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Liew, Venus Khim-Sen and Baharumshah, Ahmad Zubaidi and Puah, Chin-Hong

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Venus Khim-Sen Liew[a,*] , Ahmad Zubaidi Baharumshah[b] and Chin-Hong Puah[a]

[a] *Faculty of Economics and Business, Universiti Malaysia Sarawak, 94300 Kota Samarahan, Sarawak, Malaysia.*

[b] *Faculty of Economics and Management, Universiti Putra Malaysia, 43400 UPM Serdang, Selangor, Malaysia.*

Abstract

This paper examines the long-run relationship between exchange rate and its determinants based on the flexible-price monetary model. Multivariate cointegration approach (Johansan 1988, 1989 and Johansen-Juselius 1990) is adopted to attain our objective of study. The empirical results provide evidence favoring the monetary approach to exchange rate for a small and open emerging economy, namely Thailand. In addition, the validity of the underlying assumptions of the monetary approach to the determination of exchange rate is established. The above findings suggest that exchange rate players may effectively monitor and forecast the exchange rate movement via the money supplies, incomes, and interest rates variables of both Thailand and Japan. Besides, one has to follow the economic development of Thailand's major trading partner, Japan, to understanding the movement of exchange rate for Thailand. Moreover, our findings add new insights to accompaniment previous studies that documented the important influence of US in the emerging Asian economies.

Keywords: Exchange rate, monetary model, Thailand, cointegration

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* Corresponding author.

1. Introduction

The validity of the monetary model exchange rate has been of particular interest in the exchange rate study in the past 40 years or so. This is due to the fact that exchange rates have been notoriously volatile ever since the starting of floating exchange rate regime in the beginning of 1973. This volatile nature renders exchange rate movement somewhat difficult to be tracked. Subsequently, being able to understand or even predict the movement of exchange rate is crucial to foreign traders and investors. Policy makers also find it important to monitor and manage the movement of exchange rate so that it will not depart too far away from economic fundamentals. Otherwise, it would be harmful to the trade and disrupt economic progress.

Few models have been put forward in the literature to help understand the exchange rate movement. Among others, the flexible-price monetary model, which postulates that exchange rate may be determined by the money supplies, aggregate incomes and interest rates of domestic and foreign countries, has received much attention in the literature; see among others, MacDonald and Taylor (1991), Choudhry and Lawler (1997), Dutt and Ghosh (2000), Miyakoshi (2000), Abbot and de Vita (2002) and the more recent work Long and Samreth (2008). MacDonald and Taylor (1991) examine the long-run validity of the this model of exchange rate determination by employing the Johansen (1988, 1989) and Johansen and Juselius (1990) cointegration procedure. They provide supportive evidence for two major currencies, namely the USD and British pound. Meanwhile, Choudhry and Lawler (1997) apply the Johansen-Juselius and Engle-Granger approaches to test the validity of monetary model for the case of Canada. The outcome of their tests results reveals the existence of long-run relationship between Canadian Dollar- US Dollar and the variables of monetary models including money stocks, incomes and interest rates over the period of Canadian float spanning from October 1950 to May 1962.

Dutt and Ghosh (2000) applied KPSS and Johansen-Juselius approaches to determine the cointegration relationship between nominal Japanese Yen-US Dollar exchange rates and monetary fundamentals (money supply, interest rates and income), in the fixed exchange rate regime (1959M1 to 1973M1). Empirical evidence in favor of monetary model is obtained in this study. Miyakoshi (2000) examines the flexible-price monetary model in the case of Korea. Based on the Johansen-Juselius cointegration technique, this study finds that the Korean Won exchange rates with US Dollar, German Mark and Japanese Yen as numeraires are all cointegrated with money supplies, incomes and interest rates during the sample period 1980M1 to 1996M12. The author thus concludes that the flexible-price monetary model has long-run validity (in Korea).

The focus of previous studies on the monetary model is on the developed countries, however. Lately, Lee *et al*. (2007), Long and Samreth (2008) and Liew (2009) resuscitate the study of monetary model in the context of emerging economy. These studies are able to provide evidence favoring the long-run validity of this monetary model for the case of the Philippines[1]. Baharumshah *et al*. (2009) also find the predictive power of monetary model outperformed well. Specifically the out-sample forecast of the monetary model outperforms the naïve random walk model at four to eight quarters horizon. The overall findings from these studies have motivated us to apply the model to the Thai baht.

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[1] Lee *et al*. (2007) adopt the multivariate cointegration techniques and the VECM approach in their analysis, while the Autoregressive Distributed Lag (ARDL) approach is employed in the study of Long and Samreth (2008). Liew (2009) contributes to the literature by employing the linear and nonlinear monetary approaches.

This study attempts to contribute to the literature by adding empirical evidence for Thailand, a small and open emerging economy, which has not received much attention before the outbreak of the 1997 – 1998 Asian Currency Crisis [2]. In particular, this study aims to examine the long-run validity of the flexible-price monetary model for the baht-yen exchange rate. The Japanese yen is chosen as the base currency, since it is the traditional major trading partner of Thailand. In the year 2006, Thailand and Japan registered a bilateral trade value of 42416 million of USD rendering Japan as the leading trading partner (followed by US, China, Malaysia and Singapore in that order) of Thailand (ASEAN-Japan Centre, 2007). Various popular monetary restrictions are also tested in this study to provide further insight on the relationship among exchange rate and its determinant. For this purpose, the commonly adopted multivariate cointegration techniques in Vector Error Correction (VEC) framework are utilized to achieve these tasks.

The remainder of this study is organized as follows. Section 2 offers an overview of the Thailand exchange rate regime since 1963. Section 3 describes the model and Section 4 explains the testing procedures. Section 5 contains the data description and discussions on the estimated results. The final section spells out the implications of study and concludes.

2. Overview of Thailand exchange rate regime

[2] Thailand was said to have started off a series of failures in the financial and exchange rate markets in the Asian region through contagion effect in the second half of year 1997. Before that, Thailand has achieved a remarkable average growth rate of 8% per year for two decades, due largely to the opening of the economy to international trade (Agbola and Kunanopparat, 2005).

Thailand adopted a floating exchange rate regime before 1963. However, Bank of Thailand (BOT) abolished this regime on 20 October 1963 and baht (B) was pegged against USD at a rate of B20.80 per USD. In order to hang on to the parity, the gold content of Thai Baht had been reduced by 7.89% in December 1971. To avoid the continuous reduction of gold reserves, the BOT revised the exchange rate policy by allowing Baht to float within a 4.5% fluctuation range in May 1972. However, on the 14 February 1973, the gold content of baht was further reduced by 10% to accommodate for the USD devaluation. The implementation of the limited fluctuation range was taken into effect on 15 July 1973 when the gold content was raised by 4% and the official rate was upgraded to B20.00 per USD. The pegged sustained for few years until 8 March 1978, when an effective rate was established based on a weighted basket of currencies of U.S., West German, Swiss and Japan as major trading partners. This marked the end of baht's link to USD. On 30 October 1978, BOT placed the Effective Rate under controlled (Controlled Floating Rate) and baht was allowed to float within a limited range. Since then, baht devalued several times to improve export competitiveness. On 5 November 1984, the basket of currencies was revised to include currencies of U.S., Japanese, West German, United Kingdom (U.K.), Malaysia, Hong Kong and Singapore. To stop capital flow, a managed float was adopted to control the currency on 3 December 1987. At the same time Baht was depreciated 2% to B26.69 per USD to promote exports. In 1990, 3 more currencies (Brunei ringgit, Indonesian rupiah and Philippine peso) were added to the currencies basket. In the outbreak of the 1997 – 1998 Asian Currency Crisis, baht fell by 22% in one month, from 35.80 baht per USD in June 1997 to 43.57 baht per USD in the following month. Since 2 July 1997, Thailand had given up in defending its currency and adopted the independent float exchange rate regime, of which the value of the Baht is determined by market forces. The baht fell by 60% in just few months, to its historical low of 73.89 baht per USD in January 1998 before it reversed its course to appreciate. The exchange rate was as high as 55.90 baht per USD in April 2006, baht had not resumed its value prior to the crisis, as was the case for the currencies of the rest of ASEAN-5 countries. As of 2006, the *de facto* classification for baht exchange rate according to IMF was a managed floating regime with no pre-determined path for the exchange rate, but basically is guided by the inflation forecasts, which act as the intermediate target (IMF, 2006). As of 2008, Thailand is still pursuing this inflation targeting framework (IMF, 2008).

3. The monetary model

The flexible-price monetary model of exchange rate may be represented by:

$$e_t = (m_t - m_t^*) + \alpha_1 (y_t - y_t^*) + \alpha_2 (i_t - i_t^*) + \mu_t \qquad (1)$$

where e_t is nominal exchange rate expressed as domestic price of foreign currency. In the present case, baht-yen exchange rate is considered. Meanwhile, m_t, y_t and i_t stand for domestic money supply, aggregate national output and nominal interest rate respectively. The corresponding variables for the foreign counterpart are marked with asterisk. μ_t is the white noise.

This model postulates that the movement of exchange rate may be determined by the differentials of money, income and interest rate [$(m_t - m_t^*)$, $\alpha_1 (y_t - y_t^*)$, $\alpha_2 (i_t - i_t^*)$]. According to this model, a rise in the domestic money supply leads to a proportional rise in the price level via the quantity theory of money and to a proportional rise in exchange rate via the purchasing power parity and vice versa. The same is true for foreign money supply. Besides, income differential and interest rate differential are expected to have negative and positive impacts respectively on the exchange rate movement. Moreover, it is assumed that the elasticities for domestic and foreign money are identical. That is, these variables have effect of opposite direction but of same magnitude on exchange rate. The same assumption is apprehended for domestic and foreign interest rate variables; for more details, see, for instance, Frenkel (1976), MacDonald and Taylor (1991) and Dutt and Ghosh (2000).

4. Testing procedures and monetary restrictions

Empirically, the model is tested in the less restricted form, in which the proportionality between money differential and exchange rate, the identical elasticities of incomes and also of interests, are relaxed. The testable version can be written in the following reduced form:

$$e_t = \beta_1 m_t + \beta_2 m_t^* + \beta_3 y_t + \beta_4 y_t^* + \beta_5 i_t + \beta_6 i_t^* + \mu_t \qquad (2)$$

Conventionally, the long-run validity of this model has been tested using the Johansan (1988, 1989) and Johansen-Juselius (1990) cointegration testing procedure in the multivariate framework. This procedure requires the variables in the model to be integrated in the same order, in particular, order one (MacDonald and Taylor, 1994; Rapach and Wohar, 2002). If all variables are integrated of order one, I(1), then there is a possibility that the linear combination of this variables are stationary. This in turn may be taken as evidence of long run (cointegrating) relationship among the variables under examination. As such, the following Vector Autoregression (VAR) of order k, denoted as VAR(k), in the Vector Error Correction (VEC) framework exists and (Engle and Granger, 1987, Johansan 1991) may be represented as:

$$\Delta X_t = \Pi X_{t\text{-}k\text{-}1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t\text{-}i} + \theta D_t + \varepsilon_t \,, \tag{3}$$

where $X = (e, m, m^*, y, y^*, i, i^*)'$ is a 7×1 vector of monetary variables, D contains a set of conditional variables which may include constant, trend and dummies. $\varepsilon_t \sim iid(0, \Omega)$ is a 7×1 vector of independently and identically disctibuted (iid) Gaussian error terms. X_{t-k-1} is a 7×1 vector of the error correction terms (ECTs), that measures the long-run equilibrium relationships. $\Pi = \alpha\beta'$ is a 7×7 matrix that contains information about the of long-run relationships among the variables in the VAR system, where α and β are each of dimension $7 \times r$, where r is the number of cointegrating vectors contains in matrix β . α describes the speed of adjustment of each of the 7 individual variables in the system to deviations from the cointegration relationships. In our case, if $0 \le r \le 6$, then $\beta' X_t$ is stationary even though X_t itself is not. This may be interpreted as the existence of long-run relationship among the variables. The value of r can be determined by the trace (Johansen, 1989) and maximum eigenvalue (λ-max) (Johansen, 1988) statistics. The trace statistic is used to test the null hypothesis of there are at most r cointegrating vectors against the alternative hypothesis of more than r cointegrating vectors, and it is given by:

$$\text{Trace} = -T \sum_{i=r+1}^{k} \ln(1 - \hat{\lambda}_i), \tag{4}$$

where T is the sample size, $\hat{\lambda}_{r+1},\ldots,\hat{\lambda}_k$ are k - r smallest estimated eigenvalues.

The maximum eigenvalue statistic tests the null hypothesis of r cointegrating vectors against the alternative hypothesis of r+1 cointegrating vectors and it is computed as:

$$\lambda\text{-max}=-T(1-\hat{\lambda}_{r+1}) \qquad (5)$$

This study utilized the trace statistics as it is more robust to both skewness and excess kurtosis in the residuals than the λ-max statistic (Cheung and Lai, 1993).

In the literature, some monetary restrictions are commonly performed (MacDonald and Taylor, 1991; Long and Samreth, 2008). The restrictions to be imposed in this study are listed in the first column of Table 3. Briefly, with reference to Equation (2), the null hypothesis of $H_1: \beta_1$=-β_2=1 tests whether there is proportionally between exchange rate and the money differential. The restrictions of opposite coefficients on income and interest rate differential are tested in $H_2: \beta_3 + \beta_4 = 0$ and $H_3: \beta_5 + \beta_6 = 0$ respectively. The other null hypotheses, which include $H_4: H_1 \cap H_2$, $H_5: H_1 \cap H_3$, $H_6: H_2 \cap H_3$ and $H_7: H_1 \cap H_2 \cap H_3$, are to test the joint significance of the some of the combined effects of the first three restrictions. The likelihood ratio (LR) test of restrictions is applied for these hypotheses testing.

5. Data, empirical findings and discussions

Monthly data utilized in this study is obtained from International Financial Statistics published by International Monetary Fund. Income, money supply and interest rate are, in that order, measured by gross domestic product (GDP), M2, consumer price index and money market rate. The baht-yen exchange rate is employed in this study. The sample period runs from January 1977 (1977: M1) through March 2006 (2006: M3). All variables are log-transformed.

In this study, the semi-parametric testing procedure of Phillips and Perron (PP, 1988) are deployed to test the integration order of the variables. The results are reported in Table 1. It is evident from Table 1 that the null hypothesis of non-stationary series can be rejected in none of variables in their levels, even at 10% significance level. However, all variables are found to be stationary after first-differencing, at 1% significant level. As these variables achieved stationary only in the first differences, there are said to be integrated order one, I(1).

Having identified that all variables are I(1) variables, the VAR model is then estimated in the first difference according to Equation (3). In order to estimate Equation (3), the optimal lag length *k* must be determined in advance. Following MacDonald and Taylor (1991), *k* is set at a maximum value of 13 and for lag selection, the general to specific approach is conducted with the help of likelihood ratio (LR) statistic[3]. The VAR residuals are checked for whiteness. If the residuals are found to be non-white noise, a higher lag structure is added until they are whitened.

[3] MacDonald and Taylor (1991) examine the same monetary model for the US dollar based currency of Germany, Japan and UK using monthly data. The authors provide empirical evidence of long-run validity for the monetary model in these countries.

Table 1. Order of Integration Test Results.

Variable	Level		First difference		Decision
	Constant + Trend		Constant		
	Lag [a]	Statistic	Lag	Statistic [b]	
e_t	3	-3.125	3	-14.557*	I(1)
m_t	2	-1.723	1	-14.413*	I(1)
i_t	0	-1.927	4	-14.927*	I(1)
g_t	0	-2.530	5	-17.407*	I(1)
m_t^*	13	-3.133	1	-23.963*	I(1)
i_t^*	5	-1.944	6	-23.251*	I(1)
g_t^*	4	-2.439	3	-14.368*	I(1)
Critical value					
1%		-3.996		-3.457	
5%		-3.429		-2.873	
10%		-3.138		-2.573	

Notes: [a] A maximum of 13 lag order is considered in the PP test and the optimal lag reported is selected based on the Newey-West bandwidth criterion. [b] The null hypothesis of stationary series is rejected in favor of the alternative of non-stationary series if the test statistic is smaller than the critical value. Asterisk indicates rejection of the null hypothesis at 1% significance level.

The LR statistic suggests that VAR (13) model is optimal. However, it is observed that the residuals of VAR (13) model are not white noise. Higher lags are therefore added one by one until the residuals are whitened. This resulted in the final selection of VAR (16) model to be estimated[4]. Thus, the cointegration test is then performed and the trace test statistic is then computed. At this stage, there is a concern on the specification of deterministic component in the cointegration relation, as it can influence the outcome of the test. In this study, three practical model specifications are estimated. They are Model 2 which permits intercept in the cointegration relation; Model 3 which includes deterministic trends in the levels, and Model 4 which allows for trend in the cointegration relation.

[4] The Lagrange Multiplier (LM) test results indicate that there is no serial correlation in the VAR (16) model's residuals up to 24 lag order, implying the fulfillment of the requirement of whiteness (Dutt and Ghosh, 2000). Results are available upon request.

The estimated results are present in Table 2, in which the reported trace statistic has been adjusted for small finite biased by a correction factor of $(T\text{-}nk)/T$, where n=7 in this study, to avoid over rejection of no cointegration (Cheung and Lai, 1993; Osterholm, 2003). The optimum model may be chosen based on the Pantula principle (Johansen 1995). The selection strategy by this principle begins with looking at the most restrictive model (Model 2) for the null hypothesis of r=0 from Table 2 and comparing the test statistic with the corresponding critical values given to the right of the test statistics. If the test statistic exceeds the critical value, the model is rejected, as is the case here, we proceed on to Model 3 under the same null hypothesis of r=0. As Model 3 is also rejected here, we next consider Model 4 in the same row. Since Model 4 is also rejected, we move to the next row with null hypothesis of $r \leq 1$ and start over another round of inspection. It is clear that all Models in this row are rejected. So the process is continued with the next row with the null hypothesis null hypothesis of $r \leq 2$, until the null hypothesis is not rejected for the first time, in this case, in Model 3. Subsequently, Model 3 has been selected by the Pantula principle. Trace statistic suggests that there are at most 2 cointegration relationships among the 7 variables considered[5]. In other words, we have obtained evidence supporting the long-run validity of the monetary model under scrutinized for Thailand.

The in-sample forecasts of the baht-yen exchange rate obtained from the fitted model is plotted in Figure 1, alongside with the actual values. Figure 1 shows that the forecasted value is able to follow the movement of the actual exchange rate closely[6]. In addition, the Pearson correlation coefficient of the two series is 0.993 and it is significant at 1%

[5] Out of curiosity, the Pantula principle is also applied to the maximum eigenvalue statistic. Model 3 is consistently selected but this statistic suggests that there is only 1 cointegrating relationship. Results are not reported but are available upon request.

[6] The forecast error as measured in root mean squared error (RMSE) is 0.035.

level. This suggests that the two series are almost perfectly and positively correlated, thereby cross-validating our earlier conclusion that the forecasted values can closely mimic the actual exchange rate movement. Moreover, the R^2 value obtained from regressing the actual values on a constant and the forecasted values is 0.986, implying that 98.6% of the variation in the exchange rate can be explained by the variation in the monetary variables in the fitted model[7]. All-in-all, these findings indicate that the estimated VAR(16) model fit the data adequately.

Table 2. Model Selection and Trace Statistic

Hypothesis		Model 2		Model 3		Model 4	
Null	Alternative	Trace	1% CV	Trace	1% CV	Trace	1% CV
$r = 0$	$r > 1$	177.307*	143.09	157.672*	133.57	179.737*	158.49
$r \leq 1$	$r > 2$	128.448*	111.01	110.176*	103.18	128.270*	124.75
$r \leq 2$	$r > 3$	92.372*	84.45	75.745	76.07	92.165	96.58
$r \leq 3$	$r > 4$	60.410*	60.16	50.910	54.46	62.041	70.05
$r \leq 4$	$r > 5$	35.752	41.07	30.956	35.65	37.431	48.45
$r \leq 5$	$r > 6$	16.915	24.6	15.322	20.04	21.712	30.45
$r \leq 6$	$r = 7$	6.227	12.97	4.949	6.65	8.846	16.26

Notes: The 1% critical values are included in parentheses. Asterisk indicates rejection of the null hypothesis at 1% significance level.

[7] The intercept is insignificantly different from zero, whereas the slope coefficient is exactly one and it is significant at 1% level.

The finding of long-run validity of the monetary model as specified in Equation (2) allows us to proceed to some tests of monetary restrictions which are commonly imposed in the literature (MacDonald and Taylor, 1991). The results as shown Table 3 reveal that the outcomes depend very much on the number of cointegrating equations (r) we specified in the test. For r=1, all the restrictions can be rejected at 5% significance level or better. On the other hand, for r=2, three hypotheses, namely $H_1 : \beta_1 = -\beta_2 = 1$, $H_4 : H_1 \cap H_2$ and $H_5 : H_1 \cap H_3$, can not be rejected even at 10% significance level. However, the other two hypotheses, that is, $H_6 : H_2 \cap H_3$ and $H_7 : H_1 \cap H_2 \cap H_3$ can be rejected at 5% significance level. Meanwhile, LR test is not available for $H_2 : \beta_3 + \beta_4 = 0$ and $H_3 : \beta_5 + \beta_6 = 0$ due to not binding restrictions in the estimation.

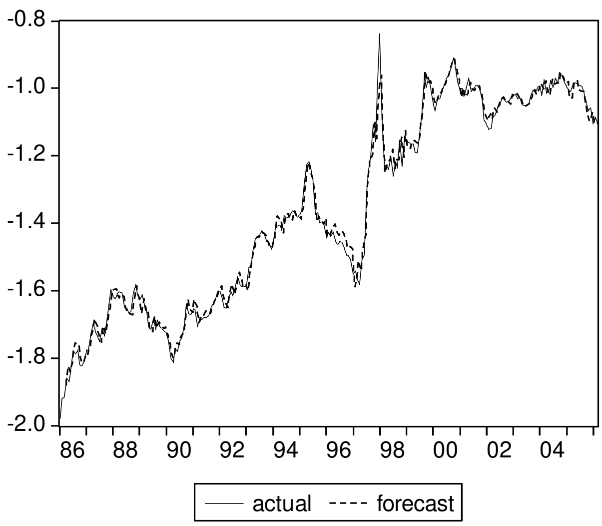

Figure 1: Forecasted and actual values of baht-yen exchange rate

As the number of cointegrating equation has been identified as 2 previously by the trace test, we should resort to the results for the case r =2. Accordingly, from the non-rejection of H_1, we can conclude the baht-yen exchange rate response proportionally to changes in money differential between Thailand and Japan. This will leave the real exchange rate unchanged and thus neutrality of money is binding for the case of Thailand. Besides, failing to reject H_4 implies that the domestic and foreign incomes do have same impact on the movement of exchange rate but they act in different direction, apart from the fact that money is neutral. By the same principle, the non-rejection of H_5 indicates that domestic and foreign interest rates move of the exchange rate in opposite direction by the same amount.

Table 3. Monetary restrictions tests

Hypothesis	r =1		r =2	
	χ^2	**Probability**	χ^2	**Probability**
$H_1 : \beta_1$=- β_2=1	8.205**	0.017	0.081	0.776
$H_2 : \beta_3 + \beta_4 = 0$	9.593***	0.002	n.a.	n.a.
$H_3 : \beta_5 + \beta_6 = 0$	5.046**	0.025	n.a.	n.a.
$H_4 : H_1 \cap H_2$	17.385**	0.001	4.576	0.101
$H_5 : H_1 \cap H_3$	8.678**	0.034	3.198	0.202
$H_6 : H_2 \cap H_3$	9.753**	0.008	4.607**	0.032
$H_7 : H_1 \cap H_2 \cap H_3$	28.162***	0.000	16.293**	0.001

Notes: χ^2 of the LR test has a $r \times m$ degrees of freedom, where r refers to the number of cointegrating vectors and m is the number of restriction in the null hypothesis. n.a. stands for unavailability of LR test due to not binding restrictions in the estimation. ** and *** indicates the rejection of null hypothesis at 5 and 1% significance levels respectively.

6. Conclusion

The monetary model of exchange rate for developed economies has received much attention from researchers. Studies using data from emerging economies are relatively limited. Long and Samreth (2008) resuscitate the study of monetary model in the context of the Philippines, an emerging economy. In the spirit of Long and Samreth (2008) and others, this study examines the validity of the flexible-price monetary model for the case of Thailand using the Johansen multivariate cointegration testing framework. Several restrictions are imposed to the estimated monetary model for further analysis.

The key findings and implications of this study include: First, there exists of two cointegrating vectors in the estimated VAR model, indicating the presence of long-run relationship among exchange rate and the monetary variables for Thailand. Moreover, it is observed that the forecasts of exchange rate generated by the fitted model tend to mimic the actual data closely. Therefore, exchange rate players may monitor and forecast the future exchange rate movement via the money supplies, incomes, and interest rates variables of both Thailand and Japan. Besides, the finding indicates that one has to follow the economic development of Japan, to understanding the movement of exchange rate for Thailand, an emerging economy. This finding adds new insights to accompaniment the majority of previous studies that documented the important influence of US in the emerging Asian economies.

Second, it has been shown in this study that the outcomes of monetary restrictions tests are sensitive to the number of cointegrating equations entered in the estimation. Hence, it is important to appropriately identify the correct number of cointegrating equations.

In this regards, this study adopts the Pantula principle and the trace statistic to conclude that there are two cointegrating vectors in the exchange rate under investigation.

Third, the proportionality of exchange rate and money differential cannot be rejected by the restriction test. This implies that money is found to be neutral in Thailand, in the sense that it influences the nominal exchange rate proportionately but the real exchange rate remains unaltered. Fourth, equal and opposite effects of income differential on the exchange rate is also found in this study, indicating that foreign and domestic economic growths are important determinants that cannot be excluded from the monetary model. Fifth, interest rates variables also have equal and opposite impact on the movement of exchange rate. This finding signifies that influences of domestic and foreign monetary policy on exchange rate of Thailand can hardly be neglected.

This study contributes to the literature by providing empirical evidence supportive of the flexible-price monetary model from Thailand, a small and open emerging economy which has not received much attention in the exchange rate literature before. This study also offers overview of Thailand exchange rate regimes.

Note that, this study fails to reject the null hypotheses of major monetary restrictions and this finding is contrasting to majority of the findings documented in the literature (see, for instance, MacDonald and Taylor, 1991; Long and Samreth, 2008). While this finding implies the validity of various underlying assumptions of the monetary approach to exchange rate determination, further researches may be conducted to explain why these assumptions are maintain in the case of Thailand exchange rate but not in others.

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