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Energy Consumption in the GCC Countries:  
Evidence On Persistence

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**ENERGY CONSUMPTION IN THE GCC COUNTRIES:  
EVIDENCE ON PERSISTENCE**

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**Abstract**

This paper examines the statistical properties of energy consumption in the GCC countries applying fractional integration methods to annual data from 1980 to 2014. The results indicate that both the raw and the logged series exhibit a (statistically significant) linear time trend in the case of Bahrain, Oman and Qatar, and the raw series only in the case of Saudi Arabia. Mean reversion (i.e., statistical evidence of  $d < 1$ ) is found in the case of Bahrain for both the raw and logged data, and in Qatar for the logged series. In the remaining cases, the  $I(1)$  hypothesis cannot be rejected except for the logged data in Saudi Arabia, since  $d$  is found to be statistically higher than 1 in that country. The implication of these findings is that in the case of Bahrain and Qatar exogenous shocks to energy consumption have transitory effects, which disappear in the long run without the need for policy action, whilst the permanent nature of the effects of shocks elsewhere means that appropriate policies have to be designed to restore equilibrium.

**JEL Classification:** C00; C22; E21; Q40

**Keywords:** Fractional integration; energy consumption; GCC countries

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## 1. Introduction

In the last two decades numerous studies have analysed the causal linkages between energy consumption and economic growth as well as other macroeconomic variables; however, many of them have not paid proper attention to the stochastic properties of the energy variables.<sup>1</sup> Narayan and Smith (2007) have stressed the key importance of testing for the possible presence of unit roots in order to design suitable energy policies based on the appropriate knowledge about the temporary or permanent nature of the effects of exogenous shocks. In particular, non-stationarity in energy consumption can lead to path dependency or hysteresis (see, e.g., Agnolucci et al., 2004; Narayan and Smyth, 2007; Smyth, 2013). Narayan et al. (2010) found that shocks have a temporary effect on energy consumption in Australia and therefore long-run stabilisation policies are not required. However, non-stationarities might spread through the economy as a result of economic relationships involving variables with a stochastic trend (see Hendry and Juselius, 2000; Narayan and Smyth, 2007; Mishra et al., 2009; Hasanov and Telatar, 2011).

Some more recent studies have addressed these issues more carefully. Narayan and Smith (2007), Narayan et al. (2010), Aslan and Kum (2011), Hasanov and Telatar (2011) among others carried out univariate unit root tests, and also tested for cointegration and Granger causality between energy consumption and economic growth. Further, Apergis and Payne (2010), Narayan et al. (2010), Ozcan (2013), Kula (2014), Lean and Smith (2014), Shahbaz and Khan (2014) and others used Lagrange Multiplier (LM) unit root tests with one or two structural breaks that provided evidence of stationarity in most cases. Narayan et al. (2008), Mishra et al. (2009), Apergis et al.

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<sup>1</sup> See Bhattacharya et al. (2016), Osman et al. (2016), Kim (2015), Al-Mulali et al. (2014), Cowan et al. (2014), Nayan et al. (2013), Shahbaz et al. (2013), Hamit-Hagggar (2012), Belke et al. (2011), Wang et al. (2011), among others.

(2010), Kum (2012), Bolat et al. (2013) extended this approach to panels and reached similar conclusions. Finally, Ozcan (2013) and Yilanci and Tunali (2014) applied Fourier ADF unit root tests, which are more powerful of standard unit root tests, and also reported that energy variables are stationary.

Determining the integration properties of energy consumption in an economy is crucial for forecasting (Chen and Lee, 2007) and policy design purposes. Following Barros et al. (2013), Apergis and Tsoumas (2012), Gil-Alana et al. (2010) and Monge et al. (2017) among others, the present paper applies fractional integration methods to examine the time series properties of energy consumption per capita in the GCC countries (namely Bahrain, Kuwait, Qatar, Oman, the Kingdom of Saudi Arabia and the United Arab Emirates) for which no previous evidence is available, using annual data for the period 1980-2014 taken from the World Bank's World Development Indicators Database. These countries are currently classified as emerging countries engaged in the transformation of their economies through the creation of new industry segments and diversification (Shkvarya et al., 2017); their average annual GDP per capita (\$69,166) is substantially higher than the world average (\$16,961). GCC members have benefited from being oil and gas producers, although they have still been affected by the worldwide slowdown caused by the 2007-8 global financial crisis. Interestingly, energy consumption and economic growth appear to be strongly linked to all sectors in the GCC, in contrast to the OECD group where energy and GDP have decoupled, which suggests that these countries should focus more on maximising the value that can be obtained from energy consumption (see Howarth et al., 2017).

Our analysis is based on the concepts of long-run dependence and long memory and allows for fractional degrees of integration. Therefore the adopted specification is

more general and has a more flexible dynamic structure than the standard AutoRegressive (Integrated) Moving Average (AR(I)MA) models only allowing for integers as the order of integration  $d$ . Note that if  $d < 1$  the effects of shocks will be transitory, whilst  $d = 1$  or  $d > 1$  imply permanent effects.

The remainder of the paper is structured as follows. Section 2 outlines the methodology. Section 3 describes the data and discusses the main empirical results. Section 4 offers some concluding remarks.

## 2. Methodology

The analysis below is based on the concept of long memory. In the time domain, a process  $\{x_t, t = 0, \pm 1, \dots\}$  with an autocovariance (or pseudo-autocovariance) function  $\gamma_u = \text{Cov}(x_t, x_{t+u})$  is said to exhibit long memory if the sum of its autocovariances is  $\gamma_u$ , i.e.,

$$\sum_{u=-\infty}^{\infty} |\gamma_u| = \infty. \quad (1)$$

In the frequency domain, consider the same process  $\{x_t, t = 0, \pm 1, \dots\}$  and assume that it has a spectral density function (or a pseudo spectral density function), which is the Fourier transform of the autocovariances, i.e.,

$$f(\lambda) = \frac{1}{2\pi} \sum_{u=-\infty}^{\infty} \gamma_u e^{i\lambda u} \quad \lambda \in [-\pi, \pi),$$

Then the process  $x_t$  is said to exhibit long memory if its spectral density function  $f(\lambda)$  has at least one singularity or pole in the spectrum, i.e.,

$$f(\lambda) \rightarrow \infty, \quad \text{as } \lambda \rightarrow \lambda^*, \quad \text{for } \lambda^* \in [-\pi, \pi). \quad (3)$$

In the time domain short memory is the property of a covariance stationary process with a finite sum of all its autocovariances, i.e.,

$$\sum_{u=-\infty}^{\infty} |\gamma_u| < \infty. \quad (4)$$

whilst in the frequency domain it is a feature of a process with a spectral density function that is positive and finite at all its frequencies on the spectrum, i.e.,

$$0 < f(\lambda) < \infty, \quad \lambda \in [-\pi, \pi), \quad (5)$$

The category of short-memory or I(0) processes includes white noise but also stationary and invertible ARMA processes.

Long memory is a property of unit-root or I(1) processes that become I(0) or stationary by taking first differences. More specifically, a process  $\{x_t, t = 0, \pm 1, \dots\}$  is said to be I(1) if it can be represented as

$$(1 - L)x_t = u_t, \quad t = 0, \pm 1, \dots, \quad (6)$$

where  $x_t = 0$  for  $t \leq 0$ ,  $L$  stands for the lag operator ( $Lx_t = x_{t-1}$ ) and  $u_t$  is a I(0) process defined as above, which can be a white noise or an ARMA process. Note that if  $u_t$  is ARMA(p,q), then  $x_t$  is said to be an ARIMA (P, 1, q) process. The differencing parameter for making a series I(0) is not necessarily an integer (e.g., 1 as in (6) above) but can be any real number. In other words, one can consider more general models of the form:

$$(1 - L)^d x_t = u_t, \quad t = 0, \pm 1, \dots, \quad (7)$$

where  $d$  can be a fraction between 0 or 1, or even exceed 1. In fact, one can use a Binomial expansion such that, for all real value  $d$ ,

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

and therefore the left-hand side of (7) can be expressed as

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

This type of processes were introduced by Granger (1980, 1981), Granger and Joyeux (1980) and Hosking (1981) after noticing that many series appeared to be overdifferentenced after differencing them to achieve stationarity. They were made popular in the nineties by Baillie (1996), Gil-Alana and Robinson (1997) and Silverberg and Verspagen (1999), and since then have been widely applied to analyse time series data in various sectors including the energy one (see, e.g., Gil-Alana et al., 2010).

In this context, the parameter  $d$  plays a very important role as a measure of the degree of persistence. In particular, if  $d$  belongs to the interval  $(0, 0.5)$   $x_t$  in (7) is covariance stationary, whereas if  $d \geq 0.5$  the process is non-stationary. Also, values of  $d$  below 1 imply mean reversion, i.e., the effects of shocks are transitory and disappear in the long run, whilst if  $d \geq 1$  they are permanent. Finally, note that if  $u_t$  in (7) is an ARMA( $p, q$ ) process, then  $x_t$  is a fractionally integrated ARMA or ARFIMA( $p, d, q$ ) process.

We estimate the fractional differencing parameter using the Whittle function in the frequency domain (Dahlhaus, 1989) applying a parametric testing procedure proposed by Robinson (1994) that is valid even in the presence of non-stationarity. This method allows to test for any real value  $d$  in the model given by (6), where  $x_t$  can be the errors of a regression model including deterministic terms such as an intercept and/or a linear trend. Moreover, the limit distribution is standard Normal and is not affected by the inclusion of deterministic components or the modelling assumptions about the  $I(0)$  disturbance term  $u_t$  in (6).

### **3. Empirical Analysis**

#### **3.1 Data**

We use data on energy consumption per capita expressed in terms of kg of oil equivalent, annually, from 1980 to 2014 for the GCC countries, namely Bahrain, Kuwait, Qatar, Oman, the Kingdom of Saudi Arabia and the United Arab Emirates. The series were obtained from the World Bank's World Development Indicators Database.

Figure 1 and 2 plot the raw and the logged data respectively. In both cases an upward trend can be noticed in all six GCC countries, which indicates an increasing average energy use per unit of output.

**[Insert Figures 1 and 2 about here]**

Following the GCC countries' ratification of the Kyoto protocol (in 2005 and 2006) calling for policy actions to limit GHG growth, there have been signs of decelerating demand growth. According to the World Oil Outlook (2017), in the non-OECD region, between 1980 and 1995, efficiency improvements were limited and energy intensity dropped at lower rates than in the OECD, on average by 0.7% per annum (p.a.). However, between 1995 and 2010, efficiency improvements in the non-OECD region accelerated. Despite the fact that the economy boomed, growing by 5.6% p.a., energy demand increased by only 3.8% p.a., which implies that energy intensity dropped on average by 1.7% p.a.

### 3.2 Empirical Results

We estimate the following model:

$$y_t = \beta_0 + \beta_1 t + x_t, \quad (1 - L)^d x_t = u_t, \quad t = 0, 1, \dots, \quad (3)$$

where  $y_t$  is the time series of interest,  $\beta_0$  and  $\beta_1$  are unknown coefficients corresponding respectively to an intercept and a linear time trend, and  $x_t$  is assumed to be integrated of order  $d$  (or  $I(d)$ ), which implies that  $u_t$  in the second equation in (3) is  $I(0)$ . Table 1 reports the estimated values of  $d$  (and their corresponding 95% intervals) for the three



cases of no regressors, an intercept, and an intercept with a linear time trend, assuming that the errors follow a white noise process.<sup>2</sup>

**[Insert Tables 1 and 2 about here]**

As can be seen, a time trend is required in four cases (Bahrain, Oman, Qatar and Saudi Arabia) for the original series and in three (the same countries except Saudi Arabia) for the logged series. The  $I(1)$  hypothesis, i.e.,  $d = 1$ , cannot be rejected in the majority of cases, the confidence interval including one, the exceptions being Bahrain with the raw data and Bahrain and Qatar with the logged ones – in these cases there is evidence of mean reversion, since the estimated value of  $d$  is significantly smaller than one. Table 2 reports the estimated values for the intercept and the time trend coefficients for each series. The latter are biggest in the case of Qatar with the raw data and Oman with the logged ones.

**[Insert Figures 3 and 4 about here]**

Figures 3 and 4 show the estimated trends for the raw and logged series respectively. There is an upward trend in the case of the raw series in Bahrain, Oman, Qatar and Saudi Arabia (Figure 3), and in the first three of these countries in the case of the logged ones (Figure 4). Note that the estimated values of  $d$  are below 1 for Bahrain and Qatar, i.e. mean reversion occurs, whilst in the remaining countries shocks appear to have permanent effects.

#### **4. Conclusions**

This paper examines the statistical properties of energy consumption in the GCC countries applying fractional integration methods to annual data from 1980 to 2014. The results indicate that both the raw and the logged series exhibit a (statistically significant)

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<sup>2</sup> The results are very similar when the errors are assumed to exhibit autocorrelation as in the model of Bloomfield (1973).

linear time trend in the case of Bahrain, Oman and Qatar, and the raw series only in the case of Saudi Arabia. Mean reversion (i.e., statistical evidence of  $d < 1$ ) is found in the case of Bahrain for both the raw and logged data, and in Qatar for the logged series. In the remaining cases, the  $I(1)$  hypothesis cannot be rejected except for the logged data in Saudi Arabia, since  $d$  is found to be statistically higher than 1 in that country. The implication of these findings is that in the case of Bahrain and Qatar exogenous shocks to energy consumption have transitory effects, which disappear in the long run without the need for policy action, whilst the permanent nature of the effects of shocks elsewhere means that appropriate policies have to be designed to restore equilibrium.

Future work will analyse possible non-linearities using the method proposed in Cuestas and Gil-Alana (2016) which estimates the order of integration of the series allowing for smooth non-linear terms in the form of Chebyshev polynomials in time - such an approach is suitable for modelling gradual changes as opposed to shifts in the parameters. Other non-linear specifications such as Fourier functions, STAR or ESTAR models could also be considered. Further, endogenous structural break tests could be carried out using the Bai and Perron's (2003) approach as well as the methods of Hassler and Meller (2004) and Gil-Alana (2008), both of which are specifically designed for the case of fractional integration; this is an important issue, since several studies have argued that long memory can be a spurious phenomenon caused by the presence of breaks in the data that have not been taken into account (see Diebold and Inoue, 2001; Granger and Hyung, 2004, etc.).

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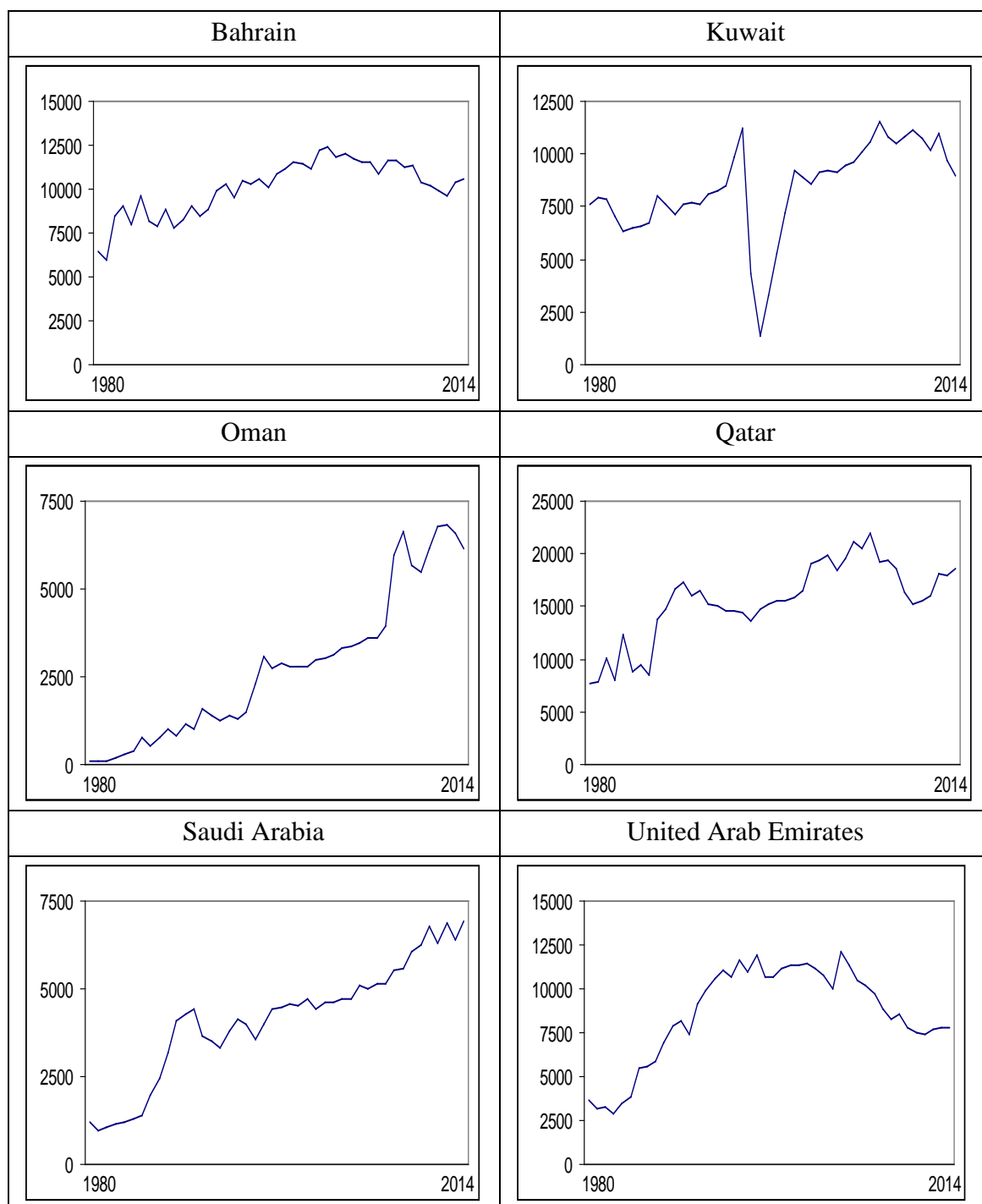
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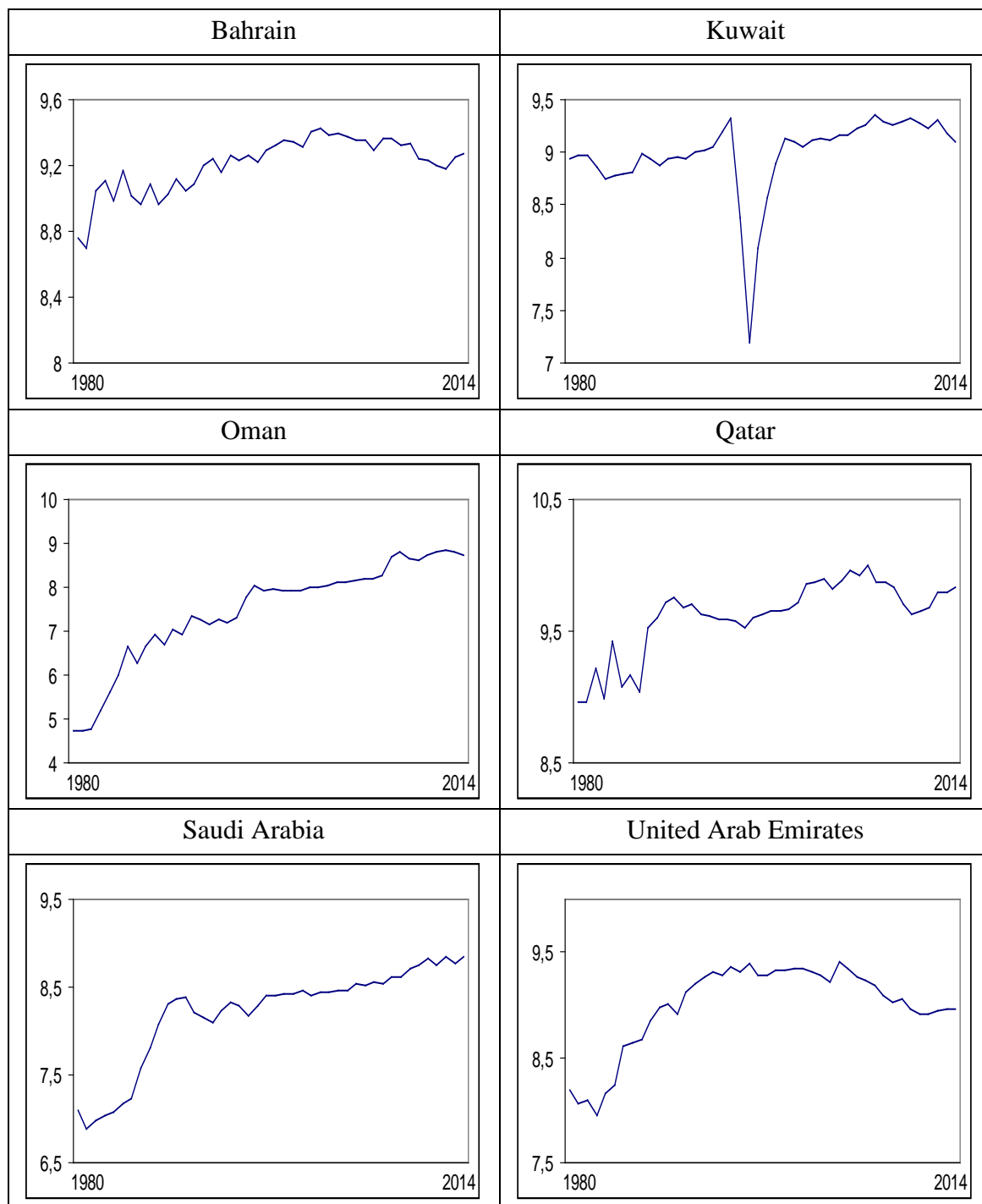
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**Figure 1: Time series plots (raw data)**



**Figure 2: Time series plots (logged data)**





**Table 1: Estimates of d and 95% confidence intervals with uncorrelated errors**

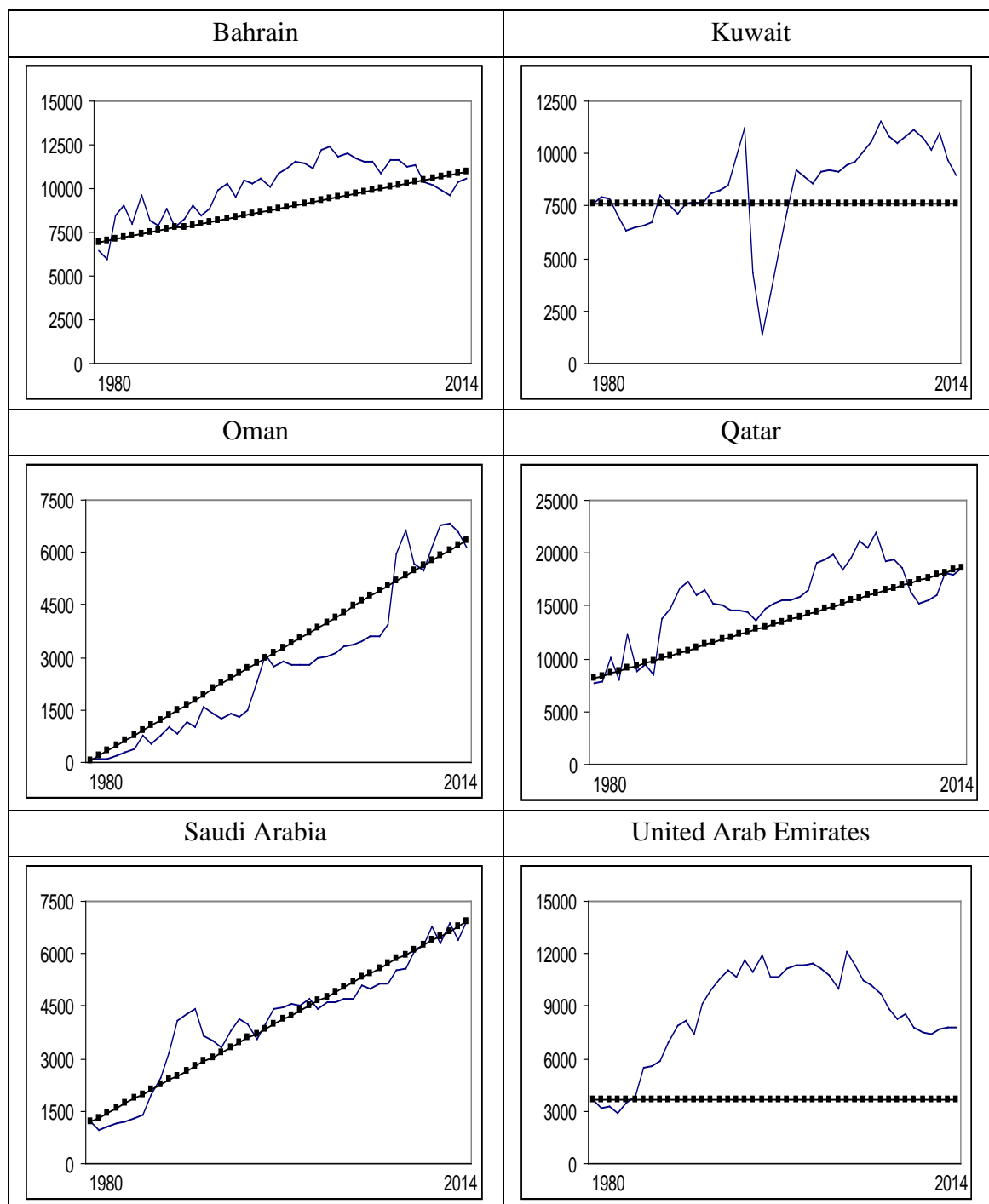
i) Raw data			
	No terms	An intercept	A linear time trend
BAHRAIN	0.89 (0.69, 1.12)	0.62 (0.50, 0.86)	<b>0.67 (0.53, 0.88)*</b>
KUWAIT	0.75 (0.40, 1.29)	<b>0.89 (0.54, 1.56)</b>	0.89 (0.43, 1.56)
OMAN	0.88 (0.72, 1.40)	0.90 (0.75, 1.42)	<b>0.80 (0.50, 1.43)</b>
QATAR	0.85 (0.61, 1.11)	0.74 (0.55, 1.00)	<b>0.77 (0.59, 0.99)</b>
SAUDI ARAB	0.89 (0.59, 1.25)	0.93 (0.70, 1.27)	<b>0.93 (0.70, 1.26)</b>
UNITED ARAB EMIRATES	0.98 (0.85, 1.16)	<b>1.04 (0.92, 1.21)</b>	1.04 (0.92, 1.21)
ii) Logged data			
	No terms	An intercept	A linear time trend
BAHRAIN	0.92 (0.73, 1.19)	0.56 (0.44, 0.82)	<b>0.64 (0.48, 0.87)*</b>
KUWAIT	0.89 (0.69, 1.19)	<b>0.76 (0.41, 1.42)</b>	0.75 (0.35, 1.42)
OMAN	0.90 (0.71, 1.15)	0.94 (0.68, 1.22)	<b>0.95 (0.80, 1.18)</b>
QATAR	0.92 (0.73, 1.20)	0.60 (0.46, 0.84)	<b>0.66 (0.50, 0.87)*</b>
SAUDI ARAB	0.90 (0.70, 1.17)	<b>1.28 (1.05, 1.59)</b>	1.26 (1.02, 1.62)
UNITED ARAB EMIRATES	0.90 (0.71, 1.18)	<b>1.10 (0.97, 1.29)</b>	1.09 (0.97, 1.28)

\*: Statistical evidence of mean reversion at the 5% level.

**Table 2: Estimated coefficients in the selected models in Table 1**

i) Raw data			
	d (95% band)	Intercept	Time trend
BAHRAIN	0.67 (0.53, 0.88)	6786.67 (10.47)	95.409 (2.57)
KUWAIT	0.89 (0.54, 1.56)	7647.36 (5.55)	----
OMAN	0.80 (0.50, 1.43)	-115.33 (-1.27)	147.24 (4.31)
QATAR	0.77 (0.59, 0.99)	7863.18 (5.14)	245.42 (2.18)
SAUDI ARAB	0.93 (0.70, 1.26)	1052.73 (3.10)	133.21 (3.28)
UNITED ARAB EMIRATES	1.04 (0.92, 1.21)	3620.85 (4.98)	----
ii) Logged data			
	d (95% band)	Intercept	Time trend
BAHRAIN	0.64 (0.48, 0.87)	8.8253 (119.96)	0.0111 (2.82)
KUWAIT	0.76 (0.41, 1.42)	8.9112 (32.60)	----
OMAN	0.95 (0.80, 1.18)	4.6475 (22.38)	0.0929 (3.50)
QATAR	0.66 (0.50, 0.87)	9.0102 (76.53)	0.0194 (2.95)
SAUDI ARAB	1.28 (1.05, 1.59)	7.1202 (74.55)	----
UNITED ARAB EMIRATES	1.10 (0.97, 1.29)	8.1998 (83.05)	----

**Figure 3: Estimated time trends (raw data)**



**Figure 4: Estimated time trends (logged data)**

