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COINTEGRATION BETWEEN ECONOMIC ACTIVITY AND OIL PRICES IN THE OPEC COUNTRIES: A TIME SERIES APPROACH

Nicola Rubino
Ph.D. Candidate, University of Barcelona, Faculty of Economics, Barcelona, Spain
Email: nicola.rubino@ub.edu.

Abstract

The aim of this paper is to study the long-term relationship between oil prices and economic activity, proxied by GDP. To account for the long accepted evidence of a non-existing long run relationship between oil prices and economic activity, we carry out unit root and cointegration tests in presence of deterministic structural breaks. Our empirical analysis concerns the OPEC group, and generally extends from 1960 to 2012. This study contributes to the extensive literature on oil prices by adding a proper supply side analysis of a possible long run equilibrium between GDP and oil prices in a group of oil exporting and producing countries, and effectively manages to find an equilibrium relationship in Saudi Arabia by taking into account possible structural breaks. Setting up a Granger causality test in presence of deterministic structural breaks, the papers concludes that, even though no short run causality linkage could be found between oil prices and GDP, the existence of such relationship holds in the long run and appears to show some degree of predictive ability on GDP growth in Saudi Arabia.

Keywords: OPEC Countries, GDP growth, Oil prices, Cointegration Analysis.

JEL Classification: C32, F43, O47.

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1. Introduction and Literature Review

The aim of this paper is to study the long-term relationship between oil prices and economic activity. Unlike most of the existing literature, which focuses on Western countries, we analyze the relationship between oil prices and GDP in a group of oil exporting countries, belonging to the Organization of the Petroleum Exporting Countries (OPEC). To describe this relationship, past literature has considered a time series framework. These studies have generally found evidence of an inverse, and generally inconsistent relationship between the two variables. In his seminal paper, Hamilton (1983) showed through an unrestricted VAR model that oil prices and U.S. GDP were negatively correlated, as every recession in the post-war era was normally preceded by a sudden spike in the oil prices. However, by the mid-1980s, the estimated linear relationship between oil prices and GDP began to loose significance: the declines in oil prices occurred over the second half of the 1980s were found to have smaller positive effects on economic activity than what was predicted by linear models. At the same time, evidence of asymmetries in the link between the two variables had been found and tested at various stages across literature. Mork (1989) re-examines results from Hamilton (1983) and finds out that rising prices appear to be more highly correlated to gross national product than price decreases in the U.S. A few years later, Mork, Olsen, and Mysen (1994) extended this analysis and found evidence of an asymmetric relationship for a group of seven industrialized countries. Exploring the linkage between monetary policy and oil prices, Lee, Ni, and Ratti (1996) switched the analysis to the role of volatility and found out that oil prices are more likely to have an impact on growth in economic contexts with more stable and predictable fluctuations. Ferderer (1996) focused on industrial production growth rather than GDP, and through a structural VAR approach found strong evidence of asymmetric behavior, as price increases explained more than twice the volatility in production growth than price decreases. Hamilton (1996) carried on from this last piece of information, and replying to Hooker (1996) introduced a new notion of net price increase arguing that it is the joint presence of volatility and asymmetry that allows oil prices to affect the economy. Recently, asymmetric cointegration modelling has been employed to confirm the evidence that economic activity responds asymmetrically to oil price shocks in order to uncover the structural relationship between oil prices and GDP[1]. Indeed, at least in countries belonging to the western hemisphere, rising oil prices appear to retard aggregate economic activity by more than falling oil prices stimulate it. However, to our knowledge, few studies have focused on the supply side of the world oil market and on the long-run relationship between GDP and oil prices in oil producing/dependent countries.

[1] To cite the most recent examples, see Lardic and Mignon (2006) and Lardic and Mignon (2008).

The rest of the paper is organized as follows: Section 2 introduces the topic and gives a literature review, together with an explanation of the theory related to the linkages between oil prices and economic activity; Section 3 introduces the data and gives an overview of the methodology; unit root test on the series are carried out in Section 4 while Section 5 contains the cointegration analysis; Section 6 reports the estimates of the weak causality tests we ran, and Section 7 finally concludes.

2. Transmission Channels

As we have already discussed in the introduction, the post-World War II relationship between oil prices and economic activity appears to have changed sometime in the 1980s. Rotemberg and Woodford (1996) state in a very clear way that sometime after 1980 OPEC lost its ability to keep the nominal price of oil relatively stable. After such date, variations in the demand for oil were reflected quickly in nominal price changes, and several statistical properties and overall behaviour of oil prices changed as a result. The meaning of the change in the oil-price GDP could thus be interpreted in two different ways. Firstly, oil might have once been able to affect GDP, or perhaps the relationship might have been absent at all, the only reason for its pre-1980 existence being that the earlier sample period of the data was not lengthy enough to expose the null character of the hypothesis. A second, possible interpretation is that the relationship was never intended to be linear, but that pricing conditions in the world oil market from just after World War II through the late 1970s let linear versions of the relationship approximate the observed behavior, while hiding the real structural form of the relationship. However, instead of resorting to an asymmetric approximation to model the relationship, our paper favors a structural break set-up in the deterministic components of the series. Our analysis thus estimates the relationship focusing on the OPEC countries, and accommodates the above first critique making use of all yearly available information for the group, and the second critique not by specifying a non-linear relation-ship between GDP and nominal oil prices through new variable definitions or asymmetric approaches, but by accounting for the impact of non-linearities on a country by country approach based on the structural break analysis of the data generating processes of GDP and nominal oil prices and on a possible change in the deterministic components in the long-run relationship between the two variables.

Oil prices are likely to affect macroeconomic activity in both the short and the medium-run. This paragraph fits in the main analysis as it indicates potential linkages between oil prices and economic activity.[2] Theory would distinguish

[2]For a valid overview of the channels of impact of energy prices shocks on aggregate economic activity, with a specific focus on importing countries, see Brown and Yucel (2002).

three main transmission channels: a traditional input-costs channel, the effects of income shifts across oil importing and oil exporting countries and finally asymmetric effects of monetary policy actions. The first channel can be considered as the industrial spill-over channel: as oil prices drop (increase) prices of competing energy commodities are driven down (up) too, and oil-intensive production sectors' prices are positively (negatively) hit as the decline in oil prices influences a range of different inputs. The income shifts channel on the other side operates as a redistribution of income from countries with a generally higher propensity to save, the exporting countries, to countries with a generally higher propensity to spend, the importing countries. This, at least in the short-medium term, assuming positive expectations of the importing countries and financial and fiscal constraints of the exporting countries remain unchanged, would result, in the case of an oil price drop, in an overall stronger demand for oil. Finally, with a specific focus on importing countries, the drop (increase) in oil prices might trigger a loosening (strengthening) of the monetary policy, fostering (depressing) economic activity, while the effect might reverse in exporting countries, where the inability of the oil industry to reach the fiscal break-even point after a decrease in oil prices might promote contractionary fiscal actions, slowing down overall economic activity.[3] In a time framework like ours, identify the impact of each channel in the short/medium run is not an easy task. Variations of the oil prices might feed into the economy through any of the channels or have a contemporary effect in two or all of them. Substantially, the presence of this three channels seems to indicate that a positive variation of the oil prices would normally involve a positive variation of economic activity. On the other side, in the long run, we can think about oil prices as an instrument to detect the presence of a commodity curse effect on long run GDP levels, and would thus expect the long run elasticity of oil prices with respect to GDP to enter the relationship with a negative sign.

3. Data and Methodology

Countries belonging to our database were selected based on data availability and their status as OPEC members. An issue we would like to address, related to the time dimension of the analysis, has to do with the sample period of the series. Many, if not the majority of the literature we previously reviewed, has up until now consistently chosen year 1980 as the initial sampling period. Historical reasons accounted for are the beginning of the globalization era, a higher degree of insulation of the world economies from commodity price shocks, as well as availability of high quality data. Given current data availability, as the GDP and the oil prices series are both available starting from 1960 for the majority of the countries belonging to the group we examined, we choose to focus on this two

[3]For a more informative overall explanation, as well as an overview of the potential short run effects of the recent oil price plunge, see Baffes, Kose, Ohnsorge, Stocker, Chen, Cosic, Gong, Huidrom, Vashakmadze, Zhang, and T. (2015).

variables to check for the existence of a long-run relationship taking into account the presence of structural breaks in their data generating processes. From an empirical viewpoint, we consider a linear cointegration framework to analyze the link between nominal oil prices and real GDP.[4] The real GDP series for our set of countries are taken from the World Bank Database (WDI) while the oil price series are taken from the British Petroleum company (BP) statistical database (where the nominal oil price series are defined as "money of the day"). The WDI contains GDP yearly series from 1960 and covered almost entirely the OPEC group. Some gaps at the beginning of the GDP series were filled in with the help of the IMF financial statistics database (IFS). The bivariate relationship between real GDP and oil prices could be implicitly represented as in equation (1):

$$GDP_t = f(oil_t) \qquad (1)$$

where GDP_t is the natural log of real GDP and oil_t is the natural log of the oil prices series. Since cointegration implies a possible error correction representation by the Granger representation theorem, we will estimate the above relationship in an error correction model specification, in order to check for long-run weak causality direction as well as short run and joint causality. Before proceeding with the weak causality tests, an analysis of the order of integration and cointegration is carried out in the next Section.

[4]The reason why we should choose nominal oil prices over real one is well explained by Hamilton (2005), according to whom endogeneity issues should be far more prevalent when employing real oil prices series as these are more likely to carry the effect of strong exogenous shocks as the Suez crisis. However, we carried on this exercise employing both a nominal and a real definition of oil prices.

Table 1. OPEC Countries, Descriptive Statistics of GDP

Country	T	Mean	Median	Variance
AGO	28	23.77	32.52	0.26
ARE	38	25.38	25.29	0.22
DZA	53	24.67	24.88	0.65
ECU	53	23.79	23.9	0.65
GAB	53	22.4	22.58	0.33
IDN	53	25.38	25.44	0.81
IRN	53	25.74	25.46	1.22
IRQ	53	23.58	23.77	0.91
KWT	53	24.68	24.63	0.15
LBY	41	24.47	24.5	0.29
NGA	49	24.76	24.74	0.26
QAT	53	24.04	23.9	0.47
SAU	50	25.74	26.01	1.01
VEN	50	25.32	25.32	0.13

Abbreviations: Ago=Angola, ARE=United Arab Emirates, DZA=Algeria, ECU=Ecuador, GAB=Gabon, IDN=Indonesia, IRN=Iran, IRQ=Iraq, KWT=Kuwait, LBY=Lybia, NGA=Nigeria, QAT=Qatar, SAU=Saudi Arabia, VEN=Venezuela

4. Unit Root Tests

4.1. Methodology

Integration analysis was conducted with the corrected Ng and Perron (2001) M-tests, were an intercept and a trend were included in every ADF regression for the time series in levels, and a constant only for the differenced series. In order to tackle the issue of trend-break stationarity, with a particular emphasis on the GDP series,we also employed the unit root tests proposed in Zivot and Andrews (1992) and Lumsdaine and Papell (1997), which account for the possibility of either one or two endogenously determined structural breaks in the series.[5] Additionally, the Perron and Vogelsang (1992) and Clemente, Montanes, and Reyes (1998) Innovation outlier (IO) and Additive outlier (AO) tests, which have been here applied to check for stationarity of the first differenced series, were carried out. Given the length of the time series at our disposal, only up to two breaks were considered in both deterministic components of the assumed data generating processes, aimed at avoiding the critique of data mining and to ensure the empirical relevance of the breaks. In order to ascertain the effective order of

[5]The necessity of looking for an endogenous methodology to test for the null hypothesis of no structural change can be spotted in Christiano (1992), where the author, based on the evidence contained in a seminal paper by Perron (1989), states that a choice of breaks independent from prior information on the data, an exogenous break-date choice, would lead to normally reject a no-break null hypothesis.

integration of the series, we followed the approach described by Dickey and Pantula (1987), starting the unit root analysis from the differenced series and then reducing the order of differentiation checking for non-rejection of the null hypothesis of unit root.

4.2 Results for GDP

Results of the M-tests in Table 2 show that only in six out of the fourteen OPEC countries[6] the GDP might actually follow an I(1) process. According to the Ng and Perron (2001) tests results, some GDPs would appear to follow an I(2) process. Since the GDP of oil-exporting countries might have been influenced by some relevant historical events,[7] we report the one and two breaks test results for the variable in first differences in Tables 6 and 8 for the innovative outlier methodology, and Tables 5 and 7 for the additive outlier methodology. The Zivot and Andrews (1992) and Lumsdaine and Papell (1997) tests, carried on the variable levels and whose output we report in Tables 3 and 4, point alternatively at two I(1) groups: the one break test would consider the GDP in Qatar, Nigeria, Libya, Iraq and Ecuador to be non-stationary, pointing at the oil crisis of the seventies as the relevant break date; the two break test would instead show a non-stationary GDP for Arab Emirates, Libya, Iraq, Qatar, Saudi Arabia and Venezuela. The same two tests, carried out following the IO methodology, would instead present a less conservative evidence: the one break test would indicate the GDP in Angola, Arab Emirates, Ecuador, Gabon, Iraq, Kuwait, Libya, Nigeria and Venezuela as I(1), while the two break IO test would point at Arab Emirates, Iraq, Kuwait, Libya, Nigeria and Venezuela as I(1) series. Based on our results, we choose to restrict our analysis to an I(1) group including Angola, Ecuador, Indonesia, Iraq, Nigeria and Venezuela, which the Ng and Perron (2001) tests indicated as having a non-stationary in level GDP,[8] and ultimately adding Qatar, Libya, Saudi Arabia and Arab Emirates, whose GDP series were indicated as non-stationary in at least one of the four unit root tests with structural breaks we employed.

[6] Angola, Ecuador, Indonesia, Iraq, Nigeria and Venezuela.

[7] Such as the first oil crisis in the late seventies, the switch from administered to free market oil prices and the oil price glut in the second half of the eighties, second order effects of the gulf war in the nineties and finally recent developments connected to the second globalization era and the increase in size of the service sector in the years 2000.

[8] It has been acknowledged how M-type tests are normally less inclined to suffer of small sample bias than other ADF based tests. However, for the purpose of this investigation, the outcome of unit root tests accounting for at least a single structural break in the form of a sudden or more gradual change cannot be neglected. For this reason, results on the integration order for the AO and IO tests were carried on to the cointegration analysis.

4.3 Results for Oil Prices

Even though the previously mentioned discussion on endogeneity issues formulated by Hamilton (2005) suggested us to evaluate nominal prices, in this Section we aim at determining the order of integration of both the nominal series and the 2013 adjusted oil price series from the BP database.[9] Structural break related analysis of price series is a relevant part of the analysis: breaks related to sudden changes in economic policies or to sudden shocks in political stability are very likely to have direct effects on oil prices, which in turn could affect real economy through the channels we already discussed in the short/medium run, as well as altering the structure of the relationship between GDP and oil prices in the long run. As we have already mentioned, we employ a bottom-up approach to test for the order of integration of the oil series, thus starting from the series first differences. Employing again the unit root tests introduced above, we test both the real and the nominal oil price in first differences with one and two endogenous breaks in Tables 11 and 12, then proceed to test the series in levels and present the output in Table 10, where we also show the output for the Ng and Perron (2001) tests both in first differences and levels. Our results show that all the modified M-tests would give evidence that the oil series, both real and nominal, are I(1). It is worth noticing that the existence of breaks in any series might result in the oil price variable appearing more integrated than it actually is. When the one break tests and the two break tests were carried out, the former tests suggested the variable might be I(2), while the latter tests, in particular the Clemente, Montanes, and Reyes (1998) tests in their additive outlier specification, again indicated oil prices as an I(1) variable, furthermore highlighting two historically relevant breaks in 1972 and 1984.[10] These results allowed us to carry out the following cointegration tests considering the oil price series as I(1).

[9] We recall that the discussion on the order of integration of macro-economic variables is mainly connected to the stream of analysis related to persistence of shocks or absence of time-wise memory in GNP and Oil price series which started with the seminal work of Perron (1989).
[10]The first date being one year before the first oil crisis price spike, the second date corresponding to the year before the Saudi Arabia induced oil price glut and sudden decline in oil prices. For a recent historic overview on the two events, see Jones (2012).

Table 2. Unit root tests, Ng and Perron (2001)

Country	**Variable**	**Mz**	**MZ_t**	**MSB**	**MP_t**
DZA	GDP_t	-0.42	-0.36	0.86	39.10
AGO	GDP_t	-4.65	-1.50	0.32	19.42
	GDP_t	-9.19**	-2.14**	0.23**	2.67**
	GDP_t	-6.19	-1.70	0.28	14.68
ARE	GDP_t	-0.40	-0.40	1.00	50.10
ECU	GDP_t	-5.57	-1.66	0.30	16.34
	GDP_t	-9.57**	-2.16**	0.23**	2.66**
	GDP_t	-3.69	-1.34	0.36	24.46
GAB	GDP_t	-1.43	-0.83	0.58	16.84
IDN	GDP_t	-5.71	-1.64	0.29	15.86
	GDP_t	-16.42***	-2.87***	0.17***	1.49***
	GDP_t	-6.68	-1.80	0.27	13.67
IRN	GDP_t	-0.26	-0.36	1.37	92.62
IRQ	GDP_t	-6.60	-1.76	0.27	13.83
	GDP_t	-22.82**	-3.38**	0.15**	1.08**
	GDP_t	-8.89	-2.10	0.24	10.27
KWT	GDP_t	-1.15	-0.76	0.66	21.30
	GDP_t	-8.61	-2.06	0.24	10.63
LBY	GDP_t	-0.82	-0.31	0.38	12.38
NGA	GDP_t	-2.52	-1.09	0.43	35.03
	GDP_t	-7.50*	-1.93*	0.26*	3.28*
	GDP_t	-2.99	-1.07	0.36	26.77
QAT	GDP_t	-2.48	-1.11	0.45	9.87
	GDP_t	-3.77	-1.21	0.32	21.87
SAU	GDP_t	-0.79	-0.63	0.80	31.02
VEN	GDP_t	-2.55	-1.08	0.42	33.94
	GDP_t	-24.66***	-3.50***	0.14***	1.02***
	GDP_t	-5.55	-1.66	0.30	16.42

***Denotes significance at the 1% level,
**Denotes significance at the 5% level,
*Denotes significance at the 10% level.
All the frequency zero spectrum calculations for the modified tests were based on an auxiliary detrended GLS autoregression of the natural log of GDP, where lagged di¤erences were chosen by the Modified Aikake selection criteria, starting from a maximum lag based on Schwert (1989) rule of thumb. A trend and a constant were included in the tests when the variables were analyzed in levels, while a constant only was included in the tests when the variables were analyzed in first differences.

Table 3. Zivot and Andrews Test (1992), One break

Country	Variable	ADF	T1	MAIC Lags
DZA	GDP_t	-4.55	1987	1
AGO	GDP_t	-4.44	1993	0
ARE	GDP_t	-3.63	1982	0
ECU	GDP_t	-5.14	1973	1
GAB	GDP_t	-4.95	1974	0
IDN	GDP_t	-5.59**	1998	0
IRN	GDP_t	-7.05***	1969	1
IRQ	GDP_t	-4.34	1991	1
KWT	GDP_t	-4.68	1990	0
LBY	GDP_t	-2.94	1973	1
NGA	GDP_t	-2.72	1980	0
QAT	GDP_t	-2.62	1984	0
SAU	GDP_t	-4.22	1972	1
VEN	GDP_t	-3.13	1980	0

***Indicates significance at 1% level, **indicates significance at 5% level, *indicates significance at 10% level. All basic specifications accomodate a trend and one break is allowed in both the intercept and the deterministic trend.

Table 4. Lumsdaine and Papell Test (1997), Two Breaks

Country	Variable	ADF	T1	T2	MAIC Lags
DZA	GDP_t	-9.61***	1983	1996	1
AGO	GDP_t	-4.74	1992	2005	0
ARE	GDP_t	-5.08	1982	1989	0
ECU	GDP_t	-7.42***	1972	1988	1
GAB	GDP_t	-7.64***	1970	1977	0
IDN	GDP_t	-8.36***	1968	1997	0
IRN	GDP_t	-7.45***	1969	1979	1
IRQ	GDP_t	-5.93	1978	1990	1
KWT	GDP_t	-6.01	1978	1992	0
LBY	GDP_t	-4.78	1978	1992	1
NGA	GDP_t	-4.94	1968	1996	0
QAT	GDP_t	-4.02	1986	1991	0
SAU	GDP_t	-5.34	1972	2003	1
VEN	GDP_t	-4.68	1979	2001	0

***Indicates significance at 1% level, **Indicates significance at 5% level, *Indicates significance at 10% level. All basic specifications accomodate a trend and two breaks are allowed in both the intercept and the deterministic trend.

Table 5. Perron and Vogelsang (1992), AO, One Break

Country	Variable	ADF	T1	P-value
DZA	GDP_t	-2.32	1985	0.284
AGO	GDP_t	-1.71	1991	0.422
ARE	GDP_t	-1.45	1984	0.18
ECU	GDP_t	-5.66**	1978	0.006
GAB	GDP_t	-3.03	1974	0.016
IDN	GDP_t	-5.56**	1996	0.027
IRN	GDP_t	-3.17	1967	0.015
IRQ	GDP_t	-7.81**	1989	0.829
KWT	GDP_t	1.96	1989	0.869
LBY	GDP_t	-9.62**	1978	0.048
NGA	GDP_t	-5.51**	2002	0.032
QAT	GDP_t	-5.49**	1990	0
SAU	GDP_t	-2.45	1974	0.001
VEN	GDP_t	-3.57	2001	0.681

**Significant at 5%. GtoS (General to specific test on significance of lagged variable) procedure employed to select number of lags. The test for first differences accommodates a single shift in the mean. Additive outlier methodology.

Table 6. Perron and Vogelsang (1992), IO, One Break

Country	Variable	ADF	T1	P-value
DZA	GDP_t	-4.10	1978	0.03
AGO	GDP_t	-6.16**	1992	0.014
ARE	GDP_t	-4.54**	1987	0.581
ECU	GDP_t	-5.80**	1974	0.03
GAB	GDP_t	-7.93**	1975	0
IDN	GDP_t	-7.51**	1997	0.609
IRN	GDP_t	-2.56	1977	0.141
IRQ	GDP_t	-11.37**	1990	0.961
KWT	GDP_t	-4.45**	1990	0.014
LBY	GDP_t	-9.12**	1973	0.008
NGA	GDP_t	-5.42**	2001	0.117
QAT	GDP_t	2.69	1992	0.224
SAU	GDP_t	-3.11	1973	0
VEN	GDP_t	-6.66**	2002	0.053

**Significant at 5%. GtoS (General to specific test on significance of lagged variable) procedure employed to select number of lags. The test for first differences accomodates a single shift in the mean. Innovation outlier methodology.

Table 7. Clemente, Montanes and Reyes (1998), AO, Two Breaks

Country	Variable	ADF	T1	T2
DZA	GDP_t	-2.06	1984	1985
AGO	GDP_t	-2.41	1991	2003
ARE	GDP_t	-7.33**	1982	1986
ECU	GDP_t	-7.26**	1970	1976
GAB	GDP_t	-2.58	1971	1974
IDN	GDP_t	-3.96	1980	1996
IRN	GDP_t	-6.74**	1974	1986
IRQ	GDP_t	-9.42**	1989	2001
KWT	GDP_t	-2.28	1980	1989
LBY	GDP_t	-10.10**	1978	1996
NGA	GDP_t	-4.40	1974	2002
QAT	GDP_t	-5.74**	1983	1994
SAU	GDP_t	-5.62**	1979	1985
VEN	GDP_t	-6.69**	1978	2001

**Significant at 5%. GtoS (General to specific test on significance of lagged variable) procedure employed to select number of lags. The test for first differences accommodates two shifts in the mean. AO methodology.

Table 8: Clemente, Montanes and Reyes (1998), IO, Two Breaks

Country	Variable	ADF	T1	T2
DZA	GDP_t	-4.42	1984	1995
AGO	GDP_t	-3.37	1992	2003
ARE	GDP_t	-7.18**	1981	1987
ECU	GDP_t	-7.54**	1971	1975
GAB	GDP_t	-9.51**	1972	1975
IDN	GDP_t	-11.40**	1966	1997
IRN	GDP_t	-7.13**	1976	1981
IRQ	GDP_t	-13.46**	1990	2002
KWT	GDP_t	-6.39**	1978	1990
LBY	GDP_t	-9.86**	1979	1994
NGA	GDP_t	-5.90**	1969	2001
QAT	GDP_t	-2.69	1984	1992
SAU	GDP_t	-3.18	1973	1981
VEN	GDP_t	-6.05**	1979	2002

**Significant at 5%. GtoS (General to specific test on significance of lagged variable) procedure employed to select number of lags. The test for first differences accomodates two shifts in the mean. IO methodology.

Table 9. Summary Results, Unit Root Tests

Tests	I(1)
Ng Perron Tests	AGO ECU IDN IRQ NGA VEN
One Break AO	ECU IRQ LBY NGA QAT
One Break IO	AGO ARE ECU GAB IRQ KWT LBY NGA VEN
Two Breaks AO	ARE IRQ LBY QAT SAU VEN
Two Breaks IO	ARE IRQ KWT LBY NGA VEN
Tests	**I(0)**
Ng Perron Tests	-
One Break AO	IDN
One Break IO	IDN
Two Breaks AO	ECU IRN
Two Breaks IO	ECU GAB IDN IRN

5. Cointegration Tests

Given oil was found to be an I(1) variable in the previous Section, we proceeded to run a series of tests in order to check whether or not the oil price might be cointegrated with GDP. The economic intuition behind such relationship has to do with the extent to which economies depend on commodities exports, which qualifies them as commodity dependent countries. For such reason, even when we might expect a spike or a decrease in oil prices to start feeding into economic activity with some delay, affecting GDP in the short and medium run, our interest lies first of all in checking whether or not oil prices and GDP in OPEC countries drift together in time following an equilibrium path. We employ the Gregory and Hansen (1996a) cointegration test, allowing for a single regime shift or a complete structural change in both deterministic and stochastic components of the cointegrating relationship, and two vector cointegration based tests, the Johansen and Juselius (1990) test and the Johansen, Mosconi, and Nielsen (2000) cointegration test, the latter allowing for an exogenous structural break in the deterministic trend.[11]

[11]Another feasible and equivalent choice while testing for cointegration would be the "de-trended" cointegration test by Saikkonen and Lütkepohl, (Saikkonen and Lütkepohl (2000a), Saikkonen and Lütkepohl (2000b), Saikkonen and Lütkepohl (2000c)) which apply a Johansen type test to a detrended series with a deterministic term. Such test, however, would have been unfeasible for a structural break in the deterministic trend term.

Table 10. Nominal and Real Oil Price Unit Root Tests

Ng and Perron (2001) Tests					
Value	**Variable**	**Mz_a**	**MZ_t**	**MSB**	**MP_t**
nominal	OIL_t	-19.33***	-3.08***	0.16***	1.37***
real	OIL_t	-18.98***	-3.05***	0.16***	1.39***
nominal	OIL_t	-5.17	-1.60	0.31	17.62
real	OIL_t	-5.52	-1.64	0.30	16.43
Zivot and Andrews (1992) Test, One Break					
Value	**Variable**	**ADF**	**T1**	**MAIC lags**	
nominal	OIL_t	-3.12	1986	0	
real	OIL_t	-3.11	1986	0	
Lumsdaine and Papell (1992) Tests, Two Breaks					
Value	**Variable**	**ADF**	**T1**	**T2**	**MAIC Lags**
nominal	OIL_t	-5.75	1973	2000	0
real	OIL_t	-6.22	1973	1999	0

***Denotes significance at the 1% level, **Denotes significance at the 5% level, *Denotes significance at the 10% level. For the Ng-Perron tests, All the frequency zero spectrum calculations for the modified tests were based on an auxiliary detrended GLS auto-regression of the natural logarithm of GDP. A trend and a constant were included in the level tests, while a constant only was included in the first differences test. Both The Zivot-Andrews and the Lumsdaine-Papell tests are based on model C specification, allowing for both a break in the intercept and the trend. In all tests, lagged di¤erences were chosen by the Modified Aikake selection criteria, starting from a maximum lag based on Schwert (1989) rule of thumb, $P_{max} = 12 * (\frac{T}{100})^{\frac{1}{4}}$

Table 11. Perron and Vogelsang (1992), One Break

Nominal Oil Price				
Specification	**Variable**	**ADF**	**T1**	**p-value**
AO	OIL_t	-2.14	1972	0.442
IO	OIL_t	-2.78	1973	0.649
Real Oil Price				
Specification	**Variable**	**ADF**	**T1**	**p-value**
AO	OIL_t	-2.07	1972	0.517
IO	OIL_t	-2.78	1973	0.717

**Indicates significance at 5%. Trend term not included (model A). AO stands for additive outlier method, IO for innovative outlier method. T1 indicates the endogenous break date.

Table 12. Clemente, Montanes and Reyes (1998), two breaks

Nominal Oil Price				
Specification	**Variable**	**ADF**	**T1**	**T2**
AO	OIL_t	-7.30**	1972	1984
IO	OIL_t	-3.32	1973	1997
Real Oil Price				
Specification	**Variable**	**ADF**	**T1**	**T2**
AO	OIL_t	-7.33**	1972	1984
IO	OIL_t	-3.51	1973	1997

**Indicates significance at 5%. Trend term not included (model A). AO stands for additive outlier method, IO for innovative outlier method. T1 and T2 indicate the endogenous break dates.

5.1. Residual Based Approach

In this Sub-section, we offer an overview of the three equation models from Gregory and Hansen (1996a) which we selected to test for cointegration in a residual based framework. The modified ADF tests proposed by the authors offer the advantage of being non-informative with respect to the time of a structural break, and as such (partially) prevent informal time series analysis such as a visual examination of the time series plot, offering a way to retrieve endogenously the suspected time of the structural break. The basic Model specification would start defining a standard single-equation cointegrating regression:

$$y_t = \mu + \alpha x_t + \varepsilon_t \tag{2}$$

where variables y_t and x_t are assumed to be I(1) and the residuals stationary. If this long run relationship were to naturally hold, the intercept and the estimated coefficient would need to be constant over time. However, such set-up does not appear to be the case in many application, such as ours.[12] For such reason, structural changes, reflected by a change in the intercept or the slope, need to be addressed. In order to do that, the authors introduce a break dummy variable:

$$\varphi_t = 0 \; if \; t \leq [T_\tau]$$

$$\varphi_t = 1 \; if \; t > [T_\tau]$$

Where τ represents the unknown relative time of the structural change, T is the number of observations in the series, and the brackets denote the integer part of the product. From the basic set-up, the authors illustrate five different specifications, three of which were considered in our analysis. The first

[12]This explains why we pointed out how the use of such methodology in presence of breaks would "partially" prevent an informal graphical analysis.

specification, defined the level shift with trend model in Gregory and Hansen (1996a) (C/T model), is represented as:

$$y_t = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha x_t + \varepsilon_t \tag{3}$$

which would basically imply a shift of the cointegrating relationship, captured by term $\mu_2 \varphi_{t\tau}$ which keeps gravitating around a non-zero mean. An alternative parametrization we considered was:

$$y_t = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha x_t + \varepsilon_t \tag{4}$$

were only a shift in the intercept in the cointegrating relationship is considered and neither the trend term nor the cointegrating relationship are able to rotate at breakpoint $t = \tau T$. Although the parametrization in (4) might appear less indicative than those in (3) and (5) after a visual inspection of the data, for the sake of a formally consistent analysis we choose to report it as well. The last speci…cation we employ is the most general one, which Gregory and Hansen (1996b) define as a complete regime shift with a shift in the trend (C/S/T model). This relationship can be formalized as in (5)

$$y_t = \mu_1 + \mu_2 \varphi_{t\tau} + \beta_1 t + \beta_2 t \varphi_{t\tau} + \alpha_1 x_t + \alpha_2 x_t \varphi_{t\tau} + \varepsilon_t \tag{5}$$

where μ_1 and μ_2 would denote respectively the original intercept and its change at time $t = \tau T$, β_1 and β_2 would denote the slope of the trend and its change at the time of the break, and α_1 and α_2 would represent the slope of the cointegrating relationship and its change. Similarly to Zivot and Andrews (1992), Gregory and Hansen (1996a) set up a series of cointegration test statistics for every possible break point in a bounded interval required for tractability of the data.[13] The smallest value of the augmented Dickey-Fuller test statistic on the residuals of the previous parametrizations across all the possible set of breaks in the selected interval is taken as the relevant statistic:

$$ADF = \inf_{\tau \in T} [\, ADF\ (\tau)]$$

with T being the compact subset inside which the minimum test statistic is taken. We next report and discuss the results of the Gregory and Hansen (1996a) and Gregory and Hansen (1996b) tests for Angola, Ecuador, Indonesia, Iraq, Nigeria and Venezuela, Qatar, Libya, Arab Emirates and Saudi Arabia. These were the countries whose GDP were found to be integrated of order one in the unit root Section. As we mentioned before, the test were carried out for a break in the

[13]Following the suggestions of the authors, we ran the tests taking for each possible break point in the interval ([:15n]; [:85n]).

intercept, trend, and cointegrating coefficient of each relationship (model C/S/T, Tables 13 and 14), a single break in the level assuming a constant and a time trend in the cointegrating relationship (model C/T, Tables 15 and 16), and a single break in the level assuming a constant only in the cointegrating relationship (model C, Tables 17 and 18[14]).

The tests show a singular result. Whenever no hint of cointegration could be found between oil prices and GDP in the majority of the analyzed sub-set of OPEC countries, one country out of the selected ten, Saudi Arabia, presents clear evidence of cointegration, depending on the set up of the deterministic component of the test, at least at the very liberal 10% level when a trend and an intercept are included alongside with what Gregory and Hansen (1996b) define as a regime shift (allowing for a break in both the intercept and the trend in the cointegrating relationship) for the case where the relationship between real GDP and nominal oil prices was analyzed, and at the 5% level with the same deterministic set-up but considering the GDP-real oil price relationship instead. Such result appears to only partially agree with Lescaroux and Mignon (2008), where the relationship could be verified as well for Iran, Iraq, and Qatar.[15] This result also appears to be in line with Saudi Arabia's position as OPEC's "swing producer", and suggests that the state of dependency of the country on oil production might be more intense than in all the other OPEC countries. A last, useful inside is that the relationship, found only in Saudi Arabia, represents a further indication that the economy of the country is still far from any attempt of diversification, evidently protected more by energy prices upswings and collapses of the oil price than in the past thanks to its price setting power and its excess reserves rather than by any attempt at structural reforms of its economy and development of alternative sectors.

[14] The possibility of a mean stationary cointegrating vector does not appear rigorous given that the GDP series would appear to show a trend while the first differenced oil series would not. Nevertheless, we report the results for model C for completeness of the analysis and to overcome, as we already stated, the limitations of a visual analysis. The number of lagged differences was taken following the MAIC criterion used in the unit root analysis.

[15]While accounting for a single structural break, we could find some pretty weak evidence of cointegration for Iraq and United Arab Emirates. However, this would pretty much depend on the lag choice. On the contrary, even when we employed a much less conservative number of lags to test for Saudi Arabia according to the unmodified Aikake criterion, the outcome of the tests did not change, staying put under the 10% critical value.

Table 14. Gregory and Hansen (1996b), GDP and Real Oil, C/S/T

Country	ADF	Lags	[T]
AGO	-3.82	0	1991
ECU	-4.54	1	2000
IDN	-3.90	0	1998
IRQ	-4.75	1	1988
NGA	-4.66	0	1979
VEN	-3.74	0	2000
QAT	-3.95	0	2002
LBY	-5.22	1	1979
ARE	-4.88	0	1988
SAU	-5.87**	1	1970

**Denotes significance at the 5% level. A trend and intercept are assumed in the cointegrating relationship, the set up includes a complete regime shift (a contemporaneous break in intercept, trend and cointegrating relationship). Critical values are -6.02 (1%), -5.50 (5%), -5.24 (10%) and were taken from Gregory and Hansen (1996b), pag. 559.

Table 13. Gregory and Hansen (1996b), GDP and nominal oil, C/S/T

Country	ADF	Lags	[T]
AGO	-3.86	0	1991
ECU	-4.13	1	2000
IDN	-3.86	0	1998
IRQ	-4.72	1	1988
NGA	-4.85	0	1979
VEN	-3.62	0	2000
QAT	-3.72	0	2003
LBY	-5.22	1	1979
ARE	-4.75	0	1988
SAU	-5.48*	1	1977

*Denotes significance at the 10% level. A trend and intercept are assumed in the cointegrating relationship, the set up includes a complete regime shift (a contemporaneous break in intecept, trend and cointegrating relationship). Critical values are -6.02 (1%), -5.50 (5%), -5.24 (10%) and were taken from Gregory and Hansen (1996b), pag. 559.

Table 15. Gregory and Hansen (1996a), GDP and Nominal Oil, C/S/T

Country	ADF	Lags	[T]
AGO	-4.04	0	1990
ECU	-4.35	1	1972
IDN	-3.61	0	1998
IRQ	-4.02	1	1988
NGA	-3.90	0	1979
VEN	-3.55	0	1980
QAT	-4.28	0	1995
LBY	-4.14	1	1971
ARE	-3.74	0	1992
SAU	-4.26	1	1970

A trend and intercept are assumed in the cointegrating relationship, the set up includes a single level shift in the mean of the cointegrating relationship. Critical values are -5.45 (1%), -4.99 (5%) and -4.72 (10%) and were taken from Gregory and Hansen (1996a), pag. 33, tab. 1A.

Table 16. Gregory and Hansen (1996a), GDP and Real Oil, C/T

Country	ADF	lags	[T]
AGO	-4.01	0	1991
ECU	-4.37	1	2000
IDN	-3.67	0	1998
IRQ	-3.88	1	1988
NGA	-4.19	0	1979
VEN	-3.44	0	1980
QAT	-4.25	0	1995
LBY	-4.47	1	1971
ARE	-3.57	0	1992
SAU	-4.36	1	1970

A trend and intercept are assumed in the cointegrating relationship, the set up includes a single level shift in the mean of the cointegrating relationship. Critical values are -5.45 (1%), -4.99 (5%) and -4.72 (10%) and were taken from Gregory and Hansen (1996a), pag. 33, tab. 1A.

Table 17. Gregory and Hansen (1996a), GDP and Nominal Oil, C

Country	ADF	lags	[T]
AGO	-4.85	0	1989
ECU	-4.28	1	1987
IDN	-3.98	0	1986
IRQ	-3.73	1	1994
NGA	-3.23	0	1974
VEN	-3.62	0	1988
QAT	-3.99	0	1995
LBY	-2.71	1	1990
ARE	-4.11	0	1991
SAU	-3.77	1	1970

An intercept is assumed in the cointegrating relationship, the set up includes a single level shift in the mean of the cointegrating relationship. Critical values are -5.13 (1%), -4.61 (5%), -4.34 (10%) and were taken from Gregory and Hansen (1996a), pag. 33 table 1A.

Table 18. Gregory and Hansen (1996a), GDP and real oil, C

Country	ADF	lags	[T]
AGO	-5.33	0	1996
ECU	-4.17	1	1987
IDN	-4.07	0	1986
IRQ	-3.67	1	1989
NGA	-2.98	0	1994
VEN	-4.04	0	1988
QAT	-3.19	0	1986
LBY	-3.50	1	2001
ARE	-3.59	0	1991
SAU	-3.62	1	1987

An intercept is assumed in the cointegrating relationship, the set up includes a single level shift in the mean of the cointegrating relationship. Critical values are -5.13 (1%), -4.61 (5%), -4.34 (10%) and were taken from Gregory and Hansen (1996a), pag. 33 table 1A.

5.2 Vector Based Cointegration Analysis

This Section shows the results based on the Johansen and Juselius (1990) trace test and the Johansen, Mosconi, and Nielsen (2000) exogenous one break test for cointegration. In particular, among the three nested long run matrix trace tests, we choose to follow the most general specification, the $H_l(r)$ test from the afore-mentioned paper, to allow for the time series in levels to have a broken trend while allowing for a broken trend in the cointegrating relationship as well. To grant the tractability of the data, the standard specification of the test was augmented with an unrestricted exogenous impulse dummy at the time of the break, an unrestricted exogenous broken trend dummy, and a restricted exogenous trend and break interaction dummy in the cointegrating relationship. Following Johansen, Mosconi, and Nielsen (2000), we start with a canonical specification of a cointegrated vector auto-regressive model:

$$\Delta X_t = \Pi X_{t-1} + \Pi_1 t + \varepsilon_t \qquad (6)$$

where, to simplify notation in (6), we have avoided any additional lag. In this model, where X_t represents the vector of variables , and the error ε_t is assumed to be identically and independently normally distributed with finite variance and a mean equal to 0, cointegration will appear if the matrix of the long run relationships has a reduced rank and could thus be described as the product $\Pi = \alpha\beta'$, where both α and β are $(p * r)$ full rank matrices. However, this would imply the presence of a quadratic trend in the level variables, which we choose not to assume due to lack of evidence on the matter. For such reason, the quadratic trend can be assumed away defining $\Pi_1 = \alpha * \gamma'$. That basically means restricting the trend to the cointegrating relationships to rule out the quadratic trend. This way, the specification for the standard Johansen trace test was

$$\Delta X_t = \alpha(\beta' X_{t-1} + \gamma' t) + \mu + \varepsilon_t \qquad (7)$$

which involved calculating the reduced rank of the combined matrix, $(\Pi, \Pi_1) = \alpha(\beta', \gamma')'$. This allowed us to formulate the cointegration hypothesis in terms of the rank of Π in conjunction with Π_1:

$$H_l(r) : \; rank(\Pi, \Pi_1) \leq r$$

If we assume the sample might have q sub-periods, or equivalently $q - 1$ breaks, the base-line model in (6) can be rewritten q times, conditional on the first k observations of each sub-sample, as q "break models":

$$\Delta Y_t = \left(\Pi, \Pi_j\right)\begin{pmatrix} Y_{t-1} \\ t \end{pmatrix} + \mu_j + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-1} + \varepsilon_t \qquad (8)$$

where $j = 1,2,\dots,q$, such that $j = 2$, and Π, Π_j and μ_j are $(p*1)$ matrices, with p being equivalent to the number of time series in Y_t. Finally, instead of writing q equations, the sub-samples are accommodated by defining the following matrices:

$$D_{j,t} = \left(1,\dots,D_{q,t}\right)', \qquad \mu = \left(\mu_1,\dots,\mu_q\right), \qquad \gamma = (\gamma_1,\dots,\gamma_q)$$

with dimensions $(q*1)$; $(p*q)$; $(q*r)$. This allows to rewrite equation (8) as:

$$\Delta Y_t = \alpha \binom{\beta}{\gamma}' \binom{Y_{t-1}}{tD_{t-k}} + \mu D_{t-k} + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-1} + \sum_{i=0}^{k-1} \sum_{j=2}^{q} \kappa_{j,i} I_{j,t-i} + \varepsilon_t \quad (9)$$

where the intervention dummies $D_{j,t}$, $D_{j,t-k}$ and $D_{j,t}$ are defined[16] as:

$D_{j,t} = 1 \text{ for } T_{j-1} + 1 > t \geq T_j$

$D_{j,t} = 0 \text{ otherwise}$

$D_{j,t+k} = 1 \text{ for } T_{j-1} + 1 + k > t \geq T_j + k$

$D_{j,t+k} = 0 \text{ otherwise}$

$for\, j = 2,\dots,q$

where, to clarify notation, t represents the time trend, T_j is the last observation of sub-sample j, and the impulse break dummy $I_{j,t}$ is defined as:

$I_{j,t} = 1 \text{ for t} = T_{j-1} + 1$

$I_{j,t} = 0 \text{ otherwise}$

Since in our case the number of sub-samples is equal to two, the impulse break dummy $I_{j,t}$ will be just equal to the first difference of the break dummy at time t: $\Delta D_{j,t} = I_t$. The use of impulse break dummies is justified by the need to restrict the residuals to 0 given the initial value in the second sub-period. Finally, the

[16] Notice that the notation we are employing here borrows extensively from Joyeux (2007) and is different from the one used by Johansen, Mosconi, and Nielsen (2000), who define the impulse break dummies as $D_{j,t}$ and the break dummies as the sub-sample sum of a set of indicator dummies, $E_{j,t}$.

likelihood ratio test devised by Johansen, Mosconi, and Nielsen (2000), and based on the squared sample canonical correlations of the long run matrix and the first differences vector of variables, will be:

$$LR\{H_l(r)|H_l(p)\} \tag{10}$$

The tests have different asymptotic properties than the standard tests without breaks given the presence of the exogenous dummies. As a consequence of that we report the different critical values for the $H_l(r)$ tests taken from Giles and Godwin (2012) in the output Tables 21 and 22, while results for the baseline specification (equation (6)) are reported in Tables 19 and 20. Results from the Johansen and Juselius (1990) and the Johansen, Mosconi, and Nielsen (2000) tests suggest the presence of a cointegrating relationship in both Angola and Saudi Arabia. However, since none of the residual based tests indicated Angola as a possible candidate to the long run equilibrium, we could conclude that only Saudi Arabia results, which consistently showed proof of cointegration between oil and GDP across all tests, could be carried on to the next Section for the weak causation tests.

5.2 Weak Causality Tests for Saudi Arabia

To complete the analysis, based on the results of the previous Sections, we present the results of the Granger causality test, a series of Wald test based on linear restrictions we specified in order to test for weak causality in the long run, short run and in the "joint run". Since the C/S/T test allowed us to find evidence of cointegration in Saudi Arabia, we take advantage of the Engle-Granger representation theorem to set up a series of Granger causality tests in order to verify whether or not the existence of a long run relationship between GDP and oil price might have some positive/negative effect on economic activity in Saudi Arabia, or better be helpful in forecasting growth in such country.[17] Following the outcome of the Gregory and Hansen (1996b) C/S/T test, we ran a first step regression of the cointegrating relationship. Recalling equation (5), we calculated the following equilibrium residuals:[18]

[17]We are aware that the Granger causality test is in reality a predictive ability test of the lags of a series with respect to another variable. For such reason, we choose to apply and interpret this test as evidence that a broken long run relationship might be able to help forecasting future growth, so that the test would at least give us an indication of weak causality direction.

[18]We report the estimates of the long run elasticities in Table 23. The oil price variable consistently enters in both the real and the nominal oil price regressions with a negative sign, confirming our expectations from Section 2.

$$\hat{\varepsilon}_t = y_t - \hat{\mu}_1 - \hat{\mu}_2\varphi_{t\tau} - \hat{\beta}_1 t - \hat{\beta}_2 t\varphi_{t\tau} - \hat{\alpha}_1 x_t - \hat{\alpha}_2 x_t\varphi_{t\tau} - \varepsilon_t \quad (11)$$

Then, using the residuals from (11), we estimated the following error correction models:

$$\Delta GDP_t = \mu_1 + \lambda_1[ECT_{t-1}] + \sum_{t=1}^{k}\beta_{i,1}(\Delta oil_{t-k}) + \sum_{t=1}^{k}\beta_{i,2}(\Delta GDP_{t-k}) + \varepsilon_t \quad (12)$$

$$\Delta oil_t = \mu_2 + \lambda_2[ECT_{t-1}] + \sum_{t=1}^{k}\beta_{i,1}(\Delta GDP_{t-k}) + \sum_{t=1}^{k}\beta_{i,2}(\Delta oil_{t-k}) + \varepsilon_t \quad (13)$$

the specifications in (12) and (13) were tested twice for $k = (1;\ 2)$, considering alternatively the error correction term from the relationship between GDP and nominal oil prices (ECT nom) and the error correction term from the relationship between GDP and real oil prices (ECT real). The error correction term, consistently with the previous cointegration analysis, contains not only an intercept and a deterministic trend, but also the structural break we identified for Saudi Arabia through the Gregory and Hansen (1996b) at time t = 1970 for the nominal oil price and t = 1977 for the real oil price. Testing for long run weak causality in (12) and (13) required us to test the speed of adjustment terms, that is $H_0: \lambda_1 = 0$ and $H_0: \lambda_2 = 0$. To test for strong Granger causation, we then carried out a joint test on the lagged first differences of the variables and the error correction coefficients, that is we tested for $H_0: \lambda_1 = \beta_{1,1} = \beta_{1,2} = \cdots = \beta_{1,k} = 0$ in equation (12) and $H_0: \lambda_2 = \beta_{1,1} = \beta_{1,2} = \cdots = \beta_{1,k} = 0$ in equation (13). The joint run test in particular aims at checking which variable bears the burden of a short-run adjustment to re-establish a long-run equilibrium after a shock to the system. Results, shown in Tables 24 and 25 show that weak causality in the long term runs uniquely from the error correction term to GDP growth, as generally all the error correction adjustment coefficients are significant at the 5% level.

Our results on Saudi Arabia are remarkably different from Lescaroux and Mignon (2008) findings, where all the countries which were initially selected by these authors for the long run causation analysis and that resulted in a single univocal direction of causation from oil prices to GDP in the long run (namely, Iran, Qatar and United Arab Emirates) were not selected by us for such analysis due to lack of sufficient evidence on cointegration. In our tests, no lagged first differences of the independent variable in any specification proved to be significant. As for the tests based on the joint causation in the long and the short run, we could gather strong evidence that changes in the oil prices appear to Granger-cause GDP growth in Saudi Arabia jointly with the long run equilibrium correction, while no reverse causation was found as the null hypotheses $H_0: \lambda_2 = \beta_{1,1} = 0$ and $H_0: \lambda_2 = \beta_{1,1} = \beta_{1,2} = 0$ in equation (13) could never be rejected at all conventional confidence levels. The joint test of Δoil_{t-1} and ECT_{t-1} shows that both real and nominal oil prices indeed affect GDP_t, confirming the existence of a mechanism through which the volatility of oil prices exerts its influence on long-run and short run economic fluctuations of GDP through at least one of the channels we discussed in Sub-section 2.1.

Table 19. Johansen Trace Tests, Real GDP and Nominal Oil Prices

H_0	AGO	ECU	IDN	IRQ	NGA
r = 0	19.54	13.3	9.31	15.62	18.74
r = 1	4.41	4.97	4.03	3.21	6.65
H_0	**VEN**	**QAT**	**LBY**	**ARE**	**SAU**
r = 0	19.29	37.84*	29.93*	16.66	24.99*
r = 1	4.34	11.63	13.01	2.24	8.51

*Denotes rejection of the null hypothesis at the 5% level. Tests are carried out assuming a continuous deterministic trend in the level data and in the cointegrating equation. Due to the results of the AIC and BIC information criteria and the general to specific analysis on a VAR with a maximum lag length set according to Schwert (1989) rule of thumb, a single lag for structure for the short run component of the model was adopted.

Table 20. Johansen Trace Tests, Real GDP and Real Oil Prices

H_0	AGO	ECU	IDN	IRQ	NGA
r = 0	20.75	11.23	9.75	13.44	21.17
r = 1	4.6	3.95	3.41	2.87	5.29
H_0	**VEN**	**QAT**	**LBY**	**ARE**	**SAU**
r = 0	18.26	38.15*	22.14	16.93	21.48
r = 1	3.63	11.56	6.22	2.1	5.34

*Denotes rejection of the null hypothesis at the 5% level. Tests are carried assuming an intercept and a continuous deterministic trend in the level data and in the cointegrating equation. Due to the results of the AIC and BIC information criteria and the general to specific analysis on a VAR with a maximum lag length set according to Schwert (1989) rule of thumb, a single lag for structure for the short run component of the model was adopted.

Table 21. Johansen, Mosconi and Nielsen (2001) Trace Test, Real GDP and Nominal Oil Prices

H_0	AGO	ECU	IDN	IRQ	NGA
r = 0	42.57*	20.31	14.89	18.62	24.14
r = 1	9.96	0.76	1.31	4.13	5.91
break date	1991	2000	1998	1988	1979
5% critical value, r = 0	35.21	35.21	35.21	37.38	36.7
5% critical value, r = 1	17.79	17.79	17.79	18.92	18.59
H_0	**VEN**	**QAT**	**LBY**	**ARE**	**SAU**
r = 0	25.72	29.39	17.28	14.67	37.64*
r = 1	4.37	8.58	7.15	4.26	8.55
break date	2000	2002	1979	1988	1970
5% critical value, r = 0	35.21	35.21	36.7	36.7	32.97
5% critical value, r = 1	15.79	15.79	18.59	18.59	16.55

**Denotes rejection of the null hypothesis at the 5% level. The tests assume both the level data and the cointegrating equation might have a broken deterministic trend. The exogenous break dates are based on the results of the Gregory Hansen tests. Due to the results of the AIC and BIC information criteria and the general to specific analysis on a VAR with a maximum lag length set according to Schwert (1989) rule of thumb, a single lag for structure for the short run component of the model was adopted.

Table 22. Johansen, Mosconi and Nielsen (2001) Trace Test, Real GDP and Real Oil Prices

H_0	AGO	ECU	IDN	IRQ	NGA
r = 0	41.83*	18.49	15.43	17.42	26.69
r = 1	9.83	0.29	1.77	3.41	5.87
break date	1991	2000	1998	1988	1979
5% critical value, r = 0	35.21	35.21	35.21	37.38	36.7
5% critical value, r = 1	17.79	17.79	17.79	18.92	18.59
H_0	**VEN**	**QAT**	**LBY**	**ARE**	**SAU**
r = 0	24.94	30.22	16.96	13.31	35.55*
r = 1	4.71	9.14	6.73	4.12	5.45
break date	2000	2002	1979	1988	1970
5% critical value, r = 0	35.21	35.21	36.7	36.7	32.97
5% critical value, r = 1	17.79	15.79	18.59	18.59	16.55

**Denotes rejection of the null hypothesis at the 5% level. The tests assume both the level data and the cointegrating equation might have a broken trend. The exogenous break dates are based on the results of the Gregory Hansen tests. Due to the results of the AIC and BIC information criteria and the general to specific analysis on a VAR with a maximum lag length set according to Schwert (1989) rule of thumb, a single lag for structure for the short run component of the model was adopted.

7. Conclusions

This paper analyzed the presence of a long run equilibrium relationship in the OPEC countries between the real GDP and a nominal and a real oil price series since the creation of the group in 1960. Our estimates show that such relationship could not be verified for the majority of the countries in the group. However, cointegration between the GDP and the nominal and real oil price series in Saudi Arabia, once the presence of a structural break in 1970 had been accounted for, was accepted by both the residual and the VAR based approach we employed in the analysis. The existence of such relationship confirms the role of the country as the determinant "swing producer" of the group, showing a consistent pattern which dates back before the oil price glut of the eighties, underlying a long run pattern of dependence of the economy of Saudi Arabia on oil prices. By analyzing a series of Granger causality tests with underlying structural breaks, we conclude that, even though no direct short run causality linkage could be proved between oil prices and GDP, the existence of such relationship holds in the long run and appears to show some degree of predictive ability on GDP growth in Saudi Arabia.

Table 23. C/S/T Specification Based Regressions, Dependent Variable GDP

	Nominal Oil	**Real oil**
oil price	-0.29	-8.63
	(0.09)	(2.82)
	[0.01]	[0.01]
$\psi_{t\tau}$ *oil price	0.49	8.88
	(0.11)	(2.82)
	[0.00]	[0.01]
intercept	22.61	45.46
	(0.09)	(7.54)
	(0.00)	[0.00]
t $\psi_{t\tau}$ *intercept	2.25	-21.07
	(0.18)	(7.54)
	[0.00]	[0.01]
trend	0.29	0.02
	(0.02)	(0.09)
	[0.00]	[0.86]
$\psi_{t\tau}$ *trend	-0.27	0.01
	(0.19)	(0.10)
	[0.00]	[0.93]
observations	50	50
R-squared	0.98	0.98

Estimation of the long run elasticities based on C/S/T specification. Standard error in parentheses, p-values in square brackets.

Table 24. Causality Tests, k = 1

	Source of Causation	
	Short Run	
	$\Delta oil_{t-1} nom$	$\Delta oil_{t-1} real$
ΔGDP_t	0.66	0.54
	(0.42)	(0.47)
	ΔGDP_{t-1}	-
$oil_t nom$	0.00	-
	(0.99)	-
$oil_t real$	0.01	-
	(0.94)	-
	Long Run	
	$ECT_{t-1} nom$	$ECT_{t-1} real$
ΔGDP_t	23.73**	23.39**
	(0.00)	(0.00)
$oil_t nom$	0.93	-
	(0.34)	-
$oil_t real$	-	0.56
	-	(0.46)
	Joint Causality	
	$\Delta GDP_{t-1}, ECT_{t-1} nom$	$\Delta GDP_{t-1}, ECT_{t-1} real$
$oil_t nom$	0.58	-
	(0.56)	-
$oil_t real$	-	0.36
	-	(0.69)
	$\Delta oil_{t-1} nom, ECT_{t-1} nom$	$\Delta oil_{t-1} real, ECT_{t-1} real$
ΔGDP_t	11.92**	11.78**
	(0.00)	(0.01)

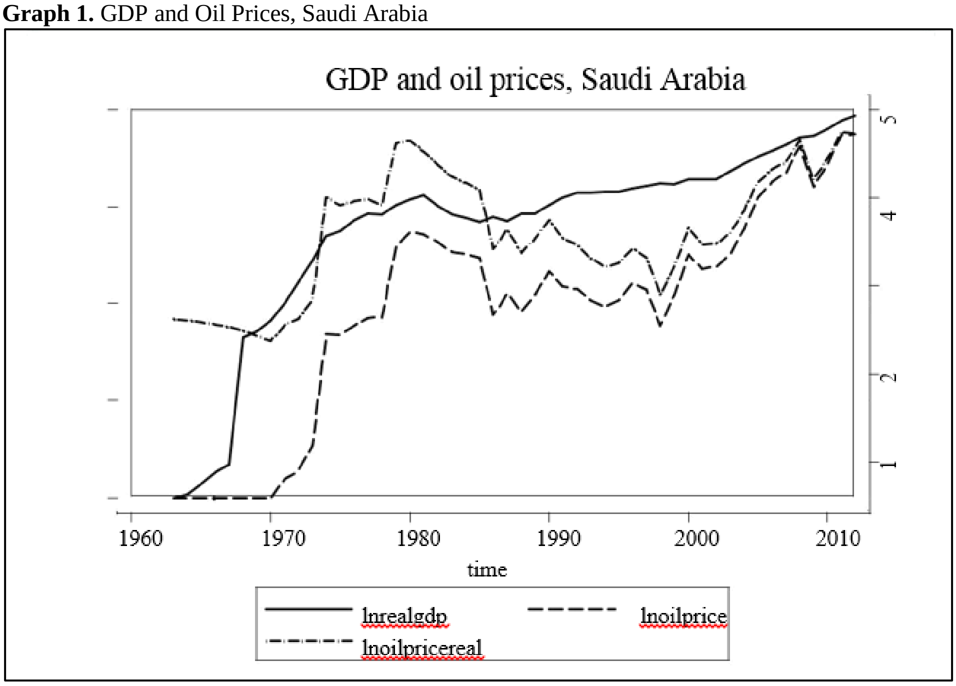
**Denotes significance at 5%. The table denotes F-statistics values. p-values are in parenthesis.

Table 25. Causality Tests, k = 2

	Source of Causation	
	Short Run	
	$\Delta oil_{t-1,t-2}nom$	$\Delta oil_{t-1,t-2}real$
ΔGDP_t	0.33	0.43
	(0.72)	(0.65)
	$\Delta GDP_{t-1,t-2}$	
$oil_t nom$	0.01	-
	(0.99)	-
$oil_t real$	0.03	-
	(0.97)	-
	Long run	
	$ECT_{t-1,t-2}nom$	$ECT_{t-1,t-2}real$
ΔGDP_t	30.37**	26.53**
	(0.00)	(0.00)
$oil_t nom$	0.86	-
	(0.36)	-
$oil_t real$	-	0.62
	-	(0.43)
	Joint Causality	
	$\Delta GDP_{t-1,t-2}, ECT_{t-1}nom$	$\Delta GDP_{t-1,t-2}, ECT_{t-1}real$
$oil_t nom$	0.41	-
	(0.75)	-
$oil_t real$	-	0.33
	-	(0.80)
	$\Delta oil_{t-1,t-2}nom, ECT_{t-1}nom$	$\Delta oil_{t-1,t-2}real, ECT_{t-1}real$
ΔGDP_t	10.17**	8.89**
	(0.00)	(0.01)

**Denotes significance at 5%. The table denotes F-statistics values. p-values are in parenthesis.

Graph 1. GDP and Oil Prices, Saudi Arabia



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