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# The Demand for Money in Asia: Some Further Evidence

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**Abstract**

This paper investigates empirically the impact of exchange rate changes on the money demand of seven Asian countries over the quarterly period, 1973 – 2009. Estimates of the cointegrating relations are obtained using different estimators and the error-correction technique was used to obtain the estimates of the short-run dynamics. The major results show that increases in the exchange rate, exert a significant positive effect upon money demand in both the long-run and the short-run in each of the seven countries. Further, domestic interest rates are found to have significant negative effect on the demand for money. These effects may result in significant reallocation of resources by monetary authorities and market participants. Our results provide justifications for the monetary authorization to pay attention to broad money.

**Keywords:** money, demand, model, exchange rate, cointegration

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## 1. Introduction

During the last two decades, exchange-rate changes have played an important role in the behavior of market participants in Asia – witness the discussions about the financial turmoil of 1997 and 1998 which led to currency devaluations and even, issuance of bailout packages by the International Monetary Fund (IMF) for some countries. Also, as the housing bubble in the United States that started in August 2007 resulted into a worst world-wide financial and economic upheaval between 2007 and 2009 (Das, 2012), it forced Asian economies into significant depreciations. Further, after the G-20 Summit in June 2010, these economies started a buildup of foreign exchange reserves. They held more than enough precautionary foreign exchange reserves in order to have greater flexibility and to minimize currency appreciations. A central question has been the effects of such movements in the exchange rate on the demand for money behavior. This paper tries to understand empirically the possible contribution of exchange-rate changes in explaining the short-and long-run money demand behavior of the Asian economies.

In conventional money demand studies for less-developed countries (LDCs), real money balances are usually modeled as a function of real domestic income and expected inflation rate; that is, the interest rate variable is generally excluded. It is often claimed that the inflation rate should be preferred because for LDCs the choice of asset holdings is limited to either real goods or money in the absence of interest bearing financial assets -- see, for instance, Bahmani-OsKooee and Rehman (2005). Others have argued that the interest rates data when available have shown very little dispersion over time due to government control. In this paper, we suggest that failure to include interest rates will lead to omitted variable bias because over the past two decades, Asian economies have instituted continuing policies of financial market liberalization and economic reforms which have led to greater interest rate liberalization as well as varying interest rate data (Arize & Shwiff, 1993). Furthermore, it can be argued that even if the deposit rates in LDCs are set by the monetary authorities, it is still a relevant alternative asset to money that the public can hold (Arize, Malindretos & Grivoyannis, 2005). Thus, the empirical importance of the interest rate variable should be treated as an empirical matter.

Concerning the impact of exchange rate on money demand, we note that with the advent of the floating exchange-rate era in 1973, a number of analysts have identified the exchange rate as a possible omitted variable in the money demand function of developed and developing countries; see, for example, Arango and Nadiri (1979, 1981), Arize (1989), Arize and Ndubizu (1990), Arize and Shwiff (1993) and Bahmani-OsKooee and

Rhee (1994). Econometric theory suggests that omission of such a variable may bias a model towards overstating the influence of the included variables. Also, McNown and Wallace (1992:107), who examined the demand for money in the U.S. have concluded that "nonstationarity in the demand for money can be resolved by inclusion of the exchange rate." Few empirical studies have investigated the impact of exchange rate on the demand for money in LDCs, and the results have been mixed.

Lee and Chung (1995) find that exchange rate exerts a negative and statistically significant effect on Korea's demand for money. Tan (1997) considers the demand for money in Malaysia and reports that exchange rate has a statistically significant effect on real M1 balances and no effect on real M2 balances. Ibrahim (2001) reexamines Malaysia and finds that exchange rate has a negative effect on demand for money. Weliwata and Ekanayake (1998) study Sri Lanka's demand for money and find that exchange rate has a significantly negative effect on real M1 balances. They failed to obtain a meaningful real M2 relation. Asian economies are largely quite open to the rest of the world and exchange-rate movements have had a bearing upon domestic money demand; see Rajan (2010) for more on this. An empirical study of the money demand function can provide monetary authorities with useful information for designing appropriate stabilization policies.

The primary purpose of this paper is to examine the long-run relations and short-run dynamics of the demand for broad money in Asian economies (India, Korea, Malaysia, Pakistan, the Philippines, Sri Lanka and Thailand) over the quarterly period 1973:1 through 2009:4. (Note 1) Our objectives are fourfold.

The first objective is to check the data for unit roots. The second objective is to determine whether there exists a stationary long-run relationship between real money balances, domestic interest rate and exchange rate. The cointegration method used is the Johansen (1995) method, which is a Full Information Maximum Likelihood (FIML) estimator. The choice of this technique follows from the Monte Carlo study by Haug (1996), which shows that the Johansen estimator has the least size distortion. The third objective is to determine, provided the long-run equilibrium exists, the sign, magnitude and statistical significance of the long-run effects of real income, interest rate and exchange rates on money demand. Recent specification suggestions (Note 2) have shown that it is better to employ single equation estimator once r =1 (cointegration) has been confirmed. We report the estimated long-run elasticities obtained after normalizing them on real money balances and our analysis is based on Fully Modified Least Squares (FIML) estimator of Phillips and Hansen (1990). The fourth objective is to examine the short-run dynamics of the money demand model for each country using the general–to-specific paradigm (see Hendry, 1987).

The features of this paper which distinguish it from the other research in this area are: (1) the inclusion of interest rate data rather than the inflation rate; (2) the focus on the movements of exchange rate as an important determinant of demand for money; (3) the use of recent quarterly data instead of data before the financial crises noted above; (4) evidence from more than one country; and (5) testing for nonlinearity in the short-run dynamics of the model.

The rest of the paper is organized as follows. Section 2 describes the money demand model. The empirical results are presented in Section 3, and concluding remarks close the paper in Section 4.

## 2. Model Specification and Theoretical Considerations

A standard simple money demand specification is as follows:

$$m_t = a_0\, y_t^{\alpha_1} i_t^{\alpha_2}; \qquad a_1 > 0 \text{ and } a_2 < 0 \tag{1}$$

where $m_t$ refers to real money balances at time t, $y_t$ is real income and at time t, and $i_t$ is domestic interest rate at time t.

Drawing on the empirical literature in this area (Arize, 1989; Arize & Shwiff, 1993) and the implications for dynamic specification of the possible existence of error-correcting mechanisms in the data-generating process, we estimate a simple, conventional model, augmented with exchange rate as the long-run equilibrium (equation 2) and the short-run relationship (equation 3) for the (desired) real money balances:

$$m_t^* = \alpha_0 + \alpha_1 y_t + \alpha_2 i_t + \alpha_3 e_t + \varepsilon_t \tag{2}$$

$$\Delta m_t - \alpha - \beta L(\Delta X_t) - \gamma \varepsilon_{t-1} = \psi_t \tag{3}$$

where $m_t$* is the logarithm of desired holdings of real money balances (real M2); (Note 3) real M2 consists of currency outside the banks and demand deposits plus quasi-money divided by the consumer price index; (Note 4) $y_t$ is the logarithm of real GDP; $i_t$ is the level of domestic interest rate which is proxied by money market rate or lending rate; $e_t$ is the exchange rate variable; and the stochastic disturbance is $\varepsilon_t$.

According to equation (2), real money balances are assumed to be an increasing function of real income (i.e., real GDP) as the usual budget conditions dictate; that is, $\alpha_1$ is expected to be positive. On the other hand, the opportunity cost of holding money relative to financial assets (i) is expected to yield a negative influence on money demand, so $\alpha_2$ is expected to be negative.

To provide a background on how variations in exchange rates could affect the demand for money in an open economy, we start by noting that foreign sector considerations, such as variations in exchange rates (Arango & Nadiri, 1981:70) play important roles in determining domestic money demand. The currency substitution literature (Agenor & Khan, 1996) suggests that portfolio shifts between domestic and foreign money provide a role for a foreign exchange variable. (Note 5) At least two effects on the demand for domestic currency will result from variations in the foreign exchange rate. On the one hand, there is a wealth effect as currency depreciation leads to increases in the domestic price of foreign securities which then causes portfolio-adjustment effects; see the discussions, for instance, by Arango & Nadiri (1979, 1981), Arize and Shwiff (1993) and Bahmani-Oskooee and Techaratanachai (2001). Assume that if wealth holders evaluate their asset portfolios in terms of their domestic currency, (Note 6) a depreciation of the domestic exchange rate that increases the value of foreign assets held by domestic residents would boost wealth (i.e., a wealth-enhancing effect). This in turn leads to a rebalancing effect because, in order to maintain a fixed share of their wealth invested in domestic assets, they will repatriate part of their foreign assets to domestic assets, including domestic currency (i.e., rebalancing). (Note 7) As Arango and Nadiri (1981:79) have argued, it is because of the "rebalancing effect" brought about by changes in the exchange rate, that an exchange rate depreciation has a positive effect on the demand for money. Hence, exchange rate depreciation would increase the demand for domestic currency.

On the other hand, variations in the exchange rate can generate a currency substitution effect in which a key role is played by investors' expectations. According to the currency substitution literature, as a weak domestic currency develops expectations for further weakening, asset holders will respond by shifting some of their portfolios away from domestic currency into foreign assets. So, if an increase in exchange rate (i.e., depreciation) induces a decline in money holding by domestic residents, the estimate of $\alpha_3$ should be negative. To summarize, based on the above discussions, an increase in the exchange (i.e., depreciation) could have a positive or negative effect on the demand for money; therefore, which effect dominates is an empirical issue.

As is customary, equation (2) has assumed that in the long run, any deviation of actual (observable) real money balances from desired (unobservable) should disappear so that it may be viewed as cointegrating model (Ericisson, 1998). The basic idea of cointegration is that two or more nonstationary time series may be regarded as defining a long-run equilibrium relationship if a linear combination of the variables in the model is stationary (converges to an equilibrium over time). (Note 8) Because the money demand function in equation (2) describes a stationary long-run relationship among the variables, this can be interpreted to mean that the stochastic trend in real money balances is related to the stochastic trends in the explanatory variables. In other words, even if deviations from the equilibrium should occur, they are mean- reverting (Arize, 1997).

## 3. Estimation Results

The empirical analysis uses quarterly frequency data for seven Asian counties, namely India, Korea, Malaysia, Pakistan, the Philippines, Sri Lanka and Thailand. The data are taken from the International Monetary Fund's *International Financial Statistics* (IFS) latest CD-ROM (2009). The data used are the logarithmic value of real broad money, real GDP and exchange rate. In general, the data for the period, 1973:1 through 2009:4 were used for our analyses.

The first step in testing for cointegration in a set of variables is to test for stochastic trends in the autoregressive representation of each individual time series using augmented Dickey-Fuller and Johansen tests. For space consideration, the empirical results are not presented here, but they suggest that all the variables in equation (2) are nonstationary.

For any set of I (1) variables, Johansen (1995) has developed a system-based cointegration procedure to test the absence or presence of long-run equilibria among the variables in equation (2). The test utilizes two likelihood ratio (LR) test statistics, the maximal eigenvalue ($\lambda$-max), and trace (Tr) to test the presence or absence of long-run equilibria between the variables in equation (2). For the $\lambda$-max, the null hypothesis of r is tested against r+1 cointegrating vector, whereas for the trace, the null hypothesis is that there are at most r cointegrating vectors and the alternative hypothesis is of a general form. The presence of a significant cointegrating vector or vectors indicates a stable relationship among the relevant variables. The number of lags used in the Vector Autoregression (VAR) is based on evidences provided by Schwarz's criterion; however, in the case of serial correlation sufficient numbers of lags are introduced to eliminate the serial correlation of the residual. For the four-variable system, we employ the lag order of two for India, Korea and Malaysia, three for the Philippines, and Sri Lanka, one for Pakistan, and five for Thailand, respectively.

Table 1. Results from cointegration tests

| | | **Maximum Eigenvalue** | | | | **Trace statistics** | | | |
|---|---|---|---|---|---|---|---|---|---|
| | **H0:** | **r=0** | **r≤1** | **r≤2** | **r≤3** | **r=0** | **r≤1** | **r≤2** | **r≤3** |
| | **H1:** | **r=1** | **r=2** | **r=3** | **r=4** | **r≥1** | **r≥2** | **r≥3** | **r≥4** |
| India | 2 | 40.30 | 15.59 | 1.42 | 0.238 | 57.54 | 17.25 | 1.66 | 0.23 |
| Korea | 2 | 38.30 | 13.26 | 11.73 | 2.81 | 66.10 | 27.80 | 14.54 | 2.81 |
| Malaysia | 2 | 42.40 | 16.47 | 7.35 | 0.84 | 67.06 | 24.66 | 8.18 | 0.84 |
| Pakistan | 1 | 40.39 | 10.61 | 5.34 | 0.0001 | 56.34 | 15.95 | 5.34 | 0.0001 |
| Philippine | 3 | 62.27 | 19.15 | 7.12 | 1.29 | 89.83 | 27.55 | 8.40 | 1.29 |
| Sri Lanka | 3 | 43.68 | 12.97 | 6.51 | 1.64 | 64.80 | 21.12 | 8.15 | 1.64 |
| Thailand | 5 | 28.38 | 10.64 | 7.40 | 2.42 | 48.83 | 20.46 | 9.82 | 2.42 |
| CV (5%) | | 27.07 | 20.97 | 14.07 | 3.76 | 47.21 | 29.68 | 15.41 | 3.76 |

Description: The critical values are from Table 1 of Osterwald-Lenum (1992). The 5%* shown above represents the critical values from table 2* of Osterwald-Lenum (1992). It is used because South Africa's model allows for a trend and constant in the cointegration space; see Johansen (1995).

Table 1 presents the cointegration test results. The results answer to two questions: whether a long-run equilibrium relationship exists and whether it is unique. Starting with the λ-max test result, the null hypothesis tested is that there is no cointegrating vector ($r = 0$) against the alternative of cointegration of order one ($r =1$). Since all test values in column 1 of Table 1 are significantly larger than the critical value of 27.07 at the five per cent level, we easily reject the null hypothesis (r = 0) of no cointegration. For space reasons, we do not report the data for comparing the null hypothesis of cointegration with $r \leq 1$ to the alternative of cointegration with $r = 2$ because none of the test statistics exceeds the critical value for this test (20.97 at the five per cent level) so we accept the null of cointegration with a single cointegrating vector.

A similar procedure applies to the trace test statistic. We first test the null hypothesis of $r = 0$ against the alternative hypothesis of r ≥ 1 in each country. Since all empirical test statistics exceed the critical value of 47.21 for the trace statistic, we reject the null of $r = 0$ for all countries. Next, we test $r \leq 1$ against the $r \geq 2$ alternative hypothesis. This time, none of test statistic exceeds the critical value of 29.68 and we thus accept the null of cointegration with a single cointegrating vector. Note that, had we excluded the exchange rate variable from our system, we would not have found evidence of cointegration at the 5 percent level. This finding implies that the effect of exchange rate is crucial to the long-run stability of the system.

### *3.1 Long-run Equilibrium Estimates*

Next, we are interested in the long-run impact of the three explanatory variables (real income, domestic interest rate and exchange rate) on money demand. In particular, we want to know whether the signs of the estimated coefficients are in line with their predicted values, whether these estimates are statistically significant, and how the explanatory variables perform in relative terms. Except the interest rate, our semi-log specification of the estimation model (see equation (2)), yield long-run elasticities in the case of real income and exchange rate. In Table 2, we show the empirical results of Equation (2) using the fully modified least squares (FMLS) estimator of Phillips and Hansen (Note 9). To increase confidence in the estimates obtained, we provide empirical results from the dynamic least squares (DLS) estimator of Stock and Watson. From these estimates, we gather that our conclusions are not particularly influenced by the method of estimation.

Table 2. Long-run Elasticities and Hypothesis tests

| Country | Phillips-Hansen Estimator (FMOLS) | | | Stock-Watson Estimator (DOLS) | | |
|---|---|---|---|---|---|---|
| | ***y*** | ***i*** | ***e*** | ***y*** | ***i*** | ***e*** |
| India | 1.17<br>(12.39) | -0.06<br>(5.56) | 0.24<br>(3.88) | 1.66<br>(11.62) | -0.02<br>(2.55) | 0.17<br>(1.81) |
| Korea | 1.36<br>(26.11) | -0.02<br>(2.1) | 0.74<br>(7.2) | 1.36<br>(25.6) | -0.02<br>(2.54) | 0.74<br>(7.11) |
| Malaysia | 1.58<br>(22.45) | -0.02<br>(1.72) | 0.18<br>(1.37) | 1.01<br>-20.46 | -0.01<br>-3.35 | 0.02<br>-1.62 |
| Pakistan | 1.68<br>(11.66) | -0.02<br>(2.62) | 0.61<br>(4.46) | 1.38<br>-9.08 | -0.06<br>-1.99 | 0.27<br>-2.31 |
| Philippine | 0.94<br>(5.41) | -0.03<br>(4.74) | 0.50<br>(7.18) | 0.64<br>(1.86) | -0.03<br>(1.78) | 0.55<br>(4.15) |
| Sri Lanka | 1.57<br>(7.64) | 0.01<br>(3.45) | 0.26<br>(1.97) | 1.86<br>(9.07) | -0.01<br>(2.73) | -0.44<br>(3.38) |
| Thailand | 1.361<br>(59.66) | -0.003<br>(1.43) | 0.899<br>(18.96) | 1.549<br>(29.35) | -0.010<br>(1.42) | 0.269<br>(1.79) |

Description: The absolute t-values are in the parentheses beneath each estimated coefficient. The critical value at 10 percent is 1.3 and 1.67 at 5 percent level.

Focusing on the results obtained from the FMLS estimator, we gather that the demand relation is estimated to be positive for real income (y) and exchange rate (e), but negative for the interest rate (i). These hold for all Asian countries in the sample. Furthermore, all coefficient estimates are statistically significant at the conventional levels. For the domestic interest rate variable, ignored in several LDCs studies, the equilibrium elasticities (calculated by multiplying the impact elasticities by the mean of the variable) are -0.624, -0.217, -0.095, -.172, -0.448, -0.184, and -0.024 for India, Korea, Malaysia, Pakistan, the Philippines, Sri Lanka and Thailand, respectively. These results provide strong support for the theoretical predictions regarding the impact of real income, interest rate and exchange rate on real money balances. As far as relative strength of the various determinants is concerned, we find that the magnitude of the exchange-rate effect is substantially smaller than the real income effect but it is generally (in absolute terms) larger than of interest-rate elasticity. This implies that the economic importance of both income levels and exchange rate effects far outweighs the importance of interest rate variations for money demand.

Table 3A. Long-run parameters using inflation rate (inf).

| Country | Phillips-Hansen Estimator (FMOLS) | | |
|---|---|---|---|
| | ***y*** | ***inf*** | ***e*** |
| India | 1.63<br>(15.69) | -0.02<br>(2.31) | 0.18<br>(2.62) |
| Korea | 1.34<br>(28.96) | 0.02<br>(3.03) | -0.01<br>(-6.00) |
| Malaysia | 1.59<br>(68.16) | -0.02<br>(1.65) | -0.03<br>(0.37) |
| Pakistan | 1.99<br>(7.37) | -0.03<br>(1.76) | 0.81<br>(3.25) |
| Philippine | 1.36<br>(8.21) | -0.02<br>(1.49) | 0.36<br>(4.72) |
| Sri Lanka | 0.19<br>(0.98) | 0.01<br>(0.72) | 0.67<br>(6.16) |
| Thailand | 1.49<br>(36.23) | -0.03<br>(2.92) | 0.39<br>(4.09) |

Table 3B. Bahmani-Oskooee and Rehman (2005) Long-run Coefficient Estimates

| **Country** | **Parameters Estimates** | | |
|---|---|---|---|
| | ***Y*** | ***inf*** | ***e*** |
| India | -3.83<br>(0.18) | -488.60<br>(0.26) | 3.49<br>(0.23) |
| Malaysia | 1.25<br>(27.04) | -5.49<br>(1.83) | -0.75<br>(3.38) |
| Pakistan | 0.58<br>(6.86) | -4.98<br>(2.31) | 0.34<br>(4.64) |
| Philippine | 1.13<br>(5.11) | -22.65<br>(2.24) | -1.09<br>(2.60) |
| Thailand | 0.90<br>(10.92) | -19.82<br>(1.63) | -0.76<br>(1.23) |

Description: See Table 2 for details, however, note that the data reported in Table 3b are taken from Bahamani-Oskooee and Rehman(2005) whose sample period is from 1972:1 through 2000:4.

For comparison, we report estimates of the cointegrating equation using the inflation rate in place of the domestic interest rate in Panel A of Table 3. Also, in Panel B of the same table, we report those taken from Bahmani-Oskooee and Rehman (2005). Without discussing each panel in detail, it is noteworthy that in Panel A, the inflation rate variable has a positive sign and significant in Korea. The same is true for Sri Lanka. For the Philippines and Malaysia, the coefficient on inflation has the correct sign but the t-values are generally small and statistically insignificant at the 5 percent level. Also, note that for Sri Lanka, our real income is only 0.19, and for Malaysia, the exchange rate is statistically non-significant at the conventional levels.

How do these results compare to those of Bahmani-Oskooee and Rehman (2005)? As can be seen in Panel B, coefficient on real income in India is wrongly signed and inflation and exchange rate are statistically insignificant. In a similar vein, the exchange rate variable is insignificant in Thailand. Further, the estimated coefficients on inflation variable, although correctly signed, are of an unusual magnitude. Nevertheless, our real income coefficients are generally consistent with theirs. All in all, the results suggest that the inflation variable is no longer adequate for these countries because their use can lead to misleading inference.

Further, we note that we experimented with the inclusion of foreign interest rate (the rate of U.S. certificate of deposit or U.S. Treasury bill rate) in the cointegration model, that is, testing for the existence of cointegration among real money balances, real income, domestic interest rate, foreign interest rate without much success. Before turning to issue of error-correction modeling, it seems prudent to check whether the variables in Equation (2) are weakly exogenous (Arize, Osang and Slottge, 2000). The results are reported in Table 4.

Table 4. Results of weak exogeneity test

| Countries | Real Balance | Real Income | Interest rate | Exchange rate |
|---|---|---|---|---|
| India | -0.024(1.81)* | 0.050(1.56) | -1.174(4.78)* | 0.022(1.12) |
| Korea | -0.057(3.15)* | 0.100(2.98)* | -2.036(4.06)* | 0.024(1.33) |
| Malaysia | -0.096(5.04)* | 0.030(0.88) | -1.590(3.10)* | 0.040(1.31) |
| Pakistan | -0.095(2.61)* | 0.003(0.28) | -1.490(1.35) | 0.027(1.53) |
| The Philippines | -0.064(2.62)* | 0.054(2.31)* | -0.940(1.32) | 0.021(1.28) |
| Sri lanka | -0.163(3.39)* | 0.090(4.41)* | -6.460(3.88)* | 0.002(0.06) |
| Thailand | -0.042(1.70)* | 0.001(0.03) | -5.170(4.79)* | 0.004(0.15) |

Description: The absolute t-values are in parentheses and the critical value is 1.645.

The data in Table 4 suggest that it may be appropriate to utilize an instrumental variable (IV) estimator when estimating our error-correction models. This is so because the null of weak exogeneity can be rejected in four out of seven cases for the real income variable and interest rate for five out of seven cases. The results also confirm that the real money demand variable should be considered endogenous.

*3.2 Error-correction Model*

The Granger representation theorem (GRT) proves that, if a congregating relationship exists among a set of nonstationary series, then a dynamic error-correction representation of the data also exists. The method used to find this representation follows the "general-to-specific" paradigm. For this purpose; the following error-correction model (ECM) exists for the variable in equation (2):

$$\Delta m_t = k_0 + \mu + \lambda EC^{*}_{t-1} + \omega D_t + \sum_{j=1}^{5}\left(\delta_j \Delta y_{t-j} + \delta_{5+j}\Delta i_{t-j} + \delta_{10+j}\Delta e_{t-j}\right) + \sum_{j=0}^{5}\beta_j \Delta m_{t-j-1}, \tag{4}$$

where $EC_{t-1}$ is the error-correction (one-lagged error) term generated from the FMLS cointegrating estimates, μ is the error term, and $D_t$ are country-specific time dummies that account for observable money demand shocks otherwise not included in the model. The existence of the error-correction term ($EC_{t-1}$) in equation (4) reflects the belief that actual real money balances do not adjust promptly to their long-run determinants; see Arize and Malindretos, (2012) for a detailed discussion.

The error-correction model results are summarized in Table 5. The list of instrumental variables consists of the constant term, the lagged EC term, five lags in the difference of the variables included in the long-run solution. In the case of Pakistan, we did not use any IV procedure. Note that the results from Sargan's test for legitimacy of the instrument set (not reported here) support the validity of the instrument set.

Table 5. Error- Correction Model Regression Results (1980:1-2010:1)

**India**

$$\Delta m_t = 0.015 - 0.024EC_{t-1} + 0.345\Delta m_{t-4} + 0.108\Delta y_{t-2} - 0.001\Delta i_{t-3} + 0.065\Delta e^*_{t-2}$$

(5.92) (1.81) (4.79) (3.23) (1.90) (1.70)

$$R^2 = 0.38, \overline{R}^2 = 0.36, F(5,135) = 16.7, DW = 1.93, BG - F = 1.2\ (0.32)$$

**Korea**

$$\Delta m_t = 0.014 - 0.057EC_{t-1} + 0.311\Delta m_{t-4} - 0.055\Delta y_{t-2} + 0.125\Delta y_{t-4} - 0.005\Delta i_t + 0.70\Delta e^*_t$$

(5.02) (3.15) (4.09) (3.25) (5.63) (3.22) (2.20)

$$R^2 = 0.51, \overline{R}^2 = 0.49, F(6,124) = 21.4, DW = 2.04, BG - F = 1.5\ (0.22)$$

**Malaysia**

$$\Delta m_t = 0.022 - 0.096EC_{t-1} + 0.162\Delta m_{t-1} + 0.146\Delta y^*_t - 0.009\Delta i_t + 0.006\Delta i_{t-1} - 0.134\Delta e_t + 0.143\Delta e_{t-1}$$

(8.30) (5.04) (2.14) (2.94) (4.14) (2.78) (2.25) (2.28)

$$R^2 = 0.33, \overline{R}^2 = 0.30, F(7,130) = 9.3, DW = 1.89, BG - F = 1.3\ (0.27)$$

**Pakistan**

$$\Delta m_t = 0.009 - 0.095EC_{t-1} + 0.085\Delta m^*_t + 0.543\Delta y_{t-2} - 0.013\Delta i_{t-1} + 0.604\Delta e^*_t - 0.742D_1$$

(0.938) (2.61) (1.82) (2.39) (1.63) (3.22 ) (9.81)

$$R^2 = 0.54, \overline{R}^2 = 0.52, F(6,138) = 26.5, DW = 1.84, BG - F = 19.2\ (0.00)$$

**The Philippines**

$$\Delta m_t = -0.005 - 0.064EC_{t-1} + 0.777\Delta m_{t-4} + 0.2\Delta y^*_t - 0.149\Delta y^*_{t-1} - 0.006\Delta i_t + 0.375\Delta e_t - 0.234D_2 - 1.343D_3$$

(1.11) (2.62) (18.70) (2.45) (2.27) (2.14) (4.83) (7.36) (28.58)

$$R^2 = 0.89, \overline{R}^2 = 0.89, F(7,126) = 149.99, DW = 1.957, BG - F = 0.567\ (0.678)$$

**Sri Lanka**

$$\Delta m_t = 0.006 - 0.163EC_{t-1} + 0.134\Delta m_{t-1} + 0.343\Delta m_{t-4} + 0.242\Delta y^*_t - 0.006\Delta i_t - 0.091\Delta e^*_{t-1} + 0.104D_4$$

(1.91) (3.39) (1.57) (4.59) (1.21) (2.33) (2.46) (4.54)

$$R^2 = 0.34, \overline{R}^2 = 0.3, F(7,114) = 8.52, DW = 1.9, BG - F = 0.9\ (0.45)$$

**Thailand**

$$\Delta m = 0.007 - 0.042EC_{t-1} + 0.084\Delta m_{t-1} + 0.567\Delta m_{t-4} + 0.093\Delta y^*_t - 0.098\Delta e^*_{t-1} - 1.15D_5$$

(2.68) (1.70) (3.05) (7.79) (2.50) (1.81) (43.20)

$$R^2 = 0.95, \overline{R}^2 = 0.94, F(7,114) = 368.67, DW = 1.86, BG - F = 0.96\ (0.43)$$

Description: The number in parentheses report absolute t-statistics. The critical value at 10 per cent is 1.67 and 1.96 at 5 per cent level. Tests: DW=Durbin-Watson test statistic, BG =Breusch-Godfrey test. $D_1$=1 2008:1,-1 in 2008:3 and zero otherwise. $D_{2=1}$ in 1984:1, 1 in 1984:3 and zero otherwise. $D_3$=1 in 2008:1, -1 in 2009:1 and zero otherwise. $D_4$ =1 in 1982:1, -1 in 1989:1 and zero otherwise and $D_5$ =1 in 2009:3 and zero otherwise. $\Delta m^*_t$ is $\Delta m_{t-2}$ +$\Delta m_{t-3}$. $\Delta y^*_t$ for Malaysia is $\Delta y$-$\Delta y_{t-4}$. $\Delta y^*_t$ for The Philippines is $\Delta y$+$\Delta y_{t-1}$. $\Delta y^*_t$ for Sri Langa is $\Delta y$+$\Delta y_{t-1}$. $\Delta y^*_t$ for Thailand is ($y_{t-1}$-$\Delta y_{t-4}$)-($\Delta y$-$\Delta y_{t-3}$). $\Delta e^*_{t-2}$ for India is ($\Delta e_{t-1}$ - $\Delta e_{t-4}$) $- \Delta e_{t-2}$ . $\Delta e^*_t$ for Korea is $\Delta e_t$ -$\Delta e_{t-4}$ . $\Delta e^*_t$ for Pakistan is $\Delta e$+$\Delta e_{t-1}$. $\Delta e^*_{t-1}$ for Sri Lanka is $\Delta e_{t-2}$+$\Delta e_{t-4}$. $\Delta e^*_{t-1}$ for Thailand is $\Delta e_{t-1}$+$\Delta e_{t-4}$.

The regress and in Table 5 is the change in real M2 balances as the dependent variable. Focusing on these results, we conclude that the short-run impact of any explanatory variable on money demand is positive (negative) if the sum of the statistically significant point estimates of the covariate in question in equation (4) is positive (negative). For each country, a parsimonious and statistically acceptable model is obtained through the simplification of a fifth-order ECM. The simplification process involves deleting successively the first-differenced variable with the lowest t-ratios.

Considering that each regress and in the table is cast in the first difference, the empirical results suggest that the statistical fit of the each model to the data is satisfactory, as indicated by the value of adjusted coefficient of determination, which ranges from 0.30 (Sri Lanka and Malaysia) to 0.94 (Thailand). The statistical appropriateness of the models is further supported by a number of diagnostic tests. In particular, we report the Durbin-Watson and the Breusch-Godfrey test to test for non-independence of the error distribution. Based on these two statistics, we can reject the null of non-independence for all countries. Parameter stability tests reveal no serious violation of stable parameters. For example, we experimented with a t-test of the null hypothesis that the mean of the recursive residuals for each of our estimated models is not statistically significant. A similar finding was obtained by using the Dufour test. Furthermore, to ensure that no serious violation of the linearity assumption in the structure of the model, we implemented two nonlinear cointegration tests-- Kapetanios, Shin and Snell (2003) test as well as the Sollis (2009) – along the lines suggested by Ghoshray (2010). The results are reported in the Appendix.

Table 6. Speed of adjustments and mean time lags for adjustments of desired real imports.

| Countries | Speed of Adjustments<br>Desired Real Imports | Response of Desired Real Imports to Each Regressor<br>Half-Life Adjustment | Mean Time Lag | | |
|---|---|---|---|---|---|
| | ***m*** | ***ψ*** | ***y*** | ***i*** | ***e*** |
| India | -0.024 (0.013) | 7.13 | 5.70 | 6.83 | 6.15 |
| Korea | -0.057 (0.018) | 2.95 | 2.71 | 3.04 | -0.05 |
| Malaysia | -0.096 (0.019) | 1.72 | 1.80 | 2.19 | 2.16 |
| Pakistan | -0.095 (0.036 | 1.73 | 0.98 | 2.44 | 0.82 |
| The Phillipines | -0.064 (0.024) | 2.62 | 0.67 | 0.89 | -0.59 |
| Sri Lanka | -0.163 (0.048) | 0.97 | 0.43 | 0.81 | 0.94 |
| Thailand | -0.042 (0.025) | 4.04 | 1.52 | 2.08 | 2.66 |
| Average | | 3.02 | 1.97 | 2.61 | 1.73 |

Description: The values in parentheses beside the speed of adjustments are the stand errors. Both the half-life and the mean time lag are in absolute terms and in quarters

Having provided evidence supporting the adequacy of the estimated equations, several features of the results in Table 5 deserve mention. First and foremost, the coefficient of the error-correction term is statistically significant in each of the seven cases and is always negative. Therefore, the validity of an equilibrium relationship among the variables in the cointegrating equation is supported. This implies that overlooking the cointegratedness of the variables would have introduced misspecification in the underlying dynamic structure and note that there is no evidence of possible nonlinear cointegration. In the case of the Philippines, the possibility of asymmetric adjustment exists given the significance of Sollis test.

Second, the change in real money balancer per quarter that is attributed to the disequilibrium between the actual and equilibrium levels is measured by the absolute values of the EC term of each equation. There is considerable intercountry variation in the adjustment speed to the last period's disequibrium, with Sri Lanka having the largest and India, the smallest. The average coefficient is -0.077 per quarter. This implies that the adjustment of real balance to changes in the regressors may take about six quarters in Sri Lanka to slightly more than thirty-eight quarters in India. In economic terms, this means that when real balances exceed their long-run relationship with their regressors, they adjust downwards at a rate of 6 to 38 percent in each quarter. Our results concerning the coefficient of the EC term for India, conflict with those of Bahmani-Oskooee and Rehman (2005), because they obtain -0.006 which is statistically insignificant with a t-value of 0.26. However, their results of -0.12, -0.15, -0.21 and -0.30 for Malaysia, Pakistan the Philippines and Thailand, respectively, are similar to those reported in Table 5 and are (in absolute terms) within the unit interval. Taken together, these results point to the existence of market forces in the money market that operates to restore long-run equilibrium after a short-run disturbance.

Finally, the dynamics of the equation show that changes in real income, interest rate and exchange rate have short-run effects on real balances in addition to their long-run effects established earlier. Therefore, they are consistent with the theory.

In sum, our findings imply (a) that not noticing the cointegration of the variables would have introduced a misspecification in the underlying dynamic structure; (b) that use of interest rate as the opportunity cost of holding money performs better than the expected inflation in the estimated Asian money demand relations; (c) that there exist market forces in the money market sector that operate to restore long-run equilibrium after a short-run deviation; (d) that, on average, it takes more than three years for 50 per cent of the deviations from long-run equilibrium (the cointegration equation) to be corrected, although much longer for India (see Table 6); (e) that the mean time lags suggest that real balances react faster to changes in domestic income than to changes in exchange rates; (f) that real balances react faster to exchange rate changes than to interest rate changes, therefore, had we focused only on the traditional variables (income and interest rate), we would have obtained more inference bias and missed the key role that exchange rates play in the demand for money function of Asian economies; (g) that linear error correction model is quite likely a good approximation of the short-run dynamics because both the KSS and Solis methods revealed insignificant t-values and (f) that the full adjustment of real balances to changes in the regressors may take between six to thirty-eight quarters, depending on the country.

## 4. Concluding Remarks

As set out in the introduction, the primary aim of this paper has been to investigate empirically the fundamental issue of whether exchange rate changes influence the demand for money in Asian countries by using recent breakthrough time series econometrics. Unlike previous research, this paper uses interest rate as the opportunity cost of holding money. Therefore, the basis of our analysis is a money demand function in which real money balances depend on real income, short-term interest rate and exchange rate.

Our results suggest that there is a unique, statistically significant long-run equilibrium relationship among the real M2 balance, real income, interest rates and exchange rate in each of the seven countries. The addition of exchange rate to the money demand model was found to be quantitatively important in improving money demand formulation. Exchange rate was positively signed in all cases and implies that an increase in exchange rate (i.e., depreciation of, for example Thai baht), would increase the demand for real money balances – that is, the value of foreign securities owned by domestic residents rises; also, the value of domestic securities in the hands of foreigners is lowered. This, in turn, increases the demand for real balances. Our results are directly relevant to concerns about which monetary aggregates best determines the long run effects of monetary policy action in Asia.

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**Notes**

Note 1. Specifically, our sample period for each country ends in 2009:4, however, for India, Malaysia and Pakistan the starting quarter is 1973:1. In the case of the Philippines and Thailand, it is 1977:1; for Sri Lanka, it is 1978:1, and for Korea, it is 1976:1. Availability of interest rate data was a major factor in determining our starting quarterly period.

Note 2. See the discussions in the *Review of Economics and Statistic*s (RESTAT), 79, May (1997: 311-328); see RESTAT, 80, August (1998: 400-403; *Journal of Business and Economic Statistics*, 16, October 1998; Arize, Osange and Slottje (2000); Funke (2001) and Arize (2012).

Note 3. It is imperative that we estimate the demand for money in real terms, because we want to understand how the real demands for money balances behave over time, so that changes in the nominal stock of money are adapted to the growing real demand at constant prices. See also Goldfeld (1976: 624).

Note 4. See Domowitz and Hakkio (1990:30) for a detailed explanation of the importance of deflating by consumer price index and using real GDP as scale measure.

Note 5. Mundell (1963) has argued that the demand for money is likely to depend upon the exchange rate, in addition to the interest rate and the level of income. A portfolio model of the financial market explored by Arango and Nadiri (1981) provide a role for the exchange rate variable in the money demand function; however they point out that the effect of the exchange is more difficult to sign. See, also McKinnon (1963).

Note 6. The exposition draws from Arango and Nadiri (1979, 1981), Arize (1989), Bahmani-OsKooee and Rhee (1994) and Bahmani-OsKooee and Techaratanachai (2001).

Note 7. Rebalancing is the process of adjusting the allocation of various assets in a portfolio to reflect the new scenario or to achieve or maintain a desired mix.

Note 8. The consequence of nonstationarity is the inapplicability of the standard sampling theory. The cointegration approach is attractive in that it can properly account for the nonstationary series.

Note 9. Following the specification suggestion in Funke (2001) we base our coefficient analysis on FMLS estimator.As Funke (2001) points out “one of the key advantages of this approved over alternative estimators, such as Johansen (1988) and Stock and Watson (1988), is that it facilitates a complete analysis of the inclusion of deterministic trends in the cointegration set. An additional benefit of the above framework is that it facilitates a test of cointegration, where cointegration is taken to be the null hypothesis. In the statistics literature this would be the natural way to test for cointegration…”

Appendix. Nonlinear Tests of the Residents of the Long-run Equation.

| Country | Sample Period | KSS | Sollis | |
|---|---|---|---|---|
| | | | $H_0: \emptyset_1 = \emptyset_2 = 0$ | $H_0: \emptyset_2 = 0$ |
| India | 1980:1-2011:10 | -0.97 | 1.14 | 1.51 (0.13) |
| Korea | 1980:1-2011:10 | -1.59 | 1.68 | -0.22 (0.82) |
| Malaysia | 1980:1-2011:10 | -1.16 | 2.33 | 0.58 (0.56) |
| Pakistan | 1980:1-2011:10 | -1.15 | 26.64 | 0.22 (0.83) |
| Philippines | 1980:1-2011:10 | -1.24 | 4.39 | -2.21 (0.03)* |
| Sri Lanka | 1980:1-2011:10 | -0.55 | 5.12 | -1.49 (0.14) |
| Thailand | 1980:1-2011:10 | -1.57 | 2.67 | 1.46 (0.15) |
| **Critical Value (10%)** | | 2.66 | 3.73 | |

Description: The values in parentheses are the p-values.

a. Kapetanios, G., Shin,Y., and Snell, A. (2003) was obtained from

$$\Delta m_t = \lambda EC_{t-1}^3 + \sum_{i=0}^{5} \theta_i \Delta y_{t-i} + \sum_{i=0}^{5} \vartheta_i \Delta i_{t-i} + \sum_{i=0}^{5} \beta_i e_{t-i} + \sum_{i=0}^{5} \gamma_i \Delta m_{t-i} + v_t$$

$H_0$: $\lambda = 0$ and $H_a$: $\lambda < 0$

b. Sollis (2009) was obtained from

$$\Delta m_t = \phi_1 EC_{t-1}^3 + \phi_2 EC_{t-1}^4 + \sum_{i=0}^{5} \theta_i \Delta y_{t-i} + \sum_{i=0}^{5} \vartheta_i \Delta i_{t-i} + \sum_{i=0}^{5} \beta_i e_{t-i} + \sum_{i=0}^{5} \gamma_i \Delta m_{t-i} + w_t$$