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AGGREGATE IMPORT DEMAND AND EXPENDITURE COMPONENTS IN JAPAN: AN EMPIRICAL STUDY

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ABSTRACT

This Paper investigates the determinants of Japan's aggregate import demand function. In contrary with traditional specification of using single real income variable, the present study examines the various components of real income that are final consumption expenditure, domestic investment and expenditure on export goods, and relative prices. The 'bounds' testing approach in Pesaran et al. (2001), which based on estimation of unrestricted error-correction model (UECM) was used for cointegration analysis over the sample period 1973-1997. The result confirms a cointegrating relation among the quantity of import and its determinants as well as various expenditure components. The various expenditure components provide different impacts on Japan's imports volume in short run. The UECM appears to track the data well.

Keywords : aggregate import demand function, expenditure components, cointegration, Bounds test, unrestricted error correction model.

ABSTRAK

Kertas ini mengkaji hubungan kointegrasi antara import agregat Japan dan penentu-penentunya. Berbeza dengan spesifikasi tradisional yang menggunakan pendapatan benar, kertas ini mengguna pelbagai komponen daripada pendapatan benar iaitu perbelanjaan penggunaan, pelaburan tempatan, dan eksport sebagai penentu-penentu permintaan import. Penentu yang lain ialah harga relatif. Pendekatan ujian 'bounds' (Pesaran et al., 2001) yang berasaskan penganggaran model pembetulan ralat tidak terbatas telah digunakan untuk analisis kointegrasi bagi data tahunan 1973 hingga 1997. Keputusan ujian 'bounds' memastikan bahawa wujudnya satu hubungan *jangka panjang antara volum import, dan penentu-penentunya. Hal ini menunjukkan bahawa fungsi permintaan import di Jepun adalah stabil sepanjang tempoh sampel. Beberapa implikasi polisi telah dibincangkan dalam kertas ini.*

Kata Kunci: fungsi permintaan import agregat, komponen perbelanjaan, kointegrasi, ujian 'bounds', model pembetulan ralat tidak terbatas.

INTRODUCTION

The trade data from World Tables (World Bank, 2002) shows that Japan has enjoyed trade surplus over the period 1960 to 1999; with only deficits in 1961 (¥ 312billion), the period 1963-1964 (¥210-258billion), in 1974 (¥1,038billion), and the period 1979-1980 (¥2,055-2.220billion). Over the period 1973-1979, trade deficits caused by the rise in oil prices led to a renewed push to increase exports. Merchandise exports rose 21 per cent over the decade, and though the volume of imports increased as well, this did not lead to huge surpluses. The new round of oil shocks in 1979 had resulted in brief trade deficits again. For the period 1987-1989, rising demand for imports had eased some trade frictions and built domestic pressure, reducing barriers in a country that had always prioritized producers over consumers. However, huge surpluses and conflicts continued, particularly with the U.S., which in 1989 categorized Japan as an unfair trading nation and had begun to hold talks to reduce structural impediments such as distribution and investment practices. In the past decade (1990-2002), the Japanese trade balance has declined annually, especially in the 1990s when it was hurt first by the increasing competition from Asian countries, and then by a regional and global economic slowdown. U.S. talks continue to include demands on the Japanese to open their markets to goods and retailers, but the talks also now include efforts to support the Japanese economy, something which was unthinkable a decade before.[1]

Table 1 reports the import demand structure of the Japanese. The country's import of agricultural raw materials reduced dramatically from 20.4 per cent in 1965 to 5.6 per cent in 1995, and a downward trend has been observed since. In addition, the percentage of imports for food, and ores and metals also showed downward trends over the period 1970-2000. Imports in the manufacturing sector increased its share from 19.4 per cent (1965) to 44.1 per cent, and 54.3 per cent in 1990 and 1995, respectively. The volume of fuel imports maintained substantially high levels for the period 1974-1985, with about 45 per cent of the total imports compared to approximately 20 per cent for other periods. Structural developments within the Japanese economy may explain the changes in import behaviour (see Ceglowski, 1996, p. 444). On the other hand, the total share of manufacturing exports compared to total exports was about 94 per cent for the period 1962 to 1997. For other categories like agricultural raw materials, food, fuel, ores and metals, the average total share was below one per cent the total exports over the period 1984 to 2000 (it was about one per cent for ores and metals). The above statements are based on the statistics from the World Tables (World Bank, 2002).

Table 1
Share of Japanese Total Imports for the Period 1965-1995
(in percentage)

Year	Agricultural raw materials	Food	Fuel	Ores & metals	Manufacturing
1965	20.4	22.7	19.9	17.3	19.4
1970	16.2	16.9	20.7	10.8	17.7
1980	8.6	12.0	50.0	10.0	18.7
1985	6.7	13.9	43.8	8.9	25.4
1990	6.5	14.7	24.5	9.0	44.1
1995	5.6	16.2	16.1	6.6	54.3

Source: World Tables (World Bank, 2002).

Empirical attempts at estimating import demand behaviour have received increasing attention in international economics because of its relevant policy implications. Heien (1968) has argued that 'for any country a value of the price elasticity (import demand) between -0.5 and -1.0 is necessary to ensure success of exchange depreciation'. According to Reinhart (1995, p. 291), relative prices play a significant role in the determination of trade flows, helping to buttress policies of devaluation as a way to correct trade imbalance which is based on the relative price variable in *static* or *long run* specifications of import demand or export supplies. It underscores the necessity to examine the presence of a long run relationship or a cointegrating relation of trade equations – imports or exports demand function. Therefore, estimation of import demand behaviour is essential for designing both the exchange rate and trade policies that can be implemented to improve a country's external balance – trade and current accounts.

Existing empirical studies which investigate the presence of a long run equilibrium relationship among the variables of Japan's import demand function are based on traditional specifications, i.e., the quantity of import demanded is determined by domestic real income and the ratio of import prices to domestic prices. Some of these studies are Mah (1994), Masih and Masih (2000), Hamori and Matsubayashi (2001), and Tang (2003). However, the results from these studies are mixed and inconsistent. Mah (1994) used biannual data from 1974:1 to 1990:2 to analyse Japan's import demand function in the long run. The results of Engle-Granger's (1987) cointegration approach, which is based on the Dickey-Fuller normalized bias and Phillip tests cannot reject the null hypothesis of no cointegration at the 10 per cent level. Furthermore, Masih and Masih (2000) have re-examined Mah's (1994) work by using Johansen's multivariate cointegration tests (Johansen, 1988; Johansen and Juselius, 1990), an approach which is more approximate than the Engle-Granger method in multivariate vectors analysis (more than two variables). They employed the same data as in Mah (1994), and the results revealed that the quantity of imports demanded, real income and relative prices were cointegrated at a 5 per cent significance level. As a result, they have pointed out that the rejection of cointegration should also be thoroughly justified in the light of certain destabilizing forces, structural breaks, and omission of relevant theoretically inferred variables.

A recent study by Hamori and Matsubayashi (2001) used the following standard cointegration tests viz. the Engle-Granger residual based approach and Johansen test, as well as the Gregory and Hansen (1996) approach which to address the issue of structural breaks in order to re-examine the long run relationship of Japan's import demand function. No long run relationship was found in the Engle-Granger test as well as in the Gregory and Hansen approach. However, the Johansen test (trace statistics) has detected at least one cointegrating vector based on one and eight lags length specification of VAR (vector autoregression). Other lags length specification - four lags failed to confirm any cointegrating vector. All of these analyses were based on a 5 per cent significance level, and the sample period was from quarter one of 1973Q1 to quarter one of 1998Q1 (Q is quarter). Hamori and Matsubayashi (2001) concluded that there was no cointegrating relations among the examined variables in Japanese imports demand function.

Recently, Tang (2003) used the bounds test procedure (Pesaran *et al.*, 2001), and found that the volume of imports, real income and relative prices term were cointegrated in Japan for the period 1973 to 1997. The estimated long run elasticities for relative price and real income are -0.82 and 0.99 respectively.

The present study will also use the bounds test procedure (Pesaran *et al.*, 2001) to examine annual time series data, and re-estimate the presence of a long run relationship of Japanese import demand function. In contrast to previous studies of using traditional import demand specifications, the present study has disaggregated the single demand variable (Real Gross Domestic Product) into various final expenditure components; namely, final consumption expenditure (private and Government consumptions), expenditure on export goods, and gross domestic investment. The need to disaggregate is based on the argument that the composition of expenditure will be important if the various components of expenditure have different import contents. The composition of expenditure is important to the extent that the import content of the different components of expenditure differs (Thirlwall & Gibson, 1992; Giovannetti, 1989; Davies, 1990). Abbott and Seddighi (1996, p. 1120) noted that in the UK and other EU member countries, the evidence available from input-output data suggests that each macro component of final expenditure corresponds to different aggregate propensities to import. It was found that if the composition of demand changes, the aggregate import propensity will change even if the disaggregated marginal propensities are unchanged (Giovannetti, 1989, p. 960). If this were true, the use of a single demand variable would lead to aggregation bias and is a possible cause of the no cointegration relationships pointed out in Mah (1994) and Hamori and Matsubayashi (2001). Among the empirical studies on import demand estimate that have considered this issue are Giovannetti (1989), Thirlwall and Gibson (1992), Abbott and Seddighi (1996), Mohammad and Tang (2000), and Tang (2001).

MODEL SPECIFICATION AND METHOD

The traditional specification for import demand function sees the quantity of import demanded as a function of domestic real income and the ratio of import prices to domestic price (Mah, 1994; Masih & Masih, 2000; Hamori & Matsubayashi, 2001; Tang, 2003). This can be expressed as Equation 1 below:

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

where at period t, M_t is the desired quantity of import demanded, and which is defined as nominal imports of goods and service divided by import price deflator; Y_t is real Gross Domestic Product (GDP), and RP_t is the ratio of the import price index and the domestic price level (referred to as relative price henceforth). In the present study, the single real income variable is disaggregated into three broad expenditure components viz., final consumption expenditure, expenditure on exports, and domestic investment (Giovannetti, 1989; Abbott & Seddighi, 1996; Mohammad & Tang, 2000; Tang, 2001). A log-linear model is specified as follows:

$$LnM_t = a_0 + a_1 LnFCE_t + a_2 LnE_t + a_3 LnGDI_t + a_4 LnRP_t + u_t \quad (2)$$

where *FCE* is the final consumption expenditure (private plus Government consumption expenditure), *E* is expenditure on exports, *GDI* is gross domestic investment, and *Ln* is natural logarithm. All of the series are in real terms, nominal value divided by GDP deflator (1995=100). u_t is a random error assumed to satisfy classical assumptions. From economic theory, it is assumed that the signs of the coefficients a_1, a_2, and a_3 are positive and a_4 is negative.

The sample period is from 1973 to 1997, and this has yielded 25 annual observations. The source of the data is from World Tables (World Bank, 2002). The present study followed the sample period found in Hamori and Matsubayashi (2001, p. 136) because it concurs with their justification that the sample period from 1973Q1 to 1998Q1, corresponded to the period of a flexible exchange rate regime. Note that Japan's fixed exchange rate of ¥360.0000 to the US dollar has been removed in 1971. In addition, Hakkio and Rush (1991) have argued that increasing the number of observations by using monthly or quarterly data do not add any robustness to the results in cointegration analysis, because the concern is the length of the period under consideration. The use of traditional cointegration techniques proposed by Engle and Granger (1987), Johansen (1988) and Johansen and Juselius (1990) to capture the long run import demand behaviour may be unreliable (Mah, 2000, p. 243). These techniques are found to be inappropriate for small samples (see Pattichis, 1999; Mah, 2000). The Monte Carlo studies have shown, however, that despite the super-consistency of the OLS (Ordinary Least Squared) estimator in a cointegrating regression, substantially biased estimates could results result in small samples (Banerjee *et al.*, 1993, p. Chapter 7). Cheung and Lai (1993) have stressed that Johansen's multivariate test would reject the null of no cointegration too often when the sample size of the data used is rather small, and when one uses the asymptotic critical values. On the other hand, Toda (1994, p.78) on the basis of the Monte Carlo evidence presented has argued that a sample size of 300 or more observations is considered necessary to ensure good performance of the Johansen's likelihood ratio test for cointegrating ranks.

Another reason to use annual data is to avoid the problems of using seasonally adjusted or unadjusted data. Variables will tend to be incorrectly regarded as random walks when using seasonally adjusted data for a unit root test, although they are actually not random walk variables (Davidson & McKinnon, 1993). Davidson and MacKinnon (1993, p. 714) have shown that to avoid the biases of using seasonally adjusted data to compute unit root tests, one has to use annual data, and the power of these tests depends more on the span of the data (i.e., the number of years the sample covers) than on the number of observations. Note that unit root equation is widely used for cointegration analysis, for example in the Engle and Granger (1987) approach. Hamori and Matsubayashi (2001, p. 136) used seasonally adjusted data in their study. In addition, Granger and Hallman (1989) have pointed out that the use of seasonal data to estimate the long run model may give rise to inconsistent estimates of the long run parameters. Charemza and Deadman (1992, p. 153) have recommended that *"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"*.

Furthermore, Pattichis (1999) and Mah (2000) have recommended the use of a robust estimation method for a small sample analysis i.e., 'bounds' testing approach (Pesaran *et al.*, 1996; 1999; 2000; 2001). This approach is based on an estimation of the UECM or conditional ECM and a critical 'bounds' test for the existence of a long run relationship as well as estimation for long and short run coefficients. The ECM-based cointegration test will be more powerful than the residual-based Engle-Granger test, and will generally give unbiased estimates of the long-run relationship and standard t-statistics for conducting statistical tests of significance (see Pattichis, 1999, p. 1062, footnote 2). The Pesaran *et al.*'s approach has two main advantages over the typical cointegration approaches (Engle & Granger, 1987; Johansen, 1988; Johansen & Juselius, 1990). Firstly, the 'bounds' test procedure can be applied irrespective of whether the explanatory variables are I(0) or I(1). Unlike standard cointegration tests, there is no need for unit root pre-testing if a conclusion can be made from the bounds test for cointegration (Pesaran *et al.*, 2001). Secondly, this method can be applied to studies that have small samples. The UECM test is likely to have better statistical properties since it does not push the short-run dynamics into the residual term as in the case of the Engle-Granger technique (Pattichis, 1999, p. 1062). Among the empirical studies with small sample sizes which have applied the Pesaran *et al.* (2001) approach for estimating import demand function, are Pattichis (1999), Mah (2000), Tang and Nair (2002), and Tang (2003). For example, using Korean annual data from 1980 to 1997 (18 observations), Mah (2000) has examined a long run relationship of disaggregated import for Information Technology products.

The UECM for the above import demand function [Equation (2)] will be expressed as follows:

$$\Delta LnM_t = b_0 + \sum_{i=0}^{k1} b_{1i} \Delta LnFCE_{t-i} + \sum_{i=0}^{k2} b_{2i} \Delta LnE_{t-i} + \sum_{i=0}^{k3} b_{3i} \Delta nGDI_{t-i}$$
$$+ \sum_{i=1}^{k4} b_{4i} \Delta nRP_{t-i} + \sum_{i=1}^{k5} b_{5i} \Delta LnM_{t-i} + b_6 \mathrm{LnM}_{t-1} + b_7 LnFCE_{t-1} +$$
$$b_8 LnE_{t-1} + b_9 LnGDI_{t-1} + b_{10} LnRP_{t-1} + u_t \qquad (3)$$

where Δ is a first difference operator.

To investigate the existence of a long run relationship of the examined variables, Pesaran *et al.* (2001) proposed the 'bounds' test based on the Wald or *F*-statistic. The asymptotic distribution of the *F*-statistic is non-standard under the null hypothesis of no cointegration relationship between the examined variables, irrespective of whether the explanatory variables are purely I(0) or I(1). The fist step is to test the null hypothesis by considering a restricted error correction model (RECM) for traditional import demand function in (3) by excluding the lagged level variables, viz., $LnM_{t-1,}$ $LnFCE_{t-1,}$ $LnE_{t-1,}$ $LnGDI_{t-1,}$ and $LnRP_{t-1}$. More formally, a joint significance test will be performed in order to test the null hypothesis of $b_6 = b_7 = b_8 = b_9 = b_{10} = 0$ (no cointegrating relation) against the alternative hypothesis of $b_6 \neq {}_0, b_1 \neq 0, b_8 \neq 0, b9 \neq 0, b_{10} \neq 0$ (a cointegrating relation).

Based on some conventionally used significance level ($\alpha = 0.1, 0.05$ or 0.01), if the computed *F*-statistic (Wald test) exceeds the upper bound (critical value), the null hypothesis of no cointegration can be rejected. This implies a long run equilibrium relationship among the examined variables. For the case when the computed *F*-statistic has fallen below the lower bounds, than the null hypothesis of no cointegration cannot be rejected. The conclusion is no cointegration. However, when the computed *F*-statistic falls between the upper and lower bounds, a conclusive inference cannot be made. Here, the order of integration for the explanatory variables, I(d) must be known before any conclusion can be drawn.

Based on the estimation of the UECM (Equation 3), the long run elasticities (coefficients) can be calculated, which is the estimated coefficient of the one lagged level explanatory variable(s) divided by the coefficient of the one lagged level dependent variable and then multiplied with a negative sign (Bardsen, 1989). For example, the long run price elasticity is $-(b_{10}/b_6)$. The coefficients of the first differenced variables in Equation (3) represents short run elasticities.

Table 2

Simplest UECM for Japan's Aggregate Import Demand Function

Dependent Variable: ΔLnM_t		
Variables:	Coefficients:	t-Statistic
Constant	1.145	1.001
$\Delta LnFCE_t$	1.546**	2.419
ΔLnE_t	0.808*	3.839
$\Delta LnGDI_t$	0.275	1.451
$\Delta LnRP_t$	-0.515*	-3.994
LnM_{t-1}	-0.329*	-3.280
$LnFCE_{t-1}$	0.340	1.394
LnE_{t-1}	-0.001	-0.160
$LnGDI_{t-1}$	-0.123	-0.845
$LnRP_{t-1}$	-0.285***	-2.092

Notes: *, **, and *** denote significance at 1%, 5%, and 10% level respectively.
Estimation method: Least Squares.
R-squared: 0.898 Adjusted R-Squared: 0.832
Sum of Squared error: 0.0135 Durbin-Watson-d: 2.061
F-statistics: 13.629 (0.000) Q-Statistics [12]: White noise
Jarque-Bera: 0.789 (0.674)
LM Test [2]: 3.743 (0.154); [3]: 6.156 (0.104)
ARCH Test [1]: 0.079 (0.778); [2]: 0.167 (0.919)
Ramsey RESET Test [1]: 1.124(0.289); [2]: 2.073 (0.355)
The numbers in parentheses are the p-values

EMPIRICAL RESULTS

The estimated UECM of Equation (3) with ARDL of $k1=k2=k3=k4=k5=0$ is reported in Table 2.[2] The UECM has passed a number of diagnostic tests. The result of Ramsey RESET test rejects the presence of a general specification error. The Jarque-Bera statistic has confirmed normality of estimated residual series. The Breusch-Godfrey LM statistic rejects the presence of second and third order serial correlation. ARCH test rejects the heteroscedasticity in the disturbance term. The plots of CUSUM and CUSUM of Squares tests (Figure 1 and Figure 2) provide evidence that the estimated parameters are stable over the sample period.

Figure 1
Plot of CUSUM Test

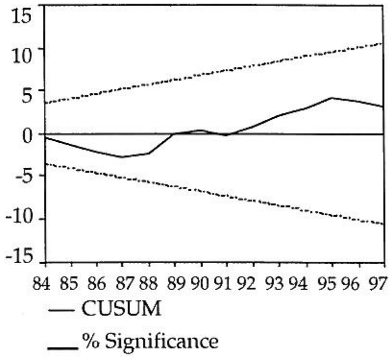

Figure 2
Plot of CUSUM of Squares Test

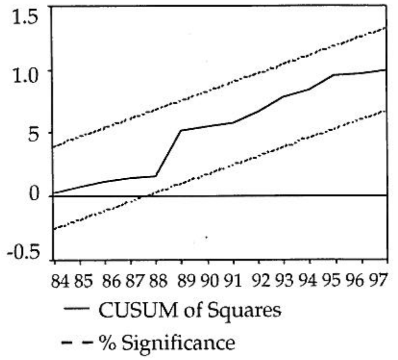

Table 3
Bounds Testing for Cointegration Analysis

Computed *F*-statistic:	**8.0507**
Critical Bounds at 1 % level:	
Lower bound	3.74
Upper bound	5.06

Notes: The *F*-statistic is a joints test for the coefficients of $LnM_{t-1,}$ $LnFCE_{t-1,}$ $LnE_{t-1,}$ $LnGDI_{t-1,}$ *and* $LnRP_{t-1}$ equal to zero. The reported critical bounds are taken from Pesaran *et al.* (2001, p. 300), Table CI(iii) Case III: Unrestricted intercept and no trend case with four regressors case.

To ascertain the presence of a long run relationship or a cointegrating relation among the interested variables in Equation (2), the 'bounds' testing approach (Pesaran *et al.*, 2001) was used. The results of the 'bounds' test are reported in Table 3. The computed *F*-statistic (Wald test) is 8.05, which exceeds the upper bound of 5.06. It is at the one per cent significance level, indicating that the null of no cointegrating relation can be rejected. This implies that the volume of imports, final consumption expenditure, expenditure on exports, domestic investment, and relative price are cointegrated for Japan. This finding is found to be consistent with Tang's (2003) study, which used a similar method (bounds test). However, he had used real GDP as a scale variable following the traditional specification for Japanese import demand.

The relative price variable was significant at the 10 per cent level in long run and its elasticity was -0.866.[3] This estimate is close to that found in Tang's (2003) work; the estimated long run price elasticity for Japanese imports demand then was –0.82. The various expenditure components of final demand were insignificant at the 10 per cent level. However, in the short run the components of final expenditure were significant, except for domestic investment. The final consumption expenditure recorded the highest elasticity, it was at 1.54 (significant at the 5 per cent level), and this was followed by expenditure on exports (0.808). The short run relative price elasticity was -0.515. The signs of the estimated elasticities, both for long and short runs were in accord with current economic theory. The R-squared showed that 89 per cent of the variations of Japanese import demand could be explained by the examined determinants. The estimated coefficient of error correction term was represented by the coefficient of the one lagged level dependent variable, b_6 that was -0.329. It had the correct sign (negative), and this indicated that the speed of adjustment from short run disequilibrium towards the long-run equilibrium state was about 32.9 per cent per year.

CONCLUSION

The main concern of this study was to ascertain the long run relationship of Japan's imports demand function. In doing so, the foregoing discussions had considered the various effects of final expenditure components on imports demand function, namely final consumption expenditure, exports, and domestic investment. The study had covered the annual period from 1973 to 1997. To provide a reliable long run estimate of Japan's imports demand function, the present study had taken into account the problem of small sample bias (see Pattichis, 1999; Mah, 2000), and had employed a recently developed estimation method, the 'bounds' testing approach (Pesaran *et al.*, 2001). The analysis carried out revealed that the volume of imports, the three components of final demand expenditure, and relative prices were cointegrated. This finding is consistent with that found in Tang's (2003) study, but contrary to the conclusions in other studies that no such long run relationship exists (Mah, 1994; Hamori & Matsubayashi, 2001). The findings presented in this paper have relevant implications for policy designed to improve Japanese trade performance in the long and short runs. The long run effects of the three broad areas of expenditure components were found to be insignificant on the quantity of imports demanded. This would seem to suggest that macroeconomic policy does not yield a favourable outcome on imports demand behaviour in the long run. However, more interesting is the suggestion that the examined series tend to move together in the long run, indicating that Japanese imports demand function was stable over the period analysed. This finding supports the recommendation that Japan should reduce her trade surplus by stimulating domestic business conditions, a recommendation that tacitly assumes stability of Japan's import demand function (Hamori & Matsubayashi, 2001, p. 135). Japan's import is closely related to domestic business conditions and the relative prices of imports (Hamori & Matsubayashi, 2001, p. 135-136).

In the short run, the final consumption expenditure variable is found to be elastic, at 1.5, and the export elasticity is 0.8. Both factors are significant at the 10 per cent level. The domestic investment variable is insignificant. It implies that fiscal policy design can contribute to a reduction of imports pressure by lowering some components of final demand, particularly final consumption goods, and expenditure on exports in the short run. Reducing exports lead to a lowering of import demand pressure in the short run, as exports variable is significant with an elasticity of 0.8. The linkages between exports and imports can explain this. The inter-relationship between exports and imports have been statistically supported in a study by Arize (2002, p. 108, Table 1), who has found that exports and imports were cointegrated in Japan, suggesting that the Japanese imports and exports have been brought into a state of long run equilibrium through the combined effects of all macroeconomic policies on trade balance. The estimated exports elasticity with respect to imports ranged from 0.9 to 1.2. The ongoing trend in Japanese production involves greater use of imported intermediate parts and supplies as substitutes for both imported raw materials and domestic intermediate goods (Ceglowski, 1996, p. 453). In addition, Hatemi-J (2002) has found that the expansion of exports is an integral part of the economic growth process in Japan.

The estimated long run price elasticity, -0.87 has fallen within the range suggested by Heien (1968) viz., -0.5 and -1.0, suggesting that exchange rate policies or devaluation can be used to influence Japanese external trade balance. This implication is supported by Arize's (2002) study that imports and exports in Japan were cointegrated. It would seem to suggest that the exchange rate policy and other macroeconomic policies have favourable combined effects on the trade balance (see Bahmani-Oskooee & Rhee, 1997). A point to note is that the Marshall-Lerner condition is not discussed here since the present study is only concerned with estimating import demand function. In addition, the relative price is significant both in long and short runs with estimated elasticity of -0.87, and –0.51 respectively, reflected in the assumption that domestic prices would increase the volume of imports. Even though, this is not a crucial issue since Japanese inflation is relatively low in the recent decade – below one percent or deflation for 1994 to 2001, except for 1.7 per cent in 1997 (World Bank, 2002).

NOTES

1. <http://www.pbs.org/wgbh/commandingheights/lo/countries/jp/jp_trade.html>
2. A set of UECMs with various lags length was estimated, and the fittest UECM was selected based on a set of diagnostic tests. First, UECMs based on autoregressive distributed lag equation, ARDL ($k1=k2=k3=k4=k5=$,1), and ARDL ($k1=k2=k3=k4=k5=$,2) were estimated. We were unable to perform lags length of three due to the limited annual observations. The general UECMs as well as final UECMs (following general-to-specific procedure, that is all those first differenced variables that have relatively small absolute t-value (less than one) were dropped sequentially) are problematic. These UECMs have failed to satisfy a number of diagnostic tests, particularly RESET for model misspecification, Q-tests and LM test for autocorrelation, and the Jarque-Bera for residuals normality. Second, a UECM with ARDL ($k1=k2=k3=k4=0,k5=1$) was estimated and it passed a number of the diagnostic tests, but showed second and third order autocorrelation, and model misspecification (RESET). Lastly, UECM with ARDL ($k1=k2=k3=k4=k5=0$) had passed the entire diagnostic tests.
3. Following Bardsen (1989), the calculation for the long run coefficient from a UECM is the estimated coefficient of one lagged level explanatory variable(s) divided by the coefficient of one lagged level dependent variable, and then multiplied by a negative sign. Thus, calculation for the long run elastic-

ity of relative price variable as in this study is $-(b_{10}/b_6)$ (as in Equation (3)), that is -(-0.285 / -0.329) = -0.866 (see the estimated coefficients from Table 2).

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REFERENCES

Abbott, A.J. & Seddighi, H.R. (1996). Aggregate imports and expenditure components in the UK: An empirical analysis. *Applied Economics, 28,* 1119-1125.

Arize, A.C. (2002). Imports and exports in 50 countries tests of cointegration and structural breaks. *International Reviews of Economics and Finance, 11,* 101-115.

Bahmani-Oskooee, M. & Rhee, Hyun-Jae. (1997). Are imports and exports of Korea cointegrated?. *International Economic Journal, 11*(1), 109-114.

Banerjee, A. Dolado, J.J., Galbraith, J.W. & Hendry, D.F. (1993). *Cointegration, error correction, and the econometric analysis of non-stationary data.* Oxford: Oxford University Press.

Bardsen, G. (1989). Estimation of long-run coefficients in error correction models. *Oxford Bulletin of Economics and Statistics, 51,* 345-350.

Ceglowski, J. (1996). The recent behavior of Japanese imports: A disaggregated analysis. *Japan and the World Economy, 8,* 443-457.

Charemza, W. W. & Deadman, D. F. (1992). *New directions in econometric practice.* England: Edward Elgar.

Cheung, Y. W. & Lai, K. S. (1993). Finite-sample sizes of Johansen's likelihood ratio tests for cointegration. *Oxford Bulletin of Economics and Statistics, 13,* 277-280.

Davies, G. (1990). The capital account and the sustainability of the UK trade deficit. *Oxford Review of Economic Policy,* 28-40.

Davidson, R. & MacKinnon J.G. (1993). *Estimation and inference in econometrics.* Oxford: Oxford University Press.

Engle, R.F. & Granger, C.W.J. (1987). Cointegration and error correction: representation, estimation and testing. *Econometrica, 55*

(2), 251-276.
Engle, R.F., Granger, C.W.J. & Hallman, J.J. (1989). Merging short-and long-run forecasts: An application of seasonal cointegration to monthly electricity sales forecasting. *Journal of Econometrics, 40*, 45-62.
Gregory, A. W. & Hansen, B. E. (1996). Residual-based tests for cointegration in models with regime shifts. *Journal of Econometrics, 70*, 99-126.
Giovannetti, G. (1989). Aggregated imports and expenditure components in Italy: An econometric analysis. *Applied Economics, 21*, 957-971.
Hakkio, C. S. & Rush, M. (1991). Cointegration: How short is the long run?. *Journal of International Money and Finance, 10* (4), 571-581.
Hamori, S. & Matsubayashi, Y. (2001). An empirical analysis on the stability of Japan's aggregate import demand function. *Japan and the World Economy, 13*, 135-144.
Hatemi-J, A. (2002). Export performance and economic growth nexus in Japan: A bootstrap approach. *Japan and the World Economy, 14*, 25-33.
Heien, D. M. (1968). Structural stability and the estimation of international import price elasticities in world trade. *Kyklos, 21*, 695-711.
Johansen, S. (1988). Statistical analysis of cointegrating vectors. *Journal of Economic Dynamics and Control, 12*, 231-254.
Johansen, S. & Juselius, K. (1990). Maximum likelihood estimation and inference on cointegration with applications to the demand for money. *Oxford Bulletin of Economics and Statistics, 52*, 169-210.
Masih, R. & Masih, A. M. M. (2000). A reassessment of long-run elasticities of Japanese import demand. *Journal of Policy Modeling, 22*, 625-639.
Mah, J.S. (1994). Japanese import demand behaviour: The cointegration approach. *Journal of Policy Modeling, 16*, 291-298.
Mah, J. S. (2000). An empirical examination of the disaggregated import demand of Korea - the case of information technology products. *Journal of Asian Economics, 11*, 237-244.
Mohammad H.A. & Tang, T.C. (2000). Aggregate imports and expenditure components in Malaysia: A cointegration and error correction analysis. *ASEAN Economic Bulletin, 17*(3), 257-269.
Pattichis, C.A. (1999). Price and income elasticities of disaggregated import demand: Results from UECMs and an application. *Applied Economics, 31*, 1061-1071.
Pesaran, M. H., Shin, Y. & Smith, R.J. (2001). Bounds testing approaches

to the analysis of level relationships. *Journal of Applied Econometrics, 16*, 289-326.

Pesaran, M. H., Shin, Y. & Smith, R. J. (2000, 1999). *Bounds testing approaches to the analysis of level relationships* (Working Paper Series, Nos. 9907). Department of Applied Economics, University of Cambridge (revised in 2000).

Pesaran, M. H., Shin, Y. & Smith, R. J. (1996). *Testing for the existence of a long-run relationship* (Working Paper Series, Nos. 9622). Department of Applied Economics, University of Cambridge.

Tang, T.C. (2003). Japanese aggregate import demand function: Reassessment from 'bounds' testing approach. *Japan and the World Economy, 15*(4), 419-436.

Tang, T.C. & Nair, M. (2002). A cointegration analysis of Malaysian import demand function: Reassessment from the bounds test. *Applied Economics Letters, 9*, 293-296.

Tang, T.C. (2001). Disaggregated import demand and expenditure components in Malaysia: An empirical analysis. *Conference proceedings of the Academy of International Business (AIB), South East Asia Region (AIBSEAR) 2001 Conference*, 4-6 July 2001, Jakarta, Indonesia [CD-Rom].

Thirlwall, A.P. & Gibson, H.D. (1992). *Balance of payments theory and the United Kingdom experience.* London: Macmillan.

Toda, H.Y. (1994). Finite sample performance of likelihood ratio tests for cointegrating ranks when liner trends are present. *Review of Economics and Statistics, 76*, 66-79.

Reinhart, C.M. (1995). Devaluation, relative prices, and international trade: Evidence from developing countries'. *IMF Staff Papers, 42* (2), 290-312.

World Bank. (2002). *World Tables.* (The time series data express, EconData Pty. Ltd.), 28 February, 2002.