ARE EXPORTS AND IMPORTS COINTEGRATED? EVIDENCE FROM NINE MENA COUNTRIES*

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Abstract

The aim of this article is to examine the long-run convergence (cointegration) between exports and imports for nine MENA (Middle East and North Africa) countries. Evidence of cointegration ensures that trade imbalances are sustainable. The article explores this issue by applying the bounds testing approach to cointegration and by using annual data. The findings of cointegration between exports and imports for Iran, Israel, Jordan, and Tunisia indicate that these countries are not in violation of their international budget constraint. In addition, the CUSUM and CUSUMSQ tests confirm stability of the estimated parameters.

JEL Codes:
Keywords: exports, imports, cointegration, ARDL

1. Introduction

The presence of a long-run equilibrium or cointegration relationship between a country's exports (*EX*) and imports (*IM*) implies that it is is not in violation of her international budget constraint and that trade imbalances are short-run phenomena and are sustainable in the long-run. In addition, cointegrated exports and imports may also suggest that macroeconomic policies may have been effective in bringing exports and imports into a long-run equilibrium relationship and that there is no productivity gap between the domestic economy and the rest of the world implying a lack of permanent technological shocks to the domestic economy (Irandoust and Ericsson 2004). Moreover, the knowledge of whether exports and imports are cointegrated is necessary in the design and evaluation of current and future macroeconomic policies aimed at achieving trade balances (Bahmani-Oskoee and Ree (1997), Arize (2002)).

Over the past decade, cointegration relationship between *EX* and *IM* has received increasing attention and has been investigated for a number of countries. Below is a brief review of recent studies that examined this long-run equilibrium relationship between exports and imports, in particular, the focus is on studies that included countries from the MENA region.

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Arize (2002) used quarterly data for the period 1973–1998 from 50 OCED and developing countries to investigate the long run relation between exports and imports (MENA countries included in his study are Egypt, Israel, Jordan, Kuwait, Morocco, and Tunisia). Using Johansen cointegration tests, he found evidence of cointegration between exports and imports for 35 of the 50 countries examined. Among these, 31 countries had a positive export coefficient (for MENA countries, cointegration was found for Egypt, Iran, Israel, Kuwait, and Tunisia). [1]

Tang (2005) examined the presence of a long run relationship between imports and exports for 27 Organization of Islamic Conferences (OIC) member nations using Engle and Granger (1987) residual based cointegration technique. Because volume of imports and volume of exports were integrated of different orders for Algeria, Bangladesh, Chad, Jordan, Mali, Niger, Pakistan, Senegal, Sierra Leone, and Syria, Tang concludes that no cointegrating relation is present between imports and exports for the these countries. His results of unit root and cointegration tests indicate no cointegration between exports and imports for Egypt, Iran, Morocco, and Tunisia.

Meanwhile, in a sample of the same 27 OIC member countries, Tang (2006) reexamined the relationship between exports and imports and found evidence of a long-run equilibrium between exports and imports for 9 of the 27 member countries examined. Using Gregory and Hansen (1996) cointegration test with structural breaks, he found a long-run equilibrium relationship between imports and exports only for Morocco among the MENA countries examined (e.g., no cointegration was found for Algeria, Egypt, Iran, Jordan, Syria, and Tunisia).

This study investigates whether a long-run equilibrium exists between *EX* and *IM* for nine MENA countries: Algeria, Egypt, Iran, Israel, Jordan, Morocco, Sudan, Syria, and Tunisia. First, we investigate the integration (stationarity) properties of the data for all countries using Philips-Perron (PP) unit root test. Second, we apply the bounds testing approach to test for long-run relation between *EX* and *IM*. This testing approach to cointegration has many econometric advantages compared to other methods. Third, we utilize three different approaches to estimate long-run elasticities to ascertain robustness of the results. Fourth, we test for parameter stability using CUSUM and CUSUMSQ as suggested by Pesaran *et al* (1999).

The rest of the study is organized as follows. Section II presents the theoretical background of the model, Section III describes the estimation technique, Section IV presents the data and the empirical results and Section V concludes.

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[1] A reservation, as pointed out by Tang (2005), on Arize's (2002) study is that he uses the ratio of imports and exports to GDP, which may not give an accurate portrayal of the trend in imports or exports. The change in GDP may increase/decrease the trend of these ratios. Arize also reports evidence of parameter instability in the estimated relationships for both Iran and Israel. If the estimated long-run relationship is not stable, conclusions can be highly misleading.

2. Theoretical background

Husted (1992) developed a simple framework for a small open economy that implies a long-run equilibrium relationship between exports and imports. In this model, the representative agent produces and exports a single composite good, has no government, can borrow and lend in international markets at the world interest rate using one-period financial instruments, and aims to maximize lifetime utility subject to budget constraints. The representative agent's current-period budget constraint in period *t* is given by:

$$C_0 = Y_0 + B_0 - I_0 - (1 + r)B_{-1} \tag{1}$$

where C_0 is current consumption; Y_0 is output; I_0 is investment; r is the one period world interest rate; B_0 is international borrowing which could be positive or negative; and $(1 + r)B_{-1}$ is the historically given initial debt of the representative agent, corresponding to the country's external debt. In deriving a testable model, Husted (1992) made several assumptions including that the interest rate is stationary with mean r and both imports and exports follow a random walk with drifts. The testable model is of the following form:

$$IM_t = \alpha + \beta_1 EX_t + e_t \tag{2}$$

Here *IM* and *EX* denotes imports and exports of goods and services. The necessary condition (weak form) for an economy to satisfy its intertemporal budget constraint is for the residual series of equation 2, e_t, to be stationary or $I(0)$ process. If this condition is not met, this indicates that the economy is not functioning properly and fails to satisfy its budget constraint and hence is expected to default on its debt (Hakkio and Rush, 1991). Such a finding is also construed as evidence against the sustainability of the current account balance (Narayan and Narayan 2005). Irandoust and Ericsson (2004) point out that lack of cointegration of the time series path of imports and exports may also indicate technological shock or productivity gap between the economy and the rest of the world. Conversely, the necessary and sufficient condition (strong form) for the intertemporal budget constraint holds when β_1 equals 1 and e_t is stationary. It follows, then, that Equation 2 provides a framework for determining the sustainability of the current account deficit (surplus).

3. Estimation technique

To examine the long-run relationship between exports and imports, we employ the autoregressive distributed lag (ARDL) cointegration procedure introduced by Pesaran *et al* (1999). The ARDL or bounds testing approach to cointegration has certain econometric advantages compared to other methods. It has better small sample properties than the Johansen and the Engle and Granger approaches.

Moreover, Pesaran *et al* (2001) show that, within the ARDL framework, the ordinary least squares estimators of the short-run parameters are consistent and the ARDL based estimators of the long-run coefficients are super consistent in small samples. In addition, the ARDL methodology is relieved from establishing the order of integration of the time series since it is applicable whether the time series are stationary, $I(0)$, nonstationary, $I(1)$, or fractionally integrated.

The unrestricted error correction (UEC) representation of the ARDL (p, q) model for equation 2 is given by:

$$\Delta LIM_t = \varphi_0 + \sum_{i=1}^{p} \beta_i \Delta LEX_{t-i} + \sum_{i=1}^{q} \delta_i \Delta LEX_{t-i} + \varphi_1 LM_{t-i} + \varphi_2 LEX_{t-i} + \mu_t \quad (3)$$

LEX and *LIM* are natural logarithm of real exports and real imports, respectively. The ARDL procedure involves two stages. At the first stage, the existence of the long-run relation among the variables of equation 3 is tested by computing *F* or Wald-statistic for testing the significance (exclusion) of the lagged levels of the variables in the error correction form of the underlying ARDL model (i.e., H_O: $\varphi_1 = \varphi_2 = 0$). Since the observations in this study are annual, we choose up to two lags of first difference of each variable and compute *F*-Statistics.

Note that the asymptotic distribution of *F*-statistic is nonstandard irrespective of whether the variables are $I(0)$ or $I(1)$. Pesaran *et al*. (1999) provide two sets of asymptotic critical values for different numbers of regressors (k), one set assumes that all variables in the ARDL are $I(0)$ and the other assumes all variables are $I(1)$ This provides a band covering all possible classifications of the variables into $I(0)$, $I(1)$, or fractionally integrated.

If the computed *F*-statistic falls outside the critical bounds, a conclusive decision can be made regarding cointegration without knowing the order of integration of the regressors.[2] For instance, if the estimated *F*-statistic is higher than the upper bound of the critical values at conventional levels of significance, then the null hypothesis of no cointegration is rejected. The test is inconclusive if the computed statistic falls within the critical value bands.[3]

Once a long-run relationship is established, then the second stage in the analysis would be to compute the long-run and the error correction estimates of the ARDL model obtained from equation 3. We estimate equation 3 using two lag selection criterions, Schwarz Bayesian Criterion (SBC) and Akaike Information Criterion (AIC).

[2] If the order of integration of a variable is greater than one, then the underlying assumptions of the ARDL are violated.

[3] Narayan (2004) argues that existing critical values are based on large sample sizes and cannot be used for small ones. He generates and reports a new set of critical values for sample sizes ranging from 30 to 80 observations. Both sets of critical values are reported in this study.

4. Data and empirical results

Exports and imports of goods and services for all countries in this study are measured in constant 2000 US dollars, obtained from World Bank-World Development Indicators CD-ROM (2008), except for Israel and Jordan where nominal exports and imports are obtained from International Financial Statistics CD-ROM (2009) and are deflated by the unit value index of exports and imports (2005=100), respectively. All series are transformed to natural logarithm. Stationarity tests were carried out for all series using Philips-Perron (PP) unit root tests.

Table 1 provides the unit root test results for *LIM* and *LEX* for all countries. In first difference, the null hypothesis that *LIM* and *LEX* series have a unit root is rejected at 5% level of significance for all countries, hence, we conclude that all first differenced variables are stationary or I(0).

Table 1: Philips-Perron (PP) unit root tests

Country	1st difference Variable	Deterministic Term *c*	*c*, *t*
Algeria	*LIM*	-8.57	-8.48
	LEX	-10.0	-10.3
Egypt	*LIM*	-8.57	-8.48
	LEX	-4.52	-4.57
Iran	*LIM*	-8.57	-8.48
	LEX	-4.28	-4.24
Israel	*LIM*	-8.57	-8.48
	LEX	-6.14	-6.55
Jordan	*LIM*	-8.57	-8.48
	LEX	-4.97	-5.89
Morocco	*LIM*	-8.57	-8.48
	LEX	-7.42	-10.1
Sudan	*LIM*	-8.57	-8.48
	LEX	-5.41	-5.61
Syria	*LIM*	-8.57	-8.48
	LEX	-6.07	-5.94
Tunisia	*LIM*	-8.57	-8.48
	LEX	-7.28	-8.34

Notes: c = intercept and t = trend; the 5% critical values for *c* and *c*, *t* are -2.93 and -3.51, respectively.

The time series for each country and the null hypothesis of no cointegration and other results are listed in Table 2. For 5 countries —Algeria, Sudan, Syria, Egypt, and Morocco— the calculated *F*-statistics are less than the critical value; hence, we fail to reject the null hypothesis of no cointegration, while the null of no cointegration is rejected for Jordan, Iran, Israel and Tunisia. As a cross check on the robustness of the bounds test for cointegration, Table 2 also reports Hansen's Instability Test. Hansen (1992) outlines a test of the null hypothesis of cointegration against the alternative of no cointegration. He notes that under the alternative hypothesis of no cointegration, one should expect to see evidence of parameter instability. He proposes the use of the L_c test statistic, which arises from the theory of Lagrange Multiplier tests for parameter instability. The L_c test results do not reject the null hypothesis that exports and imports are cointegrated at conventional levels for Jordan, Israel, Iran, and Tunisia.

Our results imply that the growth in international indebtedness is unsustainable and the discrepancy between exports and imports for Algeria, Sudan, Syria, Egypt, and Morocco may explode over time. The results also imply that the intertemporal budget constraints for Jordan, Israel, Iran, and Tunisia are not being violated.

These results are in contrast to those of Arize (2002) who found evidence of cointegration for Egypt, Iran, Israel, and Tunisia and no evidence of cointegration for Jordan and Morocco using Johansen's technique (Algeria, Sudan, and Syria were not included in Ariz's study). In addition, our results also differ with those of Tang (2006), who found evidence of cointegration between exports and imports for Morocco and no cointegration for Algeria, Egypt, Iran, Jordan, Syria, and Tunisia (Israel and Sudan were not included in Tang's study).

Table 2 also provides long-run estimates based on ARDL (SBC and AIC criterions) and two other long-run estimators: Fully Modified Phillips-Hansen, *FMPH*, and the dynamic OLS, *DOLS*. The results of the three approaches of long-run estimates are all positive, statistically significant, and very similar for all countries, demonstrating the robustness of the results.[4] Finally, Table 2 reports Wald test results obtained from testing the null hypothesis that the estimated coefficient on *LEX* is unity (H_o: β_1=1) based on ARDL regression results. The hypothesis that β_1=1 is not rejected for Iran and Israel suggests that for both countries, one dollar of imports is balanced by one dollar of exports, resulting in long-run trade balance. Nevertheless, short-run imbalances do occur since exports and imports may drift apart in the short-run.[5] The estimates of the error correction term are also reported in Table 2. They all carry the expected negative sign and are significant. This helps reinforce the findings of cointegration as reported by the *F*-test.

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[4] With the exception of Iran, where the long run elasticity is significant at 10% when estimated by the ARDL and insignificant by the other two approaches.

[5] A necessary condition for a country to obey its intertemporal budget constraint is not that β_1=1 but that a cointegration exists between imports and exports (Husted 1992).

Table 2: Tests for cointegration and long-run elasticities

	Dep. variable *F*-test			*Long-run elasticities*, β1 [t-val.]				Wald Test	H_O: β_1=1 (PHFM)	EC_{t-1} [t-value]
Country	F_{LM} (LM|LX)	F_{LX} (LX|LM)	L_c (prob.)	ARDL		PHFM	DOLS	H_O: β_1=1 (ARDL)		
				SBC	AIC					
Algeria (1960-2005)	2.195	0.589	0.724	-	-	-	-	-	-	-
			(<0.01)							
Egypt (1960-2006)	1.155	1.103	0.703	-	-	-	-	-	-	-
			(<0.01)							
Iran (1965-2006)	2.696	5.078*	0.398	1.194	0.792	0.23	0.172	0.073	5.59*	-0.138
			(0.077)	[0.09]	[0.19]	[0.75]	[0.80]	(0.78)	(0.018)	[2.323]
Israel (1960-2006)	5.864*	1.141	0.026	0.656	0.682	0.956	0.986	2.197	0.0626	-0.189
			(>20)	[3.11]	[3.181]	[13.97]	[18.37]	(0.138)	(0.802)	[1.988]
Jordan (1969-2007)	7.511*	22.51**	0.12	0.705	0.709	0.777	0.728	54.84*	155.83*	-0.467
			(>0.20)	[21.97]	[18.09]	[22.46]	[33.57]	(0.00)	(0.00)	[3.005]
Morocco (1960-2006)	3.789	1.258	0.422	-	-	-	-	-	-	
			(0.066)							
Sudan (1960-2006)	0.053	1.27	0.873	-	-	-	-	-	-	
			(<0.01)							
Syria (1975-2006)	4.074	3.434	0.298	-	-	-	-	-	-	
			(0.15)							
Tunisia (1961-2005)	6.080*	4.172	0.174	0.857	-	0.83	0.852	16.17*	18.48*	-0.32
			(>0.20)	[20.92]		[29.59]	[24.805]	(0.00)	(0.00)	[4.828]

Notes: The upper bound critical values for the *F*-test at the 5% and 1% are 5.764 and 7.815, respectively (Pesaran and Pesaran, 2009) and 4.460 and 6.193 for *n*=45, 4.523 and 6.333 for *n*=40, and 4.663 and 6.760 for *n*=30, respectively (Narayan 2004). * and ** denotes significance at the 5% and 1%, respectively.

Finally, to examine the stability of the long run coefficients together with short run dynamics, we follow Pesaran and Pesaran (1999) and apply cumulative sum (CUSUM) and cumulative sum of square (CUSUMSQ).

Specifically, the CUSUM test makes use of the cumulative sum of recursive residuals based on the first set of *n* observations and is updated recursively and plotted against break points. If the plot of CUSUM statistics stays within the critical bounds of 5% significance level (represented by a pair of straight lines drawn at the 5% level of significance), the null hypothesis that all coefficients in the error correction model are stable cannot be rejected. If either of the lines is crossed, the null hypothesis of coefficient constancy can be rejected at the 5% level of significance. A similar procedure is used to carry out the CUSUMSQ test, which is based on the squared recursive residuals. Figures 1-4 show a graphical representation of the CUSUM and CUSUMSQ plots applied to the error correction model for Iran, Israel, Jordan and Tunisia. Neither CUSUM nor CUSUMSQ plots cross the critical bounds, indicating no evidence of any significant structural instability. [6]

Figure 1: Plot of CUSUM and CUSUMSQ for Iran

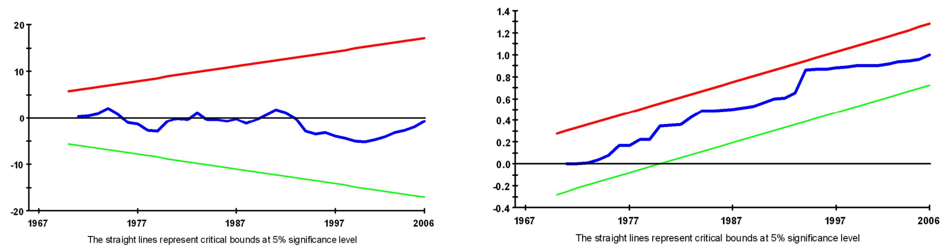
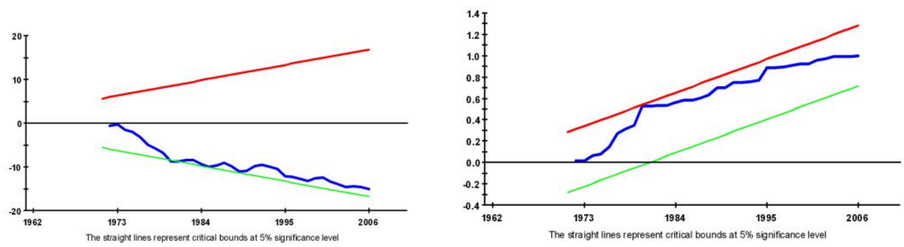


Figure 2: Plot of CUSUM and CUSUMSQ for Israel



[6] Due to instability in CUSUM and CUSUMSQ for Israel between 1973 and 1992, a dummy variable, D73, which takes the value of one from 1973-1992 and zero otherwise, is included in the estimations. Cointegration tests are not sensitive to the inclusion of this dummy variable.

Figure 3: Plot of CUSUM and CUSUMSQ for Jordan

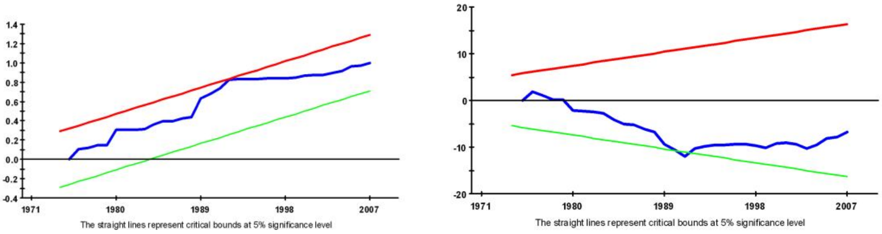
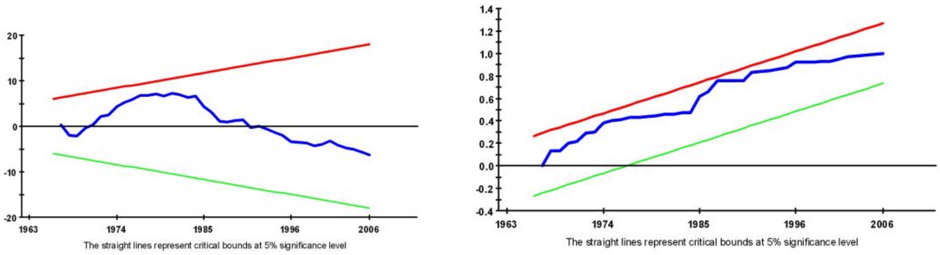


Figure 4: Plot of CUSUM and CUSUMSQ for Tunisia



5. Conclusion

The aim of this article is to apply the bounds testing approach to cointegration to investigate whether exports and imports are cointegrated for nine MENA countries. We find that for four countries examined, Iran, Israel, Jordan, and Tunisia, exports and imports are cointegrated. For these countries, we use three different techniques – the ARDL, Philip–Hansen and DOLS, to estimate the long-run elasticities. Our results show that the elasticities are positive and statistically significant in three of the four countries. While the presence of cointegration is necessary (an indication of a "weak" sustainability of current account deficits), it is not a necessary and sufficient condition for an economy to satisfy its intertemporal budget constraint since the long-run coefficient is unity only in the cases of Iran and Israel. This implies that Iran and Israel satisfy the "strong" form of their intertemporal budget constraint and hence the current account deficit is sustainable over the long-run, while Jordan and Tunisia satisfy only the weak form.

The cointegration results were complemented with the CUSUM and CUSUMSQ tests for parameter stability. The results suggest that the parameters are stable over time.

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