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Energy Consumption in the GCC Countries: Evidence On Persistence

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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. [1] 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, -10(e)-69563u2(e)11-10(e)-69563u2(e)

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

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

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

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.,

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

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 $\{x_t, t = 0, \pm 1, \ldots\}$ is said to be I(1) if it can be represented as

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

where x_t

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

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 overdifferenced 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 ≥ 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 ≥ 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

cases of no regressors, an intercept, and an intercept with a linear time trend, assuming that the errors follow a white noise process.[2]

[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 valuind [(ran)-b Td0.004<

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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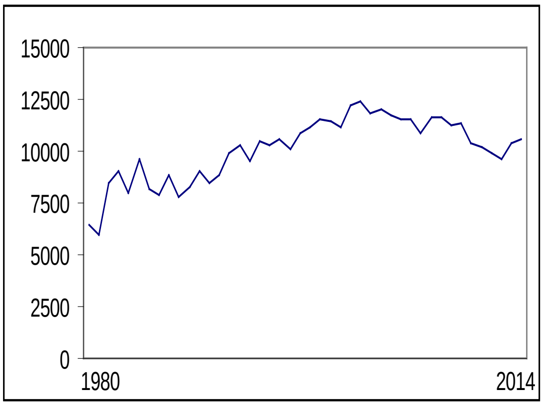
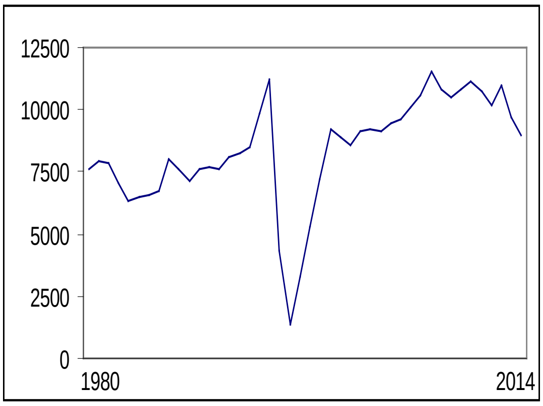
Figure 1: Time series plots (raw data)

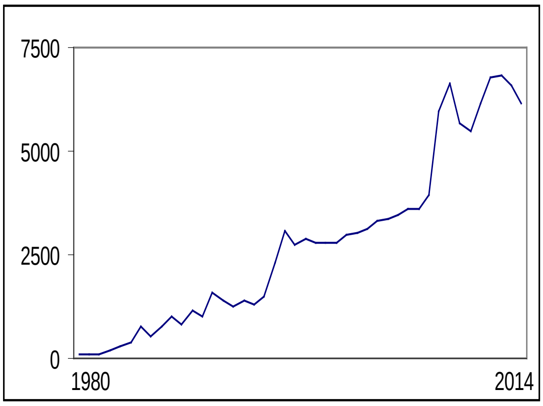
Bahrain



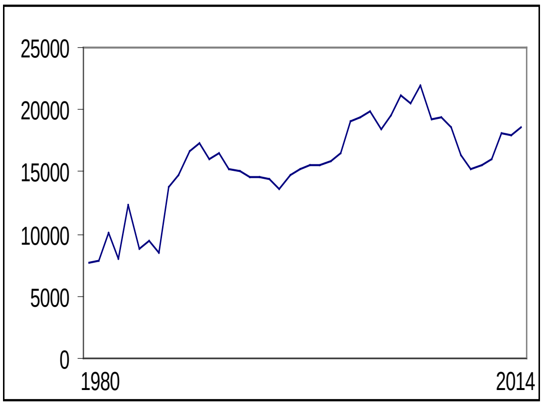
Kuwait



Oman



Qatar



Saudi Arabia

United Arab Emirates

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)

Figure 3: Estimated time trends (raw data)

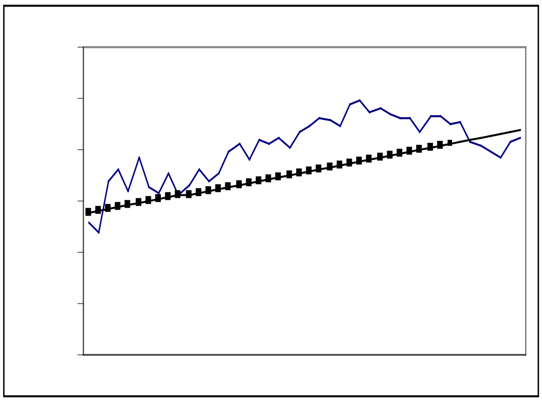
Bahrain	Kuwait
	
Oman	Qatar
Saudi Arabia	United Arab Emirates

Figure 4: Estimated time trends (logged data)

Bahrain Kuwait