Energy Policy 69 (2014) 82–91

Contents lists available at ScienceDirect

Energy Policy

journal homepage: www.elsevier.com/locate/enpol

Causal relationship between trade openness, economic growth and energy consumption: A panel data analysis of Asian countries

Samia Nasreen*, Sofia Anwar

Department of Economics, Government College University Faisalabad, Faisalabad, Punjab, Pakistan

HIGHLIGHTS

- This study analyzes causality between energy, growth and trade in the Asian region.
- Empirical results supported cointegrating relationship between variables.
- Positive impact of growth and trade openness on energy usage is found in the long run.
- Bidirectional Granger causality is observed between selected variables in the long run.

ARTICLE INFO

Article history:
Received 16 December 2013
Received in revised form
5 February 2014
Accepted 6 February 2014
Available online 12 March 2014

Keywords:
Energy consumption
Trade openness
Economic growth

ABSTRACT

This paper explores the causal relationship between economic growth, trade openness and energy consumption using data of 15 Asian countries. The study covers the period of 1980–2011. We have applied panel cointegration and causality approaches to examine the long-run and causal relationship between variables.

Empirical results confirm the presence of cointegration between variables. The impact of economic growth and trade openness on energy consumption is found to be positive. The panel Granger causality analysis reveals the bidirectional causality between economic growth and energy consumption, trade openness and energy consumption.

1. Introduction

Rapid population growth, technological development and trade expansion have increased the demand for energy consumption in the recent decade. Around the world energy consumption, economic production and international trades tend to move together so it is significant to learn more about the relationship between energy consumption, economic growth and trade openness. Energy consumption and economic growth relationship is vital because if there is a strong relationship between energy consumption and economic growth, it is very hard to change energy and environmental policies. Furthermore, if the relationship between energy consumption and economic growth is not significant, then energy conservation policy may be adopted with no adverse impact on the economy.

Trade openness is an essential component of economic growth and increase in international trade increases the economic activities and the energy demand (Sadorsky, 2012). The economic condition of the country and the extent of relationship between economic growth and trade openness determine the impact of trade openness on energy consumption (Cole, 2006). Trade openness enables developing economies to import advanced technologies from developed economies. The adoption of advanced technology lowers energy intensity and produces more output. Similarly, energy affects trade openness via various channels. Firstly, energy is an important input of production because machinery and equipment in the process of production require energy. Secondly, exporting or importing manufactured goods or raw materials requires energy to fuel transportation. Without adequate energy supply, trade openness will be adversely affected. Consequently, energy is an important input in trade expansion and adequate consumption of energy is essential for expanding trade via expanding exports and imports. The relationship between trade openness and energy consumption is important. If energy plays its key role to increase the flow of exports or imports, then any policies aiming at reduction in energy consumption such as energy conservation policies will negatively impact the flow of exports or imports, and hence reduce the benefit of trade openness.

Asian countries account for more than 50% of the global population and nature has endowed them with an array of natural energy resources such as wind, coal, water, oil, wood and solar power and a large number of these resources have remained unexploited for decades. This region accounts for just over

* Corresponding author.
E-mail addresses: sami_lcu@yahoo.com (S. Nasreen), Sofia_ageconomist@yahoo.com (S. Anwar).

http://dx.doi.org/10.1016/j.enpol.2014.02.009
0301-4215

one-quarter (28%) of the global primary energy demand, with more than half of this (17% of the global total) in China. Asian economies are relatively more coal-intensive than the rest of the world, accounting for more than half (53%) of the world's coal consumption. Electricity generation from renewable energy sources in Asia is projected to grow at an average annual rate of 5%, which would increase the renewable share of the region's total generation from 15%% in 2007 to 20% in 2035 (International Energy Agency, 2012). In the year 1980, the Asian region's average energy consumption was equivalent to 1102 Kg of oil while the amount increased to 2508 Kg of oil equivalent in 2011 (see Fig. 1).

Asian economies contribute to one-fourth of the world's trade in goods, after Europe. Exports from North America and Asia have grown faster than imports. The growth rate of Asian export was 13% while imports grew by 9%. More than 50% of Asian exports are conducted within the region. Parallel to growing intra-regional trade, Asia's inter regional trade has also increased over time. Europe (18.4%) and North America (21.4%) have become the two largest destinations of Asia's exports. The top merchandise exporter in 2011 was China (US$ 1.58 trillion). The second largest importer in 2011 was also China (US$ 1.40 trillion) (Source: World Trade Report, 2012).

Trade volume in Asia has been rising fast since the early 1970s. Asian region's merchandise trade (export plus import of goods) was worth US$ 0.8 trillion in 1980 but it has amounted to US$ 14 trillion in 2011 (see Fig. 2).

Sustained rapid growth, macroeconomic stability, and improvements in living standards are some of the remarkable achievements of the Asian economies over the past decade. Per capita income in the Asian countries has increased with the passage of time. Developing economies of East Asia and the Pacific region have become an engine of global growth, growing at 7.5% in 2012, higher than any other region in the world. The Asian region contributed to around 40% of global growth in 2012 and the global economy continues to rely on the regions's growth (World Bank Report, 2013). In 1980, the GDP per capita income was estimated as US$1155 but it increased to US$8489 in 2011 (see Fig. 3).

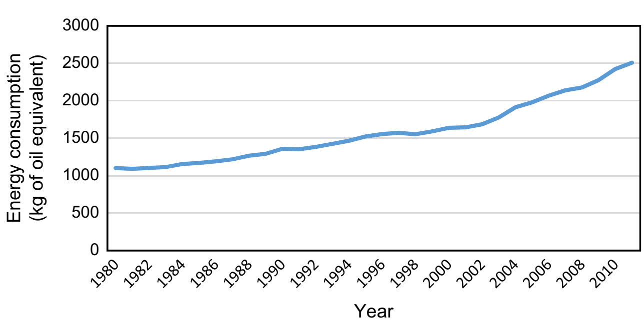

Fig. 1. Annual changes in energy consumption in Asia.
Source: World Development Indicators (CD-ROM, 2012).

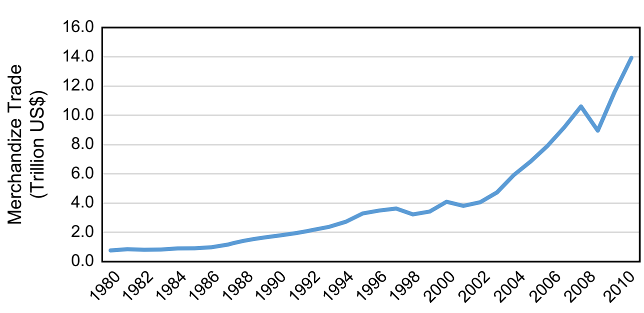

Fig. 2. Annual changes in merchandize trade in Asia.
Source: World Development Indicators (CD-ROM, 2012).

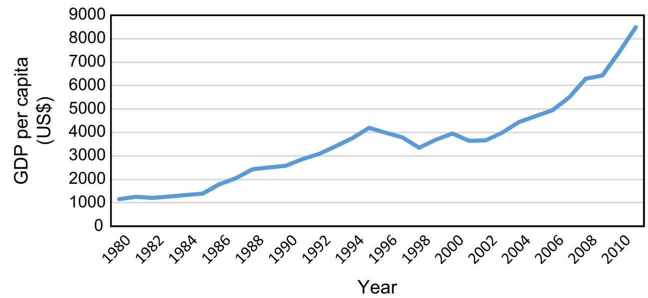

Fig. 3. Annual changes in per capita income in Asia.
Source: World Development Indicators (CD-ROM, 2012).

The existing energy economics literature seems to provide numerous studies which have investigated the causal relationship between energy consumption and economic growth (for example, see Yang, 2000; Narayan et al., 2008; Ozturk, 2010; Payne, 2010). Exports are also considered as an engine of economic growth in the theoretical growth model, and in international economics literature exports and output relationship is widely studied (Giles and Williams, 2000a, 2000b; Lean and Smyth, 2010; Halicioglu, 2010).

This paper extends the literature on energy consumption, economic growth and trade openness in three ways. First, this study uses aggregate variables for energy consumption, economic growth and trade openness so it is a more comparable study than the previous studies which used electricity consumption and exports variables to understand the relationship between energy consumption and trade. Second, in literature most of the researchers investigated only the relationship between energy consumption and economic growth or the relationship between economic growth and trade openness. But it is vital to understand the dynamic relationship of these variables; they must be taken in a combined model. Third, this paper investigates the energy consumption, economic growth and trade openness relationship for Asian countries, the area of the world's largest economies. Also, this is the first study to investigate the link between energy consumption, economic growth and trade openness in the Asian countries. This paper's results are vital for developing energy and environmental policy in the Asian countries.

The rest of the paper is organized as follows: Section 2 gives a brief review of empirical studies, Section 3 presents empirical model and data source, Section 4 provides estimation methodology, Section 5 reports the empirical analysis of results and finally Section 6 concludes the study.

2. Literature review

Theoretically there is a direct association between energy consumption and economic growth. Various studies have been conducted to support this association after the end of the 1970s energy crisis. All these studies (e.g., Kraft and Kraft, 1978; Akarca and Long, 1979, 1980; Yu and Choi, 1985; Abosedra and Baghestani, 1989) found a positive impact of energy consumption on economic growth. However, empirical evidence provided by Zahid (2008), Amirat and Bouri (2010), Noor and Siddiqi (2010), and Apergis and Payne (2010) is conflicting about the direction of causality. For instance, Yu and Choi (1985) investigated the causal relationship between national income and different forms of energy consumption by using cross-country analysis. This study was unable to find any significant relationship between energy and growth in the United States, Poland and United Kingdom. However, a significant relationship was observed between energy consumption and income growth in South Korea and Philippines.

Mazumder and Marathe (2007) observed the causal relationship between per capita GDP and per capita electricity consumption in Bangladesh. Their findings showed that energy is an important contributing factor for increasing the economic growth of Bangladesh. Chebbi and Boujelbene (2008) examined the causality between energy consumption, growth performance and environmental pollution in Tunisia. Empirical results supported long-term association between energy consumption, growth performance and environmental pollution between the period 1971–2004. Results also showed evidence of short-term unidirectional causality between energy use and economic growth in Tunisia for the same period.

Loganathan and Subramaniam (2010) investigated the sustainable relationship between total energy utilization and economic growth in Malaysia by employing auto-regressive distributed lag (ARDL) bound testing approach and the error-correction model (ECM). The result based on ARDL bound testing approach confirmed the existence of long-run relationship between energy consumption and economic growth for the period 1971–2008. The result of ECM revealed that there exists feedback causality between energy consumption and economic growth in the short run during the same period. Amirat and Bouri (2010) in their study on Algeria investigated the causality between per capita energy consumption and per capita GDP for the period 1980–2007. By applying the Engel Granger Co integration technique, the study concluded that neither of the series is co-integrated. The results of the Granger causality test revealed that energy consumption affects economic growth in Algeria.

Nondo and Kahsai (2009) investigated the long-run relationship between total energy consumption and economic growth for a panel of 19 African countries. They applied Levin et al. (1993), Im et al. (1997) and Hadri (2000) panel unit root tests to test the integrating properties of real GDP and total energy consumption. Their analysis indicated that both the variables are cointegrated for long-run relationship confirmed by the Pedroni (1999) panel cointegration approach. Moreover, they noted that economic growth is the cause of energy consumption in the long run as well as in the short run. Noor and Siddiqi (2010) investigated the causal relationship between per capita energy consumption and per capita GDP in five South Asian countries, namely, Bangladesh, India, Nepal, Pakistan and Sri Lanka. They applied panel unit root tests IPS, LLC and MW, and Pedroni cointegration as well as Kao residual cointegration approaches. They reported that energy consumption enhances economic growth. Their causality analysis reveals that economic growth Granger causes energy consumption in the South Asian countries.[1]

There are few studies investigating the relationship between trade openness and energy consumption. For instance, Cole (2006) examined the relationship between trade liberalization and energy consumption. He used data of 32 countries and found that trade liberalization promotes economic growth, which boosts energy demand. Moreover, trade liberalization stimulates capitalization which in return affects energy consumption. Jena and Grote (2008) investigated the impact of trade openness on energy consumption. They noted that trade openness stimulates industrialization via scale effect, technique effect, composite effect and comparative advantages effect which affect energy consumption. Narayan and Smyth (2009) examined the causal relationship between energy consumption and economic growth by incorporating exports as an indicator of trade openness in the production function for a panel of six Middle Eastern countries, namely, Iran, Israel, Kuwait, Oman, Saudi Arabia and Syria. They applied panel unit root test, panel cointegration and panel causality tests. Their analysis confirmed the presence of a cointegration relationship between the variables. Furthermore, they reported that a short-run Granger causality exists running from energy consumption to real GDP and from economic growth to exports but a neutral effect is found between exports and energy consumption.

Subsequently, Sadorsky (2011) examined the causal relationship between total economic growth, energy consumption and trade openness. The panel mean group cointegration and panel Granger causality approaches for the panel of 8 Middle Eastern countries, namely, Bahrain, Iran, Jordan, Oman, Qatar, Saudi Arabia, Syria and UAE. The empirical evidence reported that a long-run relationship exists between the variables. Sadorsky found that 1% increase in real per capita GDP increases per capita energy consumption by 0.62%. A 1% increase in real per capita exports increases per capita energy consumption by 0.11%, while 1% increase in real per capita imports increases per capita energy consumption by 0.04%. Panel Granger causality analysis revealed that exports Granger cause energy consumption and feedback is found between imports and energy consumption in the short run. Similarly, bidirectional causality exists between GDP and energy consumption in the short run. Sadorsky (2012) used production function to investigate the relationship between trade openness and energy consumption in the South American countries, namely, Argentina, Brazil, Chile, Ecuador, Paraguay, Peru, and Uruguay over the period 1980–2007. The panel cointegration developed by Pedroni (2004) fully modified ordinary least squares (FMOLS), and the VECM Granger causality approaches were applied. The empirical evidence confirmed the presence of cointegration for a long-run relationship between the variables. The relationship between exports and energy consumption is bidirectional and imports Granger causes energy consumption in the short run. Using data of 52 developed and developing economies, Ghani (2012) explored the relationship between trade liberalization and energy demand. The results indicated that trade liberalization has a insignificant impact on energy consumption but after a certain level of capital per labor, trade liberalization affects energy consumption.

Hossain (2012) examined the relationship between electricity consumption and exports by adding foreign remittances and economic growth as additional determinants in three SAARC countries, namely, Pakistan, India and Bangladesh. The author reported no causality between exports and electricity demand. Dedeoğlu and Kaya (2013) investigated the relationship between exports, imports and energy consumption by incorporating economic growth as an additional determinant of trade openness and energy consumption using data of the OECD countries. They applied the panel cointegration technique developed by Pedroni (2004) and used the Granger causality developed by Canning and Pedroni (2008). Their analysis showed the cointegration between the variables. They also noted that economic growth, exports and imports have a positive impact on energy consumption. Their causality analysis revealed that the relationship between exports (imports) and energy consumption is bidirectional.

3. Econometric model and data source

Following Sadorsky (2011), the relationship between total energy consumption (EN), economic growth (Y), trade openness (T) and energy price (EP) is modeled as follows:

$$EN_{it} = f\,(Y_{it}, T_{it}, EP_{it}, W_i) \tag{1}$$

$$EN_{it} = \alpha_{1i}Y_{it} + \alpha_{2i}T_{it} + \alpha_{3i}EP_{it} + W_i + u_{it} \tag{2}$$

In Eq. (2), cross-sections are denoted by subscript i (i=1, 2,…, N) and time period by subscript t (t=1, 2,…, T), W is the country fixed effect and u is the stochastic random term. Kg of oil equivalent per capita is used to measure energy consumption, real GDP per capita in constant

[1] Payne (2010) and Ozturk, (2010) presented comprehensive survey studies on the relationship between economic growth and energy consumption.

international dollar is used to measure economic growth, exports (US$) plus imports (US$) divided by population is used to measure trade openness, the price of Dubai crude oil (US$) deflated by the country's consumer price index (2005 = 100) is used as a proxy for energy price due to the unavailability of energy price data.

15 Asian countries are selected for the estimation of the econometric model on the basis of data availability and the policy to use balanced panel. Asian countries included in the balanced panel are: Pakistan, India, Bangladesh, Sri Lanka, Philippines, Thailand, Indonesia, China, Malaysia, Japan, Jordan, Iran, Korea Dem., Nepal and Vietnam. The study covers the period of 1980–2011. Data on energy consumption per capita, merchandise exports, merchandise imports, consumer price index and population are obtained from World Development Indicators (2013) of the World Bank. Data on real GDP per capita are collected from Penn World Tables Version 8.0 (Heston et al., 2013) and Dubai crude oil price data are taken from British Petroleum's 2013 statistical review of world energy.

4. Estimation strategy

4.1. Panel unit roots

We apply Levin et al. (1993) (LLC), Im et al. (1997) (IPS), Maddala and Wu (1999) (MW, ADF) and Maddala and Wu (1999) (MW, PP) panel unit root tests to check the stationarity properties of the variables. These tests apply to a balanced panel but the LLC can be considered a pooled panel unit root test, IPS represents a heterogeneous panel test and MW panel unit root test is a non-parametric test.

4.1.1. LLC unit root test

Levin et al. (1993) developed a number of pooled panel unit root tests with various specifications depending upon the treatment of the individual specific intercepts and time trends. Their test imposes homogeneity on the autoregressive coefficient that indicates the presence or absence of unit root problem while the intercept and the trend can vary across individual series. LLC unit root test follows ADF regression for the investigation of unit root hypothesis as given below step by step:

1. Implement a separate ADF regression for each country:

$$\Delta y_{i,t} = \alpha_i + \rho_i y_{it-1} + \sum_{j=1}^{p_i} \alpha_{i,j} \Delta y_{i,t-j} + \varepsilon_{i,t} \tag{3}$$

The lag order p_i is allowable across individual countries. The appropriate lag length is chosen by allowing the maximum lag order and then uses the t-statistics for ij b to determine if a smaller lag order is preferred.

2. Run two separate regressions and save the residuals $\tilde{\eta}_{it}$, $\tilde{\mu}_{i,t-1}$

$$\Delta y_{i,t} = \lambda_i + \sum_{j=1}^{p_i} \gamma_{i,t-j} \Delta y_{i,t-j} + \eta_{i,t} \Rightarrow \tilde{\eta}_{it} \tag{4}$$

$$y_{i,t-1} = \partial_i + \sum_{j=1}^{p_i} \ell_{i,t-j} \Delta y_{i,t-j} + \mu_{i,t-1} \Rightarrow \tilde{\mu}_{i,t-1} \tag{5}$$

LLC procedure suggests standardize the errors $\tilde{\eta}_{it}$, $\tilde{\mu}_{i,t-1}$ by regressing the standard error the ADF equation provided above:

$$\tilde{\eta}_{it} = \frac{\tilde{\eta}_{it}}{\hat{\sigma}_{\varepsilon i}}, \quad \tilde{\eta}_{it-1} = \frac{\tilde{\eta}_{i,t-1}}{\hat{\sigma}_{\varepsilon i}} \tag{6}$$

3. Regression can be run to compute the panel test statistics following Eq. (5):

$$\tilde{\eta}_{it} = \alpha \tilde{\eta}_{i,t-1} + \nu_{i,t} \tag{7}$$

The null hypothesis is as follows: $H_\circ : \rho_1, \ldots = \ldots \rho_n = \rho = 0$ and alternate hypothesis is $H_A : \rho = \ldots \rho_n = \rho < 0$.

4.1.2. IPS unit root test

Im et al. (IPS) (1997) introduced a panel unit root test in the context of a heterogeneous panel. This test basically applies the ADF test to individual series thus allowing each series to have its own short-run dynamics. But the overall t-test statistics is based on the arithmetical mean of all individual countries' ADF statistics. Suppose a series (TR_{ti}, EC_{ti}) can be represented by the ADF (without trend).

$$\Delta x_{i,t} = \varpi_j + \varpi_i x_{i,t-1} + \sum_{j=1}^{p_i} \phi_{i,j} \Delta x_{i,t-j} + \nu_{i,t} \tag{8}$$

After the ADF regression has different augmentation lags for each country in finite samples, the terms $E(t_T)$ and $\text{var}(t_T)$ are replaced by the corresponding group averages of the tabulated values of $E(t_T, P_i)$ and $\text{var}(t_T, P_i)$, respectively. The IPS test allows for the heterogeneity in the value ϖ_i under the alternative hypothesis. This is a more efficient and powerful test than usual single time series test. The estimable equation of IPS unit root test is modeled as follows:

$$t_{NT} = \frac{I}{N} \sum_{i=1}^{N} t_{i,t}(P_i) \tag{9}$$

where $t_{i,t}$ is the ADF t-statistics for the unit root tests of each country and P_i is the lag order in the ADF regression and test statistics can be calculated as follows:

$$A_{t^-} = \frac{\sqrt{N(T)}[\bar{t}_T - E(t_T)]}{\sqrt{\text{var}(t_T)}} \tag{10}$$

As t_{NT} is explained above and values for $E[t_{iT}(P_i, 0)]$ can be obtained from the results of Monte Carlo simulation carried out by IPS, they have calculated and tabulated them for various time periods and lags. When the ADF has different augmentation lags (P_i), the two terms $E(t_T)$ and $\text{var}(t_T)$ in the equation above are replaced by corresponding group averages of the tabulated values of $E(t_T, P_i)$ and $\text{var}(t_T, P_i)$, respectively.[2]

4.1.3. MW unit root test

The Fisher-type test was developed by Maddala and Wu (1999), which pools the probability values obtained from unit root tests for every cross-section i. This is a non-parametric test and has a chi-square distribution with 2nd degree of freedom where n is the number of countries in a panel. The test statistics are given by

$$\lambda = -2 \sum_{i=1}^{n} \log_e(p_i) \sim \chi^2_{2n}(d.f.) \tag{11}$$

[2] Karlsson and Löthgren (2000) demonstrate the power of panel unit root tests by Monte Carlo simulation. The null of all these tests is that each series contains a unit root and thus is difference stationary. However, the alternative hypothesis is not clearly specified. In LLC the alternative is that all individual series in the panel are stationary. In IPS the alternative is that at least one of the individual series in the panel is stationary. They conclude that the “presence or absence of power against the alternative where a subset of the series is stationary has a serious implications for empirical work. If the tests have high power, a rejection of the unit root null can be driven by few stationary series and the whole panel may inaccurately be modeled as stationary. If, on other hand, the tests have low power it may incorrectly concluded that the panel contains a common unit root even if a majority of the series is stationary” (p. 254). The simulation results reveal that the power of the tests (LLC, IPS) increases monotonically with: (1) an increased number (N) of the series in the panel; (2) an increased time series dimension (T) in each individual series; (3) increased proportion of stationary series in the panel. Their Monte Carlo simulations for N=13 and T=80 reveal the power of the test is 0.7 for LLC tests and approaching unity for the IPS tests.

where p_i is the probability value from ADF unit root tests for unit i. The MW unit root test is superior to the IPS unit root test because the MW unit root test is sensitive to lag length selection in individual ADF regressions. Maddala and Wu (1999) performed Monte Carlo simulations to prove that their test is more advanced than the test developed by IPS (2003).

4.2. The panel cointegration tests

Advance panel cointegration tests can be expected to have a higher power than the traditional tests. The tests applied for long-run examination are developed by Pedroni (1999, 2004) and Larsson et al. (2001).

Pedroni (1999) uses the following cointegration equation:

$$x_{i,t}=\alpha_i+\rho_i t+\beta_{1i}Z_{1i,t}+\cdots+\beta_{mi}Z_{mi,t}+\mu_{it} \tag{12}$$

where x and Z are assumed to be integrated of order one. The specific intercept term α_i and slope coefficients $\beta_{1i}, \beta_{2i}, \ldots, \beta_{mi}$ vary across individual members of the panel. Pedroni (1999, 2004) proposed seven different statistics to test for cointegration relationship in a heterogeneous panel. These tests are corrected for bias introduced by potentially endogenous regressors. In the presence of cross-sectional dependence, Pedroni suggests inclusion of common time dummies to eliminate this effect. The seven test statistics of Pedroni are classified into within dimension and between dimensions statistics. Within dimension statistics are referred to as panel cointegration statistics, while between dimension statistics are called group mean panel cointegration statistics. All statistics test the null hypothesis of no cointegration as: $H_0:\rho_i=1$ for all $i=1, 2, \ldots, N$. The alternative hypothesis for between dimension and within dimension for panel cointegration is different. The alternative hypothesis for between dimension statistics is $H_a:\rho_i<1$ for all $i=1, 2, \ldots, N$, where a common value for $\rho_i=\rho$ is not required. The alternative hypothesis for within dimension-based statistics is $H_a:\rho_i=\rho<1$ for all $i=1, 2, \ldots, N$. First we compute the regression residuals from the hypothesized cointegration Eq. (8), then follow Pedroni's seven test statistics:

1. Panel ν-statistics: $Z_v\equiv T^2N^{3/2}\left(\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{\kappa}^{-2}{}_{11,i}\hat{\mu}_{it-1}^2\right)^{-1}$
2. Panel ρ -statistics: $Z_p\equiv T\sqrt{N}\left(\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{\kappa}^{-2}{}_{11,i}\hat{\mu}_{it-1}^2\right)^{-1}$ $\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{\kappa}^{-2}{}_{11,i}(\hat{\mu}_{it-1}\Delta\ \hat{\mu}_{it}-\hat{\lambda}_i)$
3. Panel t-statistics (non-parametric): $Z_t\equiv\left(\tilde{\sigma}^2\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{\kappa}^{-2}{}_{11,i}\hat{\mu}_{it-1}^2\right)^{-1/2}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{\kappa}^{-2}{}_{11,i}(\hat{\mu}_{it-1}\Delta\ \hat{\mu}_{it}-\hat{\lambda}_i)$
4. Panel t-statistics (parametric): $Z_t^*\equiv\left(\tilde{s}_{N,T}^{*2}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{\kappa}^{-2}{}_{11,i}\hat{\mu}_{it-1}^2\right)^{-1/2}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{\kappa}^{-2}{}_{11,i}\hat{\mu}_{it-1}^2\Delta\hat{\mu}_{it}^*$
5. Group ρ-statistics: $\tilde{Z}_p\equiv TN^{-1/2}\sum_{i=1}^{N}\left(\sum_{t=1}^{T}\hat{\mu}_{it-1}^2\right)^{-1}$ $\sum_{t=1}^{T}(\hat{\mu}_{it-1}\Delta\hat{\mu}_{it}-\hat{\lambda}_i)$
6. Group t-statistics (non-parametric): $\tilde{Z}_t\equiv N^{-1/2}\sum_{i=1}^{N}\left(\hat{\sigma}_i^2\sum_{t=1}^{T}\hat{\mu}_{it-1}^2\right)^{-1/2}\sum_{t=1}^{T}(\hat{\mu}_{it-1}\Delta\hat{\mu}_{it}-\hat{\lambda}_i)$
7. Group t-statistics (parametric): $\tilde{Z}^*{}_t\equiv N^{-1/2}\sum_{i=1}^{N}\left(\sum_{t=1}^{T}\tilde{s}^{*2}\hat{\mu}_{it-1}^{2*}\right)^{-1/2}\sum_{t=1}^{N}\hat{\mu}^*{}_{it-1}\Delta\hat{\mu}_{it}^*$

where $\hat{\lambda}_i=\frac{1}{2}(\hat{\sigma}_i^2-\hat{s}_i^2)$ and $\tilde{s}_{N,T}^{*2}=\frac{1}{N}\sum_{i=1}^{N}\hat{s}^{*2}$

The first four statistics are within dimension-based statistics and the rest are between dimension statistics. In his paper Pedroni (1999) describes the seven test statistics. "The first of the simple panel cointegration statistics is a type of non-parametric variance ratio statistics. The second is a panel version of a nonparametric statistics that is analogous to the familiar Phillips–Perron rho-statistics. The third statistics is also non-parametric and is analogous to the Philips and Perron t-statistics. The fourth statistics is the simple panel cointegration statistics which is corresponding to the augmented Dickey–Fuller-statistics" (Pedroni 1999, p. 658). The rest of the statistics are based on a group mean approach. "The first of these is analogous to the Philips and Perron rho-statistics, and the last two analogous to the Phillips and Perron statistics and the augmented Dickey–Fuller statistics, respectively"(Pedroni, 1999, p. 658). Pedroni (2004) examined the small sample power properties of his seven test statistics. He found that the size distortion is small and the power is high for $T>100$. For smaller T, he shows that the group ADF test has the best power properties followed by the panel ADF test; the panel variance test and group rho test perform poorly. The panel Larsson et al. (2001) likelihood ratio test statistics is derived from the average of the individual likelihood ratio test statistics of Johansen (1995). The multivariate cointegration trace test of Johansen (1988, 1995) is engaged to investigate each individual cross-section system autonomously, in such a way that allows heterogeneity in each cross-sectional unit root for the said panel. The process of data generation for each of the groups is characterized by the following heterogeneous VAR (p_i) model:

$$Y_{i,t}=\sum_{j=1}^{p_i}\Lambda_{i,j}Y_{i,t-j}+\varepsilon_{i,t} \tag{13}$$

where $i=1,\ldots,N;\ t=1,\ldots T$

For each one, the value of $Y_{i,-j+1},\ldots Y_{i,0}$ is considered as fixed and $\varepsilon_{i,t}$ are independent and identically distributed (normally distributed): $\varepsilon\sim N_K(0,\Omega_i)$, where Ω_i is the cross-correlation matrix in the error terms: $\Omega_i=E(\varepsilon_{i,t},\varepsilon_{i,t}^{'})$. Eq. (10) can be modified in vector error correction model (VECM) as given below:

$$\Delta Y_{i,t}=\Pi_i Y_{i,t-1}+\sum_{j=1}^{p_i-1}\Gamma_{i,j}\Delta Y_{i,t-j}+\varepsilon_{i,j} \tag{14}$$

where $\Pi_i=\Lambda_{i,1}+\cdots+\Lambda_{pi}-1$and $\Gamma_{i,j}=\Lambda_{i,j}-\Lambda_{i,j-1}$, Π_i is of order $(k\times k)$. If Π_i is of reduced rank:rank $(\Pi_i)=r_i$, which can be decomposed into $\Pi_i=ab'$, where α_i and β_i are of order $(k\times r_i)$ and of full column rank that represents the error correction form. The null hypotheses of panel Larsson et al. (2001) rank test are:

$H_0=rank(\Pi_i)=r_i\le r$ for all $i=1,\ldots, N$ against

$H_a=rank(\Pi_i)=k$ for all $i=1,\ldots,N$

The procedure is in sequences like individual trace test process for cointegration rank determination. First, we test for $H_0=rank(\Pi_i)=r_i\le r,\ \ r=0$, if null hypothesis of no cointegration is accepted, this shows that there is no cointegration relationship ($rank\ (\Pi_i)=r_i=0$) in all cross-sectional groups for the 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 $(\Pi_i)=r\le 0\ \ (0<r<k)$ implies that there is at least one cross-sectional unit in the panel which has rank $(\Pi_i)=r>0$. The likelihood ratio trace test statistics for group i is as follows:

$$LR_{iT}\{H(r)/H(k)=-2\ln Q_{iT}(H(r)/H(k)=-T\sum_{l=r+1}^{p}\ln(1-\lambda_{li}') \tag{15}$$

where λ_l' is the l^{th} largest eigen value in the i^{th} cross-section unit. The LR-bar statistics is calculated as the average of individual trace statistics:

$$L\overline{R}_{iT}[H(r)/H(k)]=\frac{1}{N}\sum_{i=1}^{n}LR_{iT}[H(r)/H(k)] \tag{16}$$

Finally, a modified version of the above equation is defined as

$$\lambda_{\overline{LR}}[H(r)/H(k)] = \frac{\sqrt{N}(\overline{LR}\_{NT}[H(r)/H(k)]) - E(Z_k)}{\sqrt{Var(Z_k)}} \tag{17}$$

where $E(Z_k)$ and $\mathrm{Var}(Z_k)$ are mean and variance of the asymptotic trace statistics, which can be obtained from simulation. The Larsson et al. (2001) prove the central limit theorem for the standard LR-bar statistics that under the null hypothesis, $\lambda_{\overline{LR}} \Rightarrow N(0,1)$ as N and $T \to \infty$ in such a way that $\sqrt{NT^{-1}} \to 0$, under the assumption that there is no cross-correlation in the error terms, that is given below:

$$E(\varepsilon_{i,t}) = 0 \text{ and } E(\varepsilon_{i,t}, \varepsilon_{j,t}) = \begin{cases} \Omega_i & for\ i=j,\ i \neq j \\ 0 \end{cases}$$

Larsson et al. (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.

4.3. Estimation of panel cointegration regression

If all the variables are cointegrated, the next step is to estimate the associated long-run cointegration parameters. In the presence of cointegration, the OLS estimator is known to yield biased and inconsistent results. For this reason, several estimators have been proposed. For example, Kao and Chiang (2000) argue that their parametric panel Dynamic OLS (DOLS) estimator (that pools the data along the within dimension of the panel) is promising in small samples and performs well in general in cointegrated panels. However, the panel DOLS of Kao and Chiang (2000) does not consider the importance of cross-sectional heterogeneity in the alternative hypothesis. To allow for cross-sectional heterogeneity in the alternative hypothesis, endogeneity and serial correlation problems to obtain consistent and asymptotically unbiased estimates of the cointegrating vectors, Pedroni (2000; 2001) proposed the group mean fully modified OLS (FMOLS) estimator for cointegrated panels.

Following Pedroni (2001), the FMOLS technique generates consistent estimates in small samples and does not suffer from large size distortions in the presence of endogeneity and heterogeneous dynamics. The panel FMOLS estimator for the coefficient β is defined as:

$$\hat{\beta} = N^{-1} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} (y_{it} - \overline{y})^2 \right)^{-1} \left(\sum_{t=1}^{T} (y_{it} - \overline{y}) \right) z_{it}^{*} - T\hat{\eta}_i \tag{18}$$

where $z_{it}^{*} = (z_{it} - \overline{z}) - \frac{\hat{L}_{21i}}{\hat{L}_{22i}} \Delta y_{it},\ \hat{\eta}_i \equiv \hat{\Gamma}_{21i} + \hat{\Omega}^0_{21i} - \frac{\hat{L}_{21i}}{\hat{L}_{22i}} (\hat{\Gamma}_{22i} + \hat{\Omega}^0_{22i})$ and $\hat{L}_i$ is a lower triangular decomposition of $\hat{\Omega}_i$. The associated t-statistics give:

$$t_{\hat{\beta}^*} = N^{-1/2} \sum_{i=1}^{N} t_{\hat{\beta}^*}, i \text{ where } t_{\hat{\beta}^*}, i = (\hat{\beta}_i^* - \beta_0) \left[\hat{\Omega}^{-1}{}_{11i} \sum_{t=1}^{T} (y_{it} - \overline{y})^2 \right]^{1/2} \tag{19}$$

4.4. Panel VECM causality

If evidence of cointegration is found, a panel vector error correction model (VECM) developed by Pesaran et al. (1999) can be estimated to perform Granger causality tests. Evidence of cointegration between variables implies that there exists causality in at least one direction (Granger, 1969). Results of cointegration are important because they provide conformity about the existence of error correction mechanism by which changes in dependent variables are the function of the level of disequilibrium in the cointegrating relationship and changes in other independent variables. The following VECM models are used to test the causality relation between variables:

$$\Delta EN_{it} = \beta_{2j} + \sum_{m=1}^{p} \beta_{im} \Delta EN_{it-m} + \sum_{m=1}^{p} \vartheta_{im} \Delta Y_{it-m} + \sum_{m=1}^{p} \kappa_{im} \Delta TR_{it-m} + \sum_{i=m}^{p} \theta_{im} \Delta OP_{it-m} + \psi_1 ECT_{t-1} + \upsilon_{1t} \tag{20}$$

$$\Delta Y_{it} = \beta_{3j} + \sum_{m=1}^{p} \beta_{im} \Delta Y_{it-m} + \sum_{m=1}^{p} \vartheta_{im} \Delta EN_{it-m} + \sum_{m=1}^{p} \kappa_{im} \Delta T_{it-m} + \sum_{i=m}^{p} \theta_{im} \Delta OP_{it-m} + \psi_1 ECT_{t-1} + \upsilon_{1t} \tag{21}$$

$$\Delta T_{it} = \beta_{4j} + \sum_{m=1}^{p} \beta_{im} \Delta T_{it-m} + \sum_{m=1}^{p} \vartheta_{im} \Delta EN_{it-m} + \sum_{m=1}^{p} \kappa_{im} \Delta Y_{it-m} + \sum_{m=1}^{p} \lambda_{im} \Delta OP_{it-m} + \psi_1 ECT_{t-1} + \upsilon_{1t} \tag{22}$$

$$\Delta OP_{it} = \beta_{5j} + \sum_{m=1}^{p} \beta_{im} \Delta OP_{it-m} + \sum_{m=1}^{p} \vartheta_{im} \Delta EN_{it-m} + \sum_{m=1}^{p} \kappa_{im} \Delta Y_{it-m} + \sum_{m=1}^{p} \lambda_{im} \Delta T_{it-m} + \psi_1 ECT_{t-1} + \upsilon_{1t} \tag{23}$$

In the above models, Δ is the lag operator and ECT_{t-1} is one period lagged error term used to identify long-run causality between the variables. Short-run causality is estimated by testing various hypotheses. For example, short-run causality from Y to EN is calculated by testing hypothesis: $H_0 : \vartheta_{im} = 0$ for all i and m. The rejection of this hypothesis implies that Y is causing EN in the short run. A similar hypothesis procedure will be employed to test various hypotheses. The significance of the error correction terms in each set of equations can be tested using t-tests. Short-run dynamics can be tested using Granger causality F tests.

5. Empirical results and discussion

5.1. Panel unit root results

Table 1 presents the estimated results of unit root tests at level and first difference. These results are calculated by applying three panel unit root tests: LLC, IPS and MW on each selected variable without trend and with trend. Our empirical findings reveal that all variables are non-stationary in their level form. However, all the series are stationary at first difference. Thus, we reject the null hypothesis of non-stationary at 1% level of significance and conclude that all series are integrated of order one in the panel of 15 Asian countries.

5.2. Panel cointegration results

This unique order of integration of the variables helps us to apply panel cointegration approach to examine long-run relationship between the variables for the selected panel. The results of Pedroni (1999, 2004) panel cointegration tests are reported in Table 2. Pedroni uses four within dimension (panel) test statistics and three between dimension (group) statistics to check whether the selected panel data are cointegrated. Within dimension statistics contain the estimated values of test statistics based on estimators that pooled the autoregressive coefficient across different cross-sections for the unit root test on the estimated residuals. Between dimensions on the other hand, report the estimated values of test statistics based on estimators that average individually estimated coefficients for each cross-section. The results of within dimensions tests and between dimensions tests show that the null hypothesis of no cointegration can be rejected in most cases. Therefore, energy consumption, income growth and trade

Table 1
Panel unit root test results.

Variables	At level				At first difference			
	Without trend	*P*-value	With trend	*P*-value	Without trend	*P*-value	With trend	*P*-value
LLC test								
EN_{it}	1.045	0.852	3.218	0.999	−6.031	0.000	−6.990	0.000
Y_{it}	1.407	0.920	0.479	0.684	−5.101	0.000	−4.321	0.000
T_{it}	−0.418	0.337	0.728	0.766	−3.189	0.000	−4.416	0.000
EP_{it}	−0.528	0.635	1.108	0.866	−4.925	0.000	−4.135	0.000
IPS test								
EN_{it}	3.953	1.000	3.733	0.999	−4.359	0.000	−4.509	0.000
Y_{it}	4.805	1.000	2.868	0.997	−3.923	0.000	−4.317	0.000
T_{it}	0.262	0.603	−1.191	0.116	−4.662	0.000	−5.368	0.000
EP_{it}	2.204	0.986	1.744	0.959	−3.342	0.000	−4.945	0.000
MW (ADF) test								
EN_{it}	14.672	0.991	10.563	0.999	68.29	0.000	68.35	0.000
Y_{it}	15.135	0.989	9.9048	0.999	59.98	0.000	63.69	0.000
T_{it}	26.348	0.657	36.515	0.191	68.84	0.000	83.13	0.000
EP_{it}	13.476	0.995	14.701	0.991	53.18	0.000	77.29	0.000
MW (PP) test								
EN_{it}	14.003	0.994	14.720	0.991	239.71	0.000	218.07	0.000
Y_{it}	17.380	0.967	10.169	0.999	168.32	0.000	151.33	0.000
T_{it}	19.475	0.929	25.847	0.682	143.49	0.000	127.91	0.000
EP_{it}	11.730	0.755	15.424	0.987	153.55	0.000	118.57	0.000

Table 2
Pedroni panel cointegration test results.

Test	Panel υ-statistic	Panel σ-statistic	Panel $\rho\rho$-statistic	Panel *adf*-statistic	Group σ-statistic	Group $\rho\rho$-statistic	Group *adf*-statistic
Statistic	−0.493	−0.167	−4.043	−2.703	1.187	−2.743	−1.976
P-value	0.689	0.433	0.000	0.003	0.882	0.003	0.024

Note: an intercept and trend is included in the cointegration equations.

Table 3
Larsson et al. panel cointegration test results (variables: EN_{it}, Y_{it}, T_{it}, OP_{it}).

Country	Likelihood ratio test statistics				
	$R=0$	$R\leq 1$	$R\leq 2$	$R\leq 3$	Rank
Pakistan	84.317*	54.142*	29.699	15.324	2
India	91.096*	51.706*	21.907	11.889	2
Bangladesh	77.630*	48.866*	25.617	10.399	2
Sri Lanka	85.854*	52.774*	28.866	16.529	2
Philippines	94.864*	58.060*	27.491	12.760	2
Thailand	85.127*	49.394*	25.364	8.4742	2
Indonesia	87.708*	48.072*	29.936*	16.418	3
Malaysia	85.942*	51.962*	27.880*	11.095	3
China	88.296*	53.274*	24.831	10.268	2
Japan	89.407*	55.064*	30.596*	10.913	3
Jordan	100.788*	58.855*	32.229*	13.100	3
Iran	108.19*	61.419*	30.981*	14.164	3
Korea Dem	99.354*	64.962*	34.486*	11.968	3
Nepal	73.290*	40.811	27.030	13.568	1
Vietnam	104.909*	60.442*	29.502	10.364	2
Panel (T_{LR})	8.3603*	3.4029*	1.2492	0.9431	2

* The rejection of null hypothesis at 5% significance level.

openness are cointegrated in our selected panels of Asian countries for the period 1980–2011. Likelihood-based panel cointegration test results provide additional support for the presence of cointegration between variables.

Table 3 presents the results of Larsson et al. (2001) panel cointegration derived on the basis of likelihood test statistics by Johansen (1995). In the case of individual country likelihood ratio test statistics, the most common selected rank is $r=2$ in 8 out of 15 countries (Pakistan, India, Bangladesh, Sri Lanka, Philippines, Thailand, China, Vietnam) indicates the acceptance of the co-integrating hypothesis in these countries. However, $r=3$ in 6 countries, namely, Malaysia, Japan, Jordan, Iran, Indonesia and Korea Dem. Nepal is the only country

Table 4
FMOLS country-specific results (EN_t: dependent variable).

Country	Variables	Y_{it}	T_{it}	OP_{it}	Constant
Pakistan	Coefficient	0.287	−0.009	−0.164	3.615
	P-value	0.000	0.634	0.000	0.000
India	Coefficient	0.366	−0.040	−0.107	3.151
	P-value	0.000	0.111	0.003	0.000
Indonesia	Coefficient	0.782	0.041	−0.056	0.267
	P-value	0.000	0.522	0.326	0.749
Iran	Coefficient	0.526	0.063	−0.203	2.403
	P-value	0.003	0.094	0.000	0.082
Japan	Coefficient	0.269	0.005	−1.066	3.197
	P-value	0.484	0.936	0.033	0.257
Jorden	Coefficient	0.100	0.188	−0.289	6.669
	P-value	0.800	0.154	0.000	0.029
Korea Dem.	Coefficient	1.458	−0.311	0.189	−4.367
	P-value	0.008	0.986	0.314	0.001
Malaysia	Coefficient	0.514	0.082	−0.605	1.314
	P-value	0.050	0.363	0.007	0.395
Nepal	Coefficient	0.329	0.083	0.0003	3.475
	P-value	0.111	0.043	0.993	0.013
Philippines	Coefficient	0.220	0.105	−0.542	5.663
	P-value	0.000	0.000	0.000	0.000
Sri Lanka	Coefficient	0.613	0.051	0.023	1.049
	P-value	0.094	0.709	0.862	0.687
Thailand	Coefficient	0.353	0.161	−0.688	2.088
	P-value	0.174	0.154	0.000	0.224
Vietnam	Coefficient	0.235	0.180	0.141	4.498
	P-value	0.409	0.051	0.429	0.013
China	Coefficient	0.546	0.278	−0.293	2.202
	P-value	0.019	0.020	0.072	0.151

where $r=1$ is selected. Thus, on the basis of these results, it can be concluded that a cointegrating relationship exists between the variables in each selected Asian country. In the case of the panel, the maximum rank is $r=2$. Hence, the result of Larsson et al. (2001) panel cointegration indicates the existence of at least two cointegrating vectors in the Asian panel. Finally, panel cointegrating results confirm a stable long-run relationship between energy consumption, income per capita, trade openness, and energy prices in 15 Asian countries.

5.3. DOLS and FMOLS estimates

Tables 4 and 5 display the results of FMOLS and DOLS at individual level. The difference between the two approaches is not very marked in terms of sign, magnitude and statistical significance. For each approach the income coefficient is positive and significant in Pakistan, India, Indonesia, Iran, Korea Democratic Republic, Malaysia, Philippines, Sri Lanka and China whereas positive and insignificant in Japan, Jorden, Nepal, Thailand and Vietnam. The positive coefficient of income growth suggests that increase in income growth leads to increase in energy consumption in selected Asian countries. Trade openness increases the demand for energy consumption in Indonesia, Iran, Japan, Jorden, Malaysia, Nepal, Philippines, Sri Lanka, Thailand, Vietnam and China. However, in the case of Pakistan, India and, Korea Democratic Republic, the coefficient of trade openness is found to be negative and insignificant. Increase in oil price decreases the demand for energy consumption in almost all selected countries except Korea Democratic Republic, Nepal and Sri Lanka, where the coefficient is reported negative and insignificant. The results of FMOLS and DOLS at group level are reported in Table 6. Results show that all coefficients are statistically significant and their signs are according to economic theory. Results of FMOLS indicate that 1% increase in income growth increases energy consumption per capita by about 0.42%. Similarly, a 1% increase in trade openness increases energy consumption per capita by 0.06% in the Asian countries. The effect of oil prices on energy consumption is found to be negative and indicates that 1% point increase in real oil prices decreases energy usage by about 0.19% in a selected panel of the Asian countries.

Table 5
DOLS country-specific results (EN_t: dependent variable).

Country	Variables	Y_{it}	T_{it}	OP_{it}	Constant
Pakistan	Coefficient	0.303	−0.008	−0.166	3.487
	P-value	0.000	0.745	0.000	0.000
India	Coefficient	0.363	−0.055	0.106	3.149
	P-value	0.000	0.001	0.000	0.000
Indonesia	Coefficient	0.828	0.101	0.033	−0.065
	P-value	0.000	0.336	0.707	0.960
Iran	Coefficient	0.485	0.061	−0.207	2.776
	P-value	0.027	0.191	0.001	0.115
Japan	Coefficient	0.037	0.120	−1.477	4.257
	P-value	0.927	0.159	0.015	0.146
Jorden	Coefficient	0.043	0.217	−0.267	6.149
	P-value	0.962	0.428	0.003	0.357
Korea Dem.	Coefficient	1.433	−0.392	0.096	−3.945
	P-value	0.000	0.001	0.853	0.199
Malaysia	Coefficient	0.634	0.038	−0.508	0.602
	P-value	0.027	0.688	0.033	0.719
Nepal	Coefficient	1.171	0.108	0.178	−2.016
	P-value	0.067	0.109	0.153	0.614
Philippines	Coefficient	0.232	0.806	−0.129	5.633
	P-value	0.000	0.000	0.000	0.000
Sri Lanka	Coefficient	0.774	0.010	0.092	−0.022
	P-value	0.041	0.929	0.517	0.993
Thailand	Coefficient	1.440	−0.640	−1.038	−5.631
	P-value	0.000	0.000	0.000	0.010
Vietnam	Coefficient	1.193	0.048	−0.913	−3.245
	P-value	0.000	0.051	0.000	0.055
China	Coefficient	0.054	0.843	−1.006	5.926
	P-value	0.073	0.000	0.000	0.000

5.4. Panel causality results

Table 7 reports the results of short-run and long-run Granger causality tests. With respect to Eq. (20), the coefficient of lagged error-correction term is negative and significant at 1% level but with a relatively low speed of adjustment to long-run equilibrium. Negative error correction term confirms the existence of the long-run Granger causality running from income, trade openness and oil prices to energy consumption. With respect to short-run causality tests, there is evidence of Granger causality running from income growth to energy consumption, trade openness to energy consumption and oil prices to energy consumption. From Eq. (21), error correction term is negative and significant which suggests that income growth responds to long-run equilibrium and confirms the long-run causality running energy consumption, trade openness and oil prices to income growth. Over a short period of time, there is evidence of Granger causality running from energy to income, trade to income and energy prices to income. The significant and negative error correction term in Eq. (22) confirms the presence of long-run causality running from energy consumption, income and energy prices to trade openness. In the short run, all coefficients are found to be significant and indicate that energy consumption Granger cause trade, income Granger cause trade and energy prices Granger cause trade. In Eq. (23), both the long-run and short-run confidents are significant, thus indicating the acceptance of Granger causality running from

Table 6
FMOLS and DOLS panel results (EN_{it}: dependent variable).

Variables	**FMOLS**		**DOLS**	
	Coefficient	***P*-value**	**Coefficient**	***P*-value**
Y_{it}	0.419	0.000	0.408	0.000
T_{it}	0.058	0.002	2.158	0.031
OP_{it}	−0.191	0.000	−1.189	0.000

Table 7
VECM based Granger causality results.

Dependent variables	Source of causation (independent variables)				
	ΔEN_{it} Short run	ΔY_{it}	ΔT_{it}	ΔOP_{it}	ECT_{t-1} Long run
Eq. (20) ΔEN_{it}	–	3.79 (0.02)	5.98 (0.00)	2.87 (0.07)	−0.14* [4.36]
Eq. (21) ΔY_{it}	3.15 (0.05)	–	4.49 (0.00)	3.46 (0.03)	−0.36* [2.98]
Eq. (22) ΔT_{it}	5.43 (0.00)	4.76 (0.00)	–	5.94 (0.00)	0.18* [3.83]
Eq. (23) ΔOP_{it}	4.01 (0.00)	3.92 (0.00)	2.94 (0.06)	–	0.42** [2.90]

Note: Wald *F*-statistics reported with respect to short-run changes in the independent variables. ECT represents the coefficient of the error correction term. Values in ( ) are *p*-values and values in [ ] are *t*-ratios.

* Significance at 1% level.

** Significance at 5% level.

energy consumption, income growth and trade openness to energy prices.Results provide evidence of feedback relationship between energy consumption and income, energy consumption and trade. Evidence of feedback relationship between energy consumption and energy prices is also observed. These results suggest that economic growth and flow of trade increase the demand for energy sources in Asia. Effective utilization of energy resources, exploring new and alternative sources of energy are necessary to reap optimal fruits of trade and economic growth.

6. Concluding remarks

This paper explores the relationship between economic growth, trade openness and energy consumption using data of 15 Asian countries over the period 1980–2011. In doing so, we have applied panel unit root tests to examine the integrating properties of the variables. To examine cointegration between variables, we have applied Pedroni cointegration and likelihood-based panel cointegration approaches. The Granger causality are applied to examine the direction of causality between variables in the Asian countries.

Empirical results indicate that all variables are integrated at I(1) confirmed by panel unit root tests and the same inference is drawn about cointegration between economic growth, trade openness and energy consumption. The FMOLS and DOLS estimation analysis reveals a positive relationship between energy consumption and income growth, energy consumption and trade openness whereas an inverse relationship between energy consumption and energy prices is observed. The causality analysis confirms the existence of feedback causality between economic growth and energy consumption, trade openness and energy consumption in the Asian countries.

An important policy implication based on the general result of the study is that to regulate the flow of trade and expansion in economic growth, care should be taken in making energy conservation policies in the Asian countries. Another policy implication based on the long-run result that an increase in economic growth leads to energy consumption more than trade openness is that decision makers need to announce an integrated energy policy aimed at increasing energy use efficiently by lowering energy consumption for a given level of economic growth. For this purpose, not only national factors such as energy supply infrastructure, energy efficiency considerations or institutional constraints, but also international developments should be taken into account. The government needs to allocate more resources to the development of new sources of energy and ensure sustainability of energy use. Further, Asian economies need to increase the scale of trade openness to get the benefits of advanced technologies from other developed economies of the world. The adoption of advanced technology lowers energy intensity and produces more output. Finally, inclusion of energy price variable in the model will provide new ways in the evaluation of public policies and technological innovations in the energy sector of the Asian countries.

References

Abosedra, S., Baghestani, H., 1989. New evidence on the causal relationship between U.S. energy consumption and gross national product. J. Energy Dev. 14, 285–292.
Akarca, A.T., Long, T.V., 1979. Energy and employment: a time series analysis of the causal relationship. Resour. Energy 5, 326–331.
Akarca, A.T., Long, T.V., 1980. On the relationship between energy and GNP: a re-examination. J. Energy Dev. 5, 326–331.
Amirat, A., Bouri, A., 2010. Energy and Economic Growth: The Algerian Case. Typewritten.
Apergis, N., Payne, E.J., 2010. Energy consumption and growth in South America: evidence from a panel error correction model. Energy Eco. 32, 1421–1426.
Canning, D., Pedroni, P., 2008. Infrastructure, long-run economic growth and causality tests for cointegrated panels. The Manchester School, Wiley, 76, pp. 504–527.
Chebbi, H. E., Boujelbene, Y., 2008. CO_2 emissions, energy consumption and economic growth in Tunisia. In: Proceedings of the 12th Congress of the European Association of Agriculture Economics (EAAE2008). Ghent, Belgium 26–29 August 2008. Ghent, Belgium.
Cole, M.A., 2006. Does trade liberalization increase energy use? Eco. Lett. 92, 108–112.
Dedeoğlu, D., Kaya, H., 2013. Energy use, exports, imports and GDP: new evidence from the OECD countries. Energy Policy 57, 469–476.
Ghani, G.M., 2012. Does trade liberalization affect energy consumption? Energy Policy 43 (4), 285–290.
Giles, J.A., Williams, C.L., 2000a. Export-led growth: a survey of the empirical literature and some non-causality results. J. Int. Trade Eco. Dev. 9, 261–337.
Giles, J.A., Williams, C.L., 2000b. Export-led growth: a survey of the empirical literature and some non-causality results. J. Int. Trade Eco. Dev. 9, 445–470.
Granger, C.W.J., 1969. Investigating causal relationships by econometric models and cross-spectral methods. Econometica 37, 424–438.
Hadri, K., 2000. Testing for stationarity in heterogeneous panel data. Econom. J. 3, 148–161.
Halicioglu, F. 2010. An Econometric Study of CO_2 Emission, Energy Consumption, Income and Foreign Trade. MPRA Paper no.1 1457.
Heston, A., Summers, R., Aten, B., 2013. Penn World Table Version 8.0. Center of comparisons of production, income and prices at the University of Pennsylvania.
Hossain, Md. S., 2012. Multivariate Granger causality between economic growth, electricity consumption, exports and remittances for the panel of three SAARC countries. Global J. Manag. Bus. Res. 12, 41–54.
Im, K., Pesaran, H., Shin, Y., 1997. Testing for Unit Roots in Heterogeneous Panels. University of Cambridge, Mimeo.
International Energy Agency, World Energy Outlook, 2012. IEA, Paris.
Jena, Pradyot Ranjan , Grote, Ulrike, 2008. Growth-trade-environment nexus in India. In: Proceedings of the German Development Economics Conference, Zurich 2008 18, Verein für Socialpolitik, Research Committee Development Economics.
Johansen, S., 1988. Statistical analysis of cointegration vectors. Journal of Economic Dynamics and Control 12 (2–3), 231–254.
Johansen, S., 1995. Likelihood-Based Inference in Cointegrated Vector Autregressive Models. Oxford University Press, Oxford.
Karlsson, S., Löthgren, M., 2000. On the power and interpretation of panel unit root tests. Econ. Lett. 66, 249–255.
Kao, C., Chiang, M.-H., 2000. On the estimation and inference of a cointegrated regression in panel data. In: Baltagi B. H. (Ed.), Advances in Econometrics: Nonstationary Panels, Panel Cointegration and Dynamic Panels, vol. 15, pp. 179–222.
Kraft, J., Kraft, A., 1978. On the relationship between energy and GNP. J. Energy Dev. 3, 401–403.
Larsson, R., Lyahagen, J., Lothgren, M., 2001. Likelihood-based cointegration tests in heterogeneous panels. Econometrics J. 4, 109–142.
Lean, H.H., Smyth, R., 2010. Multivariate granger causality between electricity generation, exports and GDP in Malaysia. Energy Eco. 35, 3640–3648.
Levin, A., et al., 1993. Unit root Tests in Panel Data: Asymptotic and Finite Sample Properties. University of California, San Diego (Discussion paper).
Loganathan, N., Subramaniam, T., 2010. Dynamic co integration link between energy consumption and economic performance: empirical evidence from Malaysia. Int. J. Trade Eco. Finance 1, 3.
Maddala, G., Wu, S., 1999. A Comparative Study of Unit Root Tests with Panel Data and a new Simple Test. Oxford Bulletin of Economics and Statistics, Special Issue, November 61.
Mazumder, P., Marathe, A., 2007. Causality relationship between electricity consumption and GDP in Bangladesh. Energy Policy 35, 395–402.
Narayan, P.K., Smyth, R., 2009. Multivariate granger causality between electricity consumption, exports and GDP: evidence from a panel of middle eastern countries. Energy Policy 37, 229–236.
Narayan, P.K., Narayan, S., Prasad, A., 2008. A structural VAR analysis of electricity consumption and real GDP: evidence from the G7 countries. Energy Policy 36, 2765–2769.
Nondo C., Kahsai M., 2009. Energy Consumption and Economic Growth: Evidence from COMESA Countries. Selected Paper Prepared for Presentation at the Southern Agricultural Economics Association Annual Meeting, Atlanta, Georgia, January 31–February 3, 2009.
Noor, S., Siddiqi, M.W., 2010. Energy consumption and economic growth in south Asian countries: a co-integrated panel analysis. Int. J. Hum. Soc. Sci. 5, 14.
Ozturk, I., 2010. A literature survey on energy-growth nexus. Energy Policy 38, 340–349.
Payne, J.E., 2010. Survey of the international evidence on the causal relationship between energy consumption and growth. J. Eco. Stud. 37, 53–95.
Pedroni, P. 1999. Critical Values for Cointegration Tests in Heterogeneous Panels with Multiple Regressors, Oxford Bulletin of Economics and Statistics, Special Issue, 61, pp. 653–670.
Pedroni, P., 2000. Fully-modified OLS for heterogeneous cointegration panel. Adv. Econometrics 15, 93–130.
Pedroni, P., 2001. Purchasing power parity tests in cointegrated panels. Rev. Eco. Stat. 83, 727–731.

Pedroni, P., 2004. Panel cointegration: asymptotic and finite samples properties of pooled time series tests with an application to the ppp hypothesis. Econometric Theory 20, 597–625.
Pesaran, H.M., Shin, Y., Smith, R.P., 1999. Pooled mean group estimation of dynamic heterogeneous panels. J. Am. Stat. Assoc. 94, 621–634.
Sadorsky, P., 2011. Trade and energy consumption in the Middle East. Energy Eco. 2010 (12), 12.
Sadorsky, P., 2012. Energy consumption, output and trade in South America. Energy Econ. 34, 476–488.
World Trade Report, 2012. Trade and Public Policies: A Closer Look at Non-Tariff Measures in the 21st Century World Trade Organization, Geneva.
World Bank, World Development Report 2013, Washington, DC.
Yang, H.Y., 2000. A note on the causal relationship between energy and GDP in Taiwan. Energy Eco. 22, 309–317.
Yu, E.S.H., Choi, J.Y., 1985. The causal relationship between energy and GNP: an international comparison. J. Energy Dev. 10, 249–272.
Zahid, K., 2008. Energy-GDP relationship: a casual analysis for the five countries of South Asia. Appl. Econometrics Int. Dev. 8 (1), 167–180.