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An Analysis of Oil Production by OPEC Countries: Persistence, Breaks, and Outliers

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ABSTRACT

This study examines the degree of persistence, potential breaks and outliers of oil production for OPEC member countries within a fractional integration modelling framework using monthly data from January 1973 to October 2008. The results indicate there is mean reverting persistence in oil production with breaks identified in ten out of the fourteen countries examined. Thus, shocks affecting the structure of OPEC oil production will have persistent effects in the long run for all countries, and in some cases the effects are expected to be permanent.

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1. Introduction

According to the *Energy Information Administration*, in 2008 roughly 43% of the world's oil production was attributed OPEC member countries. Furthermore, OPEC member countries have approximately 70% of the proven oil reserves in the world. Table 1 displays country specific oil production for 1973 and 2008, as one can see oil production in Algeria, Angola, Ecuador, Iraq, Qatar, Saudi Arabia, and the United Arab Emirates (UAE) show an increase while Indonesia, Iran, Kuwait, Libya, Nigeria, and Venezuela show a decrease relative to 1973. In light of the OPEC's continued coordination of oil production policies to increase their monopoly power as a cartel and subsequent occurrences of coordination failures, understanding the time series behavior of OPEC oil production is critical in the assessment of the impact of oil shocks and structural breaks on both oil supply and the repercussions for global economic activity.[1]

Specifically, this study examines the degree of persistence, potential breaks and outliers of oil production for each OPEC member country within a fractional integration modelling framework. In particular, two important features commonly observed in oil production data are the persistence across time (Lien and Root, 1999; Kang et al. 2009) and breaks in production (Altinay and Karagol, 2004; Lee and Chang, 2005; Narayan and Smyth, 2008; Rao and Rao, 2009). Modelling the degree of persistence is important in that it can reflect the stability of production in a particular country and given the importance of oil production to other sectors of an economy the persistence of such shocks may be transmitted to other sectors of the economy and macroeconomic aggregates as well. Such transmission of shocks has implications for the effectiveness of

Table 1: OPEC Countries Oil Production

Countries	Oil production in barrels October in 2008	Oil production in barrels October in 1973
Algeria	1873.99	1059
Angola	1991	162
Ecuador	496.874	220
Indonesia	990	1447
Iran	4100	5977
Iraq	2327.578	1846
Kuwait	2628.738	3060
Libya	1745	2370
Nigeria	2185	2200
Qatar	924.756	600
Saudi Arabia	9400	7796
U. A. E.	2660.912	1669
Venezuela	2360	3381
Total	33683.85	31787

Source: Oil production from *OPEC Annual Statistics Bulletin*.

government intervention or stabilization policies. Breaks and outliers are other important features that are present in monthly oil production data which reflect shocks in oil production due to fluctuations in oil prices, changes in the world geopolitical climate, and country-specific socio-economic events, among others.[2] Indeed, if oil production is stationary I(0), shocks to oil production will be transitory and following major structural breaks in oil production, the supply of oil will return to its original equilibrium with the disruptions in oil production only having a temporary impact on economic activity. However, if oil production contains a unit root (i.e., if it is nonstationary I(1)), shocks to oil production will have persistent effects on the supply of oil with the disruptions in oil production having a permanent impact on economic activity (Narayan et al. 2008).[3] In the present paper we extend the models based on I(0) and I(1) hypotheses to the fractional I(d) case.

Despite the importance of oil as an energy source and the previous research on the oil industry, there are no studies that specifically analyse the persistence, breaks, and outliers associated with OPEC oil production. While studies consider, for example, oil consumption (Mohn and Osmundsen, 2008; Lean and Smyth, 2009), returns on investment in oil (Boone, 2001) and oil exhaustion (Tsoskounoglou et al. 2008; Höök and Aleklett, 2008; Karbassi et al. 2007), no studies have explored the long memory/persistence properties of OPEC oil production.

The remainder of this study is organised as follows. Section 2 presents a review of the previous literature. Section 3 details the methodology. Section 4 presents the data and the empirical results. Section 5 deals with the discussion of the results, while Section 6 provides concluding remarks.

2. Brief Overview of the Literature

As mentioned earlier, determining whether shocks to oil production are transitory or persistent is relevant in the formulation of energy-related policy as well as government stabilization policies. Though there have been a number of studies investigating the presence of a unit root in energy consumption (Chen and Lee, 2007; Narayan and Smyth, 2007; Hsu et al, 2008; Mishra et al, 2009; Lean and Smyth, 2009; Rao and Rao, 2009), only a few studies examine oil production.[4]

In the process of examining the economic, geological, and institutional determinants of oil production in the lower 48 U.S. states, Kaufmann and Cleveland

(2001) use the Augmented Dickey Fuller (ADF, Dickey and Fuller, 1979) unit root test with respect to oil production over the period 1938 to 1991. Their results fail to reject the null hypothesis of a unit root. Iledare and Olatubi (2006) investigate oil production in the Gulf of Mexico's Outer Continental Shelf using quarterly data from 1948 to 2000 for shallow water and from 1979 to 2000 for deep water, respectively. In both cases, the results of the ADF tests fail to reject the null hypothesis of a unit root.[5]

Narayan et al. (2008) explore the unit root properties of crude oil and NGL production for 60 countries using annual data from 1971 to 2003. Their analysis begins with the panel unit root tests by Maddala and Wu (1999), Breitung (2000), Levin et al (2002), Im et al (2003), and the panel stationarity test by Hadri (2000) each undertaken without allowance for a structural break. The results for the panel data sets without allowance for a structural break for the entire 60 country panel and the regional panels (OECD, Latin America, Central and Eastern Europe, Africa, Middle East, and Asia) provide mixed evidence of stationarity for crude oil and NGL production. However, further investigation using the panel LM unit root test by Im et al. (2005) with allowance for a structural break reveals that for the entire 60 country panel and five of the six regional panels (OECD, Latin America, Africa, Middle East, and Asia), the null hypothesis of a unit root in crude oil and NGL production is rejected at the 1% significance level while for the Central and Eastern Europe panel, the null hypothesis is rejected at the 10% level.

Maslyuk and Smyth (2009) apply the threshold unit root tests by Caner and Hansen (2001) using monthly data from January 1973 to December 2007 for crude oil production of 17 countries that include both OPEC and non-OPEC countries. Maslyuk and Smyth (2009) find the presence of threshold effects (i.e. non-linearities) in crude oil production over two regimes. Next, Maslyuk and Smyth (2009) test for a unit root against a non-linear stationary process in two regimes and a partial unit root process when the unit root is present in only one regime. Their results indicate that for 11 countries there is a unit root in both regimes (Indonesia, Kuwait, Nigeria, Qatar, Saudi Arabia, Venezuela, Canada, Norway, USSR/Russia, the UK, and the US); for two countries (China and Egypt) there is a partial unit root in the first regime; and for four countries (Iran, Iraq, Libya, and Mexico) there is a partial unit root in the second regime.

As the previous research indicates, the examination of the long memory (fractional integration) properties of OPEC oil production with the inclusion of structural breaks and outliers has not been explored in the literature.

3. Methodology

One characteristic of many economic and financial time series is its nonstationary nature. There exists a variety of models to describe such nonstationarity. Until the 1980s a standard approach was to impose a deterministic (linear or quadratic) function of time, thus assuming that the residuals from the regression model were stationary I(0). Later on, and especially after the seminal work of Nelson and Plosser (1982), there was a general agreement that the nonstationary component of most series was stochastic, and unit roots (or first differences, I(1)) were commonly adopted. However, the I(1) case is merely one particular model to describe such behaviour. In fact, the number of differences required to get $I(0)$ may not necessarily be an integer value but any point in the real line. In such a case, the process is said to be fractionally integrated or $I(d)$. The $I(d)$ models belong to a wider class of processes called long memory. We can define long memory in the time domain or in the frequency domain.

Let us consider a zero-mean covariance stationary process $\{ x_t,\ t = 0, \pm 1, ... \}$ with autocovariance function $\gamma_u = E(x_t x_{t+u})$. The time domain definition of long memory states that $\sum_{u=-\infty}^{\infty} |\gamma_u| = \infty$. Now, assuming that x_t has an absolutely continuous spectral distribution, so that it has spectral density function

$$f(\lambda) = \frac{1}{2\pi} \left(\gamma_0 + 2 \sum_{u=1}^{\infty} \gamma_u \cos(\lambda u) \right), \qquad (1)$$

the frequency domain definition of long memory states that the spectral density function is unbounded at some frequency in the interval $[0, \pi)$. Most of the empirical literature has concentrated on the case where the singularity or pole in the spectrum takes place at the 0-frequency. This is the standard case of $I(d)$ models of the form:

$$(1 - L)^d x_t = u_t, \qquad t = 0, \pm 1, ..., \qquad (2)$$

where L is the lag-operator ($L x_t = x_{t-1}$) and u_t is $I(0)$. Note that the polynomial $(1–L)^d$ in (2) can be expressed in terms of its Binomial expansion, such that, for all real d,

$$(1 - L)^d = \sum_{j=0}^{\infty} \psi_j L^j = \sum_{j=0}^{\infty} \binom{d}{j} (-1)^j L^j = 1 - d\, L + \frac{d(d-1)}{2} L^2 - \ ... \ ,$$

and thus

$$(1 - L)^d x_t = x_t - d\, x_{t-1} + \frac{d(d-1)}{2} x_{t-2} - \ ... \ .$$

In this context, d plays a crucial role since will be an indicator of the degree of dependence of the time series. Thus, the higher the value of d is, the higher the level of association will be between the observations. On the other hand, the above process also admits an infinite moving average representation such that

$$x_t = \sum_{k=0}^{\infty} a_k u_{t-k},$$

where

$$a_k = \frac{\Gamma(k+d)}{\Gamma(k+1)\Gamma(d)}.$$

Thus, the impulse responses are also clearly affected by the magnitude of d, and the higher the value of d is, the higher the responses will be. In this context, if d is smaller than 1, the series will be mean reverting, with shocks having temporary effects, and disappearing in the long run. On the other hand, if $d \geq 1$, shock will be permanent lasting forever unless strong policy measures are adopted. Processes with $d > 0$ in (2) display the property of "*long memory*", characterized because the spectral density function of the process is unbounded at the origin. However, fractional integration may also occur at some other frequencies away from 0, as in the case of seasonal/cyclical models.

In this study, we estimate d using a Whittle function in the frequency domain (Dahlhaus, 1989) along with a testing procedure developed by Robinson (1994) that permits us to test the null hypothesis H_o: $d = d_o$ in (1) for any real value d_o, where x_t can be the errors in a regression model of form:

$$y_t = \beta^T z_t + x_t, \qquad t = 1, 2, ..., \tag{3}$$

where y_t is the time series we observe; β is a (kx1) vector of unknown coefficients; and z_t is a set of deterministic terms that might include an intercept (i.e., $z_t = 1$), an intercept with a linear time trend ($z_t = (1, t)^T$), or any other type of deterministic processes like dummy variables to examine the potential presence of outliers/breaks. This method is briefly described in Appendix 1.

We also consider the possibility of structural breaks, which are endogenously determined by the model. For simplicity, we describe here the case of a single break and consider a model of the form:

$$y_t = \beta_1^T z_t + x_t; \quad (1-L)^{d_1} x_t = u_t, \quad t = 1,...,T_b, \tag{4}$$

and

$$y_t = \beta_2^T z_t + x_t; \quad (1-L)^{d_2} x_t = u_t, \quad t = T_b+1,...,T, \tag{5}$$

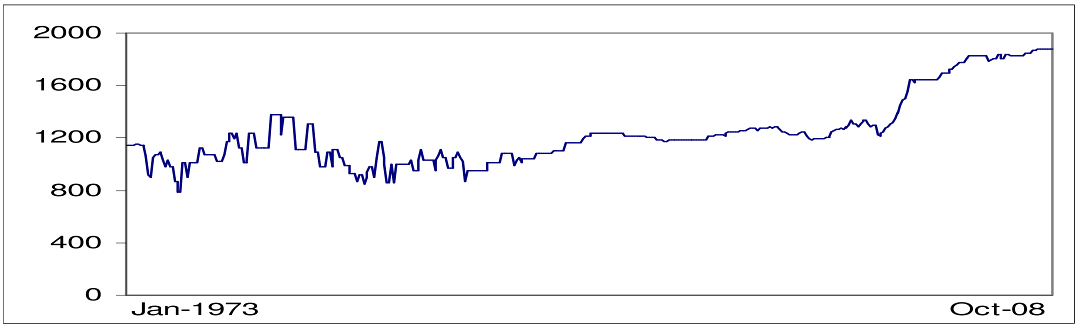
where the β's are the coefficients corresponding to the deterministic terms; d_1 and d_2 may be real values; u_t is I(0); and T_b is the time of a break that is supposed to be unknown. Note that given the difficulties in distinguishing between models with fractional orders of integration and those with broken deterministic trends (i.e., Diebold and Inoue, 2001; Granger and Hyung, 2004), we implement a recent procedure developed by Gil-Alana (2008, see Appendix 2) that is based on minimizing the residuals sum squares in the two subsamples and that can be easily extended to the case of two or more breaks.

4. Data and Results

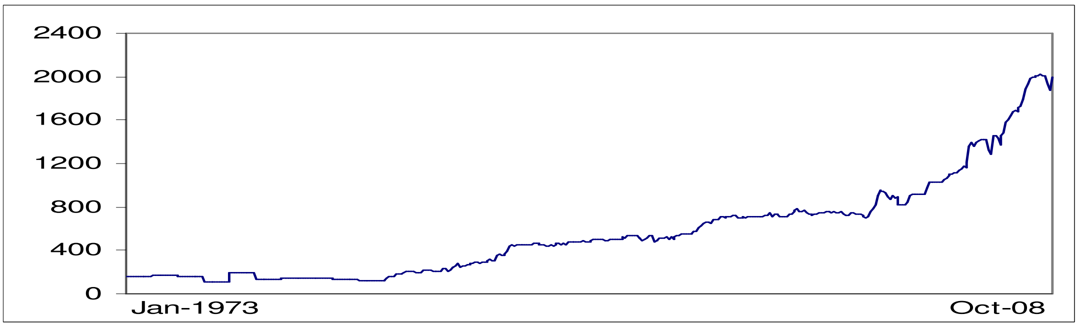
Monthly (seasonally adjusted) data of the oil production of OPEC countries were obtained from January 1973 to October 2008 from the *Energy Information Administration* web site. The total number of observations is 431 for each country: Algeria, Angola, Ecuador, Indonesia, Iran, Iraq, Kuwait, Libya, Nigeria, Qatar, Saudi Arabia, United Arab Emirates, and Venezuela. Figure 1 displays the time series plots for each OPEC country. We observe in all cases a strong persistent pattern that is changing across time, suggesting the adoption of fractional integration with and without breaks.

Figure 1: Time Series Plots

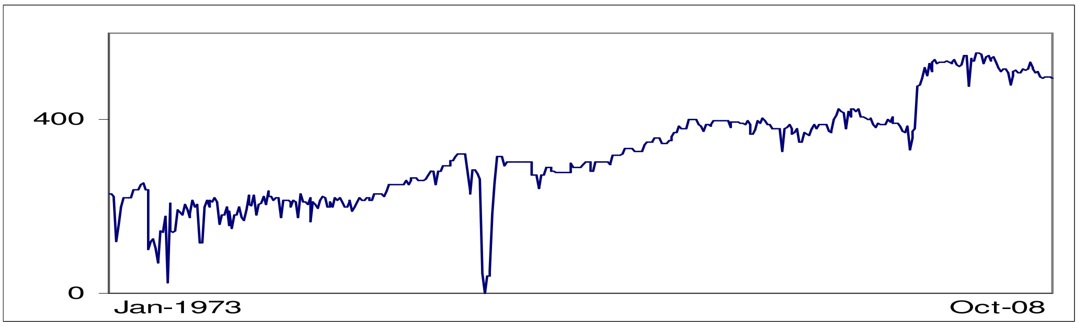
1)**Algeria**



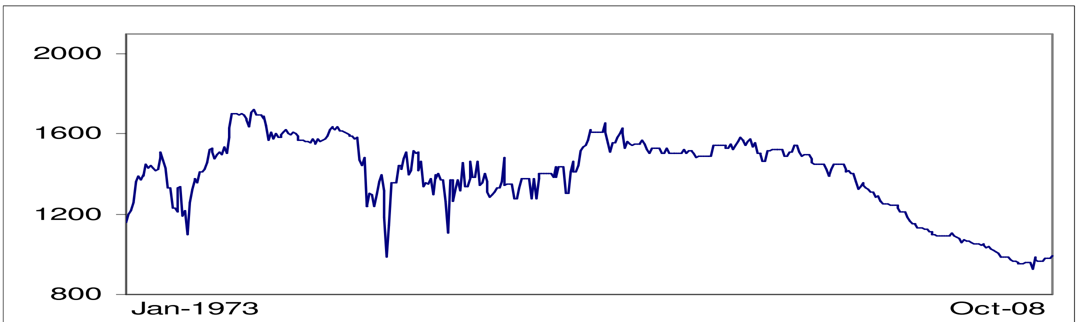
2) Angola



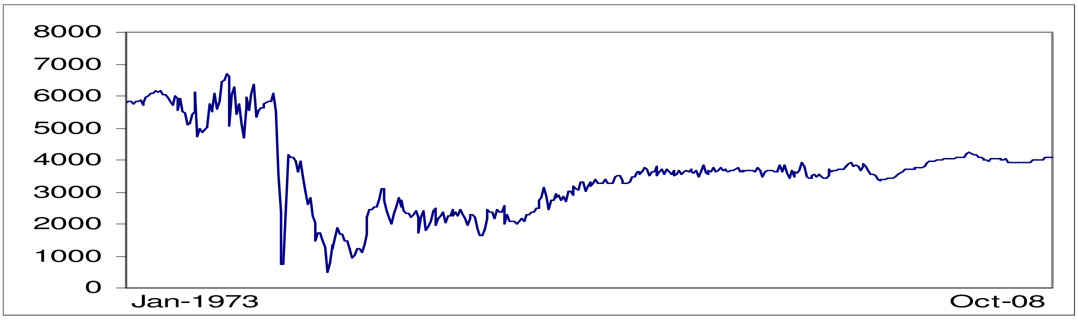
3) Ecuador



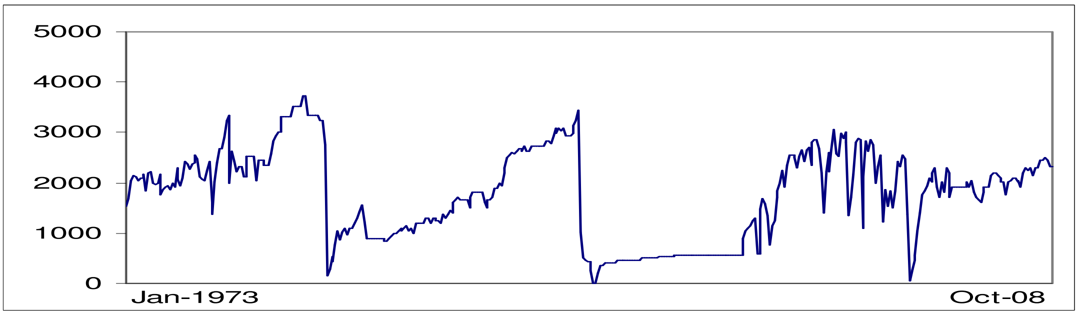
4) Indonesia



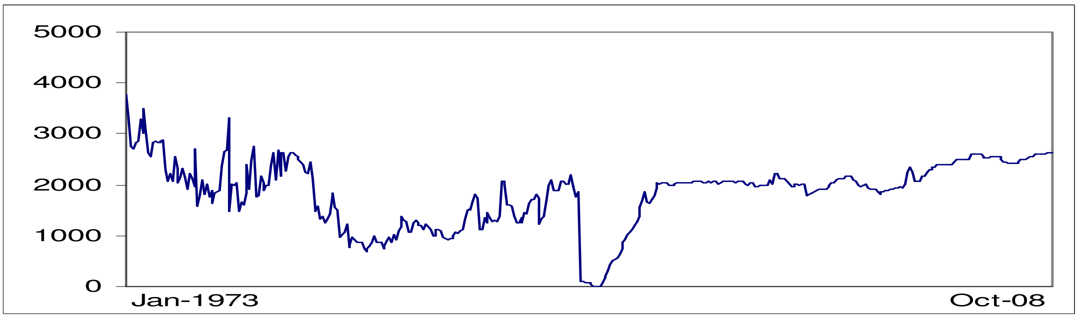
5) Iran



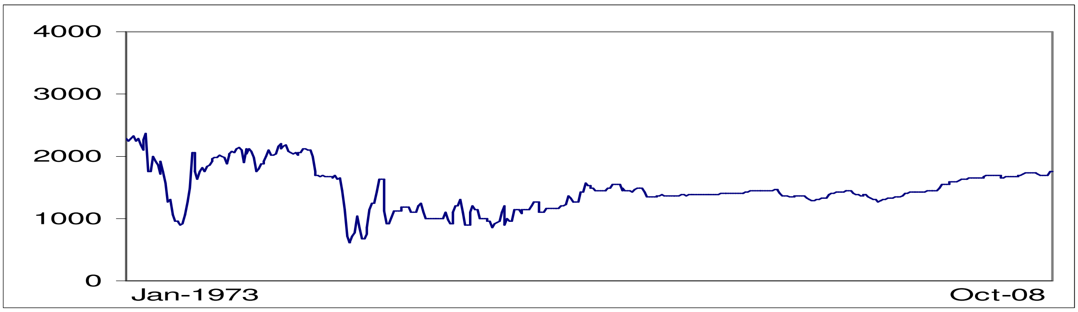
6) Iraq



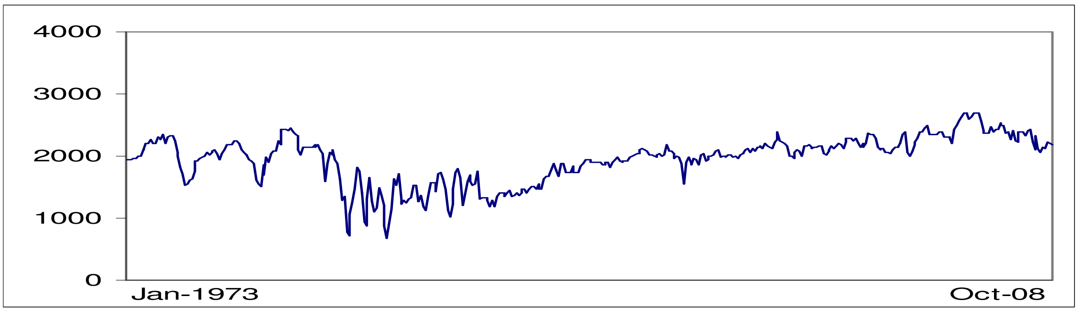
7) Kuwait



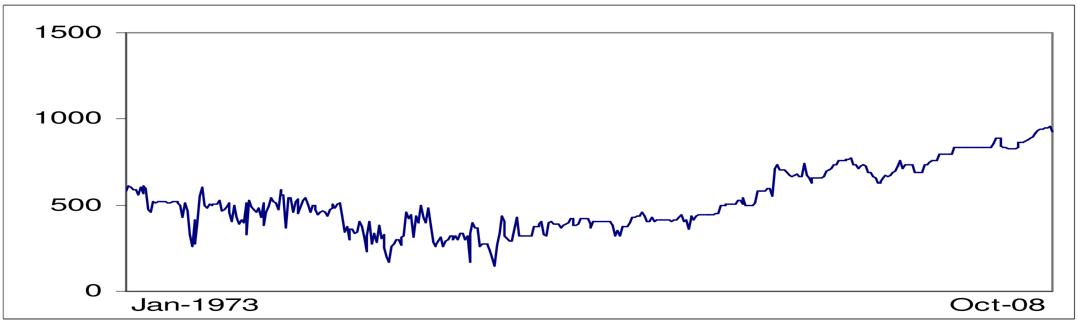
8) Libya



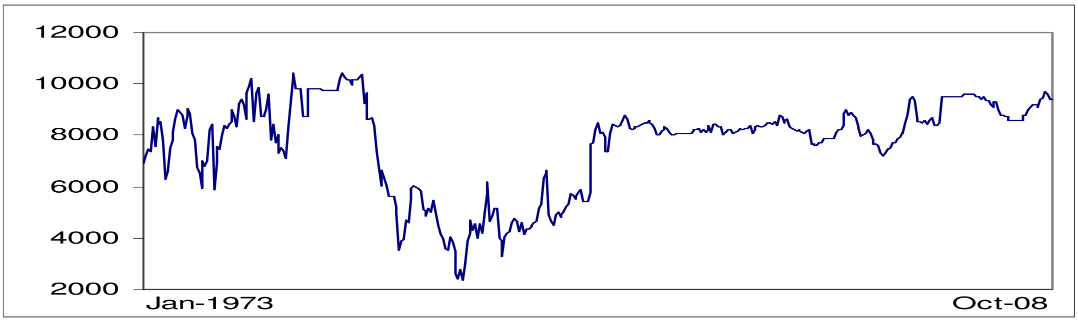
9) Nigeria



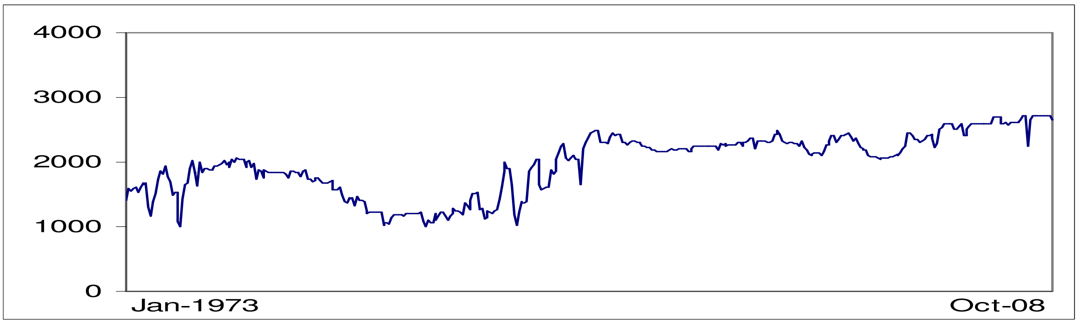
10) Qatar



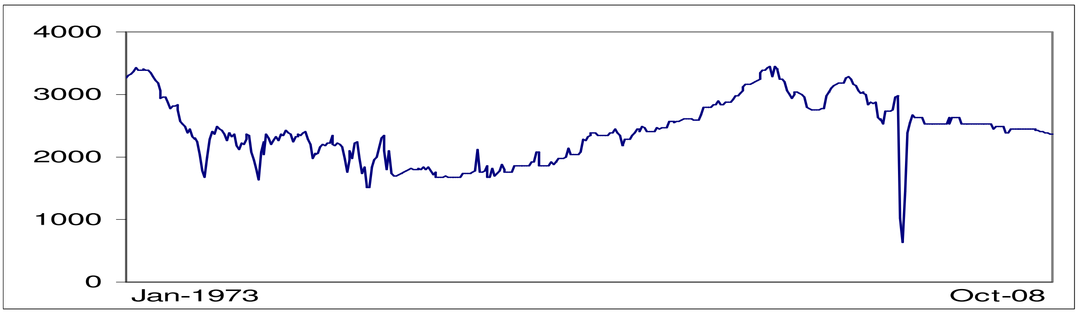
11) Saudi Arabia



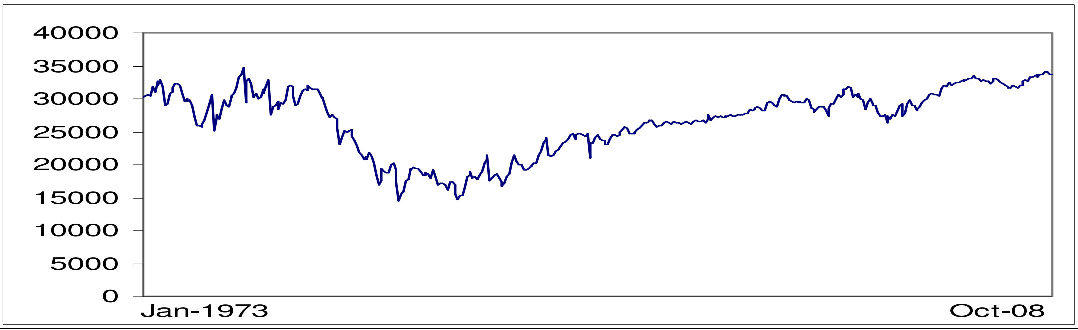
12) United Arab Emirates



13) Venezuela



14) Total



In order to take into account the main feature of the data (i.e., their degree of dependence across time), we first consider the following model,

$$y_t \;=\; \beta_1 \;+\; \beta_2 t \;+\; x_t; \qquad (1-L)^d\, x_t \;=\; u_t, \qquad (6)$$

where u_t is I(0), defined first as a white noise process, and then allowing some type of weak autocorrelation structure. The above model includes the standard cases examined in the literature. For example, if d = 0, we have a deterministic trend model with I(0) disturbances, while if d = 1, the classical unit root model. However, allowing d to be a real value, we can also examine the possibility of fractional integration. As earlier mentioned, the parameter d is an indicator of the degree of long range dependence, and the higher is the value of d, the higher is the level of association between the observations.

Table 2 displays the estimates of the fractional differencing parameter, d, in the model given by (6) assuming that the disturbances u_t are white noise. We display the estimates of d (in parenthesis within the brackets) along with the 95% confidence intervals using Robinson's (1994) tests, for the three standard cases: (1) no regressors (i.e., $\beta_1 = \beta_2 = 0$ a priori), (2) an intercept (i.e., β_1 unknown, and $\beta_2 = 0$ a priori), and (3) an intercept with a linear time trend (i.e., β_1 and β_2 unknown).

Table 2: 95% Confidence Bands and Estimates of d in a Model with White Noise u_t

	No regressors	An intercept	A linear time trend
Algeria	[0.870 (0.931)	[0.794 (0.850)	**[0.785 (0.845)***
Angola	[1.050 (1.108)	[1.036 (1.087)	**[1.038 (1.090)***
Ecuador	[0.706 (0.777)	[0.705 (0.767)	**[0.686 (0.758)***
Indonesia	[0.928 (0.980)	**[0.791 (0.842)***	[0.787 (0.840)
Iran	[0.915 (0.982)	**[0.879 (0.959)***	[0.880 (0.959)
Iraq	[0.836 (0.903)	**[0.810 (0.880)***	[0.810 (0.880)
Kuwait	[0.758 (0.811)	**[0.739 (0.791)***	[0.744 (0.793)
Libya	[0.937 (1.008)	**[0.988 (1.074)***	[0.988 (1.074)
Nigeria	[0.866 (0.937)	**[0.788 (0.876)***	[0.787 (0.876)
Qatar	[0.733 (0.782)	[0.657 (0.694)	**[0.641 (0.681)***
Saudi Arabia	[0.918 (0.975)	**[0.863 (0.924)***	[0.863 (0.924)
U.A.E.	[0.881 (0.949)	**[0.805 (0.878)***	[0.804 (0.878)
Venezuela	[0.893 (0.968)	**[0.830 (0.925)***	[0.843 (0.925)
Total	[0.916 (0.973)	**[0.838 (0.889)***	[0.837 (0.888)

In bold and with an asterisk the estimates of d where the deterministic terms are statistically significant.

We observe from Table 2 that all the estimates are above 0.5, implying long memory and a nonstationary behaviour, and the results seem to be robust across the different types of deterministic terms.[6] We present in bold type and with an asterisk the cases where the deterministic terms are statistically significant. We note that for Algeria, Angola, Ecuador, and Qatar, the time trend is significant. For the remaining cases, only an intercept appears significant in the regression models. If we focus now on the estimates of d, we observe that only for Angola, the value of d is found to be statistically significantly above 1. For another three countries, Iran, Libya, and Venezuela, the unit root null (i.e., d = 1) cannot statistically be rejected, while for the remaining nine countries (and for the total production series), the orders of integration are strictly smaller than 1, ranging from d = 0.681 (Qatar) to d = 0.924 (Saudi Arabia).

Table 3: 95% Confidence Bands and Estimates of d in a Model with Autocorrelated (Bloomfield) u_t

	No regressors	An intercept	A linear time trend
Algeria	[0.790 (0.891)	[0.720 (0.795)	**[0.694 (0.764)** *
Angola	[0.938 (0.992)	[0.977 (1.033)	**[0.977 (1.041)** *
Ecuador	[0.589 (0.722)	[0.640 (0.738)	**[0.568 (0.709)** *
Indonesia	[0.902 (0.987)	**[0.757 (0.836)** *	[0.752 (0.828)
Iran	[0.774 (0.867)	**[0.671 (0.748)** *	[0.682 (0.757)
Iraq	[0.689 (0.768)	**[0.631 (0.723)** *	[0.631 (0.723)
Kuwait	[0.761 (0.849)	**[0.790 (0.883)** *	[0.793 (0.891)
Libya	[0.792 (0.893)	**[0.692 (0.799)** *	[0.704 (0.800)
Nigeria	[0.689 (0.783)	**[0.547 (0.611)** *	[0.530 (0.607)
Qatar	[0.752 (0.837)	**[0.698 (0.753)** *	[0.676 (0.738)
Saudi Arabia	[0.872 (0.968)	**[0.798 (0.889)** *	[0.798 (0.889)
U.A.E.	[0.747 (0.828)	[0.649 (0.727)	**[0.632 (0.718)** *
Venezuela	[0.677 (0.778)	[0.532 (0.598)	**[0.548 (0.618)** *
Total	[0.867 (0.960)	**[0.800 (0.872)** *	[0.799 (0.871)

In bold and with an asterisk the estimates of d where the deterministic terms are statistically significant.

In Table 3, we report the results under the assumption that the disturbances (u_t in (6)) are weakly autocorrelated. However, instead of imposing a classical autoregressive model, whose parameters may be competing with d in describing the time dependence, we use a less conventional approach based on the exponential spectral model of

Bloomfield (1973). This is a non-parametric approach that produces autocorrelations decaying exponentially as in the AR(MA) case; however, the parameters are stationary for all values, and approximate fairly well ARMA structures with a large number of parameters.[7] The time trend coefficients are now significant in the cases of Algeria, Angola, Ecuador, UAE and Venezuela. The only estimated value of d which is above 1 is again Angola, though now the unit root cannot be rejected. This hypothesis cannot be rejected either in the cases of Kuwait and Saudi Arabia. For the rest of the countries, d is found to be strictly smaller than 1, thus implying mean reversion. Here, the lowest degree of integration occurs for Nigeria (d = 0.611), followed by Venezuela (d = 0.618) and Ecuador (d = 0.709).

Summarizing the results presented in these two tables, we can conclude that Angola is the country that presents the highest degree of nonstationarity. For some other countries (Iran, Libya, and Venezuela in case of uncorrelated errors, and Kuwait and Saudi Arabia with autocorrelated disturbances) the unit root model cannot be rejected. However, for the remaining countries (Algeria, Ecuador, Indonesia, Nigeria, Qatar, and UAE) the orders of integration are strictly smaller than 1, implying that shocks affecting the series are transitory, disappearing in the long run though taking a long time to disappear completely.

Table 4: Estimates of the Parameters in the Presence of Outliers with White Noise u_t

	With no outliers			With outliers			
	D	α	β	d	α	α^*	β (α^{**})
Ecuador	0.758 [0.686,	218.52 9	0.6964 (2.192)	0.736 [0.664,	216.726 (9.259)	- 65.280	0.6995 (2.514)
Indonesia	0.842 [0.791,	1180.6 94	---	0.837 [0.790,	1181.79 4	- 198.65	---
Iran	0.959 [0.879,	5806.5 28	---	0.885 [0.825,	5797.45 1	- 1833.4	- 540.698

Iraq	0.880 [0.810,	1581.034	---	0.877 [0.809,	1582.712	-256.52	-826.281
Kuwait	0.791 [0.739,	3528.102	---	0.788 [0.736,	3522.814	-115.82	---
Saudi Ar.	0.924 [0.863,	6928.869	---	0.923 [0.863,	7075.951	-595.47	---
Venezuela	0.925 [0.830,	3270.161	---	0.851 [0.778,	3260.510	-723.52	---

For Ecuador, the outlier takes place at April, 1987; for Indonesia at February, 1983; for Iran there are two outliers, one at January and February, 1979, and at October, 1980; for Iraq, the two outliers are at February-May, 1991, and at April, 2003; for Kuwait, at February-May, 1991; for Saudi Arabia, at August, 1985; and for Venezuela at January 2003.

Table 5: Estimates of Parameters in the Presence of Outliers with Autocorrelated u_t

	With no outliers			With outliers			
	d	α	β	d	α	α^*	β (α^{**})
Ecuador	0.709 [0.568,	214.237	0.7034 (2.897)	0.736 [0.595,	216.726 (9.259)	-65.280	0.6995 (2.514)
Indonesia	0.836 [0.757,	1182.009	---	0.885 [0.808,	1172.319	-193.36	---
Iran	0.748 [0.671,	5716.929	---	0.837 [0.770,	5782.898	-1899.2	-556.382
Iraq	0.723 [0.631,	1693.042	---	0.740 [0.652,	1678.649	-300.31	-892.617
Kuwait	0.883 [0.790,	3660.134	---	0.881 [0.781,	3657.805	-86.725	---
Saudi Ar.	0.889 [0.798,	6959.319	---	0.889 [0.798,	6959.327	-553.61	---
Venezuela	0.618 [0.548,	3127.741	-1.324 (-	0.765 [0.704,	3243.993	-801.19	-1.560 (-1.884)

For Ecuador, the outlier takes place at April, 1987; for Indonesia at February, 1983; for Iran there are two outliers, one at January and February, 1979, and at October, 1980; for Iraq, the two outliers are at February-May, 1991, and at April, 2003; for Kuwait, at February-May, 1991; for Saudi Arabia, at August, 1985; and for Venezuela at January 2003.

The results presented so far may be biased because of the presence of breaks and/or outliers (see Gil-Alana, 2003; 2005).[8] Tables 4 and 5 address the case of outliers in the data. We identify outliers in seven countries (Ecuador, Indonesia, Iran, Iraq, Kuwait, Saudi Arabia, and Venezuela) and use dummy variables to describe them. The results including these dummies in the regression model (2) and using white noise u_t are reported in Table 4 and using the autocorrelated model of Bloomfield (1973) reported in Table 5.[9]

The results in the left-hand sides of Tables 4 and 5 report the estimates of the model parameters under the assumption that there are no outliers. The results in the right-hand side refer to the estimates assuming the existence of one or two outliers depending on the series. Starting with the results based on white noise disturbances (in Table 4) we observe a reduction in the degree of integration in all the countries once the outliers are taken into account. Thus, for example, for Iran and Venezuela the unit root null hypothesis cannot be rejected if no outliers are considered, however, including them, this hypothesis is rejected in the two countries in favour of orders of integration smaller than 1. This is less clear if the disturbances are autocorrelated as shown in Table 5. Here, we note higher orders of integration with outliers for the cases of Ecuador, Indonesia, Iran, Iraq, and Venezuela; more or less the same estimates in case of Saudi Arabia, and a slight reduction only for Kuwait. In general, there are no substantial differences if outliers are taken into account. Mean reversion seems to take place in all the countries examined with the exceptions of Kuwait and Saudi Arabia where the unit root null cannot be rejected.

Table 6: Estimates of d in the Presence of Structural Breaks

<table>
<tr><th></th><th>N. Breaks</th><th>Bk.Dates</th><th colspan="4">Parameter estimates</th></tr>
<tr><td>Algeria</td><td>0</td><td>---</td><td colspan="4">$d = 0.764$; $\alpha = 1123.33$; $\beta = 1.624$;</td></tr>
<tr><td>Angola</td><td>0</td><td>---</td><td colspan="4">$d = 1.041$; $\alpha = 158.438$; $\beta = 4.316$;</td></tr>
<tr><td>Ecuador</td><td>1 + outlier</td><td>Sept-03</td><td colspan="2">d_1=0.81; α_1=225.9; α^*_1=-57.86</td><td colspan="2">$d_2 = 1.090$; $\alpha_2 = 475.610$</td></tr>
<tr><td>Indonesia</td><td>0</td><td>---</td><td colspan="4">$d = 0.842$; $\alpha = 1180.694$;</td></tr>
<tr><td>Iran</td><td>1</td><td>Nov-78</td><td colspan="2">$d_1 = 0.922$; $\alpha_1 = 5809.72$</td><td colspan="2">$d_2 = 1.023$; $\alpha_2 = 3543.88$</td></tr>
<tr><td>Iraq</td><td>2</td><td>Oct-80</td><td>$d_1 = 0.642$; $\alpha_1 = 1661.3$;</td><td colspan="2">$d_2 = 1.018$; $\alpha_2 = 109.48$;</td><td>$d_3 = 0.740$;</td></tr>
</table>

		Aug-90	$\beta_1 = -16.4$	$\beta_2 = 28.70$	$\alpha_3 = 866.9$
Kuwait	0	---	$d = 0.791$; $\alpha = 3528.102$;		
Libya	1	Jan-83	$d_1 = 1.141$; $\alpha_1 = 2300.59$	$d_2 = 0.978$; $\alpha_2 = 1125.79$	
Nigeria	1	Jul-81	$d_1 = 1.055$; $\alpha_1 = 1934.64$	$d_2 = 0.890$; $\alpha_2 = 809.569$	
Qatar	1	May-86	$d_1 = 0.69$; $\alpha_1 = 586.26$; $\beta_1 = -1.98$	$d_1 = 0.95$; $\alpha_2 = -22.02$; $\beta_2 = -1.75$	
Saudi Arabic	2	Mar-82 Sep-90	$d_1 = 0.708$; $\alpha_1 = 7249.07$	$d_2 = 0.941$; $\alpha_2 = 7259.91$	$d_3 = 1.034$; $\alpha_3 = 7651.05$
U.A.E.	1	Sep-90	$d_1 = 0.993$; $\alpha_1 = 1417.78$	$d_2 = 0.998$; $\alpha_2 = 2198.82$	
Venezuela	1 + outlier	Sep-86	$d_1 = 0.994$; $\alpha_1 = 32715992$	$d_2 = 0.97$; $\alpha_2 = 1760.7$; $\alpha^*_2 = -624.3$	
Total	1	Sep-86	$d_1 = 0.955$; $\alpha_1 = 30185.31$	$d_2 = 1.041$; $\alpha_2 = 17556.21$	

Finally, in the case of structural breaks, we implement the method of Gil-Alana (2008) described in the previous section and in Appendix 2. We employ one and two breaks and choose the appropriate number of breaks using likelihood information criteria (see Appendix 2). The results displayed in Table 6 show no breaks for the cases of Algeria, Angola, Indonesia, and Kuwait. One single break for Iran, Libya, Nigeria, Qatar, UAE, Ecuador, and Venezuela, in the latter two countries including also outliers, and two breaks for Iraq and Saudi Arabia.

Notice in Table 6 that the break dates substantially change from one series to another. Thus, there is an early break in Iran (November 1978) due to the exile of the Shah of Iran in January 1979 and the associated revolution and several other breaks in the early 1980s. For Iraq (October 1980 and August 1990), the October 1980 break is attributed to the Iranian revolution and the August 1990 break to Iraq's occupation of Kuwait (i.e. first Gulf War). In the case of Nigeria (July 1981) the observed break is attributed to the long-standing border dispute between Nigeria and Cameroon which generates the first of several crises from May 15, 1981 to July 24, 1981. In Saudi Arabia (March 1982 and September 1990), the March 1982 break is affiliated with the curtailment of production and an atmosphere of recrimination between OPEC producers over cheating on production quotas which marked the period March 1982 to March 1983 (see Ramazani, 1988) while the September 1990 break is attributed to the first Gulf War. In the case of Libya (January 1983), three breaks occurred in 1986 related to U.S. military attacks in response to Libya's support for international terrorism. For Qatar, (May 1986) the break pertains to the continuing dispute with Bahrain concerning the nearby Hawar Islands, which resulted in Qatar troops briefly occupying a coral reef which was being reclaimed from the sea. For Venezuela (September 1986), the observed break is due to the halt in production in light of low oil prices. The UAE (September 1990) also experienced a break due to the first Gulf War. Finally, the latest break occurs in Ecuador (September 2003) related to tensions between the indigenous people and oil companies located in Ecuador.

If we focus now on the orders of integration, we first notice that for the countries where there are no breaks (Algeria, Angola, Indonesia, and Kuwait), we obtain mean reversion in three of them (Algeria d = 0.764; Indonesia d = 0.842; and Kuwait, d = 0.791), however, for Angola with d = 1.041, the unit root null hypothesis cannot be rejected. For those countries with a single break, we observe a significant increase in the degree of integration in Ecuador, Iran, UAE, Venezuela, and for the total production series; in all these countries mean reversion is observed in the first subsample and the unit root cannot be rejected after the break. However, for Libya and Nigeria there is a decrease in the value of d though the unit root cannot be rejected in any of the two subsamples. Finally, there are two countries where two breaks are observed: Iraq and

Saudi Arabia. In the former country, the orders of integration are $d_1 = 0.642$, $d_2 = 1.018$ and $d_3 = 0.740$, with the presence of a unit root rejected in favour of mean reversion in the first and third subsamples, and failure to reject a unit root in the second subsample corresponding to the decade of the 1980s. In case of Saudi Arabia, the order of integration increases across time ($d_1 = 0.708$, $d_2 = 0.941$, and $d_3 = 1.034$) with the failure to reject a unit root during the third subsample.

5. Discussion

We have presented in Section 4 results based on fractional integration using three different approaches: a) a model with no outliers and no breaks, b) a model with outliers, and c) a model with outliers and breaks.

The results can be summarized as follows: if no breaks and no outliers are taken into account, most of the fractional differencing parameters are in the interval (0.5, 1) implying long memory and mean reverting behaviour, with shocks disappearing in the long run. The exceptions are Angola, and also Kuwait and Saudi Arabia (with white noise errors) and Iran, Libya and Venezuela (with autocorrelated errors) where the unit root cannot be rejected, and thus suggesting that shocks have permanent effects on these countries. Thus, according to this specification, in the event of an exogenous shock, stronger policy measures must be adopted in countries like Angola, Kuwait, Saudi Arabia, Iran, Libya and Venezuela than in others like Algeria, Ecuador, Indonesia, Iraq, Nigeria, Qatar and UAE to recover the series to its original trend. Allowing for outliers, they are found to be statistically significant in the cases of Ecuador, Indonesia, Iran, Iraq, Kuwait, Saudi Arabia and Venezuela, and the conclusions remain almost the same, with values close to but smaller than 1 in most cases. Evidence of unit roots is only obtained for Kuwait and Saudi Arabia if the disturbances are autocorrelated. Finally, we permit breaks and/or outliers and employed the methodology proposed by Gil-Alana (2008). The results here are conclusive: four countries (Algeria, Angola, Indonesia and Kuwait) do not present breaks, and evidence of mean reversion (i.e. $d < 1$) is obtained in three of them. Only Angola displays lack of mean reversion. Another group of six countries presents a single break: Ecuador, Iran, UAE, Venezuela, Libya and Nigeria, and evidence of mean reversion is only observed in the first four countries during the period previous to the breaks. Finally, two countries (Iraq and Saudi Arabia) present two breaks and evidence of permanent shocks (i.e. unit roots) are observed during the second subsample in the former country and in the last subsample in the case of Saudi Arabia.

6. Concluding Remarks

Unlike previous studies on oil production employing traditional unit root integrated models or even threshold unit root tests, this study adopts a fractional integration model adopted by Caporale and Gil-Alana (2007; 2008) which incorporates breaks and outliers in the analysis. Specifically, we present different specifications based on fractional integration, first with no breaks, and then allowing outliers and breaks to describe time series dependence and other implicit dynamics of oil production in OPEC countries. The results indicate that the standard methods employed in the literature, based on stationary I(0) or non-stationary I(1) models are clearly rejected in favor of fractional degrees of integration. Evidence of long memory ($d > 0$) is obtained in all cases, with orders of integration ranging from 0.642 (Iraq during the first subsample, January 1973 – October 1980) to 1.141 (Libya, first subsample: January 1973 – January 1981).

However, the results substantially vary from one country to another. Thus, for Algeria, Indonesia, and Kuwait, we do not observe breaks and mean reversion is obtained in the three countries with their orders of integration strictly below 1, which indicates that shocks are transitory and mean reverting, disappearing in the long run. Mean reversion is also observed in some countries in which a single break is required; for example, Iran, Qatar, UAE, Ecuador, and Venezuela during the first subsamples, and in the latter two countries outliers seem to be present as well. Finally, we observe two countries with two structural breaks, Iraq and Saudi Arabia. In the former country, mean reversion occurs during the sample period except for the period of the 1980s, and in Saudi Arabia during the period before the second break in September 1990. These results confirm the high degree of persistence in each series.

Thus, the results indicate that shocks affecting the structure of OPEC oil production (based on the estimates of d in all tables), will have persistent effects in the long run for all countries, and in some cases the effects are expected to be permanent. As a consequence, disruptions in oil production and supply will have a persistent impact on economic activity as such shocks will be transmitted to other sectors of the economy. Therefore, it is crucial for policy makers to distinguish the nature of the shock (i.e. transitory or persistent) since the policy actions may differ as to the type of shock. In case of values of d equal to or above 1, stabilization policies in restoring production levels to equilibrium levels will be required; otherwise, the implications for oil production and supply will persist forever. On the other hand, for countries with values of d below 1, shocks will disappear in the long run as production will return to equilibrium levels over time without the need for stabilization efforts.

In summary, it is clear that taking first differences in the oil production of OPEC countries under the assumption of a unit root, may lead in some cases to series that are over-differenced, and subsequently such a procedure may result in inappropriate policy actions. Second, persistence behavior is another characteristic of these data although for some countries the adjustment process takes a long time to disappear in which case an active oil policy stance is required to restore oil production levels. Third, outliers do not alter the main conclusions of this study though in two countries (Ecuador and Venezuela) outliers should be considered even in the context of structural breaks. Fourth, the breaks in oil production are specific to each country or common to OPEC policy, signifying that there are specific events that affect each country's oil production and common elements to many OPEC countries.

Appendix 1: Robinson's (1994) Parametric Approach

The LM test of Robinson (1994) for testing H_o: d = d_o in (1) and (2) is

$$\hat{r} \;=\; \frac{T^{1/2}}{\hat{\sigma}^2}\,\hat{A}^{-1/2}\,\hat{a}\,,$$

where T is the sample size and:

$$\hat{a} \;=\; \frac{-2\pi}{T}\sum_{j=1}^{T-1}\psi(\lambda_j)\,g(\lambda_j;\hat{\tau})^{-1}\,I(\lambda_j); \qquad \hat{\sigma}^2 \;=\; \sigma^2(\hat{\tau}) \;=\; \frac{2\pi}{T}\sum_{j=1}^{T-1}g(\lambda_j;\hat{\tau})^{-1}\,I(\lambda_j);$$

$$\hat{A} \;=\; \frac{2}{T}\left(\sum_{j=1}^{T-1}\psi(\lambda_j)^2 \;-\; \sum_{j=1}^{T-1}\psi(\lambda_j)\,\hat{\varepsilon}(\lambda_j)'\times\left(\sum_{j=1}^{T-1}\hat{\varepsilon}(\lambda_j)\,\hat{\varepsilon}(\lambda_j)'\right)^{-1}\times\sum_{j=1}^{T-1}\hat{\varepsilon}(\lambda_j)\psi(\lambda_j)\right)$$

$$\psi(\lambda_j) = \;\log\left|2\sin\frac{\lambda_j}{2}\right|; \qquad \hat{\varepsilon}(\lambda_j) \;=\; \frac{\partial}{\partial\tau}\log g(\lambda_j;\hat{\tau}); \qquad \lambda_j \;=\; \frac{2\pi\; j}{T}; \qquad \hat{\tau} \;=\; \arg\min\sigma^2(\tau).$$

$\hat{a}$ and $\hat{A}$ in the above expressions are obtained through the first and second derivatives of the log-likelihood function with respect to d (see Robinson, 1994, page 1422, for further details). $I(\lambda_j)$ is the periodogram of u_t evaluated under the null, i.e.:

$$\hat{u}_t \;=\; (1 \;-\; L)^{d_o}\;y_t \;-\; \hat{\beta}'\,w_t;$$

$$\hat{\beta} \;=\; \left(\sum_{t=1}^{T}w_t\,w_t'\right)^{-1}\sum_{t=1}^{T}w_t\,(1 \;-\; L)^{d_o}\;y_t; \qquad w_t \;=\; (1 \;-\; L)^{d_o}\;z_t,$$

$z_t = (1, t)^T$, and g is a known function related to the spectral density function of u_t:

$$f(\lambda;\sigma^2;\tau) \;=\; \frac{\sigma^2}{2\pi}\,g(\lambda;\tau), \qquad -\pi \;<\; \lambda \;\leq\; \pi.$$

Appendix 2: Gil-Alana's (2008) Method for Fractional Integration with Breaks

The model presented in (4) and (5) can also be written as:

$$(1-L)^{d_1} y_t = \beta_1 \tilde{z}_t(d_1) + u_t, \quad t = 1, ..., T_b,$$

$$(1-L)^{d_2} y_t = \beta_2 \tilde{z}_t(d_2) + u_t, \quad t = T_b+1, ..., T,$$

where $\tilde{z}_t(d_i) = (1-L)^{d_i} z_t$, i = 1, 2. The procedure is based on the least square principle. First we choose a grid for the values of the fractionally differencing parameters d_1 and d_2, for example, d_{io} = 0, 0.01, 0.02, …, 1, i = 1, 2. Then, for a given partition $\{T_b\}$ and given initial d_1, d_2-values, $(d_{1o}^{(1)}, d_{2o}^{(1)})$, we estimate the α's and the β's by minimizing the sum of squared residuals,

$$\min_{w.r.t.\{\alpha_1,\alpha_2,\beta_1,\beta_2\}} \sum_{t=1}^{T_b} \left[(1-L)^{d_{1o}^{(1)}} y_t - \beta_1 \tilde{z}_t(d_{1o}^{(1)})\right]^2 + \sum_{t=T_b+1}^{T} \left[(1-L)^{d_{2o}^{(1)}} y_t - \beta_2 \tilde{z}_t(d_{2o}^{(1)})\right]^2.$$

Let $\hat{\beta}(T_b; d_{1o}^{(1)}, d_{2o}^{(1)})$ denote the resulting estimates for partition $\{T_b\}$ and initial values $d_{1o}^{(1)}$ and $d_{2o}^{(1)}$. Substituting these estimated values on the objective function, we have $RSS(T_b; d_{1o}^{(1)}, d_{2o}^{(1)})$, and minimizing this expression across all values of d_{1o} and d_{2o} in the grid we obtain $RSS(T_b) = \arg\min_{\{i,j\}} RSS(T_b; d_{1o}^{(i)}, d_{2o}^{(j)})$. Next, the estimated break date, $\hat{T}_k$, is such that $\hat{T}_k = \arg\min_{i=1,...,m} RSS(T_i)$, where the minimization is taken over all partitions $T_1, T_2, …, T_m$, such that $T_i - T_{i-1} \geq |\varepsilon T|$. Then, the regression parameter estimates are the associated least-squares estimates of the estimated k-partition, i.e., $\hat{\beta}_i = \hat{\beta}_i(\{\hat{T}_k\})$, and their corresponding differencing parameters, $\hat{d}_i = \hat{d}_i(\{\hat{T}_k\})$, for i = 1 and 2.

The model can be extended to the case of multiple breaks. Thus, we can consider the model,

$$y_t = \alpha_j + \beta_j t + x_t; \; (1 - L)^{d_j} x_t = u_t, \quad t = T_{j-1}+1,...,T_j,$$

for j = 1, …, m+1, $T_0 = 0$ and $T_{m+1} = T$. Then, the parameter m is the number of changes. The break dates (T_1, …, T_m) are explicitly treated as unknown and for i = 1, …, m, we have $\lambda_i = T_i/T$, with $\lambda_1 < \ldots < \lambda_m < 1$. Following the same process as in the previous case, for each j-partition, $\{T_1, \ldots T_j\}$, denoted $\{T_j\}$, the associated least-squares estimates of α_j, β_j and the d_j are obtained by minimizing the sum of squared residuals in the d_i-differenced models, i.e.,

$$\sum_{j=1}^{m+1} \sum_{t=T_{j-1}+1}^{T_j} (1 - L)^{d_i} (y_t - \alpha_i - \beta_i t)^2,$$

where $\hat{\alpha}_i(T_j)$, $\hat{\beta}_i(T_j)$ and $\hat{d}(T_j)$ denote the resulting estimates. Substituting them in the new objective function and denoting the sum of squared residuals as $RSS_T(T_1, \ldots, T_m)$, the estimated break dates $(\hat{T}_1, \hat{T}_2,...,\hat{T}_m)$ are obtained by $\min_{(T_1,T_2,...,T_m)} RSS_T(T_1, ..., T_m)$ where the minimization is again obtained over all partition (T_1, …, T_m).

The above procedure requires the a priori determination of the number of breaks in the time series. Following standard procedures to select the number of breaks in the context of I(0) processes, Schwarz (1978) proposed the criterion:

$$SIC(m) = \ln\left[RSS_T(\hat{T}_1,...,\hat{T}_m)/(T-m)\right] + 2p^* \ln(T)/T,$$

where p^* is the number of unknown parameters. Yao (1988) used the Bayesian criterion:

$$BIC(m) = \ln\left[RSS_T(\hat{T}_1,...,\hat{T}_m)/T\right] + p^* \ln(T)/T.$$

Finally, Yao and Av (1989) proposed a third criterion based on

$$YIC(m) = \ln\left[RSS_T(\hat{T}_1,...,\hat{T}_m)/T\right] + mC_T/T,$$

where C_T is any sequence satisfying $C_T T^{-2d/k} \rightarrow \infty$ as $T \rightarrow \infty$ for some positive integer k.

The estimated number of break dates, $\hat{m}$, is then obtained by minimizing the above-mentioned information criteria given M a fixed upper bound for m.

Endnotes

1. Five founding members are Iran, Iraq, Kuwait, Saudi Arabia and Venezuela with nine other members joining later: Qatar (1961); Indonesia (1962) -- suspended its membership in January 2009; Libya (1962); United Arab Emirates (1967); Algeria (1969); Nigeria (1971); Ecuador (1973) -- suspended its membership from December 1992 to October 2007; Angola (2007); and Gabon (1975–1994), see Kaufmann et al (2008).

2. Smith (2009) provides an excellent discussion of the world oil market with respect to production decisions and its effect on price. Kaufman et al. (2008) on OPEC oil production.

3. See, for example, Lean and Smyth (2009) for the relevance of testing for unit roots.

4. There is also an enormous literature on the causal relationship between energy consumption and economic growth in which preliminary tests for unit roots are undertaken in the estimation of error correction models to infer Granger-causal relationships. See Payne (2009, 2010) and Ozturk (2009) for surveys of this literature.

5. In the context of the present work, it should be noted that the ADF test (and also other unit-root testing procedures such as Phillips and Perron (1988) and Kwiatkowski et al (1992)) have very low power if the alternatives are of a fractional form (see Diebold and Rudebusch, 1991; Hassler and Wolters, 1994; Lee and Schmidt, 1996).

6. In the I(d) context, a series is covariance stationary if $d < 0.5$. If $d \geq 0.5$, the series is no longer covariance stationary but still mean reverting if $d < 1$.

7. See Gil-Alana (2004) for a paper dealing with fractional integration in the context of Bloomfield disturbances.

8. Some authors claim that fractional integration may be a spurious phenomenon caused by the no-consideration of breaks in the data. (See, e.g., Smith, 2005).

9. We use dummies of form $D_t = 1\ I(t = T^*)$, where I is the indicator function and T^* is the time period for the outliers. Other dummies produced insignificant results.

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