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TITLE

IMPORT DEMAND WITH DOMESTIC PRICE ENDOGENEITY: THE SOUTH AFRICAN CASE

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ABSTRACT

The "elasticities" approach based on the Bickerdike-Robinson-Metzler (BRM) condition is widely used to estimate the effect of an exchange rate depreciation on the trade balance. However this approach is limited by its partial equilibrium nature and, in particular, fails to account for the inflationary effects that a depreciation has on the domestic price. This can therefore result in the BRM condition exaggerating the effect of exchange rate changes on import demand. This paper uses the imperfect substitutes model to derive an import demand function. However, here domestic prices are endogenised within a simultaneous framework in order to assess the full effect of a currency depreciation. The Johansen multivariate cointegration approach is used to estimate the import demand model as it accounts for nonstationary data and allows simultaneity between the variables. Prior to its use, the "small country assumption", which allows for import price exogeneity, is tested for South Africa. Two differing tests indicate that this assumption holds although further exploration reveals that there is scope to clarify the issue. Ultimately the paper shows that import price elasticities generally found in import demand curves (stemming from the elasticities approach) may significantly overstate the effect of an exchange rate depreciation on the trade balance. This is because the BRM condition explicitly ignores the domestic inflationary effects of an exchange rate depreciation.

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1. INTRODUCTION

There has been an increasing focus on trade and the balance of payments in South Africa due in particular to liberalization, the removal of trade barriers, the shift towards export orientation and the effects of a fluctuating exchange rate on the trade balance. Outward-looking strategies and export-oriented policies increasingly look to the exchange rate as a method of improving both firm profitability and the country's current account balance. The debate surrounding the exchange rate effect has thus come to the fore with a focus on the export side. As a result there remains scope for further research to clarify the debate surrounding the magnitude of price and income elasticities of import demand as well as the final effect of exchange rate fluctuations on import demand.

Much of the literature surrounding exchange rate effects on the trade balance is drawn from the "elasticities approach", which is based on the Bickerdike-Robinson-Metzler (BRM) condition. The BRM condition characterises a set of conditions on the size of import and export demand and supply price elasticities that result in an improvement of the trade balance. However, because of its partial equilibrium nature this approach fails to account for possible general equilibrium effects and as such can exaggerate the extent to which import demand responds to a fluctuation in import prices. In particular, exchange rate fluctuations may not only affect import prices but can also pass-through to domestic prices, reducing the incentive to select domestic over imported products. The extent to which domestic prices are affected by import prices generally rises the more open an economy becomes. This is particularly relevant to South Africa which has experienced an un-precedented period of economic liberalisation. In addition, for "small" countries that are unable to affect international market prices the effect of an exchange rate depreciation on domestic prices may be relatively greater than for countries categorised as large economies.

This paper makes a number of contributions. Firstly, a critical evaluation of the elasticities approach shows the necessity to account for general equilibrium effects, particularly the possible effects of the exchange rate on the domestic price. Secondly, using the multivariate Johansen cointegration technique, we estimate an import demand function for South Africa with endogenous domestic prices in a simultaneous equation model. This allows a more accurate understanding of the relative price effects on import demand. We are able to show that the net effect of an exchange rate depreciation on import demand is significantly lower

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once domestic price inflationary effects have been accounted for. Thirdly, the assumption of an infinite world supply of imports (which implies that South Africa is a "small" price-taking economy and allows for import price exogeneity) is tested for South Africa using a purchasing power parity test and an exchange rate pass-through model. Literature on the small country test is relatively limited in South Africa and the results show that the assumption upon which many trade studies base their estimation seems to hold for South Africa once oil price effects are accounted for. The result is additionally beneficial for this study as it simplifies analysis by 1) allowing the single equation estimation of import demand rather than a simultaneous approach involving import supply and 2) allowing the import price effect on domestic prices to be separated into exchange rate and foreign price channels.

The paper proceeds as follows: Section 2 identifies the background theory for the elasticities approach and its associated drawbacks. Section 3 describes the modelling of import demand with domestic price endogeneity and specifies the two models used to test the small country assumption. Section 4 looks at the cointegration approach used in this paper, while Section 5 looks specifically at the data used in the model. Sections 6 and 7 provide the results for the small country test and import demand framework respectively with Section 8 drawing on the findings.

2. THEORETICAL BACKGROUND

THE ELASTICITIES APPROACH

The most common approach to modelling the relationship between exchange rates and the trade balance is based on the BRM model. Empirical macroeconomic models employing import demand equations that are a function of demand and relative prices are based on this condition and are commonly termed as using the "elasticities approach". The BRM model is a partial equilibrium, two-good model which uses imperfect substitutes to illustrate the effect exchange rate changes may have on the balance of trade (with the balance of trade simply equalling exports less imports). The illustration of the BRM model using the trade balance in domestic currency *(B)* is taken from Isard (1995), with the volume of exports supplied (X^s), imports demanded (M^d) and the associated prices in domestic currency (P_X, P_M) determining the value of trade.

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With the trade balance simply seen as net exports:

$$B = P_X X^s - P_M M^d \tag{1}$$

the effect of a change in the exchange rate can be demonstrated by differentiating and converting to elasticities:

$$\frac{dB}{dE} = P_X X^s \left[\frac{(1+\varepsilon)\eta^*}{(\varepsilon + \eta^*)} \right] - P_M M^d \left[\frac{(1-\eta)\varepsilon^*}{(\varepsilon^* + \eta)} \right] \tag{2}$$

where E is identified as the exchange rate (domestic currency per unit of foreign currency). The absolute values for the import demand and export supply price elasticities are given by η and ε respectively. The asterisk denotes the foreign country. Implicitly, the exchange rate effect on the trade balance is determined by the change in the price of exports and imports and the response of export supply and import demand to these price changes. The size of this response is determined by the price elasticities of supply and demand. With the assumption of an initial equilibrium in the trade balance ($B = 0$), it is shown that a depreciation in the exchange rate will only improve the trade balance if the following condition (the BRM condition) holds:

$$\frac{\eta\eta^*(1+\varepsilon+\varepsilon^*) - \varepsilon\varepsilon^*(1-\eta-\eta^*)}{(\varepsilon+\eta^*)(\varepsilon^*+\eta)} > 0 \tag{3}$$

Of the four generalised results within the BRM condition (the small country case, elastic supply, inelastic demand and elastic demand) the small country case is often considered the most appropriate for South Africa. A small country is deemed to be a price-taker in both its import and export markets, measured in foreign currency. This would imply that both the foreign supply of exports (ε^*) and the foreign demand for imports (η^*) is infinitely elastic, reducing the BRM condition in Equation *(3)* to:

$$\varepsilon - \eta > 0 \tag{4}$$

when the trade balance is measured in domestic currency. Assuming that the domestic export supply and import demand curves behave normally (in that they are positively sloping and

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downward sloping respectively), the stated result implies that a depreciation in the exchange rate unambiguously improves the trade balance. A devaluation causes domestic currency prices of exports and imports to rise by the complete value of the currency devaluation. This causes the volume of exports to rise and the volume of imports to decline. The combination of price and volume changes results in an improvement of the trade balance. Relaxing the assumption of an elastic import demand curve alters only the size and not the overall conclusion of an improvement in the trade balance. This is because the resultant rise in both export price and volume outweighs the depreciation effect on imports. The small country (or infinite world supply) assumption implies that changes in the exchange rate and foreign price completely feed through into import prices.

Another case that may be pertinent to South African trade is the infinite elasticity of domestic and foreign export supply, resulting in the famous Marshall-Lerner condition. With an infinite domestic (ε) and foreign (ε *) supply of exports, the BRM condition simplifies to:

$$\eta + \eta^* > 1 \qquad (5)$$

Thus, in order for a depreciation to improve the trade balance either domestic or foreign import demand must be strictly elastic. However sticky prices, lags in response to exchange rate changes, a high dependence on foreign resource commodities or a lack of import substitutes can all contribute to the short-run inelasticity of import demand (Edwards and Wilcox, 2003). With inelastic import demand, the results of an exchange rate depreciation become more ambiguous, with the possibility that a currency depreciation worsens the trade balance.

DOMESTIC PRICE ENDOGENEITY

While the elasticities approach above provides a neat and fairly simple approach to identifying exchange rate effects on the trade balance, it is not without drawbacks. As Isard (1995) states, the elasticities approach ignores macroeconomic distortions that can contribute to the impact exchange rate fluctuations may have on the trade balance. The problems associated with this model arise mainly from the fact that it is a partial equilibrium framework. Other models, such as the monetarist and absorption approaches take into account this problem. In particular, the results derived above exclude the possible effects of import price changes (through foreign currency price or exchange rate changes) on the

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domestic price, which can be understood as pass-through effects. The pass-through to domestic prices can affect relative prices and therefore the demand for imports in both the long and short run.

The manner and extent to which domestic prices are influenced by import prices can be separated into supply, demand and globalisation effects. On the supply side, the higher the elasticity of factor prices with respect to domestic prices, the greater the pass-through of import price changes to domestic prices. Conversely, the larger the elasticity of domestic prices with respect to factor prices, the greater the response of domestic prices to import price changes (Goldstein and Khan, 1985). Also, the greater the use of intermediate imports in the production of domestic goods, the more responsive domestic prices will be to import prices (Edwards and Wilcox, 2003). This can be seen as the direct effect changes in import prices (due either to exchange rate fluctuations or foreign price shocks) have on the cost of production of domestic goods.

On the demand side, the extent of influence import prices have on domestic prices is determined by the substitutability between imported and domestic goods as well as the share of imports in the economy (Goldstein and Khan, 1985). Greater substitutability between domestic and imported goods has two effects: 1) it raises price competitiveness between imported and domestic goods and 2) it implies that a rise in import prices would induce a greater demand for domestic products, thereby placing upward pressure on domestic prices (Edwards and Wilcox, 2003). In addition, for "small" countries that face an infinite world supply of imports (and where a currency depreciation would have no effect on world prices) the effect of a depreciation on domestic prices is greater than would be the case for a large country, where the inflationary effect of a depreciation would be countered by a decline in the world price (McCarthy, 2000).

Mihaljek and Klau (2001) note that import prices may play an increasing role in the setting of domestic prices as firms in emerging economies begin to follow more closely the developments in foreign prices; a result of increasing globalisation, liberalisation and the shifting of production from industrialised to developing countries. The result is that cross-border price comparisons as well as inflation expectations can cause foreign prices to affect the prices of domestic goods.

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As a result domestic prices may be endogenously determined by both exchange rates and foreign prices due to possible pass-through effects. The elasticities approach (derived from the BRM condition) is distinctly limited by not accounting for domestic price changes following a currency depreciation. This paper makes an important contribution by attempting to account for the inflationary impact of a depreciation on domestic prices and import demand. In the following section the model used to estimate the import demand elasticities and the pass-through to domestic prices is outlined.

3. MODEL SPECIFICATION

The majority of import demand studies are derived from the imperfect substitutes model where there is imperfect substitution between domestic and imported goods and the consumer is subject to a budget constraint (Goldstein and Khan, 1985). From a theoretical point of view the relationship between the price and quantity of imports is determined by the simultaneous interaction of supply and demand. Within this framework domestic import demand is seen as a function of relative prices and real income (Houthakker and Magee (1969), Murray and Ginman, (1976), Carone (1996)), while the foreign supply of imports is seen as a function of the price of foreign versus exported goods (Goldstein and Khan, 1985). The imperfect substitutes model for imports can be illustrated as follows:

$$M^d = f[Y, P_D, P^*)] \qquad (6)$$

$$M^S = g\left[P^f, \frac{P^*}{E}, \Phi\right] \qquad (7)$$

$$E \cdot M^S = M^d \qquad (8)$$

where import demand, M^d, is seen as a function of real income, Y, the price of domestically produced goods, P_D, and import prices, P^*. All variables in the import demand equation are given in domestic currency. The foreign supply of imports, M^S, is a function of the price of foreign goods in foreign currency (P^f), the price of imports in foreign currency and a range of other variables that can affect supply, Φ. E is the exchange rate given in units of domestic currency per unit of foreign currency. Both the import and supply equations can be extended to account for tariff barriers and subsidies as is done by Goldstein and Khan (1985).

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IMPORT DEMAND

Equation *(6)* is most often represented in the log-linear form as Khan and Ross (1976) show that "estimating convenience and the correct functional form do coincide" by using a log-linear specification for import demand, allowing easier estimation of elasticity coefficients. Carone (1994) describes the standard model as:

$$LnM_t^d = \alpha_0 + \alpha_1 LnY_t + \alpha_2 Ln(\frac{P_t^*}{P_{Dt}}) + \mu_t \qquad (9)$$

with *Ln* the natural logarithm. Here real income is sometimes referred to as the scale or activity variable. Carone (1994) indicates that the restricted equation is based on the assumption of the absence of "money illusion". This implies that the demand function is homogenous of degree zero in prices and income. A single relative price variable is therefore used to explain why agents switch between imported and domestic goods. Murray and Ginman (1976) argue that the import demand equation should rather be specified as below in Equation *(10)* since price homogeneity is then not assumed prior to estimation. This results in separate elasticities for domestic and imported prices:

$$LnM_t^d = \alpha_0 + \alpha_1 LnY_t + \alpha_2 LnP_{Dt} + \alpha_3 LnP_t^* + \mu_t \qquad (10)$$

Using microeconomic foundations, Xu (2002) and Senhadji (1997) ultimately show the validity of the standard import demand equation estimated by equation *(9)*, with subtle differences in the activity (scale) variable used and the inclusion of a variable in an attempt to capture the dynamic effect in import demand. Thus, the majority of studies estimate import demand using equation *(9)* as the benchmark, adding variables according to the nature of the study[1].

This paper removes the assumption of price homogeneity and follows Murray and Ginman's (1976) specification. In this way the import price elasticity can be measured separately from the domestic price effect in the import demand model and the extent of domestic price endogeneity can be fully identified.

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[1]For examples of a different import demand specification relating to foreign exchange constraints see Islam and Hassan (2004) and Dutta and Ahmed (1999).

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The model is modified and follows Goldstein and Khan (1985) by including tariff rates to identify barrier effects:

$$m_t^d = \alpha_0 + \alpha_1 y_t + \alpha_2 p_{Dt} + \alpha_3 p_t^* + \alpha_4 tariff_t + \mu_t \qquad (11)$$

where, *tariff* is one plus the tariff rate, with lower case letters representing the natural log of the variables.

INFINITE IMPORT SUPPLY

Specification

As earlier mentioned the estimation of import demand elasticities implies a simultaneous equation framework (of import demand and supply) as specified by theory. Yet the standard practice has been to assume an infinite supply-price elasticity for imports (Goldstein and Khan, 1985), sometimes seen as the "small country assumption". This can be demonstrated by expanding Equation *(7)* using a log-linear form:

$$m^S = \sigma_0 + \sigma_1 p^f + \sigma_2 p^* - \sigma_2 e + \sigma_3 \phi \qquad (12)$$

where again lower case letters represent the natural log of variables. Re-arranging,

$$p^* = \frac{1}{\sigma_2} m^S - \frac{\sigma_0}{\sigma_2} - \frac{\sigma_1}{\sigma_2} p^f + \frac{\sigma_2}{\sigma_2} e - \frac{\sigma_3}{\sigma_2} \phi \qquad (13)$$

Infinite import price elasticity implies that $\sigma_2 \to \infty$. Therefore the coefficients on m^S and ϕ in Equation *(13)* will tend towards zero. The assumption of price homogeneity (the absence of money illusion) in import supply is made, implying that the sum of coefficients on p^* and p^f will equal zero i.e. $\sigma_1 = -\sigma_2$. Equation *(13)* can therefore be simplified to:

$$p^* = p^f + e \qquad (14)$$

This is sometimes considered the "law of one price" and implies that the price of imports will be the same in both domestic and foreign countries, when expressed in a common currency (Dwyer, Kent, Pease, 2001).

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Equation *(14)* can be re-formed as:

$$P^{*} = P^{f} \times E \tag{15}$$

Thus any change in either the exchange rate or foreign price is fully reflected by a change in the import price. With this assumption, the price of imports (P^*) can be viewed as exogenous and implies that South Africa is a price-taker on the world market in the long-run. This therefore implies that only the quantity of imports demanded need be determined endogenously and allows for single equation estimation of import demand elasticities. The small-country assumption is tested here using two methods. A cointegration model is used to test for complete exchange rate pass-through to import prices and a modified purchasing power parity (PPP) test is performed on import prices to indicate exchange rate pass-through.

Exchange Rate Pass-Through

Exchange rate pass-through to import prices can be defined as the percentage change in local currency import prices resulting from a one percent change in the exchange rate. With respect to imports, "the degree of pass-through will increase the greater is the elasticity of supply", where complete pass-through implies a perfect (infinite) elasticity of supply and the country is a price taker on the world market (Dwyer et al, 2001). Tariff barriers may contribute to import price changes (Rangasamy, 2003), given the higher cost faced by importers, and Equation *(15)* may be transformed to include tariffs, T:

$$P^{*} = P^{f}(1+T) \times E \tag{16}$$

From this the following model is estimated:

$$p^{*} = \theta_1 p^{f} + \theta_2 e + \theta_3 tariff + \varepsilon_t \tag{17}$$

Complete exchange rate pass-through occurs where the coefficient on the exchange rate, θ_2, is equal to 1 and is a sufficient test for the small country assumption (Menon 1995b).

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The PPP Test

The standard test for purchasing power parity (PPP) is performed by testing the real effective exchange rate for stationarity. In a similar manner one can test complete pass-through in import prices by testing for stationarity in the effective exchange rate for imports. If PPP holds the implication is that import price changes are fully explained by a change in either the exchange rate or the foreign price. Equation *(15)* therefore holds and exchange rate pass-through to import prices is complete. Banerjee, Marcellino and Osbat (2005) describe both a weak and strong test for PPP. The real effective exchange rate, *REER*, can be defined as:

$$REER = \frac{P^*}{E \times P^f} \qquad (18)$$

simply seen as the relative price of imported goods in the same currency. The strong test for PPP tests the natural log of *REER* for stationarity:

$$reer = p^* - e - p^f \qquad (19)$$

If the null hypothesis of a unit root in *reer* is rejected (i.e. the series is stationary) then PPP is said to hold. The weaker test for PPP removes the coefficient restriction of unity on the price variables and is intended to capture differences in the composition of prices indexes across countries, transaction and transportation costs and the possible effects of tariffs (Banerjee et al, 2005). The coefficients on the price variables can be estimated through single or multiple equation estimation. The weak test for PPP implies a form of cointegration analysis on import prices. Considering that a pass-through test is also done in this paper, only the strong form of PPP is tested to indicate whether South Africa can be considered a small country.

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DOMESTIC PRICES

As mentioned earlier, the need for endogenising domestic prices arises from the general equilibrium effects that foreign prices and exchange rates may have on domestic prices. A domestic price equation is therefore modelled to identify the effect these two variables may have. When dealing with these pass-through effects to domestic prices, it is found that the common approach has been to identify domestic prices as a function of import prices and a range of other variables, Z, that may affect the domestic price[2]:

$$P = f(P^{*}, Z) \tag{20}$$

Similar to that of Fedderke and Schaling (2005), the domestic price model is seen as function of wage pressures on domestic prices and excess domestic demand. This model is modified to include import price effects through the exchange rate and foreign price channels. The final domestic price equation is specified as:

$$p_{Dt} = \gamma_1 p_t^f + \gamma_2 e_t + \gamma_3 ulc_t + \gamma_4 ydgap_t + \gamma_5 tariff_t + v_t \tag{21}$$

where ulc represents (log) unit labour costs in South Africa and $ydgap_t$ denotes excess demand for domestic goods. Tariffs are included following Rangasamy's (2003) analysis of the tariff effect on both import and domestic prices.

In summary, an import demand function is estimated using Equation *(11)* with domestic prices made endogenous in a simultaneous framework using Equation *(21)*. Import prices are made exogenous by assuming an infinite supply of imports and this assumption is tested using the PPP test of stationarity (Equation *(19)*) and an exchange rate pass-through test (Equation *(17)*). The following section deals with the cointegration technique used in estimating both the import demand function (with endogenous domestic prices) and the exchange rate pass-through estimation for import price.

4. ECONOMETRIC METHODOLOGY

Cointegration techniques addressing the deficiencies of Ordinary Least Squares (such as the assumption of an equilibrium state always being maintained and its inability to deal effectively with time series) have helped develop more accurate estimates of import demand elasticities. The most popular of these are the single equation approaches of Engle-Granger and Pesaran's bounds test procedure, and the multivariate approach developed by Johansen and Juselius (1990) known as the full information maximum likelihood (ML) approach. The technique best suited to this paper is Johansen's ML approach as it allows for a system of

[2] See Ambler, Dib and Rebei (2003), Choudri and Hakura (2001) and McCarthy (2000) for micro-foundation models of pass-through to domestic price.

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equations with endogenous variables to be modelled in a simultaneous approach. A short outline of the approach, taken from Harris and Sollis (2003), is given below. Prior to performing this technique the order of integration of each time series needs to be verified to ensure that all endogenous variables are $I(1)$ at most, as the Johansen technique needs to be substantially modified in order to be used for endogenous variables that are integrated of order $I(2)$[3].

The maximum likelihood procedure has a data generating process that is represented as a p order vector autoregressive model:

$$z_t = A_1 z_{t-1} + \ldots + A_p z_{t-p} + \Psi D_t + \mu + \varepsilon_t \quad t = 1,\ldots,T; \quad \varepsilon_t \text{ is niid}(0, \Sigma) \tag{22}$$

where z_t is the vector of $I(1)$, endogenous variables, x_t, D_t is a vector of $I(0)$, exogenous (deterministic) variables and ε_t represents a white noise process. The process can also be parametized as a Vector Error Correction Model (VECM), helping to identify and account for the short run dynamics of the model:

$$\Delta z_t = \Pi z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta z_{t-i} + \Psi D_t + \mu + \varepsilon_t \tag{23}$$

where Π is a long-run multiplier matrix, Γ are the short-run coefficient matrices for each period i and Ψ is the coefficient matrix on the $I(0)$ exogenous variables. The rank of Π is determined by the number of cointegrating relations, r, between the $I(1)$ variables. Two tests developed by Johansen (1988) for the rank of Π, the Trace statistic and the Maximal Eigenvalue (λmax) statistic, are used to determine the number of cointegrating vectors within the model. If there are r cointegrating relations between the $I(1)$ variables it implies that there are r combinations between the vector of non-stationary variables, z_t, that are stationary. The matrix Π can be thereafter represented as:

$$\Pi = \alpha\beta' \tag{24}$$

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[3] For more on Cointegrating $I(2)$ variables see Johansen (1995) and Johansen (2005).

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where α is the k x r loading matrix representing the speed of adjustment to disequilibrium and β' is the r x k matrix of cointegrating vectors indicating the long-run coefficients of the cointegrating relationships (Harris and Sollis, 2003).

A trend and intercept can enter both the long- and short-run model, depending on the type of trend seen in the data. If the trend enters the model as unrestricted, it implies a quadratic trend in the levels of endogenous $I(1)$ variables and a linear trend in the short-run dynamics of the variables. Generally this is considered economically implausible, more so when using the log of variables, as this implies a continuously increasing or decreasing rate of change. Thus the trend can be excluded from the long-run model or, where there is some long-run linear growth that the model is unable to account for, the trend can be restricted to the cointegrating vectors (Harris and Sollis, 2003). The VECM can also be parametized such that some variables, x_t, in the vector z_t are treated as weakly exogenous, implying that Δx_t contains no information about the long run coefficients of the endogenous variables. The model is therefore configured such that the weakly exogenous variable remains in the long-run model but is excluded from the short-run dynamics.

In order to develop an economically meaningful model from the cointegrating vectors, one can impose a number of identifying restrictions on the coefficients within each vector. There need to be r^2 restrictions on the β coefficients for the model to be exactly identified. To comply with model structure and economic theory we can further restrict the model imposing over-identifying restrictions on the cointegrating vectors' β coefficients. Using the simultaneous framework for import demand (Equation *(11)*) and domestic prices (Equation *(21)*), the long-run parameters can be represented as:

$$\Pi z_{t-1} = \begin{bmatrix} \alpha_{11} & \alpha_{21} \\ \alpha_{12} & \alpha_{22} \\ \alpha_{13} & \alpha_{23} \\ \alpha_{14} & \alpha_{24} \\ \alpha_{15} & \alpha_{25} \\ \alpha_{16} & \alpha_{26} \\ \alpha_{17} & \alpha_{27} \\ \alpha_{18} & \alpha_{28} \end{bmatrix} \begin{bmatrix} 1 & 0 & \beta_{13} & \beta_{14} & 0 & 0 & \beta_{17} & \beta_{18} \\ 0 & \beta_{22} & 1 & 0 & \beta_{25} & \beta_{26} & 0 & \beta_{28} \end{bmatrix} \begin{bmatrix} m^d \\ p^f \\ p^d \\ p^* \\ e \\ nulc \\ y \\ tariff \end{bmatrix}_{t-1} \qquad (25)$$

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As we would expect that r = 2, there would be two just-identifying restrictions on the long-run parameters for each equation as well as additional restrictions in order to conform with the model structure (given by the ones and zeros).

Over and above the standard error statistical inspections done on the coefficients, the log-likelihood ratio test is used to assess the validity of the over-identifying restrictions. If the null hypothesis cannot be rejected when testing the above restrictions, the model is deemed to be successfully identified.

5. DATA ANALYSIS

DATA SOURCES

This study makes use primarily of two databases: The International Monetary Fund's (IMF) International Financial Statistics (IFS) database for nominal exchange rates and foreign producer and consumer prices, and the South African Reserve Bank (SARB) Quarterly Bulletin database for all South African data that was required. The study uses quarterly data starting at the first quarter of 1961 and ending in the second quarter of 2003.

The calculation of the Nominal Effective Exchange Rate (*EXRATE*) is based on weights similar to those used by the South African Reserve Bank (see Walters and de Beer (1999)). The *EXRATE* is constructed as rands per foreign currency and covers 10 major trading partners which accounts for approximately 77% of South Africa's trade in manufactured goods. The 10 trading partners include the EURO Area (Germany, Italy, France, Netherlands, Belgium, Spain, Ireland, Austria, Finland and Portugal), USA, United Kingdom, Japan, Switzerland, Korea, Canada, Australia, Sweden and Singapore. Prior to the formation of the EURO Area in 1999, each EURO country was individually weighted in the *EXRATE*, while a total EURO Area weight was used post-1999, replacing the country weights. Three variables are constructed to measure effective foreign export prices faced by South Africa. *PFOR*, *XUV* and *MUV* use foreign producer price indices, export unit values and import unit values respectively. The weights used in construction of these variables are similar to those used in construction of *EXRATE* and may be found in more detail in the appendix.

Nominal Unit Labour Costs, *NULC*, Tariff Duties, *TARIFF*, South Africa's domestic Producer Prices (goods produced in South Africa for domestic use), *PDOM*, GDP

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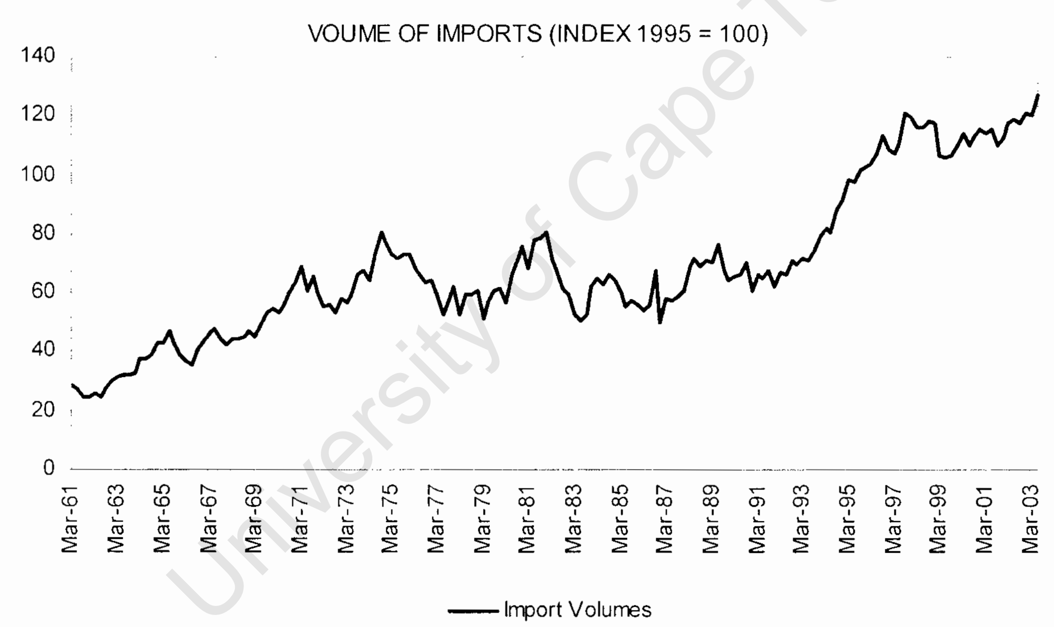
Expenditure, *Y*, Import Volumes, *M*, and Import Prices, *PIMP*, were constructed using data from the South African Reserve Bank Quarterly Bulletin. All index variables were constructed using a 1995 = 100 base and all real variables use constant 1995 prices. The world oil price (in Euros), *OILP*, was extracted from the IMF's International Key Macroeconomic Indicators (INTL) database. Further, two variables capturing excess domestic demand for South Africa, *LYG* and *LYLMG*, were created[4]. A full description of the variables may be found in the Appendix.

The model variables were analysed to assess trends, as well as identify factors that may influence the short-term fluctuations and long-term movements in import demand and prices.

IMPORT PERFORMANCE

Figures 1 and 2 below reflect import performance in South Africa for the period under study. The periods 1961-1970 and 1992-2003 exhibit a substantial increase in imports while during the period 1970-1992 imports show no clear trend with some large fluctuations.

Figure 1: Volume of Imports



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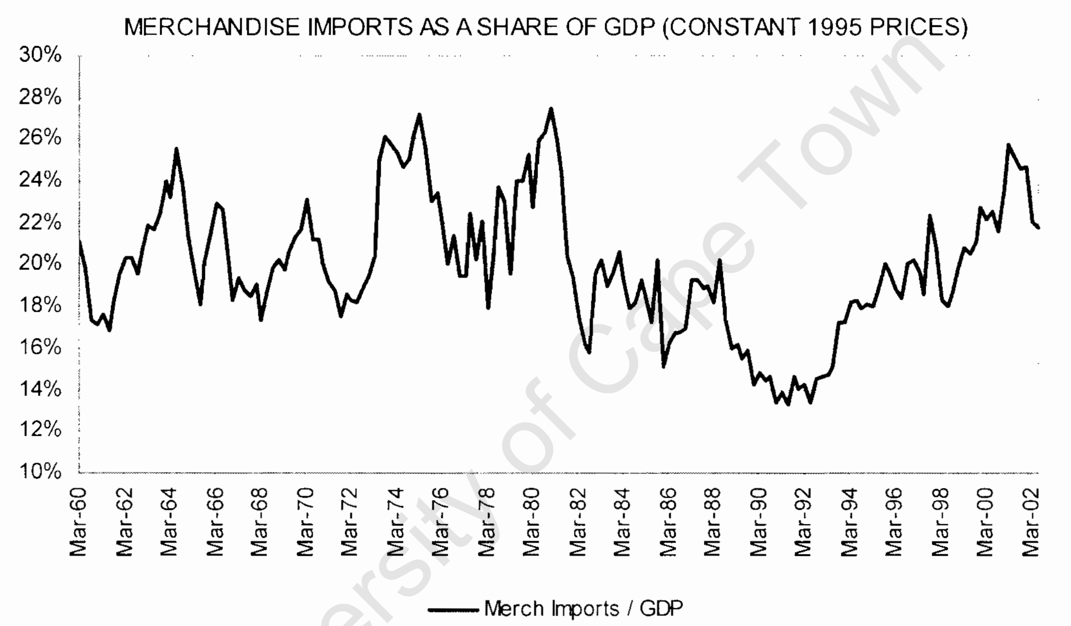
[4] The Hodrick-Prescott filter was used to create a smoothed GDP expenditure variable where $\lambda = 1600$. Excess demand is then given by the difference between GDP expenditure and this smoothed variable. *LYG* gives an indication of excess demand for all goods in South Africa, while *LYLMG* (GDP expenditure less imports) signifies excess demand for domestic goods in South Africa.

This reflects the inward-oriented policies followed by the government during the 1970s and 1980s and the associated protection afforded to domestic industries during this period. The imposing of sanctions as well as declining foreign investment in South Africa contributed to the decline in imports during the late 1980s. Merchandise imports as a share of GDP has risen substantially from 12% in 1992 to just below 25% in 2003, shown in Figure 2, indicating that there has been a substantial increase in import penetration in South Africa post apartheid.

Imports in South Africa have a clear positive trend following the 1992 period coinciding with the democratization of South Africa and the intensive liberalisation policies which followed immediately after. Factors that have affected import performance in South Africa are identified below and include the exchange rate, relative prices, factor costs and protection rates.

Figure 2: Merchandise Imports as a Share of GDP



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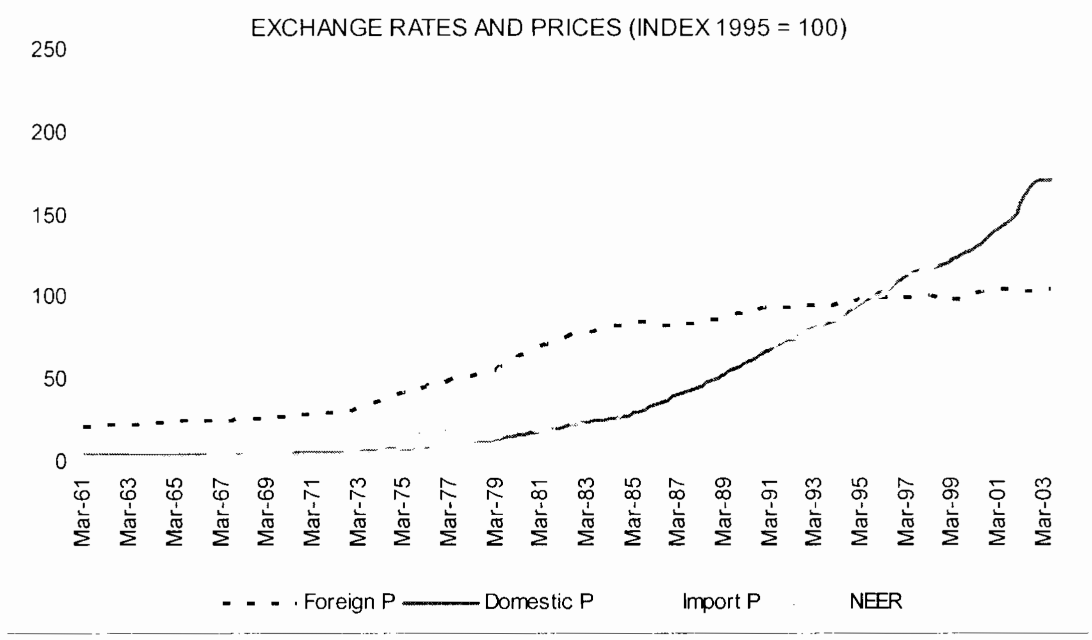
Exchange Rates and Prices

Figure 3 reflects that South African imports continued to grow despite a sharp rise (depreciation) in the exchange rate and import price between 1992 and 2003, highlighting several points.

Figure 3: Exchange Rates and Pass-through



Firstly, despite the drastic tariff reductions that have occurred during this period, import prices have nevertheless increased, implying that either an increase in foreign prices or a rise (depreciation) in the exchange rate is responsible for higher import prices. Secondly, the growth in imports despite higher prices may occur due to significant exchange rate pass-through to domestic prices, reducing any price advantage domestic goods may have over imported goods. Figure 3 reveals that there is indeed a close relation between domestic prices, the import price and the exchange rate, although there is an away movement in domestic price after 1998. Thirdly, the import price and exchange rate move closely throughout the period, suggesting that there is complete exchange rate pass-through to import prices and South Africa faces an infinite world supply of imports.

Foreign prices are fairly stable with a clear upward trend only during the 1970s and 1980s, possibly due to the two oil price shocks that occurred in this period. Oil prices affected not only foreign prices, but also the relative price of imported goods in South Africa. This is highlighted in Figure 4, which shows the relative price of imported and foreign goods measured in rand currency i.e. the real effective exchange rate (REER) of imports in rands.

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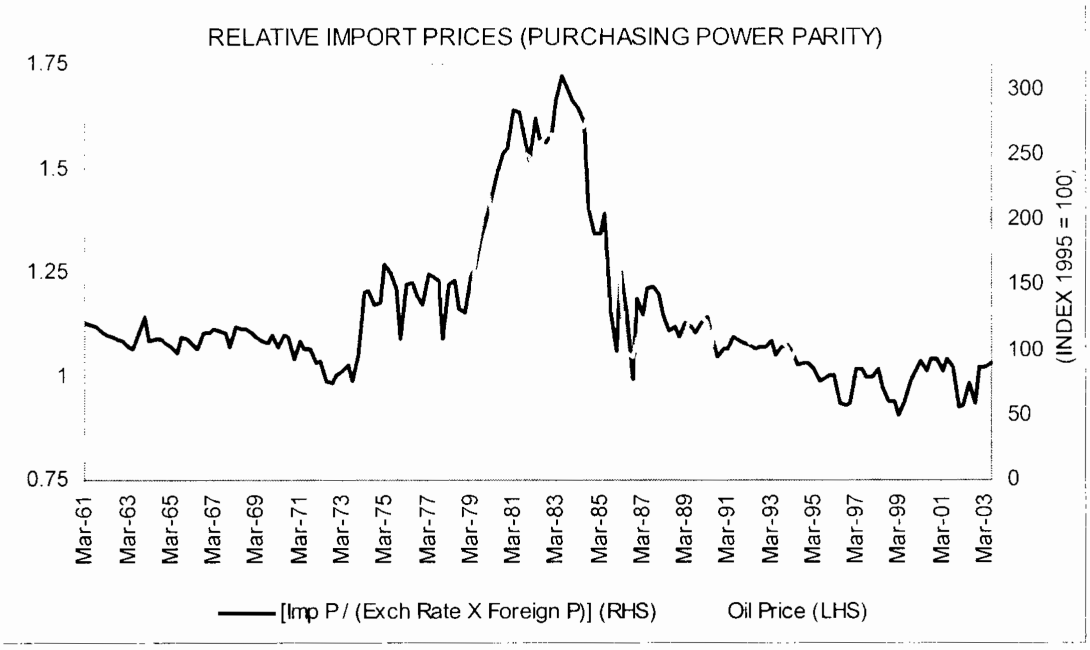
Figure 4: REER of Imports



A stable REER is an indication that PPP holds and implies a perfect pass-through of exchange rates to the import price. Figure 4 shows two periods in the 1970s and 1980s that mar the REER's relative stability. A plot of the nominal world oil price against this indicates that the oil shocks of 1973-4 and 1979-84 had a significant effect on the REER for imports. This is a combined result of two effects: 1) the relative increase in import over foreign prices and 2) the appreciation (decrease) of the nominal exchange rate. The differing bundle of goods in the import (*PIMP*) and foreign price (*PFOR*) indices results in the first effect, as South Africa would have a higher proportion of oil imports weighted in import prices than is likely in foreign producer prices, which would have a wider ranger of goods included in the index. The second effect occurred due to the rise in gold prices, a result of the higher oil prices[5] and the move to fixed exchange rates. The dollar gold price rose by 62% between the third quarter of 1973 and the fourth quarter of 1974, rising again by 102% between the third quarter of 1979 and the fourth quarter of 1981. Considering that gold averaged 31% and 38%

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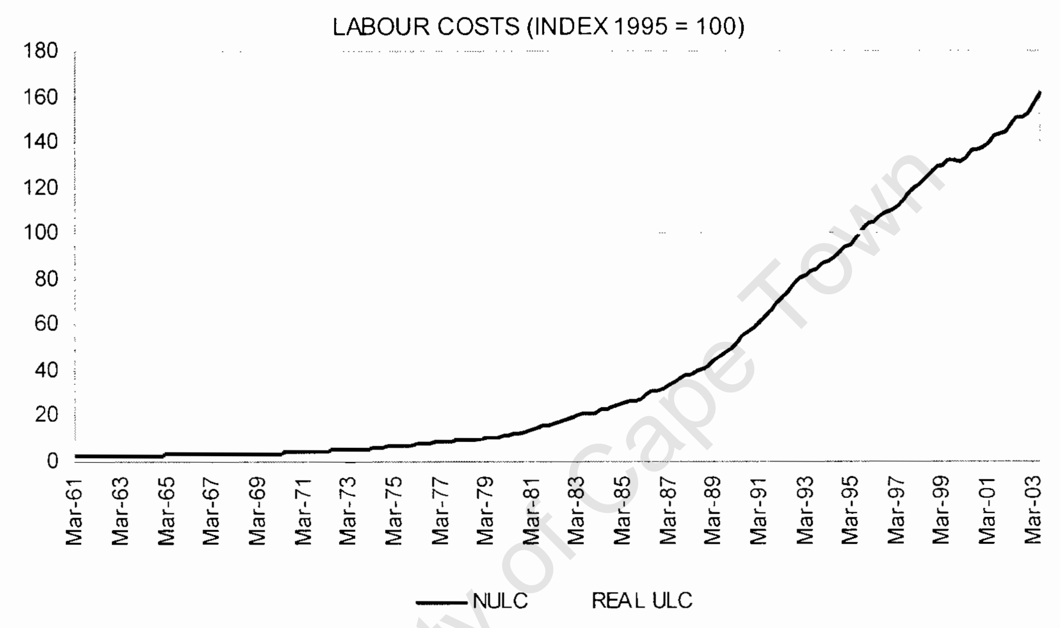
[5] This implies the assumption of the comovement of commodity prices, an assumption that has been questioned. Pindyck and Rotemberg (1990) is an early study testing commodity comovement while Palaskas and Varangis (1991) and Cashin. McDermott and Scott (1999) dispute this assumption. While the fluctuations in the gold price during the 1970s may be largely due to the end of the gold standard, oil prices shocks can have served to exacerbate this effect as countries moved to stock up on gold reserves to protect their currencies against the oil shocks.

of the total South African exports (including services) in the 1970s and 1980s, the rise in the gold price facilitated an appreciation in the nominal exchange rate in South Africa[6].

Factor Costs

Figure 5 illustrates the progressive upward trend of Nominal Unit Labour Costs (NULC) in South Africa starting in the late 1970s, a result of the feedback mechanisms between labour costs, inflationary expectations and domestic prices. Real Unit Labour Costs (RULC) by contrast have remained fairly consistent throughout the study period, indicating that once inflationary effects have been removed, the real cost of labour in South Africa has remained relatively unchanged between 1961 and 2003.

Figure 5: Labour Costs



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Tariff Protection

The liberalisation process has contributed significantly to the rise in import demand. Average tariff duties have declined by 9% between 1992 and 2003.

From Figure 6, the periods prior to 1970 and post 1992 show a negative correlation between tariff rates and import demand. This can be attributed to the direct influence tariff duties have on import prices, raising the price of imports relative to domestic goods.

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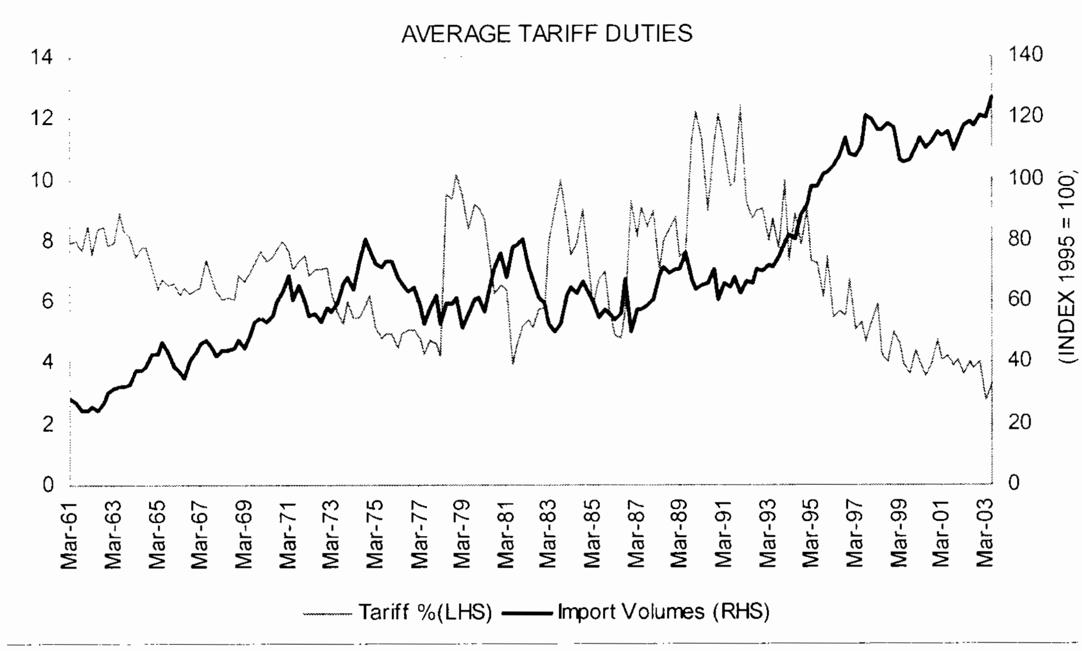
[6] Own calculations based on data from the South African Reserve Bank (SARB) and International Monetary Fund (IMF).

Figure 6: Tariff Duties



The analysis of the data indicates the strong possibility both of South Africa being an import price-taker and the likelihood that there is significant exchange rate pass-through to domestic prices which needs to be accounted for within an import demand framework. The next section deals with the issue of stationarity and ensuring all variables are $I(1)$ at most.

UNIT ROOT TESTS

Important to the Johansen technique is ensuring that all endogenous variables in the cointegrating vectors are integrated order $I(1)$ at most. Alternatively stated, first differenced endogenous variables should exhibit no unit root (should be mean reverting series) with or without a trend. Table 1 indicates that the all variables are $I(1)$ at most, with the exception of *LNULC*, nominal unit labour costs, and possibly foreign prices, *LPFOR*. A well-known weakness of the ADF test is its inability to deal with structural breaks, which are treated as evidence of nonstationarity in the series (Baum, 2001). Following this, Clemente, Montanes and Reyes (1998) propose a test that is capable of dealing with two structural breaks[7]. The structural break test can be carried out using either additive or innovative outliers. The former captures a sudden change in the series while the latter accounts for a gradual shift in the mean

[7] These tests are an extension of the single structural break tests developed by Perron (1989) and Perron and Vogelsang (1992) and are performed using Stata/SE 9.1. The AR lag order is determined by a set of sequential F-tests and the breakpoints are determined on the basis of the maximal t-statistic for the unit root hypothesis (Baum, 2001).

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of the series following the shock (Baum, 2001). The structural break test results for the above-mentioned variables are provided in Table 2.

In Table 2 breakpoints 1 and 2 are significant for both variables. Breakpoint 1 coincides with the end of the gold standard and the start of the first oil price shock in the 1970s, and may be expected given the inflationary period that directly followed the move to flexible exchange rates. For the foreign price variable, the second breakpoint corresponds to the beginning of the second oil shock in the 1970s, while the second break for nominal unit labour costs occurs at a point where a continuous period of political upheaval was coming to an end in South Africa. A graphical analysis of these variables is provided in the appendix.

A series can be considered stationary if the coefficient on $\rho - 1$ is significant and, as seen in Table 2, the null hypothesis of a unit root in both variables is rejected. Following this, all potential endogenous variables are seen as integrated $I(1)$ at most and the paper can confidently make use of the Johansen cointegration technique. The following section provides the results of firstly, the PPP test and secondly, the pass-through test for the small country assumption.

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Table 1: Augmented Dickey Fuller (ADF) Unit Root Tests, Quarterly Data, 1962Q2 – 2003Q2

	LEVELS			FIRST DIFFERENCE		
VARIABLE	TREND	LAG LENGTH	ADF TEST	TREND	LAG LENGTH	ADF TEST
LM	Upward (1961-1971, 1993-2003)	0 (SBC, HQC) 4 (AIC)	-3.201 -3.311	Level	0 (HQC, SBC, AIC)	-14.768*
LPFOR	Upward	2 (SBC, HQC) 4 (AIC)	-0.294 -0.441	Structural Breaks (1973, 1984)	0 (SBC, HQC) 3 (AIC)	-3.414* -2.813
LPDOM	Upward	3 (SBC, HQC, AIC)	-1.985	Structural Breaks (1972, 1986)	2 (HQC, SBC, AIC)	-3.025*
LPIMP	Upward	0 (SBC) 4 (AIC, HQC)	-2.334 -2.242	Level	3 (HQC, SBC, AIC)	-3.868*
LEXRATE	Upward	3 (HQC, SBC, AIC)	-2.334	Level	2 (HQC, SBC, AIC)	-5.269*
LNULC	Upward	5 (SBC) 6 (HQC, AIC)	-1.813 -1.918	Structural Breaks (1972, 1992)	4 (SBC) 5(HQC, AIC)	-2.862 -2.523
LY	Upward	0 (HQC, SBC, AIC)	-3.0601	Level	0 (SBC, HQC) 2 (AIC)	-12.814* -6.018*
LTARIFF	Downward (1961-1978, 1993-2003) Upward (1979-1992)	4 (HQC, ABC, AIC)	-2.481	Level	3 (HQC, SBC, AIC)	-6.243*
LOILP	Level, Structural Breaks (1973, 1979, 1985)	0 (HQC, SBC, AIC)	-1.424	Level	0 (HQC, SBC, AIC)	-11.504*
LYLMG	Level	0 (SBC) 4 (HQC, AIC)	-5.203* -5.763*			
LYG	Level	0 (SBC) 4(HQC, AIC)	-4.585* -5.923*			
LXUV	Upward	1 (SBC) 2 (AIC, HQC)	-0.4523 -0.7150	Level	1 (HQC, SBC, AIC)	-4.1806*
LMUV	Upward	1 (SBC) 4 (HQC) 5 (AIC)	-1.0378 0.9395 0.6775	Level	0 (SBC) 3 (HQC) 4 (AIC)	-5.4817* -4.4406* -4.7955*

Notes: An asterisk denotes that the null hypothesis of a unit root in the variable (non-stationarity) is rejected. The choice of leg length is based on the Aikaike Information (AIC), Schwarz Bayesian (SBC) and Hannan-Quinn (HQC) criteria. The 95% critical values are -2.88 for without a trend and -3.44 including a trend.

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Table 2: Clemente-Montanes-Reyes Double Structural Break Test, Quarterly Data, 1961Q2 – 2003Q2

VARIABLE	DLNULC	DLPFOR
Breakpoint 1	1972 Quarter 2 0.0218* (8.454)	1972 Quarter 1 0.0046* (3.487)
Breakpoint 2	1992 Quarter 2 -0.0165* (-6.343)	1979 Quarter 4 -0.0058* (-4.621)
ρ – 1	-0.9342* (-11.984)	-0.2843* (-5.595)

Notes: The variables *DLNULC* and *DLPFOR* were tested for stationarity using the additive outliers (AO) and innovative outliers (IO) models respectively, based on a graphical inspection for the possibility of a shifting trend. A maximal lag length of 6 was used with AR process determined by sequential F-tests. An AR(0) and AR(2) process was determined for *DLNULC* and *DLPFOR* respectively.
The t-statistics are compared with critical values provided by Perron and Vogelsang (1992). The 5% critical t-statistic for the coefficient on $\rho - 1$ is -5.490.

6. THE SMALL COUNTRY ASSUMPTION

The results for the two techniques testing the small country assumption are described below. The first section looks at the PPP test results while the second section deals with the results from the exchange rate pass-through test.

THE PPP TEST

The variable *reer* in Equation *(19)* is constructed using *LPIMP* (p^*), *LEXRATE* (e) and *LPFOR* (p^f). Table 3 shows the standard Augmented Dickey-Fuller (ADF) test for *reer.*

Table 3: Augmented Dickey-Fuller Test for a Unit Root on variable reer, Quarterly Data, 1962Q4 – 2003Q2

VARIABLE	TREND	LAG LENGTH	ADF TEST	95% CRITICAL VALUE
reer	Level with Structural Breaks (1973, 1979, 1984)	4 (SBC, HQC, AIC)	-1.9448	-2.8790

The null hypothesis of a unit root in *reer* is not rejected, implying that PPP does not hold for import price. But, as earlier mentioned, the ADF test does not account for the possibility of structural breaks. Figure 4 clearly demonstrates two, possibly three, structural breaks in the REER that seem closely correlated with the oil price shocks of 1973/4 and 1979/84. Again the Clemente-Montanes-Reyes (1998) test is used to account for these structural breaks.

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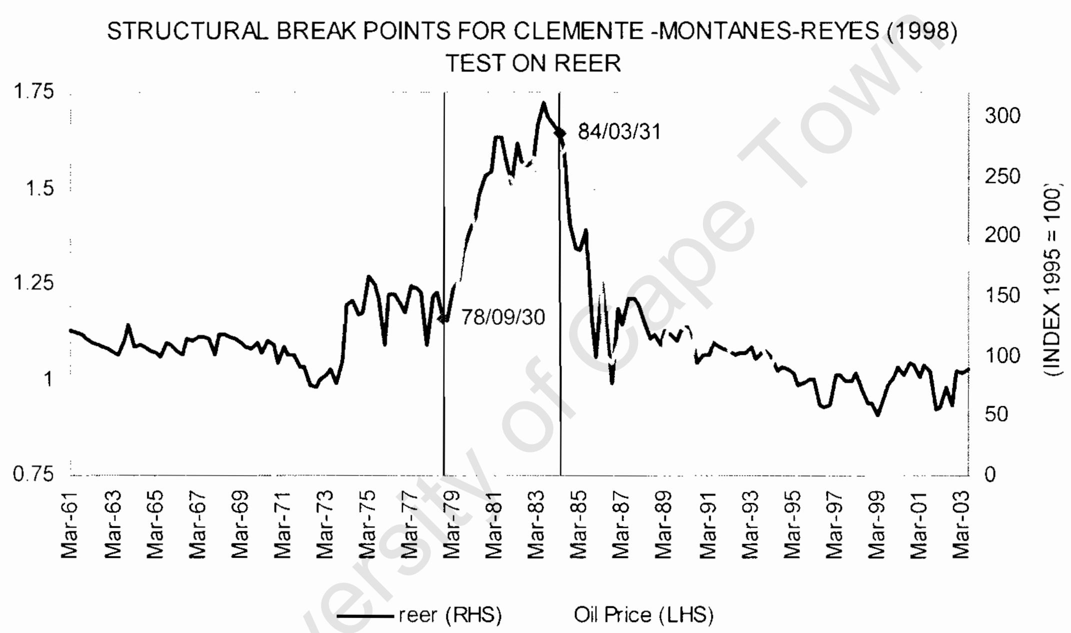
Table 4: Clemente-Montanes-Reyes Double Structural Break Unit Root Test on variable reer, IO Model, AR(2) process, Quarterly Data, 1961Q1 – 2003Q2

	BREAKPOINT 1 1978 QUARTER 3	BREAKPOINT 2 1984 QUARTER 1	P - 1
Coefficient	0.0876*	-0.1055*	-0.2108*
	(5.976)	(-6.761)	(-6.001)

Notes: A maximal lag length of 6 was used for the AR process. Values in parentheses are t-statistics, an asterisk indicates a 95% level of significance, critical values for the non-standard t-statistics are taken from Perron and Vogelsang (1992). The 95% critical value for the coefficient on ρ-1 is -5.490.

Table 4 indicates the results for the Clemente-Montanes-Reyes (1998) test. Both breakpoints are significant at the 95% level and seem to coincide with the starting and end points of the second oil shock in the late 1970s and early 1980s, as shown in Figure 7. This implies that the series does indeed have structural breaks, and is a further indication that there was significant feed through of oil-price shocks to import prices.

Figure 7: Clemente-Montanes-Reyes (1998) Structural Break Points on reer



Further, Table 4 indicates that the null hypothesis of a unit root in the real effective exchange rate is rejected, implying that purchasing power parity holds for import prices. This supports the assumption of infinite world supply for South African imports in the long-run. Over the next section, a more formal process of testing for full exchange pass-through is used to reinforce the initial results.

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EXCHANGE RATE PASS-THROUGH

The small country assumption has been tested for South Africa primarily using the exchange-rate pass through to import price model. Nell (2000) finds, for the short sample period of 1987 quarter 1 to 1997 quarter 1, that exchange rate pass-through to import prices for South Africa is incomplete, with exchange rate coefficients of 0.83 and 0.72 on two different import pricing models. Similarly, Rangasamy (2003) indicates that exchange rate pass-through to import prices is less than one for South Africa, with a coefficient of 0.67. Menon (1995a) gives a comprehensive survey of international exchange rate pass-through studies. Menon (1995a) finds that very few show complete pass-through, even those for "small" countries. These results imply that the small country assumption does not hold for South Africa and suggest that there is some market power in the purchase of imports. However the majority of these studies exclude the impact of oil prices on the price of imports in their models and hence may suffer from omitted variable bias. Figure 4 and the PPP test provide evidence that the real effective exchange rate for imports (and the import price) may have been substantially affected by oil prices fluctuations during the 1970s and early 1980s.

Oil prices may therefore act as a 'long-run forcing' or weakly-exogenous variable[8] on import prices. Equation *(17)* is therefore modified to include oil prices as such. The effects of excess demand on import prices is accounted for by including the $I(0)$ variable LYG[9]. Thus Equation *(17)* is estimated as follows:

$$LPIMP = \theta_1 LPFOR + \theta_2 LEXRATE + \theta_3 LTARIFF + \theta_4 LOILP + \varepsilon_t \qquad (26)$$

The Aikaike Information Criterion (AIC) indicated a VAR(5) while the Schwarz Bayesian Criterion (SBC) indicated a VAR of order 2. A VAR(4) was selected based on the adjusted LR test and in an attempt to minimise serial correlation. Despite this serial correlation remained in the foreign price and tariff equations.

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[8] A variable, x_t, can be considered weakly exogenous if the vector of adjustment coefficients, α, equal zero in the ECM equation for x_t, (Johansen, 1992). The ECM equation for oil prices, when included as an endogenous variable, revealed an insignificant adjustment coefficient of 0.01 with a p-value of 0.881 and is sufficient to conclude that the variable is weakly exogenous.

[9] The inclusion of dummies capturing exchange rate shocks, the fixed exchange rate period and the sanction period led to problems arising in the short-run dynamics, where the import price adjustment coefficient was found to be either insignificant or positive in value in the import price ECM. Thus the results provided exclude all dummy variables.

The maximal eigenvalue and trace test statistics used to determine the number of cointegrating vectors is shown in Table 5. The results agreed with expectations and one cointegrating relation determining import prices was selected.

Table 5: Cointegration with Unrestricted Intercepts and No Trends in the VAR; 166 observations 1962Q1 – 2003Q2; VAR Order = 4; list of variables included in the cointegrating vector: LPIMP, LPFOR, LEXRATE, LTARIFF, LOILP; list of I(1) exogenous variables included in the VAR: LOILP; list of I(0) variables included in the VAR: LYLG; list of eigenvalues in descending order: 0.17356, 0.10018, 0.039347, 0.021277, 0.0000

NULL	ALTERNATIVE	MAXIMAL EIGENVALUE STATISTIC	95% CRITICAL VALUE	TRACE STATISTIC	95% CRITICAL VALUE
$r = 0$	$r \geq 1$	31.6446*	30.7100	59.4016*	58.6300
$r \leq 1$	$r \geq 2$	17.5232	24.5900	27.7570	38.9300
$r \leq 2$	$r \geq 3$	6.6637	18.0600	10.2338	23.3200
$r \leq 3$	$r = 4$	3.5701	11.4700	3.5701	11.4700

The just-identified restriction on the import price reveals an exchange rate pass-through of 0.89 to import prices as shown in column *(1)* of Table 6. This is higher than both those found by Nell (2000) and Rangasamy (2003). In addition, the weakly exogenous oil price is significant at the 95% level, with the coefficient intimating that a 1% rise in the nominal (foreign currency) oil price will force the import price to rise by 0.24% in the long-run.

The insignificance (and incorrect sign) of tariff rates is surprising, with the coefficient sign implying that higher tariff rates reduces the price of imports. It may be that this measure of tariffs does not adequately capture effective rates for South Africa, especially during the 1970s and 1980s where the South African trade regime was characterized by sanctions, quantitative restrictions (such as quotas) and a myriad of trade barriers (surcharges) in an effort to "maintain current account surpluses in excess of required foreign debt repayments" (Aron, Muellbauer, Smit, 2004). Rustomjee (1991) also comments on the possibility of underinvoicing, especially during the sanction period, in an effort to reduce tariff charges. Additionally, the tariff variable used here is a measure of collection and not scheduled rates. This means that the effective protection measured by the tariff variable is underestimated since products with high scheduled tariffs are less likely to be imported.

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Table 6: Cointegration Results for an Import Price Function with Oil Weakly Exogenous, *LYG* included as *I*(0) variable, Unrestricted Intercepts, No Trend, Quarterly Data, 1962Q1 – 2003Q2, VAR(4)

VARIABLE	JUST-IDENTIFIED COEFFICIENTS	OVER-IDENTIFIED RESTRICTIONS		
	(1)	*(2)* $\theta_1 = \theta_2$	*(3)* $\theta_2 = 1$	*(4)* $\theta_1 = \theta_2 = 1$
LPFOR *(p^f)*	0.9076* (0.2187)	0.8960* (0.0291)	0.7256* (0.1373)	*1*
LEXRATE *(e)*	0.8914* (0.0971)	0.8960* (0.0291)	*1*	*1*
LTARIFF *(tariff)*	-0.2393 (1.4215)	*0*	*0*	*0*
LOILPRICE	0.2440* (0.0766)	0.2452* (0.0302)	0.3170* (0.0687)	0.2449* (0.0743)
ECM(-1)	-0.1322* (0.0300)	-0.1346* (0.0308)	-0.0797* (0.0239)	-0.0544* (0.0134)
LR test of restrictions (χ^2) [p-value]		(0.0385) [0.981]	(0.8590) [0.651]	(2.8469) [.416]

Notes: Values in parentheses reflect coefficient standard errors.
Prior to imposing an additional restriction on tariff rates ($\theta_3 = 0$), the restrictions of unity on the foreign price and exchange rate coefficients ($\theta_1 = \theta_2 = 1$) could not be rejected with a $\chi^2 = 2.5054$ and a p-value of 0.286.

Column *(3)* in Table 6 presents the test for complete exchange rate pass-through to import prices. The over-identifying restriction on *LEXRATE* could not be rejected at the 95% level of significance and is sufficient to conclude that the small country assumption holds for South Africa. Further, the over-identifying restrictions in column *(4)* suggest that foreign and import prices are homogenous of degree one.

While the results above imply that exchange rate pass-through is complete in the long-run, the dynamics of the system are less clear. Specifically, the speed of adjustment to equilibrium (reflected in the ECM(-1) coefficients) declines when full pass-through and price homogeneity are imposed on the system. The error correction coefficient declines from -0.132 in column *(1)* to -0.05 in column *(4)*. Although the interest of this paper is in the long-run behaviour, the declining ECM(-1) coefficient presents an opportunity to clarify the exchange rate pass-through to import prices.

The extent of exchange rate pass-through is therefore tested again using a different foreign price (p^f) variable. There is a direct trade-off between using the foreign producer price index (*LPFOR*) and either one of export and import unit values (*XUV* and *MUV*). *LPFOR* captures

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a larger proportion of South Africa's trading partners but is deficient in that it captures a range of "non-traded" goods which can influence the resulting estimation. As unit values, *XUV* and *MUV* capture not only the traded bundle of goods but also the changing composition of this "bundle" over time. However, additional measurement problems may arise from combining unit value and price indices in a single estimation[10].

Table 7: Cointegration Results for an Import Price Function using Export and Import Unit Values as the Foreign Export Price (p^f), LOILPRICE included as weakly exogenous, *LYG* included as *I*(0) variable, Quarterly Data, 1964Q1 – 2002Q4

	EXPORT UNIT VALUES (*LXUV*) (VAR (4), UNRESTRICTED INTERCEPTS, NO TREND)		IMPORT UNIT VALUES (*LMUV*) (VAR (5), UNRESTRICTED INTERCEPTS, NO TREND)	
	(1)	*(2)*	*(3)*	*(4)*
DEPENDENT VARIABLE	**JUST-IDENTIFIED COEFFICIENTS**	**SMALL COUNTRY TEST**	**JUST-IDENTIFIED COEFFICIENTS**	**SMALL COUNTRY TEST**
(p^f)	0.9900* (0.1506)	0.5439* (0.1129)	1.0707* (0.2849)	0.5596* (0.1738)
LEXRATE *(e)*	0.8657* (0.0454)	*1*	0.8618* (0.0836)	*1*
LTARIFF *(tariff)*	-2.4610 (1.3223)	-2.1280 (2.0575)	-4.3011 (2.9813)	-2.2288 (2.4507)
LOILPRICE	0.2597* (0.0552)	0.4529* (0.0944)	0.1998 (0.1154)	0.4250* (0.1599)
ECM(-1)	-0.1718* (0.0258)	-0.1095* (0.0178)	-0.0894* (0.0209)	-0.0785* (0.0196)
LR test of restrictions (χ^2) [p-value]		7.6681 [0.006]		(4.7765) [0.029]

Notes: Values in round brackets reflect coefficient standard errors. An asterisk indicates significance at the 95% level.
The just-identified coefficient on *LTARIFF* could not be restricted to zero with a χ^2 = 6.4547 and a p-value = 0.011 in the *LXUV* vector and a χ^2 = 5.9650 and a p-value = 0.015 in the *LMUV* vector.

Using the unit value measures[11], columns *(1)* and *(3)* in Table 7 reveal exchange rate pass-through that is slightly lower than earlier found but higher than both Nell (2000) and Rangasamy (2003). However, the rejection of complete exchange rate pass-through in columns *(2)* and *(4)* suggest that the small country assumption may not hold for South Africa. The contradictory results of this test reveal that there is further scope to study the small country assumption in South Africa. Nevertheless, given the PPP test results and the

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[10] For more on the use of unit values see Shiells (1991).
[11] One cointegrating vector was selected on the basis of both the maximal eigenvalue and trace statistics. These results are presented in the appendix.

extensive use of this assumption in the literature the paper continues with the small country assumption. Import prices are therefore seen as an exogenous variable in the study of import demand.

7. IMPORT DEMAND

The few published studies on import demand in South Africa have provided conflicting results on the magnitude of price and income elasticities of import demand and the exact nature of the relationship of these variables. A summary of the various studies is provided in Table 8. There is a considerable variation in the magnitude of both the price and income elasticities, with price elasticities ranging between -0.05 and -1.59 and income elasticities ranging between 0.43 and 1.92. For all seven of the studies, price homogeneity in import demand is imposed rather than tested. In a majority of the studies, the time periods are short or there are a low number of observations used in estimation. Only four of the seven studies mentioned make use of a cointegration technique in order to account for non-stationarity. In addition, not one of these studies either examines import demand with endogenous import prices or formally tests the small country assumption that allows for import price exogeneity.

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Table 8: Import Demand Studies for South Africa

STUDY	PRICE ELASTICITY	INCOME ELASTICITY	PERIOD	COMMENT
Bahmani-Oskooee and Niroomand (1998)	-0.53	0.43	1960 – 1992 Annual Data	Johansen cointegration technique for 30 countries
Edwards and Wilcox (2003)	-1.59	1.92	1972Q1 – 2001Q4, Quarterly Data	Johansen cointegration technique
Golub (2000)	Range between -0.05 and -0.32	Range between 0.93 and 1.04		Ordinary Least Squares, using varying effective exchange rate measures
Gumede (1999)	-0.71 for capital intensive goods -3.00 for labour intensive goods			Engle-Granger cointegration approach
Narayan and Narayan (2003)	-0.61	1.19	1960 – 1996 Quarterly Data	Bounds test cointegration approach
Senhadji (1997)	-1.00 in the long-run -0.44 in the short run	0.68	34 observations	Fully modified (FM) estimators using Monte Carlo method for 77 countries
Smal (1996)	-0.85	1.47	1985Q1 – 1994Q4, Quarterly Data	Ordinary Least Squares using non-oil imports

No study in South Africa has yet dealt with the possible effects of domestic price endogeneity in import demand. International literature on this is similarly poor: of the literature surveyed only one study was found to deal with the endogeneity of domestic prices in either import demand or export supply[12]. Given this, it is suspected that there is an upward bias with regard to the exchange rate effect on the trade balance inherent in the results of these studies.

The following sections deal with the deficiencies in the literature on import demand in South Africa. A forty year period of quarterly data is used to estimate an import demand function with endogenous domestic prices.

PRELIMINARY TESTS

Firstly, Equation *(21)* is adjusted to include *LYLMG* (representing $ydgap_t$) as an $I(0)$ variable. First-differenced oil prices (*DLOILP*) was also included as an $I(0)$ variable to account for the effect that commodity shocks may have on both imports and domestic prices, specifically in the late 1970s and early 1980s[13]. Dummy variables were included to account for exchange rate shocks, the fixed exchange rate period, the sanction period and the period of rapid liberalisation following 1994. These dummies were found to be effective in controlling for some of the heteroscedasticity and serial correlation experienced in both equations.[14] The simultaneous framework given by Equations *(11)* and *(21)* is therefore estimated as follows:

$$LM = \alpha_1 LY + \alpha_2 LPDOM + \alpha_3 LPIMP + \alpha_4 LTARIFF + \mu_t \qquad (27)$$

$$LPDOM = \gamma_1 LPFOR + \gamma_2 LEXRATE + \gamma_3 LNULC + \gamma_5 LTARIFF + \nu_t \qquad (28)$$

Secondly, the AIC, SBC and adjusted LR tests were used to select the lag length of the VAR. The AIC indicated a VAR of up to 6 lags while the SBC indicated a maximum of 2 lags. A VAR order of 4 was chosen for the model based on both the adjusted LR test and an analysis of the residuals of individual variable equations, though serial correlation persisted in

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[12] See Newman, Lavy and de Vreyer (1995).

[13] This follows from Johansen and Juselius (1992) and Pesaran and Pesaran (1997), who condition the data using first differenced oil prices as an $I(0)$ variable in their analysis of PPP.

[14] Following Pesaran and Pesaran (1997) the VAR equations are analysed and the "LR test of deletion" is used to identify the value and significance of a dummy variable in the VAR model and resulting ECM. The null hypothesis of deletion from the model was rejected by all dummy variables, except for the *POST94* and *FIXED* dummies. These two dummy variables were still included as they accounted for some of the convergence and serial correlation issues present in the model.

domestic expenditure. Serial correlation was induced in other variables at both a higher and lower order VAR.

Thirdly, the maximal eigenvalue and trace statistics were used to identify the number of cointegrating vectors. As shown in Table 9, both the maximal eigenvalue and trace test statistics indicated a maximum of two cointegrating relations. Following model expectations and the results of the two tests, two cointegrating relations were chosen.

Table 9: Cointegration with Unrestricted Intercepts and No Trends in the VAR; 166 observations 1962Q1 – 2003Q2; VAR Order = 4; list of variables included in the cointegrating vector: LM, LPFOR, LPDOM, LPIMP, LEXRATE, LNULC, LY, LTARIFF; list of I(0) variables included in the VAR: DLOILP, EXRSHOCK, SANCTION, POST94, FIXED, LYLMG; list of eigenvalues in descending order: 0.52274, 0.30201, 0.17141, 0.11578, 0.098926, 0.060528, 0.012912, 0.0092640

NULL	ALTERNATIVE	MAXIMAL EIGENVALUE STATISTIC	95% CRITICAL VALUE	TRACE STATISTIC	95% CRITICAL VALUE
$r = 0$	$r \geq 1$	122.788*	51.150	265.473*	157.800
$r \leq 1$	$r \geq 2$	59.687*	45.630	142.685*	124.620
$r \leq 2$	$r \geq 3$	31.212	39.830	82.998	95.870
$r \leq 3$	$r \geq 4$	20.427	33.640	51.786	70.490
$r \leq 4$	$r \geq 5$	17.292	27.420	31.359	48.880
$r \leq 5$	$r \geq 6$	10.365	21.120	14.067	31.540
$r \leq 6$	$r \geq 7$	2.157	14.880	3.702	17.860
$r \leq 7$	$r = 8$	1.545	8.070	1.545	8.070

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COINTEGRATION RESULTS

Table 10 show the cointegration results for import demand with endogenous domestic prices. The just-identified restrictions in columns *(1)* and *(2)* reveal that, with the possible exception of *LY* in the domestic price equation (*CV 2*), the signs of all coefficients are as would be expected. The LR test reveals that the over-identified restrictions could not be rejected and the model is deemed successfully identified in columns *(3)* and *(4)*.

Column *(3)* in Table 10 indicates the import demand function within the simultaneous framework. This reveals that import demand is more responsive to domestic rather than import prices in the long-run. The additional over-identifying restriction to test for equality of coefficients[15] ($\alpha_2 = -\alpha_3$) was rejected with a p-value of 0.034 indicating that import and domestic prices may not necessarily be homogenous as is usually assumed. A 1% rise in domestic prices would therefore increase the demand of imports by 0.78% while a 1% rise in import prices would reduce the demand for imports by 0.71%. A 1% rise in income increases the demand for imports by 0.84% and thus implies that aggregate imports are not necessarily luxury goods (Islam and Hassan, 2004). The tariff effect on import demand is relatively large, with 1% rise in tariffs reducing the demand for imports by 2%.

Table 10: Cointegration Results for Import Demand and Domestic Price Multivariate Framework, Unrestricted Intercepts, No Trend, Quarterly Data, 1962Q1 – 2003Q2

	Just-identified Restrictions		**Over-identified Restrictions**	
	(1)	*(2)*	*(3)*	*(4)*
DEPENDENT VARIABLE	**LM (CV 1)**	**LPDOM (CV 2)**	**LM (CV 1)**	**LPDOM (CV 2)**
LPFOR *(p^f)*	-0.2323 (0.1682)	0.9307* (0.1658)	0	0.9076* (0.0635)
LPDOM *(p_D)*	0.7295 (0.4016)		0.7833* (0.1334)	
LPIMP *(p^*)*	-0.5757* (0.1654)	-0.0286 (0.0751)	-0.7062* (0.1038)	0
LEXRATE *(e)*	0	0.6210* (0.1052)	0	0.7003* (0.0599)
LNULC *(nulc)*	-0.0366 (0.2439)	0.2229* (0.0688)	0	0.1255* (0.0579)
LY *(y)*	1.0348* (0.0956)	-0.1255* (0.0367)	0.8378* (0.0764)	0
LTARIFF *(tariff)*	-1.5169* (0.8691)	0.9611* (0.3179)	-2.0820* (0.9164)	1.2547* (0.4649)
LR test of restrictions {χ^2} [p-value]			{7.3073} [0.121]	

Notes: Values in parentheses indicate coefficient standard errors. An asterisk indicates a 95% level of significance.

Column *(4)* shows the estimation results for domestic price within the simultaneous framework. The results show that there are significant inflationary effects of both foreign prices and the exchange rate on domestic price levels. A 1% rise in the foreign price would

[15] Price homogeneity implies that $\alpha_2 = -\alpha_3$, substituting coefficients we have:

$$\alpha_3 LnP^* - \alpha_3 LnP = \alpha_3 LnP^* + \alpha_3 Ln\frac{1}{P} = \alpha_3 Ln\frac{P^*}{P}$$

Thus with price homogeneity enforced on import demand, the coefficients on both import and domestic price represent the elasticity of relative prices. Enforcing this condition on import demand resulted in a statistically insignificant relative price elasticity of -0.57.

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lead to a 0.9% increase in domestic price levels. Similarly, a 1% rise (depreciation) in the exchange rate would result in a 0.7% rise in the domestic price. The exchange rate pass-through to domestic prices is higher than those found in other studies on South Africa. Choudri and Hakura (2001) test the hypothesis that a low inflationary environment leads to low exchange rate pass-through[16] and find a pass-through coefficient of 0.02 in quarter 1 for South Africa, rising to 0.14 after 20 quarters. Devereaux and Yetman (2003) find a pass-through of 0.33 to domestic prices, concluding that short-term price rigidities lower the effective exchange rate pass-through to domestic prices. Fedderke and Schaling, also using the multivariate Johansen cointegration technique, find a long-run pass-through of only 0.23.

Wage increases represented by higher unit labour costs positively affect domestic prices, with a 1% rise in unit labour costs increasing domestic prices by 0.13%. This is smaller than the effect found by Fedderke and Schaling (2005) where a coefficient of 1.31 is estimated. The tariff effect on domestic price is positive and large, with a 1% rise in tariff rates increasing domestic prices by 1.3%. This result is in line with Rangasamy's (2003) notion that a reduction in tariff rates ultimately improves domestic price competitiveness by reducing intermediate costs (on the supply side) and competing import prices (on the demand side).

The negative adjustment coefficients in the ECM models (shown in Table 11) reveal that both import demand and domestic price converge toward equilibrium. This is presented by Figure 8 which indicates the effect of a system-wide shock on the long-run vectors. Domestic prices are seen to converge quicker than import demand, taking approximately 10-12 quarters to converge to equilibrium. Though relatively long, this is somewhat consistent with Menon's (1995a) review of lengthy lags in exchange rate pass-through. Import demand takes slightly longer, converging after only 16-18 quarters.

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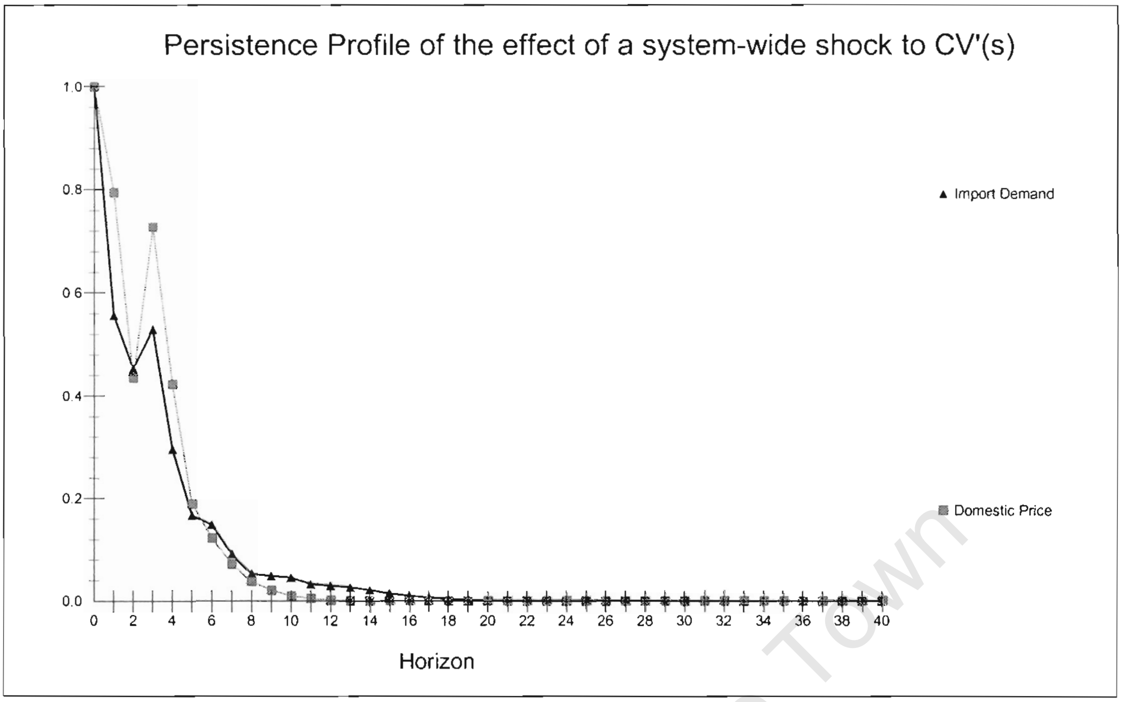
[16]The study by Choudri and Hakura (2001) find, using a database of 30 countries, that there is evidence of a positive relationship between the average inflation rate and pass-through.

Figure 8: Convergence to Equilibrium of Cointegrating Vectors in response to a System-wide Shock



A summary of the estimation results can be presented as follows:

$$p^{*} = p^{f} + e \tag{29}$$

$$p^{d} = 0.907p^{f} + 0.7e + 0.126nulc + 1.255tariff \tag{30}$$

$$m^{d} = 0.783p^{d} - 0.706p^{*} + 0.838y - 2.082tariff \tag{31}$$

Equation *(29)* describes the imposed small country assumption. Equations *(30)* and *(31)* reflect the cointegration results described in columns *(3)* and *(4)* of Table 10. Ignoring domestic price endogeneity and imposing the small country assumption, one can conclude that a 1% rise (depreciation) in the exchange rate would result in a 1% rise in the price of imports. This would therefore result in a 0.7% fall in the demand for imports. However, the net effect of an exchange rate depreciation can be found by substituting equations *(29)* and *(30)* into equation *(31)*:

$$m^{d} = 0.004p^{f} - 0.158e + 0.099nulc - 1.099tariff + 0.838y \tag{32}$$

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The net effect of a 1% rise in the exchange rate is only a 0.16% fall in the demand for imports when domestic prices are made endogenous. The importance of this study is illuminated here: the relative advantage gained by domestic goods from an exchange rate depreciation (since they become cheaper relative to imported goods) is eroded by the inflationary impact of the exchange rate. In addition a change in the import price caused by a depreciation in the currency is likely to have a lower real effect on both import volumes and the trade balance where the small country assumption does not hold (Menon (1995b)). Thus it is possible that if South Africa does not face an infinite import price elasticity the net effect of an exchange rate depreciation will be even lower than found here.

Similarly, the net effect of a 1% increase in foreign prices is negligible, raising the demand for imports by 0.004%. This is partly a result of imposing price homogeneity on import and foreign prices. Imposing this assumption is plausible if firms in emerging economies begin to follow more closely the developments in foreign prices as a result of increasing globalisation, liberalisation and the shifting of production from industrialised to developing countries (as suggested by Mihaljek and Klau (2001)). Additionally, the inflationary effect of tariffs on domestic prices reduces its effectiveness in curbing import demand. The net effect of a 1% increase in tariffs is only a 1.1% reduction in the demand for imports. This is lower than the "lone" tariff effect shown in column *(3)* of Table 10.

SHORT-RUN DYNAMICS

Insight into the short-run dynamics of both cointegrating vectors (given by the over-identifying restrictions in Table 8) is shown in the error correction equation results in Table 11. Only coefficients with a 95% level of significance are shown. The full error correction equations are presented in the appendix.

Noticeably, lagged foreign prices are insignificant in the import demand equation. This suggests that import demand is insensitive to foreign price shocks in the short run. It is immediately noticeable that the lagged changes in domestic prices are negatively related to import demand while the lagged changes in import prices have a positive effect on import demand. The coefficient signs on lagged domestic and import prices is unexpected. This may reveal the short run behaviour and expectations of consumers where a rise in import prices causes a "stocking up" of imported goods in anticipation of further price increases. The "stocking" of imported goods could also explain the positive relationship between the change

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in import demand and the lagged change in tariff rates. In the domestic price equation a positive change in both nominal unit labour costs and foreign prices feeds through into domestic prices, demonstrating the inflationary effects of wage increases and the sensitivity of domestic prices to world prices. This relationship may have strengthened following South Africa's rapid liberalisation in the 1990s.

Table 11: Error Correction Equations for Import Demand and Domestic Price Multivariate Framework, Quarterly Data, 1962Q1 – 2003Q2

DEPENDENT VARIABLE	DLM		DLPDOM	
dLM(-1)			0.0448 (0.0204)	
dLPFOR(-1)			0.3200 (0.1547)	
dLPDOM(-1)	-1.0403 (0.4945)			
dLPIMP(-1)	0.7304 (0.1697)			
dLEXRATE(-1)	-0.4875 (0.1437)			
dLY(-1)	0.4632 (0.2295)			
dLPIMP(-2)	0.5149 (0.1842)			
dlEXRATE(-2)	-0.5888 (0.1369)			
dLNULC(-2)			0.1873 (0.0541)	
dLY(-2)	0.6279 (0.2270)			
dLPDOM(-3)	-1.0091 (0.4898)			
dLPIMP(-3)	0.4489 (0.169)			
dLTARIFF(-3)	0.9507 (0.4787)			
Intercept	-6.1811 (0.9052)		-0.6275 (0.1470)	
ECM1(-1)	-0.6492 (0.0806)		-0.0398 (0.0131)	
ECM2(-1)	-0.5261 (0.1363)		-0.1061 (0.0221)	
DLOILP	-0.0648 (0.0272)			
EXRSHOCK			0.0101 (0.0039)	
FIXED			-0.0103 (0.0030)	
SANCTION	-0.0493 (0.0214)		-0.0073 (0.0035)	
POST94	0.0924 (0.0197)			
LYLMG *(ydgap)*	0.7169 (0.1609)		0.0833 (0.0261)	
	χ^2	F-test	χ^2	F-test
Serial Correlation	6.5795 {0.160}	1.3310 {0.262}	9.6312 {0.047}	1.9864 {0.100}
Functional Form	6.0654 {0.014}	5.0060 {0.027}	0.2057 {0.650}	0.1638 {0.686}
Normality	0.5607 {0.756}		8.4757 {0.014}	
Heteroscedasticity	3.5366 {0.060}	3.5700 {0.061}	1.2753 {0.259}	1.2697 {0.261}

Notes: Round brackets identify coefficient standard errors while curly brackets enclose p-values.

Oil price increases are seen to reduce import demand in South Africa, possibly due to switching of expenditure from other goods towards oil. The exchange rate shock variable

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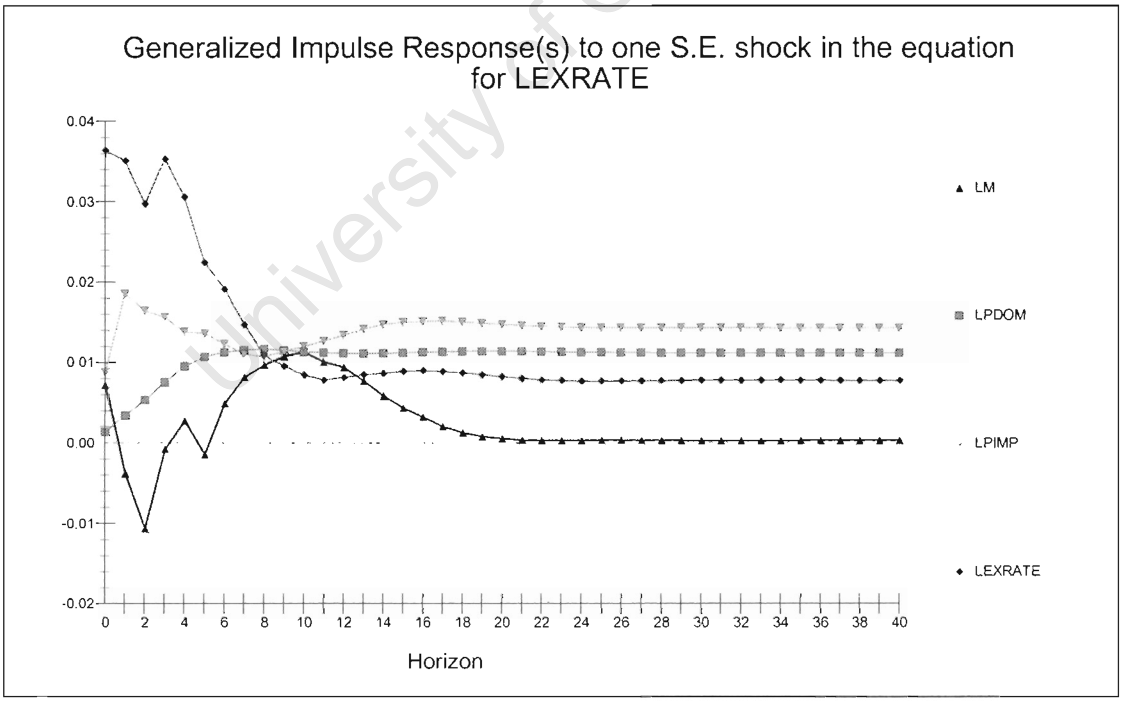
indicates the considerable impact depreciation shocks have on domestic prices, while the fixed exchange rate dummy shows the inflationary effect flexible exchange rates have had on South Africa. The sanction period is seen to be effective in reducing imports volumes as one would expect, but seems to have had a negative effect on domestic prices. This may be due to the correlation between the dummy variable and the recessionary period that South Africa experienced during this time. The liberalisation period after 1994 has seen a rise in import demand corresponding with the reduction of tariffs and trade barriers and the re-integration of South Africa into the global economy. *LYLMG*, which captures excess domestic demand, is positive and significant in both import demand and domestic price. While this effect is small in domestic prices, it contradicts Fedderke and Schaling's (2005) conclusion that economic cycles do not have an impact on inflationary pressure.

The dynamic effects of an exchange rate shock are analysed in Figure 9 below, which displays the impulse response of the endogenous variables to one standard error depreciation of the exchange rate.

Figure 9: Impulse Response of Variables to a Standard Error Shock in the Exchange Rate



The impact effect, given at period zero, indicates the correlation in errors between the exchange rate and other variables. From Figure 9 it is clear that import demand falls in response to the exchange rate depreciation in the short run. This is because import prices rise

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in the period immediately following the exchange rate shock. However, import demand actually rises in the medium term following a depreciation. This may be because import prices respond much quicker to an exchange rate depreciation (and the resultant exchange rate adjustment following a depreciation) than do domestic prices. In the medium term, import prices are falling as the exchange rate re-adjusts following the depreciation. Concurrently domestic prices are rising as the lagged effect of the depreciation feeds through. The combination of these two effects results in import demand rising in the medium term. It is clear that in the long run there is little effect of an exchange rate depreciation on import demand because of the inflationary effects a devaluation has on domestic prices. Finally, an exchange rate depreciation results in higher long run levels for both import and domestic prices.

8. CONCLUSION

There are several limitations to the elasticities approach based on the BRM condition in analysing the effect of an exchange rate change on import demand. These limitations stem from its partial equilibrium framework and its inherent inability to deal with domestic price endogeneity. Specifically, it is expected that its inability to deal with domestic price endogeneity may overstate the effect a depreciation may have on the trade balance. This paper uses the imperfect substitutes model to derive an import demand function for South Africa. The resultant elasticities can then be used in the BRM condition to determine the exchange rate effect on the trade balance.

Theory behind the imperfect substitutes model prescribes a simultaneous framework between import demand and supply to determine import prices. Yet the frequently used "small country" allows for import price exogeneity and simplifies the estimation of import demand by reducing it to a single equation model. Prior to using it, the small country assumption is tested for South Africa. By accounting for oil price effects, two different tests reveal that the assumption of infinite world supply holds for South Africa. However, the use of different foreign export price variables indicates that there is some scope for clarification of this issue.

The import demand curve is estimated with endogenous domestic prices within a simultaneous equation framework. The Johansen multivariate cointegration technique is used to account for nonstationarity and to allow for simultaneous equation analysis. The results

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show, consistent with the literature, that an exchange rate depreciation raises import prices and reduces import demand (and therefore improves the trade balance). However, once the inflationary effects of an exchange rate depreciation are accounted for, the exchange effect on import demand and the trade balance is significantly reduced.

Ultimately the paper shows that the import demand price elasticities suggest an unambiguous improvement in the trade balance when substituted into the small country variant of the BRM condition. However, this result is found to be overstated because the BRM condition explicitly ignores the inflationary effects of an exchange rate depreciation on domestic prices. In the long run this paper shows that the effect of a depreciation on import demand may be close to zero.

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APPENDIX

Table A1: SARB and Revised Weights used in Exchange Rate and Foreign Price Construction

	SARB WEIGHTS		REVISED WEIGHTS (1) USED FOR EXRATE AND LPFOR		REVISED WIEGHTS (2) USED FOR XUV AND MUV
	BEFORE 1999	POST 1999	PRIOR TO 1999	POST 1999	
Austria	0.83		0.91		
Belgium - Luxembourg	3.54		3.89		
Finland	0.81		0.89		
France	4.98		5.47		
Germany	16.91		18.56		22.84
Ireland	0.86		0.94		1.16
Italy	5.07		5.57		6.85
Netherlands	3.9		4.28		5.27
Portugal	0.34		0.37		
Spain	1.34		1.47		1.81
Australia	1.59	1.62	1.75	1.78	2.15
Canada	1.87	1.93	2.05	2.12	2.53
China Hong Kong	2.59	2.62			
China Mainland	2.91	3.11			
Euro		35.7		39.31	
Israel	1.14	1.17			
Japan	9.9	10.26	10.87	11.30	13.37
Korea	2.5	2.57	2.74	2.83	3.38
Singapore	1.55	1.62	1.71	1.78	
Sweden	1.58	1.79	1.73	1.97	2.13
Switzerland	4.99	5.28	5.48	5.82	
UK	14.09	14.91	15.47	16.41	19.03
USA	14.44	15.15	15.85	16.68	19.50
Zimbabwe	2.27	2.27			
Total	100	100	100	100	100

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Table A2: Data Sources for Exchange Rate and Foreign Price Variables

COUNTRY	EXCHANGE RATE	PPI	CPI
Austria	(IFS) 122..RF.ZF...	(IFS) 12263...ZF...	
Belgium	(IFS) 124..WF.ZF...	(IFS) 12463B..ZF...	(IFS) 12464...ZF...
Finland	(IFS) 172..RF.ZF...	(INTL) M172PJTT	
France	(IFS) 132..RF.ZF...	(IFS) 13263A..ZF...	(IFS) 13264...ZF...
Germany	(IFS) 134..RF.ZF...	(INTL) M134PJTTA	
Ireland	*(IFS) 178..RH.ZF...	(IFS) 17863...ZF...	
Italy	(IFS) 136..RF.ZF...	(IFS) 13663...ZF...	(IFS) 13664...ZF...
Netherlands	(IFS) 138..RF.ZF...	(IFS) 13863...ZF...	
Portugal	(IFS) 182..RF.ZF...	(IFS) 18263...ZF...	(IFS) 18264...ZF...
Spain	(IFS) 184..RF.ZF...	(IFS) 18463...ZF...	
EURO Area	(RBQN) RB5315M		
Australia	*(IFS) 193..RH.ZF...	(IFS) 19363...ZF...	
Canada	(IFS) 156..RF.ZF...	(IFS) 15663...ZF...	
Japan	(IFS) 158..RF.ZF...	(IFS) 15863...ZF...	(IFS) 15864...ZF...
Korea	(IFS) 542..RF.ZF...	(IFS) 54263...ZF...	
Singapore	(IFS) 576..RF.ZF...	(IFS) 57663...ZF...	(IFS) 57664...ZF...
Sweden	(IFS) 144..RF.ZF...	(IFS) 14463...ZF...	(IFS) 14464...ZF...
Switzerland	(IFS) 146..RF.ZF...	(IFS) 14663...ZF...	(IFS) 14664...ZF...
United Kingdom	*(IFS) 112..RH.ZF...	(IFS) 11263...ZF...	
South Africa (USA)	(IFS) 199..WF.ZF...	(IFS) 11163...ZF...	

IFS = IMF International Financial Statistics, INTL = IMF International Key Macroeconomics Indicators, RBQN = South African Reserve Bank Quarterly Database.
All exchange rates are in foreign currency per US dollar denomination, with the exception of exchange rates with an asterisk (*) which are in US dollar per foreign currency denomination. The EURO exchange rate denominated in South African cents per EURO.
Prior to construction of the nominal effective exchange rate variable (*EXRATE*) all currencies were converted to South African rands per foreign currency denomination and indexed to 1995 = 100.
PPI data reflects wholesale, industrial and producer prices for the countries. For countries where there are CPI codes, consumer prices were used to construct PPI data for periods between 1961 and 2003 that were not available. All country prices were indexed to 1995 = 100 prior to construction of *PFOR*.

Table A3: Data Sources for Export and Import Unit Values (IMF International Financial Statistics)

COUNTRY	EXPORT UNIT VALUES	IMPORT UNIT VALUES
Germany	13474...ZF...	13475...ZF...
Ireland	17874...ZF...	17875...ZF...
Italy	13674...ZF...	13675...ZF...
Netherlands	13874...ZF...	13875...ZF...
Spain	18474...ZF...	18475...ZF...
Australia	19374..DZF...	19375..DZF...
Canada	15674...ZF...	15675...ZF...
Japan	15874...ZF...	15875...ZF...
Korea	54274...ZF...	54275...ZF...
Sweden	14474..DZF...	14475..DZF...
United Kingdom	11274..DZF...	11275..DZF...
USA	11174..DZF...	11175..DZF...

All unit values were converted to a 1995 = 100 index prior to construction of *XUV* and *MUV*.

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Table A4: Data Sources and Description for Other Variables

	(DATABASE) CODE	DESCRIPTION
MVOLQ	(RBQN) RB5034L	Imports: Volume (Unit: Index 2000=100, seasonally adjusted (Period))
PDOMM	(RBQN) RB7048N	Production prices of goods produced for domestic use: Goods produced in SA - Total (Unit: Index June 2000=100, seasonally adjusted (Period))
PIMPQ	(RBQN) RB5035L	Imports: Price (Unit: Index 2000=100, seasonally adjusted (Period))
NOM COMPENSATION	(RBQN) RB6240L	Compensation of employees (Unit: R millions, current prices, seasonally adjusted (Period))
REAL VALUE ADDED	(RBQN) RB6645D	At constant 2000 prices: Gross value added at basic prices of all industries (Unit: R millions, seasonally adjusted at annual rate (Period))
	(RBQN) RB6645L	At current prices: Gross value added at basic prices of all industries (Unit: R millions, seasonally adjusted at annual rate (Period))
Y	(RBQN) RB6012D	Expenditure on gross domestic product at constant 1995 prices, seasonally adjusted (quarterly): Gross domestic expenditure (Unit: R millions, seasonally adjusted at annual rate (Period))
ME	(RBQN) RB6014D	Expenditure on gross domestic product at constant 1995 prices, seasonally adjusted (quarterly): Imports of goods and services (Unit: R millions, seasonally adjusted at annual rate (Period))
CUSTOMS DUTIES	kbp4005m	Customs duties (Lawrence Edwards)
SURCHARGES	kbp4006m	Customs surcharge (Lawrence Edwards)
MERCHANDISE IMPORTS	his5003j	Merchandise imports (BOP) (Lawrence Edwards)
OILPM	(INTL) M163VSPENCOBH	EURO-ZONE WHOLESALE AND RETAIL SALES WORLD MARKET PRICES OF RAW MATERIALS HWWA COMMODITY PRICE - RAW MATERIALS: CRUDE OIL (MONTHLY AVERAGES (Unit: ECU--EURO PER BARREL NSA)

Table A5: Dummy Variables

VARIABLE	DUMMY PERIOD	COMMENT
FIXED	1961Q1 – 1971Q1	President Nixon drops the fixing of the US dollar to the gold price in 1971, beginning a period of flexible exchange rates and inflation globally.
EXRSHOCK	1985Q3 – 1985Q4 1996Q2 1998Q3 2001Q4 – 2002Q1	Rubicon Speech and resulting severe capital outflow. Asian crisis impacts on South Africa. Contagion from Asian Crisis, spreads to Russia, Brazil and other developing countries. Rand experiences severe speculative shock causing sudden depreciation.
SANCTION	1986Q2 – 1993Q3	The UN Security Council imposes full sanctions on top of the arms embago in June 1986. The UN General Assembly commits to lift all sanctions against South Africa in October 1993.
POST94	1994Q3 – 2003Q2	Captures post-election liberalization period.

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Table A6: Construction of Variables used in Model Specification

VARIABLE	NAME	MODEL NAME	CONSTRUCTION OF VARIABLE / TRANSFORMATION OF RAW DATA
M_t^d	Import Volume	M	MVOLQ index changed to 1995 = 100
P_t^f	Foreign Price	PFOR XUV MUV	$PFOR = \prod_{i=1}^{n}(pp_i^{W_i})$ $XUV = \prod_{k=1}^{n}(xuv_k^{W_k})$ $MUV = \prod_{k=1}^{n}(muv_k^{W_k})$ pp_i = producer price for country I w_i = revised weight for country i For PFOR Revised Weights (1) prior to 1999 used. For XUV and MUV Revised Weights (2) were used in calculation.
P_{Dt}	Domestic Price	PDOM	PDOMM averaged from monthly data, indexed to 1995 = 100
P_t^*	Import Price	PIMP	PIMPQ index changed to 1995 = 100
E_t	Nominal Effective Exchange Rate	EXRATE	$EXRATE = \prod_{j=1}^{n}(ex_j^{W_j})$ ex_j = producer price for country j w_j = revised weight for country j Revised Weights (1) used in calculation. Euro country weights used up to 1st Janaury 1999, EURO Area weight substituted for EURO countries after this date.
ULC_t	Nominal Unit Labout Cost	NULC	$NULC = \frac{NOM\ COMPENSATION}{REAL\ VALUE\ ADDED}$ Real Value Added converted to 1995 prices using implicit deflator
Y_t	Real Gross Domestic Expenditure	Y	
	Domestic Expenditure	YLM	Y - ME
$TARIFF_t$	Average Tariff Rate	TARIFF	$TARIFF = 1 + \left(\frac{CUSTOMS\ DUTIES + SURCHARGES}{MERCHANDISE\ IMPORTS}\right)$
	Oil Price	OILP	OILPM averaged from monthly data, indexed to 1995 = 100
$ydgap_t$	Excess Demand For Domestic Goods	LYLMG	$LYLMG = LYLM - LYLMH$ where *LYLM* is the (log) domestic expenditure less imports of goods and services and (log) *LYLMH* is expenditure smoothed used the HP filter with a lamda of 1600
	Excess Demand for All Goods	LYG	$LYG = LY - LYH$ where *LYH* is (log) expenditure smoothed using the HP filter with a lamda of 1600

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Table A7: Test Statistics for Cointegrating Rank, Import Price Equation, using LXUV, Unrestricted Intercepts and No Trends in the VAR, 156 observations from 1964Q1 to 2002Q4. Order of VAR = 4

Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix

List of variables included in the cointegrating vector:
LPIMP LXUV LEXRATE LTARIFF LOILP
List of I(1) exogenous variables included in the VAR:
LOILP
List of I(0) variables included in the VAR:
LYG
List of eigenvalues in descending order:
.29758 .079965 .054439 .0023727 .0000

Null	Alternative	Statistic	95% Critical Value	90%Critical Value
r = 0	r = 1	55.1020	30.7100	28.2700
r<= 1	r = 2	13.0016	24.5900	22.1500
r<= 2	r = 3	8.7324	18.0600	15.9800
r<= 3	r = 4	.37058	11.4700	9.5300

Cointegration LR Test Based on Trace of the Stochastic Matrix

List of variables included in the cointegrating vector:
LPIMP LXUV LEXRATE LTARIFF LOILP
List of I(1) exogenous variables included in the VAR:
LOILP
List of I(0) variables included in the VAR:
LYG
List of eigenvalues in descending order:
.29758 .079965 .054439 .0023727 .0000

Null	Alternative	Statistic	95% Critical Value	90%Critical Value
r = 0	r>= 1	77.2066	58.6300	54.8400
r<= 1	r>= 2	22.1045	38.9300	35.8800
r<= 2	r>= 3	9.1029	23.3200	20.7500
r<= 3	r = 4	.37058	11.4700	9.5300

Choice of the Number of Cointegrating Relations Using Model Selection Criteria

List of variables included in the cointegrating vector:
LPIMP LXUV LEXRATE LTARIFF LOILP
List of I(1) exogenous variables included in the VAR:
LOILP
List of I(0) variables included in the VAR:
LYG
List of eigenvalues in descending order:
.29758 .079965 .054439 .0023727 .0000

Rank	Maximized LL	AIC	SBC	HQC
r = 0	1599.4	1531.4	1427.7	1489.3
r = 1	1626.9	1550.9	1435.1	1503.9
r = 2	1633.4	1551.4	1426.4	1500.7
r = 3	1637.8	1551.8	1420.7	1498.6
r = 4	1638.0	1550.0	1415.8	1495.5

AIC = Akaike Information Criterion SBC = Schwarz Bayesian Criterion
HQC = Hannan-Quinn Criterion

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Table A8: Test Statistics for Cointegrating Rank, Import Price Equation, using LXUV, Unrestricted Intercepts and No Trends in the VAR, 155 observations from 1964Q2 to 2002Q4. Order of VAR = 5

Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix

****

List of variables included in the cointegrating vector:
LPIMP LMUV LEXRATE LTARIFF LOILP
List of I(1) exogenous variables included in the VAR:
LOILP
List of I(0) variables included in the VAR:
LYG
List of eigenvalues in descending order:
.21148 .084638 .071653 .0017051 0.00

****

Null	Alternative	Statistic	95% Critical Value	90%Critical Value
r = 0	r = 1	36.8280	30.7100	28.2700
r<= 1	r = 2	13.7076	24.5900	22.1500
r<= 2	r = 3	11.5242	18.0600	15.9800
r<= 3	r = 4	.26451	11.4700	9.5300

****

Cointegration with unrestricted intercepts and no trends in the VAR
Cointegration LR Test Based on Trace of the Stochastic Matrix

****

List of variables included in the cointegrating vector:
LPIMP LMUV LEXRATE LTARIFF LOILP
List of I(1) exogenous variables included in the VAR:
LOILP
List of I(0) variables included in the VAR:
LYG
List of eigenvalues in descending order:
.21148 .084638 .071653 .0017051 0.00

****

Null	Alternative	Statistic	95% Critical Value	90%Critical Value
r = 0	r>= 1	62.3243	58.6300	54.8400
r<= 1	r>= 2	25.4963	38.9300	35.8800
r<= 2	r>= 3	11.7887	23.3200	20.7500
r<= 3	r = 4	.26451	11.4700	9.5300

****

Choice of the Number of Cointegrating Relations Using Model Selection Criteria

****

List of variables included in the cointegrating vector:
LPIMP LMUV LEXRATE LTARIFF LOILP
List of I(1) exogenous variables included in the VAR:
LOILP
List of I(0) variables included in the VAR:
LYG
List of eigenvalues in descending order:
.21148 .084638 .071653 .0017051 0.00

****

Rank	Maximized LL	AIC	SBC	HQC
r = 0	1543.8	1455.8	1321.8	1401.4
r = 1	1562.2	1466.2	1320.1	1406.8
r = 2	1569.0	1467.0	1311.8	1404.0
r = 3	1574.8	1468.8	1307.5	1403.3
r = 4	1574.9	1466.9	1302.6	1400.2

****

AIC = Akaike Information Criterion SBC = Schwarz Bayesian Criterion
HQC = Hannan-Quinn Criterion

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Table A9: ECM for Import Demand estimated by OLS, based on Cointegrating VAR(4), 166 observations used for estimation from 1962Q1 to 2003Q2

Dependent variable is dLM

Regressor	Coefficient	Standard Error	T-Ratio[Prob]
Intercept	-6.1811	.90521	-6.8284[.000]
dLM1	-.17991	.12568	-1.4315[.155]
dLPFOR1	-1.2177	.95269	-1.2782[.203]
dLPDOM1	-1.0403	.49448	-2.1039[.037]
dLPIMP1	.73043	.16968	4.3048[.000]
dLEXRATE1	-.48753	.14368	-3.3932[.001]
dLNULC1	-.045075	.33097	-.13619[.892]
dLY1	.46318	.22945	2.0186[.046]
dLTARIFF1	.93851	.49793	1.8848[.062]
dLM2	-.13698	.11733	-1.1675[.245]
dLPFOR2	.88509	1.1721	.75514[.452]
dLPDOM2	-.39234	.49085	-.79930[.426]
dLPIMP2	.51491	.18419	2.7955[.006]
dLEXRATE2	-.58883	.13688	-4.3018[.000]
dLNULC2	-.33425	.33327	-1.0029[.318]
dLY2	.62785	.22695	2.7665[.006]
dLTARIFF2	.76493	.49423	1.5477[.124]
dLM3	-.10671	.10278	-1.0382[.301]
dLPFOR3	.33392	1.0076	.33139[.741]
dLPDOM3	-1.0091	.48984	-2.0601[.041]
dLPIMP3	.44887	.16900	2.6560[.009]
dLEXRATE3	-.083628	.13512	-.61892[.537]
dLNULC3	-.0095072	.35556	-.026739[.979]
dLY3	.34858	.21437	1.6260[.106]
dLTARIFF3	.95072	.47872	1.9860[.049]
ecm1(-1)	-.64922	.080619	-8.0529[.000]
ecm2(-1)	-.52613	.13634	-3.8589[.000]
DLOILP	-.064839	.027198	-2.3839[.019]
EXRSHOCK	-.0033797	.023957	-.14107[.888]
SANCTION	-.049311	.021367	-2.3078[.023]
POST94	.092451	.019657	4.7032[.000]
FIXED	-.0064144	.018656	-.34383[.732]
LYLMG	.71694	.16091	4.4556[.000]

ecm1 = 1.0000*LM - .0000*LPFOR - .78332*LPDOM + .70617*LPIMP - .0000*LEXRATE - .0000*LNULC - .83784*LY + 2.0820*LTARIFF

ecm2 = - .0000*LM - .90761*LPFOR + 1.0000*LPDOM - .0000*LPIMP - .70034*LEXRATE - .12551*LNULC + .0000*LY - 1.2547*LTARIFF

R-Squared	.61521	R-Bar-Squared	.52263
S.E. of Regression	.049356	F-stat. F(32, 133)	6.6450[.000]
Mean of Dependent Variable	.0098697	S.D. of Dependent Variable	.071435
Residual Sum of Squares	.32399	Equation Log-likelihood	282.2964
Akaike Info. Criterion	249.2964	Schwarz Bayesian Criterion	197.9486
DW-statistic	1.8606	System Log-likelihood	3846.6

Diagnostic Tests

Test Statistics	LM Version	F Version
A:Serial Correlation	CHSQ(4)= 6.5795[.160]	F(4, 129)= 1.3310[.262]
B:Functional Form	CHSQ(1)= 6.0654[.014]	F(1, 132)= 5.0060[.027]
C:Normality	CHSQ(2)= .56065[.756]	Not applicable
D:Heteroscedasticity	CHSQ(1)= 3.5366[.060]	F(1, 164)= 3.5700[.061]

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Table A10: ECM for Domestic Price estimated by OLS, based on Cointegrating VAR(4), 166 observations used for estimation from 1962Q1 to 2003Q2

Dependent variable is dLPDOM

Regressor	Coefficient	Standard Error	T-Ratio[Prob]
Intercept	-.62749	.14700	-4.2686[.000]
dLM1	.044789	.020410	2.1945[.030]
dLPFOR1	.32000	.15471	2.0684[.041]
dLPDOM1	.050251	.080301	.62578[.533]
dLPIMP1	.034393	.027555	1.2482[.214]
dLEXRATE1	-.016241	.023332	-.69608[.488]
dLNULC1	.0077323	.053747	.14386[.886]
dLY1	-.028038	.037261	-.75248[.453]
dLTARIFF1	-.053083	.080862	-.65647[.513]
dLM2	.021053	.019054	1.1049[.271]
dLPFOR2	-.14123	.19034	-.74196[.459]
dLPDOM2	-.12297	.079712	-1.5427[.125]
dLPIMP2	.043305	.029912	1.4477[.150]
dLEXRATE2	.0090534	.022229	.40729[.684]
dLNULC2	.18731	.054121	3.4609[.001]
dLY2	-.0033420	.036855	-.090681[.928]
dLTARIFF2	-.039970	.080260	-.49801[.619]
dLM3	.0033034	.016691	.19792[.843]
dLPFOR3	.10851	.16363	.66311[.508]
dLPDOM3	.064750	.079547	.81399[.417]
dLPIMP3	.020929	.027445	.76258[.447]
dLEXRATE3	.021597	.021943	.98425[.327]
dLNULC3	.10420	.057741	1.8047[.073]
dLY3	-.036342	.034813	-1.0439[.298]
dLTARIFF3	.11209	.077741	1.4418[.152]
ecm1(-1)	-.039821	.013092	-3.0416[.003]
ecm2(-1)	-.10611	.022141	-4.7927[.000]
DLOILP	.0017567	.0044168	.39772[.691]
EXRSHOCK	.010053	.0038905	2.5839[.011]
SANCTION	-.0072539	.0034699	-2.0905[.038]
POST94	.9409E-3	.0031922	.29475[.769]
FIXED	-.010263	.0030296	-3.3875[.001]
LYLMG	.083334	.026130	3.1892[.002]

ecm1 = 1.0000*LM - .0000*LPFOR - .78332*LPDOM + .70617*LPIMP - .0000*LEXRATE - .0000*LNULC - .83784*LY + 2.0820*LTARIFF

ecm2 = - .0000*LM - .90761*LPFOR + 1.0000*LPDOM - .0000*LPIMP - .70034*LEXRATE -.12551*LNULC + .0000*LY - 1.2547*LTARIFF

R-Squared	.73482	R-Bar-Squared	.67101
S.E. of Regression	.0080151	F-stat. F(32, 133)	11.5169[.000]
Mean of Dependent Variable	.021972	S.D. of Dependent Variable	.013974
Residual Sum of Squares	.0085442	Equation Log-likelihood	584.0390
Akaike Info. Criterion	551.0390	Schwarz Bayesian Criterion	499.6912
DW-statistic	2.0776	System Log-likelihood	3846.6

Diagnostic Tests

Test Statistics	LM Version	F Version
A:Serial Correlation	CHSQ(4)= 9.6312[.047]	F(4, 129)= 1.9864[.100]
B:Functional Form	CHSQ(1)= .20568[.650]	F(1, 132)= .16376[.686]
C:Normality	CHSQ(2)= 8.4757[.014]	Not applicable
D:Heteroscedasticity	CHSQ(1)= 1.2753[.259]	F(1, 164)= 1.2697[.261]

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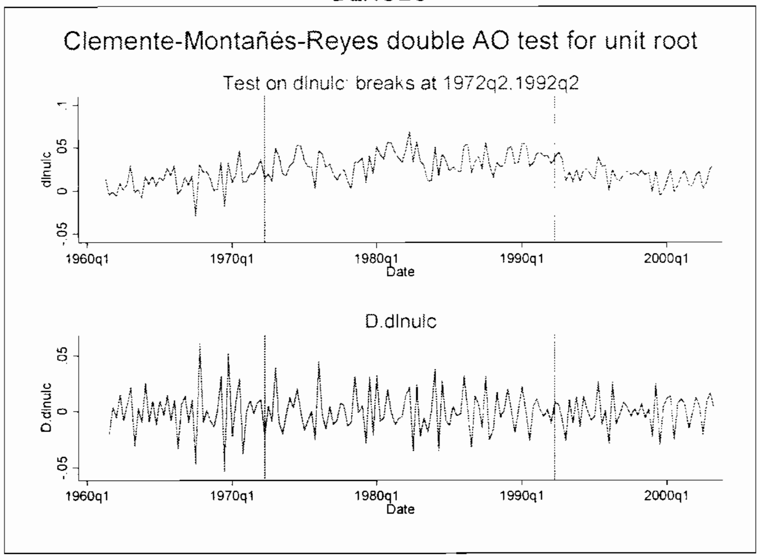
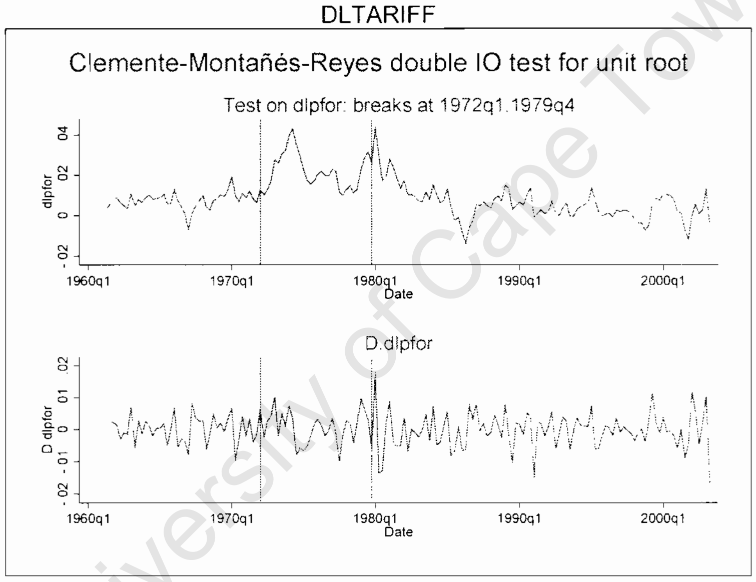
Figure A1: Clemente-Montanes Reyes Structural Break Test for DLNULC and DLTARIFF

DLNULC



DLTARIFF



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