Institute for International Economic Policy Working Paper Series
Elliott School of International Affairs
The George Washington University

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IIEP-WP-2008-10

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Estimating Import Demand Function in Developing Countries : A Structural Econometric Approach with Applications to India and Sri Lanka[1]

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Abstract

Due to the unavailability of time series data on domestic market clearing price of imports, the estimation of *notional* price and income elasticities of aggregate import demand remains a daunting task for a large number of developing countries. This paper develops a structural econometric model of a two goods representative agent economy that incorporates a binding foreign exchange constraint at the administered prices of imports. A theoretically consistent parameterization of the 'virtual relative price' of imports circumvents the data problem, and thus enables the estimation of income and price responses by cointegration approach. The price and income elasticity estimates for India and Sri Lanka, in contrast to the extant literature, have correct signs, high statistical significance, and plausible magnitudes.

Keywords: Import Demand, Foreign Exchange Rationing, Virtual price, India, Sri Lanka, Cointegration.
JEL Classification : F14; O16

[1] We would like to thank Arvind Panagariya, John Williamson, Imam Alam, Yasuyuki Sawada, and seminar participants at Stanford University for comments on earlier drafts of the paper. The standard disclaimers apply.
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Introduction

The econometric estimation of the price and income elasticity of imports has been the subject of a large literature both for developed and developing countries (see, for example, Malley and Moutos (2002), Caporale and Chui (1999), Hooper et. al. (1998), Ghei and Pritchett (1999), Faini, Pritchett and Clavijo (1992), Winters (1987), and Goldstein and Khan (1985)). Reliable estimates of the elasticity parameters are important for informed policy analysis in a number of areas, such as exchange rate policy, fiscal implications of tariff reductions under trade liberalization programs, and calculation of optimal taxes. In the context of developing countries, econometric modeling of import functions has, however, been constrained by the fact that the time series data available for most of the developing countries span periods of pervasive trade and exchange rate restrictions. To be sure, the trade and exchange rate interventions would not have created any problem for the estimation, if the right kind of data were available; most importantly, the data on the market clearing price (*virtual price* a la Neary and Roberts (1980)) of imports (administered price plus scarcity premium). In the presence of extensive secondary markets for import licenses and imported goods, the secondary market prices are the appropriate prices for imports relevant for consumer optimization. Unfortunately, for most of the developing countries, such price data are not available. This paper is concerned with modeling aggregate imports in developing countries saddled with such data problems.[2] We present a structural econometric model of a two goods representative agent economy that circumvents the data problem by parameterizing the Lagrange multiplier of a binding foreign exchange constraint at the administered prices of imports.

Although the problem of unavailability of appropriate price data is well-known and has been widely discussed, it has not yet been satisfactorily addressed, to our knowledge, especially in the context of an aggregate import demand function (see the discussion in Ghei and Pritchett (1999)). Some important progress have been made in the estimation of disaggregate import demand under foreign exchange constraint or quantitative restrictions that use the Neary-Roberts

[2] This is the "market clearing model" to use the phrase of Winters and Brenton (1993) where the secondary market clears at the equilibrium price. The model developed in this paper remains equally valid if instead the "rationing model" (in their terminology again) is applicable where the secondary market is thin or non-existent, and the appropriate scarcity prices are the "virtual prices" á la Neary and Roberts (1980), as long as the representative agent assumption is entertained.

(1980) framework where at least some of the imported goods are not constrained (see, Bertola and Faini (1990), and Winters and Brenton (1993)). In contrast, the standard model with income and relative price has been, and still is, the work-horse for modeling the aggregate import demand in developing countries. Some researchers add 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 Moran (1989)). The inadequacies of a standard demand model are obvious from the anomalous results often found in estimating the effects of relative price and income. The frequently reported 'wrong' signs of the price and income elasticities, and both economically and statistically insignificant price coefficient estimates come as no surprise from this perspective. Another way the inadequacy of the traditional model might manifest itself is in the absence of a long run relation among the variables as is the case in a number of studies.[3] 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. The estimated equation is close to an identity (*near identity* in the terminology of Emran and Shilpi (1996))[4]; the coefficients of price and income are devoid of any behavioral interpretations, and might yield nonsensical results.[5] To illustrate the force of the *near identity* problem, we report, for India, the results of the OLS regression where imports were regressed on foreign exchange availability and a constant. The coefficient of foreign exchange availability is 1.03 with a 't' value of 26.37 and $\bar{R}^2 = 0.94$. The restriction that the coefficient is equal to one can not be rejected by the Wald test with a P-value of 0.46.[6] The above results clearly demonstrate the one to one relation between imports and foreign exchange availability and are indicative of the pitfalls of using foreign exchange availability approach.

The model developed in this paper avoids the pitfall of *near identity* by parameterizing the Lagrange multiplier associated with the binding foreign exchange constraint in terms of the *ratio*

[3] Sinha (1999), for example, does not find any cointegrating relation for the traditional model in case of India.

[4] To be more precise, what is being estimated here is the foreign exchange budget constraint.

[5] For example, Mazeri (*op cit*) finds in case of Iran that the estimated price coefficient is **zero** when the oil revenue is used as a measure of foreign exchange availability. In case of Bangladesh, Emran and Shilpi (1996) find that the sign of the estimated price coefficient is **positive** when foreign exchange availability is defined to be equal to export earnings plus remittances plus disbursed foreign aid.

[6] Results for Sri Lanka are similar.

of income to foreign exchange resources available to a country. As discussed later in the paper, this parameterization is both intuitive and well grounded in the theory. This approach to modeling the effects of a binding constraint by parameterizing the associated Lagrange multiplier has an honorable pedigree in economic literature. Two of the areas where it has proven especially fruitful are the econometric modeling of (*i*) investments under credit constraints (see, for example, Hubbard and Kashyp (1992)), (*ii*) consumption under liquidity constraints (see, for example, Zeldes (1989)). We apply the model developed here to analyze the behavior of aggregate imports of India and Sri Lanka. The choice of India is particularly appropriate as an application given the extensive government interventions in trade and exchange rate throughout most of its existence as a nation state. Sri Lanka is also an interesting case study, given that the trade liberalization was implemented in 1977, much earlier than in India, and as a result the sample period characterized by an effectively free trade and exchange rate regime is much longer.[7] We compare and contrast the results of our model with those of a modified traditional model and the foreign exchange availability formulation. The empirical results from three alternative estimators of a cointegrating vector (ARDL, DOLS, and FM-AADL) clearly demonstrate the advantages of the model presented in this paper, both on theoretical and statistical grounds.

The rest of the paper is organized as follows. The first section presents a simple intertemporal optimization model of a representative consumer to derive the aggregate import demand function under trade and exchange rate restrictions. Section 2, arranged in a number of sub-sections, presents the empirical implementation of the model using annual time series data from India (1952-99) and Sri Lanka (1960-95). The sub-section 2.1 presents the estimation results of the cointegration vector for the structural import model derived in the first section. The next sub-section presents the empirical analysis of the two other competing import models extant in the literature. The sub-section 2.3 reports the estimates of the price and income elasticities during the sample period. The paper ends with some concluding remarks.

[7] For excellent discussions of the evolution of India's trade and exchange rate policies, see Bhagwati and Srinivasan (1993) and Panagariya(1999). For Sri Lanka see Cuthbertson and Athukorala (1989).

1. A Model of Aggregate Imports Under Binding Foreign Exchange Constraint

Following the analysis of Ceglowski (1991) and Clarida (1994) we use a rational expectations permanent income model of a representative agent to derive the import demand function.[8] The distinguishing feature of the model presented here is that it incorporates a binding foreign exchange constraint and thus abandons the implicit assumption of a perfect international capital market.[9] 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 describing the foreign exchange availability constraint.[10] Let P_t denote the relative price of imports at administered exchange rate; A_t, assets; $\tilde{Y}_t$, labor income; F_t, 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

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

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

[8] An early attempt to model aggregate import behavior of developing countries within an explicit intertemporal framework is that by Winters (1987). The model developed by Winters uses a utility function where imports are separable from home goods. It focuses on the intertemporal substitution of imports and there is no contemporaneous substitution (i.e. relative price effect) because home goods are not an argument in the sub-utility function.

[9] The access to international capital market is important when overvaluation of the administered exchange rate implies an unmanageable trade deficit, which, we think, is the more empirically relevant case. It is possible that a country runs a trade surplus at a overvalued exchange rate, even though the surplus is smaller than it would be at the equilibrium exchange rate. Thanks to John Williamson for pointing out this to us.

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

where a dot on a 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.[11] The current value Hamiltonian of the optimization problem can be written as:

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

where λ_t is the costate variable and μ_t is the Lagrange multiplier associated with the foreign exchange constraint. The costate variable λ_t can be interpreted as the marginal utility of wealth. The first order conditions for this optimization problem are:[12]

$$U_H = \lambda_t \tag{3}$$

$$U_M = P_t(\lambda_t + \mu_t) \tag{4}$$

$$\dot{\lambda} = (\delta - r)\lambda_t \tag{5}$$

$$[F_t - P_t M_t] \geq 0;\ \mu_t[F_t - P_t M_t] = 0 \tag{6}$$

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 can be rewritten as:

$$C_t H_t^{-\alpha} = \lambda_t \tag{7}$$

$$B_t M_t^{-\eta} = P_t \lambda_t (1 + \mu_t^*) = \lambda_t P_t^* \tag{8}$$

[11]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 exogeneous. Obviously this is a simplification 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 endogeneous, 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 correct for the endogeneity of the regressors.

[12]For simplicity, the non-negativity conditions are not explicitly considered.

where $\mu_t^* = \frac{\mu_t}{\lambda_t} = \frac{\mu_t}{U_H}$ is the scarcity premia, 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 (7) to eliminate λ_t from equation (8) and take logarithm to get the following equation:

$$b_t - \eta m_t = c_t + p_t - \alpha h_t + \ln(1 + \mu_t^*) \tag{9}$$

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

In order to derive the long -run import demand relationship, we impose the steady state conditions that $\dot{A} = \dot{\lambda} = 0$. Also, the steady state is characterized by the equilibrium price relations implying $P_t = P_t^*$. The corresponding total household income including both labor and asset income evaluated at the equilibrium price vector is denoted by Y_t^*. The steady state solution implies that:

$$Y^* = H + P^* M \tag{10}$$

Using the steady state condition and taking logarithm, we get the following expression for h_t

$$\begin{aligned} h_t &= \ln(Y_t^* - P_t^* M_t) \\ &\equiv \ln(Y_t - P_t M_t) \end{aligned} \tag{11}$$

where $Y_t = (Y_t^* - \mu_t^* P_t M_t)$ is the observed income in a foreign exchange constrained regime and P_t the observed price. Now use equation (11) to eliminate h_t from equation (9) and solve for m_t:

$$m_t = \frac{\alpha}{\eta} \ln(Y_t - P_t M_t) - \frac{1}{\eta} p_t - \frac{1}{\eta} \ln(1 + \mu_t^*) + \xi_t \tag{12}$$

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 (12) provides an import demand function which is close to the standard double-log specification estimated by numerous studies for both developed and developing countries (see the surveys by Goldstein and Khan, (1985), Faini et. al., (1992), and Ghei and Pritchett(1999)). 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 (12) can thus be defined as GDP minus exports.[13] 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 (12) 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*. To avoid this problem, we parameterize μ_t^* by the ratio of total domestic expenditure (GDP+import-export) 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 total domestic 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. In Appendix 1, we show formally that the scarcity premium on imports is a positive function of Z_t. Since $\frac{\partial \mu_t^*}{\partial Z_t} > 0$, it immediately follows that import demand will vary negatively with Z_t (assuming imports are not Giffen goods):

$$\frac{\partial M_t}{\partial Z_t} = \frac{\partial M_t}{\partial \mu_t^*} \frac{\partial \mu_t^*}{\partial Z_t} < 0$$

For empirical implementation, we use the following functional form of $\mu_t^*(Z_t)$:

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

[13]Note that GDP is expenditure on domestically produced goods including exports and thus $H_t = (Y_t - P_t M_t)$ can be defined as GDP minus export.

Both India and Sri Lanka have liberalized their trade and exchange rate regimes (India in 1991 and Sri Lanka in 1977) and the scarcity premia μ_t^* thus should be approximately zero for the sample periods after liberalization. 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 the foreign exchange constrained period (1952-1991 for India and 1960-1977 for Sri Lanka) and zero afterwards. This transformed variable is denoted as Z_t^*. The imposition of such *a priori* theoretical restrictions by transforming the data series 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). With this specification, we have the following structural import demand function that can be estimated with the data available in most of the developing countries:

$$\begin{aligned} m_t &= \frac{\alpha}{\eta}\ln(Y_t - P_t M_t) - \frac{1}{\eta}p_t - \frac{\theta_1}{\eta}Z_t^* + \xi_t \\ &= \pi_1 \ln(Y_t - P_t M_t) + \pi_2 p_t + \pi_3 Z_t^* + \xi_t \end{aligned} \tag{13}$$

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 reduced form parameters, according to the model, should satisfy the following sign restrictions: $\pi_1 > 0$, $\pi_2 < 0$, and $\pi_3 < 0$.

2. Empirical Analysis

The long run import demand equation derived in equation (13) 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.[14] We adopt the following specifications for the preference shocks b_t and c_t: $b_t = b_0 + \varepsilon_{bt}$; $c_t = c_0 + \varepsilon_{ct}$, where ε_{bt} and ε_{ct} are mean zero (strictly) stationary processes. The composite preference shock ξ_t can be rewritten as $\xi_t = \frac{1}{\eta}\left[(b_0 - c_0) + (\varepsilon_{bt} - \varepsilon_{ct})\right] \equiv \pi_0 + \varepsilon_t$. Combining this with equation (13) we get the final estimating equation for the long run import demand function:

[14] For recent contributions that use cointegration approach to estimate long run import elasticity, see, for example, Clarida(1994), Urbain(1992), Caporale and Chui (1999).

$$m_t = \pi_0 + \pi_1 \ln(Y_t - P_t M_t) + \pi_2 p_t + \pi_3 Z_t^* + \varepsilon_t \quad (14)$$

Equation (14), which forms the basis of our empirical analysis, is estimated for India using annual time series data for the period 1952-99, and for Sri Lanka using similar data for the period 1960-95. There are two central issues in the empirical analysis: (*i*) the validity of the cointegration or stationarity restriction embodied in equation (14), (*ii*) estimation of the cointegrating vector(s). To test for the existence and the number of long run relation(s), the bounds 'F' test proposed by Pesaran, Shin and Smith (2001) and the bounds 't' test based on the cointegration test of Banerjee et. al (1998)) along with the widely used Johansen approach to the determination of the cointegration rank (i.e., the maximal eigenvalue and trace tests) are employed.[15] The bounds testing approach has the advantage that the existence of a long-run relationship among a set of variables can be tested without any prior knowledge about the order of integration of the individual variables. This avoids the much discussed problems associated with the unit-roots pre-testing (for a discussion, see Maddala and Kim (1998)). Moreover, The bounds test remains valid for testing the existence of a long-run relationship under fractional integration and near unit root processes (Pesaran and Pesaran,1997).[16]

For estimation of the cointegrating vector, we use three alternative estimators: (*i*) ARDL (Pesaran and Shin (1999), and (*ii*) DOLS (Stock and Watson (1993)) and FM-AADL (Caporale and Pittis (2004)). We use alternative methods to gauge the sensitivity of the results with respect to different estimation techniques. The recent evidence shows that the ARDL and FM-AADL estimators have desirable small sample properties, and they effectively correct for potential endogeneity of the explanatory variables (see Pesaran and Shin, 1999, and Caporale and Pittis, 1999, 2004).[17] Since the ARDL approach is valid in the presence of both $I(0)$ and

[15]We employ both single equation and system-based approaches to test for the existence of a cointegrating relation to ensure the robustness of the conclusions. In a recent paper, Gregory et. al. (2002) show that the single equation and system-based approaches may yield conflicting results, even in large samples. They find that the correlation among the P-values of different single equation and system-based tests for cointegration is very low.

[16]Pesaran, Shin and Smith (2001) analyze the asymptotic power of the bounds test under a sequence of local alternatives. The distribution corresponding to the near unit root process is based on Ornstein-Uhlenbeck process. The results show that the Bounds test performs reasonably well in these cases.

[17]We include the estimates from DOLS, as it is among the most widely used estimators of a cointegrating vector in the applied literature. However, according to the recent Monte carlo evidence, the DOLS lacks desirable small sample properties (see, for example, Caporale and Pittis (2004)).

$I(1)$ processes, it suggests that it should perform well with near unit root processes. Also, as noted by Elliott (1998), one can conduct asymptotic normal inference on the long run effects in the ARDL model in the presence of near integrated processes. For ARDL approach, we use the two-step procedure suggested by Pesaran and Shin (1999) where the specification of the ARDL model is chosen by Schwartz Bayesian criterion (henceforth SBC) and then estimated by OLS. The Monte-Carlo evidence of Pesaran and Shin (1999) shows that this two-step procedure effectively corrects for endogeneity of the explanatory variables, and the estimates exhibit good small sample properties.[18] The more recent Monte Carlo studies provide strong evidence in favor of the above conclusions (see, for example, Panopoulou and Pittis (2004), Caporale and Pittis (2004)). Panopoulou and Pittis (2004) conclude that the ARDL estimator performs best among a set of widely used estimators of a cointegrating vector including FMOLS, FMGLS, DOLS, and Johansen MLE, both in terms of "estimation precision and the reliability of statistical inference" (P. 585).

The FM-AADL estimator, proposed by Caporale and Pittis (2004), is a hybrid estimator that combines ARDL and FM-OLS. It involves a two-step procedure. In the first step, the ARDL approach is used to estimate the coefficients of stationary variables including the differenced terms in the import function. These estimated coefficients are then used to net out the effect of stationary explanatory variables from aggregate imports. In the second stage, the Phillips-Hansen FM-OLS is applied to the non-stationary variables with newly defined aggregate imports (net of the effects of stationary variables) as the dependent variable. This hybrid estimator thus uses both parametric (by ARDL) and semi-parametric (by FM-OLS) approaches to take care of the second order asymptotic bias arising from serial correlation and endogeneity (for discussion, see Caporale and Pittis (2004), Pesaran and Shin (1999) and Phillips and Hansen (1990)). According to the Monte Carlo evidence due to Caporale and Pittis (2004), this hybrid estimator has the most desirable small sample properties in a set of 28 estimators of a cointegrating vector. Caporale and Pittis (2004) show that the standard asymptotic critical values are highly misleading in small to moderate samples for widely used estimators including OLS and DOLS, but the ARDL and

[18]If the ARDL model is chosen by AIC instead, the estimates lack these desirable properties. This is because while SBC is a consistent model selection criterion, AIC is not (for a discussion, see Pesaran and Shin, 1999). Also, according to the Monte Carlo evidence presented by Panopoulou and Pittis (2004), the standard information criteria like SBC and AIC select the correct lag order reliably in the ARDL model.

FM-AADL do not suffer from such problems. [19] This implies that while we can use the standard asymptotic critical values for inference in the ARDL and FM-AADL, it is not appropriate for the DOLS.

(2.1) Estimates of the Long Run Import Model

As a first step to estimating the import model, the existence of a long-run import demand relationship is tested by performing the bounds tests suggested by Persaran and Shin (1999), Banerjee et. al (1998) and the rank tests for cointegration due to Johansen (1995). The specifications of the ARDL and VAR models (lag order and deterministic part) for the tests of cointegration were determined on the basis of the modified F-test for autocorrelation (Harvey (1981)) along with the SBC. In addition, the unit root tests indicate that the relevant variables of the import model are non-stationary and integrated of order one.[20] The Johansen's $\lambda_{\max}$ and Trace tests based on the VAR model[21] indicate that there is one cointegrationg vector both in the case of India and Sri Lanka (Appendix Table A.1). The null hypothesis of no cointegration can be rejected at 5% significance level in both cases. The results of the bounds 'F' tests show that the null hypothesis of no cointegration can be rejected at 10% or less significance level for all different specifications of the deterministic terms in the case of both India and Sri Lanka irrespective of lag lengths (Appendix Table A.2). The results from the bounds 't' tests are similar (Table A.2). The overall results from the Johansen's cointegration tests and bounds tests thus provide strong evidence in favor of a unique long run relation among the variables in the import demand model.

Given the strong evidence in favor of a single cointegrating vector in the data for both countries, we estimate the long-run cointegrating relation for import using the DOLS, ARDL, and FM-AADL single equation estimation methods. The optimal lag length for the ARDL model was chosen by SBC starting from 3 lags. In the case of DOLS estimation, sufficient lags and leads of first difference terms are added to regression in order to purge the residual of the serial correlation

[19]For recent applications of ARDL and FM-AADL estimators, see, for example, Emran et. al. (2007)).

[20]The results of unit root tests for all variables except Z_t^* show that all of them can be treated as $I(1)$ variables (for both India and Sri Lanka). The transformation of the data vector for Z_t which ensures separation between constrained and unconstrained regimes 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.

[21]The lag length selected by SBC for the VAR analysis is one for both India and Sri Lanka.

problem.[22] The results from the estimation of the long run demand relationship are reported in Table 1. The regression diagnostic tests [see the bottom panel in Table 1] show that the residuals from the estimated regressions display no problems of serial correlation and/ or non-normality in the case of all three different estimation methods. The estimated coefficients for income and relative price satisfy the theoretical sign restrictions for both India and Sri Lanka regardless of the estimation technique considered. The estimated coefficients are highly statistically significant.[23] For income coefficient, the magnitude of DOLS estimate is lower than the estimates from ARDL and FM-AADL in both countries. For instance, for Sri Lanka, the estimates of income coefficient ($\widehat{\pi}_1$) vary from 0.76 (DOLS) to 0.90 (FM-AADL). The estimates of income coefficient ($\widehat{\pi}_1$) are relatively larger in India [1.17(ARDL, FM-AADL), and 1.02 (DOLS)]. Interestingly, both sets of estimates for India and Sri Lanka are reasonably close to the conventional wisdom of a long run unitary income elasticity.[24]

Similar to income coefficients, the DOLS estimate of relative price coefficient is lowest among the estimates for both of the countries. The ARDL, DOLS, and FM-AADL estimates of relative price coefficient are ($\widehat{\pi}_2 = -0.78$), ($\widehat{\pi}_2 = -0.72$), and ($\widehat{\pi}_2 = -0.82$) respectively for Sri Lanka. For India, the estimates are ($\widehat{\pi}_2 = -0.79$), ($\widehat{\pi}_2 = -0.63$) and ($\widehat{\pi}_2 = -0.71$) respectively.[25] The estimates from ARDL and FM-AADL are in general close to each other. The ARDL estimates of price elasticity are nearly identical for Sri Lanka (-0.78) and India (-0.79). Overall, the estimates from the three different estimators provide reasonably tight bounds for the relative price and income coefficients of the import model for both Sri Lanka and India. The estimates of coefficient ($\widehat{\pi}_3$) of scarcity premium variable, Z_t^*, have correct negative sign and are statistically highly significant across countries and estimators. This confirms the existence of a binding foreign exchange constraint on aggregate imports before the economic liberalization in India (1991) and

[22]The DOLS model involves three lags in the case of Sri Lanka and five lags in the case of India. The 't'statistics reported for the parameter estimates are based on standard error estimates obtained from the Newey and West (1987) adjustment, with Parzen weights and a truncation lag of 12.

[23]One should however be cautious about the significance of the DOLS estimates, as the reported critical values are likely to be much smaller than the appropriate empirical critical values.

[24]The estimates of income coefficients from Johansen's system approach are also remarkably close to ARDL, FM-AADL and DOLS estimates. The Johansen's estimates of income coefficients are ($\widehat{\pi}_1 = 0.91$) for Sri Lanka, and ($\widehat{\pi}_1 = 1.18$) for India.

[25]Similar to income coefficients, estimates of price elasticity from Johansen's system approach are comparable to the estimates reported in Table 1. The price elastictiy estimates from Johansen's approach are and ($\widehat{\pi}_2 = -0.85$) for Sri Lanka, and ($\widehat{\pi}_2 = -0.61$) for India.

Sri Lanka (1977). The estimated import model for Sri Lanka included a dummy for the devastating civil war during 1983-89. 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.

Stability of the Estimated Parameters

A concern discussed in the policy analysis is the possible instability of the estimated elasticity parameters.[26] We test for the stability of the estimated parameters by using Chow break point tests, and Chow's predictive failure tests, CUSUM and CUSUMSQ tests, recursive estimation and rolling regressions. The results for all three estimators show that over all there is no instability problems. To save space, we discuss the results only for the ARDL model. The details of the results for the other two estimators are available from the authors. We note here that the stability properties FM-AADL estimates are very close to those of ARDL, probably reflecting the fact that it is also based on the ARDL approach. According to the Chow breakpoint tests, the ARDL estimates of the parameter vectors display no instability in both India and Sri Lanka [F-statistics= 1.47 (Sri Lanka) with P-value (0.25), and F-statistics = 0.30 (India), with P-value=0.95].[27] The predictive failure tests suggested by Chow do not indicate any mis-specification for ARDL in both countries. The results from CUSUM and CUSUMSQ tests are reported in the Figure 2a - Figure 2b for Sri Lanka and in the Figure 3a-3b for India.[28] Neither of the tests (CUSUM or CUSUMSQ) show any evidence of instability in the estimated parameters at 5 percent significance level for Sri Lanka. For India, the ARDL estimates pass the CUSUM test; but there is some evidence of mild instability according to the CUSUMSQ test. However, this evidence of mild instability is not corroborated by the results from rolling regressions and recursive estimations. The evidence on the stability of individual coefficients from recursive estimation and rolling regression support the conclusion that the estimated parameters are stable irrespective of the estimation technique

[26]For instance, Marquez (2003) reports evidence of parameter instability in the case of income elasticity for U.S. imports. Such parameter instability could result from mis-specification of the long run import relationship particularly when data span over a very long time horizon.

[27]For both India and Sri Lanka, the breakpoint is assumed to be in 1985. For India, the Chow tests are conducted for other breakpoints as well, however, the results remain the same. The experimentation with different break points in the case of Sri Lanka is not possible because of the specification of civil war dummy (non-zero only for 1983-89 period).

[28]For the sake of brevity, we omit the details of results from rolling regressions and recursive estimations.

considered, for both India and Sri Lanka.[29]

[INSERT FIGURE 1 HERE]

(2.2) Comparison with Alternative Models

This sub-section reports the results of the empirical analysis of the modified traditional model (equation (14) excluding Z_t^*) and the foreign exchange availability formulation (equation (14) with log of real foreign exchange availability replacing 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 the tests for the existence and number of cointegrating vector(s) in tabular form.

Modified traditional Model

For India, the evidence from bounds tests clearly indicates the existence of a long run relation as specified in the modified traditional import model. The cointegration rank tests of Johansen also indicate the existence of a single cointegrating relationship except for the case when deterministic part includes an unrestricted intercept. In contrast, the evidence on the existence of a cointegrating relationship among the variables of the modified traditional model is weak in the case of Sri Lanka. 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 shows the absence of a cointegrating relation.[30] 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.

[INSERT TABLE 2 HERE]

Table (2) summarizes the alternative estimates of the parameters of the modified traditional model. The estimation results starkly show the problems with the traditional model as discussed earlier. When import equation is estimated by ARDL or FM-AADL, the price coefficient has a

[29] The results are available from the authors on request.

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

positive sign and is statistically irrelevant for both India and Sri Lanka. The DOLS estimates of price elasticity have the correct negative sign but are statistically insignificant (t-statistics $= -0.12$ for Sri Lanka and -0.97 for India).[31] The magnitudes of price elasticity, according to the DOLS estimates, are also implausibly small (-0.03 for Sri Lanka and -0.28 for India). While the estimate of the income coefficient has the right positive sign for both countries across all three estimators, it is statistically significant and numerically reasonable only for India [1.38 (ARDL), 1.40 (FM-AADL) 1.19 (DOLS)]. Overall, the results from DOLS are relatively better as both the price and income coefficients have the right signs. Yet, the coefficient of relative price is statistically insignificant with implausibly small numerical magnitude.

Foreign Exchange Availability Formulation

The bounds tests and Johansen's rank tests provide strong evidence in favor of the existence of a long run relation among the variables of this model both for India and Sri Lanka. However, the parameter estimates from this model are also problematic. The estimates from all three estimators bear right signs in the case of India, but are either not significant or do not have plausible numerical magnitudes or both (Table 3). For India, the income coefficient is very low for both ARDL (0.35) and FM-AADL (0.45). It is not statistically not significant in case of ARDL at conventional significance levels with a 't'-value of 1.40. The DOLS, on the other hand, yield a statistically insignificant price elasticity estimate ('t'-value of -0.93). Even though the ARDL and FM-AADL give us statistically significant relative price effect with the appropriate negative sign, the numerical magnitudes are much lower compared to the estimated price elasticity reported in Table 1 for India. In the case of Sri Lanka, the ARDL and FM-AADL estimates of relative price and income coefficients have wrong signs, and they are statistically insignificant except for the FM-AADL estimate of the relative price elasticity (please see Table 3). In contrast, the DOLS estimates of both income and price coefficients have correct signs. But, similar to the case of the traditional model, both the income and price 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

[31] Again, we note that the reported 't' statistics should be compared to the appropriate empirical critical values for DOLS, as emphasized by Caporale and Pittis (2004).

significant with correct positive sign according to ARDL and FM_AADL estimates for both India and Sri Lanka. The point estimates from ARDL and FM-AADL for Sri Lanka are virtually equal to unity (1.01 (ARDL), and 1.03 (FM-AADL)) which clearly shows the strength of the *near identity* problem. The DOLS estimate of the coefficient of foreign exchange availability is, however, much smaller (0.71 for Sri Lanka and 0.29 for India) and is statistically significant only for Sri Lanka.

(2.3) 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. 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 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:

$$E_{GDP_t} = \pi_1 \frac{GDP_t}{(GDP_t - P_t^X X_t)} \tag{15}$$

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. As the share of export in GDP varies from year to year, the estimates of income elasticity with respect to GDP also vary. Table 4 summarizes the price and income elasticity (with GDP as the scale variable) estimates for aggregate imports of India and Sri Lanka. The ARDL and FM-AADL estimates are identical for both India and Sri Lanka. According to these estimators, the income elasticity for India ranges between 1.21 to 1.28 and that for Sri Lanka between 0.98 to 1.23. The mean of income elasticity estimates is: 1.23 for India and 1.09 for Sri Lanka according to the ARDL and FM-AADL estimates. The DOLS estimates are slightly lower; it varies from 1.05 to 1.12 for India and from 0.79 to 0.98 for Sri Lanka. The mean of income elasticity, according to the DOLS estimates, is 1.08 for India and 0.88 for Sri Lanka. The estimates of average income elasticity for both countries are fairly close to unity - an estimate which is consistent with conventional wisdom about the long run elasticity of import demand (Marquez, 2003).[32] The ARDL estimates of the price elasticity are remarkably similar across both countries [-0.79 for India and -0.78 for Sri Lanka]. The price elasticity estimate from FM-AADL (-0.82) is also very close to the ARDL estimate in case of Sri Lanka, while in case of India the FM-AADL estimate falls in between the ARDL and DOLS estimates. The DOLS estimates are smaller in absolute magnitude [-0.66 for Sri Lanka and -0.63 for India]. Similar to the income elasticity estimates, the estimates of price elasticity fall within a narrow interval [-0.63 to -0.82], confirming the robustness of the estimates irrespective of the estimation technique employed.

How do the estimates from our preferred model (equation (14)) compare with those available in the literature? Estimating the import demand function for India from a sample similar to ours (1950-96), Sinha(1999) reports an income elasticity of -0.11 which has a theoretically inconsistent negative sign and is statistically insignificant (t-statistics = -0.61). For a shorter sample period(1960-92 and 1960-93 respectively), Caporale and Chui (1999) and Senhadji(1999) yield estimates of income elasticities which are positive and statistically significant. The magnitude of the income elasticities estimated by using DOLS [1.15 (Caporale and Chui, 1999)] and FMOLS [1.33 (Senhadji, 1999)] are comparable to the estimates from the structural model (equation (14)) presented in this paper (see Table 4). The ARDL estimate of income elasticity (1.55) reported by Caporale and Chui (1999) is, however, much higher than our preferred estimate of 1.09 (see Table 4). Their estimate is, however, comparable to that of the modified traditional model (1.46). For Sri Lanka, Sinha (1999) reports a negative income elasticity (-0.39).

The estimates of the price elasticity also display wide variance across studies and estimation techniques. For India, the estimate of price elasticity varies from -1.01 (ARDL) to -0.03 (DOLS) when these two estimation techniques are applied to the same data set (Caporale and Chui, 1999). Price elasticity estimates from the modified traditional model (Table 2) and foreign exchange availability formulation (Table 3) reported in this paper show similar wide variance. In contrast, the estimates from the structural econometric model developed in this paper (Table 1) show remarkable robustness across estimation techniques for both Sri Lanka and India. Although

[32] An estimate of long run income elasticity of import demand significantly different from unity will imply a changing GDP share of import which is puzzling because the GDP shares of consumption and investment are observed to be constant. For a detailed discussion of this point, see Marquez (2003).

available estimates of price elasticity for India and Sri Lanka conform with the theoretical sign restriction, some estimates for India lack statistical significance and also have implausibly low magnitude [e.g. FMOLS estimate (-0.12) with a 't'= -0.25 (Senhadji, 1999) and DOLS estimate (-0.03) with t = -0.09 (Caporale and Chui, 1999)]. Only the ARDL estimate of price elasticity (-1.01) reported by Caporale and Chui (1999) is larger in absolute magnitude compared with our estimate (-0.79) using the same estimation technique. The average estimate of price elasticity for India based on the estimates reported in Caporale and Chui (1999), Sinha(1999) and Senhadji (1999) is about -0.40 which is about half of the average of our estimates (-0.70). For Sri Lanka, the estimated price elasticity from our model (see Table 4) is much higher compared to the estimate of -0.48 reported by Sinha (*op cit*).

To summarize, the estimates from the structural import model are not only robust compared with the other available estimates, the magnitudes of price and income elasticities are also more plausible. Averaging over countries and estimation techniques (see Table 4), the structural import model of this paper provides an income elasticity estimate of 1.10 and price elasticity of -0.73. The estimate of income elasticity is thus close to unity as expected for long run import demand models. More importantly, the magnitude of price elasticity estimated from our preferred model is much higher than the available estimates for these two countries (except for the ARDL estimate reported in Caporale and Chui (1999)). As emphasized by Ghei and Pritchett (1999), this downward bias in price elasticity estimate from the traditional import demand specification results from its inability to control for the "virtual relative price" which could differ substantially from observed relative price due to foreign exchange rationing and import controls.

Conclusions

This paper presents a theoretically consistent and empirically implementable model of aggregate imports of a developing country which had historically been characterized by pervasive trade and exchange rate interventions and for which the time series data on the scarcity premium on imports are not available. The empirical results from India and Sri Lanka demonstrate the inadequacies of the extant import demand models and the superiority of the model presented in this paper, on both statistical and economic grounds. The estimates of the long run income and price elasticities derived from the model satisfy the theoretical sign restrictions and are highly significant, both economically and statistically. The parameter estimates are stable and display little variance across countries and across estimation techniques. The mean of income elasticity estimate is close to the conventional wisdom of a long-run unitary income elasticity (1.10). The mean of price elasticity estimates is about -0.73 which is nearly twice in absolute magnitude compared to the mean of the estimates available in the literature for India and Sri Lanka. The much higher price response of imports uncovered in this paper thus vindicates the long-held view in the literature that the estimate of price response of import demand is seriously biased downward in the traditional formulation of import demand function which ignores the impact of foreign exchange rationing and other restrictions in developing countries (Ghei and Pritchett, 1999). Indeed, the empirical results from this paper suggest that policy analysis such as the calculation of equilibrium exchange rate or the estimation of tariff revenue loss from trade liberalization based on the available low price elasticity estimates is likely to be off the mark by a substantial margin, and thus may lead to wrong policy prescriptions. The model of aggregate imports and the empirical methodology to implement it developed in this paper has wide applicability given the fact that a large number of developing countries had pursued restrictive trade and exchange rate regimes during the decades of 1950s through early 1970s as part of the then-in-vogue import substituting industrialization.

Appendix 1: Parameterization of Scarcity Premium (μ_t^*) and Import Demand

In the theoretical model, the scarcity premium is parameterized by assuming that it is a function of Z_t, where Z_t is defined as follows:

$$Z_t = \frac{Y_t}{F_t}, \tag{16}$$

Where Y_t is the domestic expenditure, and F_t is the foreign exchange available for import. If the foreign exchange constraint (equation (2) in the text) is binding, then

$$M_t = \frac{F_t}{P_t} \tag{17}$$

Putting together equations (7), (8), (11), and (17), and taking logarithm, we have the following expression for $\ln(1+\mu_t^*)$:

$$\ln(1+\mu_t^*) = \alpha \ln(Y_t - F_t) - \eta f_t - (1-\eta)p_t + (b_t - c_t) \tag{18}$$

Utilizing equations (16) and (18), we can derive the sign of $\frac{\partial \mu_t^*}{\partial Z_t}$. We have the following equation

$$\frac{\partial \mu_t^*}{\partial Z_t} = \frac{\partial \mu_t^*}{\partial F_t}\frac{\partial F_t}{\partial Z_t} + \frac{\partial \mu_t^*}{\partial Y_t}\frac{\partial Y_t}{\partial Z_t} \tag{19}$$

It is obvious from equation (16) that $\frac{\partial F_t}{\partial Z_t} < 0$,and $\frac{\partial Y_t}{\partial Z_t} > 0$. Also, observe that the sign of $\frac{\partial \mu_t^*}{\partial Z_t}$ is the same as that of $\frac{\partial \ln(1+\mu_t^*)}{\partial Z_t}$, and we concentrate on the latter expression. From equation (18), we have the following results:

$$\frac{\partial \ln(1+\mu_t^*)}{\partial F_t} = -\frac{\eta}{F_t} - \frac{\alpha}{(Y_t - F_t)} < 0 \tag{20}$$

$$\frac{\partial \ln(1+\mu_t^*)}{\partial Y_t} = \frac{\alpha}{(Y_t - F_t)} > 0 \tag{21}$$

From equations (19), (20),and (21) the sign of $\frac{\partial \mu_t^*}{\partial Z_t}$ is unambiguously positive.
Given that $\frac{\partial \mu_t^*}{\partial Z_t} > 0$, we can now derive the *á priori* sign restriction on $\frac{\partial M_t}{\partial Z_t}$ as follows:

$$\frac{\partial M_t}{\partial Z_t} = \frac{\partial M_t}{\partial \mu_t^*}\frac{\partial \mu_t^*}{\partial Z_t} < 0$$

References

1. Banarjee,A., J.Dolado, R. Mestre (1998), 'Error-correction mechanisms tests for cointegration in single equation framework,' *Journal of Time Series Analysis*,19, 267-283.

2. Bertola, G. and R. Faini (1990), 'Import demand and non-tariff barriers: the impact of trade liberalization: an application to Morocco,' *Journal of Development Economics*, 34, 269-286.

3. Bhagwati, J. and T.N. Srinivasan, (1993), 'India's economic reforms', New Delhi: Government of India, Ministry of Finance.

4. Caporale, G and N. Pittis (2004): " Estimator Choice and The Fisher's Paradox: A Monte Carlo Study", *Econometric Reviews*, 23 (1), 25-52.

5. Caporale, G.M and M. Chui (1999), 'Estimating income and price elasticities of trade in a cointegration framework', *Review of International Economics*, 7, 254-264.

6. Caporale, G.M and N. Pittis (1999), 'Efficient estimation of cointegrating vectors and testing for causality in vector autoregressions', *Journal of Economic Surveys*, 13, 1-35.

7. Ceglowski, J, (1991), 'Intertemporal substitution in import demand,' *Journal of International Money and Finance*, 10, 118-130.

8. Cheung, Y and K. S. Lai (1993), 'Finite-sample sizes of Johansen's likelihood ratio tests for cointegration,' *Oxford Bulletin of Economics and Statistics*, 55, 313-328.

9. Clarida, R. H. (1994), 'Cointegration, aggregate consumption and the demand for imports: A structural econometric investigation,' *American Economic Review*,84, 298-308.

10. Cuthbertson, A.G. and P. Athukorala (1989), 'Liberalizing Foreign Trade: Sri Lanka,' in D. Papageorgiou, M. Michaely and A.M. Choksi ed. *Liberalizing Foreign Trade*, Vol.5, Basil Balckwell, In., Massachusetts: USA.

11. Devarajan, S, Delfin Go and Hongyi Li, (1999), 'Quantifying the fiscal effects of trade reform: a general equilibrium model estimated for 60 countries ', Policy Research Working Paper # 2162, World Bank.

12. Emran, M. S, F. Shilpi, and I Alam (2007): "Economic Liberalization and Price Response of Aggregate Private Investment: Time Series Evidence from India", Canadian Journal of Economics, August, 2007.

13. Emran, M.S. and F. Shilpi (1996), 'Foreign exchange rationing and aggregate import demand function,' *Economics Letters*, June, 315-322.

14. Elliott, G (1998): "On the Robustness of Cointegration Methods When Regressors Have Almost Unit Roots", *Econometrica*, 66(1), 149-158.

15. Faini, R, L. Pritchett, and F. Clavijo (1992), 'Import demand in developing countries,' in M. Dagenais and P. Muet ed. *International Trade modelling*, Chapman and Hall, NY and Melbourne, 279-297.

16. Ghei, Nita and Lant Pritchett, (1999), 'The three pessimisms: Real exchange rates and trade flows in developing countries ', in L. Hinkle and P. Montiel, eds. *Exchange rate misalignment:concepts and measurement for developing countries*, Washington D.C.: World Bank.

17. Goldstein,M and M. Khan (1985), 'Income and price effects in foreign trade,' in *Handbook of international economics*, (Elsevier, Amsterdem).

18. Gregory, A, A. Haug, and N. Lomuto (2002): 'Mixed signals among tests for cointegration', Journal of Applied Econometrics, 2002.

19. Harvey, A, (1981): *The econometric analysis of time series*, Phillip Allan, London.

20. Hinkle, Lawrence and Peter Montiel, (1999), *Exchange rate misalignment:concepts and measurement for developing countries*, Washington D.C.: World Bank.

21. Hubbard, G.R., and A. K. Kashyap, (1992), 'Internal net worth and the investment process : An application to U.S. agriculture,' *Journal of Political Economy*,100, 506-34.

22. Johansen, S. (1995), *Likelihood based inference in cointegrated vector auto-regressive models*, Oxford University Press, Oxford.

23. Krueger, A.O., M, Schiff and A. Valdez (1991), *The political economy of agricultural pricing policy*, Johns Hopkins University Press for the World Bank.

24. Maddala, G. S. and I. Kim (1998), *Unit roots, cointegration and structural change*, Cambridge University Press, Cambridge, UK.

25. Mazeri, A, (1995),'Imports under a foreign exchange constraint: the case of Islamic Republic of Iran,' IMF, WP/95/97.

26. Marquez, Jaime (2003), 'U.S. Imports ", *Mimeo*, Federal Reserve Board.

27. Moran, C, (1989), 'Imports under a foreign exchange constraint,' *World Bank Economic Review*, 3, 279-95.

28. Neary, J.P. and K.W.S. Roberts (1980), 'The theory of household behavior under rationing,' *European Economic Review*, 13, 25-42.

29. Newey, Whitney K. and Kenneth D. West, (1987), 'A simple positive semi-definite heteroscedasticity and autocorrealtion consistent covariance matrix ', *Econometrica*, 55, 703-8.

30. Ogaki, M. and J.Y. Park (1997), 'A cointegration approach to estimating preference parameters,' *Journal of Econometrics*, 82, 107-134.

31. Panagariya, A. (1999), 'The WTO Trade Policy Review of India, 1998,*World Economy*, 799-824.

32. Panopoulou, E and N. Pittis (2004): "A Comparison of Autoregressive Distributed Lag and Dynamic OLS Cointegration Estimators in the Case of a Serially Correlated Cointegration Error", *Econometrics Journal*, 7, 585-617.

33. Pesaran, M. H., Y. Shin, and R. Smith (2001),'Bounds testing approaches to the analysis of long run relationships,' Journal of Applied Econometrics, 16, 289-326.

34. Pesaran, M. H and Y. Shin (1999), 'An autoregressive distributed lag modelling approach to cointegration analysis,' *Econometrics and economic theory in 20th century: theRagnar Frisch centennial symposium,* eds. S. Strom, Cambridge University Press, Cambridge.

35. Pesaran, H and B. Pesaran (1997): *Working with Microfit*, Camfit Data Limited.

36. Senhadji, A, (1998): 'Time series estimation of structural import demand equations: a cross-country analysis,' *IMF Staff Papers*, 45, 236-268.

37. Sinha, D, (1999): 'The Marshall-Lerner-Robinson condition and the trade elasticities in five Asian countries', *mimeo*, Macquarie University and Yale University.

38. Stock, James and Mark Watson,(1993), 'A simple estimator of cointegrating vectors in higher order integrated systems ', *Econometrica*, 61, 783-820.

39. Urbain, J. P., 1992, Error correction models for aggregate imports, in: M. Dagenais and P.A. Muet, ed., International trade modelling, (Chapman and Hall, London).

40. Winters, Alan, L,(1987), 'An empirical intertemporal model of developing countries' imports,' *Weltwirtschaftliches Archiv*, 123, 58-80.

41. Winters, Alan, L, and Paul A. Brenton (1993), 'Modelling quantitative trade restrictions: rationing in the Rotterdam model,' *De Economist*, 141, 112-126.

42. Zeldes, S. (1989), 'Consumption and liquidity constraints: an empirical investigation,' *Journal of Political Economy*, 97, 1196-1216.

Table 1: Estimates of Long-run Import Demamd Function, Sri Lanka and India

	Sri Lanka			**India**		
	ARDL	DOLS	FMADL	ARDL	DOLS	FMADL
h	0.85	0.76	0.9	1.17	1.02	1.17
	(4.10)	(4.65)	(5.80)	(10.43)	(23.27)	(14.67)
p	-0.78	-0.72	-0.82	-0.79	-0.63	-0.71
	(-4.14)	(-4.72)	(-6.65)	(-2.33)	(-2.80)	(-2.98)
Deterministic part						
Z^*	-0.22	-0.15	-0.22	-0.04	-0.01	-0.04
	(-5.04)	(-4.77)	(-5.04)	(-2.59)	(-2.88)	(-2.59)
Intercept	0.78	1.05	0.78	-3.9	-3.01	-3.9
	(0.50)	(0.78)	(0.50)	(-3.11)	(-5.90)	(-3.11)
Civil War Dummy	-0.43	-0.1	-0.43			
	(-2.53)	(-2.97)	(-2.53)			
Residual analysis						
Serial correlation (F)	1.18	0.45	1.18	0.76	2.00	0.76
	[0.29]	[0.52]	[0.29]	[0.39]	[0.17]	[0.39]
Normality (χ^2)	0.2	1.67	0.2	3.49	0.17	3.49
	[0.91]	[0.43]	[0.91]	[0.18]	[0.92]	[0.18]

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 for Sri Lanka

D takes a value of 1 for 1952-1991 and zero otherwise for India

t-statistics are reported in the parentheses and P-values in the brackets.

Table 2: Estimation of Long-run Relationship in Modified Traditional Model

	Sri Lanka			**India**		
	ARDL	DOLS	FMADL	ARDL	DOLS	FMADL
h	0.83	0.29	0.51	1.38	1.19	1.4
	(0.86)	(0.83)	(1.33)	(9.07)	(13.25)	(12.12)
p	0.26	-0.03	0.04	0.05	-0.28	0.1
	(0.25)	(-0.12)	(0.13)	(0.12)	(-0.97)	(0.28)
Deterministic part						
Intercept	1.7	4.87	1.7	-6.43	-4.73	-6.43
	(0.24)	(1.86)	(0.24)	(-4.17)	(-5.03)	(-4.17)
Civil War Dummy	-0.88	-0.1	-0.88			
	(-0.73)	(-2.33)	(-0.73)			
Speed of Adjustment	-0.1		-0.1	-0.22		-0.22
	(-0.88)		(-0.88)	(-2.44)		(-2.44)
Residual analysis						
Serial correlation (F)	0.13	4.48	0.13	0.34	37.72	0.34
	[0.72]	[.05]	[0.72]	[0.56]	[0.00]	[0.56]
Normality (χ^2)	2.02	1.09	2.02	0.72	0.52	0.72
	[0.36]	[0.58]		[0.70]	[0.77]	[0.70]

Note: m=log(total imports)

h= log(home good consumption)

p=log(import price index/consumer price index)

t-statistics are reported in the parentheses and P-values in the brackets.

Table 3: Estimation of Long-run Relationship in Foreign Exchange Availability Model

	Sri Lanka			**India**		
	ARDL	DOLS	FMADL	ARDL	DOLS	FMADL
h	-0.05	0.15	-0.03	0.35	0.85	0.45
	(-0.27)	(3.51)	(-0.27)	(1.40)	(4.22)	(3.11)
p	0.23	-0.01	0.22	-0.45	-0.33	-0.32
	(1.45)	(-0.27)	(2.35)	(-2.23)	(-0.93)	(-2.11)
f	1.01	0.71	1.03	0.61	0.29	0.58
	(5.94)	(10.9)	(11.08)	(3.31)	(1.60)	(5.48)
Deterministic part						
Intercept	0.36	0.81	0.36	-1.06	-3.53	-1.06
	(0.29)	(1.14)	(0.29)	(-0.82)	(-3.55)	(-0.82)
Civil War Dummy	-0.17	-0.08	-0.17			
	(-1.57)	(-2.39)	(-1.57)			
Speed of Adjustment	-0.43		-0.43	-0.36		-0.36
	(-3.66)		(-3.66)	(-3.56)		(-3.56)
Residual analysis						
Serial correlation (F)	1.03	0.27	1.03	1.83	37.82	1.83
	[0.32]	[.61]	[0.32]	[0.19]	[0.00]	[0.19]
Normality (χ^2)	0.02	0.54	0.02	0.53	3.06	0.53
	[0.99]	[0.77]	[0.99]	[0.77]	[0.22]	[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)

t-statistics are reported in the parentheses and P-values in the brackets.

Table 4: Elasticity Estimates

	Elasticity Estimates					
	Sri Lanka			**India**		
	ARDL	DOLS	FMADL	ARDL	DOLS	FMADL
Income*						
Average	1.09	0.88	1.09	1.23	1.08	1.23
Minimum	0.98	0.79	0.98	1.21	1.05	1.2
Maximum	1.23	0.98	1.23	1.28	1.12	1.28
Price	-0.78	-0.66	-0.82	-0.79	-0.63	-0.71

Note:*: Income elasticity is defined with respect to GDP by dividing elasticity estimates (with respect to expenditure on home goods consumption) in Table 1 by (1-share of export in GDP) (see formula in equation (15) in in the text).

Table A.1: Tests for Existence of Cointegrating Vectors using Johansen's FIML Approach

	Eigen Values	Null Hypothesis	Lmax	Trace	95% Critical Values[1] Lmax	Trace
India						
k=1	0.43	r=0	27.11	48.86	23.35	36.92
	0.27	r<=1	15.01	21.74	16.81	21.36
Sri Lanka						
k=1	0.72	r=0	43.77	49.88	22.99	34.28
	0.15	r<=1	5.71	6.11	16.2	19.41

k indicates the lag length of the VAR model.

1/ 95% critical values are adjusted for sample size by using Response Surface Regressions of Cheung and Lai (1993)

Table A.2: Bound Tests for Long-run Relationship in an ARDL model

Lags			Restricted Intercept	Unrestricted Intercept	Restricted Trend
			Deterministic part		
India					
1	Bound test	F-statistic	5.96*	5.03**	5.15**
	Bound test	t-statistic	-3.62**	-3.62**	-3.87***
	intercept	t-statistic	-3.43	-3.43	-2.78
	trend	t-statistic	-	-	-2.03
2	Bound test	F-statistic	5.37*	4.43**	4.54**
	Bound test	t-statistic	-3.55***	-3.55***	-3.76
	intercept	t-statistic	-3.32	-3.32	-2.71
	trend	t-statistic	-	-	-1.94
3	Bound test	F-statistic	8.18*	6.6*	5.91*
	Bound test	t-statistic	-4.19**	-4.19**	-4.24**
	intercept	t-statistic	-3.06	-3.06	-2.25
	trend	t-statistic	-	-	-1.5
Sri Lanka					
1	Bound test	F-statistic	5.86*	7.3*	6.56*
	Bound test	t-statistic	-3.95**	-3.95**	-4.21**
	intercept	t-statistic	1.09	1.09	1.69
	trend	t-statistic	-	-	1.5
	Civil War	t-statistic	-3.65	-3.65	-1.24
2	Bound test	F-statistic	5.91*	7.1*	9.82*
	Bound test	t-statistic	-4.08**	-4.08**	-5.65*
	intercept	t-statistic	0.9	0.9	3.33
	trend	t-statistic	-	-	3.2
	Civil War	t-statistic	-3.65	-3.65	-0.43
3	Bound test	F-statistic	5.62*	6.37*	8.99*
	Bound test	t-statistic	-3.96**	-3.96**	-5.43*
	intercept	t-statistic	0.84	0.84	3.24
	trend	t-statistic	-	-	3.12
	Civil War	t-statistic	-3.68	-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

Figure 1a: Plot of CUSUM of Recursive Residuals (ARDL), Sri Lanka

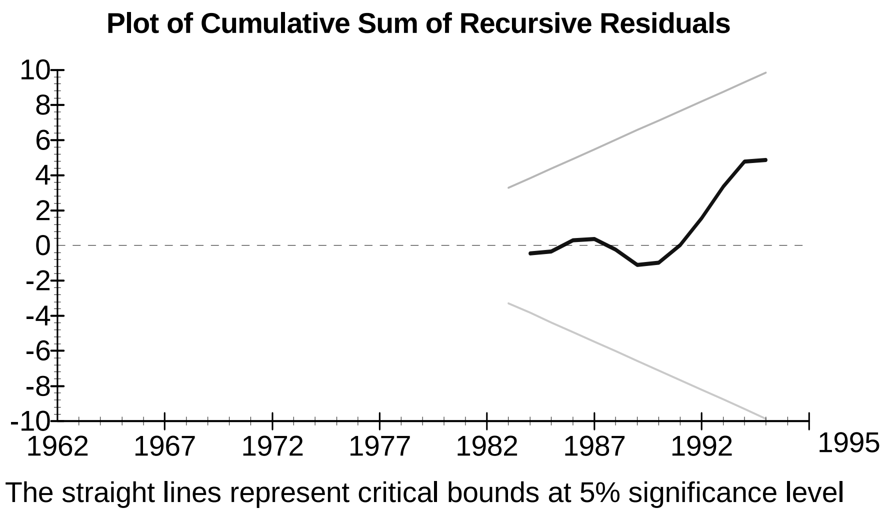

Figure 1b: Plot of CUSUMSQ of Recursive Residuals (ARDL), Sri Lanka

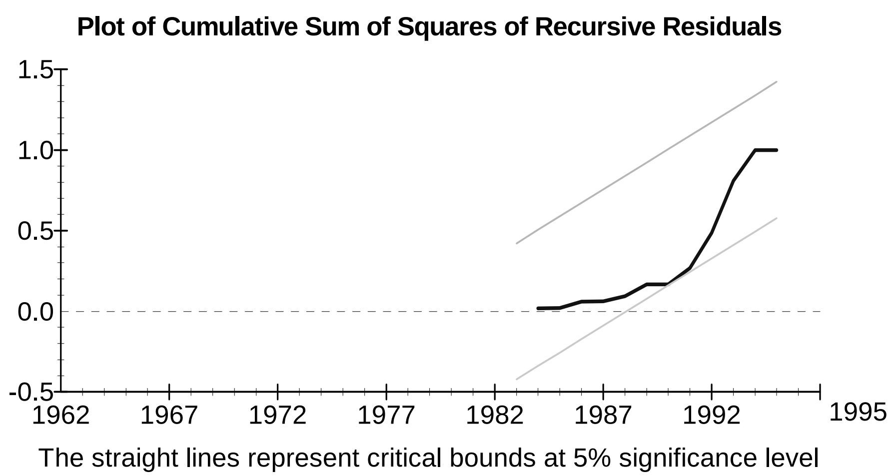

Figure 2a: Plot of CUSUM of Recursive Residuals (ARDL), India

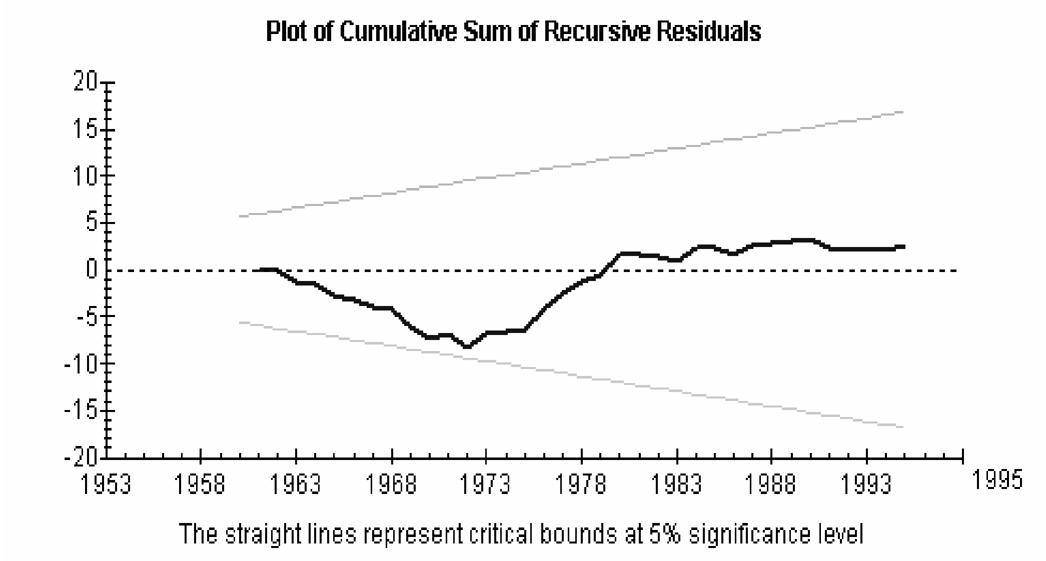

Figure 2b: Plot of CUSUMSQ of Recursive Residuals (ARDL), India