An aggregate import demand function for Australia: a cointegration approach

By

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ABSTRACT

This paper investigates the relationship among quantity of imports, relative import prices and real GDP in the aggregate import demand function for Australia during the period 1959Q3–2006Q3. Testing for cointegration, we find these variables are not stationary but are cointegrated. The results are consistent across three different cointegration tests conducted, namely the Engle-Granger's residual-based test, the Johansen and Juselius multivariate test and the Bounds Test. As only one cointegration vector is found, there is a unique long-run equilibrium relationship among the variables. In the long-run, the price elasticity is found to be close to unity and import demand is found to be fairly income elastic. The error correction model is used to investigate the dynamic behaviour of import demand. In the short-run, Australian import demand is both price and income inelastic. Price is more elastic than income in the short-run, indicating that it is the dominant determinant of Australian import demand in the short-run. Furthermore, the estimated error correction coefficient of 0.3090 suggests that the aggregated Australian import demand corrects from the previous period's disequilibrium by 31% per quarter. That is, it takes approximately 10 months to fully realign any disequilibrium that occurs. This study provides the only assessment of Australian import demand including a precise estimate for the short-run relationship, especially an estimate of the short-run adjustment term. This information will provide further input to support policy decisions relating to the management of the Australian trade balance.

Keywords: Import demand; Cointegration, Error correction model

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1 Introduction

"It is a well-known empirical fact that many macroeconomic time series are typically non-stationary, as indicated by the high serial correlation between successive observations, particularly when the sampling interval is small." (Dutta and Ahmed, 1999, p.465) Therefore, any attempt to estimate a relationship among non-stationary series will lead to spurious results. Cointegration analysis is an exception. Using cointegration analysis, we investigate the role of relative import prices and real GDP in determining the real quantity of Australian import demand. The existence of a long-run equilibrium relationship among the variables is tested by three different procedures namely, the Engle and Granger (1987) (EG) residual-based test, the Johansen and Juselius (1990) (JJ) maximum likelihood estimation (MLE) multivariate test and the Bounds Test of Pesaran (2001). The short-run relationship is estimated through an error correction model (ECM) to investigate the dynamic behaviour of Australian import demand. The findings in this paper significantly contribute in drawing up policy prescriptions relating international trade in Australia.

The objective of this paper is to investigate the behaviour of Australian aggregate imports during the period 1959Q3—2006Q3. There have been numerous empirical studies in aggregate import behaviour relating to developed countries, Latin America and Asia Pacific countries. However, there are only a handful of studies examining Australian import demand and especially using the recently developed methodology of cointegration. The present study aims to fill this gap in the literature.

Given the amazing economic growth rate in the last two decades, Australia is enjoying its most favourable terms of trade. Australia's terms of trade measures the price of goods and services exported from Australia relative to the price of goods and services imported to Australia. The terms of trade are an important economic indicator, showing the ability of a country to purchase imports for a given level of exports. During the period 2003-2005, Australia's terms of trade increased by 31%, to reach levels last observed in the early 1970s (Australian Bureau of Statistics (2005)). The current boom in Australia is primarily driven by export prices, particularly due to the increase in the prices of mining products in international markets. Between December 2002 and June 2006, the overall index of Australia's export commodity prices in US dollars increased by 88%, while the base metals component of the index increased by 171% (Reserve Bank of Australia (2005)). Although the terms of trade are also affected by a decline in import prices (mostly manufactured goods, in particular high technology goods), the effect is relatively small.

Together with rising commodity prices, other Australian industries have also been performing strongly. They generate more income and create a strong financial position for the Australian economy. All of them tend to increase import demand for manufacturing as well as for spending. It has created current account deficits. Mercereau and Rozhkov (2006, p.3) assert "since the floating of the Australian dollar and the liberalization of international capital flows in the mid-1980s these deficits have averaged 4.5 percent of GDP. This is high compared with other advanced economies, where the average current account balance is about zero. Persistent current account deficits have translated into rising net foreign liabilities, reaching 60 percent of GDP in 2005; Australia's net foreign position [assets minus liabilities] is unusually negative by OECD standards."

However, looking at compositions of Australian imports (Figure 1), there should be no relationship between the growth of Australian economy and import demand. In the last 25 years, the period in which Australian economy has experienced one of the most incredible economic growth, the share of each category in the composition namely, consumption, intermediate, capital and other goods has been relatively constant. However, there is a slight decrease in the proportion of capital and intermediate goods in Australian imports, which are offset by an increase in the proportion of consumption goods. Nevertheless, this feature in Australian imports can be explained by examining the nature of Australian economy. Given the boom in commodity prices, Australia tends to focus more on producing raw materials than on producing manufactured goods. That in turn leads to a decrease in demand for imports of intermediate goods, because such goods are used as inputs in the production process and predominantly accounted for as services. Furthermore, with the phenomenal economic growth in Australia, household consumption has been increasing, resulting in an increase in demand for imports and eventually increasing the share of consumption goods in total imports. At a time when import demand is driven by consumption, economic policy to reduce imports would not harm an economy like Australia.

With large current account deficits, the sustainability of Australian economy is questioned. As these deficits mainly originate in the private sector, it is the question of whether the sector should be trusted or whether there are risks associated with large current account deficits. This study will provide a detailed assessment of the Australian import demand function which will provide further information for the Government to address the unfavourable current account position of Australia.

The remainder of this paper is organised as follows. Section 2 gives an overview of the extant literature. Section 3 provides the theoretical and conceptual framework for the analysis presented. It is followed by an examination of the data in Section 4 and Section 5. A complete assessment of the import demand function for Australia is provided in Section 6, and following that the implications of empirical results are discussed in Section 7. Finally, some concluding remarks are discussed in Section 8.

2 Literature Review

There are numerous empirical studies investigating aggregate import behaviour, but most of them have concentrated on developed countries. Using the annual data over the period 1960-82, Arize and Afifi (1987) specify and estimate import demand functions for thirty developing countries. They are Algeria, Benin, Cameroon, the Central African Republic, Chad, Congo, Egypt, Ethiopia, Gabon, Gambia, Ivory

Coast, Israel, Kuwait, Liberia, Malawi, Mali, Morocco, Niger, Pakistan, Rwanda, Senegal, Sierra Leone, Somalia, Tanzania, Togo, Tunisia, Uganda, Upper Volta, Zaire and Zambia. The objective of their paper is to address the responsiveness of the price of aggregate imports and to determine whether the import demand relationship has shifted during the period of estimation.

Traditionally, the real quantity of imports demanded is generally determined by the ratio of import prices to domestic prices and domestic real income, in period *t*. This equation is written as follows:

$$M_t = F(P_t, Y_t) \tag{1}$$

where M is the real quantity of imports, P is the price ratio ($f_1 \leq 0$), and Y is real income ($f_2 \geq 0$). In order to find the specific model for each country, Arize and Afifi (1987) estimate four log-linear variants of the above equation. Although the inclusion of a lagged dependent variable in some of these equations implies a partial adjustment process, the validity of the equations depends on whether the variables are stationary. If such variables are not stationary, the analysis may suffer from the problem of spurious regression. As the result, the ordinary least squared (OLS) estimates of the parameters are inconsistent and less efficient unless the variables are cointegrated. Furthermore, the data generating process will not display a valid error correction presentation. The high R^2 for the estimated equations for most of the thirty countries is an indication of spuriousness.

For the case of Australia, Athukorala and Menon (1995) investigate the relationship between manufactured import flow, and relative prices and domestic economic activity net of cyclical demand effects over the period 1981Q3 to 1991Q2. This is achieved through estimation of import demand functions for total manufactured imports and nine major import categories using the general-to-specific modelling approach. The general form of their import demand function is the following:

$$MQ_t = f(RP_t, AC_t, SS_t) \qquad (2)$$

where MQ_t is real imports, RP_t is relative price derived by dividing the tariff augmented import price by the price of the domestic-competing commodity ($f_1 \leq 0$), AC_t is a measure of related domestic economic activity ($f_2 \geq 0$), and SS_t is the ratio of stocks to average sales volume as a measure of the general scarcity of domestic supplies ($f_3 \geq 0$). This model is very similar to the model presented in Arize and Afifi (1987). The relative price and real income variables are used to identify demand effects on imports. Nevertheless, the inventory-sales ratio is introduced as a control variable to capture any cyclical demand effect.

It is very important to take into account of the time series properties of the variables used. Athukorala and Menon (1995) first test the time-series properties of the data using the Dickey-Fuller (DF) and Phillips-Perron (PP) procedures (for details on these procedures, see Section 3.1 Unit-Root Tests). The test results indicate that MQ_t, RP_t, and AC_t are non-stationary process of order one—or $I(1)$ in all the cases. Guided by this finding, they test for long-run equilibrium relationships between these variables using the EG and JJ cointegration procedures (a detailed discussion of tests of cointegration is given in Section 3.2), but failed to find evidence of such a relationship. In the analysis presented in this paper, such a long-run relationship is determined with a much larger sample size.

In spite of the absence of cointegration relationships between non-stationary series, Athukorala and Menon (1995) are reluctant to ignore the long-run relationship embodied in the variables in levels. Thus, they use the general-to-specific modelling procedure, which minimises the possibility of estimating spurious relationships, while retaining long-run information to estimate the relationship. However, the above-mentioned procedure is fairly cumbersome.

In the light of cointegration, Sinha and Sinha (2000) estimate the aggregate import demand function for Greece using annual data for the period 1951-1992. Of little variation from the traditional formulation, import demand is estimated with respect to import price, domestic price, and GDP. The import demand function takes the following form:

$$M_t = f(PM_t, PD_t, Y_t) \tag{3}$$

where M_t is the import demand, PM_t is the import price ($f_1 \leq 0$), PD_t is the domestic price ($f_2 \geq 0$), and Y_t is real GDP ($f_3 \geq 0$). The empirical counterpart (log-linear form) of the import demand function given in equation (3) is detailed as follows:

$$\ln M_t = \beta_1 + \beta_2 \ln PM_t + \beta_3 \ln PD_t + \beta_4 \ln Y_t + \varepsilon_t \tag{4}$$

Furthermore, Sinha and Sinha (2000) also include a lagged dependent variable, $\ln M_{t-1}$ in equation (4) to capture the partial adjustment process, and derive the short-run equation for import demand which is as follows:

$$\ln M_t = a_1 + a_2 \ln PM_t + a_3 \ln PD_t + a_4 \ln Y_t + a_5 \ln M_{t-1} + u_t \tag{5}$$

Sinha and Sinha (2000, p.201) propose that "The coefficients of [the equation] will give the short-run elasticities because of its log-linear formulation." However, this equation is only a special case of a vector autoregression (VAR) model. When the dependent variable is in levels rather than in first differences, the estimated coefficients just provide long-run elasticities if and only if there is a cointegration vector among the variables. Therefore, the empirical results in the study by Sinha and Sinha (2000) should be interpreted from the long-run perspective rather than the short-run.

Using the cointegration and error correction modelling approaches, Dutta and Ahmed (1999) investigate the aggregate import demand function for Bangladesh. After finding the existence of unit roots, and therefore establishing non-stationarity in the levels of some variables they then apply the two commonly used procedures of cointegration tests namely, the EG test and the JJ test. For the EG test, the long-run relationship between the logarithm of the real quantity of imports and its major determinants is estimated by OLS. Then an examination for stationarity of residuals is undertaken by using the Augmented Dickey Fuller (ADF) test (see Section 3.1. Unit-Root Tests) but the results were inconclusive. Nevertheless, for the JJ test, they find a cointegrating relationship among real quantities of imports, real import prices, real GDP and real foreign exchange reserves. As the JJ procedure is superior to the EG procedure in determining cointegrating relationships, Dutta and Ahmed (1999) conclude that there exists a stable long-run relationship of aggregate import demand with its major determinants. Moreover, Dutta and Ahmed (2004) also conducts a similar investigation of import demand for India.

In addition, Masih and Masih (2000) provide a succinct description of the superiority of the JJ procedure. They assert "the JJ procedure poses several advantages over the popular residual-based EG two-step approach in testing for cointegration. Specifically, they are summarised as follows. (1) the JJ procedure does not, *a priori*, assume the existence of at most a single cointegrating vector; rather, it explicitly tests for the number of cointegrating relationships; (2) unlike the EG procedure, which is sensitive to the choice of the dependent variable in the cointegrating regression, the JJ procedure assumes all variables to be endogenous; (3) [related to (2)], when it comes to extracting the residual from the cointegrating vector, the JJ procedure avoids the arbitrary choice of the dependent variable as in the EG approach, and is insensitive to the variable being normalised; (4) the JJ procedure is established on a unified framework for estimating and testing cointegrating relations within the VECM formulation; (5) JJ provides the appropriate statistics and the point distributions to test hypothesis for the number of cointegrating vectors and tests of restrictions upon the coefficients of the vectors."

For these reasons, Masih and Masih (2000) use the JJ multivariate cointegration procedure to re-assess long-run elasticities of Japanese import demand. The analysis of Mah (1994) is based on the EG test of cointegration and fails to find evidence of a long-run relationship in the import demand function for Japan among quantity of imports, the relative price and real income. However, for the case of Japan, Masih and Masih (2000) find these variables being cointegrated, and thus share a long-run equilibrium relationship. Hence, they conclude both price and income variables do affect import demand significantly, and play an important role in explaining Japanese import demand, at least over the long-run.

By implementing the JJ multivariate cointegration procedure, Alias and Cheong (2000) examine the long-run relationship between Malaysian aggregate imports and the components of final demand expenditure (namely, public and private consumption expenditure, investment expenditure and exports) and relative prices by using annual data for the period 1970 and 1998. They argue if different components of final demand are significant in determining import demand the use of a single demand variable in the aggregate import demand function would lead to any aggregation bias. Accordingly, Alias and Cheong (2000) find the quantity of Malaysian import demand is cointegrated with its determinants, while both final consumption expenditure and investment expenditure appear to be the dominant determinants in the long run.

The traditional formulation of cointegration has still been used to examine import demand behaviour. Typically, the analysis presented in Tsionas and Christopoulos (2004) for five industrial countries namely, France, Italy, the Netherlands, the UK, and the US; in Islam and Hassan (2004) for Bangladesh, and in Anoruo and Usianeneh (2003) for Australia. The empirical results in the above mentioned cointegration assessments show significant effects of income and relative prices on import demand. But surprisingly, according to Anoruo and Usianeneh (2003), import demand in Australia is both income and price inelastic.

Recently, Pesaran et al. (2001) developed a relatively new technique for cointegration analysis which is called the Bounds Test (for detailed discussion of this technique, see Section 3.2.3.). Subsequently, given its flexibility, this approach has been used by most researchers to test for cointegration. Tang (2004) reinvestigates the empirical evidence on the long-run relationship of aggregate import demand behaviour for the ASEAN-5 founding nations. Adopting the import demand function that has been developed by Xu (2002), where "national cash flow" rather than GDP is used as the income variable, Tang (2004) find the quantity of imports, income variable, and relative price of imports are cointegrated in Malaysia and Singapore, but not cointegrated in Indonesia, Thailand, and the Philippines. Narayan and Narayan (2005) use the same technique to estimate a disaggregated import demand model for Fiji using relative prices, total consumption, investment expenditure and export expenditure variables for the period 1970 to 2000. Moreover, Narayan and Narayan (2005) claim their study is an advance over existing studies using the bounds testing approach because they use the bounds *F*-statistic critical values specific to their sample size. By using those critical values, they argue inference from their study is more appropriate. Along with evidence of a cointegration relationship among the variables, Narayan and Narayan (2005) find that total consumption expenditure, investment expenditure and export expenditure have an inelastic and positive impact on import demand while an increase in relative prices induce less imports.

Given the small sample size, Razafimahefa and Hamori (2005) decide to use the bounds test to investigate the long-run relationship among quantity of imports, level of income and relative prices of imports in the aggregate import demand functions of Madagascar and Mauritius. They find the existence of a cointegration relationship; and the long-run income and price elasticities in both countries are inelastic.

Furthermore, Tang (2005) again applies bounds testing to re-examine the long-run relationships of South Korea's aggregate import demand behaviour. The study includes four income variables namely, GDP, GDP minus exports, national cash flow and final expenditure components in the import demand formulation. More comprehensive than existing studies, Tang (2005) takes into account other techniques for cointegration analysis. The techniques include ADF test, ADF test with unknown structural break, JJ multivariate test, error correction mechanism test, and error correction model (ECM) approach. This would enable crosschecks of the consistency of the findings among different cointegration techniques. Using quarterly data for 1970-2002, Tang (2005) finds consistent evidence of a cointegrating relation in South Korea's aggregate import demand.

3 Methodology

3.1 Unit-Root Tests

The standard statistical properties of OLS hold only when the time series variables involved are stationary. A time series is said to be stationary if its mean, variance, and auto-covariance are independent of time.

Nelson and Plosser (1982) developed a test originated by Dickey and Fuller (1979) (DF) to determine whether a time series is stationary. The test is based on the model:

$$y_t = \mu + \rho y_{t-1} + \gamma t + e_t \qquad (6)$$

Subtracting both sides by y_{t-1}, we obtain:

$$\begin{aligned} &\Delta y_t = \mu + (1-\rho) y_{t-1} + \gamma t + e_t \\ \text{and} \quad &H_0: \quad p - 1 = 0 \ (\ y_t \text{ is non-stationary}) \\ &H_1: \quad p - 1 < 0 \ (\ y_t \text{ is stationary}) \end{aligned} \qquad (7)$$

The *t*-test (non-standard) on the estimated coefficient of y_{t-1} provides the DF test for the presence of a unit-root. The Augmented DF (ADF) test is a modification of the DF test and involves augmenting the above equation by lagged values of the dependent variables. It is made to ensure that the error process in the estimating equation is residually uncorrelated, and also captures the possibility that y_t is characterised by a higher order autoregressive process. Although the DF methodology is often used for unit-root tests, it suffers from a restrictive assumption that the error processes are *i.i.d.* Dutta and Ahmed (1999, p.466) assert that "When economic time series exhibit heteroskedasticity and non-normality in raw data, the PP non-parametric tests are preferable to the DF and ADF tests."

3.2 Tests for Cointegration

In the face of non-stationary series with a unit root, first differencing appears to provide the appropriate solution to our problems. However, first differencing has eliminated all the long-run information which economists are invariably interested in. Later, Granger (1986) identified a link between non-stationary processes and preserved the concept of a long-run equilibrium. Two or more variables are said to be cointegrated (there is a long-run equilibrium relationship), if they share common trend. Cointegration exists when a linear combination of two or more non-stationary variables is stationary.

3.2.1 The Engle-Granger (EG) Procedure

Once pre-testing has demonstrated that the variables are integrated of the same order, OLS is used to estimate the parameters of a cointegrating relationship. It has been shown that the application of OLS to a $I(1)$ series yields super-consistent estimates. That is estimates converge on to their true values at a faster rate than the case if $I(0)$ or stationary variables are used in estimation. Then, these parameter values are used to compute the residuals. Cointegration tests are the test for stationarity of the residuals by using DF and ADF tests. If the residuals are stationary, there exists one cointegrating relationship among variables and it will rule out the possibility of the estimated relationship being "spurious".

Since the residuals are estimated by OLS, by construction the residual variance is made as small as possible, the test is prejudiced towards finding a stationary error process. The test is also sensitive to how the equation is presented (i.e. whether x is regressed on y or vice versa). Finally, if there are more than two variables, the EG procedure will not allow discrimination between different cointegrating vectors.

3.2.2 The Johansen-Juselius (JJ) Procedure

Given these limitations of the EG procedure, several methods have been developed for testing cointegration. One of the most popular is the JJ procedure. This procedure is viewed as a generalisation of the DF testing procedure to the multivariate case. The model is written as follows:

$$\Delta Y_t = \delta + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i} + \Pi Y_{t-k} + \varepsilon_t \tag{8}$$

where Y_t is a column vector of the m variables, Γ and Π represent coefficient matrices, Δ is a difference operator, k denotes the lag length, and δ is a constant. The JJ procedure involves the identification of rank of the m by m matrix Π (or the number of its characteristic roots – Eigen values). If Π has zero rank, there is no cointegrating vector and it is the usual Vector Autoregression Model (VAR) in first difference form. If the rank r of Π is greater than zero, there are multiple cointegrating vectors, and Π may be decomposed into two matrices α and β such that $\Pi = \alpha\beta'$. In this version, β contains the coefficients of the r distinct cointegrating vectors giving $\beta' Y_t$ stationary (Y_t may not be stationary) and α contains the speed-of-adjustment coefficients.

There are two tests to determine the number of cointegrating vectors namely, the trace test and the maximum eigenvalue test. They are defined as follows:

$$\lambda_{trace}(r) = -T\sum_{i=r+1}^{n}\ln(1-\hat{\lambda}_i) \tag{9}$$

$$\lambda_{\max}(r, r+1) = -T\ln(1-\hat{\lambda}_{r+1}) \tag{10}$$

where $\hat{\lambda}_i$ is the estimated value of the characteristic roots, T is the number of usable observations, and r is the number of distinct cointegrating vectors. In the trace test, the null hypothesis (H_0) is there is at most r cointegrating vectors (i.e. $r = 0,1,2...$) is tested against a general alternative. Alternatively, in the maximum eigenvalue test, the null hypothesis ($H_0 : r = 0$) is tested against an alternative ($H_1 : r = 1$) followed by ($H_0 : r = 1$) against ($H_1 : r = 2$), and so on. The critical values for both these tests were tabulated by Johansen and Juselius (1990). The distribution of the statistics depends on the number of non-stationary components under the null hypothesis and whether or not a constant is included in the cointegrating vector.

3.2.3 The Bounds Testing Procedure

This technique for cointegration analysis was developed by Pesaran et al. (2001). It is essentially based on the estimation of the unrestricted error correction model (UECM) or error correction version of autoregressive distributed lag (ARDL) model. Other than its simplicity, the bounds test has several empirical advantages. First, according to Pesaran et al. (2001, p.315), it can be applied irrespective of whether the regressors are purely $I(0)$, purely $I(1)$, or mutually cointegrated. In other words, it is unnecessary that the order of integration of the underlying regressors be ascertained prior to testing the existence of a level relationship between two variables. Second, in the study by Pattichis (1999) and Mah (2000), the bounds testing procedure is found to be robust for small sample analysis.

Furthermore, Tang (2005, p.35) states the procedure is also applicable when the explanatory variables are endogenous and is sufficient to simultaneously correct for residual serial correlation.

The UECM is written as follows:

$$\Delta y_t = \alpha + \sum_{i=1}^{l} \beta_i \Delta y_{t-i} + \sum_{i=0}^{l} \mathrm{B}_i \Delta X_{t-i} + \pi_y y_{t-1} + \Pi_x X_{t-1} + \varepsilon_t \tag{11}$$

where l is the lag length. The absence of a level relationship between y_t and X_t is a test of the joint hypothesis, $\pi_y = 0$ and $\Pi_x = 0$, in the above equation. In other words, the Wald Test (F-statistic) is used to test the null hypotheses $H_0 : \pi_y = 0$ and $\Pi_\mathrm{x} = 0$ against the alternative $H_1 : \pi_y \neq 0$ and $\Pi_\mathrm{x} \neq 0$. However, the asymptotic distribution of the F-statistic for the bound test is non-standard under the null hypothesis among the examined variables, irrespective of whether the explanatory variables are purely $I(0)$ or $I(1)$.

Thus, Pesaran et al. (2001) developed two bounds of critical values for the different model specifications (intercept and/or trend) where the upper bound applies when all variables are $I(1)$ and the lower bound applies when all variables are $I(0)$. If for a chosen significant level, the computed F-statistic exceeds the upper bound, the null hypothesis of no cointegration is rejected. If the F-statistic is inferior to the lower bound, the null hypothesis of no cointegration cannot be rejected. When the F-statistic falls between the two bounds, conclusive inference cannot be made; and the order of integration of the variables must be known before any decision can be made.

4 Data Characteristics

The traditional formulation of import demand is specified is equation (1) as follows:

$$M_t = F(Y_t, P_t)$$

Taking all variables in log-linear form, the long-run import demand function can be written as follows:

$$\ln M_t = \beta_0 + \beta_1 \ln Y_t + \beta_2 \ln P_t + u_t \tag{12}$$

where M is real quantity of aggregate imports, Y is real gross domestic product (real GDP), and P is relative price of imports. We take data for the nominal quantities of aggregate imports (at current price), the real quantities of aggregate imports (at constant price), the real GDP (at constant price) and Consumer Price Index (CPI) from International Financial Statistics (IFS) and Australian Bureau of Statistics (ABS). As nominal values of imports are deflated by the unit value index to obtain real quantities of imports, import prices (unit value indices) are calculated by dividing the nominal quantities of aggregate imports by its real quantities. Then, import prices are deflated by CPI to obtain relative import prices. Finally, all of the above three variables are transformed into natural logarithms to interpret the coefficients as elasticities.

The summary statistics namely, means, standard deviations (SD), skewness, and kurtosis for real quantities of imports (*LRIMP*), real GDP (*LRGDP*), and relative import prices (*LRPRICE*) for the period 1959Q3 – 2006Q3 are given in Table 1. Figures 2, 3 and 4 display that the data are fit the traditional formulation of import demand adequately. *LRIMP* exhibits positive correlation with *LRGDP* and negative correlation with *LRPRICE*. *LRIMP* and *LRGDP* are trending along similar slopes, and showing an obvious pattern of seasonality. Furthermore, *LRIMP* tends to increase consistently with the growth of *LRGDP* overtime. Although being more volatile than the other two variables, *LRPRICE* shows a downward trend over the period under analysis. Since the floating of the Australian dollar in the mid-1980s, *LRPRICE* displays less volatility and a strong decreasing trend. In recent years, the stability of *LRPRICE* is a key factor to make *LRIMP* more stable and increase the correlation between *LRIMP* and *LRGDP*.

5 Seasonality

As can be seen in Figures 2 and 3, the data for *LRIMP* and *LRGDP* exhibit strong seasonality, but their amplitude remains relatively stable. Wooldridge (2006, p.869) states that seasonality is "a feature of monthly or quarterly time series where the average value differs systematically by season of the year." To test for the existence of seasonality in these variables, seasonal dummy variables are incorporated into the model.

$$\ln M_t = \alpha_{1t} D_1 + \alpha_{2t} D_2 + \alpha_{3t} D_3 + \alpha_{4t} D_4 \qquad (13)$$

The estimates of equation (13) for *LRIMP* and *LRGDP* (as well as for *LRPRICE*) are presented in Table 2. For each variable, it is characterised by 4 seasonal dummies and no constant. There is very strong evidence of seasonality in the quantity of imports and the level of GDP. All coefficients are highly significant, especially for the level of GDP. Import demand is peak in the June quarter and nadir in the March quarter, while GDP is highest at the December quarter before slump at the March quarter. Although statistically significant at 10% level, seasonality is not an explanatory factor of the relative price of imports since its associated coefficients do not have any economic significance.

When seasonality does not play a role in modelling the long-run relationship, it is the main factor which must be accounted for to obtain a plausible model in explaining the short-run behaviour of import demand.

6 Empirical Analysis

Given that time-series data for most of the variables tend to be non-stationary, there is the potential danger of capturing spurious relationships. Therefore, we begin the estimation process by testing whether there is a unit root in the above data using the ADF and PP procedures. Dutta and Ahmed (1999, p.466) state "the DF and ADF unit root tests are often applied to test whether a time series has a unit root. But the DF methodology suffers from a restrictive assumption that the error processes are *i.i.d.* When economic time series exhibit heteroskedasticity and non-normality in raw data, PP non-parametric tests are preferable to the DF and ADF tests." If there is a unit root in the variables, such variables are said to be non-stationary and the estimated relationship would be spurious. The ADF and PP unit root tests have been conducted on both in levels and first-differences for all the three variables.

The results of the unit-root tests are given in Tables 3 and 4, and regardless of whether it is the ADF or PP tests, the null hypothesis of a unit root for all the variables in levels is upheld. However, taking first differences of all the variables display stationarity under both the tests. Thus, the variables are all $I(1)$ and their first differences are $I(0)$. They are integrated of order 1.

The next step is to conduct the cointegration tests namely, the EG test and the JJ test. The first step of the EG test involves estimating equation (12) in OLS and the empirical estimates are the following:

$$\ln M = \underset{}{-6.1169} \; \underset{(53.35)}{+1.3415 \ln Y_t} \; \underset{(-12.35)}{-0.8234 \ln P_t} \qquad (14)$$

The figures in parenthesis are the respective *t*-statistics. We then check for stationary of the residuals from equation (14) by performing unit-root tests. The ADF test statistic is -3.1712 with the probability of 0.0234. Therefore, the null hypothesis of a unit root cannot be rejected at 1% level but at 5% significance level. However, the PP test statistic is -11.24 (p-values = 0.0000), and the null hypothesis is rejected even at 1% level. We conclude that there do not exist a unit root in the residuals and the variables are cointegrated of order one. Equation (14) above is the long-run relationship between the real quantity of imports and its major determinants, the GDP and relative prices.

Before undertaking the JJ cointegration test, we need to specify the relevant order of lags (p) for the VAR model. Given that the frequency of the data is quarterly, $p = 4$ is a reasonable choice. The results of the test are presented in Table 5.

At 5% level, the null hypothesis of $r \le 0$ in the trace test is rejected (p-values = 0.0371). Similarly, in the maximum eigenvalue test, the null hypothesis of $r = 0$ is also rejected at 5% level (p-values = 0.0173). Other null hypotheses in the two tests can only be rejected at 50% or higher level, suggesting that r = 1. Thus, it is concluded there is only one cointegration relationship among the variables. The estimates of the cointegrating vector are given as follows:

$$\ln M = \; -7.1950 \; +1.4338 \; \ln Y_t \; -0.7068 \; \ln P_t \qquad (15)$$

$$\qquad\qquad (28.97) \qquad (-5.26)$$

The conclusion in the JJ test is the same as in the EG test, confirming that there is a unique and equilibrium long-run relationship of Australian aggregate import demand with its major determinants of relative import prices and real GDP. The figures in parenthesis are the respective t-statistics.

Dutta and Ahmed (1999) and (2004) followed Hendry's (1979) general-to-specific approach to estimate the ECM for import demand. They first included 4 lags of the explanatory variables and 1 lag of the error correction term, and then gradually eliminate the insignificant variables. The error correction term is estimated from the cointegration equation by EG and JJ procedures. A similar approach is implemented to estimate the ECM for Australian import demand.

The general form of the ECM is written as follows:

$$\Delta \ln M_t = \beta_0 + \sum_{i=1}^{4} \beta_{1_i} \Delta \ln M_{t-i} + \sum_{i=0}^{4} \beta_{2_i} \Delta \ln Y_{t-i} + \sum_{i=0}^{4} \beta_{3_i} \Delta \ln P_{t-i} + \beta_4 EC_{t-1} + \varepsilon_t \quad (16)$$

where EC_{t-1} is error-correction term lagged one period. After experimenting with the above general form of the ECM, the following equation of the ECM is found to fit the data best.

For the EG procedure:

$$\begin{aligned} \Delta \ln M_t = \; & 0.0036 + \underset{(7.55)}{0.4686}\ \Delta \ln M_{t-4} + \underset{(3.85)}{0.2017}\ \Delta \ln Y_t - \underset{(-2.56)}{0.3457}\ \Delta \ln P_{t-3} \\ & \underset{(-5.53)}{-0.3059}\ EC_{t-1} + \underset{(4.47)}{0.3569}\ \varepsilon_{t-1} \end{aligned} \quad (17)$$

For the JJ procedure:

$$\begin{aligned} \Delta \ln M_t = \; & 0.0041 + \underset{(7.56)}{0.4713}\ \Delta \ln M_{t-4} + \underset{(3.86)}{0.2089}\ \Delta \ln Y_t - \underset{(-2.58)}{0.3505}\ \Delta \ln P_{t-3} \\ & \underset{(-5.44)}{-0.2960}\ EC_{t-1} + \underset{(4.44)}{0.3540}\ \varepsilon_{t-1} \end{aligned} \quad (18)$$

where ε_{t-1} is moving average lagged one period, which is included to account for seasonality discussed earlier. Actually, we have two variables in our equation to account for seasonality, $\Delta \ln M_{t-4}$ and ε_{t-1} and both are statistically significant. The figures in parenthesis are the respective *t*-statistics.

All estimated coefficients are statistically significant at the 5% level (or better) using the t-test and jointly significant using the F-test. For the EG procedure, we have $\overline{R}^2$ of 0.3130, DW-Stat of 2.0203 and F-statistic of 17.5872; and for the JJ procedure, getting $\overline{R}^2$ of 0.3086, DW-Stat of 2.0305 and F-statistic of 17.2461. Further diagnostic test statistics show no evidence of misspecification of functional form, no serial correlation, no problem of heteroskedasticity, and normality. Details of these diagnostic tests are given in Table 6. Our model satisfies all diagnostic tests even at 10% level (the null hypothesis of these diagnostic tests should not be rejected for a plausible model), suggesting the results of the current study gives further insight into explaining the behaviour of Australian import demand.

Alternatively, we can use the UECM to estimate the long-run and short-run relationship at the same time. If the long-run relationship is valid, the UECM will reveal the dynamic behaviour of import demand. The form of the UECM can be written as follows:

$$\Delta \ln M_t = \beta_0 + \sum_{i=1}^{4} \beta_{1_i} \Delta \ln M_{t-i} + \sum_{i=0}^{4} \beta_{2_i} \Delta \ln Y_{t-i} + \sum_{i=0}^{4} \beta_{3_i} \Delta \ln P_{t-i}$$
$$+ \beta_4 \ln M_{t-1} + \beta_5 \ln Y_{t-1} + \beta_6 \ln P_{t-1} + \varepsilon_t$$

$$\Delta \ln M_t = -1.9824 + \underset{(7.50)}{0.4694}\ \Delta \ln M_{t-4} + \underset{(3.84)}{0.2086}\ \Delta \ln Y_t - \underset{(-2.48)}{0.3419}\ \Delta \ln P_{t-3}$$
$$\underset{(-5.53)}{-0.3090}\ \ln M_{t-1} + \underset{(5.38)}{0.4228}\ \ln Y_{t-1} - \underset{(-4.17)}{0.2507}\ \ln P_{t-1} + \underset{(4.43)}{0.3560}\ \varepsilon_{t-1} \quad (19)$$

To test for cointegration in the equation (19), we use the bound test and the figures in parenthesis are the respective t-statistics. This is the procedure of using the Wald Test (F-statistic) to test the null hypotheses of $H_0 : \beta_4 = \beta_5 = \beta_6 = 0$. The F-statistic of 10.23 with a p-value of 0.00001 provides sufficient evidence to reject the null hypotheses in favour of the alternative. This outcome confirms that there is a cointegrating relationship among the variables in the Australian import demand function. This is also consistent with the results of the other two cointegration tests. Thus, we have the following long-run relationship determined from the empirical results estimated in equation (19):

$$\ln M = \ -6.4147 \ \ +1.3681 \ \ln Y_t \ \ -0.8111 \ \ln P_t \qquad (20)$$

Comparing equations (14) (15) and (20) and equations (17) (18) and (19) respectively, there are insignificant differences in the coefficient estimates, which is due to rounding errors as a result of using different procedures. More specifically there are marginal differences between equations (14) and (20), and equations (17) and (19) respectively. The estimated results of equation (14) are used to derive equation (17), from long-run to short-run, which is also known as forward procedure. Conversely, equation (20) is derived from the estimates of equation (19), from short-run to long-run which is also referred to as backward procedure. On the other hand, the difference between equation (14) and (15) which leads to the difference between equation (17) and (18) is due to statistical error. Equation (14) is estimated by Ordinary Least Squares (OLS) which minimises the sum of squared residuals while equation (15) is estimated by MLE which maximises the log-likelihood function.

7 Implications of Empirical Results

A comparative summary of our estimates using the three different techniques is represented in Table 6 for the long-run relationship and in Table 7 for the short-run relationship. This summary enables us to cross-check the consistency among the different cointegration techniques that have been implemented in this paper. Overall, the elasticities are much higher in the long-run than in the short-run. In the long-run, the two independent variables, the real GDP and the relative import prices are found to be the main determinants of Australian aggregate import demand with R^2 of 98%.

However, in the short-run (1 quarter), the effect of the two variables is diluted by other factors with R^2 of 33%.

Income is elastic in the long-run and inelastic in the short-run while price is inelastic in both the long-run and short-run. Moreover, the income elasticity is greater than the price elasticity in the long-run but not in the short-run, suggesting that price (3 periods lagged) is the most significant determinant of Australian import demand in the short-run while income is the most influential factor in the long-run. Indeed, individual preference to impulse purchasing tends to be determined by the price of goods and services. However, in longer time horizon, it tends to be determined by income.

The long-run income and price elasticities are in line with the Goldstein-Khan (1985) ranges of (1.0, 2.0) for typical income elasticity and (-0.50, -1.00) for typical price elasticity. The income elasticity is significantly greater than unity even at 1% level, owing that there is a degree of trade-off between economic growth and the trade balance. As the result, Australian balance of payments is likely to worsen with high economic growth; which is quite evident in the recent Australian economic history.

The short-run price elasticity is about 0.35 with *p*-value of 0.0141, suggesting that it is the only factor which has a reasonable effect on Australian import demand in the short-run. Therefore, to some degree the government can effectively use exchange rate policies to improve its short-term current account balance.

The estimated coefficient of error correction term is 0.3090 and statistically significant even at 1% level with the appropriate sign. This result validates the long-run equilibrium relationship among the variables. Moreover, the system tends to correct its previous period's disequilibrium by 31% a quarter; and it takes about 3.3 quarters or approximately 10 months to fully realign any disequilibrium that arises.

8 Conclusion

This paper estimates the aggregate import demand function for Australia over the period 1959Q3–2006Q3. Cointegration and error correction modelling approaches is used to estimate the long-run as well as the short-run relationships among the variables. This is the only assessment for Australian import demand which provides a precise estimate for the short-run relationship, especially estimation of the short-run adjusting term. The real quantities of aggregate import demand, relative import prices and real GDP are found to be not stationary but cointegrated of order one with only one cointegration relationship. Thus, there is unique long-run equilibrium relationship for Australian import demand. That is, income and prices are the plausible factors that affect import demand function.

The findings suggest the dominance of income factor in the long-run and price factor in the short-run for determining quantity of Australian import demand. So there should be distinct policy prescriptions relating Australian international trade over different time horizons. Furthermore, for the adjusting term of 0.3090, it takes approximately 10 months for the system to fully realign any disequilibrium from the long-run relationship. Consistent findings from the study will provide policy-makers further insight on how to improve the trade balance deficit.

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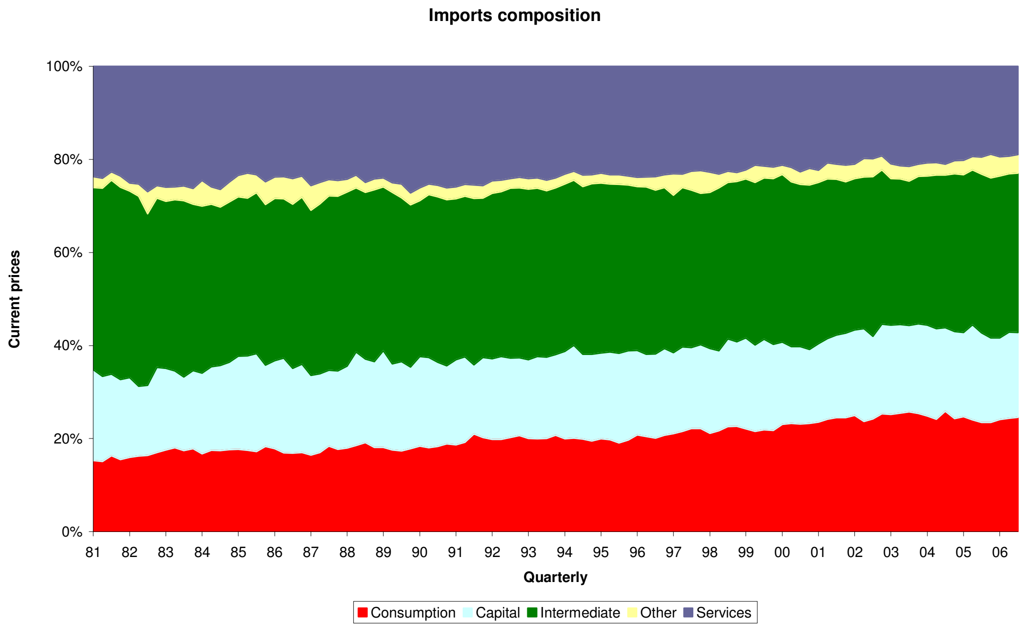
Figure 1: Composition of Australian imports



Source: Reserve Bank of Australia, Table H03 Exports and Imports of Goods and Services.

Figure 2: Natural logarithm of real quantity of import demand

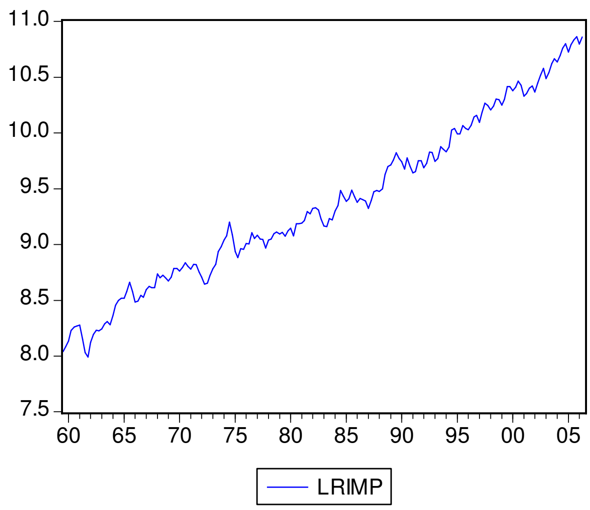


Figure 3: Natural logarithm of real gross domestic product

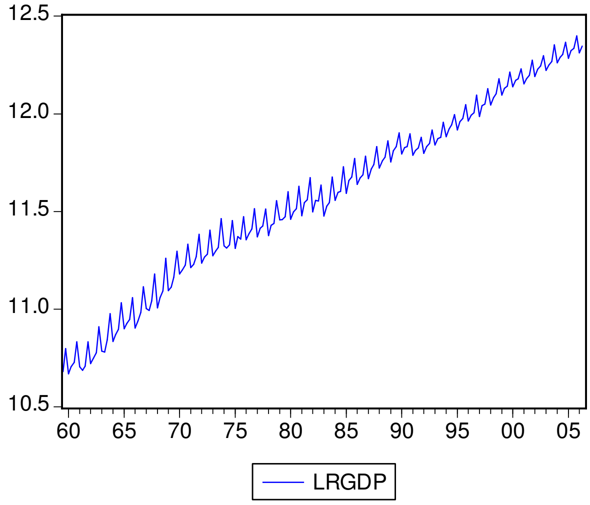


Figure 4: Natural logarithm of relative prices

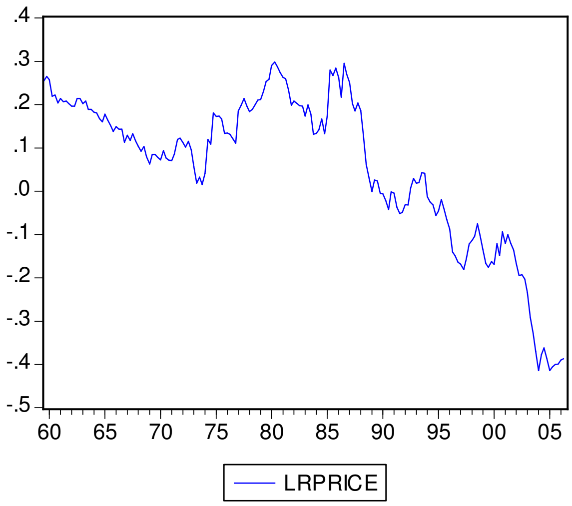


Table 1: Summary statistics of variables

Variable	Mean	SD	Skewness	Kurtosis
LRIMP	9.3708	0.7518	0.2101	2.0736
LRGDP	11.5785	0.4686	-0.1698	2.0580
LRPRICE	0.0547	0.1767	-0.9746	3.2817

Table 2: Seasonality

Variable	D1	D2	D3	D4
LRIMP	9.3556	9.3801	9.3787	9.3687
	(84.64)	(84.86)	(84.85)	(84.76)
LRGDP	11.5367	11.5702	11.5534	11.6534
	(168.21)	(168.69)	(168.45)	(169.91)
LRPRICE	0.0504	0.0497	0.0629	0.0561
	(1.94)	(1.91)	(2.42)	(2.16)

Notes: The figures in parenthesis are *t*-statistic.

Table 3: Augmented Dickey Fuller Unit Root Tests for Stationarity

Variable	Levels/First Differences	Test Statistic	Critical Value (at 5% level)	*p*-value
LRIMP	Level	0.9163	-2.8774	0.9956
	First diff.	-8.3603	-2.8774	0.0000
LRGDP	Level	-1.9922	-2.8775	0.2902
	First diff.	-4.3974	-2.8775	0.0004
LRPRICE	Level	0.2549	-2.8768	0.9753
	First diff.	-12.2495	-2.8769	0.0000

Notes: Null hypothesis: The variable has a unit root.

Table 4: Phillips-Perron Unit Root Tests for Stationarity

Variable	Levels/First Differences	Test Statistic	Critical Value (at 5% level)	*p-value*
LRIMP	Level	0.0173	-2.8768	0.9582
	First diff.	-12.8853	-2.8769	0.0000
LRGDP	Level	-1.0220	-2.8768	0.7452
	First diff.	-68.2008	-2.8769	0.0001
LRPRICE	Level	0.0650	-2.8768	0.9622
	First diff.	-12.2495	-2.8769	0.0000

Notes: Null hypothesis: The variable has a unit root.

Table 5: JJ maximum likelihood cointegration tests

Null	Alternative	Statistic	5% critical value	Probability
Trace test				
$r = 0$	$r \geq 0$	30.9073	29.7971	0.0371
$r \leq 1$	$r \geq 1$	6.6110	15.4947	0.6233
$r \leq 2$	$r \geq 2$	0.0103	3.8415	0.9189
Max eigen test				
$r = 0$	$r = 1$	24.2963	21.1316	0.0173
$r = 1$	$r = 2$	6.6007	14.2646	0.5374
$r = 2$	$r = 3$	0.0103	3.8415	0.9189

Table 6: Diagnostic tests

Diagnosis	Test	Null hypothesis	χ^2-stat		p-value	
			EG	JJ	EG	JJ
1. Functional Form	Ramsey's RESET Test (a)	Functional form is well specified	0.0795	0.0027	0.7780	0.9587
2. Serial Correlation	Breuch-Godfrey LM Test	No serial correlation in the residuals up to the specified order	2.2517	2.9583	0.6896	0.5648
3. Heteroskedasticity	White's Test (b)	No Heteroskedasticity	10.3900	12.5670	0.2387	0.1276
4. Normality	Jarque-Bera Test	Normality	0.6326	0.4513	0.7288	0.7980

Notes: [a] see White (1980)
[b] see Ramsey (1969)

Table 7: Estimates of the long-run relationship

Variable	Engle-Granger Procedure	Johansen Procedure	UECM
GDP	1.3415	1.4338	1.3681
PRICE	-0.8234	-0.7068	-0.8111

Table 8: Estimates of the short-run relationship

Variable	Engle-Granger Procedure	Johansen Procedure	UECM
GDP	0.2017	0.2089	0.2086
PRICE	-0.3457	-0.3505	-0.3419
Adjustment	-0.3059	-0.2960	-0.3090