Market Reforms in Energy Sector: Evidence from a Panel Data Based Cointegration Analysis of BRICS Countries

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Market Reforms in Energy Sector Evidence from a Panel Data based Cointegration Analysis of BRICS Countries

Aviral Kumar Tiwari, Bharti Pandey and A.P. Tiwari

1. Introduction

Recently, air pollution and global climate change issues caused by greenhouse gases have become the focus of international attention. The Inter-governmental Panel on Climate Change (IPCC) (2007) and the Stern et al. (2006) report both demonstrate that the most important environmental problem of our age is global warming. CO2 is considered to be the primary greenhouse gas responsible for global warming, and its regulation has become an important inter-governmental issue (Talukdar and Meisner 2001). Stern et al. (2006) pointed out that if no action is taken to reduce GHG emissions, the concentration of GHGs in the atmosphere could double as early as 2035 from its pre-industrial level. This implies that in the short run, global average temperature may rise by over 2°C. In the longer term, there is a greater than a 50 per cent chance that the rise in temperature would exceed 5°C. Stern et al. (2006) emphasise that this radical change in temperatures would affect all countries. Among them, the earliest and the hardest hit would be the poorest and populous nations, even though they contributed least to GHG emissions. The objective of the 1997 Kyoto protocol was to reduce greenhouse gases (GHG), which cause climate change, and it demanded a reduction of GHG emissions to 5.2 per cent lower than the 1990 level during the period from 2008 to 2012. This came into force in 2005. Though BRIC countries (Brazil, Russian Federation, India, and China) signed the Kyoto protocol to curb emission levels, there are still environmental concerns given the region’s recent economic growth. Goldman Sachs (2003) argues that BRIC economies could become a much larger force in the world economy than the G6 (United States, Japan, Germany, France, Italy, and the United Kingdom) in less than 40 years and by 2025 could account for over half the size of the G6. However, during the last few years, these economies have experienced profound structural changes that continue to influence the evolution of regional CO2 output, with potentially adverse consequences for global mitigation strategies (Tamazian, Chousa, and Vadamannati, 2009). According to World Bank (2007) in 1990, Brazil’s CO2 emissions represented 0.94 per cent of the world’s

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total, the Russian Federation’s was 3.80 per cent, India’s was 3.00 per cent and China’s was 11.00 per cent. By 2007, BRIC countries’ emissions increased; Brazil’s was 1.15 per cent of the world’s total, Russia’s was 6 per cent, India’s was 5.00 per cent and China’s was 16.00 per cent. According to the figures by the U.N. Climate Change Secretariat, the top five sources of greenhouse gases worldwide were the United States, China, Russia, India, and Japan. According to World Resources Institute estimates (2000), Brazil is the eighth largest emitter of greenhouse gases and the third largest emitter in the developing world after China and India. However, the combustion of fossil fuels is the largest single contributor to CO₂ emissions and total GHG emissions and of all the major sources has grown most rapidly from 1970 to the present.

With this backdrop, present study attempts to analyze stationarity and cointegration property of CO₂ emissions, primary energy consumption and GDP i.e., economic growth for BRICS economies. For analysis, we used a battery of unit root and cointegration test. Our results indicated that all the three variables analysed are integrated of order one i.e., I (1). Further, we find evidence of two cointegration relations in our data set of BRICS economies.

The remainder of this paper is organised in the following fashion. Section 2 presents a brief review of the economic, energy use and environmental pollution profile of the countries. Section 3 outlines the model and estimation methodology, and the econometric results are presented and discussed in the fourth section. Some policy implications and conclusions are provided in the final section.

2. Literature Review

There have been two parallel literatures on the relationship between economic growth and environmental pollution. The first set of studies has focused on the economic growth-environmental pollutants nexus and closely allied to testing the Environmental Kuznets Curve (EKC) hypothesis. The EKC hypothesis states that as income increases, emissions increase as well until some threshold level of income is reached after which emissions begin to decline. The EKC hypothesis specifies emissions as a function of income, which presumes unidirectional causality runs from economic growth to emissions. However, it is conceivable that causation could run from emissions to economic growth whereby emissions occur in the production process and, as a consequence, income increases. Considering this, some studies have examined the direction of Granger-causality between economic growth and environmental pollution (Coondoo and Dinda, 2002; Dinda and Coondoo, 2006; Akbostanci et al., 2009; Lee and Lee, 2009).*

A second set of studies on the relationship between economic growth and environmental pollution has focused on the economic growth-energy consumption nexus, as emissions are primarily generated by burning fossil fuels. Since the seminal work of Kraft and Kraft

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* Comprehensive review of literature on the relationship between energy consumption/electricity consumption and economic growth/employment is available in Tiwari (2011a, 2011b and 2011c) and references therein one may refer that.
(1978), many studies have investigated the causal relationship between energy consumption and economic growth. Griffin and Gregory (1976), Berndt and Wood (1979), and Berndt (1980, 1990) have emphasized on the substitutability or complementarity between energy and the factors of production and its interplay with technical progress and productivity within a neoclassical theory of economic growth. Bergman (1988), Jorgenson and Wilcoxen (1993), Kemfert and Welsch (2000), and Smulders and de-Nooij (2003), among others, have explored the role of energy within a general equilibrium framework.

The causal relationship among the endogenous variables can be analysed through two approaches. First is multivariate approach and second is bivariate approach. After Stern (1993) who used four-variable vector autoregressive (VAR) model- a multivariate model for the USA in the post-war period, Masih and Masih (1997, 1998), Asafu-Adjaye (2000), Stern (2000), Oh and Lee (2004), and Narayan and Smyth (2005), among others too employ the multivariate model. These studies investigated the relationship between GDP and energy within a production function model; hence, a multivariate model naturally includes GDP, energy, labour and/or capital, as well as technological change.

On the other hand, several studies used a bivariate model in detecting the causality between GDP and energy. For example, Ghosh (2002), Soytas and Sari (2003), Shiu and Lam (2004), and Yoo (2005) among others have focused just on the direction of causality.

We can classify the studies to date into four groups on the basis of their findings. First, a large number of studies found unidirectional causality running from electricity or energy consumption (both aggregate and disaggregate level) to GDP. Studies worthy of mention include those by Altinay and Karagöl (2005) for Turkey, which found strong evidence for the period 1950-2000, Lee and Chang (2005) for Taiwan for the period 1954-2003, Shiu and Lam (2004) for China for the period 1971-2000, and Soytas and Sari (2003) for Turkey, France, Germany and Japan, Wolde-Rufael (2004) for Shanghai for the period 1952-1999, Morimoto and Hope (2004) for Sri-Lanka for the period 1960-98 and Tiwari (2011d, 2011e) for India for the period 1965-2009.


A combination of these two literatures whereby the relationship between economic growth, energy consumption and pollution emissions are considered within a Granger-causality multivariate framework is a relatively new area of research. There are a limited number of studies in this direction either for developed countries, such as Ang (2007) for France and Soytas et al. (2007) for United States, or developing countries, such as Zhang and Cheng (2009) for China, Ang (2008) for Malaysia and Halicioglu (2009) and Soytas and Sari (2009) for Turkey, and Sari and Soytas (2009) for oil-rich OPEC countries.

3. Data and Methodology

In the study, we obtained data of GDP of BRICS countries from World Development Indicators and data of CO₂ emissions and primary energy consumption were sourced from BP Statistical Review of World Energy (2010). We used annual data of all variables for the period 1985-2009 and period is limited by the availability of the data.

3.1 Tests of Unit Root

Since the use of nonstationary data can produce spurious regression, therefore, it is first necessity to ensure that the panel data series are stationary. To test the stationary property of the data we have used a battery of unit root tests namely LLC (Levin, Lin, and Chu 2002), IPS (Im, Pesaran and Shin 2003), Breitung’s (2000) and MW test (Maddala and Wu 1999).

LLC (2002) developed a number of pooled panel unit root tests with various specifications depending upon the treatment of the individual specific intercepts and time trends. The test imposes homogeneity on the autoregressive coefficient that indicates the presence or absence of a unit root while the intercept and the trend can vary across individual series. LLC (2002) procedure follows Augmented Dickey Fuller (ADF) regression for the investigation of unit root hypothesis as given below step by step:

First, implement a separate ADF regression for each province

\[ \Delta y_{i,t} = \alpha_i + \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \alpha_{i,j} \Delta y_{i,t-j} + \epsilon_{i,t} \]  \hfill (1)

The lag order \( i p \) is allowable to differ across individual provinces. The appropriate lag length is chosen by allowing the maximum lag order and then using the t-statistics for \( ij \) to determine if a smaller lag order is preferred.

Second, run two separate regressions and save the residuals

\[ \Delta y_{i,t} = \lambda_i + \sum_{j=1}^{p_i} \gamma_{i,t-j} \Delta y_{i,t-j} + \eta_{i,t} \Rightarrow \eta_{i,t} \]  \hfill (2)
\[ y_{t,t-1} = \partial_t + \sum_{j=1}^{p_t} \ell_{i,t-j} \Delta y_{i,t-j} + \mu_{i,t-1} \Rightarrow \mu_{i,t-1} \]  

(3)

LLC (2002) procedure suggests more to standardize the errors \( \tilde{\eta}_i, \tilde{\mu}_{i,t-1} \) by the regression standard error according to the ADF equation above:

\[ \tilde{\eta}_i = \frac{\tilde{\eta}_i}{\sigma_{\tilde{\eta}}}, \tilde{\mu}_{i,t-1} = \frac{\tilde{\mu}_{i,t-1}}{\sigma_{\tilde{\mu}}} \]  

(4)

Third, regression can be run to estimate the panel test statistics following below given equation:

\[ \tilde{\eta}_i = \alpha \tilde{\eta}_{i,t-1} + \nu_{i,t} \]  

(5)

The null hypothesis is as follows: \( H_0 : \rho_1=....=\rho_n = \rho = 0 \) and alternate hypothesis is \( H_A : \rho = ....=\rho_n = \rho < 0 \).

IPS (2003) introduced a new and more advanced panel unit root test in the context of a heterogeneous panel. This basically applies the ADF test to individual series thus allowing each series to have its own short-run dynamics and the overall t-test statistic is based on the arithmetic mean of all individual provinces’ ADF statistic. Suppose a series (GDP) can be represented by the ADF (without trend).

\[ \Delta x_{i,t} = \sigma_j + \phi_{i,j} \Delta x_{i,t-j} + \nu_{i,t} \]  

(6)

After the ADF regression has different augmentation lags for each country in finite samples, the term \( E(t_r) \) and \( \text{var}(t_r) \) are replaced by the corresponding group averages of the tabulated values of \( E(t_r, P) \) and \( \text{var}(t_r, P) \) respectively. The IPS (2003) test allows for the heterogeneity in the value \( \sigma_j \) under the alternative hypothesis. This is more efficient and powerful test of stationarity hypothesis than usual single time series test. The econometrical equation of IPS (2003) unit root test is being modeled as given below:

\[ t_{r} = \frac{I}{N} \sum_{t=1}^{N} t_{t,i}(P_i) \]  

(7)

Where \( t_{t,i} \) is the ADF t-statistics for the unit root tests of each province and \( P_t \) is the lag order in the ADF regression and test statistic can be calculated as:

\[ A_r = \frac{\sqrt{N(T)\hat{E}(t_r)} - E(t_r)}{\sqrt{\text{var}(t_r)}} \]  

(8)

As \( t_{r} \) is explained above and values for \( E[t_r (P_t,0)] \) can be obtained from the results of Monte Carlo simulation carried out by IPS (2003). They have calculated and tabulated them for various time periods and lags. When the ADF has different augmentation lags the
two terms $E(t_i)$ and $\text{var}(t_i)$ in the equation above are replaced by corresponding ($P_i$) group averages of the tabulated values of $E(t_i, P_i)$ and $\text{var}(t_i, P_i)$ respectively.

The MW (1999) test is based on the combined significance levels (p-values) from the individual unit root tests. According to MW (1999), if the test statistics are continuous the significance levels $\pi_i$ ($i=1, 2, \ldots N$) are independent and uniform (0,1) variables. They used the combined p-values, or $P_{MW}$, which can be expressed as:

$$P_{MW} = -2 \sum_{i=1}^{N} \log \pi_i$$  \hspace{1cm} (9)

where $-2 \sum \log \pi_i$ has a $\chi^2$ distribution with the $2N$ degree of freedom. Furthermore, Choi (2001) suggested the following standardized statistic:

$$Z_{MW} = \frac{\sqrt{N \{N^{-1}P_{MW} - E[-2 \log(\pi_i)]\}}}{\sqrt{\text{Var}[-2 \log(\pi_i)]}}$$  \hspace{1cm} (10)

Since the data series are not found to be stationary at level, the same tests are performed with the first difference level of the data. The test results indicate that all the series are stationary at the first difference level.

The Breitung’s (2000) method differs from LLL (2002) in two distinct ways. First, only the autoregressive portion (and not the exogenous components) is removed when constructing the standardised proxies:

$$\Delta \tilde{y}_i = (\Delta y_i - \sum_{j=1}^{p_i} \beta_j \Delta y_{i-j}) / s_i$$  \hspace{1cm} (11)

$$\tilde{y}_{i-1} = (y_{i-1} - \sum_{j=1}^{p_i} \beta_j \Delta y_{i-j}) / s_i$$  \hspace{1cm} (12)

Where $\beta, \beta$, and $s_i$ are as defined for LLC (2002).

Second, the proxies are transformed and detrended,

$$\Delta y_i^* = \sqrt{\frac{(T-t)}{(T-t+1)}} \left( \Delta \tilde{y}_i - \frac{\Delta \tilde{y}_{i+1} + \cdots + \Delta \tilde{y}_{i+r}}{T-t} \right)$$

$$y_{i-1}^* = \tilde{y}_{i-1} - c_i$$

where

\[
c_i = \begin{cases} 0, & \text{if } n \text{ intercept or trend} \\ \tilde{y}_i, & \text{with intercept} \\ \tilde{y}_i - (t-1)/T \tilde{y}_i, & \text{with intercept and trend} \end{cases}
\]

The persistence parameter $\alpha$ is estimated from the pooled proxy equation:

$$\Delta y_i^* = \alpha y_{i-1}^* + v_i$$  \hspace{1cm} (13)
Breitung (2000) shows that under the null the resulting parameter $\alpha$ is asymptotically distributed as a standard normal. The Breitung’s (2000) method requires only a specification of the number of lag in each cross-section ADF regression, $p$, and exogenous regressors.

3.2 Tests of Cointegration

We use the Pedroni (1999) framework to test for cointegration. This formulation allows one to investigate heterogeneous panels, in which heterogeneous slope coefficients, fixed effects and individual specific deterministic trends are permitted. In its most simple form, this consists of taking no cointegration as the null hypothesis and using the residuals derived from the panel analogue of an Engle and Granger (1987) static regression to construct the test statistic and tabulate the distributions. The cointegration regression is given by:

$$
Y_i = \alpha_i + \delta t + \beta_{1i} X_{1t} + ... + \beta_{ni} X_{nt} + u_t \\
i = 1, ..., N, t = 1, ..., T \quad (14)
$$

where $Y$ is the dependent variable, $\alpha$ is the constant, $t$ associated with $\delta$ is the trend term, $X$'s are the independent variables and $u$ is the white noise process. Based on the cointegration residuals, Pedroni (1999) derives the asymptotic distribution and explores the small sample performances of seven different statistics. Of these seven statistics, four are based on pooling along what is commonly referred to as the ‘within-dimension’, and three on pooling along what is commonly referred to as the ‘between-dimension’. Pedroni (1999) describes the former and latter as ‘panel cointegration statistics’ and ‘group mean panel cointegration statistics’, respectively. The first of the simple panel cointegration statistics, the ‘$\text{panel } v$-statistic’, is a nonparametric variance ratio statistic. The second, the ‘$\text{panel rho}$-statistic’, is a panel version of a nonparametric statistic analogous to the familiar (Phillips and Perron 1988) rho-statistic. The third, the ‘$\text{panel } t$-statistic (nonparametric)’, is a nonparametric statistic analogous to the Phillips and Perron $t$-statistic. The fourth, the ‘$\text{panel } t$-statistic (parametric)’, is a parametric statistic analogous to the familiar ADF $t$-statistic. The other three panel cointegration statistics are based on a group mean approach. The first, the ‘group rho-statistic’, is analogous to the Phillips and Perron rho-statistic. The last two, the ‘group $t$-statistic (nonparametric)’ and ‘group $t$-statistic (parametric)’, are analogous to the Phillips and Perron $t$-statistic and the ADF $t$-statistic, respectively. The asymptotic distributions of these panel cointegration statistics are derived in Pedroni (1997). Under an appropriate standardization, based on the moments of the vector of Brownian motion functions, these statistics are distributed as standard normal. The standardization is given by:

$$
\kappa = [k_N - \mu \sqrt{N}] / \sqrt{N} \quad (15)
$$

Pedroni (1999) gives critical values for $\mu$ and $\nu$ with and without intercepts and deterministic trends. The small sample size and power properties of all seven tests are discussed in Pedroni (1997). He finds that size distortions are minor, and power is high for all statistics when the time span is long. For shorter panels, the evidence is more varied. However, in the
presence of a conflict in the evidence provided by each of the statistics. Pedroni shows that the group-adf statistic and panel-adf statistic generally perform best.

Further, to see the robustness of the empirical results, we also apply the alternative cointegration tests developed by Kao and Chiang (1998). Consider the following fixed-effect panel regression:

\[ Y_{it} = \beta_0 + \beta_1 X_{1it} + \ldots + \beta_m X_{mit} + u_{it}, \quad i = 1, \ldots, N, \quad t = 1, \ldots, T \]  \hspace{1cm} (16)

where \( \{Y_{it}, X_{1it}, \ldots, X_{mit}\} \) are independent across cross-sectional units and \( u_{it} \) is a linear process that satisfies the assumption in Kao and Chiang (1998). Kao (1999) presents DF and ADF types of cointegration tests in the panel data. Kao and Chiang (1998) derive limiting distributions for the ordinary least square (OLS), fully modified (FM) and dynamic ordinary least square (DOLS) estimators in a cointegrated regression and then show that they are asymptotically normal.

Further, we used Panel LLL (2001) trace test statistics as a final robustness measure that is actually derived from the average of individual likelihood ratio cointegration rank trace test statistics from the panel individuals. The multivariate cointegration trace test of Johanson (1988) is engaged to investigate each individual cross-section system autonomously, in that way, allowing heterogeneity in each cross-sectional unit root for said panel. Process of data generation for each of the groups is characterized by following heterogeneous VAR (p) model:

\[ Y_{ij} = \sum_{j=1}^{p_i} \Lambda_{i,j} Y_{i,j-1} + \varepsilon_{i,j} \]  \hspace{1cm} (17)

Where \( i = 1, \ldots, N; t = 1, \ldots, T \)

For each one, the value of \( Y_{i,-i+1}, \ldots, Y_{i,0} \) is considered fixed and \( \varepsilon_{i,j} \) are independent and identically distributed (normally distributed): \( \varepsilon \sim N_k(0, \Omega_i) \), where \( \Omega_i \) is the cross-correlation matrix in the error terms: \( \Omega_i = E(\varepsilon_{i,j}, \varepsilon_{i,j}'). \) Equation (17) can be modified in vector error correction model (VECM) model as given below:

\[ \Delta Y_{ij} = \Pi_i Y_{i,i-1} + \sum_{j=1}^{p_i-1} \Gamma_{i,j} \Delta Y_{i,j-1} + \varepsilon_{i,j} \]  \hspace{1cm} (18)

Where \( \Pi_i = \Lambda_{i,1} + \ldots + \Lambda_{p_i - 1} \) and \( \Gamma_{i,j} = \Lambda_{i,j} - \Lambda_{i,j-1}, \Pi_i \) is of order \((k \times k)\). If \( \Pi_i \) is of reduced rank: \( \text{rank}(\Pi_i) = r_i \), which can be decomposed into \( \Pi_i = \alpha_i \beta' \), where \( \alpha_i \) and \( \beta \) are of order \((k \times r_i)\) and of full rank column rank that represents the error correction form. The null hypotheses of panel LLL (2001) rank test are:
\[ H_0 = \text{rank}(\mathbf{\Pi}_i) = r_i \leq r \text{ for all } i = 1, \ldots, N \text{ against} \]

\[ H_0 = \text{rank}(\mathbf{\Pi}_i) = k \text{ for all } i = 1, \ldots, N \]

The procedure is in sequences like individual trace test process for cointegration rank determination. First, we test for \( H_0 = \text{rank}(\mathbf{\Pi}_i) = r_i \leq r, r = 0 \), if null hypothesis of no cointegration is accepted, this shows that there is no cointegration association \((\text{rank}(\mathbf{\Pi}_i) = r_i = 0)\) in all cross-sectional groups in said panel. If null hypothesis is not accepted then null hypothesis \( r = 1 \) is tested. The sequence of procedure is not disconnected and continued until null hypothesis is accepted, \( r = k - 1 \) is rejected. Accepting the hypothesis of Cointegration \( r = 0 \) along with null hypothesis of rank \((\mathbf{\Pi}_i) = r \leq 0 (0 < r < k)\) implies that there is at least one cross-sectional unit in said panel, which has \( \text{rank}(\mathbf{\Pi}_i) = r > 0 \).

The likelihood ratio trace test statistic for group \( i \) is as given below;

\[
R_{ii} \{ H(r) / H(k) \} = -2 \ln Q_{ii} (H(r) / H(k)) = -T \sum_{i=r+1}^{p} \ln(1 - \lambda_i)
\]

(19)

Where \( \lambda_i \) is the \( i^{th} \) largest Eigen value in the \( i^{th} \) cross-section unit? The LR-bar statistic is calculated as the average of individual trace statistics.

\[
L \hat{R}_{ii} [H(r) / H(k)] = \frac{1}{N} \sum_{i=1}^{n} R_{ii} [H(r) / H(k)]
\]

(20)

Finally, modified version of above equation is defined as:

\[
\lambda_{i,LR} [H(r) / H(k)] = \sqrt{N} (L \hat{R}_{ii} [H(r) / H(k)] - E(Z_k)) \over \sqrt{\text{VAR}(Z_k)}
\]

(21)

Where \( E(Z_k) \) and \( \text{VAR}(Z_k) \) are mean and variance of the asymptotic trace statistics, which can be obtained from simulation. LLL (2001) prove the central limit theorem for the standard LR-bar statistics that under the null hypothesis, \( \lambda_{i,LR} \Rightarrow N(0,1) \) as \( N \) and \( T \rightarrow \infty \) in such a way that \( \sqrt{N} \rightarrow 1 \rightarrow 0 \), under the assumption that there is no cross-correlation in the error terms, that is:

\[
E(\varepsilon_{i,t}) = 0 \text{ and } E(\varepsilon_{i,t}, \varepsilon_{j,t}) = \begin{cases} \Omega, & \text{for } i = j, i \neq j \\ 0 & \text{for } i, j \neq j \end{cases}
\]
LLL (2001) note that $T \to \infty$ is needed for each of the individual test statistics to converge to its asymptotic distribution, while $N \to \infty$ is needed for the central limit theorem.

4. Data analysis and findings

Firstly, we have analyzed the unit root property of the data series of our panel and reported results in table 1 below. For the analysis, we used two models. First includes constant term in the model and second model includes constant and trend both.

It is evident from Table 1 that there is consensus among the results of different test statistics for nonstationary of all the three variables analysed. However, in some cases we have contradictory results but preferring to our model that includes constant and trend both we can conclude that all three variables are nonstationary in their level form whereas stationary in the first difference form. This allows us to proceed for cointegration analysis. We used a number of cointegration test and different specifications in order to test the sensitivity of our results as various studies have shown that cointegration results are very much sensitive to the trend assumption used, model used and lag structure used in the analysis. Results of cointegration analysis are presented in following Table 2.

It is evident from results of Pedroni (1997) residual based cointegration test that in both specifications i.e., no deterministic term included and when deterministic intercept and trend included we have same findings i.e., in both specification and in all seven statistics we are unable to reject the null hypothesis of cointegration. Further, to test the robustness of this residual based cointegration of Pedroni (1997) we applied Kao (1999) residual based cointegration based test. Kao (1999) residual based cointegration based test is unable to reject the null hypothesis of cointegration at conventional level that is at 5 per cent level of significance. In the final step, we applied LLL (2001) test of cointegration and used different specifications and lag length in order to test the sensitivity of LLL (2001) results and to see the robustness of our previous results. Results of LLL (2001) test are found to be sensitive with the choice of lag length but not for the deterministic assumption used. In general, at 5 per cent level of significance we can conclude that there is evidence of at most 2 cointegration relations in our data set.

5. Conclusions

Recently, the concern of economists, researchers, policymakers, and environmentalists of all over the world has increased on issue related to environment, (particularly global warming, and CO2 emissions are the most responsible factor in that), energy consumption, and green economic growth. Number of studies have been conducted based on individual country and group of countries and have found conflicting results. With this respect, we conducted a study for BRICS countries in panel framework. We choose BRICS because this group includes fastest growing developing economies of the Asia on which burden has fallen to achieve sustainable development. For the analysis, panel model is preferred in order to reduce the bias which may arise due to small period (study period of our study is 1985-2009 with annual observations).
### Table 1: Results of Unit Root Analysis

<table>
<thead>
<tr>
<th>Method</th>
<th>LNCO2</th>
<th>D(LNCO2)</th>
<th>LGDGP</th>
<th>D(LNGDP)</th>
<th>LNHEC</th>
<th>D(LNHEC)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Levin, Lin &amp; Chu*</td>
<td>Statistic</td>
<td>P-value</td>
<td>Statistic</td>
<td>P-value</td>
<td>Statistic</td>
<td>P-value</td>
</tr>
<tr>
<td></td>
<td>-1.64262</td>
<td>0.0502</td>
<td>-3.15603</td>
<td>0.0008</td>
<td>-3.02333</td>
<td>0.0013</td>
</tr>
<tr>
<td>Im, Pesaran and Shin W-stat</td>
<td>1.04468</td>
<td>0.8519</td>
<td>-3.71515</td>
<td>0.0001</td>
<td>3.31970</td>
<td>0.9995</td>
</tr>
<tr>
<td>ADF - Fisher Chi-square</td>
<td>8.94881</td>
<td>0.5370</td>
<td>32.5848</td>
<td>0.0003</td>
<td>12.23406</td>
<td>0.2629</td>
</tr>
<tr>
<td>PP - Fisher Chi-square</td>
<td>9.80256</td>
<td>0.4580</td>
<td>32.9688</td>
<td>0.0003</td>
<td>33.7238</td>
<td>0.0002</td>
</tr>
<tr>
<td>Breitung t-stat</td>
<td>-0.08357</td>
<td>0.4667</td>
<td>-2.78967</td>
<td>0.0026</td>
<td>0.04733</td>
<td>0.5189</td>
</tr>
<tr>
<td>Im, Pesaran and Shin W-stat</td>
<td>0.63333</td>
<td>0.7367</td>
<td>-3.05017</td>
<td>0.0011</td>
<td>2.02378</td>
<td>0.9785</td>
</tr>
<tr>
<td>ADF - Fisher Chi-square</td>
<td>6.87067</td>
<td>0.7367</td>
<td>28.1852</td>
<td>0.0017</td>
<td>2.55159</td>
<td>0.9901</td>
</tr>
<tr>
<td>PP - Fisher Chi-square</td>
<td>3.24154</td>
<td>0.9752</td>
<td>28.7071</td>
<td>0.0014</td>
<td>2.02090</td>
<td>0.9962</td>
</tr>
</tbody>
</table>

Source: Authors’ Calculation

### Table 2: Results of Panel Cointegration Test

<table>
<thead>
<tr>
<th>Trend assumption</th>
<th>No deterministic trend</th>
<th>Deterministic intercept and trend included</th>
</tr>
</thead>
<tbody>
<tr>
<td>Method</td>
<td>Test statistic</td>
<td>P-Value</td>
</tr>
<tr>
<td>Pedroni Residual Cointegration Test:</td>
<td>-0.163352</td>
<td>0.5649</td>
</tr>
<tr>
<td>Panel v-Statistic</td>
<td>0.943061</td>
<td>0.8272</td>
</tr>
<tr>
<td>Panel rho-Statistic</td>
<td>0.611407</td>
<td>0.7295</td>
</tr>
<tr>
<td>Panel PP-Statistic</td>
<td>0.016511</td>
<td>0.5066</td>
</tr>
<tr>
<td>Panel ADF-Statistic</td>
<td>1.659635</td>
<td>0.9515</td>
</tr>
<tr>
<td>Group rho-Statistic</td>
<td>0.887312</td>
<td>0.8044</td>
</tr>
<tr>
<td>Group PP-Statistic</td>
<td>0.432669</td>
<td>0.6674</td>
</tr>
<tr>
<td>Group ADF-Statistic</td>
<td>-1.384642</td>
<td>0.0831</td>
</tr>
<tr>
<td>Kao Residual Cointegration Test:</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Johansen-Fisher tests</td>
<td>Linear deterministic trend and Lags interval (1,1)</td>
<td>Max.eigenvalue</td>
</tr>
<tr>
<td>Null hypothesis</td>
<td>At most 1</td>
<td>17.00</td>
</tr>
<tr>
<td></td>
<td>At most 2</td>
<td>14.96</td>
</tr>
<tr>
<td></td>
<td>Linear deterministic trend (restricted) and Lags interval (1,1)</td>
<td>Max.eigenvalue</td>
</tr>
<tr>
<td></td>
<td>At most 1</td>
<td>13.45</td>
</tr>
<tr>
<td></td>
<td>At most 2</td>
<td>15.29</td>
</tr>
</tbody>
</table>

Source: Authors’ calculation
We find that all variables are nonstationary in the level from whereas stationary in the first difference form. Our results of residual based cointegration analysis show the absence of cointegration whereas cointegration analysis based on Johansen-Fisher tests shows, in general, presence of two cointegration equations. Further, we have not analysed structural breaks in the unit root and cointegration process and the study can be extended in this direction. Further, more variables can be used, for example, the present analysis can be done in the production function framework.

References


