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Estimating the import demand function in the autoregressive distributed lag framework: The case of China

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Abstract

This paper uses the concept of cointegration for empirically analyzing the long-run relationship of China's import demand function. The analysis employs the annual data for the sample period from 1978 to 2009. The purpose of this study is to investigate and explain China's import demand functions and provide a more in-depth analysis of China's import behavior. The autoregressive distributed lag (ARDL) and dynamic ordinary least square (DOLS) techniques were used for estimating the long-run coefficients of price and income elasticities. The empirical results from ARDL bound testing approach and Johansen's method of cointegration provide strong evidence of the existence of a long-run stable relationship among the variables included both in the traditional model and the disaggregated expenditure model of import demand. In addition, the disaggregated import demand model estimated in this paper provides a complete description of the determinants of China's imports, and offers empirical results that are significantly different from those obtained in existing studies (Tang, 2003). This is an important finding for resolving the issue of trade imbalance from the perspective of China's policy formulation.

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1. Introduction

This paper aims to empirically analyze the long-run relationship of China's import demand function. Trade has been at the core of China's development strategy since the Communist Party Central Committee's decision in December 1978 to adopt Deng Xiaoping's program of economic reform. Over the last two decades, especially with the implementation of the world trade organization's (WTO) rules and substantial reduction in trade restrictions, the volume of China's trade has rapidly increased and it has run large trade surpluses. Its foreign reserves swelled from $21 billion in 1992 (5% of its annual GDP) to $2.4 trillion in June 2009 (approximately 50% of its GDP). The effect of this astronomical accumulation of reserves has been a source of growing public attention in the context of the debate on global imbalances. China is being criticized for the considerable trade surplus held by them and this voice gained momentum during the global crisis. Determining the manner in which China's trade imbalance problem must be resolved has become a rather difficult issue for economists across the world.

The purpose of this study is to estimate the import demand function of the Chinese economy in order to explain and investigate China's import demand functions and provide a more in-depth analysis of China's import behavior. In particular, this paper has a twofold purpose. The first is to provide estimates of the income and price elasticities of import demand using the autoregressive distributed lag (ARDL) model developed by Pesaran and Shin (1999) and the dynamic ordinary least squares (DOLS) estimator developed by Stock and Watson (1993). The second is to compare the estimates obtained by using the traditional models with those obtained by using the disaggregated expenditure models. Moreover, we have proposed pertinent recommendations for resolving the imbalance in trade in China in accordance with the results.

The remainder of this paper is organized in the following manner: Section 2 provides an overview of the import demand literatures that estimate the import determinants using certain estimation methods, which are subsequently employed in our empirical analysis. Section 3 explains the import demand model and empirical approach. Section 4 reports the empirical findings and their interpretations. Finally, the conclusion provides a summary of the empirical findings and their policy implications for China.

2. Literature reviews

Over the recent years, owing to increasing globalization, the interdependence among countries has increased. Every country wants to achieve a rapid pace of economic development by maximizing their benefits from international trade and using modern techniques in their production processes. The relationships between imports and macro-components of particular countries have become the basis for recent directions of research. Santos-Paulino (2002) highlighted that the empirical investigation of import demand functions is one of the most researched areas in international economics. The import demand specification is crucial for meaningful import forecasts, international trade planning, and policy formulation.

A plethora of studies exists on the determinants of import demand models that have been conducted over several decades. A few of the important studies are Stern et al. (1979), Gafar (1995), Carone (1996), Mah (2000), and Hamori and Matsubayashi (2001). Carporale and Chui (1999) is the seminal work on the analysis of the import demand function. They estimated income and relative price elasticity of trade in a cointegration framework for 21 countries using annual data for the period from 1960 to 1992. The ARDL and DOLS estimates confirm the existence of a cointegration relationship between growth rates and income elasticity estimates.

Now we review other empirical studies on the aggregate import demand function that have employed the bounds testing approach. Tang (2003) examined the long-run relationship of China's aggregate import demand function for the period 1970-1999 using the bounds testing approach. Several definitions have been employed in order to represent domestic demand-GDP, GDP minus exports, national cash flow, and final expenditure components. Tang (2003) validated a long-run equilibrium relationship between these measures of domestic demand and China's import demand. He found that expenditure on exports have the biggest correlation with imports (0.51), followed by investment expenditure (0.40) and final consumption expenditure (0.17); China's import demand function was found to be inelastic (–0.6) with respect to relative prices in the long run.

Narayan and Narayan (2005) found a long-run relationship between import volumes, domestic incomes, and relative prices for Fiji in a cointegration framework only when import demand is the dependent variable. They used the bounds testing approach in order to investigate the long-run as well as the short-run elasticities of Fiji's import demand. Their results confirm that although domestic income has a positive impact on import volumes, an increase in relative prices reduce import volumes. According to them, in the long run, growth in income has a significant and elastic impact on import demand.

Emran and Shilp (2010) used a structural econometric model of aggregate imports for India and Sri Lanka. In order to estimate the model, they employed the time series data for the period 1952-1999 for India and 1960-1995 for Sri Lanka. ARDL, DOLS, and FM-AADL （Fully Modified and Augmented-by-Leads Autoregressive Distributed Lag）techniques were used for estimating the log-run coefficients of price and income elasticities. The empirical results from both ARDL and Johannes' method provided strong evidence regarding the existence of a long-run relationship among the variables included in the long-run import demand models. The mean income elasticity was 1.07, which indicated a long-run unitary income elasticity. The mean of price elasticity was –0.72, and foreign exchange availability variable was highly significant with correct positive signs for both the countries.

3. Model specification and data

The import demand model adopted here is derived within the framework of the imperfect substitution theory. It typically uses a Marshallian demand function that relates the total quantity of imports demanded by a country to its real expenditure or income (or another scale variable that captures domestic demand conditions) and to the price of imports and domestic substitutes measured in the same currency.[1] According to the conventional demand theory, the demand for real imports is a function of domestic income or GDP and relative price (import price index deflated by an index of domestic prices). Microeconomic theory regards demand functions to be homogeneous of degree zero in terms of prices and money income (Deaton and Mullbauer, 1980). In accordance with the studies by Khan and Ross (1977), Salas (1982), and Gafar (1995), the traditional import demand model may be expressed in the following manner:[2]

$$\ln M_t = \alpha_0 + \alpha_1 \ln Y_t + \alpha_2 \ln RP_t + \varepsilon_t, \qquad (1)$$

where ln is the natural logarithmic form and ε_t is the error term. M_t denotes the volume of imports at time *t*, Y_t denotes real income at time *t*, and RP_t denotes the relative price (the import price index deflated by a GDP deflator) at time *t*. Generally, the hypothesized values of the coefficients of the explanatory variables are $\alpha_1 > 0$ and $\alpha_2 < 0$, which represent the income and price elasticities respectively of import demand.

The composition of expenditure is also important in that the import content of the different components of expenditure differs (Giovannetti, 1989; Davies, 1990). Indeed, if the composition of the final demand changes, the aggregate import propensity will change even if the disaggregated marginal propensities remain unchanged. Giovannetti (1989) argued that if the different components of total expenditure have different import contents, the use of a single demand variable (e.g., GDP) in an aggregate import demand function would lead to aggregation bias. Moreover, an aggregate import equation that embodies disaggregate demand variables among the regressors has a better fit and forecast than a standard specification with a single demand variable.[3] According to these studies, decomposing GDP into the following three broad categories is an alternative to the traditional approach: consumption expenditure by private and public sectors, investment expenditure (public and private), and net exports. The preference for an import demand model with disaggregated expenditure components not only eliminates aggregation bias but also can test out the impact on imports from different

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[1] For details, see, Carone (1996, p.3).

[2] A time trend is included in Tang's (2003) model in order to represent a change in consumer preferences; however, the cointegration test indicates a linear deterministic trend in the data only in our paper.

[3] The link between imports and the macro-components of final expenditure has become very popular in recent research (e.g., Abbott and Seddighi, 1996; Tang, 2003).

components of GDP.

The disaggregated expenditure model of import demand is expressed in the following manner:

$$\ln M_t = \beta_0 + \beta_1 \ln FC_t + \beta_2 \ln I_t + \beta_3 \ln EX_t + \beta_4 \ln RP_t + u_t, \quad (2)$$

where FC_t is the final consumption expenditure at time *t*, which is the sum of the real private and public consumption expenditures; I_t is the real expenditure on investment goods at time *t*; and EX_t is the real expenditure on exports at time *t*. The definitions of the other variables are the same, as defined previously. According to the equation (2), the parameter must satisfy the following sign restrictions: $\beta_1 > 0$, $\beta_2 > 0$ and $\beta_3 > 0$, $\beta_4 < 0$. With respect to details on sign restrictions, see Bahmani-Oskooee and Niroomand (1998).

We use the annual data over the period from 1978 to 2009 for empirical analysis. Each data is taken from the World Bank (2010) database.

Testing for the existence of a relationship in levels between variables is essential to empirical economics, and such testing has received considerable attention over the past decade. Generally, this analysis is based on the use of cointegration techniques. In order to test for the existence and the number of long-run relationship(s), we employ the system-based reduced rank regression approach by Johansen (1991, 1995), the bounds *F*-test proposed by Pesaran et al. (2001), and the bounds *t*-test based on Banerjee et al.'s (1998) cointegration test. We employed the bound test (Pesaran et al., 2001) for cointegration analysis because it has the following advantages: First, this test can be used irrespective of whether the regressors are purely I(0), purely I(1), or mutually cointegrated. Second, the approach of the test is such that the model takes a sufficient number of lags in order to reduce the intensity of the serial correlation of residuals in a general-to-specific modeling framework. Third, it assumes that all variables in the model are endogenous. Finally, a dynamic error correction model may be derived by making a simple linear transformation in the ARDL model (see, for example, Pesaran and Pesaran, 1997; Pesaran and Shin, 1999).

There is an important difference between our study and the extant literature on China's import demand.[4] We use the following two alternative estimators for estimating the cointegrating vector: ARDL and DOLS. We use alternative methods for gauging the sensitivity of the results to different estimation techniques. The choice of the ARDL is motivated primarily by the recent evidence that it possesses desirable small sample properties and can effectively correct for possible endogeneity of explanatory variables (see, for example, Pesaran and Shin, 1999; Panopoulou and Pittis, 2004; Caporale and Pittis, 2004). We include the estimates from DOLS, because it is among the most widely used estimators of cointegrating vectors in applied literature. However, Caporale and Pittis (2004) indicated that

[4] With respect to the empirical investigation of China's import demand function, see Moazzami and Wong (1988), Senhadji (1998), and Tang (2003).

the decision regarding the estimation method is more crucial than the actual data employed; while widely used estimators, including OLS and DOLS, have the worst performance in small samples, ARDL (and FM-AADL) does not suffer from the problem that the standard asymptotic critical values are highly misleading in small to moderate samples.

4. Empirical Results

We must investigate the existence of a long-run import demand relationship before interpreting the estimated import demand equations in Table 1. In order to investigate this, the bounds tests suggested by Persaran and Shin (1999) and Banarjee et al. (1998), and the rank tests for cointegration due to Johansen (1995) were employed. The specifications of the ARDL and VAR models are selected by the Schwartz Bayesian Information Criterion (henceforth SBIC) and then estimated by OLS.[5] In addition, the unit root tests indicate that the relevant variables of the import model are non-stationary and integrated of order one (Appendix A).[6]

4.1 Cointegration

The Johansen's maximum eigenvalue and trace tests indicate that there is one cointegrating vector in both specifications (1) and (2) (Appendix B).[7] The null hypothesis of no cointegration can be rejected at the 1% significance level in both the cases. Table 1 presents the results of a bounds test for cointegration. The cointegration test under the bounds framework involves the comparison of the *F*-statistics against the critical values, which is generated for specific sample sizes. This is also an improvement over the existing studies (Tang, 2003) that use the bounds testing approach. The results of the bounds *F* tests indicate that the null hypothesis of no cointegration can be rejected at the 1% significance level in both the models. The results from the bounds *t*-tests are similar.[8] All the three approaches provide similar results on the long-run correlation, thereby demonstrating that China's import demand and its determinants are cointegrated for the sample period.

4.2 Long-run elasticities

Since there is strong evidence of the existence of a long-run relationship among the

[5] The ARDL model selected by SBIC is more effective than alternatives like the Akaike information criterion (AIC). For more details on this, see Pesaran and Shin (1999).
[6] The result of the unit root test for the "GDP minus exports" variable, which is used as a proxy of GDP in Senhadji (1998) and Emran and Shilp (2010) is I(2) (not reported).
[7] Owing to the use of limited annual observations in this study, three lag and one lag structures of VAR were selected for specifications (1) and (2), respectively.
[8] This was because those critical values of *t*-statistics were extracted from Pesaran et al. (2001), which were based on the sample sizes of 500 and 1,000 observations; the significance levels of the t-test are higher than the corresponding significance levels of *F*-test.

variables included in the long-run import demand model, we estimate the long-run cointegration relationship (long-run coefficients) for imports using the ARLD and DOLS single equation estimation methods. Table 2 presents the long-run results.

For the income coefficient, the magnitude of DOLS estimates is lower than the estimates from ARDL in two cases. However, contrary to the income coefficients, the DOLS estimates of the relative price coefficient are higher as compared to those from ARDL. This finding is inconsistent with Emran and Shilp (2010), wherein the DOLS estimates are the lowest among the estimates for both income coefficients and relative price. In addition, the estimated coefficients are highly statistically significant except ln*FC* by both the estimates (*t*-statistic is 0.441 by DOLS and 0.805 by ARDL) in model 2.

When the import equation was estimated by ARDL, the coefficients were found to be completely different from those in Tang (2003). For example, in the traditional model, the estimated income elasticity was 2.661; this is considerably larger than Tang's (2003) estimated income elasticity of 0.73.[9] Similarly, the coefficients of the decomposed GDP are slightly higher as compared to those in Tang (2003), that is, 0.51, 0.17, and 0.4 for exports, consumption, and investment, respectively. A possible explanation for this is that over the period 1970–1999, China experienced historical trade deficits.[10]

We conducted a number of diagnostic tests including tests of autocorrelation, normality, and heteroscedasticity. The estimated residuals did not provide any significant evidence of serial correlation, nonnormality (Jarque-Bera test), or heteroscedasticity in the error term.[11] Meanwhile, the adjusted R-squared of approximately 0.995 indicated that 99.5% of the variation in import demand was explained by the variables in the models. In addition to this, the estimated coefficients for the relative price and activity variables (measured by income, consumption, investment, or exports), satisfy the theoretical sign restrictions for both models regardless of the estimation technique considered.

4.3 Stability of the estimated parameters

Model stability is necessary for prediction and econometric inference. We test for the stability of estimated parameters by using the cumulative sum of recursive residual (CUSUM) and CUSUM of square (CUSUMSQ) tests.[12] For the sake of brevity, we have only discussed the results for the ARDL model. The results of the CUSUM and CUSUMSQ tests are reported in Figures 1 and 2 for models 1 and 2, respectively. Neither tests provided any evidence of instability in the estimates at the 5 percent significance level for conventional specification. Meanwhile, for the decomposed GDP specification, although the ARDL estimates passed the

[9] The result is at odds with the conventional wisdom of long-run unitary income elasticity.
[10] See Tang (2003, p. 143)
[11] The homoscedasticity test of DOLS results may be rejected for model 2.
[12] The CUSUM test detects the systematic changes in the regression coefficients, whereas the CUSUMSQ (CUSUM of squares) test is useful for capturing the sudden departures from the constancy of regression coefficients.

CUSUM test, the CUSUMSQ test provided some evidence of mild instability.[13]

5. Conclusions and policy implications

This paper empirically analyzed the import demand function of China based on the concept of cointegration for the period 1978–2009. We contribute to the extant empirical literature by employing two different definitions of domestic real activity and two alternative estimators (ARDL model and DOLS estimator) for explaining the behavior of import demand in China. Moreover, irrespective of the estimation technique, we found strong evidence of a cointegration relationship between the income and relative price variables in both the models. However, the estimated coefficients for income and relative price variables were found to be rather different when different estimation techniques were employed. We also found that decomposing final expenditure explains China's import demand more effectively. Since the elasticity of imports significantly differs with respect to different macro expenditure components, the different macro components of expenditure have different import contents.

Pertinent policy implications may be derived on the basis of the empirical estimates. First, it is evident that prices play an important role in the determination of imports. The estimated long-run elasticity is inelastic and approximately within the range of –0.5 to –1. Similar to Tang (2003), it appears that China cannot depend on using its exchange rate policies to correct the balance of trade problem.[14] However, the long-run price elasticity is statistically significant, suggesting that if the growth in inflation in China is related to the import price, then China's import bill will increase.

Second, in model 2, the estimated coefficients of consumption were found to be statistically irrelevant for both ARDL and DOLS. This implies that increase in import does not benefit the living conditions of Chinese people. This is not a surprising result because China's currency "manipulation" is fare "effective."

Finally, contrary to Tang (2003), the growth in income has a significant and elastic impact on import demand in the long run. In addition, the estimated coefficient of investment was also found to be elastic, and larger than that of exports in the ARDL model. This indicates that higher growth especially in investment will induce higher demand for imports. Since the demand for imports rises when the level of investment increases, the balance-of-payments is expected to deteriorate if China's growth in imports exceeds their growth in exports.

[13] However, this evidence of mild instability is not corroborated by the results obtained from recursive estimations. The results are available from the authors upon request.

[14] The average long-run price elasticity of Tang (2003) is –0.51; according to Heien (1968), "exchange rate polices, which directly influence the relative price, will have little impact on China's import demand as well and trade balances."

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Table 1.Testing for the existence of a long-run relationship in autoregressive distributed lag (ARDL) models

Model 1						
Critical value bounds (case III: unrestricted intercept and no trend)						
k	90 percent level		95 percent level		99 percent level	
2	*I*(0)	*I*(1)	*I*(0)	*I*(1)	*I*(0)	*I*(1)
F-statistic	2.915	3.695	3.538	4.428	5.155	6.265
t-statistic	–2.57	–3.21	–2.86	–3.53	–3.43	–4.10
Calculated		*F*-statistic		*t*-statistic		
		6.392***		**–4.006****		
Model 2						
Critical value bounds (case III: unrestricted intercept and no trend)						
k	90 percent level		95 percent level		99 percent level	
4	*I*(0)	I(1)	*I*(0)	I(1)	*I*(0)	*I*(1)
F-statistic	2.525	3.560	3.058	4.223	4.280	5.840
t-statistic	–2.57	–3.66	–2.86	–3.99	–3.43	–4.60
Calculated		*F*-statistic		*t*-statistic		
		8.807***		**–3.709***		

Notes:

The critical values of the *F*-statistic have been extracted from Narayan (2004a, b, 2005a) and critical values of *t*-statistic are extracted from Pesaran (2001).

The optimal lag length for ARDL models that start at three lags were selected by the Schwarz-Bayesian Information Criterion (SBIC).

k is the number of regressors.

***, **, and * denote values significant at the 1 percent, 5 percent, and 10 percent levels.

Table 2. Estimates of the long-run import demand function

	Model 1		Model 2	
	DOLS	ARDL	DOLS	ARDL
ln*RP*	–0.916	–0.478	–0.397	–0.340
	(–13.811)	(–2.251)	(–2.987)	(–3.122)
ln*Y*	1.524	2.661		
	(51.588)	(4.179)		
ln*EX*			0.359	0.854
			(4.027)	(6.001)
ln*FC*			0.132	0.203
			(0.441)	(0.805)
ln*I*			0.689	1.235
			(2.656)	(4.476)
Intercept	–16.655	–9.270	–5.437	–4.064
	(20.063)	(–4.472)	(–2.617)	(–2.397)
Diagnostics				
$\bar{R}^2$	0.996	0.994	0.995	0.996
σ	0.066	0.088	0.083	0.071
Serial Correlation (F)	0.035	0.211	0.027	0.321
	[0.854]	[0.651]	[0.872]	[0.577]
Normality (χ^2)	0.387	1.130	1.441	0.383
	[0.824]	[0.568]	[0.486]	[0.826]
Heteroscedasticity (F)	0.563	0.165	4.409	0.996
	[0.809]	[0.687]	[0.004]	[0.499]

Note:

t-statistics and *P*-values are indicated in parentheses and brackets, respectively.

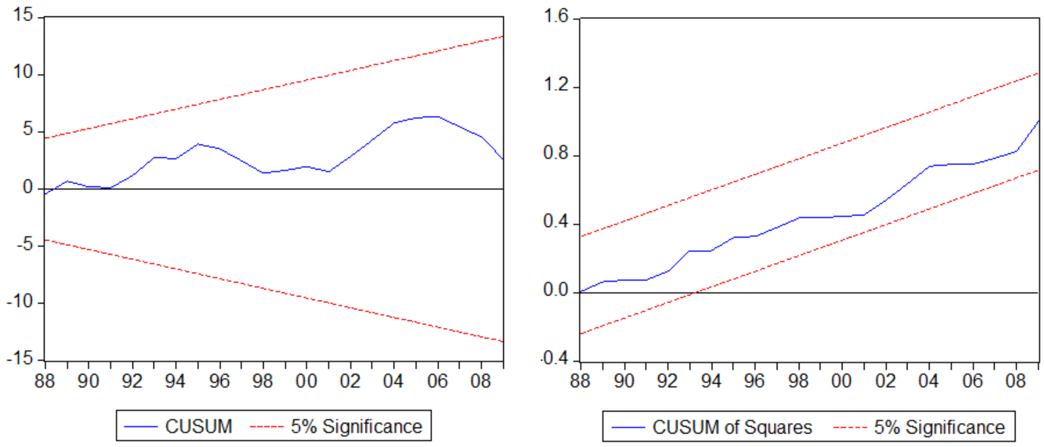

Figure 1. Cumulative sum of recursive residual (CUSUM) and CUSUM of square (CUSUMSQ) tests (Model 1)

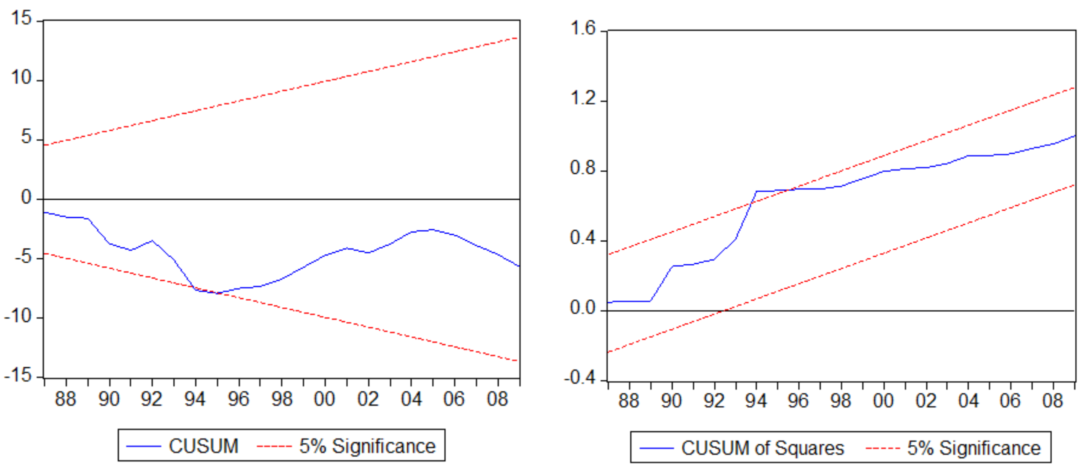

Figure 2. Cumulative sum of recursive residual (CUSUM) and CUSUM of square (CUSUMSQ) tests (Model 2)

Appendix A. Unit root test

The augmented Dickey-Fuller (ADF) (Dickey and Fuller, 1979, 1981) unit root test has been employed for each variable. The unit root test indicates that all the series employed are non-stationary and integrated of order one. The results of the ADF test are presented in the following table:

1978–2009		
Series	Level	First difference
ln*M*	0.574	–4.531***
ln*Y*	–0.270	–4.090***
ln*RP*	–2.274	–3.064**
ln*FC*	–1.251	–4.373***
ln*EX*	1.675	–3.072**
ln*I*	0.573	–3.426**

Note:

The auxiliary regression is run only with an intercept, both for the level and first differenced series.

The maximum number of lags is three, and the order of the lag length is selected by the Schwarz-Bayesian Information Criterion (SBIC).

The critical values are extracted from MacKinnon (1996).

** and *** indicate that the null hypothesis is rejected at 5% and 1% significance levels.

Appendix B. Test for existence of cointegrating vectors using Johansen approach

Test Type	Trace/ Max-Eig Statistic	5% Critical Value	Prob.	Hypothesized Number of CEs
Model 1 :	ln*M* ln*Y* ln*RP* (Conventional approach)			
Trace	50.472	29.797	0.000	none *
	9.839	15.495	0.293	at most 1
Max-Eig	40.633	21.132	0.000	none *
	7.355	14.265	0.448	at most 1
Model 2 :	ln*M* ln*EX* ln*FC* ln*I* ln*RP* (Decomposed GDP)			
Trace	79.813	69.819	0.006	none *
	39.289	47.856	0.249	at most 1
Max-Eig	40.525	33.877	0.007	none *
	19.572	27.584	0.372	at most 1

Note:

The test assumes a linear deterministic trend in the data.

The lag length selected by Schwarz-Bayesian Information Criterion (SBIC) for the VAR analysis is three for model 1 and one for model 2.

Prob indicates the MacKinnon-Haug-Michelis (1999) *P*-values.

* denote the significance at the 1% level.