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COINTEGRATION ANALYSIS ON JAPAN'S AGGREGATE IMPORT DEMAND FUNCTION: DOES DATA FREQUENCY MATTER?

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Abstract

This study aims to empirically re-examine Japan's long-run aggregate import demand function using a variety of cointegration tests. The primary contribution of this study is to compare estimates obtained from samples of quarterly, biannual, and annual data for the period 1973 to 2000. The results of bounds test and Johansen's multivariate test show that the quantity of imports, real incomes, and relative import prices are consistently cointegrated regardless of data frequency. In contrast, the Engle-Granger's residual-based and error correction mechanism tests reveal no cointegrating relationship in Japan's aggregate import equation. This study thus concludes that data frequency does not affect estimates of Japan's aggregate import demand function, but that the choice of cointegration techniques does.

JEL Classifications: C50; F10

Keywords: Import Demand; Cointegration; Data Frequency; Japan.

1. Introduction

Studies of Japan's import demand provide an important example of how the policy implications of empirical analysis can be closely related to the design of the empirical tests conducted. For example, Hamori and Matsubayashi (2001, pp. 135-136) criticise that Japan is running a trade surplus and recommend that Japan should reduce its trade surplus by stimulating domestic demand.[1] However, if stimulating domestic demand is to increase imports and reduce the trade surplus, Japan's import demand function must be stable. Correspondingly, it is important to see if the evidence suggests that Japan's import demand function is indeed stable. Hamori and Matsubayashi (2001)

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[1] Japan experienced trade deficits only in the period 1974, and 1979-1980 over the period 1973-1999 (*World Tables*, World Bank, various issues). Another study is from Tang (2004).

among others use the concept of cointegration to test the stability of Japan's import demand function. The presence of cointegration, or a stationary linear combination of non-stationary variables, between real imports and its determinants (real income and relative import prices) implies that the import demand function stable because cointegrated variables will never move too far apart, and will be attracted to their long-run relationship.

Of recent, numerous empirical studies have examined Japan's long-run, aggregate import demand function using cointegration techniques (Table 1). These studies with the exception of Tang's (2003c and 2004), use the traditional specification of import demand as a function of real income (proxied as real gross domestic product, GDP, or real gross national product, GNP) and relative import prices (the ratio of import prices to domestic prices). For example, using biannual data for 1974 to 1990 and estimation techniques developed by Engle and Granger's (1987), Dickey-Fuller (1979), and Phillips (1987), Mah (1994) failed to reject the null hypothesis of no cointegration among real imports, real income, and relative import prices. Using annual data for 1960 to 1992 and Johansen's multivariate technique (Johansen, 1991; 1995), however, Bahmani-Oskooee and Niroomand (1998) provide evidence of cointegration and thus stable long-run relationships in Japan's aggregate import demand function. Their estimates of long run income and price elasticities are of the expected sign but rather small in absolute value suggesting inelastic import demand with respect to income and relative import prices. Using similar econometric techniques, Masih and Masih (2000) also provide evidence of a cointegrating relationship in a sample of biannual data for 1974 to 1990. Saikonnen (1991) and Stock and Watson (1993) present dynamic ordinary least squares (DOLS) estimates suggesting that income and price elasticities are rather large in absolute value, 1.28 and -1.89, respectively.

Recently, Hamori and Matsubayashi (2001) reassessed the stability of Japan's aggregate import demand behaviour using quarterly data for 1973 to 1998 and Johansen's multivariate tests with different lag lengths, revealing mixed results. According to Toda (1995), the results of the Johansen test clearly depend on the lag length of vectors error correction model. In this relation, their conclusion is merely based on Engle and Granger's (1987) approach which indicated no cointegration among real imports, real income and relative prices to imports. Among others, Tang (2003a) reexamined aggregate imports demand for Japan and discovered a bounds testing procedure developed by Pesaran *et al*. (2001) and applying this bounds testing procedure using annual data for 1973 to 1997. He also cross-checked the results based on other cointegration techniques such as the Engle and Granger technique, DOLS, Error-Correction Mechanism (ECM) estimation, and Johansen's multivariate tests. In effect, a cointegrating relation among volume of imports, real income and relative import prices was suggested by results from the bounds test, DOLS, and the Johansen techniques. On the other hand, no cointegration was detected using the Engle and Granger, and ECM techniques. Estimated income and price elasticities are also varied depending on the technique used. Based on the results presented by the bounds test, Tang (2003a) concluded that there exists a long run relationship in Japan's imports demand. Here, the unrestricted error correction model yielded long-run income and price elasticities of 0.99 and -0.82 respectively. By disaggregating the components of final expenditure, and using bounds test and annual data (1973-1997), Tang (2003c) found that there is cointegration among Japan's imports, real final consumption expenditure, real domestic investment, real expenditure on export goods and relative import prices. According to another finding by Tang (2003b) using Engle and Granger approach, Johansen's multivariate test, and bounds test, this study showed no long run relationship among Japan's imports, national cash flow (see Xu, 2002) and relative import prices. The sample covers quarterly data from 1973 to 2000. Tang (2004) critiqued the bias of variable(s) omission occurs if estimating conventional import demand function, and extended the study (Tang, 2003b) by adding financial variable such as bank credit, lending rate, deposit rate, government bond yield, and share prices into Japan's import demand function. This is interesting work, and the study provides evidence of cointegration and thus stable long-run relationships in Japan's aggregate import demand function.

From the above review, it is critical in this case of Japan's import demand analyses that data frequency and different cointegration techniques yield different estimates. A matter of greater concern is the issue of data frequency in cointegration test. Cheung and Lai (1993) documented that finite sample analyses could cast some doubt on the cointegration results of likelihood ratio test (Johansen approach). Using Monte Carlo experiment, Toda (1994; 1995) revealed a minimum sample size for cointegration is around 100. Additionally, Mah (2000) argues that the conventional tests for cointegration such as Engle and Granger approach, and Johansen's multivariate test are not reliable in the case when the data set is of a small sample size. For small sample studies, Mah recommends bounds testing procedure (Pesaran *et al*., 2001) for estimating long-run relationships. Additionally, Hakkio and Rush (1991) provided empirical evidence that increasing the number of observations by using monthly or quarterly data does not add any robustness to the cointegration results, but that the span of data does i.e., the number of years the sample covers. Accordingly, Charemza and Deadman (1992. p.153) suggested, "Annual data could be used to estimate these long run parameters thereby avoiding the need to model the seasonality, and the standard tests for cointegration applied". A book by Davidson and MacKinnon (1993, p.714) documented that one possibility to avoid using seasonally adjusted data to compute unit root tests is to use annual data with sufficient sample span. Conventionally, unit root tests such as Dickey-Fuller test and augmented Dickey-Fuller test (on the residual series of a cointegrating regression) are the essence of Engle-Granger (1987) approach.

This study is motivated by the need for an in-depth empirical investigation on the long-run estimates of Japan's aggregate imports demand behaviour relating to the necessity of data frequency.[2] In the literature to date, the long run estimates of Japan's aggregate import demand behaviour obtained from samples of quarterly, biannual, and annual data are indeed inconclusive. Using various frequencies of data (quarterly, biannual, and annual data), this study has estimated Japan's aggregate import demand behaviour through cointegration approach. A bounds test and Johansen's multivariate test were initially employed in this study and, the results have been cross-checked with other methods such as Engle and Granger (1987) approach and error correction mechanism test (Banerjee *et al*., 1998). Using the dynamic OLS approach, the long run elasticities were estimated in this study. Empirically speaking, the study on import demand is deemed to provide signals about likely directions of trade policy as well as economic policy.

[2] If a GDP related measure of income is to be used, the essential choice is annual data or quarterly data.

Table 1
A Summary of Selected Empirical Studies on Japan's Aggregate Import Demand

Previous Studies	Method (sample period)	Are $\ln M_t$, $\ln Y_t$, and $\ln RP_t$ cointegrated ?	Long run Elasticities	
			$\ln Y_t$	$\ln RP_t$
Mah (1994)	Engle-Granger (1987) (1974:1-1990:2, biannual data)	No	0.781	-0.456
Bahmani-Oskooee & Niroomand (1998)	Johansen (1960-1992, annual data)	Yes	0.46	-0.97
Masih & Masih (2000)	Johansen (1974:1-1990:2, biannual data)	Yes	1.277 (Dynamic OLS)	-1.892 (Dynamic OLS)
Hamori & Mastsubayashi (2001)	1. Engle-Granger (1987) 2. Johansen – VAR(4) 3. Johansen –VAR(8) 4. Johansen – AIC & SC (1) 5. Gregory-Hansen (1996) (1973:1-1998:1, quarterly data)	No No Yes Yes No	No estimation	No estimation
Tang (2003a)	1. Bounds test 2. Engle-Granger (1987) 3. DOLS 4. ECM 5. Johansen (1973-1997, annual data)	Yes No Yes No Yes	0.985 1.05 1.25	-0.823 -0.204 -0.79
Tang (2003b)	1. Engle-Granger (1987) 2. Johansen 3. Bounds test (1973-2000, quarterly data)	No No No	Y_t is 'national cash flow' (see Xu 2002)	
Tang (2003c)	1. Bounds test (1973-1997, annual data)	Yes	Y_t is disaggregated final expenditure viz. final consumption, investment, and expenditure on export goods	-0.866
Tang (2004)	1. Engle-Granger (1987) 2. Johansen 3. Bounds test (1973-2000, quarterly data)	No Yes Yes (financial variable has been included)	Y_t is 'national cash flow' Between -1.1 and -1.3 (Autoregressive distributed lag)	Between -0.7 and -1.5

This study proceeds as follows. Section 2 gives a snapshot and review on import demand function, data, and cointegration tests. Section 3 presents a discussion on the results. Lastly, the final section provides concluding remarks on cointegration analysis on Japan's aggregate import demand.

2. Model, Data and Cointegration Techniques

The existing literature has empirically approached standard formulation of import demand equation that relating the quantity of import demanded to domestic real income and relative price of imports.[3] This specification of imports demand corresponds to that of the imperfect substitute model (Goldstein and Khan, 1985), which implies the existence of imports and domestic production as well as intra-industry trade. In this relation, Hong (1999, p. 3) has documented that "... *import demand in a market economy can be fully modelled by two determinants: income and relative prices. The other factors can all be subsumed within these two factors, at least theoretically*".[4] By assuming zero degree homogeneity, and of the supply elasticity is infinite or at least large, the following single equation of imports demand can be consistently estimated:

$$M_t = f(Y_t, RP_t) \tag{1}$$

where at period t, M_t is the desired quantity of imports demanded, Y_t is the real income (domestic real activity). RP_t is the relative price of imports that is the ratio of import price to domestic price level.

The popularity of 'traditional' imports demand model [(Equation (1)] has widely been cited from the empirical literature on import demand analyses (Mah 1994; Masih and Masih 2000; Hamori and Matsubayashi 2001; and Tang 2003a). And the double-log linear form of data-driven import demand regression is given by:

$$\ln M_t = a_0 + a_1 \ln Y_t + a_2 \ln RP_t + e_t \tag{2}$$

where at period t, M_t is quantity of imports, Y_t is real Gross Domestic Product (GDP), and RP_t is the ratio of import price to domestic price level (proxied by GDP deflator). ln is natural logarithmic. The relevant economic theory, says the signs of the parameters are expected to be $a_1 > 0$, and $a_2 < 0$ (both are expressed as income elasticity and price elasticity).

Considering the flexible exchange rate regime started in 1973, the sample period used in this study is 1973-2000 (see Hamori and Matsubayashi, 2001).[5] The quarterly data

[3] Typically, the relative import prices is the ratio of import price to domestic price level which is often used to ease estimation problems i.e. reduce multicollinearity (Houthakker and Magee, 1969).

[4] As highlighted by Hong (1999, p.3), the factors behind relative prices include: relative endowments of resources and productive factors, taste, market structure, scale, exchange rate, trade barriers etc. The impacts of changes in these factors on import demand will take place through a change in relative prices.

[5] The sample period is started from 1973 is due to the consideration of the period of the flexible exchange rate regime as in Hamori and Matsubayashi (2001, p.136).

is available from OECD Main Economic Indicators, while the import price index is from *International Financial Statistics* (International Monetary Fund). The data source for biannual and annual observations is OECD Economic Outlook. All variables are on domestic currency, and then converted into index form with 1995 prices.

The common practice for cointegration is first to test the stationality of the time series variables [or its degree of integration, I(d)]. Initially, unit root test developed by Phillips and Perron (1988) was applied on the variables used in this study. Conceptual framework and statistical property of Phillips and Perron (1988) unit root test (PP test) are not detailed here since the test is widely used in many empirical works. The PP test is robust for the presence of autocorrelation and heteroscedasticity under the null hypothesis of a unit root against the alternative of stationary (trend or level stationary (Phillips and Perron, 1988). Table 2 illustrates the results of PP test, and all the variables are statistically traced as nonstationary or I(1) even with quarterly, biannual, and annual data.

Table 2
Phillips-Perron (1988) Unit Root Tests

Variable	**Quarterly Data (1973:1 to 2000:4)**	**Biannual Data (1973:1 to 2000:2)**	**Annual Data (1973 to 2000)**
In Levels[1]:			
$\ln M_t$	-2.021 (4)	-2.147 (3)	-2.243 (3)
$\ln Y_t$	0.257 (4)	-0.007 (3)	-0.268 (3)
$\ln RP_t$	-2.889 (4)	-2.871(3)	-2.881 (3)
Critical Values: (MacKinnon, 1991)			
1%	-4.043	-4.131	-4.338
5%	-3.450	-3.492	-3.587
10%	-3.150	-3.174	-3.228
In First Differences[2]:			
$\Delta \ln M_t$	-8.625 (4)*	-3.944 (3)*	-3.739 (3)*
$\Delta \ln Y_t$	-10.993 (4)*	-4.704 (3)*	-3.162 (3)**
$\Delta \ln RP_t$	-6.155 (4)*	-4.029 (3)*	-4.567 (3)*
Critical Values: (MacKinnon, 1991)			
1%	-3.491	-3.555	-3.708
5%	-2.888	-2.916	-2.980
10%	-2.581	-2.595	-2.629

Notes: (.) is the lag truncation for Bartlett Kernel based on Newey-West's suggestion. Δ is the first difference operation. ln denotes natural logarithm.
[1] The null of a unit root is tested against the alternative of trend stationary. The constant and time trend variables are included into test equation.
[2] The null of a unit root is tested against the alternative of differenced stationary. Only constant is included into test equation.
*significant at 1% level. ** significant at 5% level. *** significant at 10% level.

Standard unit root tests such as PP test are based on the null hypothesis of a unit root and, they are thus not very powerful against relevant alternatives (Kwiatkowski *et al*., 1992, p. 160). Unlike standard unit root tests, Kwiatkowski *et al*. (1992) proposed a modified version of LM (Lagrange Multiplier) statistics for testing the null hypothesis of stationarity (trend or level stationary) against the alternative of a unit root - the test is abbreviated as KPSS stationarity test. Its statistical properties are available from Kwiatkowski *et al*. (1992). The results of KPSS stationarity tests are illustrated as in Table 3.

Table 3
KPSS Stationary Tests (Kwiatkowski *et al*., 1992)

Variable	Quarterly Data (1973:1 to 2000:4)	Bi-annual Data (1973:1 to 2000:2)	Annual Data (1973 to 2000)
In Levels[1]:			
$\ln M_t$	0.192 (8)*	0.156 (5)**	0.138 (3)
$\ln Y_t$	0.238 (9)*	0.198 (5)*	0.1477 (4)
$\ln RP_t$	0.129 (8)	0.122 (5)	0.107 (3)
Critical Values:			
1%	0.160	0.160	0.200
5%	0.145	0.145	0.171
10%	0.136	0.136	0.160
	(T=50)	(T=50)	(T=30)
In First Differences[2]:			
$\Delta \ln M_t$	0.098 (6)	0.122 (1)	0.156 (0)
$\Delta \ln Y_t$	0.400 (7)**	0.337 (4)	0.301 (2)
$\Delta \ln RP_t$	0.119 (5)	0.244 (0)	0.172 (0)
Critical Values:			
1%	0.476	0.476	0.440
5%	0.396	0.396	0.385
10%	0.344	0.344	0.347
	(T=50)	(T=50)	(T=30)

Notes: Δ is the first difference operation. ln denotes natural logarithm. (.) is the Bandwidth (Newey-West using Bartlett kernel). The test statistics were computed by Eviews 4.1 version and $l12=12[T/100)^{1/4}]$ is defaulted for selecting maximum lag length. The reported critical values are from Hornok and Larsson (2000, p.115-116). Table 4 (level-stationary case) and Table 5 (trend-stationary case) with unknown error variance and $l12$ (T is the sample size).

[1] The null of trend-stationary is tested against the alternative of a unit root that is a constant and a time trend variables are included into test equation.

[2] The null of level-stationary is tested against the alternative of a unit root that is only constant is included into test equation.

*significant at 1% level. **significant at 5% level, and *** significant at 10% level.

The results of KPSS stationary test show the relative import prices ($\ln RP_t$) trend-stationary (at 5 percent significance level) both quarterly and biannual data. For annual data, the null hypothesis of trend stationary cannot be rejected, but the tests do reject the null of level stationary. These results support the alternative hypothesis of a unit root for all annual time series data (at 1 percent significance level).[6] Next, the results of KPSS tests do not reject the null hypothesis of first difference stationary for all variables (at 1 percent significance level) where in first differences. Visual inspection from the time series plots cited in Appendix 1 that real imports and real income are likely nonstationary with upward trend, and relative price of imports seem to be stationary with a fairly downward trend. These time series plots (patterns) are visually consistent with those printed in Hamori and Matsubayashi (2001) (Figures 1, 2 and 3, p. 137-193). Hamori and Matsubayashi (2001) used augmented Dickey-Fuller test (Dickey and Fuller, 1979; 1981), and found quarterly series of real imports, real GDP and relative import prices nonstationary (a unit root). The results of KPSS, however reveal that Japan's imports demand is probably determined by both stationary ($I(0)$) and nonstatioary ($I(1)$) explanatory variables such as relative price of imports and real income. As a consequence, cointegration technique which allow $I(0)$ and $I(1)$ endogenous variables is appropriate in order to derive robust results.

A variety of tests for cointegration exist. Fortunately, econometrics literature provides techniques for cointegration with stationary ($I(0)$) and nonstationary ($I(1)$) explanatory variables such as bounds testing procedure (Pesaran *et al*., 2001) and Johansen's multivariate tests (Johansen, 1991; 1995). The following discussions give brief review on these techniques.

Bounds Testing Procedure (Pesaran et al., 2001)

Pesaran *et al*., (2001) proposed bounds testing approach which can be applied for cointegration analysis irrespective of whether the regressors are purely $I(0)$, or purely $I(1)$, or mutually cointegrated (Pesaran *et al*., 2001). Practical researchers such as Pattichis (1999) and Mah (2000) noted that a bounds test for cointegration (from unpublished version of Pesaran *et al*., 2001) is appropriate for small sample study. Additionally, bounds test is also applicable when the explanatory variables are endogenous (see Alam and Quazi, 2003, p.93).

Practically, bounds test is based on unrestricted error-correction model (UECM) estimates using Ordinary Least Squares (OLS) estimator where UECM is a simple reparameterization of a general autoregressive distributed lag (ARDL) model. The UECM form of import demand function as equation (2) can be written as follow:

$$\begin{aligned}\Delta \ln M_t = b_0 + \sum_{i=0}^{l} b_{1i} \Delta \ln Y_{t-i} + \sum_{i=0}^{l} b_{2i} \Delta \ln RP_{t-i} + \sum_{i=1}^{l} b_{3i} \Delta \ln M_{t-i} \\ + b_4 \ln M_{t-1} + b_5 \ln Y_{t-1} + b_6 \ln RP_{t-1} + e_t,\end{aligned} \tag{3}$$

[6] The LM statistics (KPSS test) for the null hypothesis of level stationary (only intercept is included in the test equation) are 0.6384, 0.656, and 0.523 for $\ln M_t$, $\ln Y_t$ and $\ln RP_t$ respectively. The Hornok and Larsson's (2000, p.115, Table 4, T=30, l12) critical values are 0.44 (1%), 0.385(5%) and 0.347(10%), thus the LM statistics do reject the null hypothesis of level stationary at 1% level, and the results indicate a unit root.

where $\Delta \ln M_t$, $\Delta \ln Y_t$ and $\Delta \ln RP_t$ are first differences of the logarithms of quantity import demand, real domestic income, and relative price, respectively. l is lag structure.

The test statistic of bounds test (F-statistic) can be derived from a restricted form of error correction model - excluding the lagged level variables $\ln M_{t-1}$, $\ln Y_{t-1}$ and $\ln RP_{t-1}$ from UECM. More formally, this is to perform a joint significance test (Wald test) for testing the null hypothesis of no cointegrating relationship ($H_0 : b_4 = b_5 = b_6 = 0$) against the alternative hypothesis of a cointegrating relationship ($H_A : b_4 \neq 0, b_5 \neq 0, b_6 \neq 0$). The decision rule is that if the computed test statistic (F-statistic) lies above the upper bound, then the null hypothesis can be rejected at a conventional level of significant (let say 1, 5 or 10%). Thus, a cointegrating relationship among the variables can be concluded. In the case when the computed F-statistic lies below the lower bound, the null hypothesis cannot be rejected - no cointegration relationships. Conclusive inference cannot be made when the test statistic falls within the lower and upper bounds. In this relation, the time series properties must be known before any conclusion can be drawn (Pesaran *et al*., 2001, p. 290). The long run elasticities of income and relative import prices are $-(b_5/b_4)$ and $-(b_6/b_4)$, respectively (Pesaran *et al*., 2001, p. 294).

Meanwhile, Pesaran *et al*. (2001, p.315) acknowledged that their single-equation approach is inappropriate in situations where there may be more than one level of relationships involving Y_t such as in this study.[7] Consequently, Johansen's multivariate approach (Johansen, 1991; 1995) is employed for crosscheck where it allows testing for the number of cointegrating relationships. The latter approach is more powerful when there are more than two variables in any reduced form model.

Johansen's Multivariate Tests (Johansen, 1991; 1995)

Johansen's multivariate test is a system approach developed by Johansen (1991; 1995) which can be used with a set of variables containing possibly a mixture of $I(0)$ and $I(1)$ regressors (Masih and Masih, 2000, p. 626; Pesaran *et al*., 2001 p. 315, footnote 39). The Johansen's multivariate test is to test the restrictions imposed by cointegration on the unrestricted VAR of the series. The likelihood ratio statistic is computed to test for the number of cointegrating vectors for a set of variables in the cointegrating system. The cointegrating vectors of the system are as $\beta' X_t = z_t$, where β is the cointegrating matrix. For m jointly determined variables, it will be of the dimension $m \times m$, but of the rank $r \leq m - 1$, where r is the number of linear independent cointegrating vectors. In specification form, the model can be written as:

$$\Delta X_t = \delta + \sum_{i=1}^{k-1} \prod_i \Delta X_{t-i} + \prod_i X_{t-k} + \varepsilon_t \quad (4)$$

[7] It must be a maximum of two cointegrating vectors for three variables case (imports, real income and relative price of imports). As dictated by economic theory, the possible cointegrating vectors are import demand equation as in this study, and price equation which as a function of imports and real income.

where Δ is a difference operator, k is the lag length, δ is a constant, $\varepsilon_t \sim niid$ and X_t is a column vector of the involved variables. If Π has zero rank, no stationary linear combination can be identified, thus, no cointegrating relation among variables in X_t. If the rank r of Π is greater than zero, there exists r possible stationary linear combinations, and Π may be decomposed into two matrices α and β. The general hypothesis of the r cointegrating vector can be formulated as $H_0: \Pi=\alpha\beta'$ where α and β are both $(m \times r)$ matrices of full rank, with β containing the r cointegrating relationships and α carrying the corresponding loadings in each of the r cointegrating vectors. The likelihood ratio test - trace statistic is used to test the null $H_0: r \leq r_0$ versus the alternative $H_A: r > r_0$ (r is the number of cointegrating vectors).[8] The trace statistic $\lambda_{Trace}(r_0)$ is calculated as $-T\sum_{i=t+1}^{p}\ln(1-\hat{\lambda}_i)$ where $\hat{\lambda}$ is the estimated eigenvalues from Π, and T is the sample size. To adjust the test with the sample size used, the test statistics (trace statistics) were adjusted by a scaling factor, $(T-nk)/T$ (where T is the sample size, n is the number of variables in the estimated system, and k is the lag of parameter), as proposed in Reinsel and Ahn (1988). According to Arize and Shwiff (1998, p.1274), "*The most delicate part of the Johansen procedure is that the estimates of one equation may be sensitive to possible misspecification in another equation*". As a consequent, this study does not estimate the cointegrarting vectors using Johansen approach. Rather, following Masih and Masih (2000), the Dynamic OLS (Saikonnen, 1991; Stock and Watson, 1993) estimator was used to estimate the Japan's aggregate import demand in the long run (see Appendix 3).

For cross-checking purpose, other cointegration techniques such as Engle and Granger (1987) approach, and error-correction mechanism test (Banerjee *et al*., 1998) were also performed for crosscheck purpose (Appendixes 2 and 4).

3. The Results

Bounds Testing Approach

The unrestricted error-correction model (UECM) was initially estimated with lag length of 12 for quarterly data, 6 for biannual data, and 3 for annual data. This study used 'general to specific' methodology in order to arrive at a parsimonious specification i.e., the first differenced lagged regressors which have relatively small absolute *t*-value (less than one) were dropped sequentially. To account for the sensitivity of the estimates from different lag structure used, different lag lengths (with a gap of a year) were performed, and an optimum lag was determined using Akaike information criterion (AIC). Tables 4, 5 and 6 report the estimates of UECM in parsimonious form associated with quarterly, biannual and annual data, respectively.

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[8] Ghirmay *et al*. (1999, p. 220) noted the trace test tends to have more power than the λ-max test since it takes account of all $m - r$ of the smallest eigenvalues. Hence, in the conflicting cases between trace and λ-max tests, the decision is made based on the trace statistic.

Table 4
Estimated Specific UECM for Quarterly Data (1973:1 to 2000:4)

Regressor	$l = 12$	$l = 8$	$l = 4^{\#}$
$\Delta \ln Y_t$	0.710***	0.487	0.436
$\Delta \ln Y_{t-1}$	0.865**	0.377	0.385
$\Delta \ln Y_{t-2}$	0.776**		0.509***
$\Delta \ln Y_{t-3}$	0.434		
$\Delta \ln Y_{t-4}$	0.621***	0.424	0.662**
$\Delta \ln Y_{t-8}$		-0.46***	
$\Delta \ln Y_{t-10}$	-0.46627		
$\Delta \ln Y_{t-11}$	-0.29		
$\Delta \ln RP_{t-1}$		-0.069	
$\Delta \ln RP_{t-2}$	0.146**	0.145**	0.054
$\Delta \ln RP_{t-3}$	0.0681		
$\Delta \ln RP_{t-4}$			-0.052
$\Delta \ln RP_{t-5}$	0.101	0.05	
$\Delta \ln RP_{t-7}$	0.075	0.05	
$\Delta \ln RP_{t-9}$	0.049		
$\Delta \ln RP_{t-10}$	0.096***		
$\Delta \ln RP_{t-11}$	-0.066		
$\Delta \ln M_{t-2}$	0.136	0.217**	0.194**
$\Delta \ln M_{t-3}$		0.2**	
$\Delta \ln M_{t-4}$	-0.267**		-0.216**
$\Delta \ln M_{t-5}$	-0.178***	-0.221**	
$\Delta \ln M_{t-6}$		-0.153	
$\Delta \ln M_{t-7}$		0.099	
$\Delta \ln M_{t-8}$	-0.201**		
$\Delta \ln M_{t-12}$	-0.193**		
Constant	0.748*	0.572*	0.455*
$\ln M_{t-1}$	*-0.093* [-1.00]*	*-0.075* [-1.00]*	*-0.043*** [-1.00]*
$\ln Y_{t-1}$	*0.041 [0.44]*	*0.027 [0.36]*	*0.006 [0.13]*
$\ln RP_{t-1}$	*-0.104* [-1.12]*	*-0.072* [-0.96]*	*-0.059* [-1.39]*
R-squared	0.545	0.448	0.417
F-statistic (p-value)	4.132 (0.000)	4.365 (0.000)	6.176 (0.000)
LM test [2] (p-value)	2.655 (0.265)	0.818 (0.664)	0.387 (0.823)
Bounds Test – F Statistics for the null of no cointegrating relation[1]			
Test statistics	9.565	6.284	6.170
Critical values	Lower bound	Upper bound	
1%	5.15	6.36	
5%	3.79	4.85	
10%	3.17	4.14	

Notes: Dependent Variable: $\ln M_t$. l is the lag length for general ECM. $^{\#}$ denotes optimum augmented lag was selected based on Akaike information criterion (AIC). *[.]* is the normalised long run elasticity. LM test is Breusch-Godfrey serial correlation test for testing the null of the errors are serially uncorrelated. *p<0.01. **p<0.05, and ***p<0.1 (two tailed t-test).

[1] Source: Pesaran *et al*. (2001), p.300, Table CI(iii) case III).

Table 5
Estimated Specific UECM for Bi-Annual Data (1973:1 to 2000:2)

Regressor:	$l = 6$ #	$l = 4$	$l = 2$
$\Delta \ln Y_t$	2.152*	1.21**	1.19**
$\Delta \ln Y_{t-1}$	2.188*	1.355**	0.44
$\Delta \ln Y_{t-2}$	-0.999***		
$\Delta \ln Y_{t-3}$		-0.511	
$\Delta \ln Y_{t-5}$	-0.938*		
$\Delta \ln Y_{t-6}$	-0.377		
$\Delta \ln RP_{t-2}$	0.173**	0.173**	
$\Delta \ln RP_{t-3}$	-0.068		
$\Delta \ln RP_{t-4}$	0.109***	0.092	
$\Delta \ln RP_{t-1}$	-0.097**		
$\Delta \ln M_{t-2}$		0.319***	0.416*
$\Delta \ln M_{t-2}$	-0.359**	-0.296**	-0.36*
$\Delta \ln M_{t-4}$	-0.417*	-0.202	
$\Delta \ln M_{t-5}$	0.371**		
$\Delta \ln M_{t-6}$	-0.222**		
Constant	0.813*	0.875*	0.525*
$\ln M_{t-1}$	*-0.114** [-1.00]*	*-0.157* [-1.00]*	*-0.093** [-1.00]*
$\ln Y_{t-1}$	*0.087*** [0.77]*	*0.127** [0.805]*	*0.0787 [0.843]*
$\ln RP_{t-1}$	*-0.142* [-1.2]*	*-0.154* [-0.979]*	*-0.096* [-1.032]*
R-squared	0.842	0.7001	0.656
F-statistic (p-value)	10.658 (0.000)	8.278 (0.000)	12.24 (0.000)
LM test [2] (p-value)	3.785 (0.151)	0.449 (0.799)	3.212 (0.201)
Bounds Test – F Statistics for the null of no cointegrating relation			
Test statistics	8.287	7.120	7.486
Critical values	Lower bound	Upper bound	
1%	5.15	6.36	
5%	3.79	4.85	
10%	3.17	4.14	

Note: Refer to Footnote to Table 4.

Table 6
Estimated Specific UECM for Annual Data (1973 to 2000)

Regressor	l=1	
$\Delta \ln Y_t$	3.886*	
$\Delta \ln M_{t-1}$	-0.454*	
Constant	0.692	
$\ln M_{t-1}$	-0.148 *[-1.00]*	
$\ln Y_{t-1}$	0.206** *[1.39]*	
$\ln RP_{t-1}$	-0.205* *[-1.38]*	
R-squared	0.795	
F-statistic (p-value)	15.558 (0.000)	
LM test [2] (p-value)	1.745 (0.418)	
Bounds Test – F Statistics for the null of no cointegrating relation		
Test statistics	17.523	
Critical values	Lower bound	Upper bound
1%	5.15	6.36
5%	3.79	4.85
10%	3.17	4.14

Notes: Only the specific UECM from general UECM of one lag structure was reported here. The specific UECMs from general UECMs of two and three lags structure show that its errors are serially correlated based on Breusch-Godfrey LM statistics (2 orders) at 10% level that are 16.237 (p-value=0.0003) and 5.609 (p-value=0.06).
Refer to Footnote to Table 4 for other denotations.

As the results reported in Table 4 (for quarterly data) show the test statistics are above the 0.05 upper bound, 4.85 and, thus the null hypothesis of no cointegrating relationship can be rejected - Japan's aggregate imports, real income and relative import prices are cointegrated. This finding is found to be insensitive to different lag structure employed (12, 8 and 4 quarters). The results reported in Table 5 (biannual data) reveal a long run relationship for Japan's aggregate import demand function where the test statistics are above upper bound at 5 percent significance level (as well as at 1 percent). Given a small sample (28 annual observations), the UECM reported in Table 6 was estimated with one lag. The test statistic lies above the upper bound at 1 percent level, the null hypothesis of no cointegrating relationship among Japan's imports, real income and relative price of imports can be rejected. Totality, the results of bounds test provides conclusive finding that of cointegration relationships among Japan's aggregate imports, real income and relative price of imports even if quarterly, biannual and annual data used.

Johansen's Multivariate Cointegration Approach

Table 7 reports the results of Johansen's multivariate cointegration test for quarterly, biannual and annual data. The likelihood ratio tests (trace statistics) do reject the null hypothesis of no cointegrating relationships among Japan's imports, real income, and relative import prices. The finding is not sensitive with lag length of VAR. The Johansen's multivariate test does confirm long run relationships for Japan's aggregate import demand function, and which is consistent with the one obtained from bounds test approach. This finding supplements the early studies done by Bahmani-Oskooee and Niroomand (1998), Masih and Masih (2000), and Tang (2003a (Bounds test, DOLS and Johansen).

Table 7
Likelihood Ratio Tests – Trace Statistics (Series: $\ln M_t$, $\ln Y_t$, and $\ln RP_t$)

Quarterly Data: 1973:1 to 2000:4	Biannual Data: 1973:1 to 2000:2	Annual Data 1973 to 2000	Hypothesized No. of Cointegrating Equation(s)
Lags: 1 to 4	1 to 2	1 to 1	
28.617***	36.771*	42.942*	None
13.409***	15.106***	10.880	At most 1
3.367***	4.803**	3.743***	At most 2
Lags:1 to 8 #	1 to 4 #	1 to 2 #	
45.226*	31.546**	27.258***	None
12.645	13.230	13.814***	At most 1
3.534***	3.76***	3.894**	At most 2
Lags:1 to 12	1 to 6	1 to 3	
29.106***	28.743***	41.596*	None
11.939	11.013	10.184	At most 1
3.458***	4.921**	3.933**	At most 2

Notes: Test assumption: Linear deterministic trend in the data. The reported test statistics (Trace statistics) were adjusted with a scaling factor, $(T-nk)/T)$, where T is the sample size, n is the number of variables in the estimated system, and k is lag parameter (Reinsel and Ahn, 1988). # denotes optimum augmented lag was selected based on Akaike information criterion (AIC). The critical values are from Osterwald-Lenum (1992).
*Rejection of the null of r at most r_0 at 1% level. The critical values are 35.65, 20.04, and 6.65 for r_0 =none, r_0 =at most 1,and r_0 =at most 2, respectively.
**Rejection of the null of r at most r_0 at 5% level. The critical values are 29.68, 15.41, and 3.76 for r_0 =none, r_0 =at most 1,and r_0 =at most 2, respectively.
***Rejection of the null of r at most r_0 at 10% level. The critical values are 26.79, 13.33, and 2.69 for r_0 =none, r_0 =at most 1,and r_0 =at most 2, respectively.

Table 8 summarises the results of cointegration analyses, and reports the estimated cointegration equations (DOLS). However the Engle and Granger's (1987) DF and ADF tests and ECM tests as well show no cointegrating relationships in Japan's aggregate import demand behaviour. This study borrows the work by Pesavento (2004) to explain this phenomenon since Pesavento (2004) proposed a theoretical explanation for the common empirical results in which different tests for cointegration give different answers. The Pesavento's (2004) study utilised the ADF test on the residuals of the cointegration regression, Johansen's maximum eigenvalue test, the *t*-test on the error correction term, and Baswijk (1994) Wald test. The study analytically showed the asymptotic distributions of these tests depend on a single nuisance parameter under the local alternative. Theoretically, this parameter is a function of the long-run correlation of the errors in the cointegration relation to the shocks to the set of independent variables. Pesavento (2004) found that when this correlation is high, a full system approach is expected to perform better, and showed that the tests have significantly different performances for different values of the nuisance parameter. Pesavento (2004), finally concluded that the error correction with redundant regressor test and the Wald test not only perform better than other tests in terms of power in large and small samples but are also not worse or better in terms of size distortions.

Table 8
Summary for Cointegration Tests and Cointegrating Equation

Techniques:-	Quarterly Data	Biannual Data	Annual Data	Cointegrating equation [$\ln M_t$, $\ln Y_t$, and $\ln RP_t$]
Engle Granger (Appendix 2)	NO	NO	NO	
Bounds Test (Tables 4,5, and 6)	YES (AIC: 4 lags)	YES (AIC: 6 lags)	YES (AIC: 1 lags)	ARDL estimates: Quarterly: [-1, 0.13, -1.39] Bi-Annual: [-1, 0.77, -1.2] Annual: [-1, 1.39, -1.38]
Johansen Approach (Table 7)	YES (AIC: 8 lags)	YES (AIC: 4 lags)	YES (AIC: 2 lags)	DOLS estimates: Quarterly: [-1, 0.68, -0.58] Bi-Annual: [-1, 0.9, -0.54] Annual: [-1, 1.04, -0.4] (Appendix 3)
ECM (Appendix 4)	NO (AIC: 4 lags)	NO (AIC: 6 lags)	NO (AIC: 1 lags)	

Notes: YES denotes exists of a cointegrating relation by rejecting the null of no cointegration at 10% significance level. NO denotes no cointegrating relation by not rejecting the null of no cointegration at 10% significance level.

From the empirical results, this study does highlight the following outcomes. The first outcome is using appropriate cointegration techniques e.g. bounds and Johansen tests which are applicable for both $I(0)$ and $I(1)$ regressors. This study supports the existence of cointegrating relationships in Japan's aggregate import demand function. Secondly, income elasticity becomes elastic when low frequency data was used via annual data. Mah (1994, p.292, footnote 2) argued that "*Since the import demand is known to adjust to a shock in the variables on the right-hand side with usually more than a quarter*,…". Thirdly, the DOLS estimates shows the relative import prices variable is inelastic to imports which between -0.4 and -0.58, but its elastic when ARLD estimator was employed (between -1.2 and -1.4). Mah (2000, p.238) argued that "*In case that the data set is of small sample size, the conventionally used cointegraton tests are not reliable, but the unrestricted error correction model can give us the long run income and price elasticities* (Pesaran *et al*., 1996)". The estimated price elasticities are quite consistent with the use of quarterly, biannual and annual data. Finally, this study finds that data frequency does not affect estimates of Japan's aggregate import demand function. Using quarterly, biannual and annual data, cointegration relationships were found among Japan's aggregate imports, real income, and relative price of imports. With this finding, this study thus concludes that data frequency does not affect estimates of Japan's aggregate import demand function, but that the choice of cointegration techniques does. A number of studies (such as Hakkio and Rush, 1991; Charemza and Deadman, 1992) documented that the number of observations used does not affect the estimates of cointegration, but that the span of data (sample period) does.

4. Concluding Remarks

Using cointegration approach, this study has documented two basic findings: Firstly, cointegration relationships exist among Japan's imports, real income and relative import prices implying the traditionally used market force variables can explain the Japanese import demand behaviour. Secondly, this study finds that data frequency does not affect estimates of Japan's aggregate import demand function, but that the choice of cointegration techniques does.

These findings help to cast a light into empirical literature on Japan's import demand. Cointegration of Japan's import demand function indicates that Japan can reduce her trade surplus by stimulating domestic business condition since this stimulation exercise requires an assumption that Japan's import demand function to be stable (cointegrated). On the other hand, the estimated long run income elasticities are found to be elastic, and ranged between 1.04 and 1.39 which is based on the estimates of annual data by considering the requirement of longer transaction periods for imports demand behaviour. Accordingly, the income elasticity tells us that stimulating domestic activities may effectively increase imports and consequently deteriorates the trade surplus. The price variable is elastic with respect to imports. The elasticities ranged between -1.2 and -1.4 (ARDL estimates, consider the argument made by Mah 2000, p.238). This shows domestic prices should be pushed up in order to stimulate imports; and thus resulting in a smaller trade surplus. With this concern, devaluation becomes unfavourable for Japan since it may improve Japan's trade balance - the major concern of Japan's trade issue is to reduce her trade surplus.[9] Arize (2002) finds that imports and exports for Japan were cointegrated implying macroeconomic policies have been effective in the long run, and has suggested that Japan is not in violation of her international budget constraint. The estimated exports coefficient is 0.92 indicating exports expansion would increase imports. A recent work by Hatemi-J (2002) supports bi-directional causal relation between export growth and economic growth in Japan over the period 1960-1999. This finding does suggest that the expansion of exports is an integral part of the Japan's economic growth process. Of course, all these issues merit further research.

No study is free from shortcomings. The major caveat of this study is that it focuses on aggregate imports alone but this generates misleading results. A general disaggregated equation explaining imports from major Asian partners for major product categories (agricultural products, mineral fuels, machinery, and other

[9] Hamori and Matsubayashi (2001, p.135-136) recommended that Japan should reduce its trade surplus by stimulating domestic demand. For a single aggregate import demand function, Heien's (1968) argued that 'for any country a value of the price elasticity (demand for imports) between -0.5 and -1.0 is necessary to insure success of exchange depreciation'. The price estimation of DOLS in this present study ranges from -0.4 to -0.58. In this relation, the Marshall-Lerner condition indicates a stable foreign exchange market if the sum of price elasticity of demand for imports and the demand for exports exceed one (in absolute terms). Thus, exchange rate policy in particular devaluation can be adopted to improve the trade balances (see for example, Bahmani-Oskooee and Niroomand, 1998).

manufactures) is recommended for future study. Nonetheless, this study does provide important policy implications. Other issue is related to the potential of mis-specified of the underlying model since Japan suffered two massive oil shocks during the period under examination, that do not appear to be accounted for in any way in the model estimated. But, this issue is beyond the scope of this study. Perhaps, the relative price of oil is potentially an important omitted explanatory variable.

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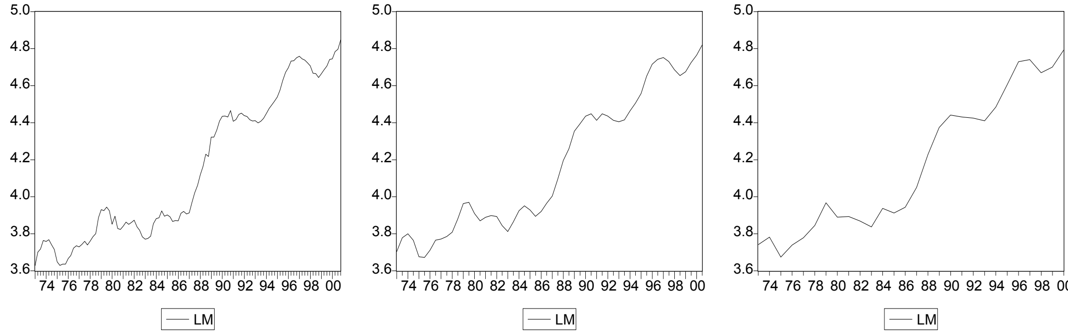
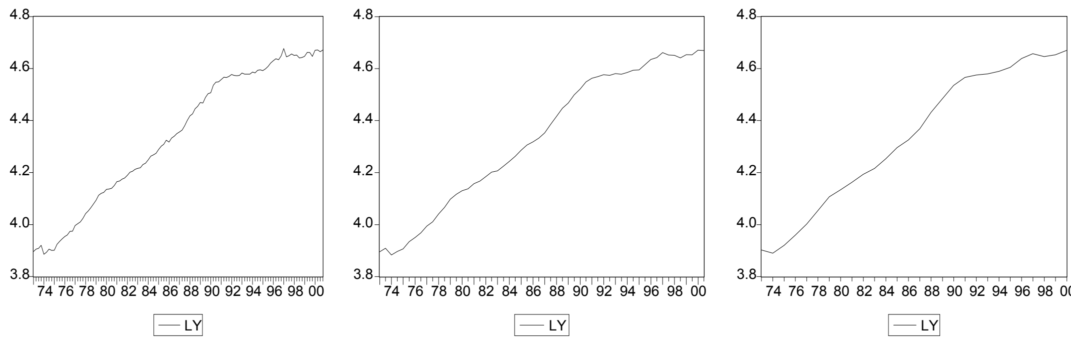
Appendix 1
Plots of Real Imports, Real Income, and Relative Prices for 1973-2000



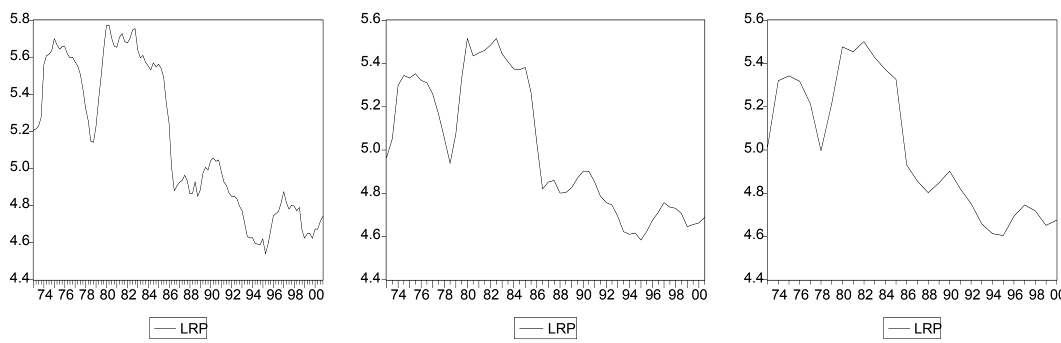
Quarterly Data Bi-Annual Data Annual Data

Real Imports ($\ln M_t$)



Quarterly Data Bi-Annual Data Annual Data

Real Income ($\ln Y_t$)



Quarterly Data Bi-Annual Data Annual Data

Relative Price of Imports ($\ln RP_t$)

Appendix 2
Residual Based Cointegration Tests (Engle and Granger, 1987)

The recent econometric literature has established that the estimation of standard ordinary least squares (OLS) estimator to be no longer consistent and may reflect a fundamental misspecification in the model, if regressing on nonstationary series that are not cointegrated. Further, an invalid inference can be drawn associated with *spurious* regression (Engle-Granger, 1987). Granger and Newbold (1974) have suggested that 'R-squared gather than the DW-d statistic' could be used as a good rule of thumb when the estimated regression (OLS) is suspected to suffer from spurious estimation. If the series have the same time series properties, $I(1)$, a stable long-run relationship can be estimated using standard ordinary least squares (OLS) technique. Based on Engle and Granger (1987) approach, a cointegrating equation for the Japanese import demand function can be specified as $\ln M_t = a + b_1 \ln Y_t + b_2 RP_t + e_t$. The nonstationary, $I(1)$ variables are said to be cointegrated, if the estimated residual from a cointegrating equation is stationary, $I(0)$ (Engle and Granger, 1987). That is to perform a unit root test (DF) on estimated cointegrating equation residuals, $\Delta e_t = c + \gamma e_{t-1} + u_t$ (without trend) in order to reject the null of no cointegration that is the e_t should be stationary at level, $I(0)$. The estimated cointegrating equation is cited below. We find R-square to be higher than the Durbin-Watson statistic. This indicates a possibilityof spurious regression, if the nonstationary series are not cointegrated. The OLS estimates and the results are reported below, no cointegrating relationships for Japan's aggregate import demand function.

The Estimated Cointegrating Equation (Dependent variable: $\ln M_t$)

Variables:	Quarterly Data 1973:1 to 2000:4	Biannual Data 1973:1 to 2000:2	Annual Data 1973 to 2000
$\ln Y_t$	1.105	1.064	1.048
	(13.32)*	(11.625)*	(7.89)*
$l \ln RP_t$	-0.221	-0.276	-0.292
	(-4.079)*	(-3.515)*	(-2.548)**
Constant	0.492	0.943	1.097
	(0.802)	(1.26)	(1.004)
R-squared	0.901	0.916	0.919
Durbin-Watson	0.079	0.196	0.525
Cointegration tests:			
Dickey-Fuller (DF)	-0.821	-1.122	-1.630
	-2.103 [4]	-1.722 [2]	-2.135 [1]
Augmented DF	-1.186 [8]#	-1.431 [4]	-1.458 [2]#
	-0.787 [12]	-0.6 [6]#	-1.16 [3]

Notes: (.) is t-statistic. [.] is the number of augmented lags in Augmented DF tests for cointegration from a maximum lag of three years. The unit root equation for cointegration test includes constant term (see Davidson and MacKinnon, 1993, p.721). The critical values are -4.29, -3.74, and -3.45 for 1, 5 and 10% levels (with three variables, k=3) (Davidson and MacKinnon, 1993, p.722 Table 20.2). Ordinary Least Squares (OLS) estimator has been used for estimations.
# denotes optimum augmented lags is selected based on Akaike information criterion (AIC).
*p<0.01 and **p<0.05 (two-tailed t-test)

Appendix 3
DOLS Estimation for Income and Price Elasticities
(Saikonnen, 1991; Stock and Watson, 1993)

The Dynamic OLS procedure (DOLS) developed by Saikonnen (1991) and Stock and Waston (1993) has the advantage that the endogeneity of any of the regressors has no effect, asymptotically, on the robustness of the estimates. It allows direct estimation of a mixture of $I(1)$ and $I(0)$ variables, which has been shown to perform well in finite samples. The procedure incorporates the lags and leads of the first differences of $I(1)$ variables. Thus, estimation of the long run relation between Y and X is carried out with a regression of the type: $Y = \lambda^{d\cdot} X + \sum_{-n}^{n} a_i \Delta X_{t-1}$, where $\lambda^{d\cdot}$ denotes the vector of long-run coefficients of X using the dynamic OLS procedure (DOLS).

One lead of first differenced explanatory variable(s) has been included in the DOLS regression due to the limited sample size used in this present study. Stock and Watson (1993) have noted that the choice of one or two leads for sample size of 100 has good size properties. General model of DOLS is based on 12 lags and 1 lead for quarterly data, 6 lags and 1 lead for bi-annual data, and 3 lags and 1 lead for annual data based on AIC. The reported DOLS estimates are in parsimony form. The general model or fairly unrestricted model was tested downwards sequentially to arrive at a parsimonious model using ‘general to specific’ methodology; that is those differenced lag and lead regressors that have relatively small absolute t-value (less than one) were dropped sequentially. The results are reported as below.

Estimated DOLS Regression (Dependent Variable: $\ln M_t$)

Variables:	Quarterly Data (1973:1 to 2000:4)	Bi-Annual Data (1973:1 to 2000:2)	Annual Data (1973 to 2000)
$\ln Y_t$	0.678*	0.90*	1.04*
$\ln RP_t$	-0.581*	-0.541*	-0.398*
$\Delta \ln Y_t$		0.957	2.225**
$\Delta \ln Y_{t-}$	1.119	1.709***	-1.12
$\Delta \ln Y_{t-2}$	1.250		1.070
$\Delta \ln Y_{t-3}$		-1.647***	-2.672*
$\Delta \ln Y_{t-6}$	-1.072	-2.111*	
$\Delta \ln Y_{t-7}$	-1.473***		
$\Delta \ln Y_{t-8}$	-1.229***		
$\Delta \ln Y_{t-12}$	-1.017		
$\Delta \ln Y_{t+1}$			-1.598***
$\Delta \ln RP_t$	0.749*	0.594*	0.475*
$\Delta \ln RP_{t-1}$	0.520*	0.526*	0.323*
$\Delta \ln RP_{t-2}$	0.596*	0.462*	0.242**
$\Delta \ln RP_{t-3}$	0.541*	0.396**	0.100
$\Delta \ln RP_{t-4}$	0.465*	0.388**	
$\Delta \ln RP_{t-5}$	0.465*	0.191	
$\Delta \ln RP_{t-6}$	0.417*	0.196***	
$\Delta \ln RP_{t-7}$	0.446*		
$\Delta \ln RP_{t-8}$	0.371**		
$\Delta \ln RP_{t-9}$	0.169		
$\Delta \ln RP_{t-10}$	0.282**		
$\Delta \ln RP_{t-11}$	0.247***		
$\Delta \ln RP_{t-12}$	0.262**		
$\Delta \ln RP_{t+1}$	0.149	0.236***	0.246**
Constant	4.265*	3.029*	1.734***
R-squared	0.981	0.897	0.996
F-statistic	177.099	183.233	187.265

Note: $*p<0.01$, $**p<0.05$ and $***p<0.10$ (two tailed t-test).

Appendix 4
Error Correction Mechanism Test (Banerjee *et al*., 1998)

Banerjee *et al*. (1998) proposed a new test for cointegration in a single-equation framework where regressors are $I(1)$ processes, which is based on the coefficient of the lagged dependent variable in an autoregressive distributed lag (ADL) model with leads of the regressors. The procedure depends upon the significance of the lagged dependent variable since this is equivalent to testing the significant of the error-correction terms in the ECM reparameterizaton of the model. Banerjee *et al*. (1998) recommended estimating the following (unrestricted) ECM regression by OLS: $\gamma(L)\Delta Y_t = \alpha(L)'\Delta X_t + \beta Y_{t-1} + \theta' X_{t-1} + \sum_1^s a'_j \Delta X_{t+j} + \varepsilon_t$, where $\gamma(L)$ and $\alpha(L)$ are polynomials in the lag operator, L. When β (or its t ratio) exceeds the critical values (provided in Banerjee et al. 1998), the null hypothesis of non-cointegration is rejected. From this procedure, the long run relationship, $Y = \lambda^e{}' X$, is also simultaneously estimated. The coefficients ($\lambda^e{}'$) of vector X from the estimated ECM procedure are given by $\lambda^e = \theta^e / \beta$.

The test statistics (t-statistics) for ECM tests using specific UECM (for the bounds test as presented in Tables 4, 5 and 6), show no cointegrating relationships among real imports, real income and relative price of imports via quarterly, biannual and annual data.

ECM Tests: t-statistic of the Coefficient of $\ln M_{t-1}$

	Quarterly data		Bi-Annual	Annual
	-2.856 (l=12)		-2.304 (l=6)	-2.445 (l=3)[a]
	-2.689 (l=8)		-2.964 (l=4)	-2.002 (l=2)[a]
	-1.684 (l=4)		-2.247 (l=2)	-1.619 (l=1)
T (Sample size)	Critical values[1]:	1%	5%	10%
25		-4.53	-3.64	-3.24
50		-4.29	-3.57	-3.20
100		-4.26	-3.56	-3.22

Notes: l is the lag structure included into general UECM. The reported t-ratio of coefficient of lagged one period level of *Ln*M is based on its parsimonious form. *a* denotes the errors are serially correlated based on Breusch-Godfrey serial correlation LM test.

[1] Source: Banerjee *et al*., 1998, p.276 Table 1.A), k = 2 (regressors)