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Long-Run Aggregate Import Demand Function in Taiwan: Evidence from the ARDL Model

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Abstract

This paper adopts the bounds test, developed by Pesaran *et al.* (2001), to determine whether there is a level long-run relationship exists between Taiwan's real import demand function and it determinants, namely real domestic income and relative prices. It is found that aggregate import quantities and their determinants, real domestic income and relative prices do indeed exhibit a level long-run relationship. The estimated equilibrium correction coefficient shows only a moderate speed of adjustment from disequilibrium to equilibrium. In addition, the empirical results show that estimated short-run elasticity and long-run income elasticity are both elastic but that short-run income elasticity is considerably greater than that of its long-run counterpart. This indicates that economic growth should have a relatively greater negative impact on trade balance in the short-run than in the long-run. Finally, with regard to estimated relative price elasticity, it is insignificantly different from zero both in the short-run and in the long-run. This suggests that import volume is insensitive to any devaluation policies.

Keywords: Aggregate import demand, Bounds test, Autoregressive distributed lag model

JEL classification: C50, F10

1 Introduction

The huge trade surplus in Taiwan has long attracted both international praise and attention, but regrettably, has oftentimes let to trade friction with her trading partners, such as the United States. From 1989 to 2003, in fact, Taiwan was on the "U.S.'s Special 301" list of the United States year-after-year, and even more serious, Taiwan was on the priority watch list in 1989, 1992, 1993 and from 2001–2003. Annual data for the 1981–2002 period show that Taiwan has enjoyed an rewarding surplus of trade (see Table 1). As shown there, especially in the 1986–1991 period trade was over 100,000 million U.S. dollars. In fact, the volume of Taiwan's imports has been escalating over the last two decades, and at the same time, the volume of exports has also been swelling at an almost equally steady pace during the same period. It is true that the latter has been slightly greater than the former, thus perpetuating Taiwan's trade surplus, and this has given rise to criticism often sever. The end result has been that Taiwan's trading partners have strongly urged it to reduce its trade surplus by opening up its domestic markets thereby sparking imports. However, one question arises: Would it really be effective to reduce its trade surplus by stimulating its domestic import demand from other countries? Theoretically, import demand is dependent upon many factors, such as real domestic income, and differences in domestic and import prices. With this in mind, this paper aims at empirically analyzing the relationship between Taiwan's import demand function and its determinants.[1]

Empirical investigations of the import demand function have been at the core of numerous research studies in international economics with many economists devoting considerable time and effort estimating the aggregate import demand function for different countries. Notable ex-

[1] See Wang (1998) for a brief review on Taiwan's economic development and foreign trade.

amples include the recent studies by Clardia (1994) and Deyak, Sawyer and Sprinkle (1993) who derived the import demand functions for the U.S. and Canada. Mah (1994), Bahmani-Oskoee and Niroomand (1998), Masih and Masih (2000), Hamori and Matsubayashi (2001) and Tang (2003) directed their attention towards Japan by estimating its import demand functions. By the same token, Bahmani-Oskooee (1997) and Mah (1992, 1993, 1997) focused on the aggregate import demands of Korea, while Tang (2002) estimated the aggregate import demand functions of Hong Kong. For Malaysia, see Tang and Alias (2000) and Tang and Nair (2002); for India, see Tang (2002) and Dutta and Ahmed (2002).[2] Going one step further, Reinhart (1995), Bahmani-Oskooee and Niroomand (1998) and Senhadji (1998) performed cross-countries analyses.

One key feature of above the studies centers on time series econometric methodology that tests the unit root and the concept of cointegration, as championed by Engle and Granger (1987), in order to analyze the long-run relationship between aggregate import demand and its explanatory variables since cointegration implies that a long-run relationship exists between regressors. When data are nonstationary, any inferences based on standard ordinary least square (OLS) results are invalidated and suffer from the problem of "spurious regrssion" (Granger and Newbold, 1974). The destination feature about the cointegration technique is that it offers a novel approach that yields empirical results that are not in any way spurious.

Engle and Granger's (1987) residual-based, two-step approach and Johansen's (1988), Johansen and Juselius's (1990) maximum likelihood method have quite often been used in estimations. Johansen's procedure does, nevertheless have several advantages over the Engle-Granger residual-based, two-step approach. The Engle-Granger procedure, for example, is highly sensitive to the

[2] There are some papers which focus on the disaggregate import demand function. Readers are referred to Deyak, Sawyer and Sprinkle (1989), Pattichis (1999) and Mah (2000), for example.

choice of the dependent variable in the cointegration regression, whereas Johansen's procedure safely assumes that all variables are endogenous.[3] However, problems inherent to both methodologies cannot be ignored. Perhaps, the most salient of these is small sample bias. In this regard, Kremers *et al.* (1992) noted that for data from a small sized sample, no cointegration relation can be determined among variables that are nonstationary. Cheung and Lai (1993) also claimed that with a finite-sample there is a tendency for Johansen's likelihood ratio test to be biased toward finding cointegration either too often or too infrequently. As shown by Shiller and Perron (1985), Hakkio and Rush (1991) and Otero and Smith (2000), it is the span of the data, and not their frequency, that determines the power of a unit root and cointegration tests. Beyond that, both methodologies require that regressors be of the $I(1)$ process and they require a unit root test for each empirical series prior to any cointegration analysis. These two shortcomings restrict empirical analysis if the data span is limited or if the variables have a different integration of orders.

Pesaran *et al.*(2001) have recently developed the bounds test procedure that overcomes the above problems. This procedure is based on the estimation of the AutoreRressive Distributed Lag (ARDL) model, and to be sure, it outperforms other estimators in small samples (see Pesaran and Shin, 1995). Besides this, when written in the Error Correction Model (ECM) form, the ARDL model is much less vulnerable to spurious regression (Pesaran and Smith, 1998). Here we use the Pesaran *et al.*(2001) bounds test approach to test for the presence of a long-run relationship between the aggregate import demand and its explanatory variables, namely, real income and relative prices in Taiwan. The bounds test procedure is applicable irrespective of whether or not the underlying regressors are integrated on the order of one or zero, or are mutually cointegrated. By contrasts, the ARDL regression yields a test statistic which can be compared to two asymptotic

[3] Masih and Masih (2000) listed the advantages of Johansen's approach over the Engle-Granger two-step approach. .

critical values. If the test statistic is above a certain upper critical value, the null hypothesis of a no long-run relationship must be rejected regardless of whether the underlying orders of integration of the regressors are zero or one. Alternatively, when the test statistic falls below a certain lower critical value, the null hypothesis of a no long-run relationship between the regressors is accepted. If the test statistic falls between these two bounds, the results are, in a word, inconclusive.

Using this approach, we find that there is a level long-run relationship exists between the aggregate real import quantities and their determinants, the real domestic income and relative prices. The estimated equilibrium correction coefficient shows a moderate speed of adjustment from disequilibrium to equilibrium. In addition, the empirical results show that both the estimated short-run and long-run income elasticity are indeed elastic but that short-run income elasticity is greater than its long-run counterpart. What the implication here is that that economic growth does have a greater negative impact on the trade balance in the short-run than in the long-run. Finally, with regard to the estimated relative price elasticity, it is insignificantly different from zero in either the short-run or the long-run. This asserts that import volume is insensitive to any devaluation policies.

The organization of the paper is as follows. Section 2 introduces the econometric methodology that we employ. Section 3 describe the methodology, the data and the empirical test results. Section 4 employs Johansen's (1988) method to check the robustness of the ARDL model. Section 5 presents the conclusions that we draw from this research.

2 Model Specification

The estimation of the aggregate import demand function is usually based on conventional demand theory which indicates that the quantity of the import demand is a function of real domestic income and relative prices (by imposing the condition of homogeneity).[4] In this study, the quantity of real import demand, M_t is considered as a function of the real domestic income, Y_t, and relative prices, RP_t, (the ratio of import prices to domestic prices). The approach of the bounds test is to determine whether there exists a single long-run relationship between the natural logarithm of the desired quantity of import demand, $m_t = \ln M_t$, and $\boldsymbol{x}_t$, where $\boldsymbol{x}_t$ is the vector time series $\boldsymbol{x}_t = \{\ln Y_t,\ \ln RP_t\}$, the method begins with an unrestricted vector autoregression:

$$\boldsymbol{z}_t = \boldsymbol{\mu} + \sum_{j=1}^{p} \boldsymbol{\phi}_j \boldsymbol{z}_t + \boldsymbol{\varepsilon}_t, \tag{1}$$

where $\boldsymbol{z}_t = [m_t\ \boldsymbol{x}_t]'$; $\boldsymbol{\mu}$ is a vector of constant terms, $\boldsymbol{\mu} = [\mu_m\ \boldsymbol{\mu_x}]'$; and $\boldsymbol{\phi}_j$ is a matrix of the Vector AutoRegressive (VAR) parameters for lag j. As noted by Pesaran *et al.* (2001), the two series m_t and $\boldsymbol{x}_t$ can be either $I(0)$ or $I(1)$. In the case where $\boldsymbol{x}_t$ is a vector time series, real domestic income and relative prices can also be of different orders of integration.[5]

The vector of the error terms $\boldsymbol{\varepsilon}_t = [\varepsilon_{m,t}\ \boldsymbol{\varepsilon}_{x,t}]' \sim N(\mathbf{0},\ \Omega)$, where Ω is a positive definite and is given by:

$$\Omega = \begin{bmatrix} \omega_{mm} & \omega_{mx} \\ \omega_{xm} & \omega_{\boldsymbol{xx}} \end{bmatrix}. \tag{2}$$

Given this, $\varepsilon_{m,t}$ can be expressed in terms of $\boldsymbol{\varepsilon}_{x,t}$ as:

$$\varepsilon_{m,t} = \boldsymbol{\omega} \boldsymbol{\varepsilon}_{x,t} + u_t, \tag{3}$$

where $\boldsymbol{\omega} = \boldsymbol{\omega}_{mx} / \boldsymbol{\omega}_{xx}$ and $u_t \sim N(0, \omega_{mm} - \boldsymbol{\omega}_{mx} \boldsymbol{\omega}_{xx}{}^{-1} \boldsymbol{\omega}_{xm})$.

[4] As noted by Hong (1999, p. 3), "...import demand in a market economy can be fully modeled by two determinants: income and relative prices. The other factors can all be subsumed within these two factors, at least theoretically."

[5] The exposition of the ARDL model in this section is basically based on Pesaran *et al.* (2001), Coe and Serletis (2002) and Atkins and Coe (2002). Readers are referred to their papers for details.

By Performing Eq. (1), this can be written as a vector error correction model as follows:

$$\Delta \boldsymbol{z}_t = \boldsymbol{\mu} + \boldsymbol{\lambda} \boldsymbol{z}_{t-1} + \sum_{j=1}^{p-1} \gamma_j \Delta \boldsymbol{z}_t + \boldsymbol{\varepsilon}_t, \tag{4}$$

where $\Delta = 1 - L$, and

$$\gamma_j = \begin{bmatrix} \gamma_{mm,j} & \gamma_{mx,j} \\ \gamma_{xm,j} & \gamma_{xx,j} \end{bmatrix} = -\sum_{k=j+1}^{p} \boldsymbol{\phi}_k. \tag{5}$$

Here, $\boldsymbol{\lambda}$ is the long-run multiplier matrix and is given by:

$$\boldsymbol{\lambda}_j = \begin{bmatrix} \lambda_{mm} & \lambda_{mx} \\ \lambda_{xm} & \lambda_{xx} \end{bmatrix} = -(\boldsymbol{I} - \sum_{j=1}^{p} \boldsymbol{\phi}_j), \tag{6}$$

where $\boldsymbol{I}$ is an identity matrix. The diagonal elements of this matrix are left unrestricted. This allows for the possibility that each of the series can be either $I(0)$ or $I(1)$. For example, $\lambda_{mm} = 0$ indicates that the quantity of import demand is $I(1)$, while $\lambda_{mm} < 0$ shows that it is $I(0)$ (see Pesaran *et al.*, 2001, p294).

This procedure enables us to test whether there exists a maximum of one long-run relationship which includes both m_t and $\boldsymbol{x}_t$, this would determine that either λ_{mx} or λ_{xm} can be non-zero, but certainly not them both. As our interest is on the long-run effect of the level of real domestic income and relative prices on the aggregate import demand, the restriction $\lambda_{xm} = 0$ is imposed, which implies that both the level of real domestic income and relative price have no long-run impact on the aggregate import demand or that the level of real domestic income and relative prices are, in the terminology of Pesaran *et al.* (2001), *long-run forcing* for aggregate import demand. Well worth noting is that this does not preclude the aggregate import demand being Granger causal for the level of real domestic income and relative prices in the short-run. These effects are captured through the short-run response coefficients described by the matrices $\boldsymbol{\phi}_1$ through $\boldsymbol{\phi}_p$.

Under the assumption $\lambda_{xm} = 0$, and using Eq. (3), the equation for aggregate import demand from Eq. (4) can be rewritten as:

$$\Delta m_t = \alpha_0 + \psi m_{t-1} + \boldsymbol{\delta x}_{t-1} + \sum_{j=1}^{p-1} \beta_{m,j} \Delta m_{t-j} + \sum_{j=1}^{q-1} \boldsymbol{\beta}_{x,j} \Delta \boldsymbol{x}_{t-j} + \boldsymbol{\omega} \Delta \boldsymbol{x}_t + u_t, \tag{7}$$

where $\alpha_0 = \mu_m - \boldsymbol{\omega}'\boldsymbol{\mu_x}$; $\psi = \lambda_{mm}$; $\boldsymbol{\delta} = \boldsymbol{\lambda}_{mx} - \boldsymbol{\omega}'\boldsymbol{\lambda}_{xx}$; $\beta_{m,j} = \gamma_{mm,j} - \boldsymbol{\omega}'\boldsymbol{\gamma}_{xm,j}$; and $\beta_{x,j} = \gamma_{mx,j} - \boldsymbol{\omega}'\gamma_{xx,j}$. This is what Pesaran *et al.*(2001) refer to the AutoRegressive Distributed Lag Model, which is denoted as ARDL(p, $\boldsymbol{q}$), or the Unrestricted Error Correction Model (UECM). Eq. (7) can be estimated by ordinary least squares and the absence of a long-run relationship between m_t and $\boldsymbol{x}_t$ can be tested by calculating the F-statistic for the null hypothesis of $\psi = \boldsymbol{\delta} = 0$. Under the alternative of interest, $\psi \neq 0$ and $\delta \neq \mathbf{0}$, there is a stable long-run relationship between m_t and $\boldsymbol{x}_t$, which is described by:

$$m_t = a_0 + \boldsymbol{a}_1 \boldsymbol{x}_t + v_t, \tag{8}$$

where $a_0 = -\alpha_0/\psi$, $\boldsymbol{a_1} = \boldsymbol{\delta}/\psi$, and v_t is a mean zero stationary process.

Pesaran *et al.*(2001) demonstrate that the distribution of the F-statistic under the null depends upon the order of integration of the empirical series. For example, in the trivariate case where all variables are $I(0)$ and where the regression includes an unrestricted intercept, the appropriate 95% asymptotic critical value is 3.79. On the other hand, when all variables are $I(1)$, this critical value rises considerably to 4.85. For cases in which one series is $I(0)$ and the other is $I(1)$, the 95% asymptotic critical value is somewhere in-between these two bounds; see Pesaran *et al.*(2001, Table CI(iii)).

If the computed F-statistic exceeds the upper critical value, then the null is rejected in favor of the alternative which is that a long-run relationship does exist between the aggregate import demand and real domestic income together with relative prices, irrespective of whether the explanatory variables are purely $I(0)$ or $I(1)$, or mutually cointegrated. In the case the computed

F-statistic falls below the lower critical value, the null hypothesis of no long-run relationship cannot be rejected. However, when the computed F-statistic falls in-between these two bounds, no conclusive inference can be made. In such circumstances, knowledge of the cointegration rank of the forcing variables $\boldsymbol{x}_t$ is required to proceed any further.

3 Methods, Data and Results

Quarterly time series data for the 1976Q1–2004Q1 period are used here, and they comprise total 113 observations. Data on the real aggregate import quantities and real domestic income (GDP) are taken from the NIAQ data of AREMOS, for the Taiwan area. The price ratio is calculated by the ratio of the import price index to the wholesale price index, both of which are taken from the PRICE data of AREMOS, for the Taiwan area. What is important note here is that Hakkio and Rush (1991) reported that increasing the number of observations by using monthly or quarterly data does not add any robustness to the results from the Engle-Granger or Johansen approaches. In that our data span is relatively small, it is expected here that Pesaran *et al.*'s (2001) methodology should provide robust empirical results.

All models in this study are estimated by the OLS (in the log-linear form) and are subjected to a number of diagnostic tests. In most cases, both the Lagrange Multiplier (LM) version of the test and the F-version are reported. As noted by Pesaran and Pesaran (1997), both versions have the same distribution asymptotically, but on the basis of the Monte Carlo results the F-version is generally preferable to the LM version in small samples. The CUSUM test and the CUSUM of the squares test, developed by Brown, Durbin and Evans (1975) are also applied to determine the stability of the parameter estimates. All estimations are computed using MICROFIT 4.0 software.

The first step in applying the bounds test is to specify an optimal lag length for the UECM, i.e., Eq. (7).[6] Table 2 gives Akaike's and Schwarz's Bayesian Information Criteria, denoted and appropriately distinguished based on the source as the AIC and SBC respectively, the Lagrange Multiplier statistics for testing the hypothesis of no residual serial correlation against the orders of 1 and 4, as denoted by $\chi^2_{SC}(1)$ and $\chi^2_{SC}(4)$, respectively. The lag order determined from the AIC is $\hat{p}_{aic} = 4$, while the SBC suggests selecting the lag order of $\hat{p}_{sbc} = 3$. The χ^2_{SC} statistics indicate that no serial correlation remains in the residual when the lag length is set at equal to either 3 or 4.

Seeing that the assumption of the serial uncorrelated errors is important for the validity of the bounds tests, it seems prudent to select p to be either 3 or 4. Here the bounds tests are conducted to confirm the existence of a long-run equilibrium relationship, and the results are reported in Table 3. It is clear that, irrespective of whether p is 3 or 4, the computed F-statistic exceeds the upper critical value, a strong indicator that the null hypothesis of a no level long-run relationship must be rejected.[7]

In practice, there is no reason that p and $\boldsymbol{q}$ in Eq. (7) to have the same value, and thus, this possibility is allowed for. The preferred ARDL model is ADRL(4, 4, 0) for the aggregate import

[6] As noted by Pesaran *et al.*(2001, p312), "in testing the null hypothesis of the absence of the level long-run relationship in Eq. (7), namely $\psi = \delta = 0$, it is important that the coefficients of the lagged change remain unrestricted; otherwise, these tests could be subject to a pre-testing problem. However, for the subsequent estimations of the level effects and short-run dynamics of the adjustments, the use of more parsimonious specifications seems advisable."

[7] The UECM model is also estimated with a deterministic trend and the bounds test is conducted. The optimal lag selected by the AIC and SBC are 3 and 4, respectively, values which are the same as those from the UECM model without deterministic trend. The computed F-statistic of the bounds test is less than the lower critical value, suggesting that there is no long-run relationship between the regressors. However, the estimate of the deterministic trend is also not significant at the conventional level, indicating that the UECM model without a deterministic trend is the preferred model.

demand function, as selected by the Akaike information criterion,[8] with the estimates of the long-run level relationship between aggregate import demand and real domestic income along with relative prices (see Table 4 being reported) by:

$$\ln M_t = \underset{[-2.59]}{-10.00} + \underset{[6.35]}{1.61} \ln Y_t \underset{[-0.25]}{-0.21} \ln RP_t + \hat{v}_t.$$

That is, the long-run real income elasticity is 1.61 (t-ratio: 6.35), and it is significant at the conventional level. This compares with the estimate of the long-run price elasticity at -0.21 (t-ratio: -0.25), which is insignificantly different from zero. These results suggest that, in the long-run, the volume of aggregate import demand is only responsive to real income but not at all to relative price. Note that the estimated long-run coefficients represent the cointegrating vector.

The conditional ECM regression associated with the above level long-run relationship is given in Table 5. The LM test, like the F-test, shows no evidence of a residual serial correlation, while Ramsey's RESET test shows no misspecification with respect to its functional form. Furthermore, there is no evidence of heteroscedasticity, of the non-normality of the residual, nor of the residual ARCH effect. Aside from this, the preferred specification passes the CUSUM and CUSUMSQ tests of parameter stability (see Figure 1). While the CUSUM test detects systematic changes in the regression coefficients, the CUSUMSQ test is particularly useful in capturing sudden departures from the constancy of the regression coefficients.

These estimates provide further substantive evidence of the complicated dynamics that seem to exist between aggregate import demand and real domestic income along with relative prices. Although changes in relative prices are insignificant, all three lagged changes in the domestic income and lagged changes in the aggregate import demand are statistically significant at the 5%

[8]The ARDL model is also estimated with $p = 3$ as obtained from the Schwarz information criterion, but the diagnostic tests indicate that the residual still exhibit a serial correlation.

level, providing further justification for the choice of $p = 4$. These results suggest that, again, the volume of aggregate import demand is only responsive to real income but not to relative prices in the short-run. The estimated short-run income elasticity is 2.05, which is higher than the long-run income elasticity. A devaluation policy can be only accomplished if the trade flows respond to relative prices in a significant and predictable manner. From the estimates of elasticity of relative prices, it can pretty much be concluded that a devaluation policy seems to be most inappropriate when it comes to improving the trade or current account balance of Taiwan.[9] Finally, the equilibrium correction coefficient is estimated at $-0.12[-4.23]$, which implies that there is only a moderate speed of adjustment from disequilibrium to equilibrium.

4 Robustness Check

In this section, Taiwan's import demand function is re-estimated using Johansen's (1988) as well Johansen and Juselius's (1990) multivariate cointegration analysis to check the robustness of the results from the ARDL model although the Johansen method is widely associated with the drawback of the problem of bias in a small sample (Cheung and Lai, 1993). The procedures of this method are as follows. First, the nonstationary property is checked using the ADF unit root test (Dickey and Fuller, 1979; 1981). The Schwarz information criterion is used to choose the lag length in the regression. The test results show that the logarithm of the volume of imports, real domestic income and relative prices are integrated of the order of one and that their first-difference series is

[9] Tang (2002) also derived similar results for Hong Kong. In particular, he found that there is no long-run relationship between Hong Kong's aggregate demand import function and real income plus relative price. The short-run relative price elasticity is 0.378, but this is insignificantly different from zero.

$I(0)$.[10]

In the next step, the lag length for the VAR model is chosen on the basis of the AIC, SBC and LR tests. All of the criteria suggest that the optimal lag length for the VAR model is five (See Table 6). In the next step, the rank of cointegration is tested by the trace λ_{trace} and maximal eigenvalue λ_{max} tests, as proposed by Johansen (1988). The results are summarized in Table 7.[11] Although the maximal eigenvalue λ_{max} tests are marginally significant at the 5% level and rejects the null hypothesis of $r \leq 1$, i.e., the alternative hypothesis of $r = 2$ is accepted. However, the results of the trace test suggest that only one cointegration relationship exists between the regressors. According to Cheung and Lai (1993) and Kasa (1992) the trace test tends to me more powerful than the maximum eigenvalue test. Thus, that only one level long-run relationship exists between the volume of real import, real domestic income and relative price in Taiwan is assumed here. After the volume of real import demand is chosen as the dependent variable, the long-run relationship between aggregate import demand and real domestic income along with relative prices is:

$$\ln M_t = \underset{(4.53)}{-10.10} + \underset{(0.30)}{1.62} \ln Y_t + \underset{(0.93)}{0.31} \ln RP_t + \hat{v}_t.$$

The numbers in brackets are standard errors. The long-run real income elasticity of 1.62 is very close to the estimate 1.61 determined from the ARDL model. The estimate of long-run price elasticity is 0.31, again, which is insignificantly different from zero at the conventional level. The insignificant estimate of the relative prices is confirmed by the over-identification restriction of $H_0 : \hat{a}_1 = 1, \hat{a}_3 = 0$, where $\hat{a}_1$ and $\hat{a}_3$ respectively are the coefficient estimates of $\ln M_t$ and $\ln RP_t$.

[10]The results of the unit root test are available from the author upon request.

[11]We test the cointegrating rank in terms of the intercept which enters the cointegrating relation and the no trend model. As argued in Pesaran, Shin and Smith (2000), this case is likely to be particularly relevant in practice and is preferable to the corresponding unrestricted case.

The computed LR statistic is 0.16 with the p-value of 0.69 compared to the $\chi^2(1)$ critical value. The associated long-run relationship imposed with the over-identification restriction is as follows:

$$\ln M_t = \underset{(1.76)}{-8.83} + \underset{(0.11)}{1.53} \ln Y_t + \underset{(None)}{0.00} \ln RP_t + \hat{v}_t.$$

Bear in mind that in the ARDL model, the long-run price elasticity is determined to have value value of -0.21, which is also insignificantly different from zero. These results reconfirm that the volume of aggregate import demand is only responsive to real income and is unresponsive to relative prices in the long-run.

Finally, the estimates of the error correction model for the variable $\ln M_t$ estimated by the OLS based on the cointegrating VAR(5) with the over-identification restriction imposed, are presented in Table 8. The LM test, like the F-test, provides no evidence of a residual serial correlation while Ramsey's RESET test shows no misspecification of the functional form. Beside this, there is no evidence of heteroscedasticity, non-normality in the residual or of the residual ARCH effects. Added to this, the preferred specification passes the CUSUM and CUSUMSQ tests for parameter stability (see Figure 2). Note that the estimates of the lagged changes in the volume of real import demand and lagged changes in real domestic income are similar to those of the ARDL model. The estimates of the lagged changes in the relative prices are also insignificantly different from zeros, a result which is also highly consistent with that of the ARDL model. The estimated equilibrium correction coefficient is -0.13, which again is very close to the estimate ($\hat{v}_{t-1} = -0.12$) from the ARDL model. It also presents a moderate speed of convergence from disequilibrium to equilibrium.

5 Concluding Remarks

The purpose of this paper is to assess the long-run relationship of the aggregate import demand function for Taiwan using quarterly data over the 1976Q1–2004Q1 period. The methodology used has only recently been developed by Pesaran *et al.*(2001) and is based on the estimation of a UECM and the bounds test. The most unique advantage of this method is that it is robust despite small sample bias irrespective of whether the regressors are $I(0)$ or $I(1)$, or are mutually cointegarted. It is found that the aggregate import volume and its determinants, real domestic income and relative prices, exhibit a level long-run relationship. The estimated equilibrium correction coefficient shows a moderate speed of adjustment from disequilibrium to equilibrium. Furthermore the empirical results from this study show that the estimated short-run elasticity and long-run income elasticity are elastic and that the short-run income elasticity is greater than its long-run counterpart. This implies that economic growth does indeed have greater negative impact on the trade balance in the short-run than in the long-run. Finally, as concerns the estimated relative price elasticity, it is insignificantly different from zero both in the short- and long-run. This strongly suggest that import volume is insensitive to any devaluation policies. The robustness of the ARDL model with respective to the Johansen (1988) and Johansen and Juselius (1990) method is also checked and the estimated results from the two methods are quite alike.

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Table 1: Value of imports and exports in Taiwan

Unit: U.S. million dollars

Year	Exports	Imports	Trade Surplus
1981	22,611	21,200	1,412
1982	22,204	18,888	3,316
1983	25,123	20,287	4,836
1984	30,456	21,959	8,497
1985	30,726	20,102	10,624
1986	39,861	24,182	15,680
1987	53,679	34,983	18,695
1988	60,667	49,673	10,995
1989	66,304	52,265	14,039
1990	67,214	54,716	12,498
1991	76,178	62,861	13,318
1992	81,470	72,007	9,463
1993	85,091	77,061	8,030
1994	93,049	85,349	7,700
1995	111,659	103,550	8,109
1996	115,942	102,370	13,572
1997	122,081	114,425	7,656
1998	110,582	104,665	5,917
1999	121,591	110,690	10,901
2000	148,321	140,011	8,310
2001	122,866	107,237	15,629
2002	130,597	112,530	18,067

Data source: Ministry of Finance, Taiwan.

Table 2: Statistics for selecting the lag order

p	AIC	BIC	$\chi^2_{SC}(1)$	$\chi^2_{SC}(4)$
1	151.14	138.95	23.03*	6.41*
2	148.30	132.10	30.24*	8.91*
3	165.80	145.61	5.33	1.15
4	167.10	142.97	3.05	0.62
5	163.26	135.20	0.97	0.18
6	158.77	126.81	4.58	0.88
7	157.62	121.79	2.00	0.36
8	154.79	115.13	1.21	0.20
9	150.69	107.22	1.85	0.30
10	147.53	100.28	8.03*	1.32

p is the lag order of the UECM.
$\chi^2_{SC}(i)$ is the LM-statistic for testing no residual serial correlation against order i.
* denotes significance at the 5% level.

Table 3: Results of bounds tests for the long-run relationship

p	F-statistic	Lower Bound, $I(0)$	Upper Bound, $I(1)$
3	5.55*	3.79	4.85
4	5.41*	3.79	4.85

p is the lag order of the UECM.
From Pesaran *et al.*(2001), Table CI(iii): unrestricted intercept and no trend (two regressors, $k = 2$).
* denotes significance at the 5% level.

Table 4: Estimated long-run coefficients using the ARDL approach

Regressor	Coefficient	t-Ratio[Prob]
$\ln Y_t$	1.61	6.35[0.00]
$\ln RP_t$	-0.21	-0.25[0.80]
CONST	-10.00	-2.59[0.01]

Table 5: Error correction representation of the ARDL(4, 4, 0) model

Regressor	Coefficient	t-Ratio[Prob]
$\Delta \ln M_{t-1}$	-0.55	-6.82[0.00]
$\Delta \ln M_{t-2}$	-0.30	-3.21[0.00]
$\Delta \ln M_{t-3}$	-0.43	-5.27[0.00]
$\Delta \ln Y_t$	2.05	9.13[0.00]
$\Delta \ln Y_{t-1}$	1.26	4.94[0.00]
$\Delta \ln Y_{t-2}$	1.32	5.03[0.00]
$\Delta \ln Y_{t-3}$	1.75	7.50[0.00]
$\Delta \ln RP_t$	0.04	0.16[0.87]
CONST	0.01	0.41[0.68]
$\hat{v}_{t-1}$	-0.12	-4.23[0.00]

	Diagnostic Tests	
Test Statistics	LM Version	F Version
A: Serial Correlation	$\chi^2(4)$= 5.51[0.23]	F(4, 95)= 1.26[0.28]
B: Functional Form	$\chi^2(1)$= 0.02[0.87]	F(1, 98)= 0.02[0.88]
C: Normality	$\chi^2(2)$= 1.33[0.51]	—
D: Heteroscedasticity	$\chi^2(1)$= 1.51[0.21]	F(1, 107)= 1.50[.22]
E: ARCH	$\chi^2(4)$= 2.71[0.60]	F(4, 94)= 0.61[0.66]

A: Lagrange Multiplier test of residual serial correlation.
B: Ramsey's RESET test using the square of the fitted values.
C: Based on a test of skewness and kurtosis of the residuals.
D: Based on the regression of the squared residuals on the squared fitted values.
E: Autoregressive conditional heteroscedasticity test of the residuals .
$\hat{v}_t = \ln M_t - 1.61 * \ln Y_t + 0.21 * \ln RP + 10.00 * CONST$

Table 6: Test statistics and choice criteria for selecting the order of the VAR model

Order	LL	AIC	SBC	LR Test	Adjusted LR Test
6	795.63	738.63	662.45	—	—
5	790.86	742.86	678.71	$\chi^2(9)$= 9.54[0.38]	7.84[0.54]
4	755.83	716.83	664.71	$\chi^2(18)$= 79.59[0.00]	65.46[0.00]
3	695.41	665.41	625.32	$\chi^2(27)$= 200.42[0.00]	164.83[0.00]
2	678.28	657.28	629.21	$\chi^2(36)$= 234.69[0.00]	193.02[0.00]
1	656.20	644.20	628.17	$\chi^2(45)$= 278.84[0.00]	229.33[0.00]
0	203.53	200.53	196.52	$\chi^2(54)$=1184.20[0.00]	973.91[0.00]

Table 7: Cointegration with restricted intercepts and no trends in the VAR

Cointegration LR Test based on the Maximal eigenvalue of the stochastic matrix				
Null	Alternative	Statistic	95% Critical Value	90% Critical Value
$r = 0$	$r = 1$	33.91*	22.04	19.86
$r \leq 1$	$r = 2$	15.89*	15.87	13.81
$r \leq 2$	$r = 3$	2.31	9.16	7.53
Cointegration LR Test based on the trace of the stochastic matrix				
Null	Alternative	Statistic	95% Critical Value	90% Critical Value
$r = 0$	$r \geq 1$	52.12*	34.87	31.93
$r \leq 1$	$r \geq 2$	18.21	20.18	17.88
$r \leq 2$	$r = 3$	2.31	9.16	7.53

* denotes significance at the 5% level.

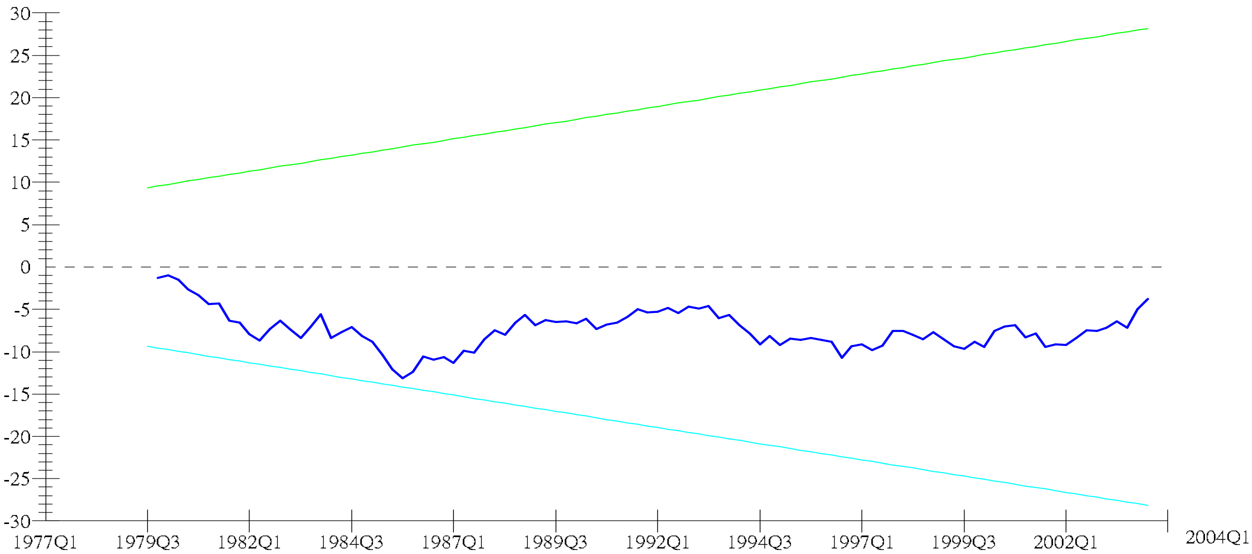
Table 8: ECM for variable LM estimated by the OLS based on cointegrating VAR(5)

Regressor	Coefficient	t-Ratio[Prob]
$\Delta \ln M_{t-1}$	-0.49	$-4.74[0.00]$
$\Delta \ln M_{t-2}$	-0.35	$-3.10[0.00]$
$\Delta \ln M_{t-3}$	-0.44	$-3.84[0.00]$
$\Delta \ln M_{t-4}$	0.07	0.68[0.49]
$\Delta \ln Y_{t-1}$	1.25	4.14[0.00]
$\Delta \ln Y_{t-2}$	1.28	4.19[0.00]
$\Delta \ln Y_{t-3}$	1.62	5.02[0.00]
$\Delta \ln Y_{t-4}$	1.76	5.31[0.00]
$\Delta \ln RP_{t-1}$	0.15	0.44[0.66]
$\Delta \ln RP_{t-2}$	-0.15	$-0.44[0.65]$
$\Delta \ln RP_{t-3}$	0.09	0.27[0.78]
$\Delta \ln RP_{t-4}$	0.27	0.83[0.40]
$\hat{v}_{t-1}$	-0.13	$-3.60[0.00]$

	Diagnostic Tests	
Test Statistics	LM Version	F Version
A: Serial Correlation	$\chi^2(4)$= 6.25[0.18]	F(4, 91)= 1.39[0.24]
B: Functional Form	$\chi^2(1)$= 0.68[0.40]	F(1, 94)= 0.60[0.43]
C: Normality	$\chi^2(2)$= 2.38[0.30]	—
D: Heteroscedasticity	$\chi^2(1)$= 0.21[0.64]	F(1, 106)= 0.21[0.64]
E: ARCH	$\chi^2(4)$= 4.03[0.41]	F(4, 94)= 0.88[0.47]

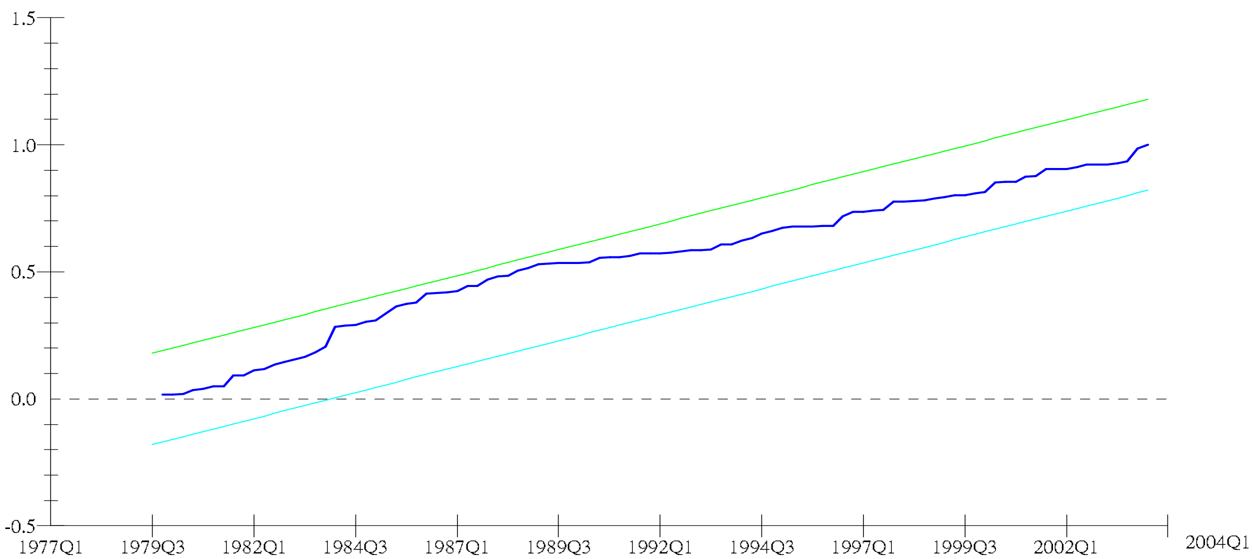
A: Lagrange Multiplier test of residual serial correlation.
B: Ramsey's RESET test using the square of the fitted values.
C: Based on a test of skewness and kurtosis of the residuals.
D: Based on the regression of the squared residuals on the squared fitted values.
E: Autoregressive conditional heteroscedasticity test of the residuals .
$\hat{v}_t = \ln M_t - 1.53 * \ln Y_t + 0.00 * \ln RP + 8.83 * CONST$

Plot of Cumulative Sum of Recursive Residuals



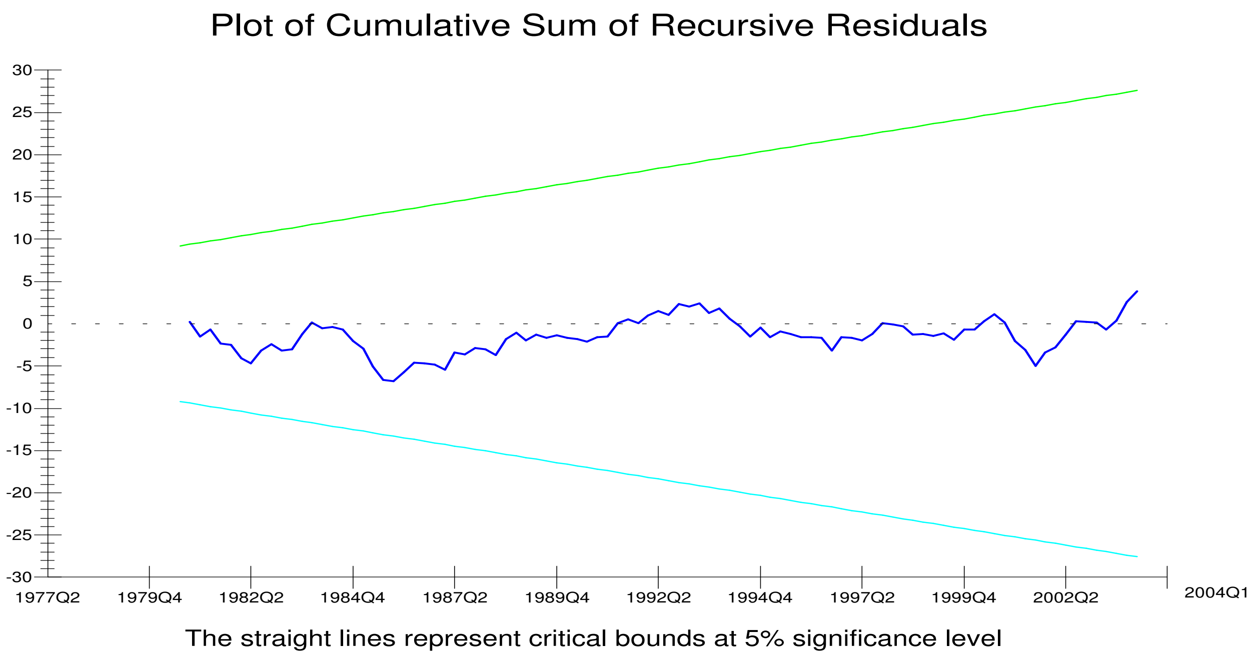
The straight lines represent critical bounds at 5% significance level

Plot of Cumulative Sum of Squares of Recursive Residuals



The straight lines represent critical bounds at 5% significance level

Figure 1: Plot of r=the results from the CUSUM and CUSUMSQ tests following the Pesaran *et al.*(2001) approach.



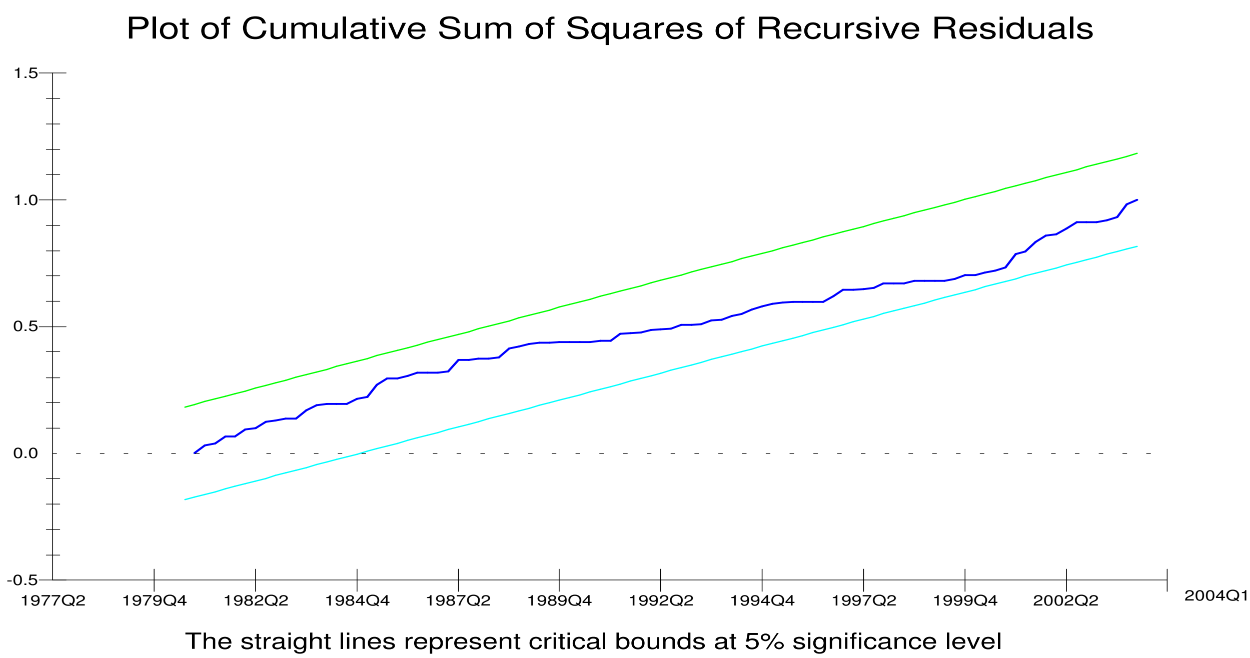


Figure 2: Plot of the results from the CUSUM and CUSUMSQ tests following Johansen's approach.