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Fractional cointegration, Energy Consumption and Growth Revisited: Evidence from Taiwan

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Abstract

Taiwan is increasingly dependent on energy for its growth and development. This is partly related to the encouragement to export and expand into the international market, which establish the foundation for its economic development since the export oriented policy in the 1960s. As a result and for realizing the future development and growth objectives, the relationship between GDP and energy consumption is of a central concern. This paper adds to the literature and its mixed results by analyzing causality between growth and energy consumption in the case of fractional cointegration using annual data for Taiwan. With some exceptions, there is evidence of a fractionally cointegrated process with a mean-reverting non-stationary long memory. The result of the causality tests incorporating fractional cointegration emphasize that an energy conservation policy implies a negative effect on growth in Taiwan with energy acting as an engine of growth. Thus, an energy conservation policy as part of a policy to optimize the use of a scarce resource as well as to reduce pollution can lead to a fall in growth. Furthermore, evidence related to weak exogeneity interpreted as a test of long-run causality can explain the evidence of a mean-reverting process, i.e. supporting a relationship between the variables in the long run but only very weakly. As a result concerning policy analysis of the empirical relationship between growth and energy consumption, the issue of fractional cointegration needs to be taken into account.

Keywords: Energy consumption; GDP; Fractional cointegration; Causality; Taiwan

JEL Classification: Q43; C32

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1. Introduction

Production and consumption of goods and services implies use of energy as an important input. It is generally recognized that the supply of energy plays a significant role in economic development. Among other things, it is assumed to enhance the productive use of factors of production such as labor, capital and land. In such a case, an adequate and regular power supply to support economic growth may be one of the most crucial factors. On the other hand, economic growth itself may imply an increased demand and use of energy. A central concern in energy policy and conservation plans is related to the direction, strength and stability of the relationship between economic growth and energy consumption. During the recent decades, this inter-relationship has been one of the most intensive areas of research in the energy economics literature.[2] However, the focus in the literature has mainly been on the direction of causality and its strength with little focus on the stability and order of integration of the relationship.

Many papers have since the seminal paper by Kraft and Kraft (1978) analyzed the relationship between economic growth and energy consumption. The interest is related to whether economic development takes precedence over energy consumption or whether energy is an input to increase economic growth. The empirical evidence of the direction of causality between economic growth and energy consumption is mixed and controversial across time-periods, countries and methodologies implying major policy inconsistencies concerning economic development and energy conservation. The mixed empirical evidence includes bi-directional and unidirectional causality to no causality as outlined in Jumbe (2004), Wolde-Rufel (2004) and Ozturk (2010). One possible explanation of the mixed results is the focus in the literature mainly on the direction of causality and its strength with little focus on the stability and order of integration. Furthermore, these conflicting results have major policy implications concerning economic development and energy conservation policies focusing on efficient use of scarce resources and environmental aspects. An energy conservation policy may be implemented with little or no adverse effect on economic growth if there is a unidirectional causality from economic growth to energy consumption. A policy to reduce energy consumption can on the other hand lead to a fall in economic growth if the causality runs from energy consumption to economic growth.

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[2] See Ghali and El-Sakka (2004) and Lee (2005).

If there is no causality in either direction, i.e. the neutrality hypothesis as discussed in Yu and Choi (1985), an energy conservation policy may not affect economic growth as outlined in Asafu-Adjaye (2000). As a result and given a causality from energy consumption to economic growth where energy conservation policies may lead to a fall in economic growth, the policy maker needs to take into account the trade-off between an increase in economic growth and material well-being and a decrease in pollution to optimize the combined growth- and environmental policy. Hence, the direction of causality is of central concern for growth- and environmental policies aiming at enhancing economic development and reducing pollution and green house gases.

The result in the paper by Kraft and Kraft (1978), finding evidence of a unidirectional causality in the USA from GNP to energy consumption over the period 1947 – 1974, implies that energy conservation policies might be enforced without affecting GNP growth. However, Akarca and Long (1980) did not detect causality in the USA when the period was shortened explained by a possible temporal time-period instability. Furthermore, Warr and Ayres (2010) for an extended period 1946 – 2000 did not detect causality in the USA. To sustain long-term growth, Warr and Ayres (2010) conclude that it is necessary to either increase energy supplies or increase the efficiency of energy usage where negative externalities of fossil fuel use implies that the latter option is preferred. Since the paper by Kraft and Kraft (1978), many papers have added to the literature with mixed results. In line with the results and conclusions in Kraft and Kraft (1978), a unidirectional Granger causality from economic growth to energy consumption is reported in e.g. Yu and Choi (1985) and Soytas and Sari (2003) for South Korea, Cheng and Lai (1997) for Taiwan, Lise and van Montfort (2007) for Turkey and Ozturk et al. (2010) for a group of low income countries. In contradiction to the results in Kraft and Kraft (1978), a unidirectional Granger causality from energy consumption to economic growth was found among others in Yu and Choi (1985) for the Philippines, Masih and Masih (1996) for India, Glasure and Lee (1997) for Singapore, Asafu-Adjaye (2000) for Indonesia, Soytas and Sari (2003) for France, West Germany, Japan and Turkey and Wolde-Rufael (2004) for Shanghai.

A unidirectional causality from economic growth to energy consumption implies that energy saving policies would not harm economic growth while a unidirectional causality from energy consumption to economic growth implies that energy conservation policies can lead to a fall in economic growth. Further papers uncover a bi-directional causality also incorporating the conclusion that a fall in economic growth can be the result of an energy conservation policy. Some papers uncovering a bi-directional causality include Stern (1993), Cheng (1995) and Stern (2000) for the USA, Masih and Masih (1996) for Pakistan, Asafu-Adjaye (2000) for Thailand and the Philippines, Yang (2000) as well as Lee and Chang (2005) for Taiwan, Soytas and Sari (2003) for Argentina, Mishra et al. (2009) for a panel of Pacific Island countries and Ozturk et al. (2010) for a group of middle income counties. Furthermore testing the energy-growth causality for G7-countries 1960 – 2006 using bootstrap non-Granger causality tests with fixed sub-samples, the paper by Balcilar et al. (2010) did not find consistent causal links but only in various sub-samples and various causality-directions corresponding to significant economic events.

As a concluding result in the survey on the energy-growth nexus by Ozturk (2010), there are some methodological reservations about the results producing the conflicting conclusions. Thus, resulting in that there is no consensus on its existence or on the direction of its causality generating policy conclusions that are controversial across time-periods, countries and methodologies. As an example, Glasure and Lee (1997) reported a bi-directional causality for South Korea and Singapore using cointegration and error-correction models. However using a standard Granger causality test, a unidirectional causality from energy consumption to economic growth as well as no causality was reported for Singapore and South Korea, respectively. The papers by Oh and Lee (2004a) and Oh and Lee (2004b) did use similar data periods[3] for South Korea together with a Vector Error Correction Model (VECM). The results in Oh and Lee (2004a) indicate a long run bi-directional causality between energy consumption and GDP with a short run unidirectional causality running from energy consumption to GDP while Oh and Lee (2004b) indicate no causality in the short run and a unidirectional causality running from GDP to energy consumption in the long run. These two contradicting results imply different policy conclusions still using a VECM and similar data periods for South Korea in both studies.

[3] In Oh and Lee (2004a) data for 1970 – 1999 was used and in Oh and Lee (2004b) data for 1981 – 2000 was used.

Furthermore using different methods, Cheng and Lai (1997) reported a unidirectional causality from economic growth to energy consumption in Taiwan using data for 1954 – 1993 while Yang (2000) using data for 1954 – 1997 and Lee and Chang (2005) using data for 1954 – 2003 reported a bi-directional causality.

The purpose of this paper is to examine the causal relationship between various aggregated and disaggregated energy consumption variables and GDP using fractionally cointegrated methods. In relation to the empirical methodologies used in previous research, the fractional cointegration methods and analysis is a more flexible methodology allowing for more subtle forms of mean reversion. Furthermore, the argument to use fractional cointegration is related to that many papers analyzing this relationship is based on unit root tests and classical cointegration procedures. However, conventional unit root tests have some restricting assumptions about the value of *d*, i.e. they do not perform well in cases of fractionally integrated processes. Furthermore, the classical cointegration procedure, such as the Johansen and Juselius (1990) procedure, assumes that the error-correction term follows an $I(0)$ process where the process instead might be of a fractional order. The fractionally integrated procedures allow this possibility. The test for causality using fractional cointegration is performed using annual data for Taiwan during 1954 – 2013.[4]

The rest of the paper is organized as follows. Section two describes the methodology while section three presents data and empirical results and section four concludes with policy implications.

2. Empirical Methodology

The classical cointegration procedure, such as the Johansen and Juselius (1990) procedure, assumes that the error-correction term follows an $I(0)$ process with moving average coefficients incorporating an exponential decay.

[4] Data for Taiwan is also used in e.g. Lee and Chang (2005). However, the empirical methodology adopted in Lee and Chang (2005) relies on the assumption that the error-correction term follows a stationary process with moving average coefficients incorporating an exponential decay. This paper adds to the analysis and policy conclusions by allowing the error-correction term to differ from the stationary process in such a way that the autocorrelations decline at a slower rate than the exponential decay. Thus, allowing for a more flexible class of processes incorporating the possibility of a mean-reverting process between economic growth and energy consumption but only in a long run perspective.

However, the estimated values of the error-correction term may differ from an $I(0)$ process in such a way that the autocorrelations decline at a slower rate than the exponential decay of the ARMA process in the classical cointegration technique. The fractionally integrated procedures allow this possibility. Thus, the cointegrating vector or error-correction term for each set of variables is used to estimate the error-correction values derived from the cointegrating relationship in line with Villeneuve and Handa (2006). Using the procedures of long-memory, the parameter of integration d is estimated testing whether the residuals of the cointegrating relationship are $I(d)$ with $0 < d < 1$ or not.

The relationship or memory between observations in different time-periods is often measured by the autocorrelation between the observations in time period t and $t + k$. The k:*th* autocorrelation, $\rho(k)$, of a stochastic process can be defined by $\rho(k) = Ak^{2d\text{-}1}$ where A is a suitable constant and d is the memory parameter.[5] The stationary short memory[6] process is characterized by $\sum_{k=-\infty}^{\infty}|\rho(k)| < \infty$ and a significant estimate within $-0.50 < d < 0$ and the stationary long memory process is characterized by $\sum_{k=-\infty}^{\infty}|\rho(k)| = \infty$ and a significant estimate within $0 < d < 0.50$. For $d = 0$, the process is stationary with a geometric decay and $d = 1$ implies a unit root in the process implying that shocks persists into the infinite future, i.e. the series follows a random walk. For $0.50 \leq d < 1$, the process possesses a long memory but is non-stationary and mean-reverting since an innovation will have no permanent effect on its value.[7] For $d > 1$ the process is non-stationary and mean-diverting.[8]

Granger and Joyeux (1980) and Hosking (1981) introduced the fractional ARFIMA (p,d,q) model, a more flexible class of processes, which includes the method of fractional differencing and incorporates long memory dynamics into time series models.

[5] The common definition of a long memory process defined by autocorrelations is that the autocorrelation function has a sufficiently slow decay or, more precisely, that it is bounded by a hyperbolic decaying function. The corresponding decay of a short memory process is at least exponential.

[6] Rosenblatt (1956) defines short memory as the case in which the dependence between two points of a process becomes trivially small as the distance between these points increases. That is, the dependence between distant observations decays exponentially.

[7] This is in contrast to an $I(1)$ process which will be both covariance non-stationary and non-mean-reverting where the effect of an innovation will persist forever.

[8] For a further discussion see Hosking (1981), Brockwell and Davis (1991), Alves et al. (2001), Gil-Alana and Toro (2002) and Tolvi (2003) among others.

The occurrence of long-term dependence in the data is conditional on the fractional differencing parameter *d* where a general class of long memory processes can be described by

$$\phi(L)(1-L)^d(y_t-\mu)=\theta(L)\varepsilon_t \quad t = 1, 2, \ldots \qquad (1)$$

where $\phi(L)$ and $\theta(L)$ are polynomials in the lag operator (L) with *p* and *q* lags, respectively, and all roots of $\phi(L)$ and $\theta(L)$ are stable. The white noise disturbance term is defined as ε_t where $(y_t-\mu)$ is the equilibrium error. The difference parameter *d* is allowed to take any real value including fractional values indicating the dynamics of the memory process in the series.[9] When $d = 0$, the usual ARMA(*p*,*q*) model is included as a special case. Furthermore, the existence of a cointegrating relationship between the variables studied requires that the equilibrium error is mean-reverting. Thus, implying that a shock to the system of variables studied will not tend to permanently drive the system out of equilibrium. Therefore as a complement to the mixed results in the literature using classical cointegration techniques, using an ARFIMA model for the relationship between economic growth and energy consumption and examining whether *d* falls in the range $0 < d < 0.50$ or $0.50 \leq d < 1$ provides a test of whether the relationship possesses a long memory relationship incorporating mean-reversion or mean-diversion.

The techniques employed in this paper include the Geweke and Porter-Hudak (GPH) (1983) frequency domain estimator and the biased-reduced technique by Andrews and Guggenberger (AG) (2003). The GPH-estimator is a semi-parametric test for fractional processes using only the lowest frequency ordinates of the log periodogram, i.e. it does not require any specification of the short memory process or the ARMA-part. This is so since only a fraction of the first frequencies are used. The GPH-procedure tests whether the error-correction term is $I(0)$ or $I(d)$ with $0 < d < 0.5$. The estimation of *d* is based on the log periodogram regression

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9 The extension to allow non-integer *d*-values raises the flexibility in modeling long-term dynamics and allows for a more rich class of spectral behavior at low frequencies. The extension to allow non-integer *d*-values raises the flexibility in modeling long-term dynamics and allows for a more rich class of spectral behavior at low frequencies. Upon using a return series as the dependent variable, $d > 0$ indicates a mean-diverting property while $d < 0$ indicates a mean-reverting property in the original series.

$$\ln(I(w_j)) = \beta_0 + \beta_1 \ln(4\sin^2(w_j/2)) + \varepsilon_j \quad j = 1, \ldots, n \tag{2}$$

where w_j is the harmonic frequency with $w_j = {2\pi j}/{T}$, $j = 0, \ldots, T-1$, with T as the size of the series, $I(w_j)$ is the periodogram at ordinate j, ε_j is independent across harmonic frequencies and equal to $\ln(I(w_j)/f_G(w_j))$ and $f_G(.)$ is the spectral density. Obtained by the ordinary least square estimator of d in equation (2), the asymptotic mean of $\ln(I(w_j)/f_G(w_j))$ and its variance $\pi^2/6$ is assumed to behave as *i.i.d.* random variables where the estimator of d is asymptotically standard normally distributed. This implies that hypothesis testing concerning d can be based on the t-statistics related to the normal distribution. As discussed in Geweke and Porter-Hudak (1983), the asymptotic distribution of d will neither depend on the order of the short memory components nor on the distribution of the error term in the ARFIMA model upon a proper choice of n in equation (2). Note also that n $(=g(T))$ is an increasing function of T, i.e. the size of the series. Usually, $g(T)$ equals T^u for $0 < u < 1$.[10]

Agiakloglou et al. (1993) criticized the GPH-procedure due to its finite-sample biases. To eliminate the first- and higher-order biases in the GPH-estimator, Andrews and Guggenberger (2003) proposed a bias-reduced log-periodogram regression estimator. It is equivalent to the GPH-estimator except that it includes frequencies to the power $2k$ for $k = 1, \ldots, r$, for some positive integer r, as additional regressors in equation (2). This does not affect the asymptotic bias, variance, mean-squared error or normality of the estimator as discussed in Andrews and Guggenberger (2003).

[10] The choice of u in the spectral regression of the GPH-estimator involves a judgment call. A too large value will cause contamination of the long memory estimate due to that medium- or high frequency components and thus more short-term influences are included in the regression. This will affect the form of the spectral density used in the estimation. As argued in Sowell (1992), using a too large bandwidth T leads to misspecification of the short-run dynamics and too strong mean-reversion not significantly in line with the data. Therefore, Sowell (1992) claims that the number of ordinates should be based on the shortest cycle associated with the long memory. Furthermore, a too small value will lead to imprecise estimates due to limited degrees of freedom as discussed in Cheung and Lai (1993). In order to be robust against the bandwidth, estimates for various values of u will be reported using 0.60 and 0.65 for u in this paper within $0.475 \leq u \leq 0.80$ as argued in Geweke and Porter-Hudak (1983), Cheung and Lai (1993) and Siokis and Christodoulou (2004).

Thus, equation (2) is evaluated by the ordinary least square estimator of the coefficient on –2log λ_j in a regression of the log I_j evaluated at λ_j on a constant, -2log λ_j, and λ_j^{2k} for j = 1, …, m where m is defined such that $m \to \infty$ and $\left(m/n\right) \to$ 0 as $n \to \infty$. Furthermore, the GPH- and AG-procedure is asymptotically equivalent when r = 0. Although the asymptotic results established in Andrews and Guggenberger (2003) hold for arbitrary large values of r, it is recommended to use small values of r such as r = 1 or r = 2. The recommendation is related to that the asymptotic properties of the estimator will not be reflected in finite samples using large values.[11] It is shown that the AG-estimator attains the optimal rate of convergence for spectral densities, which is smooth of order $s \geq 1$ at zero, when $r \geq \left(s - 2/2\right)$ and m is chosen appropriately. This property is not attained by the GPH-estimator for $s > 2$.

To test for causality between the variables, the paper follows three steps. First from the consistent and asymptotically normal estimates of d by the GPH- and AG-estimator, the series are transformed by the expansion $(1 - L)^d$ where L is the lag operator into stationary $I(0)$ series (Diebold and Rudebusch, 1991). Second, test of a cointegrating relationship using the transformed series is performed for each set of variables. Third, the VAR-system underlying the cointegration test is reduced into single equations using weak exogeneity tests.[12] The test for weak exogeneity in the system as a whole requires a test of the hypothesis $H_0 : \alpha_{ij}$ = 0 for j = 1, …. r where i only contains zeros and r is the number of cointegrating vectors. The test is based on the likelihood ratio (LR) test including the restricted and unrestricted model to analyze the validity of the restrictions.

As most economic series exhibits stochastic non-stationarity and classical inference theory generally assumes that the series are stationary, the stochastic non-stationarity in economic series undermines the foundation in classical inference theory.

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[11] This paper uses r = 1 and r = 2 being in the range prescribed by Andrews and Guggenberger (2003).

[12] A test for weak exogeneity in a cointegrated system is interpreted as a test to examine long-run causality where the null hypothesis interprets the existence of weak exogeneity (Johansen and Juselius, 1992, Hall and Alistair, 1994, and Arestis et al., 2001).

Furthermore, the Johansen and Juselius (1990) test is carried out with a prerequisite for cointegration that non-stationary series are integrated of the same order. To test for the existence of unit roots and identify the order of integration for each variable, the Augmented Dickey-Fuller unit root test (ADF) by Dickey and Fuller (1981) and the Phillips-Perron (PP) (1988) test are used. In order to take into account a possible structural break in the data, the Zivot and Andrews (ZA) (1992) unit root test will be employed. The ZA-test allows for an endogenous structural break where the test is allowed for a unit root against the alternative of a trend stationary process with a structural break. The breakpoint in the ZA-test is selected where the test statistic of the null of a unit root is the most negative for the *t*-statistic of the coefficient of the autoregressive variable. This test is included as the classical unit root tests may be suspect not taking into account that a structural break can lead to a wrong decision when the null hypothesis is not rejected. Furthermore and prior to the cointegration test, test for structural breaks by Recursive Least Square (RLS) and the step-wise Chow (1960) test to ensure within-sample coefficient constancy as well as test for seasonal-dummies and trend in data are performed.[13]

3. Data and Empirical Results

Data are taken from the AREMOS economic-statistic data base, the Enerdata data base and the UN Statistical data base including 60 annual observations for the period 1954 – 2013 for Taiwan. The data are in natural logarithms and descriptive statistics are given in table 1. Data includes the real gross domestic product per capita in 1996 prices and Taiwanese currency (GDP). Energy consumption includes aggregated as well as disaggregated data of energy consumption expressed in terms of kiloliters of equivalent oil. The aggregated data is total energy consumption (ENERGY) and the disaggregated data include coal consumption (COAL) and oil consumption (OIL) composing the majority of aggregate energy consumption in Taiwan.

[13] Dummy variables are included in the VAR-model underlying the cointegration test to represent possible structural breaks and external shocks to the markets, such as e.g. a policy change related to economic and political reforms such as the expansionary export trade policy in Taiwan adopted in the 1960s.

Table 1: Descriptive Statistics

Variable	Mean	S.D.	Min	Max	Skewness	Excess Kurtosis	JB
GDP	18.67	0.96	17.13	20.03	-0.14	-1.37	9.53**
Energy	7.18	0.97	5.40	8.54	-0.28	-1.26	9.71**
Coal	5.47	0.46	4.74	6.22	0.27	-1.29	10.21**
Oil	6.15	1.22	3.61	7.58	-0.61	-1.11	22.64**
Correlation							
	GDP		Energy		Coal		Oil
GDP	1.000						
Energy	0.997		1.000				
Coal	0.811		0.779		1.000		
Oil	0.974		0.987		0.677		1.000

Note: S.D. denotes standard deviation and JB denotes Jarque-Bera test for normality where ** indicates significance at the 1 percent.

Table 2 reports results of the unit root tests without and with a possible structural break where the lag parameters are selected based on the Akaike Information Criterion (AIC).[14] The null hypothesis of a unit root is by the ADF-test and PP-test not rejected for the series in log level but found to be stationary in first difference at the 1 percent significance level implying integration of order one. The results of the ZA-test allowing for a structural break point indicate that all variables are integrated of order one at the 1 percent significance level.[15] Thus, the results of the unit root tests are consistent also allowing for a structural break in the data indicating a unit root in the log level series but stationarity in the first difference of the log level series, i.e. $I(1)$ processes.

[14] Plot of actual values of the series indicate that they exhibit trends. Hence, the tests are run with a constant and trend included. To conserve space, the figures are not reported here.

[15] The structural breaks are as in Lee and Chang (2005) except for coal consumption in levels where the break is indicated at 1982 instead of 1971 as in Lee and Chang (2005).

Table 2: Unit Root Tests

	GDP	Energy	Coal	Oil
ADF (level)	-0.49(3)	-0.63(2)	-1.54(1)	-1.18(2)
ADF (first diff.)	-4.07(2)**	-4.49(1)**	-6.18(0)**	-3.93(1)**
PP (level)	-1.16(3)	-2.64(2)	-0.22(1)	-2.51(2)
PP (first diff.)	-4.54(2)**	-6.31(1)**	-6.91(0)**	-6.44(1)**
ZA (level)	-4.17(1)	-3.99(0)	-3.58(2)	-4.03(0)
Year of break	1995	1976	1982	1971
ZA (first diff.)	-6.25(0)**	-8.10(0)**	-6.23(1)**	-8.64(0)**
Year of break	1974	1981	1978	1981

Note: Number in parentheses are the lag order in each unit root test. Significance at the 1 percent level is indicated by **.

Since all variables are $I(1)$, the Johansen and Juselius (1990) procedure is used to check for cointegrating vectors between GDP and the various energy variables.[16] The results reported in table 3 indicate that there is one cointegrating vector for each par of variables.[17] To visually observe whether the error-correction term, i.e. the cointegrating vector, follows a mean-reverting process or not, the residuals are plotted to observe if the pattern deviates from equilibrium in the short run.[18] The pattern did not appear to be associated with a mean-reverting process. Thus, indicating at the possible non-existence of fractional cointegration for each set of variables, i.e. GDP and an energy consumption variable.

16 This procedure assumes that the error-correction term from the cointegrating relationship is $I(0)$ with moving average coefficients that decline exponentially.

17 Prior to the test, test for structural breaks are performed by RLS and step-wise Chow test as well as test for trend and seasonal-dummies in the data. In contrast to Lee and Chang (2005), no outliers or structural breaks were indicated by the 1-step Chow test. The RLS and step-wise Chow test indicates within-sample coefficient constancy and parameter stability in contrast to the findings of structural breaks in Lee and Chang (2005). To conserve space, results are not reported here.

18 To conserve space, the figures are not reported here.

Table 3: Johansen cointegration Test

Variables	VAR-model reduction, *F*-test	Hypothesis		Trace-statistic
		Null	Alternative	
GDP, Energy	1996(1)**	$r = 0$	$r >= 1$	17.03**
		$r <= 1$	$r >= 2$	5.13
GDP, Coal	2596(1)**	$r = 0$	$r >= 1$	17.84**
		$r <= 1$	$r >= 2$	2.22
GDP, Oil	3498(1)**	$r = 0$	$r >= 1$	19.61**
		$r <= 1$	$r >= 2$	5.26

Notes: The cointegration test with critical values according to Osterwald-Lenum (1992). The notation ** denotes significance at the 1 percent level. In parentheses, optimal lag according to the *F*-test. The *r* denotes the maximum number of cointegrating vectors.

However, conventional unit root tests have some restricting assumptions about the value of *d*, i.e. they do not perform well in cases of fractionally integrated processes. Furthermore, the classical cointegration procedure, such as the Johansen and Juselius (1990) procedure, assumes that the error-correction term follows an $I(0)$ process where the process instead might be of a fractional order. The fractionally integrated procedures allow this possibility. Fractional cointegration was tested for the overall period and it is divided into three steps for each set of variables studied, i.e. the GDP and an energy consumption variable. First, the values of the error-correction term are derived using the cointegrating vector for each set of variables. Second using the derived residuals and the GPH- and AG-procedure, the fractionally integrated parameter *d* is estimated, i.e. the long memory parameter. Third, the memory parameter of the residuals from the cointegrating relationship are tested whether it is an $I(d)$ process with $0 < d < 1$ or not. The sample size for the GPH-estimator is given by $n = T^u$ where the choice of *u* in this paper is 0.60 and 0.65 being in the range prescribed by Geweke and Porter-Hudak (1983), Cheung and Lai (1993) and Siokis and Christodoulou (2004). Furthermore, the asymptotic results established in Andrews and Guggenberger (2003) hold for arbitrary large values of *r* where it is recommended to use small values of *r*. This is related to that the asymptotic properties of the estimator will not be reflected in finite samples using large values. This paper uses $r = 1$ and $r = 2$ being in the range prescribed by Andrews and Guggenberger (2003).

The test for fractional cointegration by the GPH- and AG-estimator is a test of the hypothesis H_0: $d = 0$ against the alternative H_1: $d \neq 0$. By use of the cointegrating vector and the derived residuals, the hypothesis test is based on the *t*-statistic of the regression coefficient. Three hypothesizes are tested. First, the hypothesis H_0: $d = 0$ is tested to analyze if the memory parameter of the residuals from the cointegrating relationship possesses a long-memory structure. Second, the hypothesis H_0: $d = 0.50$ is tested to analyze if it is a stationary long-memory structure with mean-reversion or if it is a non-stationary process with mean-reversion. Third, the hypothesis H_0: $d = 1$ is tested to analyze if it is a non-stationary and mean-diverting process.

Table 4 reports the results from the GPH- and AG-estimator for both values of u and r as well as the *t*-statistics for H_0: $d = 0$, H_0: $d = 0.50$ and H_0: $d = 1$. For growth and the energy variable at the aggregate level, i.e. energy consumption, the memory-estimates were consistently greater than 0.50. Moreover, H_0: $d = 0$ was rejected at the 5 percent level by the GPH-estimator for both values of u and by the AG-estimator at the 1 percent and 5 percent level for both values of r, respectively. By both estimators and both values of u and r, H_0: $d = 0.50$ is not rejected. Furthermore by the AG-estimator for both values of r, the memory-estimate is significantly different from one indicating that the residuals are in the range $0.50 \leq d < 1$, i.e. a fractionally cointegrated process with a mean-reverting non-stationary long memory.[19]

[19] The GPH-estimator for both values of u indicates that the memory-estimate is not significantly lower than one. Due to the critique of the finite-sample bias in the GPH-estimator presented in Agiakloglou et al. (1993) and the elimination of the first- and higher-order biases by the AG-estimator, the conclusion related to the hypothesis is based on the AG-estimator.

Table 4: Results from the GPH- and AG-estimator

Estimator		GDP, Energy	GDP, Coal	GDP, Oil
GPH u = 0,60	d-estimate	0.69	0.42	0.52
	H_0: $d = 0$	2.35*	1.63	3.09**
	H_0: $d = 0.50$	0.61	-0.36	0.15
	H_0: $d = 1$	-1.01	-2.36*	-2.87**
GPH u = 0,65	d-estimate	0.67	0.41	0.71
	H_0: $d = 0$	2.49*	1.55	2.69**
	H_0: $d = 0.50$	0.60	-0.36	0.71
	H_0: $d = 1$	-1.25	-2.25*	-1.12
AG $r = 1$	d-estimate	0.67	0.71	0.67
	H_0: $d = 0$	6.79**	6.49**	6.78**
	H_0: $d = 0.50$	1.71	1.90	1.74
	H_0: $d = 1$	-3.37**	-2.72**	-3.36**
AG $r = 2$	d-estimate	0.60	0.86	0.55
	H_0: $d = 0$	3.23**	4.63**	2.97**
	H_0: $d = 0.50$	0.51	1.84	0.20
	H_0: $d = 1$	-2.13*	-0.76	-2.34*

Note: t-statistics for H_0: $d = 0$, H_0: $d = 0.50$ and H_0: $d = 1$ where ** and * indicates significance at the 1 percent and 5 percent level, respectively.

For growth and the energy variables at the disaggregated level, i.e. coal consumption and oil consumption, the memory-estimates were consistently greater than 0.50 except for coal consumption by use of the GPH-estimator at both levels of u. The hypothesis H_0: $d = 0$ was rejected at the 1 percent level for both variables by use of the AG-estimator by both values of r and for oil consumption by use of the GPH-estimator by both values of u. However, H_0: $d = 0$ was not rejected for coal consumption by the GPH-estimator for both values of u. Furthermore, H_0: $d = 0.50$ is not rejected for the variables by both estimators and values of u and r, respectively. The hypothesis H_0: $d = 1$ is significant for coal consumption and oil consumption at the 1 percent or 5 percent level, respectively, except for coal consumption by use of the AG-estimator at $r = 2$ and for oil consumption by use of the GPH-estimator at $u = 0.65$. The set of hypothesis tests for the two disaggregated energy variables implies similar results as to the aggregated energy variable with some exceptions.

For coal consumption by the GPH-estimator, there is evidence of a memory-parameter for the residuals in the range $0 \leq d < 1$, i.e. a long-memory process with or without a stationary property.[20] By use of the AG-estimator and $r = 1$ for coal consumption, the memory-parameter falls in the range $0.50 \leq d < 1$, i.e. a fractionally cointegrated process with a mean-reverting non-stationary long memory. By use of the AG-estimator with $r = 2$, result implies a memory-parameter in the range $0.50 \leq d \leq 1$ indicating a possible unit root in the process. Hence, implying that shocks persist into the infinite future as H_0: $d = 1$ is not rejected, i.e. the series follows a random walk. For oil consumption by the GPH- and AG-estimator by both values of u and r, respectively, except for oil consumption by the GPH-estimator and $u = 0.65$, there is evidence of a memory-parameter for the residuals in the range $0.50 \leq d < 1$, i.e. a fractionally cointegrated process with a mean-reverting non-stationary long memory.

The results show that $d > 0$ for all variables except for coal consumption by use of the GPH-estimator for both values of u. Furthermore, the hypothesis H_0: $d = 0.50$ is not rejected for the set of variables. However, H_0: $d = 1$ is rejected at the 1 percent and 5 percent level, respectively, except for energy consumption by the GPH-estimator for both values of u and for oil consumption for $u = 0.65$ as well as for coal consumption by the AG-estimator for $r = 2$. Hence, there is evidence of a fractionally cointegrated process with a mean-reverting non-stationary long memory where the error-correction term did not generally follow a stationary process. The evidence of a fractional cointegration structure implies a failure to achieve equilibrium for very long periods where shocks induce non-stationary deviations from the long-run relationship between economic growth and energy consumption established from the cointegrating vector.[21] Furthermore, an estimate significantly below 1 indicates that the error-correction term is a mean-reverting process. Hence, the estimated d values are in the non-stationary but mean-reverting range. These results support a relationship between the variables in the long run but very weakly.

Table 5 reports the results from the Augmented Dickey-Fuller (ADF) test with a constant and trend included where lag lengths are chosen by the Akaike Information Criteria (AIC). The ADF-test indicates a unit root in the transformed series, i.e. a stationary series.

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[20] A stationary process implies $0 < d < 0.50$ and a non-stationary process implies $0.50 \leq d < 1$.

[21] The existence of a cointegrating relationship between the variables requires that the equilibrium error is mean-reverting. This behaviour of the equilibrium error is of a central interest due to that without a mean-reversion of the equilibrium error, a shock to the system of variables will tend to permanently drive the system out of equilibrium.

Table 6 reports the results of the Johansen and Juselius (1990) cointegration test where lag lengths are chosen by the AIC indicating a VAR(1) model.[22] The test suggests that there exists one cointegrating vector among all set of variables. The results for weak exogeneity in the cointegrating relationship for each set of variables are outlined in table 7. There is evidence of a bi-directional causal link between GDP and both total energy and oil consumption rejecting weak exogeneity at the 5 percent and 1 percent level, respectively, at each degree of fractional integration, *d*. However, there is a unidirectional causality running from coal consumption to GDP at each degree of fractional integration, *d*. The results are similar to those reported in Lee and Chang (2005) except a bi-directional causality between GDP and oil consumption instead of a unidirectional causality and vice-versa for coal consumption.[23] Furthermore, the evidence related to the test for weak exogeneity interpreted as a test of long-run causality can explain the evidence of a mean-reverting process concerning the test for fractional cointegration, i.e. supporting a relationship between the variables in the long run but only very weakly.

In line with the results in Lee and Chang (2005) as well as Yang (2000), the results emphasize that an energy conservation policy reducing supply for production implies a negative effect on growth in Taiwan. Hence, the neutrality hypothesis of growth and energy consumption is not applicable as energy act as an engine of growth. As a result and to sustain long-term growth and development, it is necessary to increase energy supply as well as efficiency of energy usage where negative externalities on the environment of fossil fuel use implies that the latter option is preferred. However, a more efficient use of energy is not sufficient for Taiwan as a long-term solution to energy shortage. To increase energy supply, Taiwan should increase investment in energy infrastructure preferably in energy production generating low levels of CO_2 emissions and other polluting by-products, continue to promote alternative energy sources to fossil fuel combustion and improve delivery efficiency of energy.

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22 Prior to the test, test for structural breaks are performed by RLS and step-wise Chow test as well as test for trend and seasonal-dummies in the data. In contrast to Lee and Chang (2005), no outliers or structural breaks were indicated by the 1-step Chow test. The RLS and step-wise Chow test indicates within-sample coefficient constancy and parameter stability in contrast to the findings of structural breaks in Lee and Chang (2005). To conserve space, results are not reported here.

23 Although using different time-periods, the results are as well similar to those reported in Yang (2000) but contradictory to Cheng and Lai (1997).

Furthermore to reduce unnecessary wastage and pollution still not to decrease supply for production harming growth, Taiwan should put in place energy conservation policies. These strategies seek to realize the dual objectives of reducing the adverse effects of energy consumption on the environment, while avoiding the negative effect on economic growth of reducing energy consumption.

Table 5: Unit root Tests of the Transformed Series

Variable	Fractional integration, *d*	ADF (lag)
GDP	0.41	-4.17(3)*
GDP	0.42	-4.15(3)*
GDP	0.52	-4.08(3)*
GDP	0.55	-4.05(3)*
GDP	0.60	-3.99(3)*
GDP	0.67	-3.96(3)*
GDP	0.69	-3.94(3)*
GDP	0.71	-3.91(3)*
GDP	0.86	-3.79(3)*
Energy	0.60	-3.96(3)*
Energy	0.67	-4.05(7)*
Energy	0.69	-4.16(7)*
Coal	0.41	-4.09(1)*
Coal	0.42	-4.25(1)*
Coal	0.71	-3.43(1)*
Coal	0.86	-3.32(1)*
Oil	0.52	-3.58(4)*
Oil	0.55	-3.63(4)*
Oil	0.67	-3.67(4)*
Oil	0.71	-3.68(4)*

Note: Number of lags in the ADF-test in parenthesis where * denotes significance at the 5 percent level.

Table 6: Johansen's cointegration test of the Transformed Series

Model	Fractional integration, *d*	Hypothesis Null	Alternative	Trace-statistic
GDP, Energy	0.60	$r = 0$	$r >= 1$	26.34**
		$r <= 1$	$r >= 2$	2.36
GDP, Energy	0.67	$r = 0$	$r >= 1$	26.28**
		$r <= 1$	$r >= 2$	2.26
GDP, Energy	0.69	$r = 0$	$r >= 1$	26.28**
		$r <= 1$	$r >= 2$	2.25
GDP, Coal	0.41	$r = 0$	$r >= 1$	29.66**
		$r <= 1$	$r >= 2$	2.38
GDP, Coal	0.42	$r = 0$	$r >= 1$	29.35**
		$r <= 1$	$r >= 2$	2.28
GDP, Coal	0.71	$r = 0$	$r >= 1$	21.97**
		$r <= 1$	$r >= 2$	2.12
GDP, Coal	0.86	$r = 0$	$r >= 1$	19.11**
		$r <= 1$	$r >= 2$	1.29
GDP, Oil	0.52	$r = 0$	$r >= 1$	16.85**
		$r <= 1$	$r >= 2$	3.34
GDP, Oil	0.55	$r = 0$	$r >= 1$	16.99**
		$r <= 1$	$r >= 2$	3.27
GDP, Oil	0.67	$r = 0$	$r >= 1$	18.29**
		$r <= 1$	$r >= 2$	3.07
GDP, Oil	0.71	$r = 0$	$r >= 1$	18.54**
		$r <= 1$	$r >= 2$	2.95

Note: ** denotes significance at the 1 percent level.

Table 7: Test for Weak Exogeneity of the Transformed Series

Model	Fractional integration, *d*	$\alpha_{GDP} = 0$	$\alpha_{EC} = 0$
GDP, Energy	0.60	4.71(0.03)*	20.97(0.00)**
GDP, Energy	0.67	4.79(0.03)*	21.19(0.00)**
GDP, Energy	0.69	4.82(0.03)*	21.25(0.00)**
GDP, Coal	0.41	13.16(0.00)**	2.18(0.14)
GDP, Coal	0.42	13.19(0.00)**	2.03(0.16)
GDP, Coal	0.71	12.64(0.00)**	2.65(0.11)
GDP, Coal	0.86	12.25(0.00)**	3.34(0.07)
GDP, Oil	0.52	12.36(0.00)**	6.82(0.01)**
GDP, Oil	0.55	12.28(0.00)**	7.05(0.01)**
GDP, Oil	0.67	11.53(0.00)**	6.76(0.01)**
GDP, Oil	0.71	11.45(0.00)**	7.07(0.01)**

Note: EC indicates the various energy consumption variables. Numbers in parenthesis are probability values where * and ** indicates significance at the 5 percent and 1 percent level, respectively.

The mixed results in the literature analyzing the relationship between growth and energy consumption might partly be related to that many recent papers have used methodologies based on unit root tests and classical cointegration techniques, such as the Johansen and Juselius (1990) procedure. However, conventional unit root tests as well as classical cointegration procedures have some restricting assumptions about the value of d such as an $I(0)$ process of the error-correction term for the Johansen and Juselius (1990) procedure. Hence, these types of tests do not perform well in cases of fractionally integrated processes while the fractionally integrated procedures allow this possibility. This paper indicates a fractional cointegration between growth and various energy consumption variables by use of annual data for the period 1954 – 2013 for Taiwan. The general results indicate that the residuals are in the range $0.50 \leq d < 1$, i.e. a fractionally cointegrated process with a mean-reverting non-stationary long memory. Hence, the error-correction term did not follow a stationary process. These results imply a failure to achieve equilibrium for very long periods with shocks inducing non-stationary deviations from the long-run equilibrium established from cointegrating relationships. As a result in the specification of GDP and energy consumption, the GPH- and AG-procedure does suggest that the issue of fractional cointegration and long-memory structures is important and needs to be taken into account. Upon analyzing the relationship between GDP and energy consumption, a possible fractional cointegration and long-memory processes should be included in the analysis.

4. Conclusions and Policy Implications

Central to the issue of energy conservation policies is whether or not economic welfare and growth is affected by policy. Furthermore, it is generally recognized that the supply of energy plays a significant role in economic development. Among other things, it is assumed to enhance the productive use of factors of production such as labor, capital and land. In the literature, there is mixed evidences concerning the relationship and direction of causality between economic growth and energy consumption. However, the focus in the literature has mainly been on the direction of causality and its strength with little focus on the stability and order of integration of the relationship. Many studies have used conventional unit root tests and classical cointegration procedures. However, these tests have some restricting assumptions about the value of d, i.e. they do not perform well in cases of fractionally integrated processes.

Furthermore, the classical cointegration procedure, such as the Johansen and Juselius (1990) procedure, assumes that the error-correction term follows an $I(0)$ process with moving average coefficients incorporating an exponential decay. However, the estimated values of the error-correction term may differ from an $I(0)$ process in such a way that the autocorrelations decline at a slower rate than the exponential decay of the ARMA process in the classical cointegration technique. The fractionally integrated process allows this possibility. This paper adds to the literature by analyzing causality in the case of fractional cointegration between economic growth and energy consumption using annual data for Taiwan covering the period 1954 – 2013.

The result of fractional cointegration shows that $d > 0$ for all variables except for coal consumption by use of the GPH-estimator for both values of u. Furthermore with some exceptions, there is evidence of a fractionally cointegrated process with a mean-reverting non-stationary long memory, i.e. a memory-parameter for the residuals in the range $0.50 \leq d < 1$. The evidence of fractional cointegration implies a failure to achieve equilibrium for very long periods where shocks induce non-stationary deviations from the long-run relationship between economic growth and energy consumption established by the cointegrating vector. Furthermore, an estimate significantly below 1 indicates that the error-correction term is a mean-reverting process. Hence, the estimated d values are in the non-stationary but mean-reverting range. These results support a relationship between the variables in the long run but only very weakly. This behavior of the equilibrium error is of a central interest due to that without a mean-reversion of the equilibrium error, a shock to the system of variables studied will tend to permanently drive the system out of equilibrium. The test for weak exogeneity and long-run causality by the transformed series indicates a bi-directional causality between GDP and both total energy and oil consumption. For coal consumption, the causality is unidirectional to GDP. The results are similar to those reported in Lee and Chang (2005) except a bi-directional causality between GDP and oil consumption instead of a unidirectional causality and vice-versa for coal consumption. Hence, the results emphasize that an energy conservation policy implies a negative effect on growth in Taiwan. This implies that the neutrality hypothesis of growth and energy consumption is not applicable with energy acting as an engine of growth.

Furthermore, the evidence related to the test for weak exogeneity interpreted as a test of long-run causality can explain the evidence of a mean-reverting process concerning the test for fractional cointegration, i.e. supporting a relationship between the variables in the long run but only very weakly.

As a result and to sustain long-term growth and development, it is necessary to increase the energy supply as well as efficiency of energy usage where negative externalities on the environment of fossil fuel use implies that the latter option is preferred but not sufficient as a long-term solution to the energy shortage. To increase energy supply, Taiwan should increase investment in energy infrastructure preferably in energy production generating low levels of CO_2 emissions and other polluting by-products, continue to promote alternative energy sources to fossil fuel combustion and improve delivery efficiency of energy. Furthermore to reduce unnecessary wastage and pollution still not to decrease supply for production harming growth, Taiwan should put in place energy conservation policies. These strategies seek to realize the dual objectives of reducing the adverse effects of energy consumption on the environment, while avoiding the negative effect on growth of reducing energy consumption. Furthermore, the issue of fractional cointegration and long-memory structures is important and needs to be taken into account in the policy analysis of the empirical relationship between economic growth and energy consumption.

References

Agiakloglou, C., Newbold, P., Wohar, M., 1993, Bias in an estimator of the fractional difference parameter, Journal of Time Series Analysis **14**, pp. 233-246.

Akarca, A.T., Long, T.V., 1980, On the relationship between energy and GNP: a re-examination, Journal of Energy Development **5**, pp. 326-331.

Alves, D.S.C.O., Cati, R.C., Fava, V.L., 2001, Purchasing power parity in Brazil, Applied Economics **33**, pp. 1175-1185.

Andrews, D.W.K., Guggenberger, P., 2003, A biased-reduced log-periodogram estimator for the long-memory parameter, Econometrica **71**, pp. 675-712.

Arestis, P., Demetriades, P., Luintel, K., 2001, Financial development and economic growth: the role of stock markets, Journal of Money, Credit and Banking **33**, pp. 16-41.

Asafu-Adjaye, J., 2000, The relationship between energy consumption, energy prices and economic growth: time series evidence from Asian developing countries, Energy Economics **22**, pp. 615-625.

Balcilar, M., Ozdemir, Z.A. and Arslanturk, Y., 2010, Economic growth and energy consumption causal nexus viewed through a bootstrap rolling window, Energy Economics (forthcoming).

Brockwell, P.J., Davis, R.A., 1991, Time series: theory and methods, Springer Verlag, New York et al.

Cheng, B.S., 1995, An investigation of cointegration and causality between energy consumption and economic growth, Journal of Energy Development **21**, pp. 73-84.

Cheng, B.S., Lai, T.W., 1997, An investigation of co-integration and causality between energy consumption and economic activity in Taiwan, Energy Economics **19**, pp. 435-444.

Cheung, Y., Lai, K., 1993, A fractional cointegration analysis of purchasing power parity, Journal of Business and Economic Statistics **11**, pp. 103-112.

Chow, G.C., 1960, Tests of equality between sets of coefficients in two linear regressions, Econometrica **28(3)**, pp. 591-605.

Dickey, D.A., Fuller, W.A., 1981, Distribution of the estimators for autoregressive time series with a unit root, Econometrica **49(4)**, pp. 1057-1072.

Diebold, F.X., Rudebusch, G.D., 1991, Is consumption too smooth? Long memory and the Deaton paradox, The Review of Economics and Statistics **1**, pp. 1-9.

Geweke, J., Porter-Hudak, S., 1983, The estimation and application of long memory in time series models, Journal of Time Series Analysis **4**, pp. 221-238.

Ghali, K.H., El-Sakka, M.I.T., 2004, Energy use and output growth in Canada: a multivariate conintegration analysis, Energy Economics **26**, pp. 225-238.

Gil-Alana, L.A., Toro, J., 2002, Estimation and testing of ARFIMA models in the real exchange rate, International Journal of Finance and Economics **7**, pp. 279-292.

Glasure, Y.U., Lee, A.R., 1997, Cointegration, error-correction and the relationship between GDP and energy: The case of South Korea and Singapore, Resources and Energy Economics **20**, pp. 17-25.

Granger, C.W.J., Joyeux, R., 1980, An introduction to long memory time series models and fractional differencing, Journal of Time Series Analysis **1**, pp. 15-39.

Hall, S.G., Alistair, M., 1994, The relevance of p-star analysis to UK monetary policy, Economic Journal **104**, pp. 597-604.

Hosking, J. R. M., 1981, Fractional differencing, Biometrika, **68**, pp. 165-176.

Johansen, S., Juselius, K., 1990, Maximum likelihood estimation and inference on co-integration – with applications to the demand for money, Oxford Bulletin of Economics and Statistics **52(2)**, pp. 169-210.

Johansen, S., Juselius, K., 1992, Some structural hypotheses in a multivariate cointegration analysis of purchasing power parity and uncovered interest parity for the UK, Journal of Econometrics **53**, pp. 211-244.

Jumbe, C.B.L., 2004, Cointegration and causality between electricity consumption and GDP: empirical evidence from Malawi, Energy Economics **26**, pp. 61-68.

Kraft, J. Kraft, A., 1978, On the relationship between energy and GDP, Journal of Energy Development **3**, pp. 401-403.

Lee, C.C., 2005, Energy consumption and GDP in developing countries: a cointegrated panel analysis, Energy Economics **27**(3), pp. 415-427.

Lee, C-C., Chang, C-P., 2005, Structural breaks, energy consumption, and economic growth revisted: Evidence from Taiwan, Energy Economics **27**, pp. 857-872.

Lise, W. and van Montfort, K., 2007, Energy consumption and GDP in Turkey: Is there a co-integrating relationship?, Energy Economics **29(6)**, 1166-1178.

Masih, A.M.M., Masih, R., 1996, Energy consumption, real income and temporal causality: results from a multi-country study based on cointegration and error-correction modelling techniques, Energy Economics **18**, pp. 165-183.

Mishra, V., Smyth, R. and Sharma, S., 2009, The energy-GDP nexus: Evidence from a panel of Pacific Island countries, Resource and Energy Economics **31(3)**, 210-220.

Oh, W. and Lee, K., 2004a, Causal relationship between energy consumption and GDP revisited: the case of Korea 1970 – 1999, Energy Economics **26(1)**, 51-59.

Oh, W. and Lee, K., 2004b, Energy consumption and economic growth in Korea: testing the causality relation, Journal of Policy Modeling **26(8-9)**, 973-981.

Osterwald-Lenum, M., 1992, A note with quantils of the asymptotic distribution of the maximum likelihood cointegration rank test statistics, Oxford Bulletin of Economics and Statistics, **54(3)**, 461-71.

Ozturk, I., 2010, A literature survey on energy – growth nexus, Energy Policy **38(1)**, 340-349.

Ozturk, I., Aslan, A. and Kalyoncu, H., 2010, Energy consumption and economic growth relationship: Evidence from panel data for low and middle income countries, Energy Policy **38(8)**, 4422-4428.

Phillips, P.C.B. and Perron, P., 1988, Testing for a unit root in time series regression, Biometrika, **75**, 335-346.

Rosenblatt, M., 1956, A central limit theorem and a strong mixing condition, Proceedings of the National Academy of Sciences **42**, pp. 43-47.

Siokis, F., Christodoulou, C., 2004, Long memory and persistence in dollar-based real exchange rates, International Journal of Theoretical and Applied Finance **7**, pp. 31-44.

Soytas, U., Sari, R., 2003, Energy consumption and GDP: causality relationship in G-7 countries and emerging markets, Energy Economics **25**, pp. 33-37.

Sowell, F., 1992, Maximum likelihood estimation of stationary univariate fractionally integrated time series models, Journal of Econometrics **53**, pp. 165-168.

Stern, D.I., 1993, Energy use and economic growth in the USA, a multivariate approach, Energy Economics **15**, pp. 137-150.

Stern, D.I., 2000, A multivariate cointegration analysis of the role of energy in the US macroeconomy, Energy Economics **22**, pp. 267-283.

Tolvi, J., 2003, Long memory and outliers in stock market returns, Applied Financial Economics **13**, pp. 495-502.

Villeneuve, J-F., Handa, J., 2006, Purchasing power parity as a long-term memory process: evidence from Canada, Applied Financial Economics **16**, pp. 109-117.

Warr, B.S. and Ayres, R.U., 2010, Evidence of causality between the quantity and quality of energy consumption and economic growth, Energy **35(4)**, 1688-1693.

Wolde-Rufael, Y., 2004, Disaggregated industrial energy consumption and GDP: the case of Shanghai, 1952-1999, Energy Economics **26**, pp. 69-75.

Yang, H.Y., 2000, A note on the causal relationship between energy and GDP in Taiwan, Energy Economics **22**, pp. 309-317.

Yu, E.S.H., Choi, J.Y., 1985, The causal relationship between energy and GDP: an international comparison, Journal of Energy Development **10**, pp. 249-272.

Zivot, E. and Andrews, D., 1992, Further evidence on the great crash, the oil price shock, and the unit root hypothesis, Journal of Business and Economic Statistics, **10**, 936-954.