Cointegration, Structural Changes and the Relationship between Trade and Economic Growth in Tunisia

GUENICHI, Hassan

Higher Institute of Computer Sciences and Management, Kairouan University; Tunisia

December 2008

Cointegration, Structural Changes and the Relationship between Trade and Economic Growth in Tunisia

Hassen Guenichi, Mounastir University

Abstract

This paper examines the major determinants of GDP growth in Tunisia using quarterly time series data spanning from 1960 to 2003. The Tunisian economy has been subject to a multitude of structural changes and regime shifts during the sample period. Thus, time series properties of the data are first analyzed by Zivot-Andrews (1992) model. The empirical results based on this model indicate the presence of unit roots for all of the variables under investigation. Taking into account the resulting endogenously determined structural breaks; the Saikkonen and Luetkephol (2000) and Johansen and al (2001) cointegration approach is then employed to determine the long-run drivers of economic growth. This cointegration technique accommodates potential structural breaks that could undermine the existence of a long-run relationship between GDP growth and its main determinants. Empirical estimates based on Quintos (1995) and Johansen (1993) approches indicate that in the long-term, policies aimed at promoting various types of physical investment, human capital, trade openness and technological innovations will improve economic growth.

Keywords

Structural Break, Unit Root Tests, Cointegration technique, Trade and Tunisian Economic Growth

Introduction

Feder (1982), Balassa (1985) and Ghatak et al. (1997) suggested that export expansion might generate positive externality through more efficient allocation of resources, efficient management and improved production techniques, specialization, competition and the economy of scale. Hence various development theories have emerged in the literature suggesting that export expansion further accelerates economic growth due to the above-mentioned factors. This is referred to as the export-led growth (ELG) hypothesis. Endogenous growth models make use of the same idea to analyze the broad externality effects of exports on the economy, but they address the role of imports as well. These models emphasize the fact that trade works as a conduit of knowledge spillover. In turn, this knowledge spillover enables the economy to achieve increasing returns, and human capital also has a role in increasing economic growth through the same knowledge spillover effect of trade (Sengupta, 1993). In fact, according to the endogenous growth theory factors such as: physical capital (R&D effects), human capital (representing knowledge spillover effects), exports expansion (proxying positive externality effects), and capital and intermediate imports (capturing learning-by-doing effects) are the major determiners of economic growth.

Following empirical studies of the sources of growth by such researchers as Ram(1987), Sengupta (1993), Van Den Berg (1997), and Ibrahim and MacPhee (2003) and which have followed the Feder (1982) model, we include export in the typical production function. In addition, like Ven Den Berg, we include total imports as a new factor in the production function. The structure of the rest of the paper is as follows. Section II presents the review of the literature. The model, data and methodology are presented in section III. We explains first unit root test based on the Zivot-Andrews (1992) model, which take into account the existence of potential structural breaks in the data and second cointegration analysis in the presence of pre-determined structural breaks using the Saikkonen and Lutkephol (2002) and Johansen and al(2001) cointegration test and the Quintos (1995)and Johansen (1993) VECM estimation approach. Finally, section IV presents the empirical results and the economic interpretations. We ended this paper with some concluding remarks.

Review of the Literature

M. Dritsaki, C. Dritsaki and A. Adamopoulos (2004) investigated the relationship between Trade, Foreign Direct Investment (FDI) and economic growth for Greece over the period 1960-2002. Their methodology is based on VAR model and the cointegration approch. The Cointegration analysis suggested that there is a long-run equilibrium relationship. The results of Granger causality test showed that there is a causal relationship between the examined variables.

F. Abou-Stait (2005) examined the export-led growth (ELG) paradigm for Egypt, using historical data from 1977 to 2003. During this period, Egypt changed its economic philosophy from central planning and government intervention to one based on a free market economy. The paper employs a variety of analytical tools, including cointegration analysis, Granger causality tests, and unit root tests, coupled with vector auto regression (VAR) and impulse response function (IRF) analyses. The paper sets three hypotheses for testing the ELG paradigm for Egypt, (i) whether GDP, exports and imports are cointegrated, (ii) whether exports Granger cause growth, (iii) whether exports Granger cause investment. The paper fails to reject the first two hypotheses, while it fails to accept that exports Granger cause investment. In addition to the analysis of the 1977-2003 period, the paper looks briefly also at the impact of the economic reform undertaken in 1991, and weather the ELG hypothesis still holds during the 1991-2003 sub-period.

A. Abdulai and P. Jaquet (2OO2) examined the short- and long-run relationship between economic growth, exports, real investments and labor force for Cote d'Ivoire for the period 1961-97, using cointegration and error correction techniques. The results indicate that there is one long-run equilibrium relationship among the four variables, and the causal relationship flows from the growth in exports to the growth in GDP both in the short and long run, providing support for the export-led growth hypothesis. This finding suggests that the recent trade reforms aimed at promoting domestic investment and restoring international competitiveness to expand and diversify exports have the potential of increasing economic growth in the future. The same work is made by J. Balaguer and M. Cantavella-Jordá (2002) on Spanish data base during 1961-2000 periods and by E.M. Ekanayake (1999) on Asiatic countries data base (India, Indonesia, Korea, Pakistan, Philippines, Sri Lanka and Thailand) during 1960-1997 period.

Data model and methodology

Data and model

In this paper, we propose a framework based on the conventional neo-classical one-sector aggregate production technology where we treat capital, labor, total imports and total exports as separate inputs to.
That is:

$$Y = F(K, L, X, M)$$

This model is a kind of production function, which is augmented by the addition of trade factors, exports (X) and imports (M). It should be noted that in Feder-type models, GDP is considered to be simply a function of ordinary labor force growth together with other relevant factors. We follow the endogenous growth theory and consider instead, human capital (the number of employed workforce with a university degree) rather than the total labor force in our empirical models. The following modified model in logarithm form is used to examine the trade-growth nexus in developing economy like Tunisia:

$$Ln(Y_t) = \beta_0 + \beta_1 Ln(K_t) + \beta_2 Ln(L_t) + \beta_3 Ln(X_t) + \beta_4 Ln(M_t) + \varepsilon_t$$

Where *Y* = aggregate output or real GDP, *K* is the capital stock, *L* is the level of employment, M is a total imports, X is the total exports and the subscript t denotes the time period. The data are collected from the WDI CD-ROM, and the International Financial Statistics (IFS).

Methodology

We start our empirical analysis by unit root test based on the Zivot-Andrews (1992) model, which take into account the existence of potential structural breaks in the data. Then we discus the results of cointegration analysis in the presence of pre-determined structural breaks. First we test for cointégration using Saikkonen and Lütkepohl (2000a) and Johansen and al (2001) procedures. Second we estimate the VEC model using the Quintos (1995) and Johansen (1993) approaches.

Unit Roots Tests with Structural Break

The issue of structural break is of considerable importance in the analysis of macroeconomic time series. Such breaks occur in many time series for any number of reasons and this makes it difficult to test the null hypothesis of structural stability against the alternative of a one-time structural break. When present in the data generating process, but not allowed for in the specification of an econometric model, results may be biased towards the erroneous non rejection of the non-stationary hypothesis (Perron 1989; Perron 1997; Leybourne and Newbold (2003). Perron (1989, 1994, 1997) and Zivot-Andrews (1992) attempt to overcome this difficulty. In the following section, The Zivot-Andrews methodology for testing the unit root hypothesis in the presence of structural break is explained and then this method is applied for the variables under investigation.

Zivot-Andrews unit root test with structural break

Zivot and Andrews (ZA, 1992) propose a variation of Perron's (1989) original test in which the time of the break is estimated, rather than known as an exogenous phenomenon. The null hypothesis in their method is that the variable under investigation contains a unit-root with a drift that excludes any structural break, while the alternative hypothesis is that the series is a trend stationary process with a one-time break occurring at an unknown point in time. By endogenously determining the time of structural breaks, ZA argue that the results of the unit root hypothesis previously suggested by earlier conventional tests such as the ADF test may change.

In this methodology, *TB* (the time of break) is chosen to minimize the one-sided t-*s*tatistic of α=1. In other words, a break point is selected which is the least favorable to the null hypothesis. The ZA model endogenises one structural break in a series (such as y_t) as follows:

H0: $$y_t = \mu + y_{t-1} + e_t$$

H1: $$\Delta y_t = \mu + \beta t + \theta DU1_t + \gamma DT1_t + \alpha y_{t-1} + \sum_{i-1}^{k} c_i \Delta y_{t-i} + \varepsilon_t$$

Equation (4), which is referred to as model C by ZA, accommodates the possibility of a change in the intercept as well as a trend break. ZA also consider two other alternatives where a structural break impacts on the intercept only (model A) or trend only (model B). Model C is the least restrictive compared to the other two models; we thus base our empirical investigation on this model. In equation (4) $DU1_t$ is a sustained dummy variable capturing a shift in the intercept, and $DT1_t$ is another dummy variable representing a shift in the trend occurring at time *TB1*. The alternative hypothesis is that the series, y_t, is I(0) with one structural break. *TB* is the break date, and the dummy variables are defined as follows:

$$DU1_t = \begin{cases} 1 \text{ if } t > TB1 \\ 0 \text{ if } t \le TB1 \end{cases} \quad \text{and} \quad DT1_t = \begin{cases} t - TB1 \text{ if } t > TB1 \\ 0 \text{ if } t \le TB1 \end{cases}$$

The null is rejected if the α coefficient is statistically significant. The optimal lag length is determined on the basis of the t-test or SBC. The "trimming region" where we search for the minimum t-ratio is assumed to be within 0.05T-0.95T or $0.05T \le TB1 \le 0.95T$.

Cointegration Analysis with Structural breaks

Cointegration test with structural breaks

As had been noted as far back as 1989 by Perron, ignoring the issue of potential structural breaks can render invalid the statistical results not only of unit root tests but of cointegration tests as well. Kunitomo (1996) explains that in the presence of a structural change, traditional cointegration tests, which do not allow for this, may produce "spurious cointegration". In the present research, therefore, considering the effects of potential structural breaks is very important, especially because the World economy has been faced with structural breaks like revolution and war in addition to some policy changes.

Saikkonen and Lütkepohl (2000a, b, c) and Johansen and al (2001) have proposed a test for cointegration analysis that allows for possible shifts in the mean of the data-generating process. Because many standard types of data generating processes exhibit breaks caused by exogenous events that have occurred during the observation period, they suggest that it is necessary to take into account the level shift in the series for proper inference regarding the cointegrating rank of the system.

SL and Johansen argued that "structural breaks can distort standard inference procedures substantially and, hence, it is necessary to make appropriate adjustment if structural shifts are known to have occurred or are suspected" (2000b: 451). The Saikkonen and Lütkepohl (SL) test investigates the consequences of structural breaks in a system context based on the multiple equation frameworks of Johansen-Jeslius, while earlier approaches like Gregory-Hansen (1996) considered structural break in a single equation framework and others did not consider the potential for structural breaks at all.

According to Saikkonen and Lütkepohl (2000b) and Lütkepohl and Wolters (2003), an observed n-dimensional time series yt = (y1t,…., ynt), yt is the vector of observed variables (t=1,…, T) which are generated by the following process:

$$y_t = \mu_0 + \mu_1 t + \gamma_1 d_{1t} + \gamma_2 d_{2t} + \gamma_3 d_{3t} + \delta Dt_{0t} + \delta_2 Du_{1t} + x_t$$

Where *DT0t* and *DU1t* are impulse and shift dummies, respectively, and account for the existence of structural breaks. *DT0t* is equal to one, when t=T0, and equal to zero otherwise. Step (shift) dummy (*DU1t*) is equal to one when (t>T1), and is equal to zero otherwise. The parameters

$\gamma(i = 1,2,")$, μ_0, μ_1, and δ are associated with the deterministic terms. The seasonal dummy variables *d1t, d2t,* and *d3t,* are not relevant to this research since our data are yearly. According to SL (2000b), the term *xt* is an unobservable error process that is assumed to have a VAR (p) representation as follows:

$$xt = A_1 x_{t-1} + ... + A_p x_{t-p} + \varepsilon_t \qquad \text{t=1,2}$$

By subtracting xt-1 from both sides of the above equation and rearranging the terms, the usual error correction form of the above equation is given by:

$$\Delta x_t = \Pi\, x_{t-1} + \sum_{j-1}^{p-1} \Gamma j \Delta x_{t-j} + u_t$$

This equation specifies the cointegration properties of the system. In this equation, *ut* is a vector white noise process; *xt= yt -Dt* and *Dt* are the estimated deterministic trends. The rank of Π is the cointegrating rank of xt and hence of yt (SL, 2000b). The possible options in the SL procedure, as in Johansen, are three: a constant, a linear trend term, or a linear trend orthogonal to the cointegration relations. In this methodology, **t**he critical values depend on the kind of the above-mentioned deterministic trend that included in the model. More interestingly, in SL, the critical values remain valid even if dummy variablesare included in the model, while in the Johansen test; the critical values are available only if there is no shift dummy variable in the model. The SL approach can be adopted with any number of (linearly independent) dummies in the model. It is also possible to exclude the trend term from the model; that is, μ=0 maybe assumed *a priori*. In this methodology, as in Johansen's, the model selection criteria (SBC, AIC, and HQ) are available for making the decision on the VAR order. In the following section, we have applied SL tests for the cointegration rank of a system in the presence of structural breaks.

Estimation of the cointegration relationships:

The Johansen's procedure apply the likelihood maximum (LM) on VAR model assuming that errors is *iid*.

$$Y_t = A_1 Y_{t-1} + ... + A_k Y_{t-k} + U_t, \qquad t = 1,..., T$$

Where Y_t is an n-vector of I(1) variables.
We can rewrite Y_t as follow :

$$\Delta Y_t = B_1 Y_{t-1} + B_2 \Delta Y_{t-1} + ... + B_k \Delta Y_{t-k+1} + U_t$$

Where $B_1 = -I + \sum_{i=1}^{k} A_i$ and $B_j = -\sum_{i=j}^{k} A_i$ with j= 2,…, k.

The variables $\Delta Y_t,...,\Delta Y_{t-k+1}$ are all I(0) but Y_{t-1} is I(1), in order that this equation be consistent, β_1 should not be a full rank. Let its rank r and let write

$$B_1 = \alpha\beta'$$

Where α is an n×r matrix and β' is an r×n. Then, $\beta' Y_{t-1}$ are the cointegrated variables, β' is the matrix of coefficients of the cointegrating vectors and α has[2]the interpretation of the matrix of error correction terms.

Since our interest α and β' we eliminate $B_2,...,\ B_k$ first. To do this we proceed at follow. Regress ΔY_t on $\Delta Y_{t-1},...,\Delta Y_{t-k+1}$. Get the residuals. Call them R_{0t} Y_{t-1} . regress on these same variables. Get the residuals. Call them R_{1t}. Now, our regression equation is reduced to

$$R_{0t} = \alpha\beta' R_{1t} + u_t$$

This is a multivariate regression problem. Define

$$\begin{bmatrix} S_{00} & S_{01} \\ S_{10} & S_{11} \end{bmatrix}$$

As the matrix of sums of squares and sums of products of R_{0t} and R_{1t}. Johansen (1991) shows that the asymptotic variance of $\beta' R_{1t}$ is $\beta' \Sigma_{11} \beta$ the asymptotic variance of R_{0t} is Σ_{00} and the asymptotic covariance matrix of $\beta' R_{1t}$ and R_{0t} is $\beta' \Sigma_{1n}$ where Σ_{00}, Σ_{10} *et* Σ_{11} are the population counterparts of S_{00}, S_{10} *et* S_{11}.

We shall maximize the likelihood function with respect to α holding β constant and then maximize with respect to β in the second step. We get

$$\hat{\alpha}' = (\beta' S_{11} \beta)^{-1} \beta' S_{10}$$

Note that $\hat{\alpha}'$ is an r*n matrix and the conditional maximum of the likelihood function is given by :

$$[L(\beta)]^{-2/T} = \left| S_{00} - S_{01}\beta(\beta' S_{11}\beta)^{-1} \beta' S_{10} \right|$$

Maximization of the likelihood function with respect to β implies minimization of the determinant with respect to β. We will minimize

$$\frac{\left| \beta' S_{11}\beta - \beta' S_{10} S_{00}^{-1} S_{01}\beta \right| . \left| S_{00} \right|}{\left| \beta' S_{11}\beta \right|}$$

But

$$\min_X \frac{\left| X'(A_1 - A_2)X \right|}{\left| X' A_1 X \right|}$$

is given by the maximum characteristic root of the equation $\left|A_2 - \lambda A_1\right| = 0$. Thus, substituting $A_1 = S_{11}$ and $A_2 = S_{10}S_{00}^{-1}S_{01}$ we get the maximum of the likelihood function by solving the eigenvalue problem

$$\left|S_{10}S_{00}^{-1}S_{01} - \lambda I\right| = 0$$

Or finding the eigenvalue of $$\left|S_{11}^{-1}S_{10}S_{00}^{-1}S_{01} - \lambda I\right| = 0 \qquad (1)$$

But the roots of this equation are the r canonical correlations between R_{1t} and R_{0t}. If the eigenvalues of A are λ_i, the eigenvalues of $(I - A)$ are $(1-\lambda_i)$. Hence if λ_i are the canonical correlations given by solving equation (1), then $(1-\lambda_i)$ are the eigenvalues of $\left(I - S_{11}^{-1}S_{10}S_{00}^{-1}S_{01}\right)$.

The value of the determinant of the matrix is equal to the product of its eigenvalues, we have

$$\prod_{i=1}^{n}(1-\lambda_i) = \left|I - S_{11}^{-1}S_{10}S_{00}^{-1}S_{01}\right| = \frac{\left|S_{11} - S_{10}S_{00}^{-1}\right|}{\left|S_{11}\right|}$$

Hence

$$L_{\max}^{-2/T} = \left|S_{00}\right|.\prod_{i=1}^{n}(1-\lambda_i)$$

Johansen propose two statistics to determine the cointegration rank

$$\lambda_{trace} = -T \sum_{i=r+1}^{n} \ln(1 - \hat{\lambda}_i)$$

$$\lambda_{\max} = -T \ln(1 - \hat{\lambda}_{r+1})$$

In structural changes cases we follow the approach of Johansen (1993) and Quintos (1995).

The two procedures start from the equation:

$$R_{0t} = \alpha\beta' R_{1t} + u_t$$

We can rewrite this equation as follow:

$$R_{0t} = \Pi R_{1t} + u_t$$

Quintos separates the sample into different periods assuming the break dates known. For instance, let there be one break date and let Π and (Π_1, Π_2) be the parameters for the whole sample and the split samples. The hypothesis is

$$H_0 = Rang(\Pi_1) = Rang(\Pi_2) = Rang(\Pi)$$

Empirically, we estimate the model in the two regimes and show the cointegration rank in each regime.

Empirical Results

Zivot and Andrews Unit root test

Based on the results reported in Tables 1 and 2, the primary findings of the analysis are as follows. First, the results of the ZA models indicate that all series under investigation are non-stationary. Second, the timing of any structural break (T_b) for each series using the ZA approach is also shown in Table 1. The computed break dates correspond closely with the expected dates associated with the effects of the oil boom in 1974, and the effects of dept crises in developing countries in 1982. Third, the reported t statistics in Table 1 for μ, β, θ , γ and *a* are significant in the majority of cases. Given the fact that all of the estimated coefficients for the indicator and trend dummy variables are statistically significant one can argue that the estimated structural break dates are indeed statistically significant.

Variables	TB	μ	β	θ	γ	α	Causes for TBs
Ln(y)	52	-0.2159	0.1660	-0.0482	1.0359	0.0267	Oil shock
	1/10/1978	(-3.9165)	(0.6694)	(-3.5377)	(2.9637)	(2.8265)	
Ln(X)	51	-0.1157	0.1716	-0.0079	-0.5355	0.0003	Oil shock
	(1/7/1972)	(-3.1648)	(0.6211)	(-0.8439)	(-1.5359)	(0.0333)	
Ln(M)	55	-0.1700	0.7777	0.0159	-0.6019	-0.0216	Oil shock
	1/7/1973	(-4.0475)	(2.3926)	(1.2925)	(-1.6735)	(-1.6032)	

Table 1. The Zivot-Andrews test results:

Ln(K)	103 1/8/1973	-0.1518 (-3.8613)	-0.1955 (-1.0383)	0.0004 (0.1416)	0.7386 (2.4567)	0.0178 (2.1363)	Oil shock
Ln(L)	90 1/4/1982	-0.1124 (-3.6474)	-0.3803 (-1.7546)	0.0165 (2.9722)	-1.5926 (-3.961)	-0.0020 (-0.3092)	Dept crises in developing countries

Johansen cointegration test results

As explained above, Johansen (2000b) derived the likelihood ratio (LR) test in order to determine the number of cointegrating relations in a system of variables, by allowing for the presence of potential structural breaks. We now apply a maximum likelihood approach for testing and determining the long-run relationship in the model under investigation. As mentioned earlier, in this procedure Johansen assumed that the break point is known a priori. In the last section, we determined the time of the break endogenously by Zivot-Andrews (1992) procedure. The empirical result based on this method showed that the most significant break for variables of under investigation are consistent with time of oil shock. Therefore, at this stage we include one dummy variable of regime change in order to take into account the structural breaks in the system. Following the Johansen procedure we consider three cases: impulse dummy and shift with intercept included; impulse dummy and shift with trend and intercept included; and finally, impulse dummy and shift with a trend statistically independent (orthogonal) to cointegration relation included. The cointegration results in these three cases are presented in tables 2.

The optimal number of lags is determined by AIC and SC, which is more appropriate for the short span of the data. The hypothesis of the long-run relationship among non-stationary variables is tested and the result is reported in table 2. These tables indicates that the hypothesis of no cointegration r=0 and one cointegration vector r=1 are rejected at the10%, 5% and 1% significance level. The existence of two cointegration vectors is not rejected in any of the three cases mentioned above.

Table 2 :Saikkonen and Lutkephol and Johansen and al cointegration test results

Intercept included (C)						Intercept and trend included (C/T)						Trend orthogonal to cointegration relation (C/O)					
r0	LR	pval	90%	95%	99%	r0	LR	pval	90%	95%	99%	r0	LR	pval	90%	95%	99%

0	322.15	0.0000	82.19	85.80	92.83	0	382.58	0.0000	90.86	95.35	104.14	0	307.84	0.0000	65.73	69.61	77.29
1	155.28	0.0000	58.45	61.56	67.68	1	215.59	0.0000	65.49	69.35	76.98	1	144.25	0.0000	44.45	47.71	54.23
2	28.11	0.1089	38.69	41.33	46.57	2	28.38	0.1758	44.06	47.29	53.76	2	24.95	0.0743	27.16	29.80	35.21
3	15.14	0.5872	22.80	24.97	29.39	3	14.23	0.6234	26.42	29.01	34.29	3	7.79	0.4951	13.42	15.41	19.62
4	7.11	0.3449	10.82	12.65	16.57	4	6.57	0.5631	12.41	14.28	18.24	4	5.27	0.7631	11.22	14.78	17.44

Quintos(1995) and Johansen (1993) estimation approaches results

First regime

From the β vectors we can see that the coefficient on labor in the first cointegrating vector is insignificant. Testing the exclusion of labor from the first cointegrating relationship yields a likelihood ratio test = 2.54, which compared to the 5% critical value $\chi2(4) = 5.99$ enables us to easily accept the null hypothesis. The results indicate that the model is now completely identified. We estimate a vector-error-correction (VEC) model with two cointegrating vectors and two common stochastic trends. The cointegrating vectors are each indicating the direction where a stable, long-run equilibrium relationship exists and, the adjustment coefficients α are indicating the speed of adjustment of each variable to these long run equilibrium states.

Table 3: The β and α Vectors

Variables	β_1	β_2	α_1	α_2
Y	1	……….	-0.23681	0.21361
			[-2.3564]	[2.0029]
X	-1.058893	0.7234	0.199682	-0.00172
	[-5.0458]	[5.1043]	[4.9735]	[-0.7735]
M	0.13487	-3.1802	-0.235238	0.13217
	[2.6453]	[-7.1413]	[-3.22617]	[3.2285]
K	……….	1	-0.34685	-0.034685
			[-2.943]	[-5.901]
L	0.10456	-0.58456	0.06759	0.0759
	[0.1147]	[-6.6103]	[3.8133]	[7.1233]

Trend	0.001019	0.100258
	[1.1105]	[11.6653]
Constante	-2.1126	6.6296

Table 4 reports the results of the Granger-causality tests. These tests are conducted using a joint *F*-statistic for the exclusion of one variable from one equation as illustrated above. The results of these tests indicate that Granger-causality is running in both directions between, firstly output growth and imports and second between output growth and exports. Thus, our results for Tunisia indicate that trade have a causal impact on output growth.

Table4 : Test Results for Granger-causality

Null Hypothesis:	F-Statistic	Probability
X does not Granger Cause Y	40.6043	6.8E-08
Y does not Granger Cause X	17.1533	0.00014
M does not Granger Cause Y	8.50576	0.00537
Y does not Granger Cause M	1.97104	0.16678

The Granger-causality tests conducted above indicate only the existence of causality. They do not, however, provide any indication on how important is the causal impact that trade has on output growth. For example, when there is a shock to exports, it would also be interesting to know by how much this shock will affect the growth rates of output. In order to provide answers to these questions, we next decompose the variance of the forecast-error of output growth into proportions attributable to innovations in each variable in the system including its Consider again the vector error-correction model. A change in anyone of the random innovations $\eta i,t$, i=1, 2,… will immediately change the value of the dependent variable and, hence, will also change the future values of the remaining variables in the system through the dynamic structure of the model. Since changes in the random innovations produce changes in the future values of the variables, it is possible to decompose the total variance of the forecast-error in anyone of them and determine how much of this variance each variable explains. Since our interest focuses on the response of output growth to shocks in the factor inputs, in particular imports and exports, we only decompose the forecast-error variance of the output growth variable in response to a one standard deviation innovation in capital, labor imports and exports. Since the innovations are not necessarily totally uncorrelated, the residual terms are orthogonalized using a Choleski decomposition in order to obtain a diagonal covariance matrix of the resulting innovations and, therefore, isolate the effects of each variable on the other.

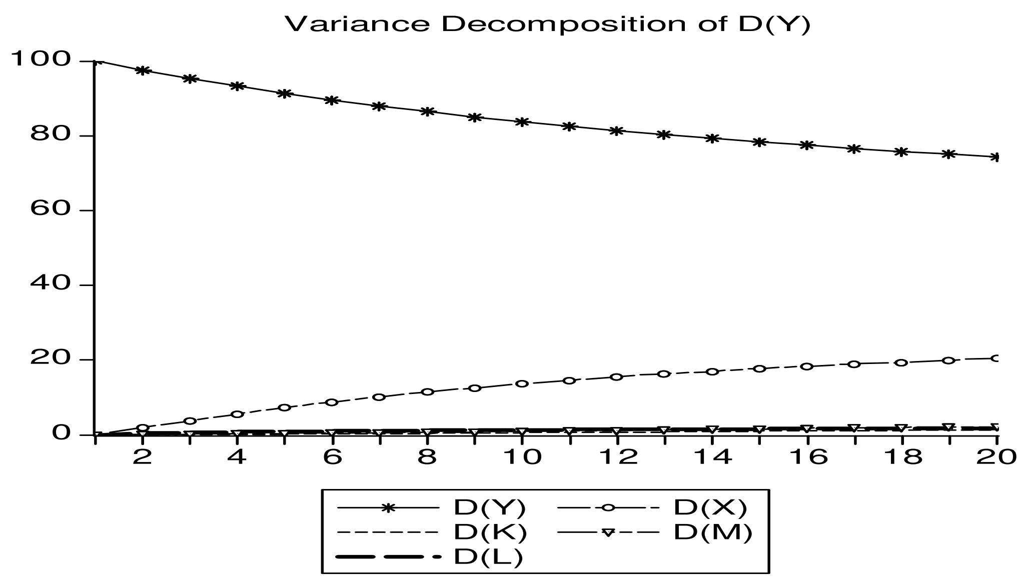
Table 5 and figure 1 report the results of the variance decomposition of output growth in Canada within a twenty period horizon. As can be seen in the table, the four factor inputs together explain about 26% of the future changes in output growth in Tunisia. The remaining 74% are due

to changes in output growth itself. Looking at the separate effects of factor inputs, exports have the highest effect on output growth followed by imports and labor then capital. In addition, shocks to imports and exports seem to generate a permanent effect on output growth.
These results confirm the assumption on the neutrality of trade and clearly illustrate how important could be the effect of imports and exports on the future growth of output.

Table5 : Results of Variance Decomposition

Periode	D(Y)	D(K)	D(L)	D(X)	D(M)
2	97.59528	0.138575	0.310929	1.882544	0.072675
4	93.29019	0.155462	0.734531	5.588535	0.231277
6	89.60841	0.202391	0.992933	8.776274	0.419992
8	86.43781	0.334636	1.170443	11.41399	0.643121
10	83.70667	0.510984	1.305788	13.58355	0.893007
12	81.34190	0.696741	1.417419	15.38595	1.157989
14	79.27655	0.873899	1.514363	16.90825	1.426934
16	77.45517	1.035563	1.601142	18.21694	1.691189
18	75.83406	1.180495	1.680093	19.36041	1.944937
20	74.37954	1.309846	1.752500	20.37332	2.184797

Figure 1. The response of output growth to a one standard deviation innovation in inputs



Second regime

In this second regime, from the β vectors we can see that the coefficient on labor in the first and second cointegrating vectors is insignificant. Testing the exclusion of labor from the first ans second cointegrating relationships yields a likelihood ratio test respectively equal to 2.54 and 1,023, which compared to the 5% critical value $\chi2(4) = 5.99$ enables us to easily accept the null hypothesis

Table 6: The β and α Vectors

Variables	β_1	β_2	α_1	α_2
Y	1	…….	-1.089456 [-5.18415]	0.142848 [3.68092]
K	……..	1	0.093745 [3.34641]	-0.010154 [-1.96274]
L	-0.03234 [-0.54451]	0.04327 [1.34641]	0.063350 [4.07936]	0.015665 [5.46250
X	0.646257 [7.16639]	-14.23574 [-4.61243]	0.303844 [3.71941]	0.089545 [5.93580]
M	-0.817596 [-4.22466]	14.74609 [5.21224]	0.153854 [1.65265]	0.110552 [6.43064]
Trend	0.001072 [4.51167]	-0.005578 [-1.60628]		
Constante	0.076356	-23.35452		

Table 7 reports the results of the Granger-causality tests. The results of these tests indicate that Granger-causality is running in both directions between, firstly output growth and imports and second between output growth and exports. Thus, in this period, our results for Tunisia indicate that trade have a causal impact on output growth

Table7 : Test Results for Granger-causality

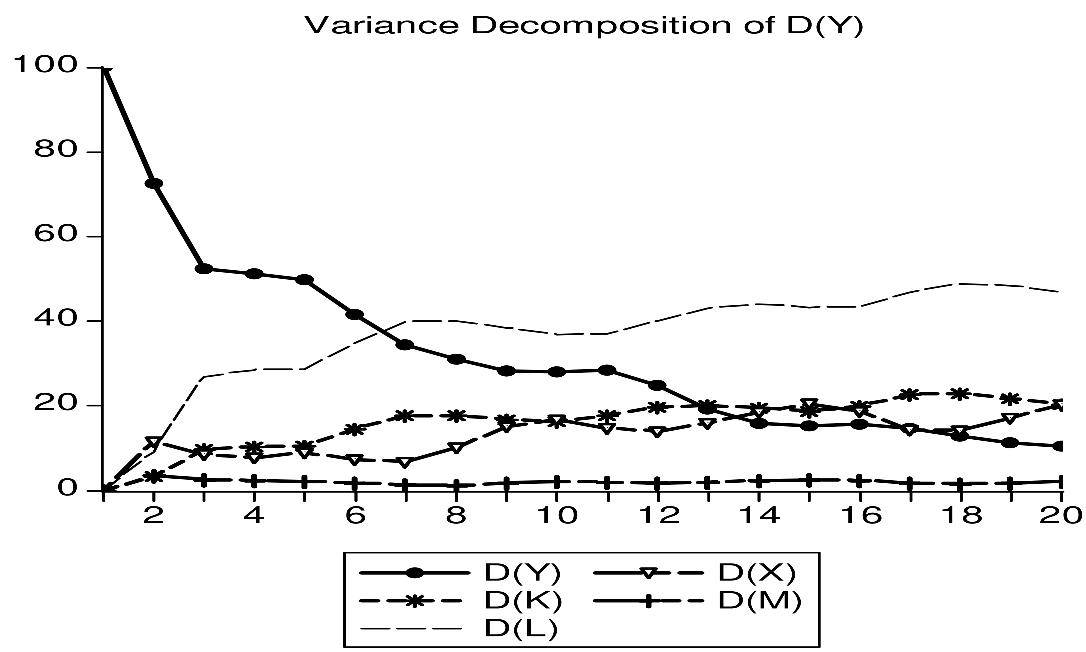
Null Hypothesis:	F-Statistic	Probability
X does not Granger Cause Y	116.914	0.00000
Y does not Granger Cause X	17.9880	1.7E-07
M does not Granger Cause Y	122.641	0.00000
Y does not Granger Cause M	30.2888	3.2E-11

Table 8 and figure 2 report the results of variance decomposition. Looking at the separate effects of factor inputs, labor has the highest effect on output growth followed by capital then exports and finally imports. About 46,8% of future changes in output growth are due to changes in labor, 20,31% due to capital, 20,18% due to exports, and 2,17 to imports.

Table8 : Results of Variance Decomposition

Period	D(Y)	D(K)	D(L)	D(X)	D(M)
2	72.60279	3.235443	9.136678	11.52853	3.496561
4	51.12988	10.35275	28.53823	7.732063	2.247072
6	41.68124	14.48786	34.82013	7.227251	1.783528
8	31.08059	17.75693	39.97063	10.05212	1.139735
10	28.00368	16.22021	36.86628	16.71802	2.191810
12	24.80359	19.66158	39.97475	13.93104	1.629040
14	15.92914	19.37813	44.04836	18.41208	2.232296
16	15.66067	19.95725	43.36729	18.65295	2.361852
18	12.84703	22.83929	48.76728	14.07589	1.470520
20	10.52492	20.31209	46.80774	20.18462	2.170630

figure 1. the response of output growth to a one standard deviation innovation in inputs



Conclusion

The objective of this paper was to examine the long-run determinants of GDP in Tunisia during the period 1960-2003 employing the Saikkonen and Lutkephol (2000) and Johansen and(2001) cointegration method. Prior to the cointegration analysis, the Zivot-Andrews (1992) test was applied in order to endogenously determine the most significant structural breaks in the major drivers of economic growth, physical and human capital, exports and imports. The empirical results based on the ZA model indicate the existence of unit root for all of the variables under investigation. Moreover, we found that the most significant structural breaks over the last forty years occurred as a result of the oil sock in 1973. These results provide complementary evidence to models employing exogenously imposed structural breaks in the Tunisian macroeconomy.

Finally, we employed the Saikkonen and Lutkephol (2000) and Johansen and al (2001.) cointegration approach to determine the long-run factors contributing to economic growth in Tunisia. It is important to use this approach in our cointegration test as during the sample period, the Tunisian economy has been subject to serious structural breaks such as: the world oil shock in 1973. In the presence of such structural breaks, the SL and Johansen cointegration tests conducted in this paper indicate that there are two cointegrating vectors which link GDP with physical and human capital, imports and exports.

Thus, based on the neo-classical one sector aggregate production technology, we developed a vector error-correction model after testing for multivariate cointegration between output, capital, labor imports and exports. The cointegration test indicates that exports and imports enter significantly the cointegration space. The study of the causal relationship between trade factors and output growth in Tunisia, the short-run dynamics of the variables show that the flow of causality is running in both directions between output growth and trade factor. Using variance decomposition of the forecast-error variance of output growth, we found that a shock to imports and exports would cause respectively a 20.37% and 2.184% changes in the future growth rates of output in the first regime and respectively 20.184% and 2.17% in the second regime. With this, our results seem to significantly reject assumption that trade is neutral to growth. Consequently, we conclude that trade is a limiting factor to output growth in Tunisia and, hence, shocks to trade factors will have a negative effect on output.

References

Balassa, B. (1985) Exports, policy choices, and economic growth in in developing countries after the 1973 oil shock, Journal of Development Economics , 18(1): 23-35.

Bruneau C., De Bandt O. (1999), « La modélisation VAR structurel : application à la politique monétaire en France », Economie et Prévision, n°131, 1, p. 67-94.
developing countries after the 1973 oil shock, Journal of Development

Dickey, D.A. and W.A. Fuller (1979): “Distribution of the estimators for autoregressive time series with a unit root,” Journal of the American Statistical Association, 74, 427-431.
Economics , 18(1): 23-35.

Engle R., Granger C. (1987), “Cointegration and error correction: representation, estimation and testing”, Econometrica, 55, p. 251-276.

Feder, G. (1982) On exports and economic growth, Journal of Development Economics, 12:59-73.

Ghatak, S., Milner.Ghatak, S., Milner, C. and Utkulu, U. (1997) Exports, export composition and growth: Cointegration and Causality evidence for Malaysia, Applied Economics, 29(2): 213-23.

Gregory, A.W., Hansen, B.E. (1996) Residual-based tests for cointegration in models with regime shifts, Journal of Econometrics, 70(1): 99-126.

Hansen, H., Johansen.S. (1993) Recursive estimation in cointegrated VAR model, Institute of Mathematical Statistics .Working paper N°1. University of Copenhagen.

Johansen, S. (1988), « Statistical analysis of cointegration vectors », Journal of Economic Dynamics and Control, 12, p. 231-254.

Johansen, S. (1991), “Estimation and hypothesis testing of cointegration vectors in gaussian vector autoregressive models”, Econometrica, 59, p. 1551-1580.

Johansen S., Mosconi R., Nielsen (2001), “Cointegration analysis in the presence of structural breaks in the deterministic trend”, Econometrics Journal, vol. 3, p. 216-249.

Johansen, S. (1995): “Identification Restrictions of Linear Equations with Applications to Simultaneous Equations and Cointegration,” Journal of Econometrics, 69, 111–132.

Johansen, S., et K. Juseluis (1994): “Identification of the Long Run and Short Run Structure. An Application to the IS-LM Model,” Journal of Econometrics, 63, 7–36.

Kunitomo, N. (1996) Tests of unit roots and cointegration hypotheses in econometric models, Japanese Economic Review, 47(1): 79-109.

Kwiatkowski, D., P.C.B. Phillips, P. Schmidt, and Y. Shin (1992): “Testing the null hypothesis of stationarity against the alternative of a unit root: How sure are we that economic time series have a unit root?” Journal of Econometrics, 54, 159-178.

Leybourne, S.J and Newbold, P. (2003) Spurious rejections by cointegration tests induced by structural breaks, Applied Economics, 35(9): 1117-1121.

Lütkepohl, H., and Wolters, J. (2003) Transmission of German monetary policy in the pre-Euro period, Macroeconomic Dynamics, **7**: 711-733

Pahlavani, M (2005) Cointegration and Structural Change in the Exports-GDP Nexus: TheCase of Iran. International Journal of Applied Econometrics and Quantitative Studies,2-4, forthcoming.

Perron, P. (1989) The great crash, the oil price shock, and the unit root hypothesis, Econometrica, 57(6): 1361–1401.

Perron, P. (1997) Further evidence on breaking trend functions in macroeconomic variables. Journal of Econometrics 80(2): 355-385.

Phillips, P. C. (1991): "To criticize the critics: an objective Bayesian analysis of stochastic trends," Journal of Applied Econometrics.

Phillips, P.C.B. et S. Ouliaris (1990): "Asymptotic properties of residual based tests for cointegration," Econometrica, 58, 165-193.

Quintos,C.E (1995) Sustainability of the Deficit Process with Structural shifts, Journal of business and Economic Statistics,13, 409-417.

Ram, R. (1987) Exports and economic growth in developing countries: evidence from time series and gross-section Data, Journal of Economic Development and Cultural Change, 36: (1): 121-30.

Saikkonen, P. and Lütkepohl, H. (2000a) Testing for the cointegrating rank of a VAR process with an intercept, Econometric Theory 16(3): 373-406.

Saikkonen, P. and Lütkepohl, H. (2000b) Testing for the cointegrating rank of a VAR process with structural shifts, Journal of Business and Economic Statistics 18(4): 451-464.

Sengupta, J. K. (1993) Growth in NICs in Asia: some tests of new growth theory. Journal of Development Studies, 29(2): 342-57.

Shin, Y. (1994): "A residual-based test of the null of cointegration against the alternative of no cointegration," Econometric Theory, 10, 91-115.

Stern, D. (2000) " A Multivariate Cointegration Analysis of the Role of Energy in the US Macroeconomy", Vol. 22, pp 267-283.

Stock, J., et M. Watson (1988): "Testing for Common Trends," Journal of the American Statistical Association, 83, 1097–1107.

Urbain, J. (1992): "On Weak Exogeneity in Error Correction Models," Oxford Bulletin of Economics and Statistics, 54, 187–207.

Van den Berg, H. (1997) The relationship between international trade and economic growth in Mexico, North American Journal of Economics and Finance, 8(1): 1-21.

Zivot, E. and Andrews, D. 1992 ‘Further evidence on the great crash, the oil price shock, and the unit root hypotheses, Journal of Business and Economic Statistics, 10(3), p. 251-270.