords: Nominal Exchange Rate, Monetary Model, Panel Unit Root; Panel Cointegration; FMOLS

INTRODUCTION

The breakdown of the Bretton Woods fixed exchange rate system in 1971 unleashed enormous interest in investigating the long-run determinants of exchange rates. The asset market approach, which asserts that financial markets detennine exchange

* Tel.: 603-89455610 (Home); 60-129075610 (H/P); Fax: 603-89486188; E-mail: yol40@yahoo.com

rates in the short run whereas in medium to long run, the goods markets play a crucial role, dominated the literature on exchange rate determination in the late 1970s and early 1980s. However, the purchasing power parity (PPP) hypothesis, which states that arbitrage tends to force prices of all goods to converge to equality internationally once measured in a common currency, is one of the oldest theories of exchange rate determination. The PPP hypothesis emphasizes the important role played by arbitrage in causing economic agents to exploit price differences so as to predict a stable long-run relationship between currency exchange rates and price levels in the countries involved. The PPP theory is based on the law of "one price", which states that, in the presence of a perfect competitive market structure, identical products should sell at the same price in different markets when expressed in one currency. Assuming that the general level of prices is a reasonable proxy for the cost of production in a country, the PPP hypothesis predicts that the ratio of price levels for any two countries should measure their relative competitiveness. Although the PPP is very appealing as a theory, empirical evidence in its support has been sketchy (e.g., see Hakkio, 1984; Abauf and Jorion, 1990; Frankel and Rose, 1996). Other approaches include the single-equation estimation technique (e.g., see Edwards, 1989; Baffes *et a!,* 1997, Feyzioglu, 1997; Mathisen, 2003; Cady, 2003; MacDonald and Ricci, 2003).

One approach that has attracted growing attention is the monetary model. This approach, which was the first[1] to relate the explanation of variations in nominal exchange rates to the balance of payments, postulates the existence of a stable long-run relationship between nominal exchange rates and a set of macroeconomic fundamentals, assuming that price levels are the same in different countries when expressed in the same monetary units and are determined by the equality between the demand for and supply of money. Since its reemergence in the early 1970s, the model has represented an important benchmark for comparing the other approaches to exchange rate determination, and hence, for measuring the degree of exchange rate misalignment. However, empirical evidence pertaining to the monetary model has not been unanimous. For example, Meese and Rogoff (1983) found that even the naYve random walk model could predict the U.S. dollar exchange rate in the

[1] See Copeland (1994) for detailed discussion.

12-month horizon during the late 1970s and early 1980s better than the structural models in general and monetary model in particular. On the other hand, Baillie and Selover (1987), McNown and Wallace (1989), Groen (2000) and others detected little evidence of cointegration between exchange rates and the independent macroeconomic variables in the monetary model although Mark (1995), in a desperate attempt to save the monetary model, found that both relative money supplies and relative real output levels were important variables in predicting U.S. exchange rates at longer horizons over the 1981-1991 period. However, Mark's failure to report evidence of cointegration among the variables attracted severe criticism from a number of researchers. For example, Kilian (1999), Berkowitz and Giorgianni (2001), Faust *et al.* (2001) particularly criticized the methodology, conclusions and the assumptions about how the long-run behavior of the data series influences the evidence on predictability. In contrast, Woo (1985) and MacDonald and Taylor (1994), among others, found that structural models outperformed the random walk models at all five forecasting horizons.

The reported failure to find a long-run relationship between nominal exchange rates and independent fundamentals has been attributed to a number of problems that range from simultaneous equation bias to the employment of insufficient data samples that characterized the post-Bretton Woods floats, resulting in low power of standard tests. As a result, Husted and MacDonald (1998, 1999), Groen (2000), Mark and Sul (2001) and Crespo-Cuaresma *et al.* (2003), among others, have resorted to employing the panel data approach to test the relationship between nominal exchange rates and monetary fundamentals using the post-Bretton Woods floats. The favorable evidence provided by the panel data approach has certainly rekindled growing research interest in the monetary model.

The objective of this paper is two-fold; first, to test whether exchange rates are cointegrated with the set of macroeconomic fundamentals as the theory predicts and secondly, to examine whether the flexible-price monetary model can be successfully used to predict and forecast future exchange rates.

In many ways this work is more motivated by recent developments in time series econometric literature rather than by any new theoretical advancement in lllodelling the determinants of long-run exchange rates. The panel data approach employed in this paper, by generating more observations from the pooled data, exploits the cross-country variations of the data in estimation, thereby yielding sensible long-run relationships and higher test power than standard unit root tests based on pure time-series data (e.g., the multivariate method of Johansen, 1988 and Johansen and Juselius, 1990). Furthermore, empirical investigation of monetary models of exchange rate determination in general and those employing panel data approach in particular, in the developing world, are meagre. The few exceptions include those by Husted and MacDonald (1999) which examined the extent to which a number of currencies central to the Asian currency crisis were misaligned at the end of 1996. The study, which includes seven Asian countries, excluding China, and including Australia and New Zealand, employed the OLS fixed-effects model on time series data over the period 1974-1996. The empirical evidence derived from the model's panel estimates was in favor of panel estimates but not in the individual country estimates. The authors argued that, with the exception of Indonesia and Malaysia where the ringgit and rupiah were experiencing substantial appreciation and undervaluation respectively, exchange rates at the end of 1996 were close to their actual values, concluding, therefore, that the currency falls that beset these countries were due to shifts in long-run mean values rather than in the underlying fundamentals.

This paper employs the cointegration tests in heterogeneous panels (Pedroni 1995, 1999) for testing cointegration among the series while employing the multivariate group mean panel FMOLS tests in heterogeneous panels (Pedroni 1996, 2000) to estimate the panel parameter coefficients on time series which extend over the period 1980-2005. To estimate the country's long-run parameter coefficient estimates, the paper employs the fully modified OLS (Phillips and Hansen, 1990). Panel models make more information available, thereby allowing for more degrees of freedom and more efficiency in addition to allowing control for individual heterogeneity. Additionally, the technique is capable of correcting for possible endogeneity and serial correlation effects in addition to asymptoticallY eliminating sample bias. For these reasons, we believe that the results from the panel data approach fitted on data from longer floating periods (1980-2005) will hopefully make this paper different from many previous studies and be a valuable contribution to extant literature on the performance of the monetary model of exchange rate determination in these emerging countries.

The remainder of the paper is organized as follows: Section 2 presents the formulation of the monetary model of exchange rate determination; Section 3 describes the econometric methodology; Section 4 presents the discussion of the empirical results and finally, section 5 summarizes the findings, conclusions and Policy implications of the paper.

THE MONETARY MODEL

The monetary model of exchange rate determination is a composite framework that can accommodate many variants of the model, depending on the number of restrictions imposed by individual researchers. Groen (2000) formulated the basic Variant in terms of three relationships: the quantity theory relationship, the uncovered interest parity (UIP) relationship and the money velocity relationship. A. reduced form of the monetary model, commonly known as the flexible-price monetary approach (FLMA) estimated by a number of authors (for example Bilson, 1978; Frenkel, 1978, Hodrick, 1978) is written as follows:

$$e_t = \alpha_0 + \beta_1 \left(m_t - m_t^*\right) + \beta_2 \left(y_t - y_t^*\right) + \beta_3 \left(i_t - i_t^*\right) + \varepsilon_t \qquad (1)$$

Where $e,$ is the spot exchange rate (number of domestic currency units per unit of foreign currency), $m''y'$ and $1;$ denote domestic money supply, income and long-term domestic interest rates whereas the corresponding foreign variables are identified by asterisks and £,is the disturbance term. All the variables, except the interest rates, are in log form. The theory predicts that $/3_1 = 1, /3_2 < 1,$ and $/3_3 > 0.$

The rationale behind these hypothesized parameter coefficients is that a given Percentage increase in domestic money supply is expected to cause an exactly equivalent depreciation of domestic currency whereas a given percentage increase in foreign money supply leads to an exactly equivalent percentage appreciation of domestic currency. As the purchasing power parity (PPP) holds continuously, this implies an equivalent depreciation of the currency and vice-versa. Similarly, an increase in domestic income relative to foreign income increases the transaction demand for money. Assuming that domestic money stock and interest rates are held constant, this increased demand for money can be satisfied only by the fall in dolll.estic prices which, in tum, requires an appreciation of the domestic currency to maintain the PPP. Conversely, an increase in foreign income leads to a fall in foreign price levels and consequently a depreciation of domestic currency is necessary to maintain the PPP. Finally, an increase in domestic interest rates causes depreciation of the domestic currency. Since the nominal interest rate is a component of real interest rate and expected inflation, assuming that the real interest rates are identical between the home country and the rest of the world, an increase in domestic nominal interest rates could be attributed to domestic inflation expectations which can, in turn, cause domestic money to decrease and expenditure on goods to increase. This rise in domestic prices entails a depreciation of the currency to maintain the PPP and vice-versa.

The fundamental premise upon which the flexible-price monetary model is based is that all prices are totally flexible and domestic and foreign bonds are perfect substitutes so that the demand for money in relation to its supply plays a pivotal role in exchange rate determination. This simple notion implies that countries with high monetary growth rates run the risk of high inflationary expectations, leading to a fall in demand for real money balances, increase in expenditure on goods, rise in domestic price levels and a depreciation of domestic currency to maintain the PPP.

Equation (1), which posits a strong long-run equilibrium relationship between nominal exchange rates and relative money supplies, relative output levels and interest rate differentials establishes the long-run flexible-price monetary model as developed by Bilson (1978), Frenkel (1978) and Hodrick (1978). It is equation (1) that will be estimated in this study.

THE ECONOMETRIC METHODOLOGY

Prior to conducting the panel cointegration tests (Pedroni, 1995, 1997), we employ panel unit root tests developed by Levin *et al.* (1992, 2002) and Im *et al.* (1997, 2003) to examine the level of integration among the variables. The starting point for these tests is the classification of the unit root tests on the basis of whether there are restrictions on the autoregressive process across cross-sections or series. The general AR(l) process considered for panel data is:

$$Y11 = PiY1t\text{-}1 + \quad ,oi + Ei, \tag{2}$$

where i = 1, 2, . . ., are N cross-section units or series that are observed over periods t = 1, 2, . . ., r. The ' represent the exogenous variables in the model, including any fixed effects or individual trends, *p;are* the autoregressive coefficients, and £;, are assumed to be mutually independent idiosyncratic disturbances. If IP,! < 1, then *y* is considered to be weakly trend-stationary. However, if *IP)* == 1, then y contains a unit root. To conduct a panel unit root test, two important assumptions regarding *P;* are made. It is assumed (for example, Levin *et al.,* 2002; Breitung, 2000 and Hadri, 2000) that the persistence parameters are common across cross-sections so that *P;* = *p* for all i: Alternatively, p;can be allowed to vary freely across cross-sections. (e.g., Im *et a!.,* 2003).

The Levin *et al.* (1992, 2002, hereafter LL) test assumes a common unit root Process so that *P;* is identical across cross-sections. LL specified the basic ADF specification:

$$4Y_{,1} == a_{;,}Y_{;\_1} + \sum_{JI}^{p,} /3/Y;-;+ \quad 8 + £_{1} \quad i = 1,...,N: t = 1,...,T \tag{3}$$

Where *a== p-* 1, but allows the lag order for the difference tenns, $p_{1'}$ to vary across cross-sections, and £;, – *1ia(O,a;)*. The null and alternatives hypotheses tested by LL are written as: $H_0 : a = 0$, against $H_1 : a < 0$. These hypotheses imply that, Under the null hypothesis, there is a unit root otherwise there is no unit root. LL showed that under the null, a modified t-statistic for the resulting *a* is asymptotically normally distributed

$$t = \frac{t_a - (Nf)s.rr^{-2} srn(a)\mu^*_{mJ} ...}{IV\text{-}^* \; amr} \quad N(O,l) \tag{4}$$

Where *t* is the standard t-statistic for *a* = 0,cT^2 is the estimated variance of the '<tort.:„ *STD(d)* ;,th"tandirrd orrornf *d* and *T* *T-(* *p,t N*)-1.

As the LL test was considered[2] to be too restrictive and biased[3], Im *et al.,* (2003) proposed another panel unit root test based on the average of augmented Dickey-Fuller tests computed for each panel unit in the model as:

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[2] Breitung (2000) argued that, in the first place, the bias correction used by LL implied a severe Joss of Power and secondly LL statistic would be distorted if an ADF were used in the LL instead of i\DF.

[3] ln dynamic panel data, Nickell (1981) showed that the presence of heterogeneous intercept (fixed-effects) causes OLS estimator of a common autoregressive coefficient to be biased as T oo for fixed l'. Only when T tends to infinity does the correlation disappear. However, in many practical applications Where the time period is relatively short the LSDV estimators suffer from severe bias.

$$\Delta y_{it} = \alpha_i y_{i,t-1} + \sum_{j=1}^{p_i} \beta_{ij} \Delta y_{i,t-j} + X'_{it}\delta + \varepsilon_{it} \tag{5}$$

where ε_{it} can be serially correlated and heteroscedastic, but cross-sectionally independent. The IPS technique allows heterogeneity in the short-run dynamics, in the error structure and in the form of fixed effects and linear trend coefficients.

The null hypothesis is: $H_0 : \alpha_i = 0$, for all i, while the alternative hypothesi is given as

$$H_1 : \begin{cases} \alpha_i = 0 & for\ i = 1,2,\ldots,N_1; \\ \alpha_i < 0 & for\ i = N_1+1, N_1+2, \ldots, N \end{cases} \tag{6}$$

To compute the standardized test after estimating the separate ADF regressions, IPS computed the average of the t-statistics for α_i from the individual ADF regressions, $t_{iT_i}(p_i)$ as follows:

$$\bar{t}_{NT} = \left(\sum_{i=1}^{N} t_{iT_i}(p_i) \right) / N \tag{7}$$

In the general case where the lag order in Equation (5) may be non-zero for some cross-sections, IPS showed that a properly standardized t-statistic, which has an asymptotic standard normal distribution, could be calculated as:

$$W_{IPS} = \frac{\sqrt{N}\left(\bar{t}_{NT} - N^{-1} \sum_{i=1}^{N} E\left(\bar{t}_{iT}(p_i)\right) \right)}{\sqrt{N^{-1} \sum_{i=1}^{N} Var\left(\bar{t}_{iT}(p_i)\right)}} \to N(0,1) \tag{8}$$

where $E(\bar{t}_{iT}(p_i))$ and $Var(\bar{t}_{iT}(p_i))$, the expected mean and variance of the ADF regression t-statistics, are provided by IPS via Monte Carlo simulation for various values of T and p and differing test equation assumptions. In addition, IPS developed the standardized LR-bar as presented by:

$$\bar{L}_{IPS} = \frac{\sqrt{N}\left(\overline{LR}_{NT} - N^{-1} \sum_{i=1}^{N} E\left(LR_{iT}(p_i)\right) \right)}{\sqrt{N^{-1} \sum_{i=1}^{N} Var\left(LR_{iT}(p_i)\right)}} \to N(0,1) \tag{9}$$

'Y.here $E(ZR,T(P;))$ and $Var(ZR;r(P;))$ are the asymptotic values of the mean and variance, respectively, of average LR statistic also tabulated by IPS. The test uses the average of log-likelihood ratio statistics for testing the null of a unit root in individual ADF equations. Under the .null ofunit root both the t-bar and LR-bar statistics have a standard normal distribution as N ∞ and T ∞ and for N/T N 0. Under the alternative hypothesis of stationarity, they diverge to minus and plus infinity, respectively. Hence, acceptance of both tests of the null implies strong evidence of panel non-stationarity.

As a common practice, if the series are not 1(0), the next step is to test for cointegration among nonstationary variables, using the Johansen cointegration technique. However, the power of Johansen's test in multivariate systems with small sample sizes can be severely limited. This is the problem which the panel cointegration test was developed to address. In panel cointegration regression models, the asymptotic properties of the estimators of the regression coefficients and the associated tests are different from those of time series cointegration models since pooling cross-section and time-series dimensions together increases data and the power of the test, thus, causing the panel test statistics and estimators to converge in distribution to normally distributed random variables.

Pedroni (1997) proposed seven test statistics for the null of no cointegration, four of which are pooled or "within" tests while the other three are group-mean tests or "between" tests. Some of the tests are completely parametric while others involve nonparametric estimation of long-run covariances. The test is based on a Panel ADF regression on the residuals estimated via a statistic LSDV regression:

$$y_{it} = \alpha_i + \beta x_{it} + \varepsilon_{it} \quad \text{(IO)}$$

The first set of four statistics is based on pooling along the within-dimension Which is constructed by summing both the numerator and denominator terms over the N dimension separately, whereas the other set of three statistics is based on Pooling along the between-dimension which is constructed by first dividing the numerator by the denominator prior to summing over the N dimension[4]. The first of these tests is a type of non-parametric variance ratio statistic. The second and third statistics are panel versions of non-parametric statistics that are similar to

'For complete derivation of the seven tests please see Pedroni (1999).

Phillips and Perron rho-statistics and t-statistics, respectively. The fourth panel cointegration statistic is a parametric statistic, which is analogous to the conventional Augmented Dickey-Fuller t-statistic. The other three panel cointegration statistics are based on a group mean approach. The first is similar to the Phillips and Perron rho-statistic while the last two are analogous to the Phillips and Perron t-statistic and the Augmented Dickey-Fuller t-statistic, respectively.

FULLY MODIFIED ORDINARY LEAST SQUARES (FMOLS) ESTIMATION

In time series as well as panel series, when designing an estimator, two issues must be borne in mind: potential non-exogeneity of the regressors and heterogeneity of the covariance of the error processes. In order to obtain asymptotically efficient consistent estimates in time series case, non-exogeneity and serial correlation problems are tackled by employing fully modified OLS (FMOLS). In this study, we employ the FMOLS. The starting point is the Phillips and Moon (1999) OLS as given by the equation:

y,, =a;+ <5,.t+e,+ /3 +e;;
X;, =X;,_1+e;, (11)

where the vector error process " = (e;;.e,1) 1 is stationary n,. with asymptotic covariance matrix n,. The estimator f3s are consistent when the error process ;=[eH,e;J satisfies the assumption of cointegration between y 11 and x,,.

DATA DESCRIPTION AND THE EMPIRICAL RESULTS

The data employed in this study consist of annual observations of the nominal exchange rate (domestic currency units per US dollar), the money supplies, output levels (GDP as the scale variable) and long-term interest rates from ASEAN 5 (Indonesia, Malaysia, the Philippines, Singapore and Thailand) plus two (China and Japan relative to U.S.). With the exception of nominal exchange rates and interest rates, the other series were converted into US dollar using respective exchange rates. We employ seasonally adjusted money for all the countries. Interest rate differentials for each country were derived by subtracting international interest rates from each country's long-term interest rates. The data were obtained from the IMF International Financial Statistics Website and cover the period 1990-2005. The sample sizes were determined by the availability of relevant data.

As a standard practice, we apply the panel testing techniques for panel unit roots. When the test contains only the constant, both tests show that the series are .1{)). Hfilwever, when the test contains constants and deterministic trends, only the LL test indicates that relative money supply is 1(0) whereas interest rate differential is trend non-stationary at any level. Therefore, we conclude that the series are difference-stationary processes and therefore, could be feasibly employed in cointegration/VECM estimation processes[5].

Given that the series are 1(1), we perform the panel cointegration test using the Pedroni cointegration test with the results reported in Table 1. When the test contains only the constant, four statistics are statistically significant (panel pp-statistic, panel adf-statisitic, group pp-statistic and group adf-statistic). However, When the test contains both the constant and deterministic trend only panel v-statistic and both the panel and group adf-statistics are statistically significant. Bence the Pedroni cointegration test rejects the null hypothesis of no cointegration among the panel variables, implying the existence oflong-run relationships among Panel nominal exchange rates, relative money supplies, relative real output levels

Table 1 Panel Cointegration Tests

Panel Cointegration Test Results							
	Within group				Between group		
Statistics	v-stat.	rho-stat	pp-stat	Adf-stat	rho-stat	pp-stat	Adf-stat
Constant	0.483	-0.578	-2.389	-2.327	-0.488	-1.821	-3.037
Constant + Trend	2.120	-0.306	-1.179	-2.426	-1.302	0.601	-1.985

The critical values for the panel cointegration tests are based on Pedroni (1995). It is to be noted that Pedroni (1997) statistics have the critical values of -1.64 (that is, k < 1.64 implies a rejection of the null) while v-statistic has a critical value of 1.64 (that is, k > 1.64 implies a rejection of the null). 8Ymbols '' and 'denote I% and 5% level of significance.

[5] To conserve a space, the unit test results are not reported here but available upon request from the author.

and interest rate differentials in this seven-country group. These results corroborate those reported, for example, by Abauf and Jorion (1990), Lothian and Taylor (1992) and MacDonald (1993), which supported the existence of long-run reversion to PPP.

PANEL GROUP FMOLS ESTIMATION RESULTS

The panel group FMOLS estimation results are reported in Table 2 in which most panel coefficients carry correct expected signs and are statistically significant at least at the 1% level. In the first place, the coefficient of relative money supplies is positive as the theory predicts. This implies that a 10% increase in money supplies in this seven-country group, relative to US money supply, will cause nominal excl:tange rates to depreciate by approximately 5% per annum. Furthermore, the coefficient of relative real output levels is correctly negative as expected, suggesting that a 10% increase in real output in the region relative to the US real output will appreciate nominal exchange rate by about 11.3% per year. Similarly, interest rate differential is positive as conjectured, implying that a 10% increase in interest rate in this region, relative to the world interest rates, causes nominal exchange rates to appreciate by approximately 0.13% per annum. In a nutshell, the panel domestic money supplies and real output levels relative to the US money supply and real output levels as well as domestic interest rates relative to international interest rates are important and significant factors that determine nominal exchange rates in these seven Asian countries as a group in the long run.

The individual country FMOLS model estimation results reported in Table 3 indicate that the coefficients of money supplies relative to the US money supply are statistically significant in all cases at the 1% level. However, with the exception

Table 2 Panel Group FMOLS Estimation Results

$$e_t=\beta_0+\beta_1 Ln(m_t-m_t^*)+\beta_2 Ln(y_t-y_t^*)+\beta_3(i_t-i_t^*)+\mu_t$$

Variable	β_1	β_2	
Coefficient	0.49"	-1.13"	0.013"
	(42.80)	(90.09)	(573.59)

Note: ("') refer to 1% significance level. Figures in parentheses denote t-Statistics.

of Japan, the coefficients carry correct positive signs and are close to the theoretical value of 1 only in Malaysia. Similarly, the coefficients ofreal output levels relative to the US output are statistically significant and carry correct negative signs consistent with the theory, suggesting that an increase in real output in each country relative to that of the US causes nominal exchange rates in that country to appreciate. The coefficients are less than 1 as the model predicts except in China, Indonesia, Malaysia and the Philippines. Finally, with the exception of Singapore, the coefficients of interest rate differentials are statistically significant at least at the 10% level. The coefficients exceed zero in all cases except in the Philippines, as the theory predicts.

To compute the dynamic multivariate forecasts of nominal exchange rates, We employ the error-correction model derived from the Johansen and Juselius (1990) system. The results for both actual and forecasted nominal exchange rates for the seven countries are plotted in Figures 1-7. The plots exhibit the exceptional ability of the fundamentals to track nominal exchange rate movements across the

Table 3 Individual FMOLS Estimation Results

$$e_t = \beta_0 + \beta_1 Ln(m_t - m_t^*) + \beta_2 Ln(y_t - y_t^*) + \beta_3 (i_t - i_t^*) + \mu_t$$

Countries	β_0	β_1	β_2	β_3
China	0.849...	0.781...	-1.64'..	0.025'''
	(0.000)	(0.000)	(0.000)	(0.000)
Japan	6.277'''	-0.247'''	-0.694'''	0.001··
	(0.000)	(0.000)	(0.000)	(0.689)
Indonesia	6.789'''	0.741'''	-1.245'''	0.021'''
	(0.000)	(0.000)	(0.000)	(0.000)
Malaysia	14.040'''	1.035'''	-1.829'''	0.026'''
	(0.000)	(0.000)	(0.000)	(0.000)
Philippines	0.530'	0.200'''	-1.130'''	-0.016'''
	(0.053)	(0.000)	(0.000)	(0.000)
Singapore	-4.370'''	0.289'''	-0.544'''	0.006
	(0.000)	(0.000)	(0.000)	(0.390)
Thailand	3.302'''	0.688'''	-0.788'''	0.021•
	(0.000)	(0.000)	(0.000)	(0.052)

Notes: (''') , ('') and (') refer to I%, 5% and I0% significance levels. Figures in parentheses denote P·vaJues.

study period, revealing remarkably even the episodes of exchange rate evolution across the 1997 financial crisis. With the exception of Malaysia, Singapore and Thailand, the plots show a consistent pattern of convergence between the fitted and the actual values of long-run exchange rates in other five countries. For these three countries, there was substantial evidence in favor of misalignment especially at the earlier periods. For example, while the Malaysian ringgit experienced volatile adjustments in the period prior to 1999, it became overvalued in the post-1999 period. Similarly, while the Singaporean dollar was overvalued in the period prior to 1995, it has been undervalued for most of the 1995-2005 period. A similar volatile adjustment characterized the Thai bhat, which became undervalued in the period 1999-2002 although it exited and entered an overvaluation episode in the period between 2002 to 2004. In short, the findings show that the monetary model could be an important tool for successfully explaining and forecasting the exchange rates of the East Asian countries individually and as a group in the long run.

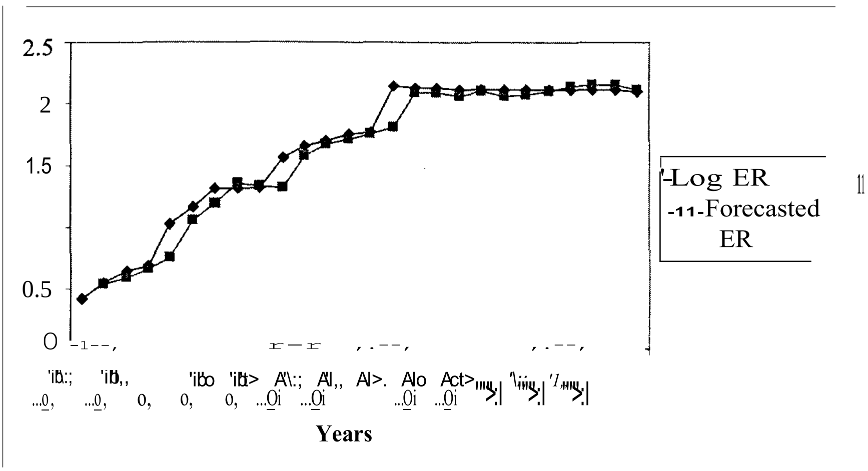

Figure 1 Log of Actual and Forecasted Exchange Rates for China (Yuans/US$)

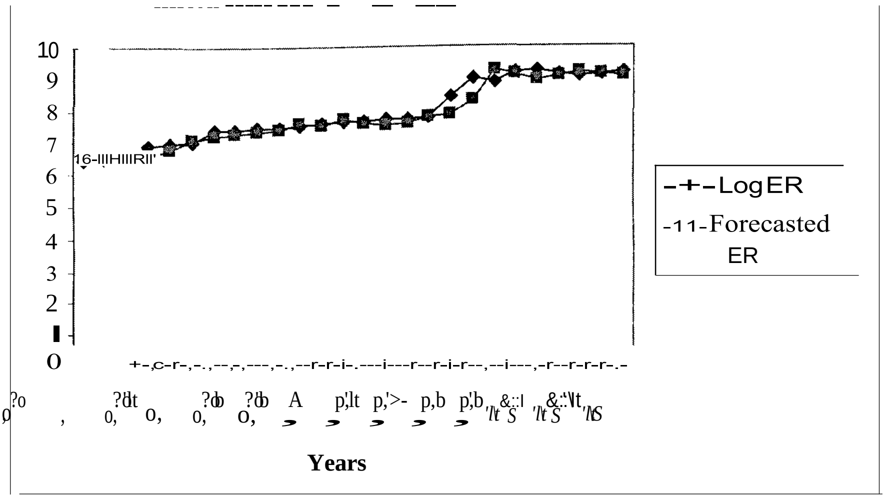

Figure 2 Log of Actual and Forecasted Exchange Rates for Indonesia (Rupiahs/ US$)

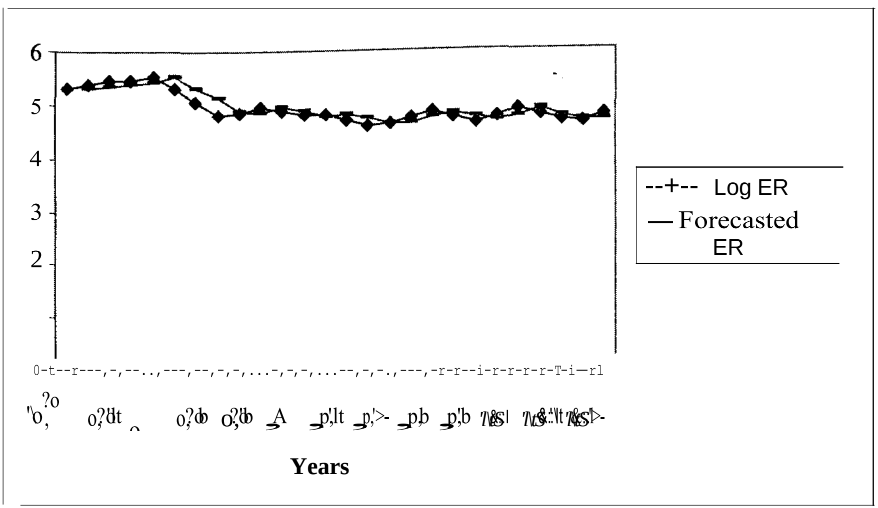

Figure 3 Log of Actual and Forecasted Exchange Rates for Japan (Yens/US$)

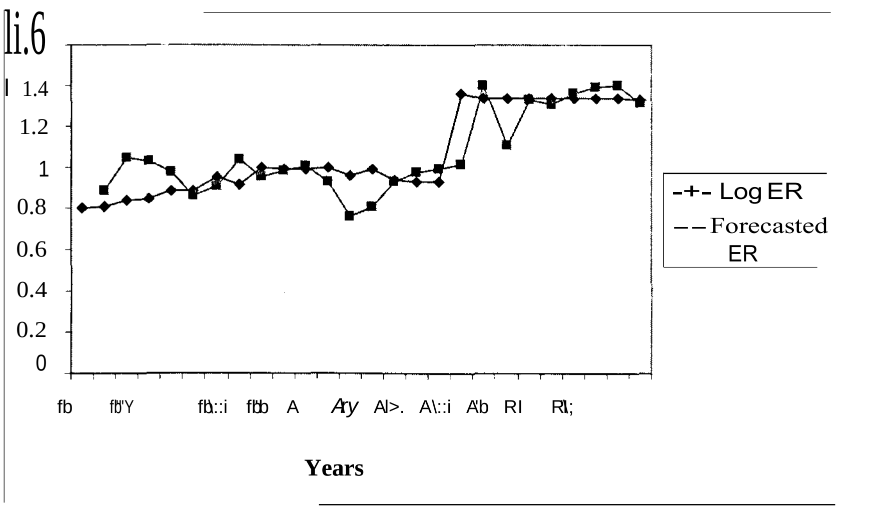

Figure 4 Log of Actual and Forecasted Exchange Rates for Malaysia (Ringgits/ US$)

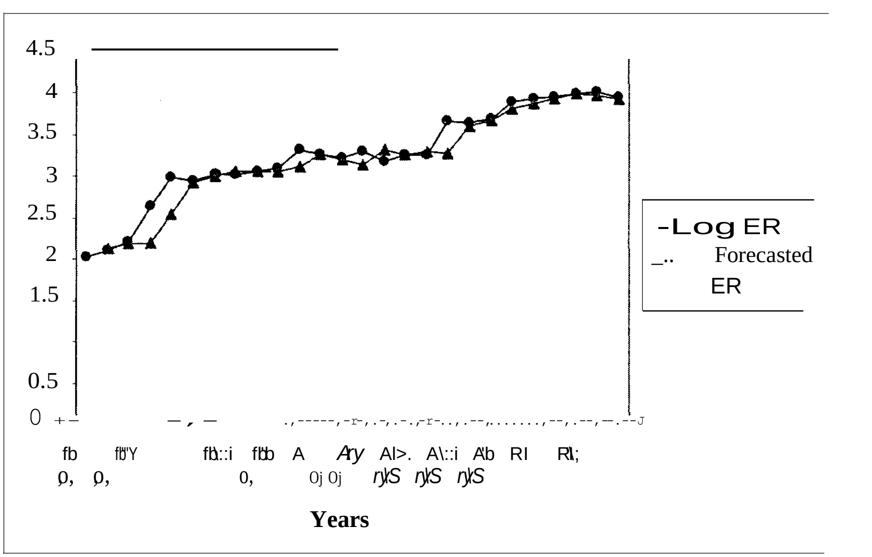

Figure S Log of Actual and Forecasted Exchange Rates for the Philippines (Pesos/ US$)

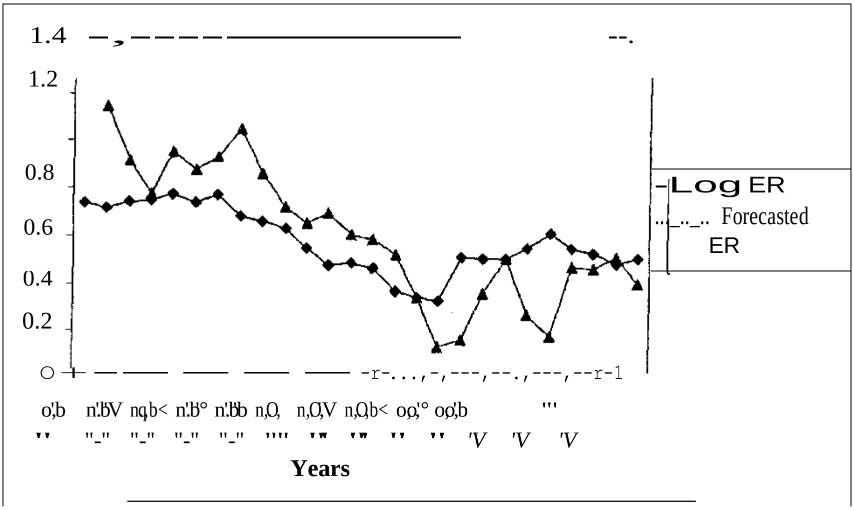

Figure 6 Log of Actual and Forecasted Exchange Rates for Singapore (Singapore $/US$)

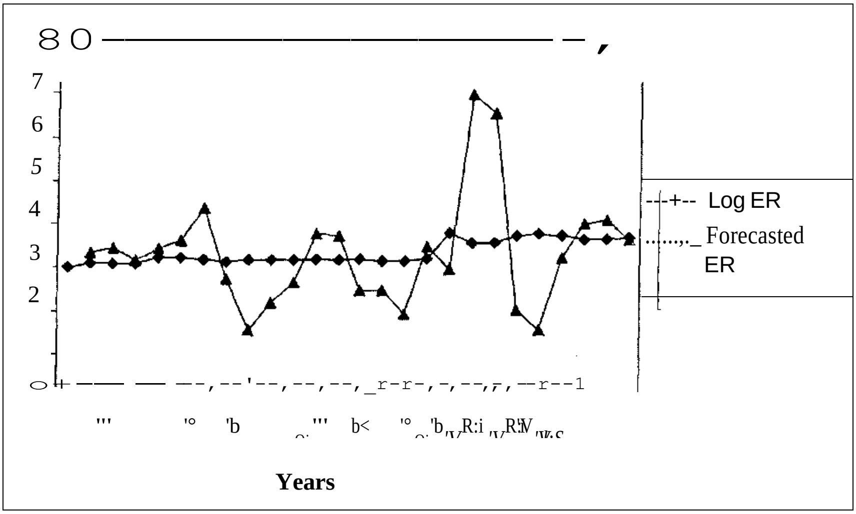

Figure 7 Log of Actual and Forecasted Exchange Rates for Thailand (Bahts/ US$)

CONCLUDING REMARKS

The objective of this paper is two-fold; first, to test whether exchange rates are cointegrated with macroeconomic fundamentals as the theory predicts and secondly, to examine whether the flexible-price monetary model can be used to predict and forecast future exchange rates successfully. The paper applies the panel data approach on annual time-series data from seven Asian countries covering the period 1980-2005. While the panel unit root tests find the series to be difference-stationary processes, the panel cointegration tests reject the null hypothesis of no cointegration among the series.

These results can be summarized as follows: First, the panel group FMOLS estimation results show that the estimated coefficients of panel money supply and real output levels relative to the US as well as the interest rate differentials are statistically significant and carry correct positive, negative and positive signs respectively as the theory predicts. Secondly, the individual country FMOLS model estimation results indicate that the estimated coefficients of money supplies relative to the US money supply are statistically significant in all cases at the 1% level. Moreover, with exception of Japan, the estimated coefficients carry correct positive signs and are close to the theoretical value of 1 only in Malaysia. Thirdly, the estimated coefficients of real output levels relative to the US are statistically significant and carry correct negative signs in all cases, although less than 1,just as the flexible-price monetary model predicts in all countries except China, Indonesia, Malaysia and the Philippines. Finally, except in case of Singapore, the estimated coefficients of interest rate differentials are statistically significant at least at the 10% level. The coefficients exceed zero in all cases except in the case of the Philippines.

With the exception of Malaysia, Singapore and Thailand, the plots show a consistent pattern of convergence between the fitted and the actual values of long-run exchange rates in five countries. For these three countries, there was substantial evidence in favor of misalignment especially in the earlier periods. For example, while the Malaysian ringgit experienced volatile adjustments in the period prior to 1999, it became overvalued in the post-1999 period. Similarly, while the Singaporean dollar was overvalued in the period prior to 1995, it has been undervalued for most of the 1995-2005 period. A similar volatile adjustment characterized the Thai baht, which became undervalued in the period 1999-2002 although it exited and entered an overvaluation episode in the period 2002-2004.

The important finding of this study is that the monetary model could be an important tool for tracking and forecasting the long-run movements of nominal exchange rates in these five East Asian countries plus China and Japan individually and as a group. Perhaps the widespread financial intermediation and foreign exchange market liberalization policies that have been rigorously pursued for the last two decades, resulting in the prevalence of flexible exchange rate regimes in some of these countries, might explain the efficacy of the monetary model in explaining significantly and forecasting successfully the nominal exchange rate movements in these East Asian countries.

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