IIUM Journal of Economics and Management 10, no.2 (2002):

AGGREGATE IMPORT DEMAND BAHAVIOR FOR INDONESIA: EVIDENCE FROM BOUNDS TESTING APPROACH

Tuck Cheong Tang*

School of Business and Information Technology, Monash University Malaysia, 2 Jalan Kolej, Bandar Sunway, 46150 Petaling Jaya, Selangor Darul Ehsan, Malaysia (e-mail: tang.tuck.cheong@ busit.monash.edu.my)

ABSTRACT

This paper empirically re-estimates the long-run relationship between Indonesian aggregate import demand and its determinants; namely, real income and relative import prices. In contrast to previous studies (Reinhart, 1995; and Senhadji, 1998), the result of the bounds test (Pesaran et al., 2001) reveals that import volume, real income and relative import prices are cointegrated. This is an important finding from the viewpoint of the Indonesian economic policy. The estimated long-run elasticity of real income and relative price are 0.98 and –0.4 respectively.

JEL Classification: C22, F14

Key words: Aggregate import demand, Bounds testing approach, Indonesia

1. INTRODUCTION

The primary objective of this paper is to ascertain the existence of a long-run relationship between Indonesia's aggregate import demand and its determinants, real income and relative import prices, using recent cointegration approach; namely, the bounds testing procedure (Pesaran et al., 2001).[1] The present study is motivated by the following considerations. Firstly, an understanding of import demand behavior

*The author would like to thank the anonymous referees' comments and suggestions on the earlier version of this paper. However, all remaining errors are entirely the author's.

of a small developing nation such as Indonesia is necessary for trade policy design. That is, devaluations have often been used by developing countries to reduce a large external imbalance (current account or trade account), correct perceived *overvaluations* of the real exchange rate, increase international competitiveness, and promote export growth (Reinhart, 1995, 290). According to Ariff and Khalid (2000, 151), Indonesia's trade structure became unfavorable as exports began to decline in mid-1995. The exchange rate overshot hugely, and the current account problem then became a nightmare in 1998 and was made worse by the fall of authoritarianism. However, the cash flows from the International Monetary Fund (IMF) rescue have been maintaining the current account since August 1998. The trade statistics for the period 1981 to 1999 (World Bank, 2002) have shown that Indonesia only enjoyed current account surpluses of US$4,096 million in 1998 and US$5,785 million in 1999, respectively. In the meantime, the changes in the Indonesian trade structure reflect a structural change of the Indonesian economy and trade policies. The share of manufactured goods in total exports recorded 3.7 percent in 1965 and it increased significantly from 13 percent in 1985 to 54.4 percent in 1999. The export of agricultural goods declined sharply from 32.2 percent in 1965 to 5 percent in 1990 and 3.8 percent in 1999. According to figures shown in Table 1, manufactured imports contributed the largest portion of total imports, followed by food and fuel. The structure of import demand indicates that Indonesia's industrialization relies highly on imported goods, especially manufactured items.

It is well-documented in trade literature that one of the major concerns in formulating a commercial policy or an exchange rate policy to correct the trade imbalance is the responsiveness of trade flows to relative prices changes. Heien (1968) stated that a price elasticity of –0.5 to –1.0 is necessary to ensure success of an exchange rate depreciation. A study by Gafar (1988) on import demand in Trinidad and Tobago found that price elasticity was in the range suggested by Heien, suggesting that exchange rate policies can be used to correct for balance of payment disequilibrium. In addition, Reinhart (1995, 291) suggested that relative prices play a significant role in the determination of trade flows, buttressing policies of devaluation as a way to correct trade imbalance, which was based on the relative price variable in static or long-run specifications of import demand or export supplies. In other words, a devaluation policy can only be accomplished if the trade flows

TABLE 1
Indonesia's Import Demand by Category (in percent), 1965 to 1998

Year	Agricultural raw materials	Food	Fuel	Ores & metals	Manufactures
1965	0.7	6.5	2.8	1.4	88.1
1970	1.2	15.4	1.6	2.1	79.5
1975	2.8	12.5	5.4	1.9	77.3
1980	3.5	12.7	16.2	2.4	64.9
1985	4.1	6.9	12.5	3.9	72.1
1990	4.7	5.1	8.9	4.3	76.9
1995	6.2	8.8	7.5	4.4	72.9
1996	5.4	10.8	8.7	3.7	71.2
1997	4.8	8.8	9.8	3.4	73.1
1998	6.7	10.5	10.0	3.3	69.1

Source: *World Tables*, World Bank (2002).

respond to relative prices in a significant and predictable manner. Another justification for implementing a favorable exchange rate policy is that if the sum of import and export demand price elasticity is greater than unity, indicating that the Marshall-Lerner condition is satisfied. Therefore, a devaluation or depreciation will have favorable effects on the external balance (Bahmani-Oskooee and Niroomand, 1998). The above discussions reveal that to rectify any trade imbalance, a knowledge of import demand behavior is essential in order to formulate both exchange rate and trade policies.

Secondly, limited empirical studies have been published on estimating aggregate import demand function for Indonesia. Among them are Reinhart (1995) and Senhadji (1998). Reinhart (1995) estimated an import demand function for Indonesia using annual data from 1970 to 1992. The results of Johansen's cointegration test indicated that there has been no cointegration relation between quantities of import demanded and its determinants; namely, real income and relative price. On the other hand, Senhadji (1998) estimated Indonesia's aggregate import demand function, using annual data from 1960 to 1993. The results of Phillips-Ouliaris (1990) residual-based cointegration test showed that the volume of imports, real income, and relative price were not cointegrated. Both of these studies used limited annual observations for conventionally used cointegration approaches, and concluded that the aggregate import demand function for Indonesia was unstable as there was no long-run equilibrium relationship among quantities of import demanded, real income and relative price terms. However, those results would be unreliable and should be interpreted with caution as Kremers et al. (1992) noted that for nonstationary data with a small sample size, no cointegration relationship can be made. Mah (2000) stated that the conventionally used cointegration methods such as residual-based (Engle and Granger, 1987) or system-based (Johansen, 1988; and Johansen and Juselius, 1990) cointegration tests are not reliable for studying small samples. Thus, this study uses a more robust cointegration method for a small sample study; that is the bounds test procedure proposed by Pesaran et al. (2001). Mah (2000, 243) recommended that Pesaran et al.'s (1996) bounds test is appropriate for small sample study.

The structure of this paper is organized as follows. The next section discusses the literature on empirical works of import demand function for selected Less Developed Countries (LDCs) and developing countries. Section 3 provides model specification, data and method. Section 4 reports the results, and the last section concludes this paper.

2. LITERATURE REVIEW

A number of studies have been conducted to empirically investigate the major determinants of import demand behavior in LDCs as well as developing countries. The conventionally used import demand function relates quantities of import demanded to real income and relative price (ratio of import prices to domestic prices). Dutta and Ahmed (1999) used Engle-Granger's (1987) and Johansen's multivariate approaches to estimate the aggregate import demand function for Bangladesh using quarterly data from 1974 to 1994. Their study found that Bangladesh's aggregate import demand and its determinants, real import prices, real gross domestic product (GDP) and real foreign exchange reserves, were cointegrated. The estimated long-run elasticities of the explanatory variables based on Engle-Granger's (1987) approach were –0.52 (for relative prices), 1.63 (for real GDP) and –0.10 (for real foreign exchange reserves, but insignificant at the 10 percent level). A dummy variable was introduced to reflect the liberalization policies, but it was found to be insignificant.

Tang and Nair (2002) re-investigated the aggregate import demand behavior for Malaysia using the bounds testing approach (Pesaran et al., 1996). The study involved annual data from 1970 to 1998 as employed by Tang and Alias (2000). The result of the bounds test indicated that volume of imports, real income and relative price were cointegrated. The estimated income and price elasticities were 1.5 and –1.3, respectively. The estimated parameter elasticities were consistent with those of Tang and Alias (2000). However, Tang and Alias (2000) found that import volume, real income and relative price were not cointegrated based on the insignificance of the estimated error correction term (see Kremers et al., 1992).

Sinha's (1997) study found one cointegrating vector in Thailand's aggregate import demand function using Johansen's multivariate procedure for the period 1953 to 1990. The study found that Thailand's aggregate import demand was price inelastic (-0.77) and cross-price inelastic (0.3) but highly income elastic (2.15). Rijal, Koshal and Jung (2000) estimated an aggregate import demand function for Nepal based on an annual series from 1968 to 1997. They applied the unit root test and Johansen-Juselius (1990) multivariate cointegration analysis. The results showed that real imports, real income, import price index and domestic price index were nonstationary but cointegrated. A partial adjustment mechanism showed that Nepalese import was inelastic with respect to its own price and cross-price, both in the short-run and the long-run. The Nepalese long-run import demand was found to be income elastic (2.13). A recent study by Tang (2002) investigated the aggregate import demand behavior in India using Johansen's multivariate cointegrating approach over an annual period from 1970 to 1999. He found that the volume of imports, real income and relative price were cointegrated. The estimated income and price elasticities were 1.4 and -0.34, respectively.

A study by Bahmani-Oskooee and Niroomand (1998) using Johansen's multivariate cointegration method found that there existed at least one cointegrating vector among the variables of Philippines' import demand function (volume of imports, relative prices and domestic income). The sample covered annual data from 1960 to 1992. The estimated long run elasticities of income and relative prices were 1.35 and -1.01, respectively. Mah (1993) estimated an import demand function for Korea using quarterly data for the period 1971 to 1988.

The study found that the estimated long-run income and price elasticities of import demand to be 0.658 and –1.029 respectively. However, acknowledging the evidence of structural instability between the import demand in the 1970s and that in the 1980s, Korean imports have become more responsive to income changes (from 0.802 to 1.246) and relative price changes (from –0.89 to –3.046).

Bahmani-Oskooee (1998) estimated the import and export demand equations for six developing countries (Greece, Korea, Pakistan, the Philippines, Singapore and South Africa) using a long-run approach (Johansen's approach). Contrary to traditional formulation, the volume of imports was related to relative prices, domestic income and nominal effective exchange rate. The sample period covered quarterly data from 1973 to 1990. The study concluded that the Marshall-Lerner condition was satisfied and revealed that devaluations could improve the countries' trade balances. The import demand for these countries was found to be price elastic except for Singapore (0.15, with incorrect sign). The real income variable was elastic, but not for Korea's case (0.31, and insignificant). The import demand for these developing countries was exchange rate inelastic (between 0.002 and 0.33) except for Singapore (-1.66, with incorrect sign).

3. MODEL SPECIFICATION, DATA AND METHODOLOGY

This study uses the traditional formulation of the aggregate import demand function as in equation (1) that relates the quantity of import demanded to domestic real income and the ratio of import prices to domestic prices (relative price term).

$$(1) \qquad M_t = f(Y_t, RP_t)$$

where M_t is the desired quantity of import demanded at period t, Y_t is real GDP, and RP_t is a relative price term that is the ratio of import price index to domestic price level. According to Doroodian et al. (1994), the log-linear formulation for the traditional import demand function is deemed to be more appropriate than the linear one. In addition, Gafar (1998) stated that the use of the log-linear specification also avoids some estimation problems, particularly multicollinearity. Thus, the log-linear specification used in this study is shown in equation (2).

(2) $$lnM_t = b_0 + b_1 lnY_t + b_2 lnRP_t + e_t$$

where e_t is white noise and normally distributed residuals, and '*ln*' indicating the natural logarithmic form. Based on economic theory, the signs of the coefficients are expected to be as follows: $b_1 > 0$ and $b_2 < 0$.

Firstly, this paper tries to include an exchange rate variable into the analysis. However, the cointegration test (bounds test procedure) fails to detect a long-run relationship in Indonesia's import demand function by including the exchange rate variable. The estimated model and results are cited in Appendix 1. According to Perman (1991, 20), cointegration analysis can be used as a form of misspecification test, or equivalently as a guide to variable selection. On the other hand, as noted by Hong (1999, 3), ". . . 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."[2] The relative price measure is often the ratio of the import price to the domestic price index for the commodity adjusted for the exchange rate, which gives a measure of the real exchange rate. Reinhart (1995, 297) noted that a scale variable (permanent income or wealth) and relative prices are both necessary and sufficient to define the long run behavior of imports, and this would argue against the inclusion of any other variable in an *ad hoc* manner. Based on the above justifications, this study uses the traditional import demand function as stated in equation (1).

The sample used in the present study covers annual observations from 1960 to 1999. The raw data are obtained from *World Tables* (World Bank, 2002). The quantity of import demanded is nominal imports of goods and services deflated by import price index. An income variable is proxied by real GDP that is nominal GDP deflated by the GDP deflator. The relative import price variable is the ratio of import price to domestic price (GDP deflator). All implicit price indexes are based on 1995 prices.

Here, the order of integration, *I*(d) for the variables (import volume, real income and relative price) must be known before applying cointegration analysis, like the bounds testing procedure, to ensure that no variable is integrated of order greater than one (see Abbott et al., 2001, 48). The Phillip-Perron (1988) unit root test (PP) was employed and was designed to be robust for the presence of autocorrelation and heteroscedasticity. The results of the PP unit root test are shown in Table 2. Contrary to previous studies (Reinhart, 1995; and Senhadji, 1998), the relative price series (*RP*) is stationary, *I*(0), at the one percent significance level.[3] The volume of import and real income are nonstationary, *I*(1). As noted, including a stationary, *I*(0), regressor series might yield biased results when using conventionally-used cointegration approaches like Engle-Granger (1987) and Johansen's multivariate (Johansen, 1988; and Johansen and Juselius, 1990) approaches. These methods are applicable for series which have the same degree of integration, *I*(1). Enders (1995, 396) stressed that although forms of the Johansen tests can detect differing orders of integration, it is wise *not* to mix variables with different orders of integration. This might be a possible explanation as to why previous studies like Reinhart (1995) and Senhadji (1998) failed to detect a long-run relationship of import demand function in Indonesia. This study uses the bounds test procedure proposed by Pesaran et al. (2001) which provides a feasible application on mixed *I*(0) and *I*(1) regressors.

TABLE 2
Phillip-Perron (PP) Unit Root Test

Variable	Level	First Difference	Conclusion, *I*(d)
ln M	-1.251210 (3)	-3.701046 (3)*	*I*(1)
ln Y	-2.185765 (3)	-4.004826 (3)*	*I*(1)
ln RP	-4.701557 (3)*	-	*I*(0)
MacKinnon's (1991) Critical Values			
*1%	-4.2092	-3.6117	
**5%	-3.5279	-2.9399	
***10%	-3.1948	-2.6080	

Note: The unit root equation includes the time trend for level analysis, but was excluded in first difference analysis. The reported critical values (MacKinnon, 1991) permit the tests for any sample size and for any number of right-hand variables. ( ) is the truncation lag of Bartlett Kernel based on Newey-West suggestion. In order to test its sensitivity, truncation lags of one, two and three were included into the unit root equation. The results are consistent with the three-truncation lag.

The Pesaran et al.'s approach has two major advantages over the conventionally used cointegration methods (Engle-Granger, 1987;

Johansen, 1988; and Johansen and Juselius, 1990). First, the bounds test procedure is robust for small sample size. Mah (2000) stated that the ECM, Johansen (1988) and Johansen and Juselius (1990) methods are not reliable for studies that have small samples. Mah (2000) used Pesaran et al's approach for examining disaggregated import demand (information technology products) for Korea with 18 annual observations. Other examples are from Pattichis (1999) and Tang and Nair (2002). Secondly, the bounds test procedure can be applied irrespective of whether the explanatory variables are $I(0)$ or $I(1)$ as in the present study. An application of bounds test procedure on this issue can be found in Abbott et al. (2001).[4]

In order to investigate the presence of a long-run equilibrium relationship among the variables in aggregate import demand equation (2), an unrestricted error correction model (UECM) as in (3) can be estimated for the bounds test procedure. The ordinary least squares (OLS) method is used in the estimation.

$$(3) \qquad \Delta lnM_t = b_0 + \sum_{i-0}^{n} b_{1i} \Delta lnY_{t-i} = \sum_{i-0}^{n} b_{2i} \Delta lnRP_{t-i} + \sum_{i=1}^{n} b_{3i} \Delta lnM_{t-i} + b_4 lnM_{t-1} + b_5 lnY_{t-1} + b_6 lnRP_{t-1} + e_t$$

where *DlnM*, *DlnY*, and *DlnRP* are first differences of the logarithms of quantity of import demand (*lnM*), real GDP (*lnY*), and relative price (*lnRP*), respectively.

Pesaran et al. (2001) proposed that the bounds test is based on the Wald or *F*-statistics for cointegration analysis. The asymptotic distribution of the *F*-statistics 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 test is conducted in the following way. The null hypothesis considers the UECM (3) by excluding the lagged level variables, lnM_{t-1}, lnY_{t-1} and $lnRP_{t-1}$. More formally, a joint significance test is performed. The null and alternative hypotheses are, H_0: $b_4 = b_5 = b_6 = 0$, and H_A: b_4^1 0, b_5^1 0, b_6^1 0.

For conventional significance levels of 1, 5, and 10 percent, if the *F*-statistic falls outside the critical value, a conclusive inference can be made without considering the order of integration of the explanatory variables. If the *F*-statistic is higher than the critical bound, then the null hypothesis of no cointegration can be rejected. In the case when the *F*-statistic falls between the upper and lower bounds, a conclusive inference cannot be made. Here, the order of integration for the explanatory variables must be known before any conclusion can be drawn. From the estimated UECM, the long-run elasticities are the coefficients of the one lagged explanatory variables (multiplied with a negative sign) divided by the coefficient of the one lagged dependent variable. Thus, the long-run relative price elasticity and income elasticities are - (b_6 / b_4) and - (b_5 / b_4), respectively. The short-run effects are captured by the coefficients of the first difference variables in (3).

4. THE RESULTS

In selecting an appropriate lag length for the UECM, a set of UECMs as in equation (3) was estimated using three, two and one lag length on its first difference variables. As a rule of thumb, the maximum lag length for annual data is three (Charemza and Deadman, 1992). The three lag lengths were selected in order to minimize the Akaike information criterion (AIC); and the residuals are white noise (based on the *Q*-statistics) and normally distributed (based on the Jarque-Bera test). Further, a parsimonious specification of the UECM was arrived at by adopting a model selection procedure, which moves from *general to specific* (see Pattichis, 1999, 1063). That is, all those first difference variables that have relatively small absolute *t*-ratio (less than one) were dropped sequentially.

A parsimonious specification of the UECM is presented in equation (4). The lagged one, two and three period first difference real import variables ($DlnM_{t-1}$, $DlnM_{t-2}$, and $DlnM_{t-3}$) were dropped during the *general to specific* process. This implies that growth of Indonesia's import demand does not depend on its own past growth.[5] The estimated preferred UECM passes a battery of diagnostic tests. The Breusch-Godfrey LM test shows no evidence of residual serial correlation up to third order. No evidence of autoregressive conditional heteroskedasticity (ARCH) was also found in the residual based on the ARCH LM test. The Jarque-Bera test confirms that the estimated residual is normally distributed. Ramsey's RESET test confirms no functional form misspecification. The plots of CUSUM of Squares and CUSUM (see Figure 1) reveal that the estimated parameters of equation (4) are stable over the sample period.

$$\Delta lnM_t = \quad -7.03 + 0.423\Delta lnY_t + 2.06\Delta lnY_{t-1} + 2.045\Delta lnY_{t-2}$$
(*t*-ratios) (-2.149) **(1.13) (5.54)* (2.40)**

$$+ 2.057\Delta lnM_{t-3}$$
(2.80)*

(4) $$+ 0.199\Delta lnRP_{t-1} + 0.141\Delta lnRP_{t-2} + 0.101\Delta lnRP_{t-3}$$
(1.75)*** (1.33) (1.44)

$$-\ 0.256\Delta lnM_{t-1} + 0.251\Delta lnY_{t-1} - 0.103\Delta lnRP_{t-1}$$
(-3.01)* (2.20)** (-0.998)

Note: *, **, and *** denote significance at the 1%, 5% and 10% levels.

Sample (adjusted): 1964-1999

R-squared: 0.79 Adjusted *R*-squared: 0.707

Sum of squared residuals: 0.170 *F*-statistic (*p*-value): 9.432 (0.000)

Q-Statistics [20]: White noise Jarque-Bera: 0.646 (0.724)

Breusch-Godfrey Serial Correlation LM Test [1]:0.05 (0.822); [2]:1.3 (0.522); [3]:5.42 (0.144)

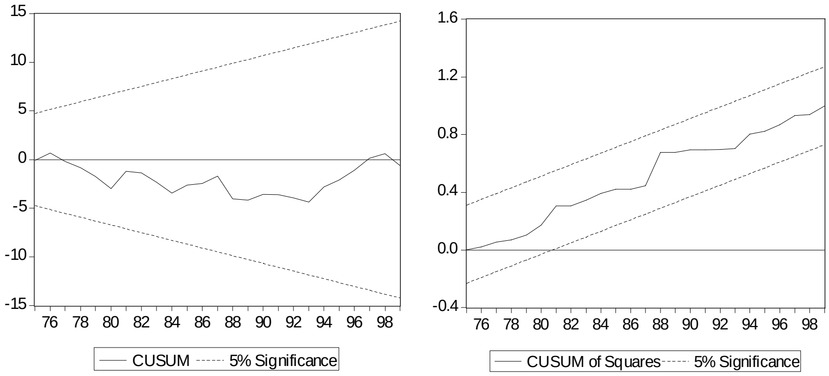
Ramsey RESET Test [1]: 3.565 (0.06); [2]:4.103 (0.129); [3]: 4.104 (0.25)

ARCH Test [1]: 0.315 (0.575); [2]: 4.073 (0.131); [3]: 4.94 (0.176)

In order to ascertain the existence of a long-run relationship among

FIGURE 1

Plots of CUSUM and CUSUM of Squares



the variables in model (2), the *F*-statistic (Wald test) for the bounds test (Pesaran et al., 2001) was computed. The *F*-statistic and critical bounds values for testing the null of no cointegrating relationship are reported in Table 3. The computed *F*-statistic of 8.228 was found to exceed the upper bounds critical value, 4.85 at the 5 percent significance level. Therefore, the null of no cointegration is rejected. The result is unlike the findings from previous studies that failed to reject the null of no long-run relationship among the volume of import, real income and relative price. The finding implies that the volume of imports and its determinants, namely real income and relative price, are cointegrated. This indicates that Indonesia's import demand function is stable over the sample period.

TABLE 3
Bounds Test for Cointegration Analysis

F-Statistic (Wald test): 8.228 (for $H_0: b_4 = b_5 = b_6 = 0$)			
Critical Bounds[#] at:	1%	5%	10%
Lower bounds, $I(0)$:	4.41	3.79	3.17
Upper bounds, $I(1)$:	5.52	4.85	4.14

[#]Pesaran et al. (2001, 300) Table CI(iii) case III: unrestricted intercept and no trend (k=2).

The estimated long-run elasticities for income and relative price are 0.98 and -0.402, respectively. Both of the estimated elasticities have the expected signs. However, the estimated long-run elasticity for relative price variable ($lnRP_{t-1}$) is insignificant at the 10 percent level. The insignificant relative import price variable is maintained in the import demand analysis in this study, because of its relevant policy implication especially in favoring a devaluation to improve trade balance. The short-run price and income elasticities are 0.199[6] and 2.06, respectively.

5. CONCLUSION

The present study aims at empirically re-investigating the long-run relationship of Indonesia's import demand function using the bounds testing approach (Pesaran et al., 2001). The sample period covers annual data from 1960 to 1999. In contrast to previous studies (Reinhart, 1995; and Senhadji, 1998), the results of the bounds test suggest a cointegrating relationship among volume of imports, real income and relative import prices. Thus, the results indicate that Indonesia's aggregate import demand function is stable. Therefore, some relevant policy implications can be discussed.

In the long-run, the estimated relative price elasticity is –0.4, but insignificant at the 10 percent level, implying that the Indonesian import demand is insensitive to increases in domestic price levels. Its estimated that elasticity falls outside of Heien's (1968) range of -0.5 to –1.0, suggesting that the exchange rate policy is inappropriate to correct for a balance of payment deficit. On the other hand, the estimated long-run relative price for Indonesian export demand function was inelastic, -0.015 (Reinhart, 1995). Meanwhile, the results in this study have been found to be insufficient to satisfy the Marshall-Lerner condition for favoring a devaluation policy to correct Indonesia's external imbalance. This finding is consistent with Tang's (2003) study where the volume of imports and exports for Indonesia are not cointegrated, indicating that exchange rate policy is inappropriate to rectify the country's trade imbalance. In addition, the relative price elasticity is well below unity, suggesting that large relative price swings are necessary to produce an appreciable reallocation of trade flows (Reinhart, 1995, 308). The estimated long-run income elasticity is closer to unity (0.98), and does suggest that Indonesia's import demand is significantly driven by economic growth. If imports are biased towards imports of consumption goods, *ceteris paribus*, the country may face problems in the balance of payments in the longer run. Thus, policy designs in influencing the pattern of various final expenditures (real income) would seem most effective. The present study suggests that an effective domestic demand management is necessary to be viewed as part of a comprehensive trade stabilization plan for Indonesia.

ENDNOTES

1. Indonesia is a member of the Organization of Islamic Conferences (OIC).

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

3. Senhadji (1998, 247) found that the relative price (*RP*) variable for Indonesia was nonstationary using the augmented Dickey-Fuller test – ADF (Dickey and Fuller, 1979) test over an annual period from 1960 to 1993. In this study, the *RP* variable was re-tested based on the ADF test for annual data from 1960 to 1999. An optimal lag for the ADF test is one that minimizes the Akaike information criterion (AIC). The computed ADF statistics for *RP* are –2.8282 (in level) and –2.917 (in first difference). The 10 percent MacKinnon critical values are –3.1968 (in level) and –2.608 (in first difference). The result of the ADF test shows that *RP* is of *I*(1) process. The ADF results are not sensitive with two, three, and four augmented lag lengths. As stated by Obben (1998, 114), where there is inconsistency between the ADF results and the PP results, the conclusion from the PP test is preferred. The rationale is that the DF or ADF test is based on the assumption that the series is generated by an autoregressive (AR) process whilst the PP test is based on the more general autoregressive integrated moving-average (ARIMA) process. Hallam and Zanoli (1993, 160) stated that the PP test is more powerful over the DF test especially for small samples.

4. The existing studies used bound testing approach from Pesaran et al. (1996) (unpublished version).

5. This is not an unusual result as Pattichis (1999) encountered a similar issue, i.e., the lagged first difference real import variable was dropped in estimating a parsimonious UECM for Cyprus's disaggregated import demand function using Pesaran et al.'s approach (1996). See Pattichis (1999), Table 2 (p. 1065), Table 3 (p. 1066), and Table 4 (p. 1067). Tang and Nair (2002) found that the one period lagged import growth was insignificant at the 10 percent level for Malaysian import demand (see Table 1, UECM, p. 295). Mohammad and Tang (2000) estimated an error correction model (Engle and Granger, 1987) for Malaysian aggregate import demand and found that the one period lagged import growth was dropped from the general to specific exercise (equation 6, p. 264). However, in this study, one lagged period real imports, lnM_{t-1} influences its current import growth, $\Delta lnM_{1.}$ The estimated coefficient of lnM_{t-1} is –0.256. This is the speed of adjustment, or the coefficient of error-correction term as defined by Engle and Granger (1987). In particular, the coefficient of –0.256 indicates that 25.6 percent of the disequilibrium is corrected by the volume of imports in the last year towards a long-run equilibrium state.

6. The estimated short-run relative import price variable is insignificant at

the 5 percent level. Abbott and Seddighi (1996) estimated an aggregate import demand behavior for the UK. The estimated error correction model, ECM (Engle and Granger, 1987) showed the short-run elasticities of the relative price variable having positive signs (equation 6). In addition, Mohammad and Tang (2000) estimated the ECM for the Malaysian aggregate import demand and found that the short-run relative price elasticity showed a positive sign, 0.68 (equation, p. 264). Both ECMs passed a battery of diagnostic tests and there was no further explanation on the estimated positive sign of relative price elasticity in the short-run. Other empirical studies examined import demand behavior in the long-run, due to its relevance in improving trade balance. Among them are Reinhart (1995); Bahmani-Oskooee and Niroomand (1998); and Bahmani-Oskooee (1998).

REFERENCES

Abbott, A., A. C. Darnell, and L. Evans. "The Influence of Exchange Rate Variability on UK Exports." *Applied Economic Letters* 8 (2000): 47-9.

Abbott, A. J., and H. R. Seddighi. "Aggregate Imports and Expenditure Components in the UK: An Empirical Analysis." *Applied Economics* 28 (1996): 1119-25.

Alias, H. M., and T. C. Tang. "Aggregate Imports and Expenditure Components in Malaysia: A Cointegration and Error Corretion Analysis." *ASEAN Economic Bulletin* 17, no. 3 (2000): 257-69.

Ansari, M. I. "Testing the Relationship Between Government Expenditure and National Income in Canada, Employing Granger Causality and Cointegration Analysis." *Managerial Finance* 19, no. 7 (1993): 31-47.

Ariff, M., and A. M. Khalid. "Indonesia: Liberalization Amidst Persistent Exchange Rate Instability." Chap. 6 in *Liberalization, Growth and the Asian Financial Crisis*. UK: Edward Elgar, 2000: 147-71.

Arize, A. C., T. Osang., and D. J. Slottje. "Exchange-Rate Volatility and Foreign Trade: Evidence From Thirteen LCDs." *Journal of Business and Economic Statistics* 18, no. 1 (2000): 10-7.

Bahmani-Oskooee, M. "Cointegration Approach to Estimate the Long-Run Trade Elasticities in LDCs." *International Economic Journal* 12, no. 3 (1998): 89-96.

_______, and F. Niroomand. "Long-Run Price Elasticities and the Marshall-Lerner Condition Revisited." *Economic Letters* 61 (1998):

101-9.

Charemza, W. W., and D. F. Deadman. *New Directions in Econometric Practice.* England: Edward Elgar Publishing Ltd., 1992.

Dickey, D. A., and W. A. Fuller. "Distribution of the Estimators for Autoregressive Time Series with a Unit Root." *Journal of the American Statistical Association* 74 (1979): 427–31.

Doroodian, K., R. K. Koshal, and S. Al-Muhanna. "An Examination of the Traditional Aggregate Import Demand Function for Saudi Arabia." *Applied Economics* 26 (1994): 909-15.

Dutta, D., and N. Ahmed. "An Aggregate Import Demand Function for Bangladesh: A Cointegration Approach." *Applied Economics* 31 (1999): 465-72.

Enders, W. *Applied Econometric Time Series*. USA: John Wiley & Sons, Inc., 1995.

Engle, R. F., and C. W. J. Granger. "Co-Integration and Error Correction: Representation, Estimation and Testing." *Econometrica* 55, no. 2 (1987): 251-76.

Gafar, J. S. "The Determinants of Import Demand in Trinidad and Tobago: 1967-84." *Applied Economics* 20 (1988): 303-13.

Hallam, D., and R. Zanoli. "Error Correction Models and Agricultural Supply Response." *European Review of Agricultural Economics* 20, no. 2 (1993): 151-66.

Heien, D. M. "Structural Stability and the Estimation of International Import Price Elasticities in World Trade." *Kyklos* 21 (1968): 695-711.

Hong, Pingfan. *Import Elasticities Revisited*. Department of Economic and Social Affairs Discussion Paper No. 10, United Nations, 1999 (http://www.un.org/esa/papers.htm).

Johansen, S. "Statistical Analysis of Co-Integrating Vectors." *Journal of Economic Dynamics and Control* 12 (1988): 231-54.

_______. and K. Juselius. "Maximum Likelihood Estimation and Inference on Cointegration with Applications to the Demand for Money." *Oxford Bulletin of Economics and Statistics* 52 (1990): 169-210.

Kremers, J. J. M., N. L. Ericsson, and J. Dolado. "The Power of Cointegration Tests." *Journal of Econometrics* 52 (1992): 389-402.

Mah, J S. "Structural Change in Import Demand Behavior: The Korean Experience." *Journal of Policy Modeling* 15, no. 2 (1993): 223-7.

_______. "An Empirical Examination of the Disaggregated Import Demand of Korea–The Case of Information Technology Products."

Journal of Asian Economics 11 (2000): 237-44.

MacKinnon, J. G. “Critical Values for Cointegration Tests.” Chap. 13 in *Long-run Economic Relationship: Readings in Cointegration*, edited by R.F. Engle and C.W.J. Granger, Oxford University Press, 1991.

Obben, J. “The Demand for Money in Brunei.” *Asian Economic Journal* 12, no. 2 (1998): 109-21.

Pattichis, C. A. “Price and Income Elasticities of Disaggregated Import Demand: Results from UECMs and an Application.” *Applied Economics* 31 (1999): 1061-71.

Pesaran, H., Y. Shin, and R. J. Smith. *Testing for the Existence of a Long-Run Relationship*. Department of Applied Economics Working Paper No. 9622, University of Cambridge, 1996.

______. “Bounds Testing Approaches to the Analysis of Level Relationships.” *Journal of Applied Econometrics* 16 (2001): 289-326.

Perman, R. “Cointegration: An Introduction to the Literature.” *Journal of Economic Studies* 18, no. 3 (1991): 3-30.

Phillips, P. C. B., and S. Ouliaris. “Asymptotic Properties of Residual-Based Tests for Cointegration.” *Econometrica* 58 (1990): 165-93.

Phillips, P. C. B., and P. Perron. “Testing for a Unit Root in Time Series Regression.” *Biometrika* 75, no. 2 (1988): 335-46.

Reinhart, C. M. “Devaluation, Relative Prices, and International Trade: Evidence from Developing Countries.” *IMF Staff Papers* 42, no. 2 (1995): 290-312.

Rijal, A., R. K. Koshal, and C. Jung. “Determinants of Nepalese Imports.” *Journal of Asian Economics* 11 (2000): 347-54.

Sinha, D. “Determinants of Import Demand in Thailand.” *International Economic Journal* 11, no. 4 (1997): 73-83.

Senhadji, A. “Time-Series Estimation of Structural Import Demand Equations: A Cross-Country Analysis.” *IMF Staff Papers* 45, no. 2 (1998): 236-68.

Tang, T. C. “Aggregate Import Demand in India: Stable or Unstable?” *The Economic Challenger* 14, no. 4 (2002): 35-7.

______. “Are the Imports and Exports of the Five ASEAN Economies Co-Integrated: An Empirical Study.” *International Journal of Management* 20, no. 1 (2003) (forthcoming).

______, and M. Nair. “A Cointegration Analysis of Malaysian Import Demand Function: Reassessment from the Bounds Test.” *Applied Economic Letters* 9, no. 5 (2002): 293-6.

______, and H. M. Alias. "An Aggregate Import Demand Function for Malaysia: A Cointegration and Error Correction Analysis." *Utara Management Review* 1, no. 1 (2000): 43-57.

World Bank. "World Tables." *dx for Windows version 3.0*, The Time Series Data Express, EconData Pty. Ltd., 2000 (Released data: February 28, 2002).

APPENDIX 1

ESTIMATED IMPORT DEMAND MODEL WITH EXCHANGE RATE VARIABLE

A log-linear model for the import demand function which includes the exchange rate variable (*ER*) as an explanatory variable is specified as $lnM_t = a + b_1 lnY_t + b_2 ln RP_t + b_3 ln ER_t + e_t$. The nominal exchange rate variable is measured in local currency per \$US. The sample period is from 1967 to 1999. The Phillip-Perron unit root test finds *lnER* to be nonstationary, *I*(1). The result is consistent based on 1, 2, 3, and 4 truncation lag specifications. Further, a general UECM with 2 lag length of first difference variables was estimated considering residuals are white noise (*Q*-statistics) and normally distributed (with the highest *p*-value of the Jarque-Bera test). In selecting a parsimonious specification, all those first difference regressors that have relatively small absolute *t*-ratios (less than one) were dropped sequentially. The estimated preferred UECM is as follows:

$$\begin{array}{lllll}
\Delta lnM_t = & -1.33 & + 0.96\Delta lnY_t & - 0.78\Delta lnRP_t & + 0.25\Delta lnER_t \\
\text{(t-ratios)} & (-2.53)^{***} & (1.39) & (-3.89)^{*} & (1.69)
\end{array}$$

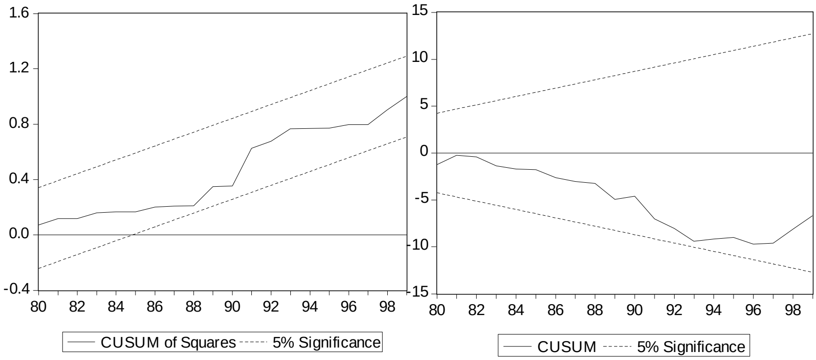
$$\begin{array}{l}
-2.057\Delta lnER_{t-1} \\
(1.95)^{***}
\end{array}$$

$$\begin{array}{llll}
+ 0.18\Delta lnM_{t-1} & + 0.227\Delta lnM_{t-2} & - 0.45lnM_{t-1} & + 0.544lnY_{t-1} \\
(1.3) & (1.66) & (-2.94)^{*} & (2.18)^{**}
\end{array}$$

$$\begin{array}{ll}
- 0.48lnRP_{t-1} & + 0.019lnER_{t-1} \\
(-1.98)^{***} & (0.162)
\end{array}$$

Note: *, **, and *** denote significance at the 1%, 5% and 10% levels.
Sample (adjusted): 1969-1999 *Q*-Statistics [20]: White noise

R-squared: 0.887 Adjusted *R*-squared: 0.831
Sum of squared residuals: 0.083 *F*-statistic: 15.76 (0.000)
Jarque-Bera: 3.155 (0.206)
Ramsey RESET Test [1]: 2.88 (0.09); [2]: 3.03 (0.22); [3]: 3.73 (0.29)
Breusch-Godfrey serial correlation LM Test [1]: 1.75 (0.186); [2]: 2.51 (0.28); [3]: 3.1 (0.38)
ARCH Test [1]: 1.75 (0.186); [2]: 2.51 (0.29); [3]: 3.1 (0.38), where ( ) is the *p*-value.



The computed *F*-statistic (bounds test) for the null of no cointegration is 3.293. It falls between the upper and lower critical bounds at the 10 percent (2.72 and 3.77) and 5 percent levels (3.23 and 4.35). Therefore, the hypothesis testing is found to be inconclusive at the 10 percent and 5 percent significance levels. However, the null hypothesis cannot be rejected at the one percent level, since the *F*-statistic is lower than the lower critical bounds' values (4.29 and 5.61).