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AGGREGATE IMPORT DEMAND FUNCTION FOR EIGHTEEN OIC COUNTRIES: A COINTEGRATION ANALYSIS

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ABSTRACT

This is an empirical investigation on the long-run relation of the aggregate import demand function for Organization of Islamic Conferences (OIC) economies, using the recently developed import demand equation that is derived from the dynamic-optimizing intertemporal approach (Xu, 2002). We include only 18 of the 27 OIC founding countries for analysis due to data unavailability for a sufficient sample span. The results of the bounds test (Pesaran, Shin and Smith, 2001) indicate that the volume of import demanded, domestic real activity, and relative prices for 10 of the 18 sample countries are cointegrated. Overall, the estimated price and domestic activity variables are inelastic in the long-run. Some policy implications on trade balance have been drawn in this study.

JEL classification: C51, F14

Key words: Bounds testing approach; Import demand function; Organization of Islamic Conferences

1. INTRODUCTION

The main objective of this study is to investigate the long-run relationship between import demand and its determinants for 18 OIC[1] countries using the import demand specification proposed by Xu (2002). The determinants are the real activity variable and relative prices. The bounds

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test procedure (Pesaran, Shin and Smith, 2001), which is based on the estimation of an unrestricted error correction model (UECM) or conditional ARDL (autoregressive distributed lag)-ECM is employed to examine the presence of a level relationship between the examined variables. The estimated conditional ARDL-ECM is also used to capture the long-run relation of import demand to its determinants. The evidence presented will provide an additional dimension to the empirical literature.

In this study, we focus on the long-run relationship of the aggregate import demand function instead of in the short-run, mainly considering the issues of policy implication. Reinhart (1995) documented that the relative price variable played a significant role in the determination of trade flows, buttressing policies of devaluation as a way to correct trade imbalances. Further, Reinhart (1995) added that the relative price estimate was based on '*static*' or '*long-run*' specifications of import demand or export supplies. The author (Reinhart, 1995) estimated the long-run import demand, and export demand functions for developing countries using the cointegration technique. Another example is Bahmani-Oskooee and Niroomand (1998). For a single aggregate import demand function, Heien (1968) argued that ". . . for any country a value of the price elasticity (demand for imports) between –0.5 and –1.0 is necessary to insure success of exchange depreciation." Gafar (1988) estimated an import demand for Trinidad and Tobago. The price elasticity obtained from a single import equation is -0.5316, the weighted average price elasticity is -0.5665, and aggregate price elasticity using equation 6 (as in Gafar, 1988, 309) is -0.5383 (Gafar, 1988, 311). Finally, Gafar (1988) has concluded that the estimated price elasticities have fallen within the range suggested by Heien, concluding that exchange rate policies could be used to correct the balance of payments disequilibrium.

The reason for choosing the OIC member countries lies basically in the presence of their trade deficits. By observing the period 1990-2000, most of the 18 selected OIC member countries experienced unfavorable trade deficits. These countries include Bangladesh, Chad, Egypt, Guinea, Jordan, Mali, Mauritania, Morocco, Niger, Pakistan, Senegal, Tunisia and Turkey. However, trade deficits are also experienced by Algeria in 1990, 1993-1995 and 1998; Indonesia in 1995-1997; Iran in 1990-1993 and 1998; Malaysia in 1991 and 1993-1995; and Syria in 1991-1998 (World Bank, various issues). Another motivation for this study is the importance of imports to the individual countries as indicated by the percentage of imports in relation to national income (see Table 1). The share of imports to Gross Domestic Product (in 1999) ranged from 18 percent (Iran) to 97 percent (Malaysia). The average growth of exchange rate (in local currency per US$) in 1990-1999 varied from 1.4 percent (Morocco) to 72.8 percent (Turkey). A positive growth rate indicates devaluation of domestic currency. Here, exchange rate plays a significant role in determining trade flows. Devaluation would affect domestic prices through its impact on import prices, affecting costs of production directly through an increase in the prices of imported industrial supplies, and indirectly through an increase in wage-rates claimed and paid in compensation for the rise in the cost of living. The end result of devaluation, therefore, depends on the net outcome following the fall in the external price and the increase in the domestic cost of exports (Briguglio, 1989, 327). As noted by Bahmani-Oskooee (1998, 90) a decrease in nominal effective exchange rate (units of foreign currency per unit of domestic currency), which indicates a depreciation of domestic currency, is expected to discourage imports and encourage exports. The analysis of import demand is meaningful since the import level is reasonably high and exchange rate is sufficiently flexible (except Syria). Syria has fixed the exchange rate to 11.225 local currency/US$ since 1988. Their import structure is highly dominated by manufactured goods with the share of total manufactured imports in the range of 41 percent (in 1998) for Niger to 85 percent (in 1999) for Malaysia (Table 1).

Other than that, some macroeconomic characteristics of these countries are illustrated in Table 1. Almost all of the examined OIC countries are in the low and low middle income categories, except Malaysia (upper middle income). Secondly, in terms of economic structure, the agricultural sector is the major contributor to Gross Domestic Product (GDP) over these countries compared to the manufacturing sector. The countries with the share of agriculture as percentage of GDP (in 1999) under the range of 30 percent and 47 percent per annum are Chad, Mali, and Niger. And the lowest is 2 percent for Jordan. In addition, the value-added of manufacturing as a percentage of GDP (in 1999) is found to be higher than agriculture in Egypt, Indonesia, Jordan, Malaysia, Syria, and Tunisia. Here, the macroeconomic performance of the selected OIC countries is also briefly illustrated. The average growth of real GDP (1990-1999) has ranged between 1.6 percent (Algeria) to 6.3 percent (Malaysia) per

TABLE 1
Economic Characteristics of the Selected OIC Countries

Country	% of Imports to GDP in 1999	Value added as % of GDP in 1999		Average Growth (%) for the Period 1990-1999				
		Agriculture	Manufacturing	GDP	Exports	Imports	Inflation (CPI)	Exchange Rate[1]
Algeria[b]	24 [67]	13	11	1.6	2.2	2.8	21.4	21.7
Bangladesh[a]	19 [69]	21	17	4.8	13.2	9.2	4.33	4.2
Chad[a]	30 [56]	38	11	2.3	5.0	4.9	6.5	6.6
Egypt[b]	24 [59]	17	27	4.4	3.1	2.8	9.9	13.6
Guinea[a]	26 [64]	23	4	4.2	4.7	2.1	N.A.	9.05
Indonesia[a]	27 [58]	20	25	4.7	9.2	4.8	16.0	14.9
Iran[b]	18 [73]	N.A.	N.A.	3.4	0.2	-5.5	25.5	31.9
Jordan[b]	62 [62]	2	15	4.8	7.4	-1.3	4.3	2.1
Malaysia[c]	97 [85]	14	35	6.3	11.0	10.9	3.7	3.4
Mali[a]	36 [58]	47	4	3.6	9.6	2.7	6.4	6.6
Mauritania[a]	49 [44]	25	10	4.1	1.6	1.1	5.9	9.2
Morocco[b]	34 [69]	17	17	2.3	3.0	6.6	3.4	1.4
Niger[a]	22 [41]	40	6	2.5	1.7	-4.6	4.5	6.6
Pakistan[a]	20 [53]	26	17	4.0	2.7	1.9	10.0	8.8
Senegal[a]	38 [57]	18	17	3.2	2.6	1.6	4.0	6.6
Syria[b]	32 [59]	24	27	5.7	4.7	1.0	8.0	0[2]
Tunisia[b]	44 [80]	13	18	4.6	5.1	3.9	4.7	2.2
Turkey[b]	34 [74]	18	16	4.1	11.9	12.2	2.2	72.8

Notes: The reported figures are taken and calculated from the data from World Tables (World Bank, various issues). N.A. stands for not available. The figure in [.] is the share of total imports of manufactures (%) in 1999, except for Bangladesh (1998), Chad (1995), Guinea (1997), Mali (1997), Mauritania (1996), and Niger (1998) due to data availability. The country classification is by income, based on the 1999 GNP per capita (World Bank, 2001): 'a' for low income, 'b' for lower middle income, 'c' for upper middle income. [1]based on local currency per US$. [2]fixed the exchange rate to 11.225 local currency/US$ since 1988.

annum. The selected OIC countries have recorded a positive average growth of exports for the period 1990-1999. The average export growth for Bangladesh is the highest, that is 13.2 percent per annum. This can be explained by the export-led growth strategy implemented by the country. Iran, Jordan, and Niger have negative growth of imports over the period 1990-1999. This indicates the effectiveness of Government policies (fiscal or monetary) in improving the country's trade balances. We also observe that the import growth is found to be higher than export growth in Algeria, Chad, Morocco, and Turkey. Inflation is a major determinant of trade flows. It measures the growth of domestic prices (proxied by the consumer price index, CPI), and an increase in domestic price will make imports cheaper, and consequently demand for imports will rise. An average growth of the inflation rate above 10 percent per annum for the period 1990-1999 can be found in Algeria (21.4 percent), Indonesia (16 percent), Iran (25.5 percent), and Pakistan (10 percent).

Using available annual data (1960 to 1993) from the World Bank databases, Senhadji (1998) estimated a structural import demand model for 77 countries. Senhadji (1998) adopted an import demand equation that was close to the standard import demand function except for the use of the correct activity variable, GDP minus exports rather than GDP as a proxy for the activity variable. Among the OIC countries included in the analysis were Algeria, Chad, Egypt, Indonesia, Malaysia, Mauritania, Morocco, Pakistan, Sudan, Tunisia and Turkey. The results of the unit root test showed that the activity variable for Tunisia was stationary, $I(0)$. The cointegration test revealed that the volume of imports, domestic real activity and relative price were cointegrated for Mauritania, Morocco, Pakistan, and Turkey.[2] This brings into focus the need for further investigation of country studies on the long-run relationship of aggregate import demand behavior in OIC countries, particularly its invaluable formulation of policy.

In addition, this study is justified by the following considerations. Firstly, many of the existing empirical exercises adopted the traditional form of the import demand equation. It relates the quantity of import demanded to domestic real activity and relative prices (the ratio of import prices to the domestic price level) (Gafar, 1988). The relative prices variable is the ratio of import price to domestic price. Meanwhile, the activity variable is always proxied by domestic real income; real GDP (Reinhart, 1995) or GDP minus exports (Senhadji, 1998). This traditional specification of the long-run equilibrium in the empirical import demand is standard and needs no elaboration here. The studies which used the traditional import demand function can be found in Bahmani-Oskooee and Niroomand (1998), Gafar (1988), Pattichis (1999), Mah (2000), Sinha (1997), Reinhart (1995), Tang and Mohammad (2000), and Tang and Nair (2002). Xu (2002), however, noticed that the conventional import demand equations derived from either the imperfect substitute model or perfect substitute model, suffer from several drawbacks. First, they are partial and static in nature and, therefore, lack intertemporal elements, particularly the current income variable is typically used without any justification from intertemporal optimization theory. Second, the empirical implementation is somewhat *ad hoc*. Typically, a log-linear relationship is assumed but its inconsistency with consumer demand theory has long been well-known (Xu, 2002, 265-6). The present study has considered the above issues since the formulation of policy might prove very costly if the estimated import demand equation is inappropriate. Fortunately, Xu (2002) has derived a structural import demand function using an intertemporal optimization approach that was close to the conventional one, but more flexible, and therefore provided theoretical foundation for the estimation of an import demand equation. Xu's import demand equation takes into account a growing economy rather than an endowment economy, investment and government activity. Xu (2002, 269) proposed that the use of a correct activity variable, namely, '*national cash flow*'[3] rather than GDP, relative prices and a time trend that captures any trend-stationary shocks to consumption, is necessary and sufficient to define the long-run behavior of imports. Consequently, this would argue against the inclusion of any other variable in an *ad hoc* manner, for example, the current income variable, or supply-side variables. The inclusion of investment and the government sector also produces an 'activity variable' that is different from the 'activity variable' (GDP minus exports) as suggested by Senhadji (1998) (Xu, 2002, 269).

Secondly, this study has employed the recently developed technique for cointegration *viz.* the bounds testing procedure (Pesaran, Shin and Smith, 2001). This time series econometric technique has several advantages. Given the uncertainty concerning the time series properties of the variables in question, we view this method as the most appropriate in this context. Unlike standard cointegration tests, there is no need for unit root pre-testing if a conclusion can be made from the bounds test for cointegration (Pesaran, Shin and Smith, 2001). Meanwhile, the bounds test allows testing for cointegration when it is not known with certainty whether the regressors are $I(0)$, or $I(1)$ (Pesaran, Shin and Smith, 2001). Some applications in this context can be found in Tang (2001), Vita and Abbott (2002), and Coe and Serletis (2002). On the other hand, as stated by Pattichis (1999, 1062) and Mah (2000, 243), the conventionally used cointegration tests like Engle-Granger (1987) or Johansen-Juselius (1990) are not reliable for a study that involves a small sample size. Monte Carlo studies have shown, however, that despite the super-consistency of the Ordinary Least Squares (OLS) estimator in a cointegrating regression, substantially biased estimates could result in small samples (Banerjee et al. (1993, Chapter 7). Meanwhile, Toda (1994, 78) has documented that the available Monte Carlo evidence on the sample size of 300 or more observations is considered necessary to ensure good performance of the Johansen's likelihood ratio test for cointegrating ranks. Further, Pattichis (1999, 1062) and Mah (2000, 238 and 243) recommended the bounds test procedure and UECM for cointegration analysis, which are found to be appropriate for small sample studies. The UECM test is likely to have better statistical properties since it does not push the short-run dynamics into the residual term as in the case of the Engle-Granger technique (Pattichis, 1999, 1062). In addition, the ECM-based cointegration test will be more powerful than the residual-based Engle-Granger test, and will generally give unbiased estimates of the long-run relationship and standard t-statistics for conducting statistical tests of significance (see Pattichis, 1999, 1062, footnote 2). Several studies have recently employed the bounds test procedure (Pesaran, Shin and Smith, 2001) to examine the long-run relationship of the aggregate import demand function based on small sample size annual data. They are Pattichis (1999) with 20 observations for 1975-1994; Mah (2000) with 18 observations for 1980-1997, and Tang (2002a) with 26 observations for 1973-1998; Tang (2002b) with 34 observations for 1965-1998; Tang (2002c) with 40 observations for 1960-1999; Alam and Quazi (2003) with 27 observations for 1973-1999. Thus, this method is deemed to be suitable for the present study since we employ limited annual data from World Tables (1960 to 2000) with 41 observations. Another important advantage of the bounds test procedure is that estimation is possible even when the explanatory variables are endogenous.

We do not further review other empirical studies on the import demand function of individual countries considering the misspecification of the traditional import demand function used in previous studies as noticed by Xu (2002). The studies which used the traditional import demand equation can be found in Bahmani-Oskooee and Niroomand (1998), Gafar (1988), Pattichis (1999), Reinhart (1995), Sinha (1997), Senhadji (1998), Tang and Mohammad (2000), Mah (2000), and Tang and Nair (2002).[4]

In the next section, we discuss the specification of the aggregate import demand function, the data and the method used in estimation. Section 3 reports and discusses the results of the bounds test for cointegration analysis. The last section comprises the concluding remarks.

2. MODEL, DATA AND ESTIMATION METHOD

In the present study, we adopt the structural import demand specification proposed in Xu (2002, 269), and the long-run aggregate import demand equation can be written as:

$$\ln M_t = a_0 + a_1 \ln NCF_t + a_2 \ln RP_t + a_3 Time + e_t \tag{1}$$

where at period t, M_t is the volume of imports, NCF_t is the activity variable that is the '*national cash flow*' proposed in Xu (2002), RP_t is the relative price that is the ratio of import prices to the domestic price level, *Time* is the time trend variable that captures any trend-stationary shocks to consumption (Xu, 2002, 269), and e_t is the residual, while 'ln' indicates natural logarithms. Following the Keynesian line of argument, we expect $a_1 > 0$, that is, an increase in domestic activity will stimulate imports. An increase in the import price relative to the domestic price level will hurt import volume, thus $a_2 < 0$.

In this study, we consider 18 of the 27 OIC founding countries for analysis due to data unavailability for a sufficient sample span (World Bank, 2002) (see Appendix A (A.1)). These countries are Algeria, Bangladesh, Chad, Egypt, Guinea, Indonesia, Iran, Jordan, Malaysia, Mali, Mauritania, Morocco, Niger, Pakistan, Senegal, Syria, Tunisia and Turkey. The data definitions are cited in Appendix A (A.2). Most of the sample countries cover annual data from 1960 to 2000, which is found to be sufficient for cointegration analysis (see Sinha, 1997, 78) and applicable for the bounds test (see Mah, 2000, 243). Despite the unavailability of quarterly data, particularly the components of GDP, we chose to use annual data since we believe that the interaction of economic variables cannot work in short periods of a few quarters (see Tao and Zestos, 1999, 122). There is a gestation period for the macroeconomic variables to work through the import demand. In addition, Engle, Granger and Hallman (1989) have pointed out that the use of seasonal data to estimate the long-run model may give rise to inconsistent estimates of the long-run parameters. According to Charemza and Deadman (1992, 153), "Annual data could be used to estimate these long-run parameters thereby avoiding the need to model the seasonality, and the standard tests for cointegration applied."

Next, we briefly describe the method used in cointegration analysis, that is, the bounds testing approach (Pesaran, Shin and Smith, 2001).[5] The bounds test is computed based on an estimate of the conditional ARDL-ECM equation by using the Ordinary Least Squares (OLS) estimator (Pesaran, Shin and Smith, 2001). The conditional ARDL-ECM equation for equation (1), aggregate import demand function, can be written as below:

$$\begin{aligned} (2) \qquad \Delta \ln M_t = b_0 &+ \sum_{i=0}^{l} b_{1i} \Delta \ln NCF_{t-i} + \sum_{i=0}^{l} b_{2i} \Delta \ln RP_{t-i} \\ &+ \sum_{i=1}^{l} b_{3i} \Delta \ln M_{t-i} + b_4 \ln M_{t-1} + b_5 \Delta \ln NFC_{t-1} \\ &+ b_6 \ln RP_{t-1} + b_7 Time + u_t , \end{aligned}$$

where D is the series in first-differences, l is the lag length, and u_t is the residual.

Furthermore, from the estimated ARDL-ECM (2), we compute the bounds test for cointegration analysis that is to calculate the F-statistic (Wald test) for testing the null of non-cointegration (H_0: $b_4 = b_5 = b_6 = 0$) against the alternative of a stable long-run relationship between the volume of imports, relative price, and real activity variable (H_0: at least one of b_4, b_5, or b_6 is not zero). As mentioned by Pesaran, Shin and Smith (2001), the asymptotic distribution of the F-statistic is non-standard under the null hypothesis of no level relationship, irrespective of whether the regressors are $I(0)$ or $I(1)$. On the other hand, the asymptotic critical value bounds for the F-statistic computed by Pesaran, Shin and Smith (2001, 300-1) provide a band covering all possible classifications of the regressors into purely $I(0)$, purely $I(1)$ or mutually cointegrated.

For statistical inference at the conventional levels, $\acute{a}$ =10 percent, 5 percent or 1 percent, if the computed F-statistic (Wald test) falls outside the critical value bounds, lower bound and upper bound, a conclusive inference can be made without considering the order of integration, I(d) of the regressors. If the computed F-statistic lies above the upper bound, the null hypothesis can be rejected, revealing a cointegrating relation among the examined variables. If the test statistic (F-statistic) lies below the lower bound, the null of no cointegrating relation cannot be rejected, thus no long-run relationship among the examined variables can be made. In the case where the F-statistic falls between the upper and lower bounds, a conclusive inference cannot be made. Here, the order of integration, I(d) for the regressors must be investigated before any conclusion can be drawn (see Pesaran, Shin and Smith, 2001).

From the UECM equation (2), the long-run coefficients can be derived that are $-(b_5 / b_4)$ and $-(b_6 / b_4)$ for the real activity variable and relative prices, respectively (Pesaran, Shin and Smith, 2001, 294). The short-run effects are captured by the estimated coefficients of the first differenced variables. We also perform parameter stability tests for the estimated model using the CUSUM or CUSUM of squares tests.[6]

3. THE EMPIRICAL RESULTS

This section reports the results of the bounds test for cointegration analysis. First, we estimate a conditional ARDL-ECM in general form as in equation (2). An appropriate lag-length, l, for the conditional ARDL-ECM (equation 2) is selected to minimize the Akaike information criterion (AIC) from three, two and one-year lag order that we consider appropriate given the small sample in this study (see Pattichis, 1999, 1063). A higher lag order, l=4, is usually not feasible given that we usually have 41 annual observations except for Mauritania.[7] Considering the degree of freedom and the problem of over-parameterization in UECM (equation 2), a parsimonious specification of ARDL-ECM can be re-estimated, that is all of these first difference regressors that have relatively small absolute t-ratio (less than one) are dropped sequentially.[8]

TABLE 2
The Results of Bounds Test for Cointegration

Country	Bounds test (F-statistic)	Conclusion	Normalized Cointegrating Relation[2] [ln$M$, lnRP, lnNCF]
1. Algeria	13.24[c]	CI	[-1.00, -1.66^{+}, -0.92^{+}]
2. Bangladesh	2.05[a]	NC	
3. Chad	13.30[c]	CI	[-1.00, -0.93^{+}, 0.09^{+}]
4. Egypt	7.10[c]	CI	[-1.00, -0.24^{+}, -0.84^{+}]
5. Guinea	41.59[c]	CI	[-1.00, -0.11^{+}, -0.50^{+}]
6. Indonesia	5.17[c]	CI	[-1.00, -0.6^{+}, -1.25^{+}]
7. Iran	10.72[c]	CI	[-1.00, 0.33^{++}, 0.73]
8. Jordan	2.78[a]	NC	
9. Malaysia	4.24[d]	NC	
10. Mali [1]	2.75[a]	NC	
11. Mauritania	1.05[a]	NC	
12. Morocco	2.74[a]	NC	
13. Niger	7.90[c]	CI	[-1.00, 0.24^{++}, -1.05^{+}]
14. Pakistan	8.25[c]	CI	[-1.00, 0.44^{+++}, -0.89^{+++}]
15. Senegal	4.33[d]	NC	
16. Syria	5.7[c]	CI	[-1.00, -2.69, -3.39]
17. Tunisia	2.38[a]	NC	
18. Turkey	8.48[c]	CI	[-1.00, 0.22^{+}, -0.84^{+}]
#Critical value bounds:	lower bound	upper bound	
10%	4.19	5.06	

Notes: # The critical bounds are obtained from Pesaran, Shin and Smith (2001, 301), Table CI(v) Case V: Unrestricted intercept and unrestricted trend with two regressors. CI denotes Cointegrated, and NC denotes Non-cointegrated. ‘a’ indicates that the statistic lies below the 0.10 lower bound, ‘b’ that it falls within the 0.10 bounds, ‘c’ that it lies above the 0.10 upper bound, and ‘d’ indicates that the statistic lies below the 0.05 lower bound. The bounds at 0.05 are 4.87 and 5.85.

[1] The null hypothesis of all the estimated coefficients in the final ARDL-ECM are zero has not been rejected at the 10 percent level based on *F*-statistic since its probability value (*p*-value) is 0.28 (see Appendix B, “*F*-statistic” for Mali case). It indicates that the estimated ARDL-ECM model is meaningless for Mali. This null hypothesis, however, has been rejected at the 10 percent significance level for other countries examined with *p*-value less than 0.10 (see Appendix B, “*F*-statistic”).

[2] In order to make inferences on whether import demand is price and income elastic, tests have been performed to examine if the coefficients of elasticity are significantly different from one. +, ++, +++ denote rejection of the null of long-run coefficient is one based on *p*-value at the 1%, 5% and 10% significance levels, respectively. Here, I would like to thank an anonymous referee who has raised this point.

The estimated final ARDL-ECMs for the 18 OIC countries are reported in Appendix B. In addition, the test statistics for serial correlation, and autoregressive conditional heteroskedasticity (ARCH) in the residuals are all satisfactory (see Appendix B). A possible concern raised from the estimated models is the stability of its coefficients. However, the CUSUM or CUSUM of squares tests are inside the 5 percent critical bounds indicating the estimated coefficients of ARDL-ECM are stable over the period analyzed (See Appendix C). Table 2 reports the results of the bounds test for testing the presence of a long-run relationship among the volume of imports demanded, the domestic real activity variable and relative prices term.

Several observations can be made regarding the obtained output. The cointegration tests reported in Table 2 show that the test statistics (bounds test, *F*-statistic) for Algeria, Chad, Egypt, Guinea, Indonesia, Iran, Niger, Pakistan, Syria, and Turkey exceed the upper bound, 5.06, at the 10 percent significance level, and have to reject the null of non-cointegration. This reveals that the volume of imports, domestic real activity, and relative price are cointegrated in these countries (10 countries). Thus, we can say that there is a stable import demand function in these countries during the period analyzed.

For the remaining countries (Bangladesh, Jordan, Malaysia, Mali, Mauritania, Morocco, Senegal and Tunisia), no cointegrating relation of import demand behavior can be concluded since the computed *F*-statistics (bounds test) lie below the lower bound of 4.19 at the 10 percent, or 4.87 at the 5 percent significance levels, respectively. Once the conclusion of cointegration or non-cointegration can be made from the bounds test, the investigation for the order of integration, I(d) of the involved time series is not necessary (Pesaran, Shin and Smith 2001). Some possible reservations on the finding of non-cointegrating relations can be justified in the light of certain destabilizing forces, structural breaks and the omission of relevant theoretically inferred determinants (Masih and Masih, 2000, 634). However, the present study does not investigate further these issues in detail. Perhaps, this can be left for future research.

From the normalized long-run relation of the import demand function reported in Table 2 (last column), the relative price variable is found to be elastic in Algeria and Syria as the estimated elasticity is more than unity (with the correct sign, i.e., negative). However, in the cases of

Iran, Niger, Pakistan and Turkey, the estimated long-run relative price elasticity is in the opposite sign, i.e., positive.[9] This result reveals a question about the substitutability of these countries for imports with domestically produced goods.[10]

In the long-run, the real activity variable is found to be elastic for Indonesia, Niger, and Syria. However, a negative sign of real activity's elasticity is found for Algeria, Egypt, Guinea, Indonesia, Niger, Pakistan, Syria and Turkey. However, it is not surprising to find negative income elasticity for import demand, if increases in domestic activity are due to the production of import-substitute goods, then imports may actually fall resulting in a negative elasticity (Bahmani-Oskooee and Niroomand, 1998, 102). This means that if economic growth leads to an increase in the production of goods, which would have been otherwise imported then imports will be decreased, resulting in negative income elasticity to import demand. On the one hand, the income elasticity of import demand shows a negative sign if increases in domestic output exceed the increase in the domestic demand for the types of product imported or if imports from certain countries tend to be inferior goods (Tegene, 1989, 1449, footnote 7).

We do not discuss further the estimated short-run elasticities of the import demand function since the policy implication is mainly based on '*static*' or '*long-run*' estimates as mentioned in Section 1 (also see the study by Bahmani-Oskooee, 1998).

4. CONCLUDING REMARKS

The present study provides a piece of empirical work to investigate the presence of a cointegrating relation between the volume of imports demanded and its determinants, namely relative prices and the real activity variable for 18 OIC countries. Using the recently developed import demand specification (Xu, 2002), the results of the bounds test indicate that a long-run relation of import demand function exists for 10 of the 18 sample countries, namely, Algeria, Chad, Egypt, Guinea, Indonesia, Iran, Niger, Pakistan, Syria and Turkey. The estimated relative price and real income elasticities for Chad, Egypt, Guinea, Pakistan, and Turkey are both found to be inelastic. The price elasticity is only elastic for Algeria. Meanwhile, Indonesia, and Niger are found to be income elastic (see Table 2).

From a policy perspective, the results presented in this study are important. Among the sample countries for which the estimated price elasticity is within the range suggested by Heien (1968), i.e., -0.5 to -1.0, are Algeria, Chad, Indonesia, and Syria, indicating that the devaluation policy may yield a favorable outcome on a country's trade balances. Algeria and Syria have price elasticities of more than unity and outside the suggested range of -0.5 to -1.0 (Heien, 1968), i.e., -1.66 and -2.69, respectively, suggesting that the Marshall-Lerner condition may be satisfied even though this study does not estimate export price elasticites.[11] Based on the rule of Heien and the Marshall-Lerner condition, the results indicate that devaluation or depreciation will have favorable effects on the trade balance for these countries. The exchange rate data (positive average growth of the exchange rate based on local currency/US$) shows that all of these countries have devalued their currencies over the period 1990-1999. However, devaluation is not applicable for Syria since the exchange rate has been fixed at 11.225 local currency/US$ in 1988. On the other hand, the estimated long-run price elasticities for Chad, Egypt, Guinea, Indonesia, Iran, Niger, Pakistan and Turkey are well below unity. This suggests that large relative price swings are necessary to produce an appreciable reallocation of trade flows (Reinhart, 1995, 308). For Algeria, domestic inflation needs to be kept in check, as the estimated price elasticity for these countries is elastic implying that domestic inflation would increase the demand for imports. Over the period 1990-1999, Algeria has faced inflationary pressures with 31.7 percent in 1992, but with inflation dropping to 8.5 percent and 2.6 percent in 1997 and 1999, respectively (*World Tables*, World Bank, various issues).

On the other hand, the domestic real activity or economic growth may not have significant negative implications on the trade balance for Chad and Iran as import demand is inelastic to domestic activity. For Algeria, Egypt, Guinea, Indonesia, Niger, Pakistan, Syria and Turkey (with negative domestic activity elasticity), government strategies on the production of import-substitute goods, particularly the development of resource-based industries that have high import contents, may be used to dampen the increase in import demand. For Bangladesh, Jordan, Malaysia, Mali, Mauritania, Morocco, Senegal and Tunisia whose import demand behavior is unstable (no long-run relationship), monetary and fiscal policies may still be appropriate to improve the country's trade balances. What we can observe is that the import structure of the examined countries is highly dominated by manufactured goods with the share of total imports of manufactures in the range of 41 percent (in 1998) for Niger to 85 percent (in 1999) for Malaysia (see Table 1).

The above discussed policy issues can be linked to a recently conducted study by Arize (2002). Arize (2002) has found that imports and exports are cointegrated in Egypt, Indonesia, Iran, Jordan, Malaysia, Morocco, Pakistan, and Tunisia using Stock and Watson's (1988) cointegration test. Arize's (2002) study is based on 50 developing and developed countries. The study has concluded that macroeconomic policies have been effective in the long-run and suggests that these countries are largely not in violation of their international budget constraints.

However, the present study has several drawbacks. Firstly, we have reservations on the policy implications discussed above, which are merely based on the import demand estimations. They may be too strong in some cases since most of the examined countries are found to be income and price inelastic – only Algeria is price elastic, and Indonesia and Niger are income elastic. However, we have not made any comment on the countries' trade policies due to lack of available information about the trade and macroeconomic policies implemented in all of the selected OIC member countries from available published materials as well as soft materials through internet search. This limitation is commonly acknowledged by many researchers for studying low-income (or less developing) countries like most of the OIC member countries in the present study. Secondly, the bounds test technique is based on a single-equation approach. Consequently, it is inappropriate in situations where there may be more than one level relationship, i.e., more than two variables are involved (Pesaran, Shin and Smith, 2001, 315). In future studies, a system-based approach like that of Johansen and Juselius (1990) can be used to check its consistency. Finally, the present study only involves 18 OIC member countries out of a total of 57 countries. This is due to unavailability of annual data for a sufficiently long time span for all the variables that enter our import demand function from the published source - *World Tables* (World Bank, various issues). For further studies, more countries should be included in the analysis by using data collected by other relevant agencies.

ENDNOTES

1. The Organization of Islamic Conferences (OIC) was established in 1969. It is an international grouping of fifty-seven States which have decided to pool their resources, combine their efforts and speak with one voice to safeguard their interests, and to secure the progress and well-being of their people and of all Muslims in the world.

2. Senhadji (1998, 245) only reported the countries that had correct estimated signs for the price and income elasticities; negative sign for relative price and positive sign for activity variable.

3. This correct activity variable proposed by Xu (2002) that is the '*national cash flow*' variable is derived from GDP_t-I_t-G_t -EX_t, where GDP_t is Gross Domestic Product; I_t is investment; G_t is government spending; and EX_t is exports (Xu, 2002, 269).

4. The present study does not report the estimated elasticities obtained from these studies considering the limited space available in this study for a large number of countries. For example, the study by Bahmani-Oskooee and Niroomand (1998) involved almost 30 countries, and 66 countries were involved in Senhadji's (1998) study. However, more literature review can be found in Tang (2002c, 182-4).

5. See Coe and Serletis (2002, 181-3), for the technical expression of the bounds testing approach (Pesaran, Shin and Smith, 2001).

6. The CUSUM of squares test (Brown, Durbin, and Evans, 1975) is one of the diagnostic tests. If movement of the CUSUM of squares statistic is outside the area between the two critical lines, it shows parameter or variance instability.

7. For Mauritania, a four lag structure (l=4) has been adopted to account for the autocorrelation problem for the UECM. The AR(2) autocorrelation in the UECM has not been corrected via the Generalised Least Squares method (GLS) since this is not acceptable as the dependent variable is already first differenced, and lags of the dependent as well as independent variables are included in the model to take care of autocorrelation. The correction of using GLS would cause "over-differencing" and as a result, the findings would not be reliable. Variables could be artificially made to be significant when they are rightly not significant. I thank an anonymous referee for raising this point.

8. Verbeek (2000, 53) was concerned that "In presenting your estimation results, it is not a 'sin' to have insignificant variables included in your

specification. The fact that your results do not show a significant effect on y_i of some variable x_{ik} is informative to the reader and there is no reason to hide it by re-estimating the model while excluding x_{ik}." By considering this issue, the paper chooses to drop regressors with *t*-ratios of less than one in absolute values but not all of those that are not significant. This method has recently been practiced in Tang and Mohammad (2000), Mohammad and Tang (2000), and Tang (2002c) in estimating a parsimonious specification of the UECM for import demand analysis.

9. Bahmani-Oskooee and Niroomand (1998) estimated the long-run import and export demand behavior for 30 countries using annual data from 1960 to 1992. They found that some estimated relative price elasticities were in positive sign (see Bahmani-Oskooee and Niroomand, 1998, 107, Table 5).

10. The relative price variable is based on the assumption that there is some degree of substitutability between imports and domestically produced goods (Bahmani-Oskooee and Niroomand, 1998, 102, note 1).

11. The Marshall-Lerner condition indicates a stable foreign exchange market if the sum of price elasticity of demand for imports and the demand for exports exceeds one (in absolute terms). Thus, exchange rate policy in particular devaluation can be adopted to correct for balance of payments disequilibrium (see for example, Bahmani-Oskooee and Niroomand, 1998).

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APPENDIX A
Data Appendix

A.1. Sample of Countries (sample period)

Algeria (1960 to 2000)	Mali (1967 to 2000)
Bangladesh (1960 to 2000)[1]	Mauritania (1960 to 2000)
Chad (1960 to 2000)	Morocco (1960 to 2000)
Egypt (1960 to 2000)	Niger (1960 to 2000)
Guinea (1986 to 2000)	Pakistan (1967 to 2000)
Indonesia (1960 to 2000)	Senegal (1960 to 2000)
Iran (1974 to 2000)	Syria (1975 to 2000)[2]
Jordan (1976 to 2000)	Tunisia (1961 to 2000)
Malaysia (1960 to 2000)	Turkey (1963 to 2000)

Notes: [1]As part of Pakistan was one of the OIC founders. It became full member after independence in 1974.
[2]As part of the United Arab Republic (UAR) was one of the OIC founders. It became full member after leaving the UAR in 1970 (www.sesrtcic.org/oic/oicaccda.shtml).

We include only countries in Table A.1 for analysis due to data unavailability for a sufficient time span as well as all variables used in the import demand function.

All data are collected from *World Tables*, World Bank (2002). The data for Turkey are obtained from OECD Economic Outlook (1987=100). All monetary variables are measured in the domestic currency. Afghanistan, Jamahiriya, Kuwait, Lebanon, Palestine, Saudi Arabia, Somalia, Sudan, and Yeman have been excluded from the analysis due to data unavailability for a sufficiently long period and for some variables such as import price, GDP deflator or Consumer Price Index, CPI (World Bank, 2002). The exclusion of the above countries is found to be consistent with Senhadji (1998) due to data unavailability from the World Bank's database.

A.2. Data Definitions

1. *M* is the real imports for goods and services. The nominal values are deflated by the implicit import price index, 1995=100 (in the domestic currency).

2. *RP* is the ratio of import price to domestic price. The domestic price is proxied by the GDP deflator (1995=100).

3. *NCF* is the real activity variable, ‘*national cash flow*’ that is proposed in Xu (2002). This variable is defined as GDP-I-G-EX, where GDP is Gross Domestic Product, I is investment, G is government expenditure and EX is exports. The GDP deflator (1995=100) has been used to deflate the nominal series.

APPENDIX B

Estimated Import Demand Function - Final Unrestricted Error Correction Model

Dependent variable: $\Delta \ln M_t$

Country:	Algeria	Bangladesh	Chad	Egypt	Guinea	Indonesia	Iran	Jordan	Malaysia
$\Delta\ln NCF_t$	-0.56*		0.25	-0.23*		-0.56*		-0.096*	-0.37**
$\Delta\ln \mathrm{NCF}_{t-1}$	0.16**		0.23	0.32*		0.77*	-0.63**		0.35
$\Delta\ln NCF_{t-2}$		-0.74		0.14		0.41			
$\Delta\ln NCF_{t-3}$			-0.40**	0.21**			-0.45***		
$\Delta\ln RP_t$		-0.66*				-0.37*		-0.81*	-0.97*
$\Delta\ln RP_{t-1}$	0.93*	-0.27***	0.60*			0.18			
$\Delta\ln RP_{t-2}$	0.38		0.41**				0.35**		0.95*
$\Delta\ln RP_{t-3}$				0.16			0.21		
$\Delta\ln M_{t-1}$	0.13	-0.21	0.15	0.42*	0.29***	0.55**	0.24		0.41**
$\Delta\ln M_{t-2}$		0.36**	0.52*	-0.22		0.35***	0.33		0.36**
$\Delta\ln M_{t-3}$			0.45*						
$\ln M_{t-1}$	-0.86*	-0.32**	-0.75*	-0.64*	-1.54*	-0.86*	-0.67*	-0.36**	-0.36**
$\ln NCF_{t-1}$	-0.8*	0.60	0.06	-0.53*	-0.77*	-1.08*	0.49	-0.04	0.05
$\ln RP_{t-1}$	-1.44*	-0.15	-0.69*	-0.15**	-0.17***	-0.52**	0.23	-0.14	-0.81**
Time Trend	0.06*	-0.002	0.02*	0.05*	0.07*	0.12*	-0.08*	0.006	0.04*
Constant	15.01*	-1.58	6.05*	3.76*	16.39*	7.31*	2.05	3.74	4.05**
Lag length l^1	2	2	3	3	1	3	3	2	2
Adjusted-R^2	0.761	0.681	0.51	0.572	0.914	0.541	0.624	0.680	0.577
F-statistic	14.1	9.77	4.13	5.38	26.45	4.96	4.66	9.15	6.04
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
LM test	0.74 [1]	0.15 [1]	0.52 [1]	3.75 [2]	0.57 [1]	0.533 [1]	4.86 [3]	0.86 [1]	0.25 [1]
	(0.39)	(0.70)	(0.47)	(0.15)	(0.45)	(0.465)	(0.182)	(0.35)	(0.614)
ARCH test	0.09 [1]	0.023 [1]	1.22 [1]	0.32 [1]	0.19 [1]	0.06 [1]	0.22 [1]	0.62 [1]	0.023 [1]
	(0.77)	(0.88)	(0.27)	(0.57)	(0.66)	(0.80)	(0.64)	(0.43)	(0.88)

APPENDIX B (continued)

Country:	Mali	Mauritania[2]	Morocco	Niger	Pakistan	Senegal	Syria	Tunisia	Turkey
$\Delta\ln NCF_t$	-0.24	-0.16**	-0.69*			-0.15	-0.63**	-0.21*	
$\Delta\ln NCF_{t-1}$	-0.20	-0.36***	-0.28***	0.31	0.76	0.11***	0.94**		
$\Delta\ln NCF_{t-2}$		-0.19			0.72		0.55**		
$\Delta\ln NCF_{t-3}$		-0.07		-0.34			0.21	-0.09***	
$\Delta\ln RP_t$	-0.40	-0.07	-0.86*	-0.56**	-0.15	-0.34*	-0.69***	0.29**	
$\Delta\ln RP_{t-1}$	0.27	-0.89		-0.78**	-0.29**		0.68**		
$\Delta\ln RP_{t-2}$		0.36		-0.64**	-0.23		0.61**	0.26**	
$\Delta\ln RP_{t-3}$		0.62		-0.40	-0.53*			-0.16***	
$\Delta\ln M_{t-1}$	0.12	-1.06**		-0.24			-0.77		0.27***
$\Delta\ln M_{t-2}$		-0.48		-0.19	-0.31***		-0.67	0.17	
$\Delta\ln M_{t-3}$		-0.18		-0.17	-0.29***		-0.45		
$\ln M_{t-1}$	-0.62**	0.12	-0.19***	-0.67*	-0.499**	-0.43*	-0.58	-0.11	-0.92*
$\ln NCF_{t-1}$	0.38	0.16	-0.17	-0.7*	-0.45	-0.05	-1.97*	0.03	-0.77*
$\ln RP_{t-1}$	-0.36	0.67	0.13	0.17	0.22***	-0.03	-1.56**	0.15**	0.2
Time	0.03***	-0.02	0.01	-0.006	0.03	0.008	0.15**	0.002	0.1*
Constant	2.17	-3.46	0.64	7.7*	4.25	3.18***	17.18*	0.04	26.15*
Lag length, l	1	4	1	3	3	1	3	3	1
Adjusted-R^2	0.085	0.848	0.613	0.494	0.768	0.455	0.603	0.618	0.383
F-statistic	1.32	11.85	9.59	3.71	9	5.53	3.28	6.67	5.34
	(0.282)	(0.00)	(0.00)	(0.003)	(0.00)	(0.00)	(0.06)	(0.00)	(0.00)
LM test	0.19 [1]	1.54 [1]	0.0 [1]	1.55 [1]	0.0 [1]	0.32 [1]	0.47 [1]	0.0 [1]	0.14 [1]
	(0.66)	(0.21)	(0.99)	(0.213)	(0.99)	(0.57)	(0.49)	(0.98)	(0.71)
ARCH test	0.28 [1]	0.00 [1]	0.13 [1]	0.005 [1]	0.12 [1]	1.51 [1]	1.55 [1]	2.61 [1]	0.36 [1]
	(0.59)	(0.98)	(0.714)	(0.941)	(0.73)	(0.22)	(0.21)	(0.11)	(0.55)

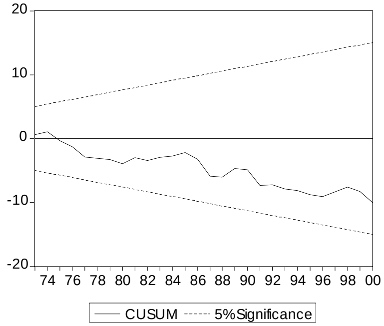
Notes: *p<0.01, **p<0.05, and ***p<0.1 for the estimated coefficient being different from zero. Δ is the first difference operator. The value in ( ) is p-value. LM test is Breusch-Godfrey serial correlation LM test, under the null of no serial correlation up to lag order p, where p is a pre-specified integer. The ARCH test is computed to test the null hypothesis that there is no ARCH up to order q in the residuals. The figure reported in [ ] is p or q.

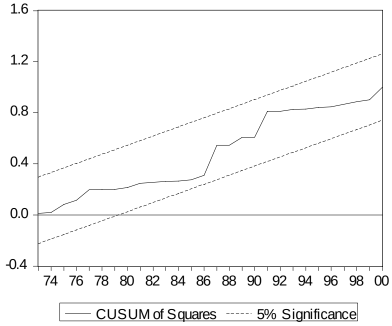
[1] The reported figure is the optimum lag length for general UECM. In the case of Guinea, we use a one-year lag due to limited annual observations, 1986 to 2000.

[2] The model is estimated with a four-lag structure, l=4 in order to remove the autocorrelation problem. Due to the limited space, the estimated coefficients for the forth-lagged regressors are reported in the table. The estimated coefficients for $\Delta\ln NCF_{t-4}$, $\Delta\ln RP_{t-4}$, and $\Delta\ln M_{t-4}$ are 0.03, 0.48, and -0.05 but all are insignificant at 10% level.

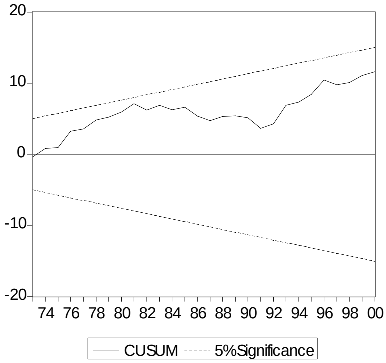
APPENDIX C
CUSUM and CUSUMSQ Tests

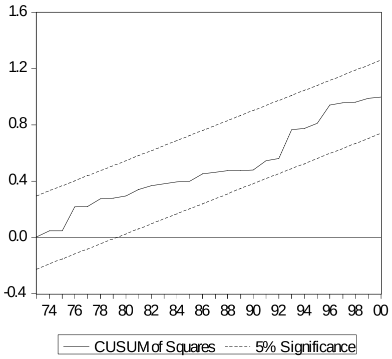
Algeria



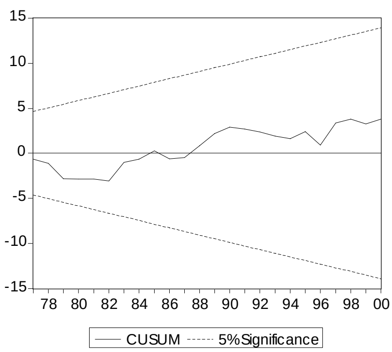


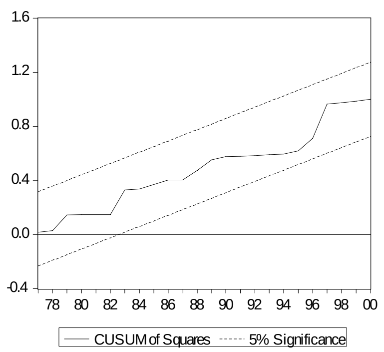
Bangladesh



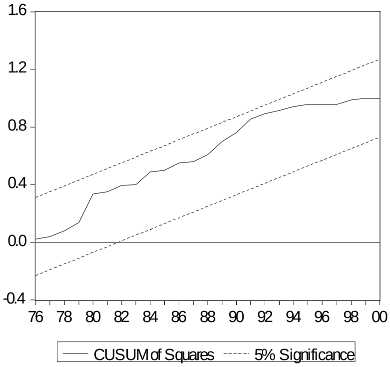


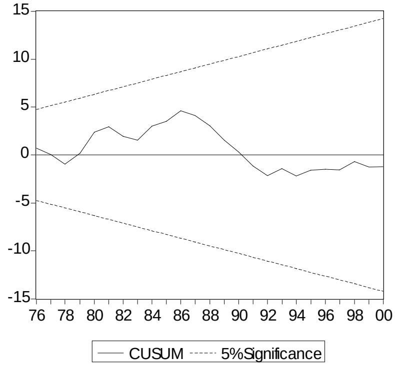
Chad



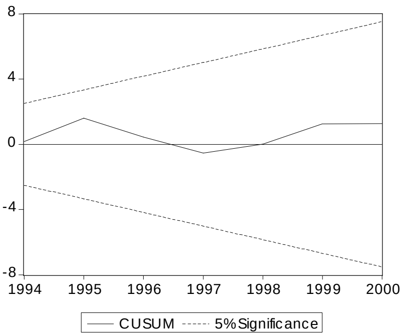


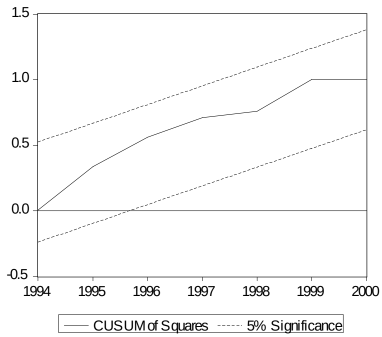
Egypt





Guinea





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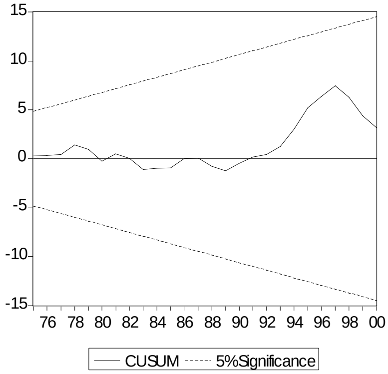

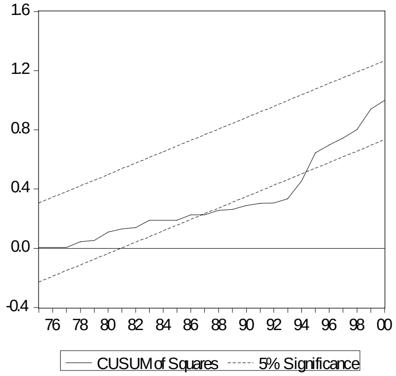

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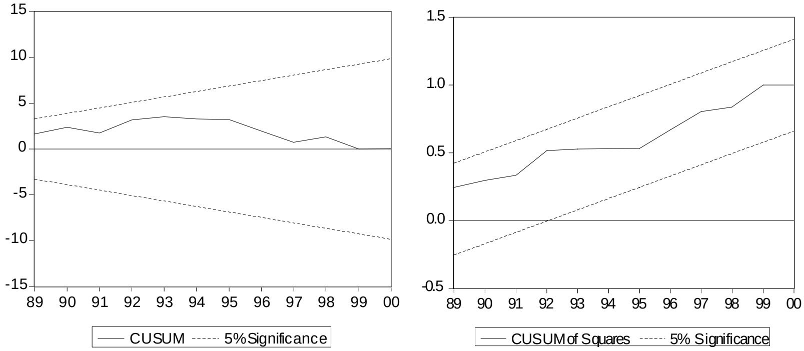

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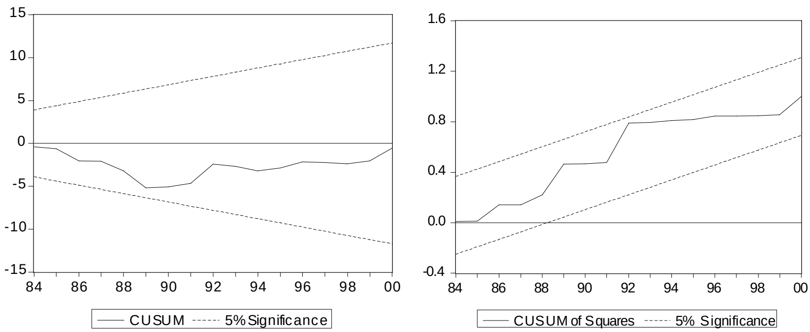

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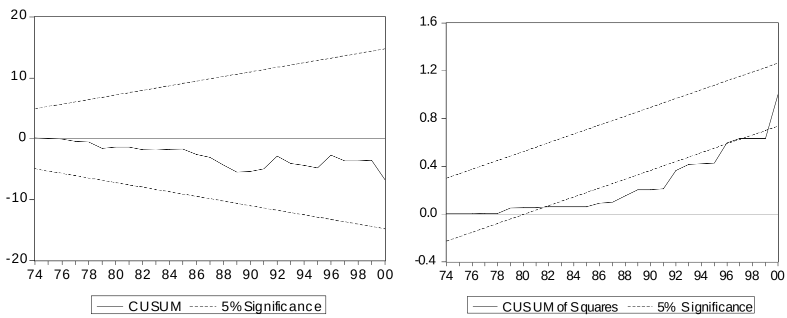

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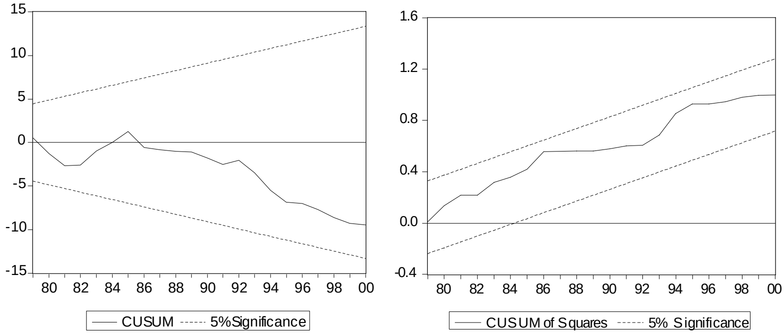

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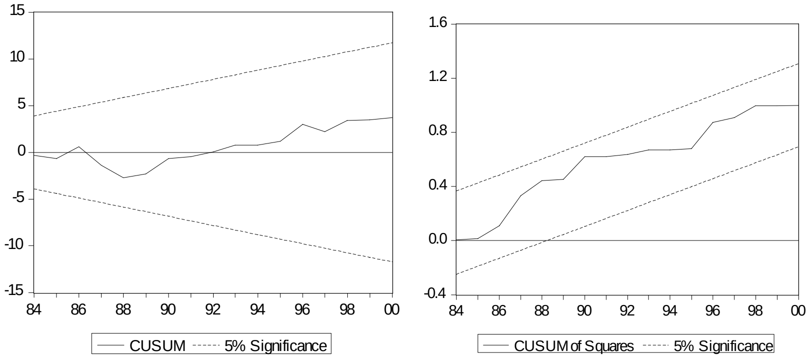

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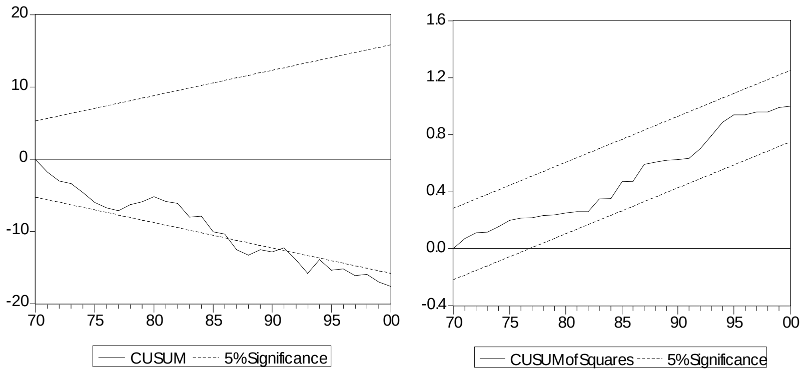

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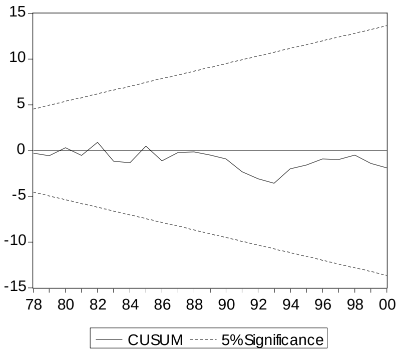
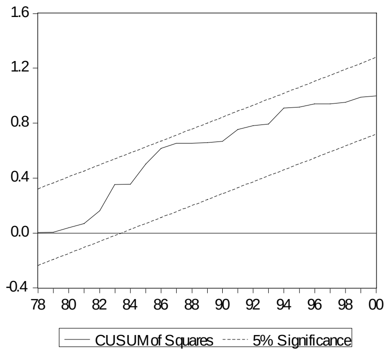
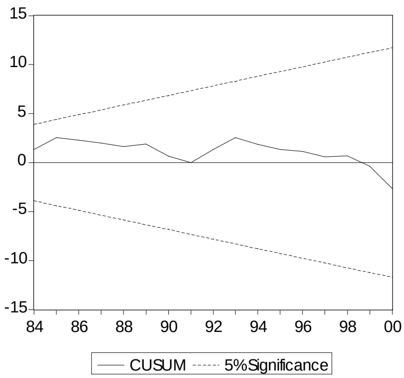
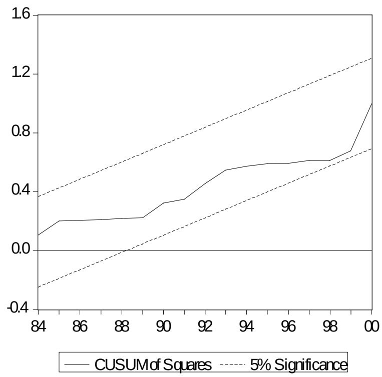




Pakistan





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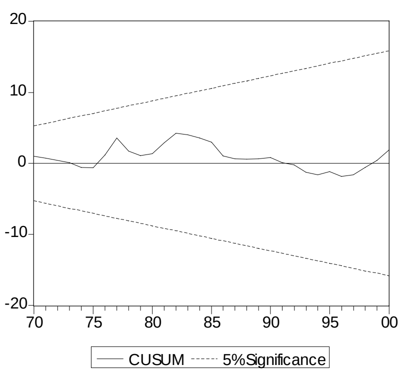

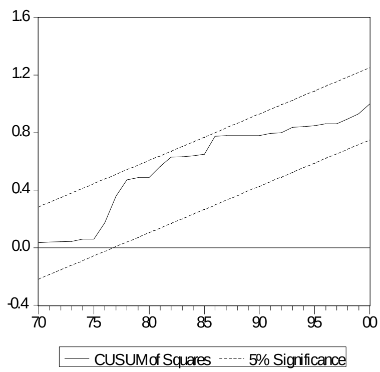

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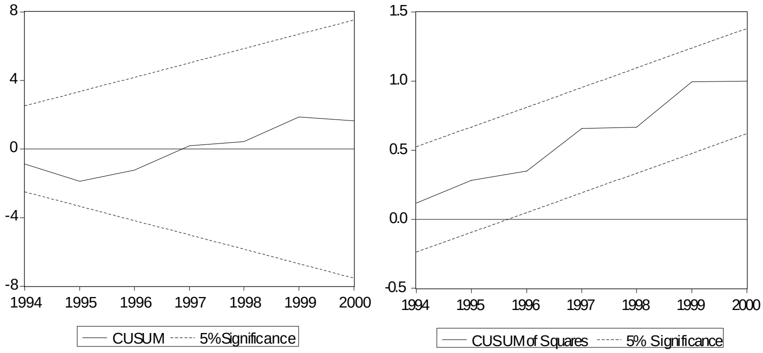

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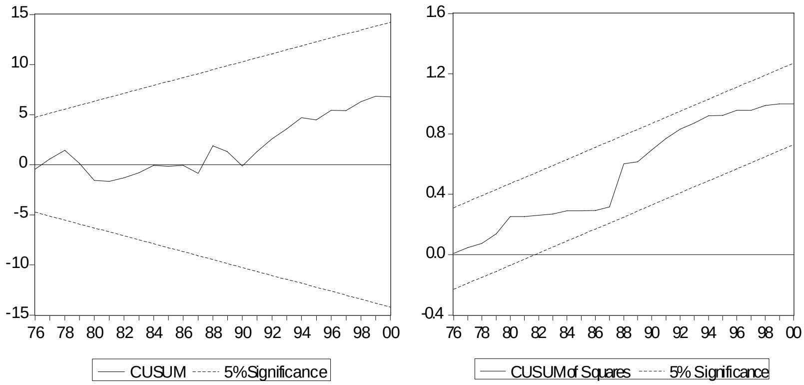

Turkey