Foreign Trade Regimes and Import Demand Function: Evidence from Sri Lanka

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Abstract

Time series data for Sri Lanka span periods of pervasive trade and exchange restrictions along with periods of liberalized trade. This paper implements a structural econometric model of aggregate imports which incorporates the implications of the shifts in the policy regime. The results demonstrate that the model outperforms the existing alternatives both on statistical and economic grounds. The estimated elasticities, in contrast to the available evidence, have correct signs, high statistical significance, and plausible magnitudes. The implications for policy analysis like calculation of equilibrium exchange rate are discussed.

Keywords: Trade Policy, Import Demand, Sri Lanka, Cointegration, Intertemporal Substitution

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Introduction

The econometric estimates of price and income elasticities of imports (both aggregate and disaggregate) are critical inputs to important policy analyses such as calculation of equilibrium exchange rate, the design of optimal trade taxes, and estimation of the fiscal implications of trade liberalization.[2] The use of inappropriate estimates of the elasticity parameters in the analysis and formulation of policy might prove very costly. For example, a gross miscalculation of the extent of the overvaluation of exchange rate due to a significantly biased estimate of the price elasticity of aggregate imports could result in the loss of export competitiveness.

In this paper, we argue that the changing trade and exchange rate policies have critical bearings on the econometric modeling of aggregate imports as they determine a country's overall capacity to import. To adequately capture the implications of the changing trade and exchange rate regime of a country, an in-depth and detailed study of the history of the policy regime is necessary. It is, however, extremely difficult, if not impossible, to give sufficient attention to the country-specific policy changes when working with a large number of countries, as is the case in a number of recent studies. This is why the same generic model is implemented across different countries, ranging from a country like India where until recently pervasive import controls had been the norm, to a country like USA where import regime is completely liberalized (see, for example, Caporale and Chui, (1999); Bahmani-Oskooee and Niroomand, (1998), Senhadji, (1998)). In this paper, our focus is on a single country, Sri Lanka. We explore the implications of the changing policy regime in a structural model of aggregate imports.

The time series data available for Sri Lanka, as for most developing countries, span the historical periods of pervasive trade and exchange restrictions along with periods of liberalized trading regime.[3] To be sure, the trade and exchange rate interventions would

[2]For instance, Krueger, Schiff and Valdez(1991) report that the estimates of the extent of overvaluation of exchange rates are highly sensitive to the assumed magnitude of the price elasticity of aggregate imports.

[3]During the 1960's and 1970's, Sri Lanka implemented a host of increasingly restrictive protectionist policies which severely restricted its foreign trade. The sweeping economic reforms implemented in Novem-

not have created any problem for a researcher, if the period covering liberalized trade were sufficiently long enough to allow estimation with some statistical confidence. In this case, one could simply exclude the periods of constrained imports from the sample.[4] On the other hand, estimation could proceed for the sample period including the constrained regime, if right kind of data were available; most importantly, data on the scarcity price (administered price plus scarcity premium) of imports. In the presence of extensive secondary markets for import licenses and imported goods, the secondary market prices are the appropriate scarcity prices for imports relevant for consumer optimization. Unfortunately, for Sri Lanka, such price data are not available for the relevant sample period.

Although there is a large literature on the estimation of price and income elasticities of aggregate imports of developing countries, the problem of unavailability of appropriate price data for the constrained import regime has not been satisfactorily addressed. While the standard import model with income and relative price has been the work-horse in the literature (for recent examples, see Bahmani-Oskooee and Niroomand (op cit), Senhadji (op cit), Sinha (1999), Bahmani-oskooee (1986)), some researchers have added a foreign exchange availability variable on an *ad hoc* basis to an otherwise standard import demand model to reflect a binding foreign exchange constraint (for example, see Mazeri (1995), Moran (1989)). The inadequacies of a standard demand model might manifest themselves in (*i*) the absence of a long run equilibrium relationship among the variables and (*ii*) the theoretically inconsistent signs and economically implausible magnitudes of the effects of relative price and income. For example, Sinha (1999) reports a negative income elasticity (-0.39) for Sri Lanka, and a number of studies find that the price elasticity is both statistically and economically insignificant for India (estimated elasticity is -0.13 with a 't' statistic of -0.25 in Senhadji (1998); and -0.03 with a standard error of 0.35 (DOLS esti-

ber 1977, and its continuation in the ensuing years, transformed Sri Lanka from a virtually closed economy to a highly liberalized and outward oriented one (Athukorala and Rajapatirana(1999)).

[4] This is not an attractive option even in a country like Sri Lanka which implemented trade liberalization early on compared to most of the developing countries. Only a very small sample size starting from 1978 is available if one restricts to the liberalized period alone, and it severely compromises the usefulness of the cointegration approach to uncover a long run relation.

mate) in Caporale and Chui (1999)). The other approach which we call *foreign exchange availability formulation* suffers from the problem that if foreign exchange availability is used as a regressor when the foreign exchange constraint is binding, it alone determines the volume of imports completely resulting in a *near identity problem.* For example, Emran and Shilpi (1996) find that the estimated price elasticity of aggregate import demand of Bangladesh is positive when foreign exchange availability is defined as export earnings plus remittances plus disbursed foreign aid. To appreciate the pitfalls involved in using the foreign exchange availability formulation in case of Sri Lanka, consider the simple OLS regression where imports are regressed on foreign exchange availability and a constant. The estimated coefficient on foreign exchange availability is 0.85 with a 't' value of 15.37 and a $\bar{R}^2 = 0.87$. As is clear, the foreign exchange availability almost completely determines the imports.

To analyze the aggregate imports of Sri Lanka, we use a structural econometric model of a two goods representative agent economy that incorporates a binding foreign exchange constraint at the administered prices of imports. By parameterizing the Lagrange multiplier associated with the binding foreign exchange constraint in terms of the *ratio* of income to foreign exchange resources available to a country, the model avoids the pitfalls of both traditional and foreign exchange availability models. In this paper, we apply that model to the case of Sri Lanka taking into account the changing trade and exchange rate regime. We compare and contrast the results of our preferred model with those of a modified traditional model and the foreign exchange availability formulation.

The rest of the paper is organized as follows. The first section presents a simple intertemporal optimization model of a representative consumer to derive a theoretically consistent and empirically implementable specification of aggregate imports under significant policy shifts. Section 2 presents the empirical implementation of the model developed in section 1 with data from Sri Lanka. The paper ends with some concluding remarks on the implications of the elasticity estimates for Sri Lanka.

(1) A Model of Aggregate Imports Under Binding Foreign Exchange Constraint

In this section, we present a brief exposition of the model of aggregate imports under a binding foreign exchange constraint used to model the imports of Sri Lanka . [5] The rational expectations permanent income model of a representative agent is used to derive the import demand function. The representative agent consumes two composite goods: a home good (H_t) and an imported good (M_t). The feasibility set of the optimization problem is defined by two constraints: a dynamic budget constraint describing the asset accumulation, and an inequality constraint describing the foreign exchange availability constraint.[6] Let P_t denote the relative price of imports at administered exchange rate; A_t, assets; $\tilde{Y}_t$, labor income; F_t, the total amount of foreign exchange available; and r, the constant real interest rate. We take home goods as the numeraire and all the variables above are expressed in terms of it. The representative agent discounts the future by the subjective rate of time preference δ. The optimization problem of the representative agent is as follows:

$$Max_{[H_t,M_t,A_t]}V = E\int_{t=0}^{\infty} e^{-\delta t}U(H_t, M_t)dt$$

subject to

$$\dot{A} = rA_t + \tilde{Y}_t - H_t - P_tM_t \tag{1}$$

$$P_tM_t \leq F_t \tag{2}$$

where a dot above any variable denotes a time derivative, i.e., $\dot{A} = \frac{dA_t}{dt}$. If constraint (2) is binding then the volume of imports is equal to the foreign exchange available and the standard price and income variables are irrelevant.[7] The current value Hamiltonian of the

[5] For details of the theoretical model, see Emran and Shilpi (2000).

[6] This subsumes the effects of both the quantitative restrictions and foreign exchange overvaluation in a single foreign exchange constraint.

[7] This is the source of the near-identity problem in the standard foreign exchange availability approach. Also, observe that foreign exchange availability is treated as exogenous. Obviously this is a simplification

optimization problem can be written as:

$$L = U(H_t, M_t) + \lambda_t[rA_t + \tilde{Y}_t - H_t - P_tM_t] + \mu_t[F_t - P_tM_t]$$

where λ_t is the costate variable and μ_t is the Lagrange multiplier associated with the foreign exchange constraint. Following Clarida (1994), we assume that $U(.)$ is an addilog utility function:

$$U(H_t, M_t) = C_t\frac{H_t^{1-\alpha}}{1-\alpha} + B_t\frac{M_t^{1-\eta}}{1-\eta}$$

where C_t and B_t are random, strictly stationary shocks to preference.

With the above utility function, the first order conditions of the optimization problem are as follows:

$$C_tH_t^{-\alpha} = \lambda_t \quad (3)$$

$$B_tM_t^{-\eta} = P_t\lambda_t(1+\mu_t^*) = \lambda_tP_t^* \quad (4)$$

$$\dot{\lambda} = (\delta - r)\lambda_t \quad [F_t - P_tM_t] \geq 0;\ \mu_t[F_t - P_tM_t] = 0 \quad (5)$$

where $\mu_t^* = \frac{\mu_t}{\lambda_t} = \frac{\mu_t}{U_H}$ is the scarcity premium, and P_t^* is the scarcity price at which transactions occur at the shop floor in the secondary market or the *virtual price* in the terminology of Neary and Roberts (1980) if the secondary market fails to clear. Use equation (3) to eliminate λ_t from equation (4) and take logarithm to get the following equation:

$$b_t - \eta m_t = c_t + p_t - \alpha h_t + \ln(1+\mu_t^*) \quad (6)$$

where the lower case letters denote natural logarithm of the corresponding upper case letters.

that helps to focus on the modeling of scarcity premia on imports. In a fully specified general equilibrium model, the decisions of exporters and of international migrants (for remittances) will be endogenous, and a full macro-econometric model needs to be estimated. In the empirical work, we define foreign exchange availability as disbursed foreign aid plus exports plus remittances plus foreign exchange reserve. The econometric approaches used are robust to the endogeneity of the regressors.

The steady state solution implies that:

$$Y^* = H + P^* M \tag{7}$$

where Y^* is the total household income including both labor and asset income evaluated at the equilibrium price P^*. Using this steady state condition and taking natural logarithm, we get the following expression for h_t :

$$\begin{aligned} h_t &= \ln(Y_t^* - P_t^* M_t) \\ &\equiv \ln(Y_t - P_t M_t) \end{aligned} \tag{8}$$

where $Y_t = (Y_t^* - \mu_t^* P_t M_t)$ is the observed income in a foreign exchange constrained regime and reported in the national income accounts. Now use equation (8) to eliminate h_t from equation (6) and solve for m_t:

$$m_t = \frac{\alpha}{\eta} \ln(Y_t - P_t M_t) - \frac{1}{\eta} p_t - \frac{1}{\eta} \ln(1 + \mu_t^*) + \xi_t \tag{9}$$

where $\xi_t = \frac{1}{\eta}(b_t - c_t)$ is the composite preference shock. Note that if the foreign exchange constraint is not binding, then μ_t^* is equal to zero, and equation (9) provides the standard double-log specification similar to those estimated by numerous studies for both developed and developing countries (see the surveys by Goldstein and Khan, (1985), and Faini et. al., (1992)). Observe that Y is the total expenditure by *domestic* consumers on both domestically produced goods and imports. The scale variable $\ln(Y_t - P_t M_t)$ in the right hand side of equation (9) can thus be defined as GDP minus exports. When the foreign exchange constraint is binding, the Kuhn-Tucker theorem requires that $\mu_t > 0$, and hence $\mu_t^* > 0$.

The problem with equation (9) for econometric implementation is that time series data on μ_t^*, the scarcity premia on imports, are not available for most of the developing countries. To arrive at an estimable import equation, we need a theoretically consistent parameterization of μ_t^* in terms of the observed variables. Since μ_t^* represents the scarcity premia on foreign exchange, it should be, *ceteris paribus*, a negative function of the amount of foreign exchange available. So one would tend to think that a good proxy for μ_t^* can be the availability of foreign exchange, thus providing an *ex-post* rationalization of the widely used foreign exchange availability approach. But, as we emphasized in the introduction, using foreign exchange availability as a regressor leads to the problem of *near identity* in a foreign exchange constrained regime. To avoid this problem, we parameterize μ_t^* by the ratio of total domestic expenditure ($= GDP + import - \exp ort$) to the available foreign exchange resources (denoted below as Z_t). The intuition behind this parameterization is that given a price vector determined by the world prices and the administered exchange rate, the excess demand for (and hence the scarcity premia on) the imported goods is (i) a negative function of foreign exchange availability keeping expenditure fixed, and (ii) a positive function of expenditure keeping foreign exchange availability fixed provided that imports are not inferior goods. More importantly, there is no one to one relation between imports and Z_t in a foreign exchange constrained regime, and it is not subject to the problem of near identity. Also, since the import regime in Sri Lanka was unconstrained after 1977, the scarcity premium is zero for the subsample of 1977-95 [see appendix 1 for a brief description of trade and exchange rate policy regimes in Sri Lanka]. To incorporate this a priori restriction, we transform Z_t by multiplying it by a dummy variable that takes on the value of 1 for 1960-1977 and zero afterwards. This transformed variable is denoted as Z_t^*. [8] Although the sign of the effects of a marginal change in Z_t^* on imports follow from the theory, it provides no guide as to the specification of the functional form of $\mu_t^*(Z_t^*)$ which might vary across different countries. To determine the appropriate functional form, we use a semiparametric approach. The results of the semiparametric analysis, the details of

[8] Since Z_t in an unconstrained regime is lower than in a constrained regime, the relationship between Z_t^* and μ_t^* stays the same as the relationship between Z_t and μ_t^*. The imposition of *a priori* theoretical restrictions by transforming the data series as is done above is a widely used practice in the empirical modeling of investment and consumption under imperfect credit and capital markets (See, for instance, Hubbard and Kashyap (1992) for an application to investment).

which are omitted for brevity, show that, for Sri Lanka, the relationship between Z_t^* and μ_t^* can be adequately represented by the following functional form:

$$\mu_t^*(Z_t^*) = e^{\theta_1 Z_t^*} - 1\ ;\ \theta_1 \geq 0 \tag{10}$$

With this specification, we have the following structural import demand function that can be estimated with available data for most of the developing countries:

$$\begin{aligned} m_t &= \frac{\alpha}{\eta}\ln(Y_t - P_t M_t) - \frac{1}{\eta}p_t - \frac{\theta_1}{\eta}Z_t^* + \xi_t \\ &= \pi_1 \ln(Y_t - P_t M_t) - \pi_2 p_t - \pi_3 Z_t^* + \xi_t \end{aligned} \tag{11}$$

Note that the parameters (α, η, θ_1) are just identified in the above model because we can recover them from the reduced form coefficients π_1, π_2, and π_3. The estimate of $\theta_1$can be used to derive an estimate of the scarcity premia on imports using the function $\mu_t^*(Z_t^*)$. However, since we are interested in the elasticity estimates, the parameters π_1 and π_2 are the relevant ones for our purpose.

(2) Empirical Analysis

The long run import demand relation derived in equation (11) implies that m_t, $\ln(Y_t - P_t M_t)$, $p_{t,}$ Z_t are cointegrated under the assumption that the random preference shocks b_t and c_t are strictly stationary. We adopt the following specifications for the preference shocks $b_t = b_0 + \varepsilon_{bt}$; $c_t = c_0 + \varepsilon_{ct}$,where ε_{bt} and ε_{ct} are mean zero (strictly) stationary processes. So the composite preference shock ξ_t can be rewritten as $\xi_t = \frac{1}{\eta}\left[(b_0 - c_0) + (\varepsilon_{bt} - \varepsilon_{ct})\right] \equiv \pi_0 + \varepsilon_t$. By using this equation and incorporating a dummy to capture the disruptions due to the civil war in 1983-89, we get the following estimable long run import demand function for Sri Lanka [9]:

$$m_t = \pi_0 + \pi_1 \ln(Y_t - P_t M_t) - \pi_2 p_t + \pi_3 Z_t^* + \pi_4 `civil' + \varepsilon_t \tag{12}$$

There are two central issues in the empirical analysis: (i) the validity of the cointegration or stationarity restriction embodied in the equation(12), (ii) estimation of the cointegrating vector if there exists adequate evidence in favor of one or more long run relation(s). To test for the existence and the number of cointegrating relation(s), the recent bounds tests approach proposed by Pesaran, Shin and Smith (2001) (both the 'F' test and the 't' test based on the cointegration test of Banerjee et. al (1998)) along with the widely used Johansen procedure for determination of the cointegration rank (i,e. the maximal eigenvalue and trace tests) are employed. For estimation of the cointegrating vector, we use two alternative approaches : (i) ARDL approach (Pesaran and Shin (1999)), (ii) DOLS (Stock and Watson, 1993). The choice of the estimation methods is motivated by strong evidence in favor of better small sample properties of ARDL and DOLS (for a discussion, see Caporale and Pittis (1999)). We also perform parameter stability tests.

(2.1)The Existence and Number of Cointegrating Relation(s)

The specification of the ARDL and VAR models (lag order and the deterministic part) for tests of cointegration was determined on the basis of the modified F test for autocorrelation along with the Schwartz Bayesian criterion (henceforth SBC). The modified F tests indicate the absence of serial correlation for all specifications of the deterministic part of the ARDL model for all lags above one. Indeed, when insignificant lagged terms are dropped, the serial correlation problem disappears even in the case of one lag. Thus, we perform the bounds tests for all different specifications of the deterministic part and for all

[9]Observe that the following estimating equation does not contain a time trend in the long run relation implying that only the deterministic cointegration is considered. This is motivated by both theory and evidence. First, the stationarity restriction implied by equation (11) is that of deterministic cointegration. Second, in our empirical analysis, the time trend when restricted to be in the cointegration space was found to result in implausible parameter estimates.

three lags (1-3). The global maximum of the SBC selects a VAR model with an intercept and no trend, and involving two lags. The Johansen's Trace and $\lambda_{\max}$ tests are performed for this specification of the VAR model.

For the bounds 'F'and 't' tests, we initially include the full set of lagged differenced variables. Then we omit the statistically insignificant lagged differenced variables which seems eminently desirable to minimize the problem of over-parameterization given the small sample size of the data.[10] Table 1 summarizes the results of alternative tests of the validity of the cointegrating relation specified in equation (12). The results of the bounds 'F' Tests (Table 1) show that the null hypothesis of no long run relation among the variables of the import model can be rejected at 1 *percent* significance at all three lags for all different formulations of the deterministic part. The results from the bounds t tests are similar though they seem to depend on the selected lag and the specification of the deterministic part. The bounds t tests at all three lags also indicate the existence of a long run relationship at 5 percent or lower significance level (see Table 1). The overall results of the bounds tests thus provide strong evidence in favor of a long run relation in the data for Sri Lanka.

Unlike the bounds tests procedure which obviates the need for unit roots pre-testing, Johansen's eigenvalue and trace tests are conditional on the order of integration of individual variables. Results of unit root pre-tests (DF/ADF) for all variables except Z_t^* show that they can be treated as $I(1)$.[11] The sample period for which Z_t^* has a value different from zero is too small to allow a proper unit root testing. The transformation of the data vector for Z_t which ensures separation between constrained and unconstrained regimes also introduces a lower bound to the value of Z_t^*. As Z_t^* decreases with time in our data, and is bounded below by zero, we treat it as an $I(0)$ variable.

The results of the bounds tests are supported by the Johansen's $\lambda_{\max}$ and Trace tests based on the VAR model selected by the SBC (with intercept but no trend and a lag length

[10]Following Pesaran et. al. (2000) we omit only those lagged differenced variables which are statistically insignificant in all estimated regressions.

[11]The unit root tests results are omitted for the sake of brevity.

of two). Both the $\lambda_{\max}$ and Trace tests indicate that there is only one cointegrating vector among the variables of the import model (see Table 2). The loading factors and their respective t values show that none of the $I(1)$ variables can be treated as weakly exogenous to the system. The multivariate diagnostics suggest that there are no serial correlation, heteroscedasticity and/or non- normality problems in the residuals. The same results are also supported by the univariate statistics of each equation in the VAR.

(2.2) Estimates of the Price and Income Elasticities

In this section, we present alternative estimates of the elasticity parameters using ARDL and DOLS methods.[12] As we noted earlier, the cointegarting restriction implied by our model is that of a deterministic cointegration and thus the estimating equation (12) does not contain a time trend. But if we rely on statistcial criteria for slection of the ARDL model, the trend is found to be significant. The results are, however, implausible as the income coefficient has a negative sign. So we estimate a model which includes an intercept but no trend.[13]

The estimated coefficients not only meet the theoretical sign restrictions but are also highly statistically significant with a P-value of 0.00 for both price and income elasticity coefficients. (Table 3). The ARDL estimates of income and relative price elasticities are: $\hat{\pi}_1 = 0.85$ and $-\hat{\pi}_2 = -0.78$ respectively.[14] The corresponding price elasticity estimate ($-\hat{\pi}_2 = -0.78$) from DOLS is identical to the ARDL estimate, while the income elasticity ($\hat{\pi}_1 = 0.75$) is smaller in magnitude. Both the ARDL and DOLS estimates of coefficient

[12]The specification of the ARDL model was chosen by AIC, as the specification slelected by SBC shows strong evidence of serial correlation.

[13]A conflict between the statistical and economic model selection criteria is, however, not alien to the literature. Croix and Urbain (1998), for example, find that, in a model of non-durable imports for France, the inclusion of a trend is dictated by the statistical significance, but the resulting estimates are not plausible. So they override the statistical evidence and exclude the trend as dictated by the economic plausibility of the estimates

[14]Note that we are concerned with the *notional* price and income elasticity parameters that correspond to an unconstrained regime, i.e., where $\mu_t^* = 0$. Thus the coefficients of income (π_1) and price ($-\pi_2$) give us the appropriate elasticity estimates. Also observe that the concept of *constrained* elasticity is not meaningful for an aggregate import function. The concept of constrained elasticity can, however, be useful in a disaggregate import demand model, as in Bertola and Faini (1990).

of scarcity premium variable, Z_t^*, have correct signs and are statistically highly significant, although the numerical magnitudes are slightly different: ($-\hat{\pi}_3 = -0.22$ (ARDL); $-\hat{\pi}_3 = -0.17$ (DOLS)). The statistical and economic significance of the coefficient of the civil war dummy, both in ARDL and DOLS, show that the disruptions during 1983-89 period had significant negative impact on Sri Lanka's imports.

A recurring concern in the policy analysis is the question of possible non-constancy of the estimated elasticity parameters due to structural breaks. We test for the structural stability of the estimated elasticity parameters from both ARDL and DOLS by using CUSUM and CUSUMSQ tests. For the ARDL estimates, neither of the tests show any evidence of instability in the estimated parameters at 5 *percent* significance level (see Figures 1a and 1b). In case of DOLS estimates, the CUSUM show that there are no evidence of any parameter instability (see Figure 2a), while, according to CUSUMSQ there are instabilities towards the end of the sample period (see Figure 2b). However, the parameter estimates become stable even in case of DOLS if the intercept is excluded from the regression.[15]

(2.3) Elasticity Estimates From Alternative Models

This section reports the results of the empirical analysis of (i) a modified traditional model which estimates equation (12) while ignoring the variable Z_t^*, and (ii) the foreign exchange availability formulation which uses the log of real foreign exchange availability as a regressor in equation (12) instead of Z_t^*. The general empirical strategy is the same as that followed above, but for the sake of brevity we do not report the results of tests of cointegration in tabular form.[16]

[15] Recall that the specification of the estimated ARDL model (Table 3) is selected by AIC, as the model selected by SBC shows evidence of serial correlation. However, if the model is selected by SBC, the intercept term does not belong to the regression. The resulting income and price elasticity estimates become slightly larger than those reported in Table 3. The estimates of income elasticity are: 0.95 (ARDL) and 0.91 (DOLS). The price elasticity estimates are : -0.85 (ARDL) and -0.93 (DOLS). The coefficient of the foreign exchange premium variable Z^* bears correct sign and is highly statistically significant irrespective of estimation techniques.

[16] The details can be obtained from the authors.

The Modified Traditional Model

For the modified traditional model, the bounds 'F' tests indicate the existence of a long run relation only at 10 percent significance level for the specification without trend or intercept at one and two lags, and with an intercept at three lags. For all other specifications of the deterministic part and lags, the evidence show the absence of a cointegrating relation. [17] The λ_{max} and Trace tests also support the conclusion that there is only very weak evidence, if any, in favor of a cointegrating relation. This evidence is indicative of the inadequacy of the traditional model for modelling aggregate imports of Sri Lanka which is confirmed by the anomalous estimates of the parameters reported in Table (4). When the ARDL estimator is used, the price coefficient has a positive sign and is statistically irrelevant (t value is 0.25).[18] The income coefficient has the correct sign, but it is statistically insignificant ($t = 0.86$). The DOLS estimates have right signs but both the price and income elasticity estimates are statistically insignificant and implausibly small in magnitude. Also, the magnitude of the estimates are extremely sensitive to the estimation method used (see Table 4).[19] The results clearly demonstrate that the traditional model is ill-suited for estimating the elasticity parameters in case of Sri Lanka.[20]

Foreign Exchange Availability Formulation

The foreign exchange availability (FAV) formulation replaces Z_t^* in equation (12) by a variable measuring total foreign exchange availability. In contrast to the traditional

[17] According to bounds 't' tests, there is no long run relationship in the modified traditional model.

[18] In table 4, we reported the estimates which we judged to be best in terms of the signs and magnitudes.

[19] Such dramatic change in the magnitude of a parameter across ARDL and DOLS estimates is observed in case of India by Caporale and Chui (*op cit*). The price elasticity estimates for aggregate imports of India from the traditional model obtained by them are: $-0.03(0.35)$ (DOLS) and $-1.01(0.40)$ (ARDL). Also, contrast this fragility of the estimates with the estimates from our preferred model where ARDL and DOLS give identical estimates for the price elasticity parameter.

[20] The inadequacy of the traditional model in estimating import elasticities is evident even if one includes a dummy to capture the shift in policy regime in 1977. When the traditional model is modified to include such a policy shift dummy, both the price and income elasticity estimates bear correct signs and become statistically significant. However, magnitudes of the parameter estimates still remain implausible. For instance, the estimates of the income coefficient (0.45 (DOLS) and 0.62 (ARDL)) are rather low, particularly compared with those obtained from the structural model presented in the preceding section.

model, the evidence from both bounds F and t tests, and the $\lambda_{\max}$ and Trace tests show that there is a single cointegrating vector in the foreign exchange availability model. The ARDL estimates of relative price and income coefficients have wrong signs, and they are statistically insignificant (Table 5). [21] In contrast, the DOLS estimates of both income and price coefficients have correct signs, and the income coefficient is also statistically significant. But, similar to the case of the traditional model, both the price and income elasticity estimates are implausibly low ($\hat{\pi}_1 = 0.15$ and $-\hat{\pi}_2 = -0.01$). The coefficient of foreign exchange availability is highly statistically significant with correct positive sign according to both ARDL and FIML estimates. The point estimate in the case of ARDL is almost equal to unity (1.01) which clearly shows that the strength of the *near identity* problem is not diluted by the addition of the standard price and income variables. The DOLS estimate of the coefficient of foreign exchange availability is, however, much smaller (0.71).[22] But in both cases, it is clear that the foreign exchange availability dominates the price and income variables in explaining the variations in imports, and that the estimates of the elasticity parameters from this model are unacceptable, both on economic and statistical grounds.

(2.4) Comparison With Other Available Elasticity Estimates

In this sub-section, we compare and contrast the estimated price and income elasticities from our preferred model with the other estimates available in the literature, and also discuss the implications of the estimated parameters for intertemporal elasticity of substitution. Observe that the income variable in our model is GDP minus exports and thus the income elasticity estimate is, in strict sense, not comparable to other estimates in the

[21] This result is quite robust, as it holds true in all different formulations of the deterministic part.

[22] If we introduce a dummy in the FAV formulation to capture the policy regime shift since 1977, the estimates of the income and price elasticities confirm the theoretical sign restrictions in both ARDL and DOLS models, but the price elasticity is statistically insignificant, and both income and price elasticities are implausibly small in magnitudes (income elasticitiy :0.2 (ARDL), and 0.19 (DOLS), and price elasticity: -0.23 (ARDL) and -0.18 (DOLS)). According to both ARDL and DOLS estimates, foreign exchange availability continues to explain most of the variations in imports. This finding highlights the inadequacy of using a dummy to account for policy shifts in the presence of foreign exchange constraint.

literature where GDP is used as the income variable. We can, however, derive an estimate of elasticity of aggregate imports with respect to GDP from our model. The following formula gives us the elasticity of aggregate imports with respect to GDP:

$$E_{GDP_t} = \pi_1 \frac{GDP_t}{(GDP_t - P_t^X X_t)} \tag{13}$$

Where E_{GDP_t} is the elasticity of aggregate imports with respect to GDP at time period t and $P_t^X X_t$ is the export earnings denominated in terms of home goods. Table 6 summarizes the available price and income elasticity estimates for aggregate imports of Sri Lanka. Unfortunately, there are only a few estimates of price and income elasticities of aggregate imports of Sri Lanka are available in the literature.

As the share of export in GDP varies from year to year, the estimates of income elasticity with respect to GDP also vary. The range of income elasticity, estimated from our model using DOLS and ARDL, is [0.87, 1.23] [Table 6]. The mean of income elasticity estimates with GDP as the scale variable is: 0.96 (DOLS) and 1.09 (ARDL). The price elasticity estimates are identical at -0.78 regardless of the estimation techinique used (ARDL or DOLS). Note that the estimated price elasticity is much higher compared to the available estimates (about three times the estimate of -0.30 reported by Reinhart (*1995*) and nearly twice as large as the estimate of -0.48 reported by Sinha (*op cit*)) (Table 6). The muted price response found in these studies is probably due to the fact that no account was taken for the subsample of period with trade and exchange interventions. Also, in contrast to the negative sign of the income elasticity (-0.39) reported by Sinha (*op cit*), the income elasticity has the expected positive sign in both estimation techniques used. The magnitude of income elasticity is smaller, nearly half of the estimate of 1.98 reported by Reinhart. Our estimate is, however, close to the conventional wisdom of a long run unitary income elasticity of aggregate imports.

As is well-known, the inverse of the parameters of the addilog utility function can be interpreted as measures of intertemporal elasticity of substitution. Our results in this regard are interesting in that they imply very different relative magnitudes compared to the available evidence on developed countries. For example, the available estimates for USA show that the intertemporal elasticity of substitution is two (Clarida(1994)) to three (Ceglowski, *op cit*, and Croix and Urbain, *op cit*) times higher for imports when compared to that of home goods consumption. In contrast, our estimates for Sri Lanka suggest that the magnitudes of intertemporal elasticity of substitution are only slightly higher for home goods consumption compared with imports (for imports, the estimates are 0.78 (both ARDL and DOLS); for home goods: 0.92 (ARDL) and 1.04 (DOLS). Further work is needed to see if this is true for other developing countries as well.

Conclusions

Sri Lankan economy had been characterized by pervasive trade and exchange rate interventions during the 1960's and 1970's. The protectionist policies were, however, almost completely dismantled in 1977-78, much earlier than most of the developing countries. The shift in policy implies that the time series data on imports cover periods of both constrained and unconstrained trade and exhange regimes. In this paper, we show that the estimates of critical import elasticity parameters may be implausible and significantly biased if one does not take proper account of the changing policy regime. The traditional import model which treats the constrained regime as if there is no foreign exchange constraint produces theoretically inconsistent estimates, with wrong signs and implausible magnitudes. The foreign exchange availability formulation fares no better because it treats the unconstrained sub-sample as if it is also constrained. In contrast to these two benchmarks, the estimates from our model not only satisfy the theoretical sign restrictions but are also economically and statistically highly significant. The results show that while the conventional wisdom of a unitary income elasticity might be almost right on the mark, the price elasticity is less than a half of the estimate of -2.0 used in Bhalla(1991) for the estimation of equilibrium exchange rate. The assumption of a price elasticity of -2.0 in case of Sri Lanka is likely to have introduced a substantial downward bias in the estimate of the equilibrium exchange rate given the rather robust evidence that the elasticity is in the neighborhood of -0.80.

The more general lesson from this exercise is that an appropriate treatment of the policy regime in a country is of paramount importance for reliable estimates of the price and income elasticities of import demand. This brings into focus the need for more in-depth country studies as invaluable tools for analysis and formulation of policy.

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Appendix 1: Trade and Exchange Rate Policy in Sri Lanka

Starting from the late 1950s, Sri Lanka pursued increasingly interventionist policies of import and exchange controls in order to cope with balance of payments difficulties. By 1965, quantitative restrictions and bans on imports, foreign exchange controls and restrictions on capital movements virtually insulated Sri Lankan economy from rest of the world (Cuthbertson and Athukorala(1989)). In November 1977, the new United National Party(UNP) government replaced all quantitative restrictions with tariffs, revised tariff structure to achieve greater uniformity and removed most restrictions on foreign capital movements. The dual exchange rates system was abolished and the new unified rate was placed under a managed float. The trade account transactions have been free from any restrictions since November 1977. The first wave of reforms ($1977-82$) was followed by a second wave ($1990-1995$) in which tariff structure was further simplified, rates were reduced and the remaining restrictions on the current account transactions were removed (Dunham and Kelegama (1997)). Sri Lanka accepted the IMF Article VIII obligations in March, 1994. Sri Lanka also dealt with a devastating civil war during $1983 - 89$ period which not only stalled the first wave of reforms but also inflicted enormous human and economic costs.

Appendix 2: Data Source and Definition of Variables

The data used in the empirical analysis were taken from *Central Bank of Sri Lanka Annual Reports*, *International Development Statistics* (OECD) and *International Financial Statistics* (IMF). Annual data for the sample period 1960 to 1995 were used for empirical analysis. Microfit and CATS programs are used for the econometric analysis.

Definitions:

M = Import payments in domestic currency (Rupee) deflated by import price index in Rupee.

H = Gross Domestic Product (GDP) minus export payments deflated by consumer price index (CPI).

P =import price index divided by CPI.

F = Foreign exchange available which consists of export earnings, remittances and total foreign aid disbursement and beginning of the period foreign exchange reserves, deflated by CPI.

f = foreign exchange availability(F) divided by import price index.

Z = [*real* expenditure((GDP+imports-exports)/cpi)]/Foreign exchange availability(F)).

$D = 1$ for the sample period 1960-1977 and zero otherwise.

$Z^* = Z * D$.

Table 1: Bound Tests for Long-run Relationship in an ARDL model: Sri Lanka

Lags			Deterministic part No Intercept No Trend	Restricted Intercept	Unrestricted Intercept	Restricted Trend
1	Bound test	F-statistic	6.99*	5.86*	7.3*	6.56*
	Bound test	t-statistic	-3.88**	-3.95**	-3.95**	-4.21**
	intercept	t-statistic	-	1.09	1.09	1.69
	trend	t-statistic	-	-	-	1.5
	Civil War	t-statistic	-3.61	-3.65	-3.65	-1.24
2	Bound test	F-statistic	7.24*	5.91*	7.1*	9.82*
	Bound test	t-statistic	-4.08*	-4.08**	-4.08**	-5.65*
	intercept	t-statistic	-	0.9	0.9	3.33
	trend	t-statistic	-	-	-	3.2
	Civil War	t-statistic	-3.7	-3.65	-3.65	-0.43
3	Bound test	F-statistic	6.93*	5.62*	6.37*	8.99*
	Bound test	t-statistic	-4.07*	-3.96**	-3.96**	-5.43*
	intercept	t-statistic	-	0.84	0.84	3.24
	trend	t-statistic	-	-	-	3.12
	Civil War	t-statistic	-3.87	-3.68	-3.68	-0.56

Note: Critical values for Bound tests (both F and t-tests) are taken from Pesaran et al (2001)
'Civil War' is a dummy for the civil war years (1983-89)
* : significant at 1 percent level
** : significant at 5 percent level
*** : significant at 10 percent level

Table 2: Tests for Existence of Cointegrating Vectors and Weak Exogeneity using Johansen's Approach, Sri Lanka

Full System

Eigen Values	Null Hypothesis	Lmax	Trace	90% Critical Values[1] Lmax	90% Critical Values[1] Trace
0.81	r=0	57.33	66.21	22.85	34.46
0.21	r<=1	7.98	8.88	15.59	18.86
Loading Factors					
	Coefficient	t-value			
Δm	-0.11	-1.88			
Δh	0.11	4.57			
Δp	-0.34	-5.43			
Residual analysis					
	Statistics	p-value			
LM(1)	5.25	0.81			
Normality	12.1	0.06			
Equation	ARCH(1)	Normality	R^2		
m	0.74	2.62	0.45		
h	1.93	3.03	0.54		
p	0.46	1.82	0.55		

Note: m=log(total imports)
h= log(home good consumption)
p=log(import price index/consumer price index)

1/: 90% critical values are adjusted for sample size by using Response Surface Regressions of Cheung and Lai(1993)

Table 3: Estimates of Long-run Relationships, Sri Lanka

	ARDL model		DOLS Model	
	Coefficient	t-value	Coefficient	t-value
h	0.85	4.1	0.75	6.01
p	-0.78	-4.14	-0.78	-6.98
Deterministic part	Coefficient	t-value		
Z^*	-0.22	-5.04	-0.17	-11.19
Civil War Dummy	-0.43	-2.53	-0.19	-4.83
Intercept	0.78	0.5	1.26	1.32
Speed of Adjustment	-0.38	-2.85		
Residual analysis for ARDL model				
	χ^2	p-value	χ^2	p-value
Serial correlation (F)	1.18	0.29	0.01	0.93
Normality	0.2	0.91	0.02	0.99

Note: m=log(total imports)
h= log(home good consumption)
p=log(import price index/consumer price index)
Z^* =(real domestic expenditure/real foreign exchange availability(f))*D
D takes a value of 1 for 1960-1977 and zero otherwise.

Table 4: Estimation of Long-run Relationship in Modified Traditional Model

	ARDL model		DOLS Model	
	Coefficient	t-value	Coefficient	t-value
h	0.83	0.86	0.29	0.83
p	0.26	0.25	-0.03	-0.12
Deterministic part	Coefficient	t-value		
Civil War Dummy	-0.88	-0.73	-0.1	-2.33
Intercept	1.7	0.24	4.87	1.86
Speed of Adjustment	-0.1	-0.88		
Residual analysis for ARDL Model				
	χ^2	p-value	χ^2	p-value
Serial correlation (F)	0.13	0.72	4.48	0.05
Normality	2.02	0.36	1.09	0.58

Note: m=log(total imports)
h= log(home good consumption)
p=log(import price index/consumer price index)

Table 5: Estimation of Long-run Relationship in Foreign Exchange Availability Model

	ARDL model		DOLS Model	
	Coefficient	t-value	Coefficient	t-value
h	-0.05	-0.27	0.15	3.51
p	0.23	1.45	-0.01	-0.27
f	1.01	5.94	0.71	10.9
Deterministic part	Coefficient	t-value		
Civil War Dummy	-0.17	-1.57	-0.08	-2.39
Intercept	0.36	0.29	0.81	1.14
Speed of Adjustment	-0.43	-3.66		
Residual analysis for ARDL Model				
	χ^2	p-value	χ^2	p-value
Serial correlation (F)	1.03	0.32	0.27	0.61
Normality	0.02	0.99	0.54	0.77

Note: m=log(total imports)
h= log(home good consumption)
p=log(import price index/consumer price index)
f=log(real foreign exchange availability)

Table 6: Comparison of Elasticity Estimates

	Elasticity Estimates		
	This paper	Reinhart (1995)	Sinha(1999)
Income*			
Average	0.96-1.09	1.98	-0.39
Minimum	0.87-0.98	-	-
Maximum	1.09-1.23	-	-
Price	-0.78	-0.3	-0.48

Note:*: Income elasticity is defined with respect to GDP by dividing elasticity estimates (with respect to expenditure on home goods consumption) in Table 3 by (1-share of export in GDP) (see formula in equation (13) in the text).

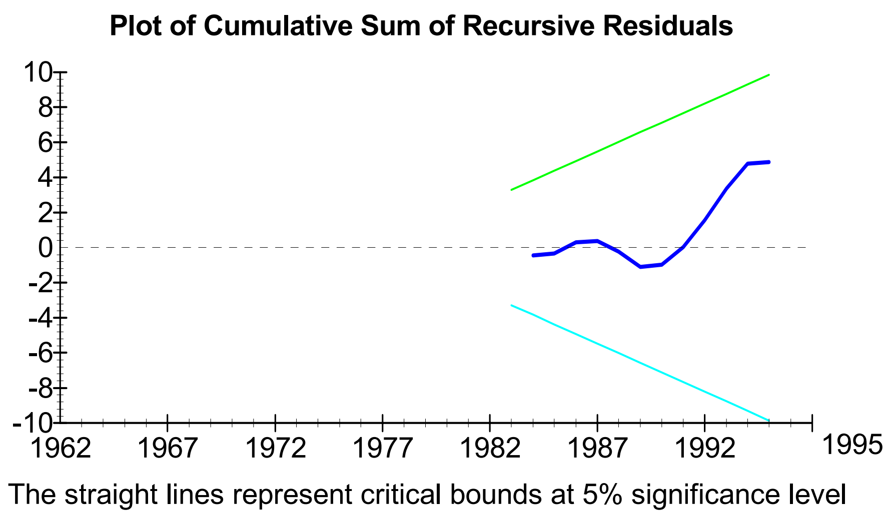

Figure 1a: Cusum tests (ARDL)

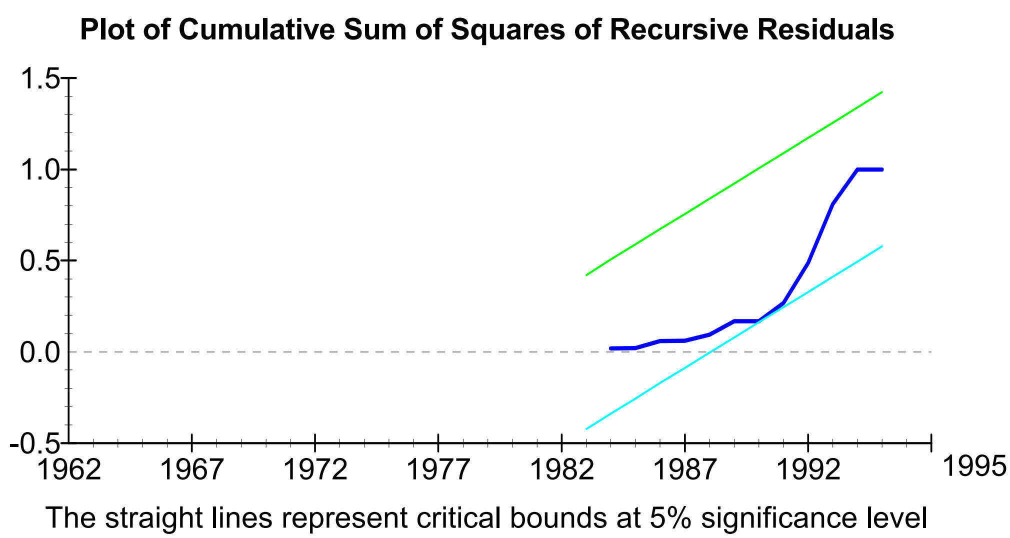

Figure 1b:Cusum Square tests (ARDL)

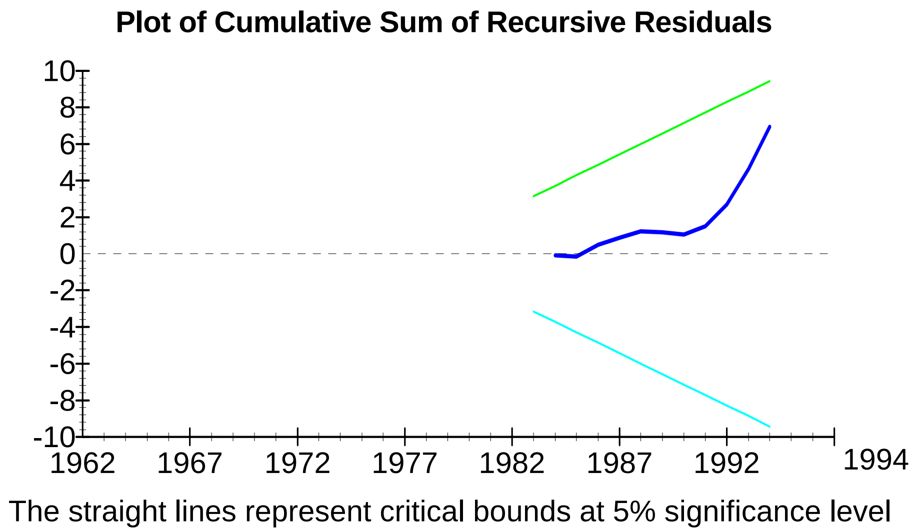

Figure 2a: Cusum tests(DOLS)

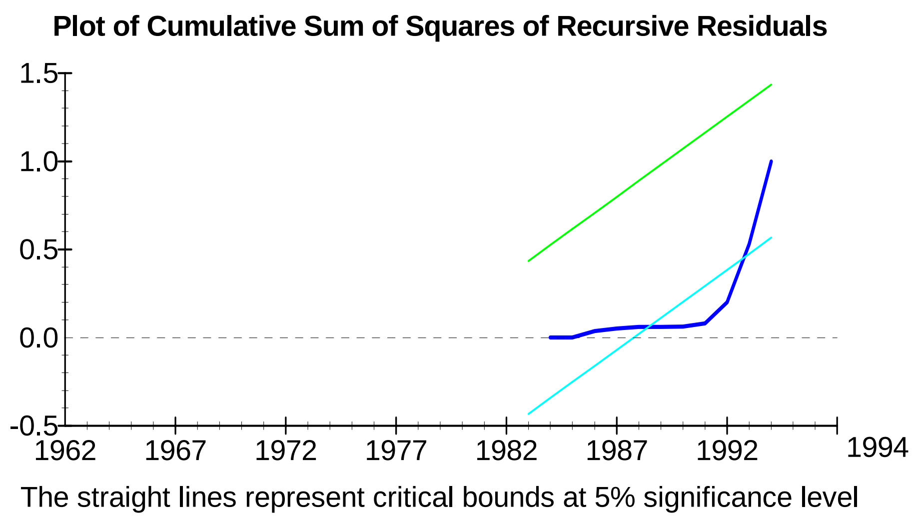

Figure 2b: Cusum Square tests (DOLS)