Environmental Science and Pollution Research (2019) 26:14627–14637
https://doi.org/10.1007/s11356-019-04699-w

RESEARCH ARTICLE

Electricity consumption and economic growth nexus in China: an autoregressive distributed lag approach

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Received: 15 January 2019 / Accepted: 25 February 2019 / Published online: 15 March 2019

Abstract

This study attempts to investigate the relationship among electricity consumption, economic growth, and employment in China. Distinct from most of the previous studies, our empirical research identifies a long-run equilibrium cointegration relationship among the three covariates during the period of 1971–2009 with the recently developed autoregressive distributed lag (ARDL) bounds testing approach. The parameters are estimated through a long-run static solution of the estimated ARDL model and short-run dynamic solutions of the error correction model. The estimated models successfully pass diagnostic tests and both the long-run and short-run elasticities are found to be statistically significant. The study also indicates the existence of short-run and long-run causalities from electricity consumption and employment to economic growth. Results of this study show that electricity serves as an important driver of economic growth. Based on these results, several policy prescriptions on energy use and economic development are suggested for China.

Keywords Electricity consumption · Economic growth · China · Granger causality · ARDL model · Cointegration

Introduction

As the most flexible form of energy, electricity plays a vital role in the modern socioeconomic development (Shahbaz and Lean 2012). In recent years, the nexus of economic growth and electricity consumption has emerged as an issue of immense interest among economists and policymakers. This is largely triggered by the concerns about the increasing demand for electricity around the world, the environmental implications of electricity usage, and the ensuing need for conservation policies (Chandran et al. 2010; Hu and Lin 2008; Narayan and Prasad 2008). As regards policy making, the direction of causal relationship among these variables could have significant bearing (Asafu-Adjaye 2000; Ghosh 2002; Narayan and Prasad 2008; Narayan and Smyth 2005; Yoo 2005). No causality in either direction, or just a unidirectional causality from economic growth or working force to electricity depletion, implies that policies related to electricity conservation would not affect the economic growth. However, if this causal link is found in the reverse direction, then electricity consumption reduction could have a detrimental effect on economic growth and/or employment.

Since the initiation of market reforms in the late 1970s, China's economy has presented a rapid increase at an annual growth rate of around 9.7%. Together with the rapid economic growth, the demand for electricity has continuously maintained a highly increasing rate. Table 1 shows the trends in electricity consumption and economic growth in China during the period of 1971–2009. Although in the periods of 1980–1985 and 1995–2000, the growth of electricity consumption trends down slightly, it surged ahead at a remarkable pace in the shadow of its average 9.8% economic growth entering the new millennium. With electricity consumption rising from 127.1 TWh in 1971 to 3503.4 TWh in 2009, China has become the second largest electricity consumer in the world, next only to the USA.

Responsible editor: Muhammad Shahbaz

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Table 1 Growth of electricity generated and real GDP, China 1971/72–2008/2009

Period	Growth in electricity consumption (%)	Growth in real GDP (%)
1971/72–1974/75	9.1	5.6
1975/76–1979/80	9.0	6.5
1980/81–1984/85	6.1	10.7
1985/86–1989/90	9.3	7.9
1990/91–1994/95	9.8	12.3
1995/96–1999/2000	6.2	8.6
2000/01–2004/05	13.1	9.8
2005/06–2008/09	10.8	11.4

China has the world's third-largest coal reserves behind the USA and Russia, and massive hydroelectric resources. According to the World Energy Council (2010) statistics, China's verified coal reserves stood at 114.5 billion tons and accounted for 13.3% of the world's total reserves. Due to China's resource endowments, the electricity supply structure of China is dominated by coal. As of the year 2009, thermal power accounted for around 80.6% of the bulk of the electricity generated, followed by 16.7% from hydro, 1.9% from nuclear, and less than 0.8% from wind and solar power. China's coal-dominated energy structure has also led to a serious atmospheric environmental degradation. In 2009, the entire industrially emitted SO_2 amounted to 16.94 million tons, 55% of which is emitted from the electricity sector (National Bureau of Statistics of China 2010). Hence, China's authorities emphasized the importance of the energy efficiency and environmental conservation and they proposed the so-called "20/10" targets, a reduction of 20% of the energy intensity and 10% of the SO_2 emission during the 11th 5-year plan. In this situation, there is a need to reconsider the development strategy for the electricity sector of China and suggest a set of policy that can be adaptive to address short- and long-term electricity development by coordinating economic growth and environmental conservation in China.

Literature review

With motivating roots in energy conservation policies, a large body of literature has evolved to explore the causal direction between energy consumption and economic growth. During the last three decades, these studies originated from the paper by Kraft and Kraft (1978) to the recent studies such as Tsani (2010), Wang et al. (2011a, b), and Fuinhas and Marques (2012). Our empirical study is mainly focused on electricity sector, and for this reason, we will solely concentrate on the literature review about electricity consumption-economic growth nexus.

In general, empirical studies on the causality between electricity consumption and economic growth can be divided into two major groups, according to econometric methodologies employed in their analyses. Table 2 presents a summary of the selected empirical studies on the causal relationship between electricity consumption and economic growth. Based on the literature review of earlier empirical studies, the residual-based cointegration test associated with Engle and Granger (1987) and the maximum likelihood test based on Johansen (1988) and Johansen and Juselius (1990) have been widely employed to test for cointegration relationship between the two variables (Odhiambo 2009b). Empirical studies on electricity-growth nexus by employing these conventional econometric methods are summarized in Panel A of Table 2.

Given that these cointegration techniques may not be appropriate if data spans are too short (Lee and Chang 2005), empirical evidences from these studies have been mixed and remain ambiguous even though some of those studies are for the same country or region.

In more recent studies, the bounds testing approach to cointegration within an autoregressive distributed lag (ARDL) framework suggested by Pesaran (Pesaran and Shin 1999; Pesaran et al. 2001) has become a popular approach pertaining to causal relationship investigation. Compared to the Engle and Granger (1987) and Johansen and Juselius (1990) methods, the ARDL bounds testing approach has the advantage particularly when the sample size is small. In this section, we are mainly focused on reviewing the new developments in empirical studies during the last 10 years and bring the literature survey up to date. An overview of the findings of those studies that employ the ARDL approach is provided as follows (Panel B of Table 2).

Narayan and Smyth (2005) employed the ARDL approach to investigate the relationship between electricity consumption, employment, and real income in Australia during the period from 1966 to 1999. They found there to be a Granger causality running from employment and real GDP to electricity consumption. In the same manner, Narayan and Singh (2007) detected in the relation of GDP and the electricity consumption in the Fiji Islands and revealed a long-run unidirectional Granger causality performing from electricity consumption towards GDP. Wolde-Rufael (2006) examined the causal relationship between electricity consumption per capita and real GDP per capita for 17 African countries during the period of 1971 to 2001. His findings identified the Granger causality for only 12 countries—(1) unidirectional causality running from real GDP per capita to electricity consumption per capita in Cameroon, Ghana, Nigeria, Senegal, Zambia, and Zimbabwe; (2) reverse causality in Benin, Congo DR, and Tunisia; and (3) bidirectional causality in Egypt, Gabon, and Morocco. Squalli (2007) investigated the relationship between electricity consumption and economic growth for 11 OPEC member countries from 1980 to 2003, finding evidence

Table 2 Summary of selected empirical studies on electricity consumption-growth nexus

Author	Country/region	Period	Model	Result
Panel A: Engle–Granger or Johansen–Juselius cointegration test				
Abosedra et al. (2009)	Lebanon	Jan. 1995–Dec. 2005, monthly data	Bivariate model	$EC \rightarrow Y$
Akinlo (2009)	Nigeria	1980–2006	Bivariate model	$EC \rightarrow Y$
Altinay and Karagol (2005)	Turkey	1995–2000	Bivariate model	$EC \rightarrow Y$
Ghosh (2002)	India	1950–1997	Bivariate model	$Y \rightarrow EC$
Golam Ahamad and Nazrul Islam (2011)	Bangladesh	1971–2008	Bivariate model	$EC \rightarrow Y$ in the short run
Ho and Siu (2007)	Hong Kong SAR, China	1966–2002	Bivariate model	$EC \leftrightarrow Y$
Jumbe (2004)	Malawi	1970–1999	Bivariate model	$EC \leftrightarrow Y$
Lai et al. (2011)	Macao SAR	Q1 1999–Q4 2008, quarterly data	Bivariate model	$Y \rightarrow EC$
Lee and Chang (2005)	Taiwan	1954–2003	Bivariate model	$EC \rightarrow Y$
Lorde et al. (2010)	Barbados	1960–2004	Multivariate model of production side	$EC \rightarrow Y$ in the short run; $EC \leftrightarrow Y$ in the long run
Mozumder and Marathe (2007)	Bangladesh	1971–1999	Bivariate model	$Y \rightarrow EC$
Odhiambo (2009a)	South Africa	1971–2006	Multivariate model of production side	$EC \leftrightarrow Y$
Pao (2009)	Taiwan	Q1 1980–Q4 2007, quarterly data	Bivariate model	$Y \rightarrow EC$
Shiu and Lam (2004)	China	1971–2000	Bivariate model	$EC \rightarrow Y$
Soytas and Sari (2007)	Turkey	1968–2002	Multivariate model of production side	$EC \rightarrow Y$
Yoo (2005)	Korea	1970–2002	Bivariate model	$EC \leftrightarrow Y$
Yoo (2006)	4 ASEAN countries	1971–2002	Bivariate model	$Y \rightarrow EC$ (Indonesia and Thailand); $EC \leftrightarrow Y$ (Malaysia and Singapore)
Yoo and Kwak (2010)	7 South American countries	1975–2006	Bivariate model	$EC \rightarrow Y$ (Argentina, Brazil, Chile, Colombia and Ecuador); $EC \leftrightarrow Y$ (Venezuela); $EC \sim Y$ (Peru)
Yuan et al. (2007)	China	1978–2004	Bivariate model	$EC \rightarrow Y$
Yuan et al. (2008)	China	1963–2005	Multivariate model of production side	$EC \rightarrow Y$
Wang et al. (2011a)	East China: Jiangsu	1990–2007	Bivariate model	$EC \rightarrow Y$
Panel B: ARDL bounds testing approach				
Chandran et al. (2010)	Malaysia	1971–2003	Bivariate model and multivariate model of demand side	$EC \rightarrow Y$ in the short run
Ghosh (2009)	India	1970–2006	Multivariate model of production side	$Y \rightarrow EC$ in the short run
Jahangir Alam et al. (2012)	Bangladesh	1972–2006	Bivariate model	$EC \leftrightarrow Y$
Kouakou (2011)	Cote d'Ivoire	1971–2008	Bivariate model	$EC \rightarrow Y$
Kumar Narayan and Singh (2007)	Fiji	1971–2002	Multivariate model of production side	$EC \rightarrow Y$
Narayan and Smyth (2005)	Australia	1966–1999	Multivariate model of production side	$Y \rightarrow EC$
Odhiambo (2009b)	Tanzania	1971–2006	Bivariate model	$EC \rightarrow Y$
Ouédraogo (2010)	Burkina Faso	1968–2003	Multivariate model of production side	$EC \leftrightarrow Y$
Ozturk and Acaravci (2011)	11 MENA countries	1971–2006	Bivariate model	Israel: $Y \rightarrow EC$ in short run; Oman: $Y \rightarrow EC$ in short and long run; Egypt and Saudi Arabia: $EC \rightarrow Y$ in long run
Shahbaz and Lean (2012)	Pakistan	1972–2009	Multivariate model of production side	$EC \leftrightarrow Y$
Shahbaz et al. (2011)	Portugal	1971–2009	Multivariate model of production side	$EC \leftrightarrow Y$
Squalli (2007)	11 OPEC member countries	1980–2003	Bivariate model	$EC \rightarrow Y$ (Indonesia, Nigeria, UAE and Venezuela); $Y \rightarrow EC$ (Algeria, Iraq, Kuwait and Libya); $EC \leftrightarrow Y$ (Iran, Qatar and Saudi Arabia)
Wolde-Rufael (2006)	17 African countries	1971–2001	Bivariate model	$EC \rightarrow Y$ (Benin, Congo DR, Tunisia); $Y \rightarrow EC$ (Cameroon, Ghana, Nigeria, Senegal, Zambia and Zimbabwe); $EC \leftrightarrow Y$ (Egypt, Gabon and Morocco); $EC \sim Y$ (Algeria, Congo Rep., Kenya, South Africa and Sudan)
Wang et al. (2011b)	China	1972–2006	Multivariate model of production side	$EC \rightarrow Y$

Note: EC and Y denote electricity consumption and economic growth, respectively. ↔, →, and ~ denote bi-directional causality, unidirectional causality, and neutral causality, respectively of the energy-dependent economy for Indonesia, Nigeria, UAE, and Venezuela, the less energy-dependent economy for Iran, Qatar, and Saudi Arabia, and the energy-independent economy for Algeria, Iraq, Kuwait, and Libya. Ghosh (2009) explored the relationship between electricity supply, employment, and real GDP in India from 1970 to 1971 to 2005–2006, finding a unidirectional short-run causality of growth-led-electricity supply. Odhiambo (2009b) investigated the relationship of the two variables in Tanzania during the period of 1971–2006, reporting the result of causal flow from electricity consumption to economic growth. Chandran et al. (2010) modeled the electricity-growth nexus in Malaysia during the period of 1971–2003, showing a unidirectional causal flow of electricity consumption-led-growth. Ouédraogo (2010) explored the direction of causality between electricity consumption and economic growth in Burkina Faso for the period of 1968–2003, detecting a long-run bidirectional causal relationship between electricity use and real GDP. Other empirical researches include Kouakou (2011), which examined the causal relationship between the electric power industry and the economic growth of Cote d'Ivoire from 1971 to 2008; Ozturk and Acaravci (2011), which explored the short-run and long-run causality between electricity consumption and economic growth in the selected 11 Middle East and North Africa (MENA) countries during the period of 1971–2006; Shahbaz et al. (2011), which reconsidered the relationship among electricity consumption, economic growth, and employment in Portugal for the period of 1971–2009; Jahangir Alam et al. (2012), which examined the dynamic causality among economic growth, depletion of energy resource, electricity requirement, and carbon emissions in Bangladesh during the period of 1972–2006; and Shahbaz and Lean (2012), which re-examined the economic growth and electricity consumption nexus in Pakistan during the period of 1972–2009.

Following the evolvement of these recent researches, the present study employs the ARDL approach within a trivariate framework to re-examine the electricity-GDP nexus for China. The present study differs from earlier studies in two important aspects. Firstly, to the best of our knowledge, studies on electricity consumption are relatively few for China, limited to Shiu and Lam (2004), Chen et al. (2007), and Yuan et al. (2007). Furthermore, these earlier studies all tested for Granger causality between electricity consumption and real GDP within a bivariate framework. Although bivariate models can be employed when only scarce data are available, recently, its limitation in examining interactions has been criticized for specified bias due to the omission of relevant variables (Asafu-Adjaye 2000; Glasure 2002; Stern 1993, 1997, 2000). The most common approach in more recent studies is to incorporate capital and/or labor variables in modeling electricity consumption and economic growth within a multivariate causality framework. In line with the studies of Narayan and Smyth (2005), our empirical study includes a variable for employment in addition to electricity consumption and real GDP. Second, this study employs the ARDL bounds testing approach to cointegration, which is preferred over other alternative methods such as Engle and Granger (1987) and Johansen and Juselius (1990) cointegration tests for the simple reason that the bounds test has a better performance on small sample and can potentially produce more robust results from data taken over short-time spans. On this understanding, the present study has twofold novelty and its findings should be able to provide recommendations in regard to viable policy options for China.

The remainder of the paper is structured as follows. "Data and methodology" describes the model, econometric methodology, and data used in this study. "Empirical results" presents the unit root test results, the cointegration results, and the Granger causality test results. This is followed by policy analysis in "Policy implications" and finally "Conclusions".

Data and methodology

Data description

This empirical study adopts annual time series data in China spanning the period of 1971–2009, including total employment, electricity consumption (in kWh), and real GDP (base year 2000) that are denoted as *EM*, *EC*, and *GDP*, respectively. The data on real GDP and electricity consumption series comes from the World Development Indicators database, while the data on total employment series is collected from National Statistics Bureau of China. All variables are transformed into the natural logarithms to make first differences approximating growth rates.

Stationary

Although the ARDL bounds test to cointegration can be applied irrespective of whether the underlying variables are integrated to order $I(0)$ or $I(1)$, unit root tests might still be employed to ensure that none of the variables are integrated to order $I(2)$ or higher. Thus, our analysis conducted the unit root tests by employing the augmented Dickey-Fuller (ADF), Phillips-Perron (PP), and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) methods. In addition to the conventional unit root tests, we also use the Zivot and Andrews (ZA) method (Zivot and Andrews 1992), a structural break unit root test, to access the order of integration of underlying variables and to check the robustness of the results.

Perron (1989) pointed out that many macroeconomic time series with structural breaks indicate stationary fluctuation around a deterministic trend function if allowing a possible change in intercept and slope. Since standard unit root tests fail to test whether the series is a trend stationary process with a structural break, we should take the structural break into account when conducting unit root tests. For this reason, we use the endogenous break unit root test developed by Zivot and Andrews (1992), in order to capture the effect of any possible structural shift of underlying series over the studying period. There are two models used in our empirical study. Model A tests for a unit root against the alternative of a trend stationary process with a structural break in the intercept, while Model C tests for a unit root against the abovementioned alternative in both slope and intercept. Specifically, Model A takes the following form (Eq. (1)):

$$y_t = \mu + \alpha y_{t-1} + \beta t + \theta DU_t + \sum_{j=1}^{k} c_j \Delta y_{t-j} + \varepsilon_t \tag{1}$$

while Model C has the following form (Eq. (2)):

$$y_t = \mu + \alpha y_{t-1} + \beta t + \theta DU_t + \gamma DT_t + \sum_{j=1}^{k} c_j \Delta y_{t-j} + \varepsilon_t \tag{2}$$

where Δ denotes the difference operator, ε_t is a white noise term with variance σ^2, and $t = 1, ..., T$ denotes an index of time. The terms Δy_{t-j} allow for serial correlation and ensure that the disturbance term is white noise. DU_t and DT_t are dummy variables representing mean and trend shifts, respectively; $DU_t = 1$ if $t > T_B$, and 0 otherwise; $DT_t = t\text{-}T_B$ if $t > T_B$, and 0 otherwise. The breakpoint is estimated by the minimum t-statistic on the coefficient of the autoregressive variable (t_α). The asymptotic critical values for t statistics are provided by Zivot and Andrews (1992). If the computed t statistics in absolute value are higher than Zivot and Andrews (1992) critical values, one can reject the null hypothesis and conclude that the series is a trend stationary process with a structural break.

Cointegration

The ARDL bounds testing approach is employed to test for cointegration relationship among the variables when the order of integration is I (0) or I (1). An ARDL model is a general dynamic specification that is formulated with the lags in the dependent variable and the lagged and contemporaneous values of the independent variables. In doing so, short-run dynamic effects can be directly estimated and the long-run equilibrium relationship can also be estimated indirectly. For the present study, the ARDL method is adopted in investigating the presence of a long-run equilibrium relationship employing the following unrestricted error correction model (UECM):

$$\begin{aligned}\Delta \ln GDP_t = \alpha_0 &+ \sum_{i=1}^{n} \alpha_{1i} \Delta \ln GDP_{t-i} + \sum_{i=1}^{n} \alpha_{2i} \Delta \ln EC_{t-i} \\ &+ \sum_{i=1}^{n} \alpha_{3i} \Delta \ln EM_{t-i} + \alpha_4 \ln GDP_{t-1} + \alpha_5 \ln EC_{t-1} \\ &+ \alpha_6 \ln EM_{t-1} + \mu_{1t}\end{aligned} \tag{3}$$

$$\begin{aligned}\Delta \ln EC_t = \beta_0 &+ \sum_{i=1}^{n} \beta_{1i} \Delta \ln EC_{t-i} + \sum_{i=1}^{n} \beta_{2i} \Delta \ln GDP_{t-i} \\ &+ \sum_{i=1}^{n} \beta_{3i} \Delta \ln EM_{t-i} + \beta_4 \ln EC_{t-1} + \beta_5 \ln GDP_{t-1} \\ &+ \beta_6 \ln EM_{t-1} + \mu_{2t}\end{aligned} \tag{4}$$

$$\begin{aligned}\Delta \ln EM_t = \gamma_0 &+ \sum_{i=1}^{n} \gamma_{1i} \Delta \ln EM_{t-i} + \sum_{i=1}^{n} \gamma_{2i} \Delta \ln GDP_{t-i} \\ &+ \sum_{i=1}^{n} \gamma_{3i} \Delta \ln EC_{t-i} + \gamma_4 \ln EM_{t-1} + \gamma_5 \ln GDP_{t-1} \\ &+ \gamma_6 \ln EC_{t-1} + \mu_{3t}\end{aligned} \tag{5}$$

Where Δ represents the first difference operator, μ represents the white noise error term, lnGDP denotes the nature logarithm of the real GDP, and similarly for lnEC and lnEM. The parameters $\alpha_{1,\ ...,3}$, $\beta_{1,\ ...,3}$, and $\gamma_{1,\ ...,3}$ are the short-run dynamic coefficients, while $\alpha_{4,\ ...,6}$, $\beta_{4,\ ...,6}$, and $\gamma_{4,\ ...,6}$ are the corresponding long-run multipliers of the underlying ARDL model.

To determine whether there is a cointegration relationship among the variables, we test for the joint significance of the lagged levels of the variables with the aid of the F test. The null hypotheses of no cointegration among the variables in Eqs. (3), (4), and (5) are H_0: $\alpha_4 = \alpha_5 = \alpha_6 = 0$, $\beta_4 = \beta_5 = \beta_6 = 0$, and $\gamma_4 = \gamma_5 = \gamma_6 = 0$ against the alternative hypotheses H_1: $\alpha_4 \neq \alpha_5 \neq \alpha_6 \neq 0$, $\beta_4 \neq \beta_5 \neq \beta_6 \neq 0$, and $\gamma_4 \neq \gamma_5 \neq \gamma_6 \neq 0$, respectively. Under the null hypothesis, the computed F statistics are represented as $F_{\ln GDP}(\ln GDP|\ln EC, \ln EM)$, $F_{\ln EC}(\ln EC|\ln GDP, \ln EM)$, and $F_{\ln EM}(\ln EM|\ln GDP, \ln EC)$, respectively.

Pesaran et al. (2001) showed that, under the null hypothesis, the F test has a non-standard distribution. They proposed a set of asymptotic critical F values in each significance level for a large sample size. While Narayan (2005) developed a set of asymptotic critical F values for a small sample size ranging from 30 to 80 observations. It should be emphasized that critical values based on the large sample size deviate significantly from those for the small sample size. Given that the sample size of our empirical study is relatively small and only about 40 observations, we use the critical F values from Narayan (2005) instead of references from Pesaran et al. (2001). A judgment on whether there exists a co-integration relationship between the dependent variable and its regressors is then to be made according to the computed F statistics. If the computed F statistics falls below the lower bounds value, then one fails to reject null hypothesis of no cointegration. If the computed F statistics exceed the upper critical-bound value, then the null hypothesis can be rejected. It could be concluded that there is a co-integrated relationship between the dependent variable and the regressors. However, if the computed F statistics falls between the bounds, then the test for cointegration becomes inconclusive.

If evidence for a long-run equilibrium relationship can be found, the associated ARDL and error correction mechanisms (ECMs) are employed to estimate the long-run and short-run coefficients and examine the existence of Granger causality among the variables.

Long-run and short-run Granger causality

This stage uses augmented constructing standard Granger-type causality tests with a lagged error correction term. When a cointegration relationship exists, a multivariate pth-order vector error correction model (VECM) can be used to investigate Granger causality, as follows.

$$\begin{aligned}\Delta \ln GDP_t = \alpha_0 + \sum_{i=1}^{p} \alpha_{1i} \Delta \ln GDP_{t-i} + \sum_{i=1}^{p} \alpha_{2i} \Delta \ln EC_{t-i} \\ + \sum_{i=1}^{p} \alpha_{3i} \Delta \ln EM_{t-i} + \alpha_{4i} ECT_{t-1} + \mu_{1t}\end{aligned} \tag{6}$$

$$\begin{aligned}\Delta \ln EC_t = \beta_0 + \sum_{i=1}^{p} \beta_{1i} \Delta \ln EC_{t-i} + \sum_{i=1}^{p} \beta_{2i} \Delta \ln GDP_{t-i} \\ + \sum_{i=1}^{p} \beta_{3i} \Delta \ln EM_{t-i} + \beta_{4i} ECT_{t-1} + \mu_{2t}\end{aligned} \tag{7}$$

$$\begin{aligned}\Delta \ln EM_t = \gamma_0 + \sum_{i=1}^{p} \gamma_{1i} \Delta \ln EM_{t-i} + \sum_{i=1}^{p} \gamma_{2i} \Delta \ln GDP_{t-i} \\ + \sum_{i=1}^{p} \gamma_{3i} \Delta \ln EC_{t-i} + \gamma_{4i} ECT_{t-1} + \mu_{3t}\end{aligned} \tag{8}$$

Where, α, β, and γ are parameters to be estimated; the lag length p is based on Schwarz–Bayesian (SBC) and/or Akaike information criteria (AIC); μ_t is the serially-uncorrelated error terms; and ECT_{t-1} is the lagged error-correction term (ECT) obtained from the long-run equilibrium relationship. Theoretically, the coefficient of the ECT should be negative and less than one in absolute value, which ensures the ECT maintains the equilibrium relationship between the cointegrated variables over time. To test for long-run and short-run Granger causality, one can check the F statistics on the lagged explanatory variables of the ECM, which shows the significance of short-run causal effects. Similarly, the t statistics on the coefficients of the lagged error-correction term shows the significance of the long-run causal effect.

Empirical results

Unit root test

As for the time series properties of the underlying variables, we conduct the different unit root tests, namely those listed in "Stationary", to obtain robust results. It should be noted that the unit root tests vary from different lag structures. Accordingly, the Schwarz information criterion could be applied to the lag selection in unit root test. The ADF and PP tests (but not the KPSS test) have a null hypothesis that the investigated series has a unit root against the alternative hypothesis of stationarity. The null hypothesis of KPSS test, in contrast, states that the variable is stationary.

Table 3 presents the results of the ADF, PP, and KPSS tests on the integration properties of the real GDP (ln*GDP*), electricity consumption (ln*EC*), and employment (ln*EM*) concerning China. As evidence from the results, it turns out that all of the variables are non-stationary in their level data from the three tests (Panel A). After taking the first difference of the variables, nevertheless, the stationary property can be shown at 10%, 5%, or the stricter 1% critical levels (Panel B). Though one conflicting result on the stationary property in the first difference is found for the series ln*EM* in the intercept and trend model of the KPSS test, in general, the results from conventional unit root tests support the argument that all the variables are integrated at order one (i.e., $I(1)$).

The results of the ZA Model A and Model C unit root tests for ln*GDP*, ln*EC*, and ln*EM* are reported in Table 4. No additional evidence exists against the unit root hypothesis in the

Table 3 Augmented Dickey-Fuller, Phillips-Perron, and Kwiatkowski-Phillips-Schmidt-Shin tests for unit root

	Variables	ADF test	PP test	KPSS test
Panel A: level				
Intercept	ln*GDP*	1.1378 (2)	2.4284	0.7600[a]
	ln*EC*	1.1902 (1)	1.2064	0.7645[a]
	ln*EM*	− 1.6778 (0)	− 1.6207	0.7259[b]
Intercept and trend	ln*GDP*	− 2.9646 (1)	− 3.0786	0.1516[b]
	ln*EC*	− 1.4334 (1)	− 1.2605	0.1446[c]
	ln*EM*	− 0.0922 (0)	− 0.1714	0.1660[b]
Panel B: first difference				
Intercept	ln*GDP*	− 3.0011 (1)[b]	− 4.1215[a]	0.3116
	ln*EC*	− 3.3370 (4)[b]	− 4.6094[a]	0.2050
	ln*EM*	− 5.3812 (0)[a]	− 5.4443[a]	0.3297
Intercept and trend	ln*GDP*	− 3.3223 (1)[c]	− 4.2882[a]	0.1027
	ln*EC*	− 3.4086 (4)[c]	− 4.7697[a]	0.0630
	ln*EM*	− 5.8399 (0)[a]	− 5.8389[a]	0.1261[c]

Notes:

1. The null hypothesis for the augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests indicates that a unit root exists. The optimal lag lengths for the ADF test on the variables are adopted with the Schwarz information criterion. The maximum number of lags is set at five and the computed lag lengths are shown in parentheses. The bandwidth for the PP test was adopted with the Newey-West Bartlett kernel.

2. The null hypothesis for the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test is stationary. The spectral estimation utilizes the Barlett-Kernel method. The Newey-West method is used for the bandwidth.

3. Superscripts a, b, and c denote significance at 1%, 5%, and 10% levels, respectively.

Table 4 Zivot and Andrews test for unit root with a structural break

	ln*GDP*		ln*EC*		ln*EM*	
	Model A	Model C	Model A	Model C	Model A	Model C
Panel A: level						
Test value	− 3.231(1)	− 3.630(1)	− 3.687(2)	− 4.013(2)	− 1.474(0)	− 3.086(0)
Break	1990	1982	2002	1997	1984	1996
Panel B: first difference						
Test value	− 4.781(0)[c]	− 4.914(0)[c]	− 5.385(0)[a]	− 6.129(0)[a]	− 7.299(0)[a]	− 8.840(0)[a]
Break	1978	1988	2001	2002	1992	1992

Note: The critical values for the 1%, 5%, and 10% levels are − 5.34, − 4.80, and − 4.58 for Model A. The critical values for the 1%, 5%, and 10% levels are − 5.57, − 5.08, and − 4.82 for Model C. Superscripts a, b, and c denote significance at 1%, 5%, and 10% levels, respectively. The numbers in parentheses are the lag order

level data. However, a stationary trend with structural break is found in the first difference of the variables at 10% or the stricter 1% critical levels, confirming that all the three series are at $I(1)$.

Bounds cointegration test

Before implementing bound-testing procedures, optimal lag orders on the first difference variables in Eqs. (3)–(5) should be determined from the unrestricted models by following the minimum values of the SBC. The results show that the optimal lags in Eqs. (3)–(5) are all four. With these optimal lag lengths, a bounds F test is applied to examine the long-run equilibrium relationship in Eqs. (3)–(5).

The results of the cointegration bounds test are listed in Table 5, together with the relevant critical value bounds. Concerning Eq. (3), the computed $F_{\ln GDP}(\ln GDP|\ln EC, \ln EM)$ is 6.098, and higher than the upper bound critical value at the 1% significance level. This suggests that there exists a long-run relationship between ln*GDP*, ln*EC*, and ln*EM*. The variables ln*EC* and ln*EM* can be treated as the long-run forcing variables for the explanation of ln*GDP*. However, the bounds test indicates that when ln*EC* and ln*EM* are the dependent variables $F_{\ln EC}(\ln EC|\ln GDP, \ln EM)$ and $F_{\ln EM}(\ln EM|\ln GDP, \ln EC)$ are lower than the lower bound critical value at the 10% significance level. It shows that there is no cointegration relationship when the variables ln*EC* and ln*EM* are used as dependent variables. From the above test results, we can conclude that there is one single long-run cointegration relationship among the variables under investigation.

Referring to Table 6, the long-run parameters related to real GDP in China present to be positive as expected. Concerning electricity consumption, its long-run impact in real GDP is around 1.00 at statistically significant level of 1%, implying that a 1% growth in electricity consumption will lead to a 1% increase in GDP. Similarly, it is found that employment has a positive impact on real GDP in the long run. The elasticity is

Table 5 Bounds F test for cointegration

Dependent variable		Function		F test statistic	
ln*GDP*		$F_{\ln GDP}(\ln GDP\|\ln EC, \ln EM)$		6.098[a]	
ln*EC*		$F_{\ln EC}(\ln EC\|\ln GDP, \ln EM)$		1.442	
ln*EM*		$F_{\ln EM}(\ln EM\|\ln GDP, \ln EC)$		0.695	
Asymptotic critical value					
10%		5%		1%	
$I(0)$	$I(1)$	$I(0)$	$I(1)$	$I(0)$	$I(1)$
2.835	3.585	3.435	4.260	4.770	5.855

Note: Superscript a denotes significance at 1% level. Asymptotic critical values of the lower $I(0)$ and upper $I(1)$ bounds are taken from Narayan (2005, Appendix: Case II). As real GDP is the dependent variable, a long-run equilibrium exists in Eq. (3). The procedure in the next step uses the associated ARDL and ECM to estimate long-run and short-run coefficients. Finally, ARDL (4,4,1) is adopted. Along with several diagnostic tests for the underlying ARDL model, the derived long-run elasticities are shown in Table 6

Table 6 Estimated long-run coefficients based on autoregressive distributed lag (4,4,1)

Dependent variable: ln*GDP*		
Panel A: estimated long-run coefficients		
Regressor	Coefficient [standard error]	T ratio [p value]
ln*EC*	0.9972[0.0343]	29.0975[0.000]
ln*Em*	0.3733[0.1136]	3.2873[0.003]
Intercept	− 5.4320[1.4558]	− 3.7313[0.001]
Panel B: diagnostic tests		
Test statistics	Statistic value	p value
Serial correlation	0.1949	0.659
Functional form	0.2271	0.634
Normality	0.2446	0.885
Heteroscedasticity	0.6541	0.419

Table 7 Estimated short-term coefficients based on autoregressive distributed lag (4,4,1)

Dependent variable: ΔlnGDP

Regressor	Coefficient	Standard error	T-ratio [*p* value]
Δln*GDP*(− 1)	0.6554	0.1242	5.2757[0.000]
Δln*GDP*(− 2)	− 0.2064	0.1494	− 1.3812[0.179]
Δln*GDP*(− 3)	0.3996	0.1207	3.3110[0.003]
Δln*EC*	0.7703	0.0909	8.4710[0.000]
Δln*EC* (− 1)	− 0.5960	0.0997	− 5.9774[0.000]
Δln*EC* (− 2)	0.0788	0.1342	0.5871[0.562]
Δln*EC* (− 3)	− 0.5120	0.1114	− 4.5945[0.000]
Δln*Em*	− 0.1890	0.1070	− 1.7680[0.089]
ecm(− 1)	− 0.3206	0.0493	− 6.4981[0.000]

ecm = ln*GDP*-0.9972*ln*EC*-0.3733*ln*EM*+5.4320*Intercept

around 0.37 with statistically significant level of 1%, implying that a 1% growth in employment will cause a 0.37% increase in GDP. Finally, a battery of diagnostic tests for the underlying ARDL model that includes testing for serial correlation, misspecification of functional form, normality of the residuals, and heteroscedasticity, do not find any significant evidence for a departure from the standard assumptions.

The short-run equilibrium relationship is derived from the error correction model, and the results are displayed in Table 7. The parameters of short-run elasticities are lower in absolute value than those in the long-run. With the exception of the coefficients of Δln*GDP*(−2) and Δln*EC*(−2), all the other coefficients are statistically significant. As expected, it is found that the lagged error correction term (denoted as ecm(−1) in Table 6) is negative with statistical significance.

This reveals the dynamic of the endogenous variable adapt for changes of the explanatory variables before converging to its equilibrium level. In addition, the results imply that the adjustment to restore equilibrium is greatly effective. In the present study, the relatively high coefficient in absolute magnitude of the error correction term shows a more rapidly adjustment dynamic. According to the coefficient of − 0.32 at a statistically significant level of 1%, it suggests that the process of converging to equilibrium need over 3 years after a shock to GDP in China.

In addition, the stability of the short-run and long-run coefficients is checked through the cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) tests to confirm

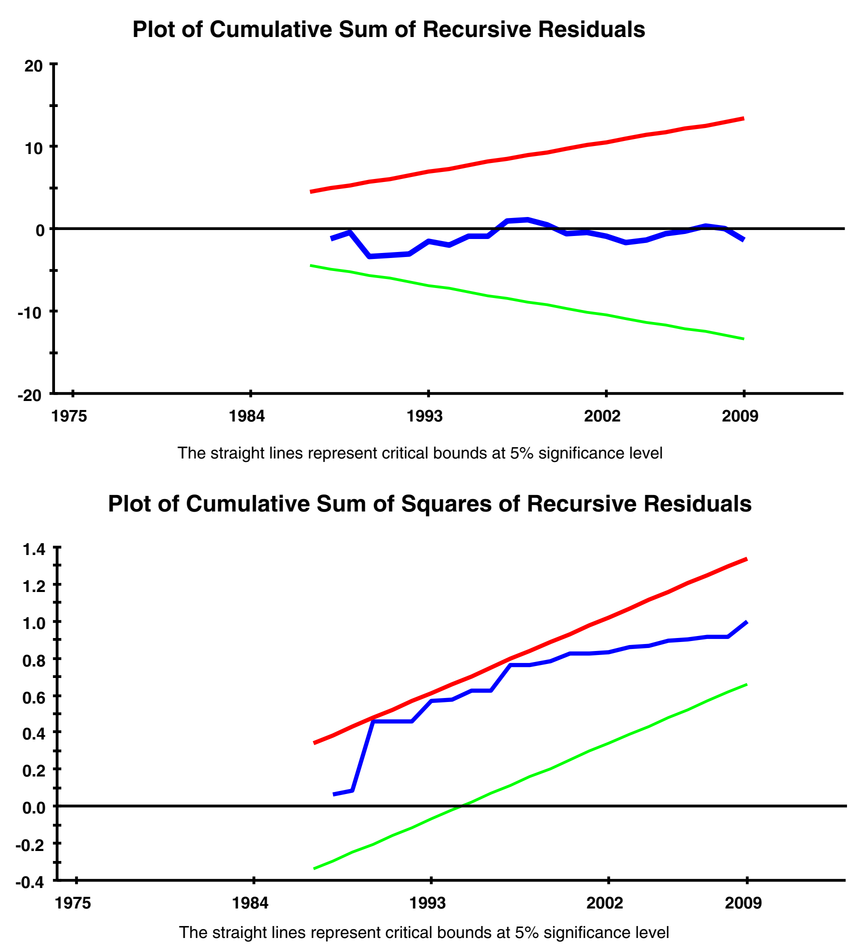

Fig. 1 Plot of CUSUM and CUSUMQ tests for the parameter stability

Table 8 Results of Granger causality

Dependent variable	F statistics [p value]			t statistic [p value]
	$\sum\Delta\ln GDP_{t-1}$	$\sum\Delta\ln EC_{t-1}$	$\sum\Delta\ln EM_{t-1}$	ECT_{t-1}
$\Delta\ln GDP_t$	–	12.7340 [0.000]	5.0484 [0.034]	−6.4981 [0.000]
$\Delta\ln EC_t$	6.4525 [0.001]	–	0.0217 [0.844]	–
$\Delta\ln EM_t$	0.9876 [0.432]	0.2709 [0.894]	–	–

the goodness of model fit. Figure 1 shows the plot of CUSUM and CUSUMSQ test statistics that both fall within the critical bounds of the 5% significance level. This implies that the estimated parameters are all stable over time and policy conclusion can be inferred from the model.

Granger causality test

The cointegration among electricity consumption, real GDP, and labor force suggests that there could exist Granger causality at least in one direction. But it could not reveal the concrete direction of causality among those variables. Therefore, we use the error correction mechanism (ECM) to simultaneously test short-run and long-run Granger causality. The results listed in Table 8 show the short-run and long-run causal effects are both significant according to the F statistics on the lagged explanatory variables and the t statistic on the coefficient of the lagged error-correction term. For the short-run causality, the results show that employment and electricity consumption are both significant at 5% or 1% level in the real GDP equation. Besides, the electricity consumption equation which presents real GDP is also significant at the 1% level. Nevertheless, the employment equation presents no significance in real GDP or electricity consumption. It suggests that the bidirectional Granger causality of real GDP and electricity consumption is weak in the short run; and there existed only unidirectional Granger causality from employment towards real GDP.

With regard to the t statistic on the coefficient, the lagged error-correction term is found to be statistically significant in the real GDP equation at 1% level, which confirms the result of the bounds test to cointegration. Since the variables are not cointegrated in Eqs. (4) and (5), a lagged error correction cannot be included when either electricity consumption or employment is used as the dependent variable. The Granger causality test presents that there is no equilibrium in the cointegration relationship, but long-run causality running interactively from electricity consumption and employment to real GDP.

The short-run and long-run Granger causality tests appear to identify a significant impact of electricity consumption on economic growth. Compared to previous works on the direction of causal relationship between electricity consumption and economic growth in China, the evidence of Granger causality from electricity consumption to real GDP in this study appears to be consistent with the research results by Shiu and Lam (2004) and by Yuan et al. (2007). However, the evidence differs from empirical results of Chen et al. (2007), in which no cointegration is found between electricity consumption and economic growth for China while estimated with a single country data set. Considering the evidence of a short-run causal relationship running from real GDP to electricity consumption, our empirical result also differs from the findings by Shiu and Lam (2004) and Yuan et al. (2007). Although there appear differences in the empirical results for the same country, the results in our study appear to be more robust due to the two developments in our model and method used, which have been mentioned in "Literature review".

Policy implications

The long-run causality suggests that to some extent, China is an energy-dependent country. The implication is that any extreme conservation policy or shock to electricity supplies will have a significantly adverse effect on economic growth. Hence, the measures focusing only on electricity consumption reduction could not be easily implemented in China. Efficient implementation of policies and measures should be considered to balance the electricity consumption reduction and economic growth.

First, electricity consumption has played an important role in China's economic development. Therefore, policies are required to ensure that the electricity is sufficient to keep regular economic growth. Guaranteeing supply and enhancing security are prerequisite if the functioning of economic activities is to be maintained. China's increasing demand for electricity needs adequate generating preparation and speeding up nationwide inter-connectivity by upgrading power distribution networks, thereby undertaking increasing electrical supplies.

Second, improving energy efficiency can be a viable policy. China's current per unit GDP energy consumption is obviously higher compared to international levels (Wang et al. 2017). The energy consumption intensity is 3.8 times of global averages, 4.3 times of that in the USA, and 11.5 times of that in Japan, all of which indicate that there is much space for improvements in energy efficiency. Hence, China could make further efforts through the implementation of the Clean

Development Mechanism, focusing on reinforcing technology innovation in the electrical power sector, accelerating popularization of clean production, recycling material, and cascading utilization of heat technologies to improve energy efficiency in the manufacturing sector. In addition, accelerating electricity structure amelioration and optimization can be another positive policy as well. The emphasis placed on cultivating vigorously new energy industry that would not have significant adverse influence in the economy in the long term (Weidou and Johansson 2004). For this reason, China could get over traditional dependence on fossil fuel in the long run by diversifying electricity supplies with a preference for cleaner, renewable, and cost-effective energy such as hydropower, nuclear power, and solar and wind force, to relieve environmental pressure. It is worth noting that China has regarded developing new energy and renewable energy as an effective strategy for national economic and social development and has proposed for the first time the target of improving the non-fossil energy ratio in energy depletion to 15%.

Moreover, to succeed in the sustainability of the economy in the long term, China should adopt an alternative economic growth model and develop new strategic industries to readjust its economic structure. In 2009, secondary and tertiary industries accounted for around 46.3% and 43.4% of the total economy in China while these ratios in the main developed countries accounted for less than 30% and over 70%, respectively. The nature of its economic structure is one of the key reasons why China's energy consumption intensity is relatively high. Thus, it is important to increase the share of the tertiary sector in aggregate output. In secondary industries, China also needs to improve industrial structure adjustment and technological advance, in an attempt to promote manufacturing industry transformation from the industries highly dependence on energy (such as those involved in ferrous metals, nonferrous metals, textiles, chemicals, and nonmetal mineral products). Thus, if heavy industrialization mode can be transited to a high-efficiency and cleaner production development pattern, energy can be saved with lower gas emission in the course of economic growth.

Conclusions

This article has reassessed the dynamic nexus of economic growth, electricity consumption, and labor force in China during the period of 1971–2009 by the ARDL approach proposed by Pesaran et al. (2001). The contribution of this research to the field of energy-GDP nexus is reflected in two important aspects. First, following recent trends, our empirical study tests the relationship within a multivariate framework instead of using a bivariate model. Second, this study employs a relatively new time-series approach (ARDL) capable of uncovering relationships that might otherwise be missed using conventional estimation techniques.

The results of the bounds testing procedure confirm the presence of cointegration between electricity consumption, labor, and economic growth in China. Over the long run, electricity and labor are key determinants for economic growth. Our empirical study finds a robust result that electricity consumption has a significantly positive impact on economic growth. The magnitude of the impact is 1.00, implying that a 1% increase in electricity consumption leads to 1% increase in GDP. Our empirical results also provide insights on the short-run speed of the adjustment process to restore long-term equilibrium in the real GDP equation. With a coefficient of −0.32, the significance is that it could take a little over 3 years to restore to an equilibrium after China's GDP unrest.

The findings presented in this paper imply that electricity serves as an important contributor to economic growth in China and that binding electricity consumption constraints may prevent the economy from moving forward. To tackle this dilemma between energy consumption and economic development in China, the government should concentrate on such aspects as energy efficiency enhancement, electricity structure amelioration, and economic structure optimization, with the premise of enhancing electricity security and guaranteeing electricity supplies.

Acknowledgments This work was funded by the Fundamental Research Funds for Fujian Province Public-interest Scientific Institution, Grant 2017R1101003-7, and the National Natural Science Foundation of China (NSFC), Grant 40701063. We also sincerely thank Prof. Yusheng Yang of Fujian Normal University for your kind support in this work.

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