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IMPORT DEMAND FUNCTIONS: EVIDENCE FROM CIBS

YAN ZHOU AND SMILE DUBE*

California State University, USA

This paper adopts the bounds testing approach to test for the validity of the cointegration or stationarity restriction embodied in five import demand model specifications for CIBS during the period 1970-2007. It identifies long-run relationships in a subset of the five models for each CIBS countries. We find that long-run income elasticities are much higher compared to earlier studies and are higher than the short-run counterparts for CIBS. In addition, contrary to the traditional wisdom, price elasticities are not significantly negative for these countries.

Keywords: Import Demand, ARDL, Bounds Test, Income Elasticities, Price Elasticities
JEL classification: F43, O11, O24

1. INTRODUCTION

According to Nayyar (2008), China, India, Brazil and South Africa (CIBS) can potentially become southern engines of global economic growth with increasing capacities to support prosperity in other countries. The four countries together account for 60% of the population of developing countries and over 40% of the GDP of the developing world. One important way that these four countries can support other countries' economic growth is through their substantial increase in imports. As shown in Table 1, CIBS' imports as a percentage of the world increased from 3% during 1970-80 period to 8% in the 2001-07 period, which has almost tripled. The merchandise trade by the CIBS accounts for more than one-fourth of the overall merchandise trade in developing countries.

In addition, Harry Broadman, Economic Adviser for the Africa Region at the World Bank, believes that the recent boom of sub-Saharan Africa's exports to some of the CIBS countries is a potentially pivotal opportunity for sub-Saharan African countries to move beyond their traditional reliance on single-commodity exports and move up from the bottom of the international production chain. He points out that some of the CIBS

* CIBS stands for China, India, Brazil, and South Africa. The authors would also like to thank anonymous referees for useful suggestions. Usual disclaimers apply.

countries have burgeoning middle classes whose members are increasingly buying sub-Saharan Africa's light manufactured products, household consumer goods and processed foods, unlike in the past when the imports from the sub-Saharan African countries were mainly natural resources (see Broadman, 2008). For example, sub-Saharan Africa's exports to China increased at an annual rate of 48% between 2000 and 2005, two and a half times as fast as the rate of the region's exports to the United States and four times as fast as the rate of its exports to the European Union over the same period.

Table 1. Total Imports and Exports of CIBS (in millions of U.S. dollars)

Countries	Total Imports			Total Exports		
	1970-80	1981-91	2001-07	1970-80	1981-91	2001-07
Brazil	132,894	205,038	525,770	102,737	297,495	705,332
China	88,685	450,016	3,920,071	85,587	424,477	4,572,576
India	44,868	191,240	812,826	34,990	130,891	595,257
South Africa	83,326	187,523	373,399	108,800	221,011	321,198
Totals of above	349,773	1,033,817	5,632,066	332,114	1,073,874	6,194,363
CIBS as a percentage of World	3%	4%	8%	3%	4%	9%
CIBS as a percentage of Developing Countries	15%	16%	27%	12%	16%	27%

Source: IMF Financial Statistics CD Rom. Calculations by authors.

Therefore, it is interesting to investigate the import demand behavior of the CIBS, given its importance and potential in supporting prosperity in other countries. A large number of studies have been done on estimating the aggregate import demand functions for different countries. The import demand specification is important for informed policy analysis in many areas (Tang, 2003; Emran and Shilpi, 2008). Yet none of the existing studies on import demand functions have looked at the CIBS as group, nor have they investigated the most recent import booms in the CIBS.

This paper is the first study to estimate the import demand functions of the CIBS using recent data. In addition, the paper is the most complete study of the import demand specifications for these countries in the sense that we explore all the existing import demand models in the literature, including the most recent model developed by Emran and Shilpi (2010). We find that the four fast-growing developing countries share similar patterns in their import demand behavior. Specifically, income elasticities are very high and almost always statistically significant for all four countries. Our finding suggests that a 1% increase in income is likely to induce a 1.5% - 4.2% increase in imports in the long-run. Our long-run income elasticity estimates of CIBS using most recent data are much higher compared to earlier studies about other developing nations or about some of the CIBS countries using earlier data, which are usually less than 0.85% (See Goldstein and Khan, 1985; Senhadji, 1998; Emran and Shilpi, 2010). On the other hand, the income elasticities in advanced nations are usually greater than one, the majority of which falls in the range of 1.2% - 2.3% (Caporale and Chui, 1999; Chang, Ho and Huang, 2005). This seems to support the argument in favor of CIBS' potential to become southern engines of global economic growth. We also find higher income elasticities in the long run as compared to the short run, which implies that continued economic growth and import growth in the CIBS countries are likely to have negative impacts on their balance of payments in the long run. More importantly, we find that price elasticities are either significantly positive or statistically insignificant for the four countries, unlike the traditional wisdom that the aggregate imports are negatively associated with current relative prices. This may be attributed to the fact that for a fast-growing open economy, certain goods important for development or exports, such as capital goods or intermediate goods, are increasingly imported and consumed even if their relative prices increase. In terms of policy implications, the finding suggests that trade negotiations that aim to lower or remove tariff and nontariff barriers in the CIBS will not necessarily lead to a rise in the flow of their imports.

The contributions of this paper are as follows. First, it is the first study to analyze the import demand behavior of four potential southern engines of global economic growth and to detect some similar patterns among them. Second, by using the most recent data and exploring all the existing model specifications in the literature, this paper finds that price elasticities are no longer significantly negative for these countries, which is contrary to the traditional wisdom.

The paper is organized as follows. Section 2 is a brief review of literature on import demand in developing countries. Section 3 presents an empirical model and Section 4 provides the results. Finally, Section 5 gives the concluding remarks.

2. BRIEF REVIEW OF RELEVANT LITERATURE ON IMPORT DEMAND FUNCTION ESTIMATION

So far there are mainly five types of models with regard to aggregate import demand function estimation: the traditional model with income measured as real GDP, the revised traditional model with income measured as the real value of GDP minus exports (or the Senhadji model), the disaggregated or decomposed GDP model, the dynamic structural import demand model (or the "National Cash Flow" model) and the structural model that incorporates a binding foreign exchange constraint (or the "Emran & Shilpi" model).

The specifications of the traditional model and the Senhadji model are relatively straightforward and will be discussed in Section 3-2. The ideas of the three other models are briefly discussed as follows. The disaggregated or decomposed GDP model is adopted by many studies to take into account the fact that different macro components of final expenditure have different import contents (Giovannetti, 1989; Abbott and Seddighi, 1996; Min, Mohammad, and Tang, 2002; Mohammad and Tang, 2000; Tang, 2002; and Tang, 2003). This approach decomposes GDP into three categories: final consumption expenditure, expenditure on investment goods, and exports.

The dynamic structural import demand model is developed by Xu (2002). It takes into account a growing economy, rather than an endowment economy, and investment and government activity. The model replaces real GDP with a "national cash flow" variable (See section 3-2 for a detailed definition of the "national cash flow" variable).

Finally, the Emran & Shilpi (2010) model is a structural econometric model of a two goods representative agent economy. They circumvent the issue of unavailability of data on the domestic market clearing price of imports by parameterizing the Lagrange multiplier of a binding foreign exchange constraint at the administered prices of imports. A group of studies added a foreign exchange availability variable on an *ad hoc* basis to a standard import demand model to reflect a binding foreign exchange constraint (Moran, 1989). However, Emran and Shilpi (2010) point out that these studies suffer 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. They avoid the problem 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.

Existing studies of the import demand functions of these four countries usually focus on only a single country or adopt a subset of the models and hence present mixed results for elasticity estimates. Tang (2003) analyzed the long-run relationship of China's aggregate import function for the period 1970-1999 using four models: namely the traditional model, the Senhadji model, the "National Cash Flow" model, and the disaggregated model. He found the long-run price elasticity to be in the range of -0.45 to -0.6 and the long-run income elasticity to be in the range of -0.19 to 0.73. Dutta and Ahmed (2004) examined Indian aggregate import demand using the traditional model for the period 1971-1995, where they found long-run price and income elasticities to be -0.37 and -0.03, respectively. Emran and Shilpi (2010) adopted a structural model incorporating a binding foreign exchange constraint and examined the Indian aggregate import demand for the period 1952-1999. They found a long-run price elasticity of -0.78 and a long-run income elasticity of 0.85. Senhadji (1998) studied the import demand for a group of countries for the period 1960-1993 using the revised traditional model with income measured as the real value of GDP minus exports, or the Senhadji model. He found that the long-run price elasticity and income elasticity were -0.30 and 0.20, respectively for Brazil; varied in the ranges -0.44~-0.53 and 0.30~0.33, respectively for South Africa; and varied in the ranges -0.03~-0.14 and 0.28~0.49, respectively for India. All the above studies on the four countries' import demand focus only on time periods before year 2000, therefore cannot contribute to explaining the recent import booms in these countries. Although the magnitudes of income and price elasticities vary, by and large, the studies found income elasticities to be less than one (i.e., inelastic), and price elasticities to be negative.

This paper estimates the import demand functions of the CIBS using more recent data (1970-2007) and examines all five model specifications for the four countries. The next section provides a description of the empirical methodology and the specifications of empirical models.

3. METHODOLOGY: MODEL SPECIFICATION

3.1. Methodology and Data

We use bounds test for the validity of the cointegration or stationarity restriction embodied in the import demand function. We adopt the bounds testing approach mainly for two reasons. First, this approach is applicable irrespective of whether the explanatory variables are stationary or nonstationary (Pesaran *et al*., 2001). It avoids the much discussed problems associated with the unit-roots pre-testing (See Maddala and Kim, 1998). Second, Mah (2000) and Pattichis (1999) point out that the small sample bias of cointegration analysis could be addressed by employing the unrestricted-error-correction model (UECM) and bounds test. Given our small samples for each country, the bounds test produces more accurate estimates than the usual residual-based Engle-Granger test (1987) or the VAR-based Johansen test.

For estimation of the cointegrating vector, we adopt the Autoregressive Distributed Lag (ARDL) approach. Specifically, we use the two-step procedure suggested by Pesaran and Shin (1999). We choose the single equation estimation method from the large number of estimation techniques available in the literature for two reasons: its desirable small sample properties and its ability to address potential endogeneity problems. 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.

We use annual data from 1970 to 2007 for all countries. The use of annual data does not discount the robustness of the cointegration test (Tang, 2003). Hakkio and Rush (1991) find that increasing the number of observations by using monthly or quarterly data does not make the cointegration tests more robust and that the length of the period under consideration is more important than the number of observations. Our paper extends the length of the period under study as much we can for each country in order to improve robustness. We also test the stability of the estimated parameters by employing the Cumulative Sum of Recursive Residuals (CUSUM) and Cumulative Sum of Squares of Recursive Residuals (CUSUMSQ) tests. The data we use come from the International Financial Statistics data provided by the International Monetary Fund and World Development Indicators provided by the World Bank.

3.2. Model Specifications

We experimented with all the existing import demand model specifications in the literature and identified the model specifications where a cointegrating relation can be found for the CIBS countries. According to Perman (1991), cointegration analysis could serve as a misspecification test to guide variable selection in empirical macroeconomics. Tang (2003) also chooses the appropriate import demand model specification based on the presence of a cointegrating relation. Therefore, we adopt this criterion when we choose the appropriate model specifications for each CIBS country. The details of each of the five important demand model specifications are given as follows.

First, the traditional model suggests that import demand can be modeled by two determinants: relative prices and real domestic activity (Gafar, 1988; Carone, 1996; Hong, 1999; and Tang, 2003). One can specify the import function in log-linear form as

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln p_t + \beta_3 D + \varepsilon_t \,, \tag{1}$$

where m_t, y_t, and p_t represent the volume of imports, real GDP, and the relative price, respectively. The relative price is measured as the import price index deflated by an index of domestic prices. The variable *D* is a dummy variable capturing structural changes. ε_t is a random error. In earlier studies, the error term is assumed to be orthogonal to all determinants. However, in our paper, we adopt the two-step procedure suggested by Pesaran and Shin (1999) and allow for the possibility of endogeneity.

The second model is a variation of the traditional model, which replaces real GDP with real GDP minus exports (gex) as a measure for real domestic activity (Senhadji, 1998).

$$\ln m_t = \beta_0 + \beta_1 \ln(gex_t) + \beta_2 \ln p_t + \beta_3 D + \varepsilon_t \,. \tag{2}$$

The real GDP minus exports variable presumably is a more precise measure of real domestic activity. This model is often referred to as the Senhadji model.

Third, the real domestic activity variable is decomposed into three broad categories to capture the possibility that the components of real domestic activity have different import contents. This model specification is often referred to as the disaggregated model. It addresses the possible aggregation bias when the different macro components have different import contents (see Abbot and Seddighi, 1996; Tang and Mohammad, 2000; and Tang, 2003). The model specification is given as follows:

$$\ln m_t = \beta_0 + \beta_1 \ln ei_t + \beta_2 \ln fc_t + \beta_3 \ln ep_t + \beta_4 D + \varepsilon_t \,, \tag{3}$$

where ei_t, fc_t, and ep_t represent expenditures on investment goods, final consumption expenditure, and exports, respectively.

The fourth model is the dynamic structural import demand model proposed by Xu (2002). Xu took into account economic growth and derived a structural import demand function using an intertemporal optimization approach. The model specification is:

$$\ln m_t = \beta_0 + \beta_1 \ln ncf_t + \beta_2 \ln p_t + \beta_3 D + \varepsilon_t, \tag{4}$$

where ncf_t is the national cash flow. The national cash flow variable is calculated as ($gdp_t - i_t - g_t - ex_t$), where i_t, g_t, and ex_t represent investment, government expenditure, and exports, respectively. Again, D represents a dummy variable capturing structural changes.

Finally, a number of studies have identified real foreign exchange reserves as an additional variable affecting import demand (see Moran, 1989; Faini *et al*., 1992; Dutta and Ahmed, 1999; and Arize *et al*., 2004). Omission of foreign exchange reserves may bias a model's empirical estimates and overstates the influence of the included explanatory variables. In addition, a more recent study by Emran and Shilpi (2010) finds that simple inclusion of foreign exchange reserves as a regressor creates new problems. If foreign exchange is used as a regressor when the foreign exchange constraint is binding, it alone determines the volume of imports completely. That is, the estimated equation is close to an identity and the coefficients of relative price and real domestic activity are devoid of any behavioral interpretations. As a result, Emran and Shilpi (2010) propose the following model:

$$\ln m_t = \beta_0 + \beta_1 \ln(Y_t - P_t M_t) + \beta_2 \ln p_t + \beta_3 \ln(1 + \mu_t^*) + \xi_t, \tag{5}$$

where m_t and p_t denote the volume of imports and the relative price, respectively. ξ_t is a mean zero (strictly) stationary process. The variable ($Y_t - P_t M_t$) refers to real home good consumption and is denoted by h_t in the regression results. The variable μ^* is the scarcity premium on imports and the availability of foreign exchange reserves is used as a proxy for μ^*. To avoid the problem of near identity, μ^* is parameterized by the ratio of total domestic expenditure (GDP+import-export) to the available foreign exchange reserves (denoted below as Z_t). According to Emran and Shilpi, the intuition behind this parameterization is that given the prices, the excess demand for (and hence the scarcity premium 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. Unlike the case when foreign reserve availability is directly included in the regression, this parameterization avoids the one to one relation between imports and Z_t in a foreign exchange constrained regime and is not subject to the problem of near identity. For empirical implementation, the functional form of $\mu_t^*(Z_t)$

is assumed to be $\mu_t^*(Z_t) = e^{\theta_1 Z_t} - 1$; $\theta_1 \geq 0$.

In addition, a priori restriction about trade liberalization and exchange rate regime changes is incorporated by multiplying the parameterized variable by a dummy variable that takes on the value of 1 for the foreign exchange constrained period and zero afterwards. As Emran and Shilpi (2010) point out, the scarcity premia should be approximately zero for the sample periods after liberalization or regime changes. The term capturing the foreign exchange constraint with the above a priori restriction is denoted by *fer* in the regression results.

To carry out the bounds test, import demand Equations (1) - (5) are converted into UECM forms.[1] Instead of listing the UECM form for each model, we showcase the idea by converting the Emran and Shilpi model (i.e., Equation (5)) into the following:

$$\Delta \ln m_t = \alpha_0 + \sum_{i=0}^{k} \alpha_{1i} \Delta \ln h_{t-i} + \sum_{i=0}^{k} \alpha_{2i} \Delta \ln p_{t-i} + \sum_{i=1}^{l} \alpha_{3i} \Delta \ln m_{t-i}$$
$$+ \alpha_4 ECM_{t-1} + \alpha_5 \Delta fer_t + \alpha_6 \Delta intercept_t + e_t, \tag{6}$$

where $ECM_{t-1} = \ln m_{t-1} - \varphi_1 \ln h_{t-1} - \varphi_2 \ln p_{t-1} - \varphi_3 fer_{t-1} - \varphi_4 intercept_{t-1}$, $k = 0, 1, 2, \ldots, s$, $l = 0, 1, 2, \ldots, s$, and s is the maximum lag. Here *fer* and *Intercept* are deterministic variables in an augmented autoregressive distributed lag model (see Pesaran and Pesaran, 2009; for more details). The variables $\ln m_t$ and $\ln p_t$ represent the logarithm of real imports and relative price in period *t*. The variable *h* represents real home good consumption, which corresponds to $(Y_t - P_t M_t)$ in Equation (5). The variable *fer* is calculated as Z_t^* (the dummy variable that captures trade liberalization and exchange rate regime changes). Emran and Shilpi (2010) used the sum of exports, remittances, and disbursed foreign aid as a proxy for real foreign exchange availability. However, data on disbursed foreign aid are not available for all countries for the 1970-2007 period examined in this paper. Hence we used foreign exchange reserves data as a proxy for real foreign exchange availability instead.

The null hypothesis of no cointegration is $H_0: \alpha_4 = \alpha_5 = \alpha_6 = 0$ and the alternative hypothesis is that at least one of the parameters is nonzero and a cointegrating relation exists. The bounds test is a Wald test (F-statistic) testing the joint significance of the coefficients of the lagged variables in the UECM. The asymptotic distribution of the F statistic is nonstandard under the null hypothesis of no cointegration, irrespective of whether the explanatory variables are I(0) or I(1). The F statistic is compared to the non-standard critical bounds.[2] If the F statistic falls below the lower bound, then we

[1] The Error Correction Representation for the first four models, i.e., the UECM forms, are given in Appendix B.

[2] The critical values reported in Pesaran *et al*. (2001) are for samples of 500 and 1000 observations and

cannot reject the null hypothesis of no cointegration. If the F statistic falls between the lower and upper bounds, then no conclusion can be drawn about the cointegration without an explicit knowledge of integration of the variables.[3] If the F statistic exceeds the upper bound, then we can reject the null hypothesis of no cointegration. The appropriate lag structure is determined by using the Schwartz Bayesian Criterion (SBC).[4]

Once a long-run relationship is identified amongst the variables, the estimated long-run elasticity is the negative value of the coefficient for the lagged explanatory variable divided by the coefficient for the lagged dependent variable (Bardsen, 1989; Pesaran *et al*., 2001). Short run elasticities are simply the estimated coefficients of the first differenced variables in the UECM.

4. ESTIMATION RESULTS

The computed F statistics from the Wald tests for restrictions imposed for each model are shown in Table 2. Note that not all five models are reported for each country in Table 2. This is because the cointegrating relation exists only in a subset of the five models for each CIBS country. As mentioned before, we follow the literature and choose the appropriate import demand model specification based on the presence of a cointegrating relation. Specifically, the traditional model, the Senhadji model, and the Emran & Shilpi model are adopted for China; the disaggregated model and the Emran & Shilpi model are adopted for India; the traditional model and the Emran & Shilpi model are adopted for Brazil; and the national cash flow model and the Emran & Shilpi model are adopted for South Africa. As shown in Table 2, the computed F-statistics for these models lie above the 5% or 10% upper bound and hence we can reject the null hypothesis of no cointegration at these levels.

The next step is to estimate the cointegrating vector and calculate the estimated long-run and short-run elasticities of the import demand function. For estimation of the cointegrating vector, we use the two-step procedure suggested by Pesaran and Shin (1999). Specifically, the specification of the Autoregressive Distributed Lag model is chosen by the Schwarz Bayesian Criterion (SBC) and then estimated by OLS.

20,000 and 40,000 replications respectively. Nayaran and Nayaran (2004) argue that these critical values (CVs) are not suitable for small samples such as, one used in this paper. Given our sample size (96 observations); we use the appropriate CVs from Nayaran (2004).

[3] In the inconclusive case, Kremers *et al*. (1992) and Bannerjee *et al*. (1998), suggest that a significant error term can be used to establish cointegration.

[4] If the ARDL model is chosen by AIC instead, the estimates lack the desirable small sample properties (see Pesaran and Shin, 1999).

Table 2. Bounds Tests for Long-run Relationship for All Reported Models

Country	Models	F Stat [p. value]	Upper Bound (%) (Intercept and no trend)
China	Traditional: Fm (m|y,p)	F(3,20)=4.13 [0.078]	3.80 (10%)
	Senhadji: Fm (m|gex, p)	F(3,20)=5.2513 [0.008]	4.378 (5%)
	Emran & Shilpi: Fm (m|h,p,fer)	F(4,17)=4.0956 [0.017]	4.049 (5%)
India	Disaggregated:Fm (m|ei,fc,ep,p)	F(5,16)=3.3667 [0.089]	3.367 (10%)
	Emran & Shilpi: Fm (m|h,p,fer)	F(4,17)=3.7433 [0.069]	3.574 (10%)
Brazil	Traditional: Fm (m|y,p)	F(3,22)=3.7934 [0.063]	3.800 (10%)
	Emran & Shilpi: Fm (m|h,p,fer)	F(4,19)=3.5911 [0.074]	3.574 (10%)
South	National Cash Flow: Fm (m|ncf,p)	F(3,22)=6.1704 [0.003]	4.378 (5%)
Africa	Emran &Shilpi: Fm (m|h,p,fer)	F(4,17)=5.7072 [0.005]	4.049 (5%)

Note: The definitions of the variables in the models are given in Appendix A.

Tables 3A - 3B, 4A - 4B, 5A - 5B, and 6A - 6B report the empirical results for China, India, Brazil, and South Africa, respectively. Due to the small sample size, a maximum lag structure of three is considered for the UECM estimation. Tables 3A - 6A report long-run elasticities for the determinants of import demand in each relevant model for each country. Tables 3B - 6B report the corresponding UECM model estimation for each adopted model of the CIBS countries.

Table 3A. Estimated Long Run Elasticities of Import Demand of China Using the ARDL Approach (Dependent Variable: $\ln(m_t)$)

	ARDL(), (Observations), [maximum lags], {Income Measure}	Regressors	Coefficient	Standard Error	T-ratio [Prob]
Model 1A	ARDL (2,0,0), (27), [3],	ln*y*	4.21	0.244	17.27 [0.000]
(Traditional	{*y*=Real GDP},	ln*p*	0.09	0.250	0.38 [0.710]
Model)	where *D*=1 for1970-1994,	*D*	0.48	0.126	3.77 [0.001]
	and *D*=0, for 1995-2004	*intercept*	-11.25	1.22	-9.20 [0.000]
Mode1 2A	ARDL(2,0,0), (27), [3]	ln*gex*	2.27	0.297	7.66 [0.000]
(Senhadji	{*gex*=Real GDP -Exports},	ln*p*	0.38	0.528	0.72 [0.479]
Model)	where *D*=1 for1970-1994,	*D*	0.55	0.199	2.74 [0.013]
	and *D*=0, for 1995-2004	*intercept*	-15.77	3.316	-4.75 [0.000]
Model 3A	ARDL (2,0,0), (27), [3]	ln*h*	2.244	0.318	7.06 [0.000]
(Emran &	{*h*=log(home good	ln*p*	0.370	0.542	0.68 [0.504]
Shilpi	consumption)}	*fer*	-0.002	0.006	-0.29 [0.776]
Model)		*intercept*	-15.448	3.550	-4.35 [0.000]

Note: The numbers in the parenthathese following ARDL refer to the optimal number of lags of import volume, income and the relative price, respectively, based on Schwarz Bayesian Criterion.

Table 3B. Error Correction Representations for the Selected ARDL for Import Demand of China (Dependent Variable: $\Delta \ln m_t$)

	ARDL(), (Observations), [maximum lags], {Income Measure}	Regressors	Coefficient	Standard Error	T-ratio [Prob]
Model 1B (Traditional Model)	ARDL (2,0,0), (27), [3], {*y*=Real GDP}, where *D*=1 for1970-1994, and *D*=0, for 1995-2004	Δln*m*(-1)	0.51	0.137	3.70 [0.001]
		Δln*y*	2.99	0.536	5.59 [0.000]
		Δln*p*	0.07	0.179	0.37 [0.712]
		Δ*D*	0.34	0.091	3.73[0.001]
		Δ*intercept*	-8.01	1.588	-5.05[0.000]
		ecm(-1)	-0.71	0.125	-5.70[0.000]
	ecm = ln*m* - 4.21ln*y* - 0.09ln*p* - 0.48 *D* + 11.25*intercept* Serial Correlation (*LM*): $\chi^2(1)$=0.06[0.808]: Adjusted R^2 = 0.58 Functional (Ramsey's Reset): $\chi^2(1)$=0.47[0.495]: AIC= 23.09 Normality: $\chi^2(2)$=1.84[0.399]: SBC=19.20 Heteroscedasticity: $\chi^2(1)$=0.25[0.617]: DW=2.08				
Model 2B (Senhadji Model)	ARDL(2,0,0), (27), [3], {*gex*=Real GDP -Exports}, where *D*=1 for1970-1994, and *D*=0, for 1995-2004	Δln*m*(-1)	0.50	0.176	2.82[0.011]
		Δln*gex*	1.31	0.350	3.75[0.001]
		Δln*p*	0.22	0.309	0.71[0.486]
		Δ*D*	0.32	0.116	2.72[0.014]
		Δ*intercept*	-9.09	2.771	-3.28[0.004]
		ecm(-1)	-0.58	0.147	-3.91[0.001]
	ecm = ln*m* - 2.27ln*gex* - 0.38ln*p* - 0.55*D* + 15.76*intercept* Serial Correlation (*LM*): $\chi^2(1)$=0.64[0.424]: Adjusted R^2 = 0.40 Functional (Ramsey's Reset): $\chi^2(1)$=0.55[0.458]: AIC= 14.81 Normality: $\chi^2(2)$=0.37[0.833]: SBC=11.28 Heteroscedasticity: $\chi^2(1)$=0.01[0.927]: DW=2.11				
Model 3B (Emran & Shilpi Model)	ARDL (2,0,0), (27), [3], {h=log(home good consumption)}	Δln*m*(-1)	0.510	0.187	2.73[.014]
		Δln*h*	1.297	0.362	3.59[.002]
		Δln*p*	0.214	0.318	0.67[.510]
		Δ*fer*	-0.001	0.003	-0.29[.776]
		Δ*intercept*	-8.929	2.896	-3.08[.007]
		ecm(-1)	-0.578	0.151	-3.82[.001]
	ecm = ln*m* - 2.24ln*h* - 0.37 ln*p* + 0.002*fer* - 0.56*D*+ 15.45*intercept* Serial Correlation (*LM*): $\chi^2(1)$= 1.00[.316]: Adjusted R^2 = 0.37 Functional (Ramsey's Reset): $\chi^2(1)$= 0.91 [0.340]: AIC= 13.87 Normality: $\chi^2(2)$=0.39[0.823]: SBC=9.77 Heteroscedasticity: $\chi^2(1)$=0.002[0.964]: DW=2.14				

Note: The numbers in the parenthathese following ARDL refer to the optimal number of lags of import volume, income and the relative price, respectively, based on Schwarz Bayesian Criterion.

As shown in Table 3A, the estimate of income elasticity is well above 1 for China, ranging from 2.24 in the Emran & Shilpi model to 4.21 in the traditional model. The results suggest that China's imports are very responsive to income, regardless of how income is defined. This implies that if China maintains its economic growth, it can be a potentially big buyer in the world market. The price elasticity appears to be positive, though statistically insignificant in all three estimated models. This suggests that China's imports are not responsive to relative price changes. Yet the positive sign is a little counter-intuitive. The explanation may be found through China's trade policies during the period under study. In the late 1980s, China formalized the duty drawback system, which rebates import duties on raw materials, parts and components, and so forth used for export processing, allowing export processing to take place at world prices, free from tariff or domestic pricing distortions (Lardy, 2003). This policy to promote export may have distorted the importing behavior and contributed to the positive price elasticity in the aggregate import demand estimation for China. In addition, in all three models, the dummy D (=1 for the period before1994) is significantly positive, which is consistent with the fact that this trade policy was removed in late 1994.

Table 3B shows the results of the corresponding UECM model estimation for China. The short-run income elasticities are significantly positive, ranging from 1.297 to 2.99, which are still very high, though not as high as those in the long run. The short-run price elasticities are again insignificant. In Table 3B, the diagnostic tests indicate that there are no problems with respect to serial correlation, normality, heteroscedasticity, and functional form. The stability of the coefficients is tested using Cumulative Sum of Recursive Residuals (CUSUM) and Cumulative Sum of Squares of Recursive Residuals (CUSUMSQ) tests. All regression models are found to be stable within the 5% bounds level of significance.[5]

In sum, the results for China suggest that both in the short run and in the long run, China's imports are very responsive to income changes, but not responsive to relative price changes. As long as China's income keeps growing, its imports will keep growing at a faster pace.

Table 4A reports the long-run elasticities for India, and Table 4B reports the corresponding UECM estimations. As shown in Table 4A, the long run income elasticity is estimated to be 2.24 in the Emran & Shilpi model (Model 2A). Compared to earlier studies, this income elasticity estimate is slightly higher, partially because we use more recent data.[6] The result indicates that India's imports are also very responsive to income. A 1% increase income is associated with an over 2% increase in imports in India. In

[5] The CUSUM and CUSUMQ figures for all models and for all CIBS countries are available from the authors upon request.

[6] For India, Dutta and Ahmed (2004) report an income elasticity of 1.48, Goldstein and Khan (1985) report a range of 1.0 to 2.0, and Emran and Shilpi (2010) report an income elasticity of 1.09 for the period 1952-1999.

addition, the variable *fer* has correct negative sign and is statistically significant. This confirms the existence of a binding foreign exchange constraint on aggregate imports before the economic liberalization in India in 1991.

Table 4A. Estimated Long Run Elasticities of Import Demand of India Using the ARDL Approach (Dependent Variable: $\ln(m_t)$)

	ARDL(), (Observations), [maximum lags], {Income Measure}	Regressors	Coefficient	Standard Error	T-ratio [Prob]
Model 1A (Disaggregated Model)	ARDL(3,1,0,0,2)*, (35), [3], {*fc*=real final consumption, *ei*=real expenditures on investment, *ep*=real expenditures on exports}	ln*fc*	0.629	0.611	1.03 [0.314]
		ln*ei*	0.695	0.370	1.88 [0.073]
		ln*ep*	0.259	0.189	1.37 [0.183]
		ln*p*	0.464	0.152	3.06[0.006]
		D	-0.512	0.160	-3.20 [0.004]
		intercept	-3.107	1.840	-1.69[0.105]
Mode1 2A (Emran & Shilpi Model)	ARDL(1,0,2)**, (34), [3] {*h*=log(home good Consumption)}	ln*h*	2.243	0.269	8.35 [0.000]
		ln*p*	0.314	0.175	1.80 [0.084]
		fer	-0.003	0.001	-2.51 [0.018]
		intercept	-7.884	1.424	-5.54[0.000]

Notes: *: The numbers in the parenthathese following ARDL refer to the optimal number of lags of import volume, real final consumption, real expenditures on investment, real expenditures on exports, and the relative price, respectively, based on Schwarz Bayesian Criterion. **: The numbers in the parenthathese following ARDL refer to the optimal number of lags of import volume, income and the relative price, respectively, based on Schwarz Bayesian Criterion.

The disaggregated model (Model 1A in Table 4A) breaks income into three components: final consumption expenditure, investment expenditure, and exports (all in real terms). It turns out that the association between import demand and investment expenditures in India is significantly positive. Specifically, a 1% increase in investment expenditure is associated with a 0.695% increase in import demand. However, the import demand is not responsive to changes in the other two components, final consumption expenditure and exports. This is consistent with another recent study by Goldberg *et al*. (2009). According to their study, over the period 1987-2000, the import growth of final products (i.e., consumer durables and non-durables) was 90%, but the import growth of inputs (i.e., capital goods, basic goods and intermediate products) was 227% for India.

The relative price variable is significantly positive in both import demand models for India. This can arise from the positive association between imports and investment expenditures that we mentioned above. For a fast-growing open economy, certain inputs important for development or exports, such as capital goods or intermediate goods, are increasingly imported regardless of their price increase. If the majority of the import increase lies in the category of inputs (which is indeed the case for India), then it is not surprising to observe the perverse relationship between imports and the relative price.

Table 4B. Error Correction Representations for the Selected ARDL for Import Demand of India (Dependent Variable: $\Delta \ln m_t$)

	ARDL(), (Observations), [maximum lags], {Income Measure}	Regressors	Coefficient	Standard Error	T-ratio [Prob]
Model 1B (Disaggregated Model)	ARDL(3,1,0,0,2)*, (35), [3], {*fc*=real final consumption, *ei*=real expenditures on investment, *ep*=real expenditures on exports}	Δln*m*(-1)	-0.273	0.133	-2.05[0.051]
		Δln*m*(-2)	-0.240	0.119	-2.02[0.054]
		Δln*fc*	-0.186	0.328	-0.57[0.575]
		Δln*ei*	0.309	0.166	1.87[0.074]
		Δln*ep*	0.115	0.091	1.26[0.219]
		Δln*p*	0.395	0.086	4.61[0.000]
		Δln*p*(-1)	0.280	0.088	3.18[0.004]
		Δ*D*	-0.228	0.063	-3.61[0.001]
		Δ*intercept*	-1.383	0.835	-1.66[0.110]
		ecm(-1)	-0.445	0.075	-5.91[0.000]
	ecm = ln*m* – 0.629ln*fc* - 0.695ln*ep* - 0.464ln*p* + 0.512*D* + 3.107*intercept* Serial Correlation (*LM*): $\chi^2(1)$=0.16[0.692]: Adjusted R^2 = 0.73 Functional (Ramsey's Reset): $\chi^2(1)$=1.28[0.259]: AIC= 52.89 Normality: $\chi^2(2)$=0.74[0.692]: SBC=43.56 Heteroscedasticity: $\chi^2(1)$=3.66[0.056]: DW=1.75				
Model 2B (Emran & Shilpi Model)	ARDL(1,0,2)**, (34), [3] {*h*=log(home good Consumption)}	Δln*h*	0.5379	0.1749	3.08[0.005]
		Δln*p*	0.3946	0.0915	4.31[0.000]
		Δln*p*(-1)	0.2145	0.0918	2.34[0.027]
		Δ*fer*	-0.0006	0.0003	-2.16[0.039]
		Δ*intercept*	-1.8911	0.6867	-2.75[0.010]
		ecm(-1)	-0.2399	0.0781	-3.07[0.005]
	ecm = ln*m* - 2.243ln*h* - 0.314ln*p* + 0.002*fer* + 7.884*intercept* Serial Correlation (*LM*): $\chi^2(1)$=0.2495[0.617]: Adjusted R^2 = 0.48 Functional (Ramsey's Reset): $\chi^2(1)$=0.0002[0.988]: AIC= 43.62 Normality: $\chi^2(2)$=1.1131[0.573]: SBC=38.28 Heteroscedasticity: $\chi^2(1)$=1.1047[0.293]: DW=2.13				

Notes: *: The numbers in the parenthathese following ARDL refer to the optimal number of lags of import volume, real final consumption, real expenditures on investment, real expenditures on exports, and the relative price, respectively, based on Schwarz Bayesian Criterion. **: The numbers in the parenthathese following ARDL refer to the optimal number of lags of import volume, income and the relative price, respectively, based on Schwarz Bayesian Criterion.

As shown in Table 4B, the short-run income elasticity is lower compared to the long run, but is also significantly positive. The diagnostic tests indicate that there are no problems with respect to serial correlation, normality, heteroscedasticity, and functional form. Again, the stability of the coefficients is tested using Cumulative Sum of Recursive Residuals (CUSUM) and Cumulative Sum of Squares of Recursive Residuals (CUSUMSQ) tests and found to be stable.

Table 5A and 5B contain the results for Brazil. As shown in Table 5A, the income variable is significant in the traditional model (Model 1A) and shows a long-run income elasticity of 2.13. That is, a 1% increase in income is associated with a 2.13% increase in imports in Brazil for the period 1970-2007. This is higher compared to estimated income elasticities for Brazil in earlier periods using the same model.[7] In the Emran & Shilpi model (Model 2A), long-run income elasticity is significantly positive, yet slightly lower compared to the traditional model. This may be attributed to the fact that the income variable in the Emran & Shilpi model excludes exports and measures only home good consumption. The foreign reserve constraint variable *fer* is of the correct sign, but not statistically significant, possibly implying no binding foreign exchange constraint on aggregate imports. Just like the case of China, the relative price variable is not statistically significant, indicating that Brazil's import demand is not responsive to price increases in the long run.

Table 5A. Estimated Long Run Elasticities of Import Demand of Brazil Using the ARDL Approach (Dependent Variable: $\ln(m_t)$)

	ARDL(), (Observations), [maximum lags], {Income Measure}	Regressors	Coefficient	Standard Error	T-ratio [Prob]
Model 1A (Traditional Model)	ARDL (1,1,1), (33), [3], {*y*=Real GDP}	ln*y*	2.13	0.55	3.91[0.001]
		ln*p*	-0.22	0.73	-0.30[0.770]
		intercept	-27.55	12.16	-2.27[0.320]
Mode1 2A (Emran & Shilpi Model)	ARDL(1,1,1), (35), [3] {*h*=log(home good Consumption)}	ln*h*	1.47	0.68	2.16[0.039]
		ln*p*	-0.46	1.60	-0.29[0.777]
		fer	-0.02	0.02	-0.78[0.444]
		intercept	-12.02	14.95	-0.80[0.428]

Note: The numbers in the parenthathese following ARDL refer to the optimal number of lags of import volume, income and the relative price, respectively, based on Schwarz Bayesian Criterion.

[7] Thirlwall and Hussain (1982) report an income elasticity of 2.05, while Perraton (2003) reports an income elasticity of imports of 1.77 for Brazil.

Table 5B shows the results of the corresponding UECM model estimation for Brazil. As shown in the table, the short-run income elasticities are significantly positive in both models, though lower than their long-run counterparts. Similar to India, the short-run relative price elasticities are significantly positive. This again can arise from the inelastic demand for inputs by Brazil ever since its trade liberalization in the late 1980s. The diagnostics indicate no problems with respect to the model specification.

Table 5B. Error Correction Representations for the Selected ARDL for Import Demand of Brazil (Dependent Variable: $\Delta \ln m_t$)

	ARDL(), (Observations), [maximum lags], {Income Measure}	Regressors	Coefficient	Standard Error	T-ratio [Prob]
Model 1B (Traditional Model)	ARDL (1,1,1), (33), [3], {*y*=Real GDP}	Δln*y*	0.65	0.09	6.93[0.000]
		Δln*p*	0.63	0.18	3.53[0.001]
		Δ*intercept*	-3.65	1.12	-3.26[0.003]
		ecm(-1)	-0.13	0.06	-2.33[0.002]
	ecm = ln*m* - 2.13ln*y* - 0.22ln*p* + 27.54*intercept* Serial Correlation (*LM*): $\chi^2(1)$=1.57[0.211]: Adjusted R^2 = 0.79 Functional (Ramsey's Reset): $\chi^2(1)$=0.11[0.740]: AIC= 26.46 Normality: $\chi^2(2)$=1.57[0.456]: SBC=19.77 Heteroscedasticity: $\chi^2(1)$=0.16[0.692]: DW=1.56				
Model 2B (Emran & Shilpi Model)	ARDL(1,1,1), (35), [3] {*h*=log(home good Consumption)}	Δln*h*	0.564	0.097	5.8208[0.000]
		Δln*p*	0.983	0.170	5.7716[0.000]
		Δ*fer*	-0.002	0.002	-0.99566[0.327]
		Δ*intercept*	-0.972	1.159	-0.83900[0.408]
		ecm(-1)	-0.081	0.061	-1.3328[0.193]
	ecm = ln*m* - 1.47ln*h* + 0.46ln*p* + 0.02*fer* + 12.02*intercept* Serial Correlation (*LM*): $\chi^2(1)$=0.29[0.587]: Adjusted R^2 = 0.77 Functional (Ramsey's Reset): $\chi^2(1)$=2.29[0.130]: AIC= 32.68 Normality: $\chi^2(2)$=1.33[0.513]: SBC=27.23 Heteroscedasticity: $\chi^2(1)$=1.95[0.162]: DW=1.76				

Note: The numbers in the parenthathese following ARDL refer to the optimal number of lags of import volume, income and the relative price, respectively, based on Schwarz Bayesian Criterion.

Table 6A and 6B contain the results for South Africa. Long run relationships are detected in the following models for South Africa: the National Cash Flow (Model 1) and the Emran & Shilpi model (Model2). The long-run income elasticity is 1.36 in the National Cash Flow model (Model 1A) and is 1.45 in the Emran & Shilpi model (Model 2A), indicating that imports in South Africa are also very responsive to changes in income. The estimated long-run income elasticities are higher than those estimated using earlier data for South Africa by other studies.[8]

Table 6A. Estimated Long Run Elasticities of Import Demand of South Africa Using the ARDL Approach (Dependent Variable: $\ln(m_t)$)

	ARDL(), (Observations), [maximum lags], {Income Measure}	Regressors	Coefficient	Standard Error	T-ratio [Prob]
Model 1A (Dynamic Structural Import Demand Model, or National Cash Flow Model)	ARDL (1,0,1), (34), [3], {*ncf*=national cash flow}	ln*ncf*	1.36	0.44	3.09[0.004]
		ln*p*	-0.13	0.33	-0.40[0.692]
		intercept	-8.40	9.00	-0.93[0.358]
Mode1 2A (Emran & Shilpi Model)	ARDL(1,0,2,1), (33), [3] {*h*=log(home good Consumption)}	ln*h*	1.453	0.410	3.54[0.002]
		ln*p*	-0.573	0.508	-1.13[0.269]
		fer	0.001	0.002	0.62[0.539]
		intercept	-10.317	8.431	-1.22[0.232]

Note: The numbers in the parenthathese following ARDL refer to the optimal number of lags of import volume, income and the relative price, respectively, based on Schwarz Bayesian Criterion.

As shown in Table 6A, the coefficient of relative price is insignificant in both models, which is consistent with some earlier studies on South Africa.[9] This indicates that South Africa's aggregate import demand is not so responsive to changes in overall relative prices. Indeed, at the industry level, Gumede (2000) found that an increase in the costs of capital equipment, machinery, transport, or oil does not deter imports in these sectors in South Africa because such imports are necessary for economic growth. Our study confirms the inelasticity of South Africa's import demand with regard to relative prices at the aggregate level, which indicates South Africa's potential to support other countries' exports, particularly in the sectors that are important for its economic growth. The coefficient on *fer* is insignificant, possibly implying no binding foreign exchange constraint on aggregate imports for South Africa.

Table 6B shows the results of the corresponding UECM model estimation for South Africa. Similar to the other three countries, the short-run income elasticities are much lower compared to the long run counterparts, even though they are also significantly positive. The short-run relative price elasticities are also significantly positive, which is

[8] Senhadji (1998), Houthakker and Magee (1969), Bahman-Oskoosee and Niroomand (1998) reported South African income coefficients of 0.33, 1.13, and 0.43 respectively.

[9] See Senhadji (1998), Houthakker and Magee (1969), and Bahman-Oskoosee and Niroomand (1998).

similar to India and Brazil. This reinforces the perverse relationship between imports and the relative price identified in the long run. The diagnostic tests also indicate no problems with respect to the model specification.

Table 6B. Error Correction Representations for the selected ARDL for Import Demand, ARDL (.) Selected based on Schwarz Bayesian Criterion (Dependent Variable: $\Delta \ln m_t$)

	ARDL(), (Observations), [maximum lags], {Income Measure}	Regressors	Coefficient	Standard Error	T-ratio [Prob]
Model 1B (Dynamic Structural Import Demand Model, or National Cash Flow Model)	ARDL (1,0,1), (34), [3], {*ncf*=national cash flow}	Δln*ncf*	0.30	0.07	3.96[0.000]
		Δln*p*	1.17	0.17	6.83[0.000]
		Δ*intercept*	-1.87	1.49	-1.26[0.216]
		ecm(-1)	-0.22	0.09	-2.44[0.021]
	ecm = ln*m* - 13.6ln*ncf* + 0.13ln*p* + 8.40*intercept* Serial Correlation (*LM*): $\chi^2(1)$=0.29[0.587]: Adjusted R^2 = 0.89 Functional (Ramsey's Reset): $\chi^2(1)$=0.09[0.760]: AIC=31.52 Normality: $\chi^2(2)$=0.83[0.660]: SBC=27.71 Heteroscedasticity: $\chi^2(1)$=0.018[0.895]: DW=2.11				
Model 2B (Emran & Shilpi Model)	ARDL(1,0,2,1), (33), [3] {*h*=log(home good Consumption)}	Δln*h*	0.3578	0.0813	4.40[0.000]
		Δln*p*	0.8654	0.1604	5.40[0.000]
		Δln*p*(-1)	0.4661	0.1400	3.33[0.003]
		Δ*fer*	0.0007	0.0003	2.75[0.010]
		Δ*intercept*	-2.5412	1.6426	-1.55[0.133]
		ecm(-1)	-0.2463	0.0739	-3.33[0.002]
	ecm = ln*m* - 1.45ln*h* + 0.57ln*p* - 0.001*fer* + 10.32*intercept* Serial Correlation (*LM*): $\chi^2(1)$=0.15[0.703]: Adjusted R^2 = 0.93 Functional (Ramsey's Reset): $\chi^2(1)$=0.67[0.413]: AIC= 49.17 Normality: $\chi^2(2)$=0.84[0.658]: SBC=37.26 Heteroscedasticity: $\chi^2(1)$=0.04[0.844]: DW=2.60				

Note: The numbers in the parenthathese following ARDL refer to the optimal number of lags of import volume, income and the relative price, respectively, based on Schwarz Bayesian Criterion.

5. CONCLUSION

This paper uses bounds test for the validity of the cointegration or stationarity restriction embodied in five import demand model specifications for CIBS. It applies the two-step procedure appropriate for small sample studies and examines the short-run and long-run income and price elasticities for these countries using recent data.

We find that the four fast-growing developing countries share similar patterns in their import demand behavior. First, based on recent data, we find that long-run income elasticities are higher for all four CIBS countries compared to earlier studies and are almost always statistically significant. Specifically, in the long-run a 1% increase in income is likely to induce a 1.5% - 4.2% increase in imports. Our estimates using most recent data are much higher than either the results in studies about some of the CIBS countries using earlier data or the estimates in other developing countries. On the other hand, our results are more in line with the income elasticity estimates in advanced nations, though slightly higher. This seems to suggest that the CIBS countries indeed have the potential to become southern engines of global economic growth given their fast economic growth and even faster import growth lately. In addition, long-run income elasticities are higher than short-run income elasticities. This suggests that over time import demand in CIBS becomes more and more responsive to changes in income, which implies that continued economic growth and import growth in the CIBS countries are likely to have negative impacts on their balance of payments in the long run. In other words, there is a long-run trade-off between economic growth and balance of payments for CIBS countries. Second, price elasticities are either significantly positive or statistically insignificant for the four countries, unlike the traditional wisdom that the aggregate imports are negatively associated with current relative prices. This finding implies that trade negotiations that aim to lower or remove tariff and nontariff barriers in the CIBS will not necessarily lead to a rise in the flow of their imports. However, certain goods important for development or exports, such as capital goods or intermediate goods, are increasingly imported and consumed by the CIBS even if their relative prices increase.

APPENDIX

A. Variable Definitions and Data Sources

Annual data for the period 1970-2007 are used. All the variables are expressed in real terms. Natural logarithms are taken on all variables. All data were taken from the IMF's International Financial Statistics database plus the World Bank's Development Indicators.

Definitions

m =real imports for goods and services. The nominal series is deflated by the import price index (base year: 2000).

y = real GDP (base year: 2000).

p = relative price term obtained as the ratio of the import price index to GDP deflator (or CPI, depending on data availability).

gex = real (GDP minus exports), as suggested by Senhadji (1998). The GDP deflator is used to derive real value.

ncf = the real national cash flow obtained by calculating real (GDP: investment - government expenditure - exports) (see Xu, 2002; and Tang, 2003).

fc = real final consumption expenditure obtained by calculating real (Household Expenditures + General Government Expenditures).

ei = real expenditures on investment on goods and services.

ep = real expenditures on exports goods and services.

fer = (real domestic expenditure/real foreign exchange availability)**D*, where *D* takes a value of 1 for 1970-1994 and zero otherwise for China; *D* takes a value of 1 for 1970-1988 and zero otherwise for Brazil; *D* takes a value of 1 for 1970-1991 and zero otherwise for India; *D* takes a value of 1 for 1970-1990 and zero otherwise for South Africa.

h = (*y-pm*) is real home good consumption, calculated as (Real GDP - Exports) (Emran & Shilpi, 2008).

B. The Error Correction Representation for the First Four Models

B1. The Traditional Model (corresponding to Equation (1) in the paper)

$$\Delta \ln m_t = \alpha_0 + \sum_{i=0}^{k} \alpha_{1i} \Delta \ln y_{t-i} + \sum_{i=0}^{k} \alpha_{2i} \Delta \ln p_{t-i} + \sum_{i=1}^{l} \alpha_{3i} \Delta \ln m_{t-i} \\ + \alpha_4 ECM_{t-1} + \alpha_5 \Delta D_t + \alpha_6 \Delta intercept_t + e_t ,$$

where $k = 0, 1, 2, \ldots, s$, $l = 0, 1, 2, \ldots, s$, and s is the maximum lag;

$$ECM_{t-1} = \ln m_{t-1} - \varphi_1 \ln y_{t-1} - \varphi_2 \ln p_{t-1} - \varphi_3 D_{t-1} - \varphi_4 intercept_{t-1} .$$

B2. The Senhadji Model (corresponding to Equation (2) in the paper)

$$\Delta \ln m_t = \alpha_0 + \sum_{i=0}^{k} \alpha_{1i} \Delta \ln gex_{t-i} + \sum_{i=0}^{k} \alpha_{2i} \Delta \ln p_{t-i} + \sum_{i=1}^{l} \alpha_{3i} \Delta \ln m_{t-i} \\ + \alpha_4 ECM_{t-1} + \alpha_5 \Delta D_t + \alpha_6 \Delta intercept_t + e_t ,$$

where $k = 0, 1, 2, \ldots, s$, $l = 0, 1, 2, \ldots, s$, and s is the maximum lag;

$$ECM_{t-1} = \ln m_{t-1} - \varphi_1 \ln gex_{t-1} - \varphi_2 \ln p_{t-1} - \varphi_3 D_{t-1} - \varphi_4 intercept_{t-1} .$$

B3. The Disaggregated Model (corresponding to Equation (3) in the paper)

$$\Delta \ln m_t = \alpha_0 + \sum_{i=0}^{k} \alpha_{1i} \Delta \ln fc_{t-i} + \sum_{i=0}^{k} \alpha_{2i} \Delta \ln ei_{t-i} + \sum_{i=0}^{k} \alpha_{3i} \Delta \ln ep_{t-i} + \sum_{i=0}^{k} \alpha_{4i} \Delta \ln p_{t-i}$$
$$+ \sum_{i=1}^{l} \alpha_{5i} \Delta \ln m_{t-i} + \alpha_6 ECM_{t-1} + \alpha_6 \Delta D_t + \alpha_7 \Delta intercept_t + e_t ,$$

where $k = 0, 1, 2, \ldots, s$, $l = 0, 1, 2, \ldots, s$, and s is the maximum lag;

$$ECM_{t-1} = \ln m_{t-1} - \varphi_1 \ln fc_{t-1} - \varphi_2 \ln ei_{t-1} - \varphi_3 \ln ep_{t-1} - \varphi_4 \ln p_{t-1}$$
$$- \varphi_5 D_{t-1} - \varphi_6 \, intercept_{t-1} .$$

B4. The Dynamic Structural Import Demand Model, or National Cash Flow Model (corresponding to Equation (4) in the paper)

$$\Delta \ln m_t = \alpha_0 + \sum_{i=0}^{k} \alpha_{1i} \Delta \ln ncf_{t-i} + \sum_{i=0}^{k} \alpha_{2i} \Delta \ln p_{t-i} + \sum_{i=1}^{l} \alpha_{3i} \Delta \ln m_{t-i}$$
$$+ \alpha_4 ECM_{t-1} + \alpha_5 \Delta D_t + \alpha_6 \Delta intercept_t + e_t ,$$

where $k = 0, 1, 2, \ldots, s$, $l = 0, 1, 2, \ldots, s$, and s is the maximum lag;

$$ECM_{t-1} = \ln m_{t-1} - \varphi_1 \ln ncf_{t-1} - \varphi_2 \ln p_{t-1} - \varphi_3 D_{t-1} - \varphi_4 \, intercept_{t-1} .$$

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Mailing Address: Yan Zhou, Department of Economics, California State University, Sacramento (CSUS), USA. Fax: 916 278 5768. E-mail:yzhou@csus.edu.

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