Energy Studies Review Vol. 17, No.2, 2010

WORLD OIL PRICE, ECONOMIC GROWTH, INFLATION & INTEREST RATE RELATIONSHIPS IN TUNISIA

HASSEN GUENICHI & SALWA BENAMMOU

ABSTRACT

The aim of this paper is to examine the long run relationship between the world oil price, economic growth demand, and inflation in the developing country of Tunisia, by means of annual data base (1970-2008), univariate and multivariate tests of structural breaks, and cointegration analysis with multiple structural changes. Our empirical results indicate that by positively impacting the price level, oil price negatively impacts real output. The results also indicate that in Tunisia the monetary policy responds to a surge in the oil price in order to reduce or sustain any growth consequences. The ensuing higher inflation however prompts a subsequent tightening of monetary policy leading to a further decline in output. In addition, output does not revert quickly to its initial level after an oil price shock, but declines over an extended period.

Keywords
Economic growth, inflation, interest rate, world oil price, cointegration, structural changes

Hassen Guenichi *and* ***Salwa Benammou*** *University of Sousse, Tunisia*
guenichihassen@yahoo.fr

INTRODUCTION

Oil prices may have an impact on economic activity through various transmission channels. To begin with, there is the classic supply-side effect according to which rising oil prices are indicative of the reduced availability of a basic input to production, leading to a reduction of potential output (see, among others, Barro (1984), Brown and Yücel (1999), Abel and Bernanke (2001)). Consequently there is a rise in production cost, and the growth of output and productivity are slowed. Secondly, an increase of oil prices causes the terms of trade for oil importing countries to deteriorate (see Dohner (1981)). Thus, there is a wealth transfer from oil-importing countries to oil-exporting ones, leading to a fall of the purchasing power of firms and households in oil-importing countries. Thirdly, an increase in oil prices would lead to increase in money demand. Due to the failure of monetary authorities to meet growing money demand with increased supply, there is a rise of interest rates and a slowing down of economic growth (for a detailed discussion on the impact of monetary policy, see Brown and Yücel (2002)). Finally, a rise in oil prices generates inflation.

Since the seminal contribution of Hamilton (1983), several recent studies based on both theoretical and empirical models have made lucid insights into the macroeconomic consequences of oil price shocks. From these studies, centered mainly on the U.S. economy, a general finding is that post-shock recessionary movements of GDP are largely attributable to oil price shocks, although a strand of studies (Ferderer, 1996; Bernanke et al(1997); Hamilton and Herrera, (2004); and Balke et al., (2002)) have provided mixed evidence about the role of post-shock monetary policy. In addition, respective non-linear (Hamilton, 2001) and asymmetric (Davis and Haltiwanger, 2001) specifications of oil price shocks have been found that yield stable oil price–GDP relations over the entire post World War II period.

Davis and Haltiwanger distinguish between aggregate and allocative channels of the effects of oil price shocks. Their analysis suggests that the aggregate channels would increase job destruction and reduce job creation in response to an oil price increase, while the allocative channels would increase both job creation and destruction. Their discussion also emphasizes the view that the aggregate channels should operate symmetrically while the allocative channels would operate asymmetrically because both oil price increases and decreases would alter firms' desired employment structures. Thus, if oil price shocks operate predominantly through aggregate channels, employment would respond roughly symmetrically to positive and negative oil price shocks.

Some studies have argued that the possible impact of energy use on growth will depend on the structure of the economy and the stage of economic growth of the country concerned. Solow (1978), Berndt (1980), Denison (1985) and Cheng (1995) among others suggest that as the economy grows, its production structure is likely to shift towards services, which are not energy intensive activities. In this regard, developing country economies are expected to be more vulnerable to oil price shocks than those of the developed countries. Over time however, as developing countries' technologies improve, conversion processes and end-use devices would progress along their learning curves. As inefficient technologies are retired in favour of more efficient ones, the amount of primary energy needed per unit of economic output would cause the energy intensity to decrease, making their economies less vulnerable to oil price shocks (see, Nakicenovic et al., 1998)

This current study examines the relationship between the world oil price and aggregate demand in the developing country, Tunisia, via the interest rate channel by means of a full systems multivariate cointegration analysis with multiple structural changes. Specifically, we will explore the following issues: what is the relationship between the nominal world oil price, the price level and real domestic output; how does monetary policy influence the said relationship; what are the short-run responses to disequilibrium and long-run behaviour; and how long do the effects of a shock to the world oil price last. In particular, we would like to ascertain whether output tends to revert quickly to its initial level after an oil price shock, or whether the effects of the shock persist (or to lead to a changed level of output for an extended period). Consideration for the role of monetary policy distinguishes the current study from earlier research that has dealt with oil price-output relationships in the context of developing countries, notably North Africa.

News in energy price trends have become topical in recent years and, combined with the ongoing crises in the Middle East, a great deal of uncertainty abounds concerning future oil price movements. Although Africa is endowed with the widest possible range of energy resources that would far exceed its energy requirements, many African countries are reliant on oil imports, Tunisia being one of them. Tunisia has made enormous economic progress in the last two decades. In recent years both political and macroeconomic stability have enabled the country to achieve annual growth rates of over five percent, making it one of the early developing nations to receive debt relief from the international finance institutions. In 2005, oil imports alone accounted for about 20 percent of the country's merchandise imports. The recent hikes in global oil prices, if sustained, could eventually jeopardize the accumulated economic gains. It is the presence of this oil price risk that provides the motivation for the current study.

The remainder of this paper is organized in the following fashion. Section 1 presents the methodology employed in the empirical part. Section 2 discusses our empirical results - unit root test and cointegration analyzed in multiple structural breaks cases. Section 3 concludes.

1. METHODOLOGY

We start our empirical analysis by unit root test and break dates estimations based on the univariate case (Lumsdain and Papell (1997) and multivariate case (Zhongjun Qu and Pierre Perron (2007) approaches, which take into account the existence of potential multiple structural breaks in univariate and multivariate regressions. We then discuss the results of cointegration analysis in the presence of pre-determined structural breaks. First we test for cointegration using Saikkonen and Lütkepohl (2000a) and Johansen and al (2001) procedures, and secondly we estimate the VEC model using Johansen's (1993) approach.

Structural Break dates determination

The Lumsdain and Papell (1997) approach

This approach calculates the minimum value of the LM statistic of the unit root test and determines two potential structural break dates. To correct for serial correlations by including k first differenced lagged (augmented) terms, the program first determines the optimal lag length (k) at each combination of two breaks. The optimal lag length is determined by a general to specific procedure. Starting with the maximum number of lags, max k, the t-statistic on the maximum lagged term is examined to see if significant at the asymptotic 10% level. If not, the maximum lagged term is dropped and the test is repeated until the maximum lagged term is significant or no lags are found (see, for example, Ng and Perron, 1995). Once the optimal lag length at each combination of two breaks is determined, the program searches for the two break points where the unit root t-test statistic is minimized.

In their methodology, Lumsdain and Papell use the two following models:

Model 1 includes two changes in intercept or level of the time series.
Z(t) = [y(t-1),(lags..omit), 1,t,B1(t),B2(t),D1(t),D2(t),DT1,DT2]

Model 2 includes two changes in intercept and trend slope
Z(t) = [y(t-1),(lags..omit), 1,t,D1(t),D2(t),DT1,DT2]

The Zhongjun Qu and Pierre Perron (2007) (ZP) approach

ZP provide a comprehensive treatment of issues related to estimation, inference and computation with multiple structural changes occurring at unknown dates in linear multivariate regression models that include VAR,

ZP consider testing for structural changes. Their setup is quite general in that they shall consider tests that allow for changes in the coefficients of the conditional mean, or in the variance of the error term, or both. Also, they allow only a subset of coefficients to change across regimes, hence partial structural break and block partial structural break models are permitted. They first consider using a likelihood ratio test for the null hypothesis of no change in any of the coefficients, versus an alternative hypothesis with a pre-specified number of changes, say m.

Test of *l* versus *l*+1 breaks

It is often the case that we do not know the number of changes in the system, and a statistical procedure to determine it is needed. For this purpose, information criteria such as those proposed by Liu, Wu and Zidek (1997) and Bai (2000) are possible. But as argued by Perron (1997), these perform rather poorly, especially in models involving lagged dependent variables. Hence, it is useful to have a complementary test-based procedure. Following Bai and Perron (1998), ZP consider a sequential testing procedure based on the estimates of the break dates obtained from a global maximization of the likelihood function.

Consider a model with l breaks, with estimated break dates denoted by $(\hat{T}_1, \dots, \hat{T}_l)$, which are obtained by a global maximization of the likelihood function. The procedure to test the null hypothesis of l breaks versus the alternative hypothesis of $l + 1$ breaks is to perform a one break test for each of the $(l + 1)$

segments defined by the partition $(\hat{T}_1, \dots, \hat{T}_l)$ and to assess whether the maximum of the tests is significant. More precisely, the test is defined by

$$supSEQ_T = (l + 1/1)$$
$$= \max_{1 \le j \le l+1} \sup_{\tau \in \Lambda_{j.\varepsilon}} lr_T(\hat{T}_1, \dots \hat{T}_{j-1}, \tau, \hat{T}_j, \dots, \hat{T}_l) - lr_T(\hat{T}_1, \dots, \hat{T}_l)$$

Where

$$\Lambda_{j.\varepsilon} = \{\tau : \hat{T}_{j-1} + (\hat{T}_j - \hat{T}_{j-1})\varepsilon \le \tau \le \hat{T}_j - (\hat{T}_j - \hat{T}_{j-1})\varepsilon\}$$

Note that this is different from a purely sequential procedure since for each value of l the estimates of the break dates are re-estimated to get those that correspond to the global maximizers of the likelihood function.

Double maximum tests

As in Bai and Perron (1998), Zhongjun Qu and Pierre Perron (2007) also consider a test of the null hypothesis of no break versus the alternative hypothesis of some unknown number of breaks between (1) and some upper bound M. These are called double maximum tests since they are based on the maximum of the (possibly weighted) individual tests for the null of no break versus m breaks (m = 1, ..., M). These are particularly useful to determine whether some structural change is present since a sequential testing procedure can be unreliable for particular forms of multiple changes (Bai and Perron, 2004). More precisely, the test and its limiting distribution are given by

$$Dmax\ LR_T(M) = \max_{1 \le m \le M} a_m \sup LR_T(m, p_b, n_{b0}, \varepsilon)$$

$$\Rightarrow \max_{1 \le m \le M} a_m \sup_{(\lambda_1, \dots, \lambda_m) \in \Lambda_\varepsilon} \sum_{j=1}^{m} LR_j(\lambda, p_b, n_b^*)$$

With $LR_j(\lambda, p_b, n_b^*)$ as defined in Theorem 5 (Zhongjun Qu and Pierre Perron (2005)). They consider an equally weighted version defined by $a_m = 1$, denoted UD max LR_T (M), and a second version that applies weights to the individual tests such that the marginal p-values are equal across values of m, denoted W D max LR_T (M). More precisely, $a_1 = 1$ and for m > 1, $\boldsymbol{a_m = c(\alpha, 1)/c(\alpha, m)}$ where c(α, m) is the asymptotic critical value of the test $\boldsymbol{sup\ LR_T\ (m, p_b, n_{bd}, n_{bo}, \varepsilon)}$ at significance level α.

In practice, ZP suggest to use the following procedure to determine the number of structural breaks. First, use either UD max LR_T (M) or W D max LR_T (M) to test if at least one break is present. If the test rejects, then apply the test SEQT (l + 1 | l) sequentially, for l = 1, 2,..., until the test fails to reject the null hypothesis of no additional structural break.

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”. Therefore, in the present research, 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 (SL) (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 an appropriate adjustment if structural shifts are known to have occurred or are suspected” (2000b: 451). The 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 SL (2000b) and Lütkepohl and Wolters (LW) (2003), an observed n-dimensional time series $y_t = (y_{1t},\ldots,y_{nt})$, y_t is the vector of observed variables (t=1,…, T) which are generated by the following process

$$y_t = \mu_0 + \mu_1 t + \gamma_1 d_{1t} + \gamma_2 d_{2t} + \gamma_3 d_{3t} + \delta Dt_{0t} + \delta_1 Du_{1t} + x_t$$

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 $\gamma(i = 1,2,"), \mu_0,\ \mu_1$, 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:

$$x_t = A_1 x_{t-1} + \cdots + A_p x_{t-p} + \varepsilon_t \quad t = 1,2$$

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). There are three possible options in the SL procedure (as in Johansen): a constant, a linear trend term, or a linear trend orthogonal to the cointegration relations. In this methodology, the critical values depend on the kind of the above-mentioned deterministic trend that is included in the model. More interestingly, in SL the critical values remain valid even if dummy variables are 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.

Causality

Having established the number of cointegrating vectors, we performed Granger-causality tests (Granger, 1969) in order to verify the informational relationships between the four variables. Granger-causality from x_k to x_j means that the conditional forecast for x_j can be significantly improved by adding lagged x_k to the information set. The feasibility of the Granger-causality tests depends on the stationary features of the system. If the series are stationary, the null hypothesis of no Granger causality can be tested by standard Wald tests (Lütkepohl, 1991).

In a cointegration model as in (1) however, two sources of causation come to light, either through the error-correction term (ECT),$\Pi Y_{t-k-1} = \alpha\beta' Y_{t-k-1}$, if, α≠0, or through the lagged dynamic terms,$\sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i}$ if all Γ_i= 0 (see Toda and Phillips, 1993). The ECT measures the long-run equilibrium relationship while the coefficients on lagged difference terms indicate the short-run dynamics. Thus in a cointegration model (1) the proposition of x_k not Granger-causing x_j in the long run is equivalent to$\alpha_{jk} = 0$. In this context x_j is said to be weakly exogenous for the parameter β, x_j does not react to the equilibrium errors. Also from model (1) the proposition of x_k not Granger-causing x_j in the short run is equivalent to $\Gamma_{jk}(L) = 0$, where (L) is the lag operator.

Impulse response

We determine how each endogenous variable responds over time to a shock in oil price variable, i.e., the effects of a shock or change in the error term associated with the oil price equation in model (1). To calculate these impulse responses, we increase for one period only the error term in the equation for the world oil price by one standard deviation and then calculate the immediate and then future effects of this change on output, world oil price, interest rate and the price level. For this exercise, estimates of the covariances among the four error terms would be required.

2. EMPIRICAL RESULTS

The Model

We estimate a four-dimensional macro-econometric model that represents a vector of real output (GDP), world oil price in nominal terms (OP), nominal interest rate (IR) and the price level as indicated by the consumer price level(CPL) in the form of a vector error correction (VEC) representation (see Johansen (1988; 1991)

$$\Delta Y_t = \Pi Y_{t-k-1} + \sum_{i=1}^{k-1}\Gamma_i \Delta Y_{t-i} + \Psi D_t + \varepsilon_t \tag{1}$$

where the reduced rank, r, of the 44× matrix of Π equals the number of cointegration vectors in the system, and n equals four, the number of (endogenous) series in cointegration equation (1). Thus, Π can be written as$\Pi = \alpha\acute{\beta}$, where α and β are each of the dimension r×4 and rank r. The matrix β contains the cointegrating vectors β, $\beta = (\beta_t, \dots, \beta_\tau)$ while the matrix of the adjustment coefficients α describes the speed of adjustment of each of the four individual series in Y_t to deviations from the cointegration relationships.

Data description

We used annual time series data covering the period 1970–2008. Real GDP at constant 2000 prices in cedis as well as the consumer price index data were obtained from the World Bank's World Development Indicators data-base. As a measure of monetary policy stance, we used the discount rate obtained from the IMF International Financial Statistics data-base. We also used oil price and interest rate data bases in our study.

All the variables were used in their logarithmic form. The logarithmic data were deemed to permit more parsimonious dynamics than non-logarithmic data (see Jumah and Kunst, 1996). Figure 1 illustrates changes in the respective logarithmic data series. The figure shows that all the variables are upward trending with the price level showing the most upward trend.

Figure 1: Trend of real output, oil price, interest rate and price level in logs

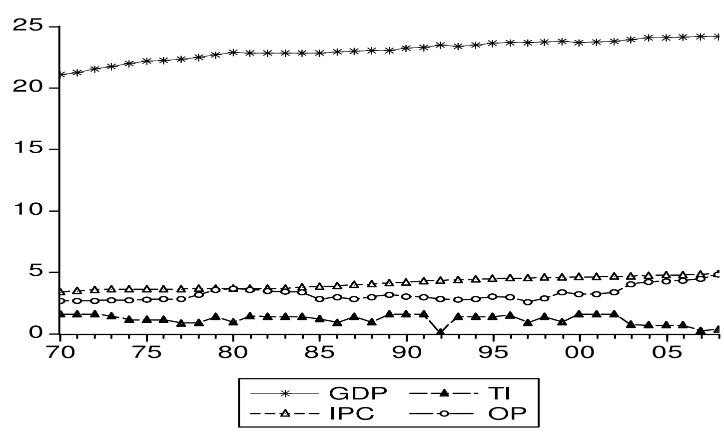


Table 1: Lumsdain and Papell (1997) results.

variables	Time break	Causes of TB	model	Min LM	B_1	B_2	D_1	D_2	DT_1	DT_2
GDP	Obs(17) 1986 Obs(22) 1991	Debt crises in developing countries Gulf war effect	1	-6.09	-2.10	1.14	-1.64	1.34	0.007	0.0306
			t-sat		-2.09	3.25	-1.63	3.32	0.84	2.21
			2	-5.75	….	……	1.4626	1.39	0.0080	0.030
			t-stat		…..	……	3.85	3.507	0.93	2.25
CPL	Obs(4) 1973 Obs(22) 1991	Embargo of the OPEP countries Gulf war effect	1	-6.50	-1.36	2.33	-1.96	0.049	0.073	-0.051
			t-stat		-1.40	4.86	-3.075	0.148	3.9162	-3.927
			2	-5.57	……	……	2.265	1.423	0.0688	-0.044
			t-stat		……	……	3.682	0.465	3.5268	-3.389
IR	Obs(16) Q2 1985 Obs(30) Q1 1999	Debt crisesin developing countries Second war in the Gulf	1	-5.77	1.24	-1.04	-0.274	1.800	-0.015	-0.010
			t-stat		1.22	-3.02	-.2551	3.353	-1.923	-0.410
			2	-6.81	……	……	-1.066	1.752	-0.015	-0.008
			t-stat		……	……	-3.136	3.510	-1.956	-0.353
OP	Obs(18) 1988) Obs(32) 2001	Debt crises in developing countries War in Afghanistan (World Trade Center attack)	1	-4.93	0.13	-0.08	3.1942	-0.989	-0.171	0.0744
			t-stat		3.52	-4.11	3.3522	-3.066	-2.049	5.0333
			2	-5.64	……	……	0.8665	-1.503	-0.029	0.0512
			t-stat		……	……	2.5017	-3.235	-3.185	4.1376

Structural break dates determination

Based on the Lumsdain and Papell (1997) method our empirical results are in Table (1).

In this univariate case, the results (Table1) show two different structural breaks in such variables. Here we find only one significant break date (1980-1988) which explains the debt crises in most developing countries. But we cannot easily find the second significant break date. The Lumsdain and Papell (1997) test results give many dates such as the Gulf war effect, the second war in the Gulf, war in Afghanistan and the World Trade Center attack. Then we must take the multivariate case as a possible solution for this problem which determines the possible break dates in a VAR model.

We apply the Zhongjun Qu and Pierre Perron (2007) approach to determine the possible multiple structural breaks in the multivariate cases (VAR model).

Table2: LR test results

Breaks	SupLR	Critical value (10%)	Critical value (5%)
0 versus 1	76.203	13.711	15.662
0 versus 2	68.123	22.588	25.090

Table 3: WDmax test results

	Statistic value	Critical value (10%)	Critical value (5%)	Critical value (2,5%)
WDmax (2breaks)	73.223	14.845	16.879	18.772
Seq(L+1/L) : (2/1)	120.01	15.858	17.823	19.409

Table 4: Break dates estimation

Break date	Confidence intervals (CI)
The date of the first break is 17th observation (1986)	CI= [15.000 18.000] at 95% CI= [16.000 18.000] at 90%
The date of the second break is 32th observation (2001)	CI= [31.000 34.000] at 95% CI=[31.000 33.000] at 90%

Based on the results reported in Tables 2, 3 and 4, the primary findings of the analysis are as follows. The results of the Zhongjun Qu and Pierre Perron (2007) models indicate that the timing of any structural break (TB) for the multivariate model using the ZP approach is also shown in Table 4. The computed break dates correspond closely with the expected dates associated with the debt crises in most developing countries (1986), the effects of the attack of World Trade Center and the beginning of the war in Afghanistan in 2001.

Cointegration analysis results

We now apply a maximum likelihood approach for testing and determining the long-run relationship under investigation in the model. As mentioned earlier, in this procedure, Johansen assumed that the break point is a known priori. In the last section, we determined the time of the break endogenously by the Zhongjun Qu and Pierre Perron (2007) procedure. The empirical result based on this method showed two significant structural breaks in the model under investigation, which are consistent with the time of the debt crises in most of the developed countries, and the war in Afghanistan. Therefore, at this stage we include two dummy variables of regime change in order to take into account the two 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 table 5.

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 5. These tables indicate that the hypothesis of no cointegration r=0 is rejected at the10%, 5% and 1% significance level. The existence of one cointegration vector is not rejected in any of the three cases mentioned.

Table 5: cointegration test results

Intercept included (C)						Intercept and trend included (C/T)						Trend orthogonal to cointegration relation (C/O)					
r	LR	pval	90%	95%	99%	r	LR	pval	90%	95%	99%	r	LR	pval	90%	95%	99%
0	89.33	0.0000	69.17	72.80	79.96	0	111.14	0.0000	79.99	84.35	92.95	0	70.56	0.0011	44.45	47.71	54.23
1	46.55	0.0789	46.90	50.00	56.15	1	42.35	0.1014	55.25	58.95	66.32	1	20.18	0.1702	27.16	29.80	35.21
2	22.99	0.2244	28.49	31.03	36.19	2	19.55	0.3338	34.28	37.28	43.36	2	10.05	0.6928	13.42	15.41	19.62
3	11.10	0.2289	13.91	15.97	20.32	3	15.13	0.2501	16.65	18.80	23.28						

Table 6 presents evidence for the long-run behaviour of the variables. Oil price is positively related to the price level but negatively related to interest rate and output. The results may be interpreted as implying that Tunisian monetary policy is eased in response to a surge in the price of oil in order to lessen any growth consequences, but at the cost of higher inflation. An alternative interpretation might be that monetary policy is tightened in response to a rise in inflation emanating from a surge in the price of oil, resulting in a crowding out of the private sector. Whichever explanation predominates, it will be certified by the results of our impulse response analysis.

Also from Table 6, the magnitudes of the short-run coefficients, i.e., estimates of the ∝ coefficients attached to the error correction terms, confirm that the speed of adjustment to the long-run change in real output is slow, while those of the other three variables are moderate. In addition, the interest rate coefficient is insignificant, implying that this variable is weakly exogenous to the cointegration vector, which explains that in Tunisia the interest rate is maintained and fixed by the government.

Table 6: Long-run cointegration vector estimates

	GDP	CPL	IR	OP	constant
β p-value t-stastic	1.000	-5.771 {0.000} [-8.323]	0.426 {0.007} [2.713]	0.995 {0.000} [6.905]	-43.683 {0.000} [-17.159]
∝ p-value t-stastic	0.088 {0.012} [2.820]	0.235 {0.000} [4.964]	0.287 {0.644} [0.462]	-0.331 {0.000} [-3.523]	

Formally, the test of H: $\alpha_i = 0$for i = 3, signifying that the interest rate is weakly exogenous involves: $\dot{\alpha}$= [*,*,0, *] for interest rate, where * denotes an unrestricted coefficient. Results of the corresponding likelihood ratio (LR) tests based on 2 degrees of freedom are presented in Table 7. As can be seen from the table, the null hypothesis of the presence of weak exogeneity is not rejected at the 5 percent significant level.

Table 7: Restricted long-run cointegration vector estimates

	GDP	CPL	IR	OP	constant
β p-value t-stastic	1.000	5.671 {0.000} [8.323]	0.436 {0.007} [2.713]	0.993 {0.000} [6.900]	-13.683 {0.000} [-11.159]
∝ p-value t-stastic	0.028 {0.874} [0.820]	-0.036 {0.000} [-4.967]	0 [0]	-0.331 {0.000} [-3.523]	

The results of the short-run causality tests as shown in Table 8 clearly indicate that the price level Granger causes real output. Also, oil price is seen to Granger Cause the price level. The insight here is that energy costs influence the firm's price setting and even the relation between output and employment. Given wages, an increase in the price of oil increases the cost of production forcing firms to increase prices, leading to an increase in the price level. To the extent that increases in the price of oil lead to a rise in the price level, they also reduce consumer-spending power through a reduction in the real money stock. The result is a fall in output.

Table 8: VEC Granger Causality

Null Hypothesis:	F-Statistic	Probability
IR does not Granger Cause GDP	1.99196	0.15299
GDP does not Granger Cause IR	1.22177	0.30809
CPL does not Granger Cause GDP	5.60584	0.00818
GDP does not Granger Cause IPC	3.03016	0.06235
OP does not Granger Cause GDP	1.24335	0.30198
GDP does not Granger Cause OP	0.60462	0.55240
CPL does not Granger Cause IR	0.89870	0.41713
IR does not Granger Cause IPC	0.05875	0.94304
OP does not Granger Cause IR	2.14272	0.13387
IR does not Granger Cause OP	1.08408	0.35031
OP does not Granger Cause IPC	0.17026	0.84421
CPL does not Granger Cause OP	2.86790	0.07151

The conclusion to be drawn from the weak exogeneity and Granger causality tests is that the interest rate and the world oil price are strongly exogeneous. This inference has important implications for econometric modelling of oil markets. For instance, Cavallo and Wu (2006) have observed that conventional measures of oil-price shocks based on oil-price changes suffer from the two obvious flaws of endogeneity and forecastability.

Table 9: Residual covariance matrix

	GDP	CPL	IR	OP
GDP	0.0059652	-0.0002997	-0.0146331	-0.007587
CPL	-0.00029	0.000326	0.000974	0.00109
IR	-0.014633	0.000974	0.13160	-0.030489
OP	-0.00758	0.001092	-0.030489	0.05422

In order to examine the impulse responses or the short-run adjustments to a shock in the world oil price (i.e., a standard error increase in the structural residual of the equation determining the world oil price), we need to know the covariances among the four error terms. The estimated covariances are shown in Table 9. Note that in general, the correlations among the residuals are low. The residual in the oil price equation is most correlated with that of the interest rate equation. Thus, a shock to oil price will have a more common component with the interest rate.

Figures 2- 4 show the response of each variable to a one-standard-deviation in the oil price.

In the next ten periods, output declines due to the negative covariance and then more gently thereafter. Regarding the interest rate, there is a rise in the first and second period, followed by a decline in the third period due to the negative covariance between oil price and the interest rate. The price level also rises slightly in the second period because of the positive covariance between the two variables, followed by a decline in the third and fourth periods and a rapid rise in all next periods. The results imply that the effects of an oil price shock persist, leading to a decline in output over an extended period.

Figure 2: Response of output to 1-standard deviation shock in oil price

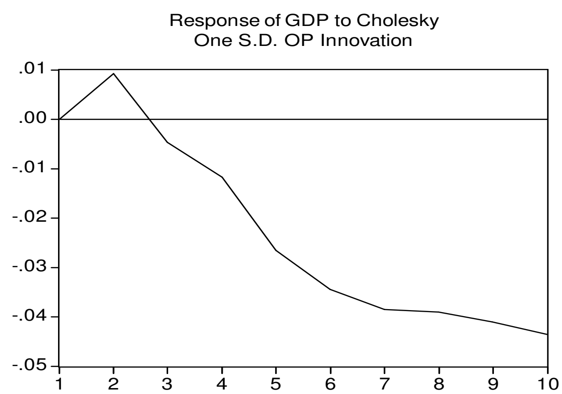


Figure 3:
Response of nominal interest rate to 1-standard deviation shock in oil price

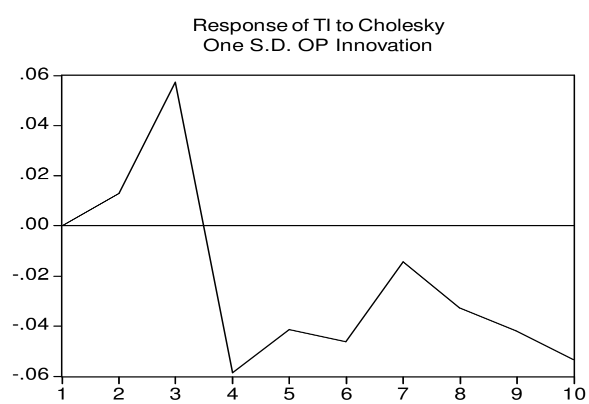
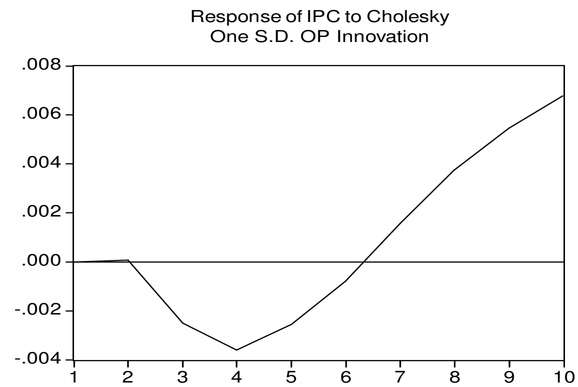


Figure 4:
Response of the price level to 1-standard deviation shock in oil price



CONCLUSION

It is important to understand the relationship between the world oil price and real GDP because oil shocks immediately raise prices of petroleum products that are key production inputs as well as essential consumer goods. Additionally, oil shocks are likely to push up prices in other energy markets as has been witnessed in recent times. These price increases are considerable enough that they characteristically show up as temporary spikes in the overall rate of inflation and may even get passed through to continuing rates of inflation. While increases in the price of oil lead to a rise in the price level, they also reduce consumer-spending power through a reduction in the real money stock. According to LeBlanc and Chinn (2004), the implications of higher energy prices on inflation depend in part on how important energy is in the economy. The importance of oil in production and consumption across countries may however be offset by differences in inflationary transmission mechanisms with respect to wage setting institutions. But that is not the essential point here.

This paper attempts to analyze the long run relationship between oil price and real GDP in a developing country, Tunisia, via the interest rate channel by means of cointegration analysis output growth in Tunisia. We examine the long-run determinants of GDP during the period 1970-2008, employing the Saikkonen and Lutkephol (2000) and Johansen and (2001) cointegration method. Prior to the cointegration analysis, the Lumsdain and Papell (1997) and Zhongjun Qu and Pierre Perron (2007) tests are applied in order to endogenously determine the multiple structural breaks in the major drivers of economic growth, interest rate, oil price and level price. The empirical results based on the unit root tests indicate the existence of the unit root for all of the variables under investigation. Moreover, we find that the structural breaks over the last forty years occurred as a result of the debt crises in developing countries in 1980"s and the World Trade Center attack in 2001. These results provide complementary evidence to models employing exogenously imposed structural breaks in the Tunisian macro-economy.

The results of the study indicate that oil price impacts the price level positively and negatively impacts real output. The results also indicate output does not revert quickly to its initial level after an oil price shock, but declines over an extended period.

The study suggests that the adverse effect of higher oil prices on growth can be mitigated through prudent fiscal policies and structural reforms, such as the dismantling of oil price subsidies, as well as credible monetary policy. Lower government-controlled petroleum prices may not promote an efficient use of resources and may lower the incentives for households and firms to switch to other energy sources. By reducing the demand for oil products, higher petroleum prices could improve the foreign trade balance and reduce the harmful environmental effects that stem from oil consumption. Nevertheless, the gains in efficiency may be reduced when households switch expenditures to other goods that are heavily subsidized or that may have a greater environmental effect.

The Tunisian government tries to reduce the oil consumption by introducing other kinds of energy (solar energy, electrical energy) into consumer's traditions. Moreover, it maintains the local oil prices to protect the consumers from sudden surge of the world oil prices. Also the interest rate is maintained by a prudent monetary policy. Until now, all of these efforts are still not enough to get a durable economic growth.

REFERENCES

Akçay, O.C., C.E. Alper, and S. Özmucur (2002): "Budget Deficit, Inflation and Debt Sustainability: Evidence from Turkey (1970-2000)." In: Kibritcioglu, A., Rittenberg, L., Selcuk, F. (eds.), Inflation and Disinflation in Turkey, Aldershot, U.K., Ashgate, pp. 77-96.

Atkeson, A. and L.H. Ohanian (2001): "Are Phillips Curves Useful for Forecasting Inflation?" *Federal Reserve Bank of Minneapolis Quarterly Review*, 25:2-11.

Balke, N.S., S.P.A. Brown, and M.K. Yucel (2002): "Oil Price Shocks and the U.S. Economy: Where Does the Asymmetry Originate?" *Energy Journal*, 23:27-52.

Bernanke, Ben S., M. Gertler, and M. Watson (1997): "Systemic monetary policy and the effects of oil price shocks." *Brookings Papers on Economic Activity* 1:91-157.

Brendt, E. (1980): "Energy Price Increases and Productivity Slowdown in US Manufacturing." In: Federal Reserve Bank of Boston, Decline in Productivity Growth, Federal Reserve Bank of Boston, Conference Series, USA, Boston.

Cavallo, M., and T. Wu (2006): "Measuring Oil-Price Shocks Using Market-Based Information." Working Paper 2006-28, Federal Reserve Bank of San Francisco.

Cheng, B. (1995): "An Investigation of Cointegration and Cauasality Between Energy Consumption and Economic Growth." *Journal of Energy Development*, 21:73-84.

Coady, D., M. El-Said, R. Gillingham, K. Kpodar, P. Medas, D. Newhouse (2006): "The Magnitude and Distribution of Fuel Subsidies: Evidence from Bolivia, Ghana, Jordan, Mali, and Sri Lanka." IMF Working Paper WP/06/247, International Monetary Fund, Washington, D.C.

Davis S.J. and J. Haltiwanger (2001): "Sectoral Job Creation and Destruction Response to Oil Price Changes." *Journal of Monetary Economics*, 48:465-512.

Denison, E. (1985): "Trends in American Economic Growth, 1929-1982." Brookings Institution, Washington, D.C.

Ferderer, J. P. (1996): "Oil Price Volatility and the Macroeconomy: A Solution to the Asymmetry Puzzle." *Journal of Macroeconomics*, 18:1-26.

Gramlich, E.M. (2004): "Oil Shocks and Monetary Policy, Annual Economic Luncheon." Federal Reserve Bank of Kansas City.

Granger, C.W.J. (1969): "Investigating Causal Relations by Econometric Models and Cross-spectral Methods." *Econometrica,* 37:424-438.

Hamilton, J. (2005): "Oil and the Macroeconomy." Manuscript Prepared for: *Palgrave Dictionary of Economics* (forthcoming).

Hamilton, J.D. (1983): "Oil and the Macroeconomy since World War II." *Journal of Political Economy*, 91:228-248.

Hamilton, J.D. (2001): "A Parametric Approach to Flexible Nonlinear Inference." *Econometrica*, 69:537-573.

Hamilton, J.D. and A.M. Herrera (2004): "Oil Shocks and Aggregate Macroeconomic Behaviour: The Role of Monetary Policy. *Journal of Money, Credit, and Banking*, 36:265-286.

Humpage, O. and E. Pelz (2003): "Do Energy Price Spikes Cause Inflation? Economic Commentary." Federal Reserve Bank of Cleveland.

Johansen, S. (1988): "Statistical analysis of co-integrating vectors." *Journal of Economic Dynamics and Control,* 12:231-254.

Johansen, S. (1991): "Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models." *Econometrica*, 59:1551-1580.

Jumah, A. and R.M. Kunst (1996): "Forecasting Seasonally Cointegrated Systems: Supply Response of the Austrian Breeding Sow Herd." *European Review of Agricultural Economics,* 23:487-507.

Kpodor, K. (2006): "Distributional Effects of Oil Price Changes on Household Expenditures: Evidence from Mali." IMF Working Paper WP/06/91, International Monetary Fund, Washington, D.C.

LeBlanc M., and M.D. Chinn (2004): "Do High Oil Prices Presage Inflation? – The Evidence from G-5 Countries." UC Santa Cruz Economics Working Paper No. 561; SCCIE Working Paper No. 04-04.

Lütkepohl, H. (1991): "Introduction to Multiple Time Series Analysis." Springer-Verlag, Berlin.

Miller, P.J. (1983): "Higher Deficit Policies Lead to Higher Inflation." Federal Reserve Bank of Minneapolis *Quarterly Review,* 7(1):8-19.

Mishkin, F.S. (2007): "Inflation Dynamics, Annual Macro Conference." Federal Reserve Bank of San Francisco.

Nakicenovic, N., M. Jefferson, and A. McDonald (1998): *Global Energy Perspectives.* Cambridge University Press, Cambridge.

Qu, Z. and P. Perron (2007): "Estimating and Testing for Structural Changes in Multivariate Regressions", *Econometrica.*

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

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.