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ENERGY-GROWTH CAUSALITY: A PANEL ANALYSIS

Abstract

Understanding the impact of energy consumption on economic growth is an important consideration in the formulation of both energy and environmental policies. Motivated by this development, this paper empirically re-examines the direction of causality and the sign (in the panel sense) between energy consumption (EC) and the gross-domestic product (GDP) for seventeen selected Asian countries. Results reveal long-run stable equilibriums in these countries, while the EC brings about a positive impact on GDP. Causality runs from EC to GDP in the short-run, while the long-run causal linkage exists from GDP to EC. This indicates that energy is a force for economic growth in the short-run, but in the long-run, the EC is fundamentally driven by economic growth. Efficient coordination and cooperation towards the implementation of energy conservation policies to support sustainable economic development should be in the regional agenda.

Key words:

Energy consumption, Panel analysis, Economic growth.

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1. Introduction

Energy consumption has steadily increased over the past few decades in Asian countries due to the population increment and industrial expansion[1]. Energy consumption is expected to increase to 159.3 quadrillion BTU in 2015, 187.8 quadrillion BTU in 2020, 217.0 quadrillion BTU in 2025, 246.9 quadrillion BTU in 2030 and 277.3 quadrillion BTU in 2035. The average annual percentage change from 2007 to 2035 in Asia is 2.8 percent, which is higher than other regions, such as the Middle Eastern countries (2.2 percent), Central and South America (1.8 percent), and Africa (1.8 percent) (EIA, (2010; Table 1, pp. 9).

The major users of energy were China and India, who continue to lead the world in relation to economic growth and energy demand growth. Together, China and India accounted for about 10 percent of the world's total energy consumption in 1990 and 20 percent in 2007 (EIA, 2010). China and India's other significant increases include a fast-paced growth in population, rapid economic growth and industrial expansion into other areas of the Asian region.

The episodic energy crisis, coupled with depleting energy sources, environmental costs and high-energy consumption, has forced governments around the globe to more intently monitor and manage energy markets (ECSSR, 2004). Growing concerns had attracted the interest of the government in Asian countries. These measures include cooperation for energy conservation and the efficient usage of energy policies. In this context, the long-run relationship between energy consumption and economic growth has been a lively topic of empirical assessment. In the energy economics literature, the direction of causality as to whether the adoption of energy savings inhibits or stimulates economic growth has been a much debated matter[2].

Understanding the impact and causality patterns of energy consumption on economic growth is an important consideration in the formulation of both energy and environmental policies. Accordingly, Squalli (2007, pp. 1193–1194), Payne (2010a, pp. 54–55) and Ozturk (2010, pp. 340–341) provide excellent de-

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[1] According to EIA (2008), Asia's total primary energy consumption in 1990 was 47.4 quadrillion British Thermal Units (BTU). This number doubled to about 127.1 quadrillion BTU in 2007

[2] Literature about the energy-growth causality was coined from the seminal work of Kraft and Kraft (1978). Since then, impressive volumes of papers were dedicated to this genre. Ozturk (2010) and Payne (2010a) conducted an excellent survey, while Payne (2010b) and Narayan *et al.* (2010) investigated the electrical consumption and growth literature. On the Asian side, Yu and Choi (1985), Masih and Masih (1996, 1998), Asafu-Adjaye (2000) Soytas and Sari (2003) and Lee and Chang (2008) were among the champions.

scriptions of four hypotheses related to the relationship between energy consumption and economic growth. Briefly, the four patterns include: (1) The «growth» hypothesis, where the causality runs from energy consumption to growth. This pattern exists in energy dependent countries (Yu and Choi, 1985 for the Philippines, Masih and Masih, 1996 for India, Asafu-Adjaye, 2000 for India and Indonesia, Soytas and Sari, 2003 for Turkey, France, Japan and Germany, Lee, 2005 for a panel of eighteen developing countries and Tsani, 2010 for Greece); (2) The «conservation» hypothesis, where GDP Granger-causes energy consumption (Kraft and Kraft (1978) for the United States (US), Abosedra and Baghestani (1989) for the US, Cheng and Lai (1997) for Taiwan, Cheng (1999) for India, Ang (2008) for Malaysia and Zhang and Cheng (2009) for China). For this purpose, policies such as the reduction in greenhouse emissions designed to reduce energy consumption and waste may not adversely affect real GDP. (3) The «neutrality» hypothesis views the absence of Granger-causality between energy consumption with GDP (Yu and Hwang (1984); Altinay and Karagol (2004); Halicioglu (2009) and Payne (2010a). (4) The «feedback» hypothesis suggests that energy consumption and GDP are interdependent and support the existence of bi-directional causality (Hwang and Gum, 1991; Yang, 2000; Oh and Lee, 2004; Climent and Pardo, 2007; Apergis and Payne, 2009 and Ozturk and Acaravci 2010). The literature has not come to a general agreement on the nature of causal relationships between energy consumption and economic growth. In this context, policies aiming at the gradual curtailing of energy need to consider the potential causal linkages between economic growth and energy consumption.

Motivated by this development, the goal of this study is to empirically re-examine the direction of causality and sign (in the panel sense) between energy consumption (EC) and real GDP for seventeen Asian countries. Once the causality is ascertained, appropriate energy development policies in these countries can be adopted. As such, the structure of the rest of this paper is as follows. A brief and intuitive account of the econometric methodology employed is provided in Section 2, before discussing the results in detail in Section 3. Some policy implications and conclusions are made in Section 4.

2. Econometric Modeling

2.1. Panel Unit Root and Stationary Tests

The first step in the estimation of dynamic panels is to test whether the variables at hand contain unit roots. Studies that have used joint panel unit root tests include Maddala and Wu (1999, MW), Hadri (2000, HADRI), Levin et al., (2002, LLC) and Im *et al.* (2003, IPS). The null hypothesis in all joint panel unit root tests, except the HADRI test, is that the panel series has a unit root (non-stationary). Unlike the augmented Dickey Fuller (ADF) test, the HADRI test is similar to the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS – based LM) statistic, which has a null hypothesis of level (trend) stationarity and an alternative of difference stationarity in the panel. A comparison of the results obtained from the wide range of panel unit root tests can provide some insight into the stationarity properties of the data. If both procedures fail to reject the null hypothesis (or if both reject), we have mixed results and can only conclude that the data are not informative enough. On the other hand, if an ADF type panel unit root test rejects the null and the KPSS type test fails to reject it, we have greater confidence that the series under consideration is in fact stationary. As these panel-based unit root tests are becoming common in the literature, interested readers may refer to their original articles for a more comprehensive discussion.

2.2. Panel Cointegration

We then proceed to examine whether there exists any long-run equilibrium relationship between the variables under investigation. We resort to Pedroni (1999, 2001, 2004) and Kao (1999) panel cointegration tests. Pedroni considers seven different statistics, four of which are based on pooling the residuals of the regression along the within-dimension (panel test) of the panel. The other three are based on pooling the residuals of the regression along the between-dimension (group test) of the panel. The within-dimension tests take into account common time factors and allow for heterogeneity across countries. The between-dimension tests are the group-mean cointegration tests, which allow for the heterogeneity of parameters across countries.

Kao (1999) proposed **Dickey Fuller (**DF) and ADF-type tests for ε_{it}, where the null is specified as no cointegration. In this study, we only report the ADF-type test. The details of these tests are discussed in Appendix 1.

2.3. Panel Fully Modified OLS (FMOLS) Estimates

To obtain the long-run estimates for the cointegrating relationship (the coefficients of EC), we adopt the panel group mean Fully Modified OLS (FMOLS), following the work by Pedroni (2000). The FMOLS procedure accommodates the heterogeneity that is typically present, both in the transitional serial correlation dynamics, and in the long-run cointegrating relationships. The FMOLS estimator is described in detail in Appendix 1.

2.4. Granger Causality Tests

To test for panel causality, we estimate a panel based vector error correction model (VECM) with a dynamic error correction term based on the analysis in Holtz-Eakin *et al.* (1988, 1989). The empirical models are as follows:

$$\Delta GDP_{it} = \pi_{1j} + \sum_{p=1}^{m} \pi_{11ip} \Delta GDP_{it-p} + \sum_{p=1}^{m} \pi_{12ip} \Delta EC_{it-p} + \mu_{1i} ECT_{it-1} + \zeta_{1it} \quad (1a)$$

$$\Delta EC_{it} = \pi_{2j} + \sum_{p=1}^{m} \pi_{21ip} \Delta EC_{it-p} + \sum_{p=1}^{m} \pi_{22ip} \Delta GDP_{it-p} + \mu_{2i} ECT_{it-1} + \zeta_{2it} \quad (1b)$$

where: Δ is the lag operator and p denotes the lag length. The specification in Equation 1 allows for testing the causality direction. For example, in the short-run, the EC does not Granger cause GDP where $H_0 : \pi_{12ip} = 0$ for all i and p, while $\mu_{1i} = 0$ in Equation (1a)[3]. The rejection implies that EC $\longrightarrow$ GDP, supporting the growth hypothesis. Similar analogous restrictions and testing procedures can be applied in testing the hypothesis that GDP does not Granger cause movement in EC, where the null hypothesis H_0: $\pi_{22ip} = 0$ for all i and p, while $\mu_{2i} = 0$ in Equation (1b).

3. Empirical Results

3.1. Data sources

Annual data from 1980 to 2006 for the 17 Asian countries were utilized for the study[4]. Per capita total primary energy consumption (EC) data were obtained from the International Energy Annual 2006 of Energy Information Administration (EIA). Real GDP data were obtained from the World Development Indicators (WDI) 2008 of the World Bank. All variables were transformed into the logarithmic form.

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[3] The F-test or Wald χ^2 of the explanatory variables (in the first differences) indicates the short-run causal effects ($\pi_{12ip} = 0$ for all i and p), while the long-run causal (μ_{1i} =0) relationship is implied through the significance of the lagged ECT, which contains the long-run information.

[4] The Asian countries included Bangladesh, Bhutan, Brunei Darussalam, China, Hong Kong, India, Indonesia, Japan, Korea, Malaysia, Maldives, Nepal, Pakistan, Philippines, Singapore, Sri Lanka, and Thailand.

3.2. Panel Unit Root and Stationary Results

The results, made available upon request, illustrate that the series of the variables are of an *I(1)* process, as the pooled data are stationary in their first differences. These results enable us to test the cointegration among EC and GDP.

3.3. Panel Cointegration Results

From the panel cointegration results in Table 1, we find strong evidence to reject the null hypothesis of no cointegration for all seven statistics provided by Pedroni (1999, 2001, 2004). Similarly, we reject the null hypothesis of no cointegration using the ADF-type statistics from the Kao (1999) panel cointegration tests, suggesting that that the two-dimensional model for the Asian countries is cointegrated and moves together in the long-run. Thus, we find that GDP and EC are cointegrated in the multi-country panel setting for the sample period.

3.4. Panel FMOLS Estimates

Having established cointegration in the long-run, we estimate the long-run parameters of the model by using the FMOLS technique. The FMOLS corrects the standard OLS for bias induced by the endogeneity and serial correlation of the regressors (Lee, 2005). The elasticity of energy consumption is important for understanding the past and assessing future economic dynamics. It represents the weights with which the marginal relative changes of the energy consumption contributes to the relative change of output (Lee *et al.,* 2008).

Table 2 reports the results of the long-run estimates for seventeen Asian countries and the panel estimates based on Pedroni's group mean FMOLS estimator. The panel results of the regression equation with GDP as the dependent variable illustrate that the coefficient of the EC is positive and statistically significant at the 5 percent significance level. A one percent increase in energy consumption leads to a 0.21 percent increase in GDP for these seventeen Asian countries. This positive coefficient on EC implies that more energy results in greater outputs, as suggested by Lee, (2005), Narayan and Smyth (2008), Lee and Chang, (2008) and Ozturk, (2010).

Turning to the country specific evidence, the results also indicate a positive and significant relationship between EC and GDP for all countries. The elasticity estimates range from 0.10 (Hong Kong) to 0.94 (Philippines). The results suggest that the EC contributes most to the Philippines' output, whereas it contributes least to Hong Kong's output. Having inelastic coefficients on EC suggests that the vulnerability of energy prices would not have a significant impact on the consumption patterns in these countries, as it would be considered necessities for the society as a whole.

3.5. Panel Granger Causality Test Results

Once the long-run estimates have been determined, we turn to the causality linkages. The empirical results presented in Table 3 illustrate that the coefficient of the error correction term (ECT) is not statistically significant in the GDP equation, indicating the absence of a long-run causality relationship running from EC to GDP. However, we note the existence of a significant short-run causal relationship running from EC to GDP, since the estimated coefficients of the explanatory variables are statistically significant. The short-run results are supported by Asafu-Adjaye (2000), Soytas and Sari (2003), Lee (2005), Narayan and Smyth (2008) and Tsani (2010), who established evidence of a short-run Granger causality running from EC to GDP.

On the other hand, we find evidence of the existence of a long-run relationship running from GDP to EC, in which the coefficient of the error correction term (ECT) is statistically significant in the EC equation. This result illustrates that energy consumption is determined by economic growth; supporting the conservation hypothesis. This pattern is similar to results from developing countries (Cheng and Lai, 1997; Cheng, 1999; Mahadevan and Asafu-Adjaye, 2007; Ang, 2008 and Ozturk, 2010).

4. Concluding remarks

Using panel estimation for seventeen Asian countries, this paper empirically examines the relationship between energy consumption and the gross domestic product (GDP). We find that the variables were in a stationary fashion in their first differences or were in an $I(1)$ process. The panel cointegration results reveal a long-run equilibrium relationship among the two variables. The results of the FMOLS show that the energy consumption variable has a positive sign. This indicates that an increase in GDP would lead to a greater use of energy. From the Granger causality test, there is a short-run unidirectional causal relationship running from energy consumption to GDP. This implies that in the short-run, energy consumption leads to economic growth, since the economies in these 17 Asian countries are energy-dependent economies. Additionally, in the long-run, GDP Granger causes energy consumption for the panel. This provides additional evidence in support of the proposition that energy consumption is a result of economic activity, rather than being an essential input to production.

In the short-run, the implementation of energy conservation policies might lead to a significant, but temporary, negative impact on economic growth in these Asian countries. However, economic development in the Asian countries is less dependent on energy in the long-run. Cooperation for energy conservation policies among the Asian countries would be an imperative move that would not harm GDP. Proactive agendas of research and development on renewable technologies in response to depleting supplies of energy sources would be another avenue that could be used to improve energy transportation facilities and infrastructure development to improve delivery efficiency. Niu *et al.* (2011)[5] argued that developing countries may benefit from their developed nations counterparts, where they may fetch advanced technology and capital to facilitate efficient energy use, while reducing energy consumption and carbon emissions. Efforts have also been made in pursuit of more environmentally-friendly and resource-saving societies to promote energy efficiency in the face of concern about the effects of global warming for the Asian region (Chang, 2010; Lean and Smyth, 2010 and Li *et al.*, 2011). With the recent experience of unprecedented high levels of energy prices, depleting energy sources and international initiatives such as Kyoto protocol, the commitment needs to be established to facilitate successful energy conservation policies.

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[5] Accordingly, as per Table 2 (pp. 2123) in the paper, four developed countries (Australia, New Zealand, Japan and South Korea) have the technological advancement therefore, their energy efficiency is higher comparatively to the four developing countries (China, India, Indonesia and Thailand).

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Appendix 1

Panel Cointegration and Fully Modified OLS

Pedroni panel cointegration test

There are in all seven panel cointegration tests. Detailed description of the formulae for the seven panel cointegration statistics, are given in Pedroni (1999: 660–661).

A. Within-dimension (panel tests):

a) Panel ν-Statistic

b) Panel Phillip-Perron (PP) type ρ -Statistics

c) Panel Phillips-Perron (PP) t -Statistic (non-parametric)

d) Panel Augmented Dickey Fuller (ADF) t -Statistic (parametric)

B. Between-dimension (group tests):

e) Group Phillip-Perron (PP) type ρ -Statistics

f) Group Phillips-Perron (PP) t -Statistic (non-parametric)

g) Group Augmented Dickey Fuller (ADF) t -Statistic (parametric)

These seven statistics are based on the estimated panel cointegration regression residuals of the likely cointegrating vector

$$GDP_{i,t} = \alpha_i + \phi_i t + \beta_1 EC_{i,t} + \varepsilon_{i,t} \qquad \text{(A.1)}$$

varying across countries, thus permitting full heterogeneity (β_i), fixed effects (α_i) and individual specific deterministic trends ($\phi_i t$) across individual members of the panel

Pedroni (1999) shows that under appropriate standardization based on the moments of vector of Brownian motion function, each of these statistics converges weakly to a standard normal distribution when both the T and N of the panel grow large. The standardized distributions for the above mentioned seven panel and group statistics can be expressed in the form of

$$\frac{e_{N,T} - \mu\sqrt{N}}{\sqrt{\nu}} \Rightarrow N(0,1) \qquad \text{(A.2)}$$

where e_{NT} is the respective panel/group cointegration statistic and μ and ν are the expected mean and variance of the corresponding statistics. They are computed by Monte Carlo stochastic simulations and tabulated in Pedroni (1999, Table 2).

Kao panel cointegration test

Unlike Pedroni test, Kao (1999) test specifies cross-section specific intercepts and homogeneous coefficients on the first-stage regressors. In this case, we specified the panel regression model as

$$y_{it} = x_{it}^{'}\beta + z_{it}^{'}\gamma + \varepsilon_{it} \tag{A.3}$$

where y_{it} and x_{it} are I(1) and non cointegrated. For $z_{it} = \{\mu_i\}$ Kao (1999) proposed DF and ADF-type unit root tests for ε_{it} where the null is specified as no cointegration.

The DF-type test can be calculated from this regression of:

$$\hat{\varepsilon}_{it} = \rho\hat{\varepsilon}_{it-1} + v_{it} \tag{A.4}$$

while the augmented version of the pooled specification:

$$\hat{\varepsilon}_{it} = \rho\hat{\varepsilon}_{it-1} + \sum_{j=1}^{p}\varphi_j\Delta\hat{\varepsilon}_{it-j} + v_{itp} \tag{A.5}$$

where $\hat{\varepsilon}_{it} = \tilde{y}_{it} - \tilde{x}_{it}\hat{\beta}$ and $\tilde{y} = y_{it} - \bar{y}_i$. The OLS estimate of ρ and the t-statistics are given as

$$\hat{\rho} = \frac{\sum_{i=1}^{N}\sum_{t=2}^{T}\hat{\varepsilon}_{it}\hat{\varepsilon}_{it-1}}{\sum_{i=1}^{N}\sum_{t=2}^{T}\hat{\varepsilon}_{it}^{2}} \text{ and } t_{\rho} = \frac{(\hat{\rho}-1)\sqrt{\sum_{i=1}^{N}\sum_{t=2}^{T}\hat{\varepsilon}_{it-1}^{2}}}{s_{\varepsilon}}.$$

In this case, $s_{\varepsilon}^{2} = \frac{1}{NT}\sum_{i=1}^{N}\sum_{t=2}^{T}(\hat{\varepsilon}_{it} - \hat{\rho}\hat{\varepsilon}_{it-1})^{2}$. Under the null of no cointegration, Kao (1999) shows that following the statistics:

$$DF_{\rho} = \frac{\sqrt{NT}(\hat{\rho}-1) + 3\sqrt{N}}{\sqrt{10.2}} \tag{A.6}$$

$$DF_{t} = \sqrt{1.25}t_{p} + \sqrt{1.875N} \tag{A.7}$$

$$DF_{\rho}^{*} = \frac{\sqrt{NT}(\hat{\rho}-1)\frac{3\sqrt{N}\hat{\sigma}_{v}}{\hat{\sigma}_{0v}^{2}}}{\sqrt{3 + \frac{36\hat{\sigma}_{v}^{4}}{5\hat{\sigma}_{0v}^{4}}}} \tag{A.8}$$

$$DF_t^* = \frac{t_\rho + \frac{\sqrt{6N}\hat{\sigma}_v}{2\hat{\sigma}_{0v}}}{\sqrt{\frac{\hat{\sigma}_{0v}^2}{2\hat{\sigma}_v^2} + \frac{3\hat{\sigma}_v^2}{10\hat{\sigma}_{0v}^2}}} \tag{A.9}$$

where $\hat{\sigma}_v^2 = \hat{\Sigma}_{yy} - \hat{\Sigma}_{yx}\hat{\Sigma}_{xx}^{-1}$ and $\hat{\sigma}_{0v}^2 = \hat{\Omega}_{yy} - \hat{\Omega}_{yx}\hat{\Omega}_{xx}^{-1}$. For ADF can be constructed as:

$$ADF = \frac{t_{ADF} + \frac{\sqrt{6N}\hat{\sigma}_v}{2\hat{\sigma}_{0v}}}{\sqrt{\frac{\hat{\sigma}_{0v}^2}{2\hat{\sigma}_v^2} + \frac{3\hat{\sigma}_v^2}{10\hat{\sigma}_{0v}^2}}} \tag{A.10}$$

where t_{ADF} is the t-statistics of ρ in equation A.5.

Fully Modified OLS Estimates

Following Pedroni (2000, 2001), we consider the following cointegrated system for panel data of

$$Y_{it} = \alpha_i + \beta_i X_{it} + \mu_{it} \tag{A.11}$$

$$X_{it} = X_{i,t-1} + e_{it} \tag{A.12}$$

where, $i = 1,2,...,N$ countries over the time period of $t = 1,2,...M$. In addition, $Z_{it} = (Y_{it}, X_{it})' \sim I(1)$ and $\zeta_{it} = (\mu_{it}, e_{it})' \sim I(0)$ with covariance matrix of $\Omega_i = \Omega_i^0 + \Gamma_i + \Gamma_i^{'}$, where Ω_0^i is the contemporaneous covariance, Γ_i is the weighted sum of autocovariances while $\Omega_i = L_i L_i'$ in which L_i is the lower triangular decomposition of Ω_i. For simplicity, we assume that Y = GDP while X [EC] of A.1 in this study. The panel FMOLS estimator for coefficient β is given as:

$$\beta_{FM}^* = N^{-1}\sum_{i=1}^{N}\left(\sum_{t=1}^{T}(X_{it} - \overline{X}_{it})^2\right)^{-1}\left(\sum_{t=1}^{T}(X_{it} - \overline{X}_{it})Y_{it}^* - T\hat{\gamma}_i\right) \tag{A.13}$$

where

$$Y_{it}^* = \left(Y_{it} - \overline{Y}\right) - \frac{\hat{L}_{21i}}{\hat{L}_{22i}}\Delta X_{it} \text{ and } \hat{\gamma}_i = \hat{\Gamma}_{21i} + \hat{\Omega}_{21i}^0 - \frac{\hat{L}_{21i}}{\hat{L}_{22i}}\left(\hat{\Gamma}_{22i} + \hat{\Omega}_{22i}^0\right).$$

Likewise, the associated t-statistics for the estimator can be constructed as:

$$t_{\hat{\beta}_{FM}^*} = N^{-1/2}\sum_{i=1}^{N} t_{\hat{\beta}_{FM,i}^*} \text{ where } t_{\hat{\beta}_{FM,i}^*} = (\hat{\beta}_{FM,i}^* - \beta_0)\left(\hat{\Omega}_{11i}^{-1}\sum_{t=1}^{T}(X_{it} - X_i)^2\right)^{1/2}.$$