

Foreign Exchange Reserves in Asia and Its Impact on Import Demand

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Abstract

Some Asian countries have experienced increases in the level of their foreign exchange reserves as well as increases in their import volume. Theory suggests that as the level of exchange reserves increases, it may affect the demand for imports since more funds will be available for imports. In this paper, we employ quarterly data of five Asian countries and test the null hypothesis that the import demand behavior in India, Japan, Korea, Singapore and Thailand are not determined by real income, relative import price and foreign exchange reserves. The empirical analysis of import demand behavior is presented using the dynamic error-correction model, which allows an explicit parameterized division of effects into long-run influences, short-term adjustment and error-correction term. It uses econometric techniques organized around Johansen, Harris-Inder, and Hansen Lcointegration analyses; fully modified OLS, dynamic OLS and ARDL to estimate long-and-short run demand elasticities.

Keyword: Import demand, Foreign exchange reserves, Cointegration

JEL: D53, E44

1. Introduction

Over the past one and half decades, Asian economies have exhibited the most noticeable increase in foreign exchange reserves as well as increases in their import volume. For example, besides China, Japan leads the world in foreign reserve ranking with over \$1 trillion in 2010. The foreign reserves (rank) for India, Korea, Singapore and Thailand are \$287 billion (8th), \$295 billion (6th), \$226 billion (11th) and \$176 billion (13th), respectively.¹ The sources of these reserves have been export earnings, remittances of Asians residing abroad, and in few cases, foreign assistance. As a consequence, in the period 1973 - 2010, imports by these countries grew at a higher rate. Imports averaged 18.88, 4.83, 12.20, 7.15 and 14.25 percent per year in India, Japan, Korea, Singapore and Thailand, respectively, and the variation in imports has been large and wide with the coefficient of variation ranging from a high of 1.64 in India to a low of 0.44 in Japan.

Foreign reserves play an important role in the design and evaluation of current and future macro policies aimed at achieving the trade balance. In countries with fixed or partially flexible exchange rates, the reserves are mainly used to maintain competitiveness of the tradable sectors. They achieve this by preventing the appreciation of their currencies and by keeping the exchange rate at or close to the official target level. Beyond exchange rate stabilization, foreign reserves are generally viewed as indicators of the strength of an economy, especially in particular itsexporting industries. From a policy perspective, foreign reserves influence trade policies. A high level

of foreign exchange reserves is associated with less restrictive policies. With regard to international trade, foreign currency is often an indispensable requirement to finance imports of goods and services. In this sense, anecdotal evidence suggests that foreign reserves play the role of an international liquidity constraint and any increase in reserves should thus have a positive impact on import demand. The degree to which this hypothesis, among others, is 'genuine' for Asia is what this study will attempt to test empirically using data from five Asian countries (ACs): India, Japan, Korea, Singapore and Thailand.

Empirical estimation of import-demand function has generally related the quantity of imports to real domestic income and relative import prices (see, for example, Thursby and Thursby, 1984; Bahamani-Oskooee, 1986; Arize and Afifi, 1987; Arize and Ndubizu, 1992; Gafar, 1995; Senhadji, 1998; Masih and Masih, 2000 and Chen, 2008). Following the literature, we thus include a measure of domestic income as well as the relative price of imports in our empirical model, in addition to holdings of foreign exchange reserves. A number of studies have identified real foreign exchange reserves as an additional variable influencing import demand (see, for example, Moran, 1989; Faini et al. (1992); Dutta and Ahmed, 1999; and Arize et al., 2004; Arize and Osang, 2007; Sultan, 2011). Thus, omission of exchange reserves may bias a model's empirical estimates towards overstating the influence of the included variables, in particular domestic income and relative prices.

Given the time-series nature of the data, we are interested in differentiating between the long-run and the short-run impact of reserves, income and relative import prices on real imports provided there exists a long-run equilibrium relationship among real imports, real income, relative import prices and real foreign exchange reserves for each ACs in our sample. Previous studies for India by Dutta (1964), Kantu (1972) and Pani (1977) conclude that a positive relationship exists between real imports and real foreign exchange reserves; however, these studies assumed data stationarity, and recent developments in econometrics suggest that macroeconomic data are non-stationary; hence, their findings need to be reexamined. For a more detailed discussion of these issues, see Arize, Osang and Slottje (2000), Dutta and Ahmed (1999), who examine the demand for imports in Bangladesh, confirm the existence of a long-run equilibrium relationship between real imports, real import prices, real GDP and real foreign exchange reserves. Arize et al. (2004), who investigate import demand for Pakistan, also find that real imports, real income, relative price and real foreign reserves share a long-term relationship. Furthermore, they find a significant positive long-run relation between import demand and foreign exchange reserves, but conclude that in the short run the impact of foreign reserves on import demand is statistically insignificant. Arize and Osang (2007) show that the real foreign reserves coefficient is positive and statistically significant with an elasticity of about one fourth and their results also suggest that real foreign reserves have long and short term effects on import volume in the case of seven Latin American countries. Further empirical support for the effect of real foreign exchange reserves on real imports has also come from the work of Sultan (2011). His estimates show a fairly large response of real imports to changes in real foreign exchange reserves using India's annual data. Further, he presents evidence that real foreign reserves exert both a significantly positive effect on real imports in both the short-run as well as in the long-run.

The purpose of this paper is fourfold. The first is to check the data for unit roots. Second, we examine whether a long-run relationship exists among real imports, real domestic income, relative import price and real foreign exchange reserves for each country and whether this relationship is unique. Third, provided the long-run equilibrium exists, we determine the sign, magnitude and statistical significance of the long-run effects of real income, relative price and real foreign exchange reserves on import demand. Fourth, we investigate the short-run dynamics of the import demand function for each country. We use data for quarterly period 1973:2 through 2005:3, (i.e., 130 observations) for estimation, and forecasting is performed employing data for period, 2006:1 through 2008:4 (12 observations). Finally, we draw policy conclusions.²

The rest of this paper is as follows. Section 2 describes the empirical import demand model. Section 3 reports the empirical results. Section 4 contains the major conclusions and policy implications of the paper.

2. Methodology

2.1 Model Specification and Theoretical Considerations

Drawing on the empirical literature in this area (Arize, Malindretos and Grivoyannis, 2004) 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 (see Thursby and Thursby, 1984; Arize and Afifi, 1987), augmented with real exchange reserves as the long-run equilibrium (equation 1) and the short-run relationships (equation 2) for the (desired) real imports:

$$m_t^* - \delta_0 - \delta_1 y_t - \delta_2 p_t - \delta_3 IR_t = \varepsilon_t \quad (1)$$

$$\Delta m_t - \alpha - \beta L(\Delta X_t) - \gamma \varepsilon_{t-1} = \psi_t \quad (2)$$

where m_t^* is the logarithm of (desired) real imports; y_t is the logarithm of real domestic income (i.e. real GDP); p_t is the logarithm of the relative price of imports, defined as the ratio of the import price index (P_t^m) to the domestic wholesale price index (P_t^d); IR_t is the logarithm of the real foreign exchange reserves, defined as foreign exchange reserves deflated by the wholesale price index; and ε_t is a stochastic disturbance term.

Equation (1) embodies the hypothesis that the demand for imports with respect to real domestic income would be positive. Real imports would be expected to increase with real income for two reasons. First, if an increase in real income leads to an increase in real consumption, with an unchanged distribution of income, more foreign goods will be purchased. Second, if an increase in income also leads to an increase in real investment, then investment goods not domestically produced must be bought from abroad.² On the other hand, the effect of the relative price on the demand for imports is expected to be negative, as consumers substitute domestic products for imports when the price of imports increases. Inasmuch as foreign reserves constitute a limit on the size of excess import demand, an increase in real foreign exchange reserves is expected to have a positive on the demand for imports.

Before presentation of the empirical results, two technical notes regarding equation (1) and the method of estimation are in order. The first issue is whether import demand function should be specified in log-linear or linear terms. Following the 'best practice' in most of the empirical literature, we estimate the import equations in log-linear format.

The second issue is that equation (1) assumes that the import market is in a (long-term) equilibrium so that it may be viewed as a cointegrating model. The basic idea of cointegration is that two or more non-stationary 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). Thus, if the import demand function describes a stationary long-run relationship among the variables in these equations, this can be interpreted to mean that the stochastic trends in real imports are related to the stochastic trends in the real income, real foreign exchange reserves and relative prices. In other words, even if deviations from the equilibrium should occur, they are mean reverting, which means that the variable will revert back to its (long-term) equilibrium level over time (Arize, Osang and Slottje, 2000).

3. Empirical Results

3.1 Data Issues

The Empirical analysis uses quarterly frequency data for India, Japan, Korea, Singapore, 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 import, real GDP, the relative import price (import prices divided consumer prices) and real foreign exchange reserves (exchange reserves divided by wholesale prices). The observations covering the period 1973:2 through 2005:4 were used as the estimation period whereas the data from 2006:1 to 2008:4 were used as the forecast period. While the plots of the level of the logs of all variables are available from the authors, for space considerations, we only present those of the log of real exchange reserves for each country. This is shown in Figure 1 where real exchange reserves trend upwards in each case.

Insert Figure 1 Here

A prerequisite for testing the existence of a long-term equilibrium (cointegration) among the four variables is to test for stochastic trends in the autoregressive representation. The common practice is to use the augmented Dickey-Fuller (ADF) to investigate this issue. The ADF test statistics – in levels and in differences- for the null hypothesis of non-stationary (including a constant and time trend in the regression equation) for m_t , y_t , P_t and IR_t are given in Table 1.

Insert Table 1 Here

The critical value is -3.41 at the five per cent level. For the ADF test statistics in levels, this implies that if the value of the test statistic is smaller than -3.41, we would reject the null hypothesis on non-stationarity for the variables in levels. Since none of the values in Table 1 happens to meet this threshold, we conclude that all variables are non-stationary in all countries. This conclusion is supported by the test statistics in differences. This time, all test statistics for the differenced variables are smaller than -2.86, so the null hypothesis of nonstationary can be rejected.⁴

Having established the non-stationarity of variables in all countries, the next step is to test for the presence of absence of a long-run equilibrium among the four variables. This so-called cointegration test can be performed in various ways, forms and shapes -- see the results in Table 2. Here, we have performed four of these tests since we do

not wish to rely on only one. The four cointegration tests are the Johansen test, the Harris and Inder test, the Shin's test and the Hansen Lc test.

Insert Table 2 Here

The Johansen procedure employs two likelihood ratio (LR) test statistics, the maximal eigenvalue (λ -max), and trace (Tr) to test the presence or absence of long-run equilibria between the variables in Eq. (2). For λ -max and Tr statistics, the hypotheses are $H_0: rk(\Pi) = r$ against $H_1: rk(\Pi) = r + 1$ and $H_0: rk(\Pi) = r$ against $H_1: rk(\Pi) = r + 1$, respectively. Note that rk is rank, Π is a matrix of long-run responses, and the matrix Π has rank r , $rk(\Pi) = r$. If the data cointegrate, Π must be of reduced rank, $r < N$, where N is the number of variables.

For the implementation of the Johansen cointegration test, we determine the optimal number of included lags in each cointegration equation by employing several information criteria tests.⁵ For the four-variable system, we employ the lag order of two for India and Japan, three for Korea and four for Thailand and Singapore, respectively. Without discussing each of the remaining test statistics in detail, we also examined whether any variables can be considered weakly exogenous using procedures discussed in Arize, Osang and Slottje (2000).

Table 2 presents the cointegration test results, where r denotes the number of cointegrating vectors.⁶ The Johansen cointegration test procedure gives answer to two questions: whether a long-run equilibrium relationship exists and whether it is unique. To this end, it uses two related test statistics: the maximum eigenvalue and the trace statistics. For both statistics, the null hypothesis is that there are (at most) r cointegrating vectors, whereas under the alternative hypothesis there are (at least) $r + 1$ cointegrating vectors. The test procedure is thus a nested sequence of test for various levels of r , where each level of r is associated with a different critical value of the test statistic.

Focusing first on the maximum eigenvalue 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 2 are significantly largest than the critical value of 27.42 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 $r \geq 1$ 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.

Concerning the three remaining tests in Table 2, we note that the Harris-Inder test as well as the Shin's test is conducted with three, six and eight lags, and the overall results suggest that the null hypothesis of stationarity (or cointegration) is not rejected at all lag lengths in the case of Singapore for these tests. However, as the authors suggest some degree of augmentation in the tests is needed for better results. As the data show in all cases at higher lags, the null hypothesis of cointegration is retained. In other words, our findings indicate the existence of a unique long-run equilibrium relationship among real imports, real income, relative import prices and real foreign exchange reserves for each ACs in our sample.

3.2 Long-run Dynamics

Next, we are interested in the long-run impact of the three explanatory variables (real income, relative import price and reserves) on import 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. Due to our log-log specification of the estimation model (see equation (1)), the resulting values for real income, relative import price and real foreign exchange reserves are long-run elasticities. In Table 3, we show the empirical results of Equation (1) using three different approaches, the fully modified least squares (FMLS) estimator of Phillips and Hansen, the dynamic least squares (DLS) estimator of Stock and Watson and the Autoregressive-distributed lag (ARDL) estimator.⁸

Insert Table 3 Here

Focusing on the results obtained from the FMLS estimator⁹, we gather that the demand relation is estimated to be positive for real income (y) and real foreign exchange reserve (IR), but negative for the relative price of imports (p). This holds for all ACs in the sample. Furthermore, all coefficient estimates are statistically significant at the conventional levels. These results provide strong support for the theoretical predictions regarding the impact of income, relative prices and foreign reserves on imports. As far as relative strength of the various determinants is

concerned, we find that the magnitude of the foreign exchange effect is substantially smaller than the income effect and the relative price effect. The average income elasticity across all ACs is 1.03 and the average price elasticity is -0.302. Both averages exceed (in absolute terms) the average of 0.152 obtained for foreign reserve elasticity. This implies that the economic importance of both income levels and relative price effects far outweighs the importance of foreign reserve variations for import demand.

As a cross-check, we also obtained long-run estimates by using two alternative estimation procedures, namely the dynamic OLS estimator of Stock and Watson and the autoregressive distributed lag estimator. The obtained elasticities are consistent with those of FMLS also reported in reported in Table 3. From these estimates, we gather that our conclusions are not particularly influenced by the method of estimation.

Table 4 summarizes elasticities from previous studies of Asian countries. In general, the elasticity estimates obtained in this study accords well with some previous studies. For example, the income elasticities in Table 4 reported in Bahmani-Oskooee and Niroomand (1998) for Korea and the results in Senhadji (1998) including Tang (2006) for Singapore support at least in spirit the findings of our study. However, they differ markedly from the two vectors obtained by Bahmani-Oskooee and Niroomand (1998) for Japan and Bahmani-Oskooee (1998) for Korea and Singapore, respectively. Since all our income elasticities are positive, they differ markedly from the negative one (-1.032) obtained by Sinha (1997). Our real foreign exchange reserves elasticity estimates are less than 0.3 for all countries and are thus similar to the ones reported by Khan and Knight (1988), Dutta and Ahmed (1999), Arizeet al. (2004) and Sultan (2011).

Insert Table 4 Here

3.3 Error-correction Model

The Granger representation theorem (GRT) proves that, if a cointegrating relationship exists among a set of nonstationary series, then a dynamic error-correction representation of the data also exists. The existence of GRT involves two doctrines by Engle and Granger (1987) and Johansen (1991).¹⁰The methodology used to find this representation follows the “general-to-specific” paradigm (see Hendry and Doornik, 1994). For this purpose, the following error-correction model (ECM) exists for the variable in equation (1):

$$\Delta m_t = \mu + \lambda EC_{t-1} + \omega D_t + \sum_{j=1}^5 (\delta_j \Delta y_{t-j} + \delta_{5+j} \Delta p_{t-j} + \delta_{10+j} \Delta IR_{t-j}) + \sum_{j=0}^5 \beta_j \Delta m_{t-j-1}, \quad (3)$$

where EC_{t-1} is the error-correction (one-lagged error) term generated from the FMLS cointegrating estimates, μ is a constant, and D_t are country-specific time dummies that account for observable import demand shocks otherwise not included in the model. The presence of the error-correction term (EC_{t-1}) in equation (3) reflects the presumption that actual real imports do not adjust instantaneously to their long-run determinants. Therefore, the equation gives the short-run determinants of import demand and embodies both the short-run dynamics and the long-run relation of the series. In the short run, adjustments are made to correct any disequilibrium in the long-run import demand.

The parameter λ is the error-correction coefficient and measures the response of the real imports in each period to departures from equilibrium conditions. The ECM therefore reflects how the system converges to the long-run equilibrium implied by equation (1), with convergence being assured when λ is between zero and minus one. In addition, the value of λ depends on the normalization of the cointegrating vector.

Table 5 presents the results for the change in real imports as the dependent variable. We conclude that the short-run impact of any explanatory variable on import demand is positive (negative) if the sum of the statistically significant point estimates of the covariate in question in equation (3) is positive (negative).

Insert Table 5 Here

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. Our findings are reported in Table 5.

Considering that each regressand in Table 5 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 0.372 (Thailand) to 0.636 (India). 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. Furthermore, Ramsey's RESET test reveals no serious omission of variables or violations of the linearity assumption in the structure of the model.

Having provided evidence supporting the adequacy of the estimated equations, we can make the following observations regarding the obtained estimates:

First and foremost, since the sum of the estimates on current and lagged values of (ΔIR_t) is positive for all countries, we conclude that exchange reserves has a positive short-run effect on real imports in addition to its long-run effect established earlier. Second, concerning the short-run dynamics of the relevant variables they are consistent with the theory and in line with their long term effects.

Third, it is worthwhile to interpret the finding with respect to the error-correction coefficient's term. From the data in Table 5, this coefficient term is statistical significant in each of the five cases and is always negative. The average coefficient is -0.2278 per quarter. For example, only about twenty-three percent occurs in one quarter. In addition, since the coefficient estimates range from -0.128 to -0.320 and are thus within the unit interval, convergence to the long-run equilibrium is assured for each country. These findings support the validity of an equilibrium relationship among the variables in each cointegrating equation. In economic terms, this means that when real imports exceed their long-run relationship with real income, relative import prices and real foreign exchange reserves, they adjust downwards at a rate of 13 to 32 percent in each quarter.

Insert Table 6 Here

In sum, our findings imply (a) that overlooking the cointegration of the variables would have introduced a misspecification in the underlying dynamic structure; (b) that there exist market forces in the imports market sector that operate to restore long-run equilibrium after a short-run deviation; (c) that it takes less than a year for 50 per cent of the deviations for long-run equilibrium (the cointegration equation) to be corrected, but a little above one year in Singapore (see Table 6); (d) that the mean time lags suggest that imports react faster to changes in domestic income than to changes in exchange reserves, therefore, ignoring exchange reserves can produce biased results due to misspecification; and (e) that the full adjustment of real imports to changes in the regressors may take between two and five quarters, depending on the country.

To provide the reader with some insight into the behavior of real imports over time, we first use the Chow test and treat the break date as known. We chose the breakpoint to keep each subsample roughly the same size. The test was implemented using intercept and slope dummies (see Gujarati, 2002, for details). Our results suggest that only the model for Singapore is unstable at the 10 percent level though not at the 5 percent level. Tests suggested by Nyblom (1989) and Hansen (1992) also find instability in Singapore at the ten percent level. To examine the forecasting ability of the error-correction models, we split our data into estimation and forecast periods. The start of the sample until 2005:4 is treated as the estimation period, whereas the forecast period (out-of-sample period) starts from the sample period 2005:4 through 2008:4. The Theil U test results are as follows: the Theil U coefficients are 0.011, 0.003, 0.021, 0.33, and 0.01 for India, Japan, Korea, Singapore, and Thailand, respectively. A Theil U statistic value between zero and one ($0 < U < 1$) would suggest that the model yields more accurate prediction than would be a naïve model.

4. Conclusions and Policy Implications

This paper explores empirically the long-run and short-run impact of domestic income, relative import prices and foreign reserves on real imports of selected Asian economies. After examining this relationship, previous research, and nonstationarity of data, we use several techniques to explore empirically this possible relationship. Our result indicates that an increase in foreign reserves may have a positive effect on the demand for imports since it relaxes the excess demand liquidity restriction, at least in theory. It tends to influence trade policies, especially those associated with trade liberalization. The paper examines whether this hypothesis holds over both the long and short term using quarterly data over the 1973:2 – 2005:4 period. Forecasting was conducted using data from 2006:1-2008:4.

Our empirical results suggest the following conclusions. First, the evidence points to the nonstationarity of all variables involved in the model. Therefore, the treatment of this nonstationarity is vital for meaningful subsequent results. In addition, further test results suggest that there is a unique, statistically significant long-term equilibrium relationship among real imports, real income, relative price and real foreign exchange reserves.

Second, we find that foreign exchange reserves matter for import demand both in the long and in the short run. In addition, the sign of the estimated coefficient of reserves is positive, as expected. While the statistical impact is significant, the economic impact of foreign reserves on import demand is small, in particular in comparison to the size of the estimated elasticities for real income and relative prices. Nevertheless, policies which focus on increasing foreign exchange reserves should be pursued, as they are likely to influence import behavior. For instance, policies would have to be oriented towards export promotion. Exports are likely to increase reserves and can thereby provide

greater access to international markets (Esfahani, 1991). Trade policy measures also should discourage efforts aimed at borrowing given that reserves are available. Also, had we neglected foreign exchange reserves, our estimated model would have overstated – in absolute terms – the elasticity estimate of the income variable and the relative price variable.

Third, our empirical estimates show that real income is a significant variable in explaining the demand for imports, and that income elasticity is highly elastic in India, Korea and Thailand, whereas, it is inelastic in Japan and Singapore. The high income elasticity in India, Korea and Thailand implies that increases in real income are likely to lead to rising imports and ultimately to large trade imbalances. The relevant policy implication is that actions are needed in these countries to bring the income coefficient to less than or equal to unity. It is important that the management of import demand has to be viewed as part of a comprehensive stabilization plan. As part of this effort, imports should be targeted to offset shortfalls in domestic production or to change its nature and composition. Further, income growth could be cut by strategies that reduce government expenditure or those that raise taxes. For Japan and Singapore, our results indicate that import demand is inelastic with respect to real income and thus submits that a growth income may not lead to the expectation of trade deficits. Also, taken together, the estimates for the timing of import responses to changes in income are similar across most of the countries. The overall mean time-lag is 3.14 quarters so that a strong domestic growth in these five countries is likely to stimulate the export activities of their respective trading partners.

For the relative price variable, the evidence suggests that they play a vital role in the determination of imports. The long-run elasticity is less than one (in absolute terms) in each country and would tend not to enhance an expansionary policy even if such a policy is accompanied by depreciating exchange rate. Further, the evidence indicates that it would take a little above a year (4.93 quarters) for a change in relative price to influence the quantity imported in these countries. To the extent that price lags provide information about the shape of the “J-curve” these findings suggest that currency depreciation will be reflected with a significant delay in import volumes.

Finally, the significance of relative prices in the our import demand equations has an important implication for the effectiveness of exchange rate policy or commercial policy aimed at correcting trade imbalances and promoting export growth. A fall in the quantity of imports will largely compensate for the increase in price of imported items. This implies that the Marshall-Lerner (M-L) condition is very likely achieved, if the demand price elasticity for exports is substantially large as expected.

References

- Arize, A. C., & Ndubizu, G. (1992, December). Error-Correction Representation and Import Demand Function With Implications for International Finance and Accounting. *Review of Quantitative Finance and Accounting*, 2, 359-376. <http://dx.doi.org/10.1007/BF00939017>
- Arize, A.C., Osang, T., & Slottje, D. J. (2000). Exchange-Rate Volatility and Foreign Trade: Evidence from Thirteen LDCs. *Journal of Business and Economic Statistics*, 18, 10-17. <http://dx.doi.org/10.2307/1392132>
- Arize, A.C., & Afifi, R. (1987). An Econometric Examination of Import Demand Function in Thirty Developing Countries. *Journal of Keynesian Economics*, 4, 604-616.
- Arize, A.C., Malindretos, J., & Grivoyannis, E. C. (2004). Foreign Exchange Reserves and Import Demand in a Developing Country: The case of Pakistan. *International Economic Journal*, 18(2), 259-74. <http://dx.doi.org/10.1080/1016873042000228376>
- Arize, A.C., & Nippiani, S. (2010). Import Demand Behavior in Africa: Some New Evidence. *The Quarterly Review of Economics and Finance*, 50, 254-263. <http://dx.doi.org/10.1016/j.qref.2010.02.003>
- Arize, A. C., & Osang, T. (2007). Foreign Exchange Reserves and Import Demand: Evidence from Latin America. *The World Economy*, 30, 1477-1489. <http://dx.doi.org/10.1111/j.1467-9701.2007.01052.x>
- Bahmani-Oskooee, M. (1986). Determinants of International Trade Flows: The case of Developing Countries. *Journal of Development Economics*, 20(1), 107-23. [http://dx.doi.org/10.1016/0304-3878\(86\)90007-6](http://dx.doi.org/10.1016/0304-3878(86)90007-6)
- Bahmani-Oskooee, M. (1998). Cointegration Approach to Estimate the Long-Run Elasticities in LDCs. *International Economic Journal*, 12(3), 89-96.
- Chen, S. (2008). Long-Run Aggregate Import Demand Function in Taiwan: An ARDL Bounds Testing Approach. *Applied Economic Letters*, 15, 731-735. <http://dx.doi.org/10.1080/13504850600749032>
- Dutta, M. (1964). A prototype model of India's foreign sector. *International Economic Review*, 5, 82-103. <http://dx.doi.org/10.2307/2525635>

- Dutta, D., & Ahmed, N. (1999). An Aggregate Import Demand Function for Bangladesh: A Cointegration Approach?, *Applied Economics*, 31(4), 465-72. <http://dx.doi.org/10.1080/000368499324174>
- Engle, R. F., & Granger, C. W. J. (1987). Cointegration and Error Correction: Representation, Estimation and Testing. *Econometrica*, 55, 251–276. <http://dx.doi.org/10.2307/1913236>
- Esfahani, H. S. (1991). Exports, Imports, and Economic Growth in Semi-Industrialized Countries. *Journal of Development Economics*, 35, 93–116. [http://dx.doi.org/10.1016/0304-3878\(91\)90068-7](http://dx.doi.org/10.1016/0304-3878(91)90068-7)
- Faini, R., Pritchett, L., & Clavijo, F. (1992). Import Demand in Developing Countries, in M.G. Dagenais and P.A. Muet (eds), *International Trade Modelling*, (New York: Chapman & Kegan Paul).
- Funke, M. (2001). Money demand in Euroland. *Journal of International Money and Finance*, 20, 701–713. [http://dx.doi.org/10.1016/S0261-5606\(01\)00017-1](http://dx.doi.org/10.1016/S0261-5606(01)00017-1)
- Gafar, J. (1995). Some Estimates of the Price and Income Elasticities of Import Demand for Three Caribbean Countries. *Applied Economics*, 27(11), 1045-1048. <http://dx.doi.org/10.1080/00036849500000087>
- Gujarati, D. N. (2002). *Basic Econometrics*, New York, McGraw Hill, New York.
- Hansen, B. E. (1992). Testing for Parameter Stability in Linear Models. *Journal of Policy Modeling*, 14, 517-533. [http://dx.doi.org/10.1016/0161-8938\(92\)90019-9](http://dx.doi.org/10.1016/0161-8938(92)90019-9)
- Hansen, P.R. (2005). Granger's Representation Theorem: A closed-form expression for I(1) processes. *Econometrics Journal*, 8, 23-39. <http://dx.doi.org/10.1111/j.1368-423X.2005.00149.x>
- Hendry, D. F., & Doornik, J.A. (1994, February). Modelling Linear Dynamic Econometric System. *Scottish Journal of Political Economy*, 41, 1-33. <http://dx.doi.org/10.1111/j.1467-9485.1994.tb01107.x>
- International Financial Statistics. (2009), IFS CD-ROM, International Monetary Fund, Washington D.C.
- Johansen, S. (November 1991). Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models. *Econometrica*, 1551-1580. <http://dx.doi.org/10.2307/2938278>
- Johansen, S. (1995). Likelihood –based Inference in Cointegrating Vector Autoregressive Models (New York: Oxford University Press).
- Kantu, M. (1972). An Econometric Model for India. *Indian Economic Review*, 7, 53-91
- Khan, M.S., & Knight, M. D. (1988). Import Compression and Export Performance in Developing Countries. *Review of Economics and Statistics*, 70(2), 315-21. <http://dx.doi.org/10.2307/1928316>
- Masih, R., & Masih, A. M. M. (2000). A Reassessment of Long-run Elasticities of Japanese Import Demand. *Journal of Policy Modelling*, 22, 625-639. [http://dx.doi.org/10.1016/S0161-8938\(98\)00014-3](http://dx.doi.org/10.1016/S0161-8938(98)00014-3)
- McAleer, M. (1987). Specification Test for Separate Models: A survey, in M.L. King and D.E.A. Giles (eds.), *Specification Analysis in the linear Model* (London: Routledge & Kegan Paul).
- Moran, C. (1989). Imports under a Foreign Exchange Constraint. *The World Bank Economic Review*, 3(2), 279-95. <http://dx.doi.org/10.1093/wber/3.2.279>
- Nyblom, J. (1989). Testing for the Constancy of Parameters over Time. *Journal of the American Statistical Association*, 84, 223-230. <http://dx.doi.org/10.2307/2289867>
- Pani, P. K. (1977). *Amacroeconomic model for the Indian economy*. New Delhi: MacMillan.
- Senhadji, A. (1998). Time-series Estimation of Structural Import Demand Equations: a Cross-Country Analysis. *International Monetary Fund Staff papers*, 45(2), 236-68.
- Sinha, D. (1997). Determinants of Import Demand in Thailand. *International Economic Journal*, 11(4), 73–83.
- Sultan, Zafar A., (2011). Foreign Exchange Reserves and India's Import demand: A Cointegration and Vector Error Correction Analysis. *International Journal of Business and Management*, 6(7), 69-76.
- Stock, J., & Watson, M. (1988, December). Testing for Common Trends. *Journal of the American Statistical Association*, 83, 1097-107. <http://dx.doi.org/10.2307/2290142>
- Tang, T.C. (2006). Aggregate Import Demand and Bank Credit in Southeast Asia: An Empirical Study. *International Journal of Economics and Management*, 1(1), 117-139.
- Thursby, J., & Thursby, M. (1984). How Reliable Are Simple, Single Equation Specifications of Import Demand? *Review of Economics and Statistics*, 66(1), 120-8. <http://dx.doi.org/10.2307/1924703>

Zarembka, P. (1974). Transformation of Variables in Econometrics, in P. Zarembka (ed.), *Frontiers in Economics* (New York: Academic Press).

Notes

Note 1. See also The World Factbook, CIA, especially <http://www.cia.gov/library/publications/the-world-factbook/rankorder/2188rank.html> and World Economic Outlook (various issues, especially 2003)

Note 2. This section draws on Arize and Nippani (2010).

Note 3. Alternatively, we could have tested for the appropriate functional form of the import demand using the following tests: (i) Box-Cox (B-C) procedure as discussed in Zarembka (1974), (ii) Bera and McAleer (BM) as discussed in McAleer (1987) and (iii) the MacKinnon, White and Davidson (MWD) procedure as discussed in Gujarati (2002:281)

Note 4. The results from Table 1 represent strong evidence that all four variables from equation (1) are not only non-stationary, but are likely to be integration of order one occurs if a variable in levels is found to be nonstationary, while the same variable in differences is found to be stationary.

Note 5. To be precise, we use the Sims likelihood ratio test (SLR), the Akaike Information Criterion (AIC), the Schwarz Bayesian criterion (SBC) and the Ljung-Box test.

Note 6. Following Johansen (1995), we also include an (unrestricted) constant term in the model.

Note 7. To provide complementary evidence and check the robustness of the Johansen results, we employ three other tests.

Note 8. Recent specification suggestions have shown that it is much 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 imports.

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...”

Note 10. See Hansen (2005) for further description.

Table 1. Augmented Dickey-Fuller Unit Root Tests

Country	Levels				Differences			
	m	y	p	IR	m	y	p	IR
India	-2.75	-2.16	-3.65	-2.72	-5.12	-6.97	-4.38	-4.38
Japan	-2.39	-1.57	-2.47	-2.43	-5.59	-3.59	-4.59	-4.47
Korea	-2.84	-0.90	-2.36	-2.02	-4.65	-4.52	-5.58	-5.81
Singapore	-2.19	-2.14	-1.63	-0.84	-4.33	-4.59	-3.96	-3.45
Thailand	-1.96	-1.35	-1.31	-2.48	-3.59	-3.58	-5.74	-4.93

Note: The critical value of the ADF statistic at the 5 per cent level is -3.41.

Table 2. Test for Cointegration among Import Demand, Domestic Income, Relative Prices and Foreign Exchange Reserves

Country	Johansen		Harris-Inder			Shin's Test			Hansen's Lc
	H0: $r=0$	H0: $r=0$	H0: Cointegration			H0: Cointegration			H0:Cointegration
	Ha: $r=1$	Ha: $r>=1$	Ha: No Cointegration			Ha: No Cointegration			Ha:No Cointegration
	λ -Max	Trace	3	6	8	3	6	8	Lc
India	46.01	65.18	0.11	0.08	0.07	0.13	0.09	0.08	0.55(0.11)
Japan	35.58	61.40	0.21	0.16	0.14	0.15	0.12	0.11	0.49(0.19)
Korea	77.58	92.84	0.28	0.2	0.17	0.14	0.11	0.10	0.52(0.16)
Singapore	51.75	79.73	0.18	0.12	0.1	0.12	0.09	0.08	0.33(0.18)
Thailand	44.17	52.96	0.09	0.06	0.06	0.09	0.07	0.06	0.45(0.20)

Notes: The critical values for $p-r = 4$ in the case of Johansen are 27.07 for λ_{max} and 47.21 for the Trace test. Since we test for one cointegration vector; hence p , the number of variables, is 4, and r , the number of cointegrating vector is zero. The critical values are from Osterwald-Lenum (1992, Table 1.1*). For the Harris-Inder test, the critical value is 0.32 at the 5 percent level. In the case of the Shin test, it is 0.159. The p -values for Hansen's test are in parentheses beside the Lc test statistic

Table 3. Long-run Elasticities and Hypothesis tests

Country	Phillips-Hansen Estimator (FMOLS)			Stock-Watson Estimator (DOLS)			Autoregressive Distributed Lag (ARDL)		
	y	p	IR	y	p	IR	y	p	IR
India	1.45 (43.48)	-0.39 (4.97)	0.09 (4.83)	1.44 (43.03)	-0.38 (4.70)	0.11 (6.43)	1.44 (28.70)	-0.46 (3.56)	0.10 (3.60)
Japan	0.74 (10.94)	-0.16 (3.95)	0.19 (11.27)	0.77 (14.53)	-0.14 (3.87)	0.19 (15.35)	0.78 (6.31)	-0.20 (2.38)	0.18 (6.12)
Korea	1.20 (26.58)	-0.45 (3.27)	0.05 (2.28)	1.02 (20.46)	-0.81 (3.35)	0.06 (1.62)	1.19 (16.62)	-0.43 (2.16)	0.06 (2.23)
Singapore	0.73 (7.42)	-0.27 (1.94)	0.22 (3.20)	0.84 (9.08)	-0.23 (1.99)	0.15 (2.31)	0.88 (12.06)	-0.23 (1.91)	0.12 (2.04)
Thailand	1.02 (15.74)	-0.24 (2.21)	0.21 (6.35)	1.06 (17.71)	-0.29 (3.40)	0.20 (6.38)	1.13 (6.47)	-0.70 (2.14)	0.19 (2.25)

Notes: The numbers in parentheses report absolute t-statistics. The critical value at 10 percent is 1.3 and 1.67 at 5 percent level.

Table 4. Comparison of long-run import elasticities.

Study	Data period	Country	Income	Price
Bahmani-Oskooee and Niroomand (1998)	73:80	Korea	1.83	-0.13
	73:80	Japan	0.46	-0.97
	73:80	Japan	1.17	-0.09
Bahmani-Oskooee (1998)	73:1-80:4	Korea	2.17	-1.37
	73:1-80:4	Korea	0.402	-1.91
	73:1-80:4	Singapore	1.26	-0.15
	73:1-80:4	Singapore	1.054	-0.3
Senhadji (1998)	60:93	India	1.33	-0.12
	60:93	Korea	1.32	-0.84
	60:93	Japan	1.04	-0.52
		Thailand	1.67	-0.43
Tang (2006)	74-02	Singapore	1.077	-0.463
		Thailand	0.917	-0.78
Sinha (2006)	53-90	Thailand	2.147	-0.768
Sinha (2010)	51-96	Japan	0.841	-0.908
	53-95	Thailand	0.902	-0.199
	50-96	India	-0.1134	-0.5142

Table 5. Short-run Dynamics: Error Correction Model

India										
$\Delta m_t = 0.024 - 0.296EC_{t-1} - 0.213\Delta m_{t-1} - 0.131\Delta m_{t-2} + 0.137\Delta m_{t-4} + 0.401\Delta^2 y_{t-1} + 0.346\Delta^2 y_{t-2} + 0.360\Delta^2 y_{t-3} - 0.670\Delta p_t - 0.289\Delta p_{t-1} + 0.048\Delta^2 IR_{t-3}$										
$R^2 = 0.667, \bar{R}^2 = 0.636, F(10,109) = 21.79, DW = 2.07, BG - F = 1.12(0.35), RESET - F = 0.004(0.95)$										
Japan										
$\Delta m_t = 0.043 - 0.186EC_{t-1} + 0.126\Delta^2 y_t - 0.121\Delta p_t - 0.154\Delta p_{t-4} - 0.032\Delta^2 IR_t + 0.046\Delta IR_t - 0.057S_1 - 0.031S_2 - 0.041S_3$										
$R^2 = 0.595, \bar{R}^2 = 0.563, F(9,115) = 18.74, DW = 1.98, BG - F = 0.30(0.87), RESET - F = 5.10(0.03)$										
Korea										
$\Delta m_t = 0.011 - 0.320EC_{t-1} - 0.178\Delta m_{t-1} + 0.482\Delta y_t + 0.254\Delta y_{t-2} + 0.257\Delta y_{t-3} - 0.226\Delta p_t - 0.277\Delta p_{t-2} - 0.085\Delta IR_{t-3} + 0.095\Delta IR_{t-4}$										
$R^2 = 0.645, \bar{R}^2 = 0.617, F(9,115) = 23.22, DW = 1.94, BG - F = 1.16(0.34), RESET - F = 1.73(0.19)$										
Singapore										
$\Delta m_t = -0.002 - 0.128EC_{t-1} + 0.895\Delta y_t - 0.243\Delta y_{t-2} + 0.501\Delta p_{t-4} + 0.159\Delta IR_{t-3}$										
$R^2 = 0.459, \bar{R}^2 = 0.436, F(5,115) = 19.52, DW = 2.09, BG - F = 0.62(0.65), RESET - F = 0.09(0.77)$										
Thailand										
$\Delta m_t = -0.005 - 0.209EC_{t-1} + 0.153\Delta m_{t-4} - 0.623\Delta y_{t-1} + 0.775\Delta y_{t-3} - 0.477\Delta p_t - 0.182\Delta p_{t-1} + 0.128\Delta IR_{t-4}$										
$R^2 = 0.408, \bar{R}^2 = 0.372, F(7,118) = 11.60, DW = 1.98, BG - F = 0.17(0.95), RESET - F = 2.45(0.12)$										

Notes: 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, RESET=Ramsey RESET normality test.
 Variable names: EC=error correction term, m=desired real imports, y=real domestic income, p=domestic wholesale price, IR=real foreign exchange reserves.

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

Countries	Speed of adjustments		Response of desired real imports to each regressor		
	Desired real imports	Half-life adjustment	Mean time lag		
			m	ψ	y
India	-0.296 (0.08)	1.97	0.34	5.05	3.92
Japan	-0.186 (0.05)	3.38	4.71	6.22	5.56
Korea	-0.320 (0.07)	1.80	2.08	4.54	3.65
Singapore	-0.128 (0.05)	5.08	5.94	3.92	6.59
Thailand	-0.209 (0.05)	2.96	2.64	4.93	3.44
Average		3.04	3.14	4.93	4.63

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

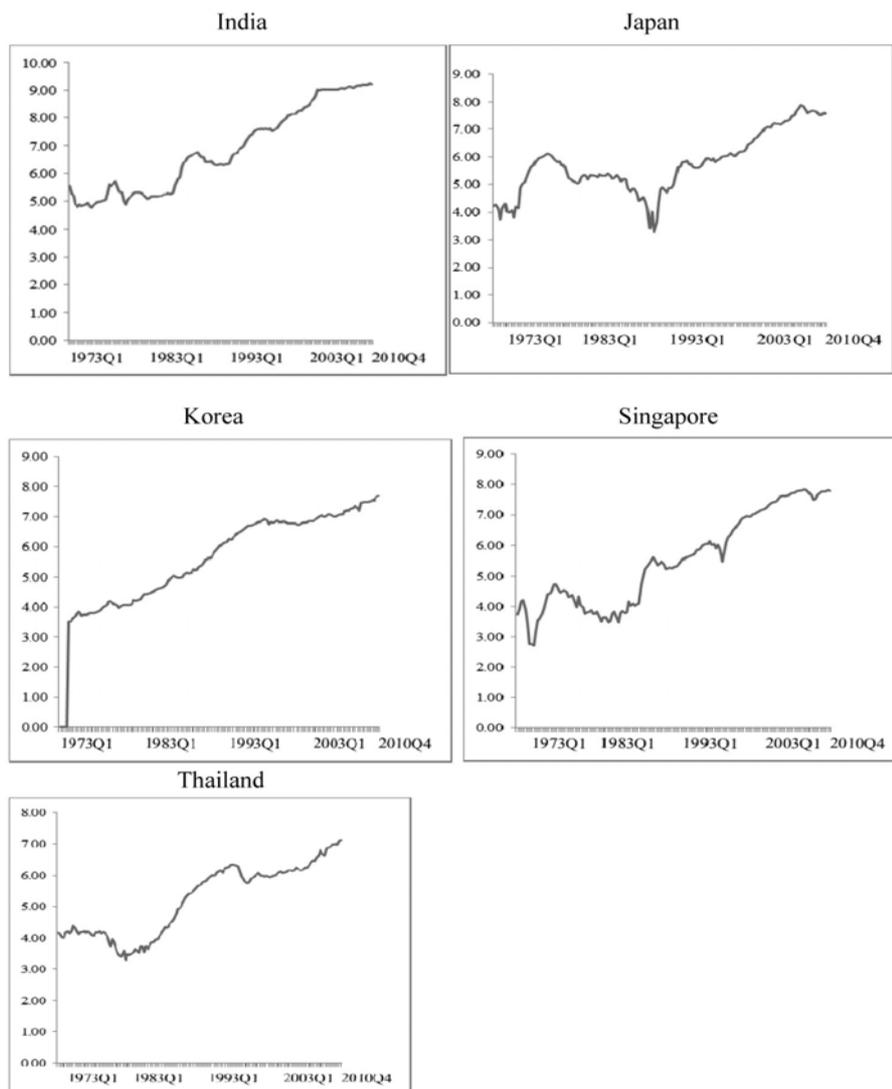


Figure 1. Plots of Real Foreign Exchange Reserves