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Validating the environmental Kuznets curve hypothesis in India and China: The role of hydroelectricity consumption

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ABSTRACT

The aim of this research is to examine the link between CO_2 emissions, hydroelectricity consumption, urbanisation and real GDP in China and India during the period of 1965–2013. The long-run cointegration is investigated by the autoregressive distributed lag (ARDL) bounds testing approach, which is augmented with structural breaks. We employ the ARDL cointegration test to establish long run relationship in the variables. Furthermore, we use the ARDL to show that real GDP and urbanisation have long-run positive impact on emission, while hydroelectricity consumption exerts long-run negative impact on emission in both countries. The results support the existence of environmental Kuznets curve (EKC) hypothesis in China and India. Besides, the paper assesses the causal link between the variables by using Granger causality procedures and the results show that there is long-run bidirectional relationship between the variables in both countries.

1. Introduction

It is well known that China and India achieved a significant economic development in the last three decades. Both countries' gross domestic product (GDP) collectively account for 15% of the global GDP [59]. Since the 1990s, the GDP growth rates of the two countries are perceived to be among the fastest in the world. The increase in the living standard in these economies encouraged most of their population to move towards urban areas (which are more energy intensive than rural areas) for better quality of living and higher job opportunities and this caused urban population to grow to high levels. For instance, urban population represents 16% of total population in 1960 in China. However, in 2015 urban population represents 56% of total China's population. Likewise, India's urban population increased significantly from 18% of total India's population in 1960 to 33% in 2015 [60]. Associated with this economic growth, the combined share of the two countries' energy consumption in the global energy consumption increased from 10% in 1990 to 25% in 2013. This figure is projected to increase further to 32% in 2035 (Energy Information Administration, 2015). Moreover, in 2014 and 2016 most of China's and India's consumption of energy comes from the industrial and construction sectors, which represent 70% of total energy consumed ([15]; Energy Statistics, 2016). Fig. 1 shows clearly that the gap between the amount of energy consumed in both countries and the world is getting smaller during the last two decades and it is expected to get closer in the future.

The increase in energy consumption in the two countries significantly increased their air pollution levels. Hence, India and China's contribution to the world's CO_2 emission increased from 15% of the world's CO_2 emission to 33% in 2013. This large production of CO_2 emission is due to the rise in fossil fuels energy consumption which constitute about 77% and 81% of the energy mix in China and India, respectively in 2013 [20]. This increase in the air pollution levels that both countries are witnessing increased efforts by the two governments to implement projects that promote renewable energy. For China's case the country has implemented different policies to promote renewable energy such as the Policy Processes of 2006 and 2009, subsidies schemes of 2003 and 2010, which were all designed to encourage renewable energy use in the country [51]. In addition, India also made serious steps to promote the role of their renewable sources of energy including the incorporation of Indian Renewable Energy Development Agency Limited (IREDA) in 1992. IREDA was created as a reliable institute for promoting and funding self-sustaining projects in energy generation from renewable sources, environment technologies and energy efficiency for sustainable and inclusive development. Moreover, among the recently introduced blueprints is the National Clean Energy Fund (NCEF) which is aimed to provide fund for renewable energy projects including hydropower in the country Ministry of New and Renewable Energy India [32]. These policies played important roles in the increase of renewable energy share of

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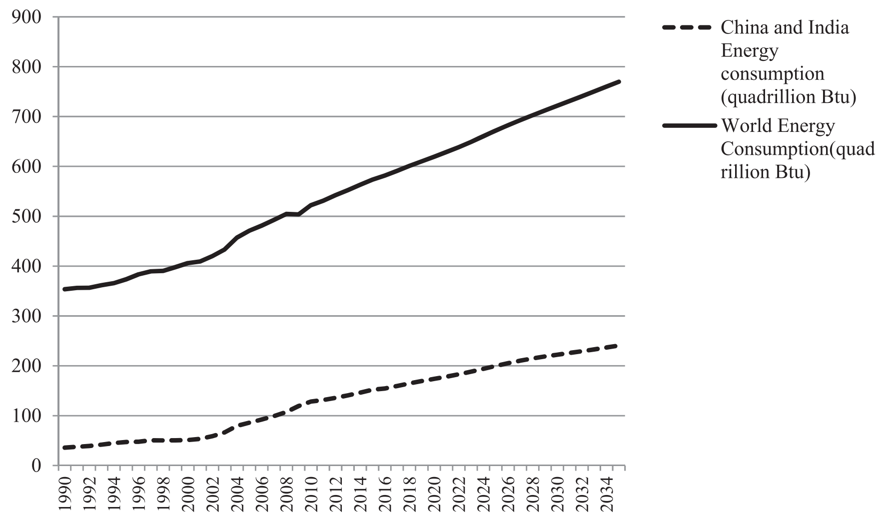

Fig. 1. Energy consumption between India and China and the World (1990–2035).

total electricity generation by over 45% [20]. Hydroelectricity is the dominant source of renewable energy in China and India. Basically in 2014 other renewable sources represent only 3.5% in China and 4.9% in India while hydroelectricity in China and India represents 16.7% and 11.8% respectively of the total energy mix [60]. Hydroelectricity use in particular and electricity use in general are increasing due to the growing economic activities in the two countries. Moreover, these countries are in the top 10 hydroelectricity producers in the world. This source of renewable energy can aid in reducing the levels of air pollution in both countries. During the period of 1980–2012, hydroelectricity role increased in these countries as its levels of production almost doubled [59]. Therefore, the rise in the levels of production and consumption of this source of energy might have a noteworthy effect in plummeting air pollution. However, despite the well-established literature (see Table 1), the effect of hydroelectricity on air pollution in China and India has gained little attention by the scholars.

The previous studies utilized different determinants of pollution, such as GDP and energy consumption ([12,13,21,34,37,56,5,3,50,6,53]; and so forth.), urban population or population density (Al-mulali et al., 2015; [26,36,11,6,64,61]), trade openness ([21,28,48,49,3,4,26]), and financial development ([4,35,48,49,62]. Moreover, since CO_2 emission accounts for over 50% of greenhouse gas emission, most of the scholars used CO_2 emission as an indicator of air pollution ([12,13,21,24,27,28,37–39,50,56,5,7]; Baek, 2015; [26,54,11,61]; and so forth).

The environmental Kuznets curve (EKC) hypothesis became popular among scholars as it represents an important tool for environmental policy. The hypothesis states that in the initial stages of economic expansion of a country, the increase in GDP growth will cause more environmental pollution until it reaches a certain point where the relationship between GDP growth and environmental pollution is negative. There are several aspects of EKC hypothesis that have not been adequately addressed by the existing literature in China and India. The existing papers have largely used energy consumption (and in some cases coal consumption) as proxy for the energy use. The role of hydroelectricity consumption in the EKC has not been adequately provided for, in the EKC hypothesis. Considering aggregate renewable energy without minding the diverse nature of its components can hide the differing impacts of different kinds of energy consumption and might cause wrong policy inferences for each component, especially for hydroelectricity which is characteristically different from other kinds of energy and has different impacts on emission in the case of China and India. In addition, the issues of structural breaks have not been adequately incorporated in the model on China and India. Ignoring the possibility of structural breaks in the analysis may affect the ability of rejecting the null hypothesis of nonstationarity [41]. We aim to contribute to the current literature by examining the consequences of hydroelectricity consumption on air pollution in China and India. We also contribute to the literature by allowing for structural breaks in the estimation process including the causality test.

2. Methodology and data

The conventional EKC hypothesis implies that the environmental degradation is dependent on GDP and square of GDP. Energy use and urbanisation have been used as additional factors of emissions [22,24,5,55,64]. Consistent with the foregoing papers, the following model is considered:

$$\ln C_t = \alpha_1 + \alpha_2 \ln Y_t + \alpha_3 \ln Y_t^2 + \alpha_4 \ln H_t + \alpha_5 \ln U_t + \alpha_6 T + \alpha_7 D_1 + \alpha_8 D_2 + \upsilon_t \tag{1}$$

C_t is per capita CO_2 emission (tonnes of oil equivalent of carbon dioxide emissions by population), Y_t is real gross domestic product per capita (constant 2005 US$), Y_t^2 is the square of real gross domestic product per capita (constant 2005 US$), H_t is per capita hydroelectric use (tonnes of oil equivalent of hydroelectricity consumed as a fraction of the total population), and U_t is urbanisation population ratio (urban population as a fraction of the total population). T is the time trend and D_1 is the first dummy variables, which captures the first structural break. D_2 is the second dummy variable, which captures the second structural break. The structural shifts are selected based on the structural breaks in the unit root analysis of emission at level. The first sets of structural breaks are mostly in the early 1980s and the second sets are mostly in the late 1990s and therefore we pick the most recurring in each set. Our dataset is for the period of 1965–2013. The data for CO_2 emission and hydroelectricity consumption were retrieved from BP Statistical Review of World Energy, 2014, while their divisor-population is retrieved from World Development Indicators of the World Bank. The data for GDP per capita and the share of urban population in the total population were acquired from World Development Indicators of the World Bank.

Using the autoregressive distributed lag (ARDL) method, we formulate the following error correction model:

$$\Delta \ln C_t = \alpha_1 + \sum_{i=1}^{k} \alpha_2 \Delta \ln C_{t-i} + \sum_{i=1}^{k} \alpha_3 \Delta \ln Y_{t-i} + \sum_{i=1}^{k} \alpha_4 \Delta \ln Y_{t-i}^2 + \sum_{i=1}^{k} \alpha_5 \Delta \ln H_{t-i} + \sum_{i=1}^{k} \alpha_6 \Delta \ln U_{t-i} + \alpha_7 C_{t-1} + \alpha_8 Y_{t-1} + \alpha_9 Y_{t-1}^2 + \alpha_{10} H_{t-1} + \alpha_{11} U_{t-1} + \alpha_{12} T + \alpha_{13} D_1 + \alpha_{14} D_2 + \upsilon_t \quad (2)$$

Δ is the difference operator. We have previously defined all the variables.[1] The use of GDP (and the square of GDP) can be supported with specific studies that have used GDP (and the square of GDP) as determinant(s) of emission including Apergis and Ozturk [6], Begum et al. [11] and Ozturk and Al-mulali [4]. The adoption of hydroelectricity consumption in the model is consistent with the studies of Saboori and Sulaiman [46] Kivyiro and Arminen [27] and Tang and Tan [54]. The use of urbanisation can be supported with studies that have used urbanisation as determinant of emission include Begum et al. [11], Apergis and Ozturk, [6], Zhang et al. [64] and Yin et al. [61]. According to Pesaran et al. [42], the cointegration test can be conducted by examining the F-test on the lagged levels of the variables. Therefore, the null hypothesis of no-cointegration $\alpha_7 = \alpha_8 = \alpha_9 = \alpha_{10} = \alpha_{11} = 0$ can be tested against the alternative hypothesis of $\alpha_7 \neq \alpha_8 \neq \alpha_9 \neq \alpha_{10} \neq \alpha_{11} \neq 0$.[2] After analysing the long-run correlation between the series and finding the long-run coefficients, we investigate the short-run coefficients as follows:

$$\Delta \ln C_t = \alpha_1 + \sum_{i=1}^{k} \alpha_2 \Delta \ln C_{t-i} + \sum_{i=0}^{k} \alpha_3 \Delta \ln Y_{t-i} + \sum_{i=0}^{k} \alpha_4 \Delta \ln Y_{t-i}^2 + \sum_{i=0}^{k} \alpha_5 \Delta \ln H_{t-i} + \sum_{i=0}^{k} \alpha_6 \Delta \ln U_{t-i} + \alpha_7 \Delta T + \alpha_8 \Delta D_1 + \alpha_9 \Delta D_2 + \alpha_{10} ECT_{t-1} + \upsilon_t \quad (3)$$

α_9 is the speed of adjustment parameter and error correction term (ECT) in Eq. (3) is the residuals generated from the estimated cointegration regression of Eq. (1). In order for ECT to be reliable, it must produce statistically significant negative coefficients. We additionally test the causal link among the variables by using the following regression equations:

$$\Delta \ln C_t = \alpha_{11} + \sum_{i=1}^{k} \alpha_{12} \Delta \ln C_{t-i} + \sum_{i=1}^{k} \alpha_{13} \Delta \ln Y_{t-i} + \sum_{i=1}^{k} \alpha_{14} \Delta \ln Y_{t-i}^2 + \sum_{i=1}^{k} \alpha_{15} \Delta \ln H_{t-i} + \sum_{i=1}^{k} \alpha_{16} \Delta \ln U_{t-i} + \alpha_{17} T + \alpha_{18} \Delta D_1 + \alpha_{19} \Delta D_2 + \alpha_{110} ECT_{t-1} + \upsilon_t \quad (4)$$

$$\Delta \ln Y_t = \alpha_{21} + \sum_{i=1}^{k} \alpha_{22} \Delta \ln C_{t-i} + \sum_{i=1}^{k} \alpha_{23} \Delta \ln Y_{t-i} + \sum_{i=1}^{k} \alpha_{24} \Delta \ln Y_{t-i}^2 + \sum_{i=1}^{k} \alpha_{25} \Delta \ln H_{t-i} + \sum_{i=1}^{k} \alpha_{26} \Delta \ln U_{t-i} + \alpha_{27} T + \alpha_{28} \Delta D_1 + \alpha_{29} \Delta D_2 + \alpha_{210} ECT_{t-1} + \upsilon_t \quad (5)$$

[1] Check the explanations of the variables in the definitions used for Eq. (1).

[2] We also test for cointegration with the GDP, hydroelectricity and urbanisation expressed as dependent variables. However, we do not express the equations here in order to conserve space.

Table 1
An overview of the environmental Kuznets curve literature.

Author	Period	Country/ region/ organization	Methodology	Variables used in the study	EKC Hypothesis
Ang [6]	1960–2000	France	Johansen cointegration and VECM Granger causality.	CO_2 emission, GDP, GDP square and energy consumption.	Yes
Akbostancı et al. [2]	1992–2001	Turkey	Johansen cointegration and pooled generalized least square.	CO_2 emission, SO_2 emission, PM_{10}, GDP, GDP square, GDP cubic, population.	No
Acaravci and Ozturk [1]	1960–2005	European countries.	ARDL modelling approach and VECM Granger causality.	CO_2 emission, GDP, GDP square and energy consumption.	Yes
Pao and Tsai [37]	1971–2005	Brazil, Russia, India and China (BRIC).	Pedroni, Kao and Johansen cointegration, ordinary least square (OLS) and VECM Granger causality.	CO_2 emission, GDP, GDP square, energy consumption	Yes
Pao and Tsai [38]	1980–2007	BRIC countries.	Pedroni, Kao and Fisher cointegration, OLS and VECM Granger causality.	CO_2 emission, GDP, GDP square, energy consumption and FDI.	Yes
Pao and Tsai [39]	1980–2007	Brazil	Grey prediction model (GM).	CO_2 emission, GDP, GDP square and energy consumption.	Yes
Wang et al. [56]	1995–2007	China	Pedroni cointegration and VECM Granger causality.	CO_2 emission, GDP, GDP square and energy consumption.	No
Arouri et al. [7]	1981–2005	Middle East and North African (MENA) countries.	The Cross Correlated Effects (CCE) estimation procedure for cointegration and using the common correlated effects mean group (CCE-MG) estimates.	CO_2 emission, GDP, GDP square and energy consumption.	Yes
Esteve and Tamarit [19]	1857–2007	Spain	Kejriwal–Perron tests for testing multiple structural breaks in cointegrated regression, Estimation of long-run relationships: Stock–Watson–Shin cointegration test.	CO_2 emission, GDP and GDP square	Yes
Jayanthakumaran et al. [24]	1971–2007	China and India.	ARDL modelling approach	CO_2 emission, GDP, GDP square, and energy consumption	Yes
Pao et al. [40]	1980–2009	China	Grey prediction model (GM).	CO_2 emission, GDP, GDP square and energy consumption.	No
Chandran and Tang [13]	1971–2008	Association of South East Asian Nations (ASEAN).	Johansen cointegration and VECM Granger causality.	CO_2 emission, GDP, GDP square, transport energy consumption and foreign direct investment (FDI).	No
Govindaraju and Tang [22]	1965–2009	India and China	Bayer and Hanck combine cointegration test, VECM Granger	CO_2 emission, GDP, GDP square, coal consumption.	No

(continued on next page)

Table 1 *(continued)*

Author	Period	Country/ region/ organization	Methodology	Variables used in the study	EKC Hypothesis
			causality.		
Kohler [28]	1960–2009	South Africa	ARDL modelling approach, Johansen cointegration and VECM Granger causality.	CO_2 emission, GDP, GDP square, energy consumption and trade openness.	Yes
Ozcan [34]	1990–2008	Middle East countries.	Westerlund panel cointegration test, FMOLS and VECM Granger causality.	CO_2 emission, GDP, GDP square and energy consumption.	Yes
Saboori and Sulaiman [45]	1980–2009	Malaysia	ARDL modelling approach, Johansen cointegration and VECM Granger causality.	CO_2 emission, GDP, GDP square, electricity consumption, oil consumption, coal consumption and gas consumption.	Yes
Saboori and Sulaiman [46]	1971–2009	ASEAN countries	ARDL modelling approach and VECM Granger causality.	CO_2 emission, GDP, GDP square and energy consumption.	Yes for Thailand and Singapore.
[48,49]	1965–2008	South Africa	ARDL modelling approach and VECM Granger causality.	CO_2 emission, GDP, GDP square, coal consumption, financial development and trade openness.	Yes
Bella et al. [12]	1965–2006	Organization for Economic Cooperation and Development	Larsson, Lyhagen, and Löthgren (LLL) cointegration and VECM Granger causality.	CO_2 emission, GDP, GDP square and electricity consumption.	Yes
Farhani et al. [21]	1971–2008	Tunisia.	ARDL modelling approach to cointegration and VECM Granger causality.	CO_2 emission, GDP, GDP square, energy consumption and trade openness.	Yes
Kivyiro and Arminen [27]	1971–2009	Sub-Saharan	ARDL modelling approach and VECM Granger causality.	CO_2 emission, GDP, GDP square, energy consumption and FDI.	Yes in most countries.
Saboori et al. [44]	1977–2008	OPEC countries	ARDL modelling approach and Toda–Yamamoto–Dolado–Lütkepohl (TYDL) causality tests.	CO_2 emission, GDP, GDP square, capital, labor, oil price and oil consumption.	Yes
Al-mulali et al. [2]	1980–2008	Ninety-three countries base on income level.	Panel fixed effects and the generalized method of moments (GMM).	Ecological footprint, GDP, GDP square, energy consumption, trade openness, urbanisation and financial development.	Yes for the upper-middle and high income countries.
Al-mulali et al. (2015b)	1981–2011	Vietnam	ARDL modelling approach and VECM Granger causality.	CO_2 emission, GDP, GDP square, capital, labor, exports, imports, fossil fuels and renewable energy consumption	No
Apergis and Ozturk [6]	1990–2011	Asian countries.	Nyblom–Harvey Fisher–Johansen and Pedroni cointegration, FMOLS, DOLS, pooled mean group estimator (PMGE) and mean group MG.	CO_2 emission,GDP, GDP square, GDP cubic, population density, land, industry shares in GDP and quality of institutions.	Yes
Baek [9]	1960–2010	Arctic countries.	Autoregressive distributed lag (ARDL) modelling approach to cointegration.	CO_2 emission, GDP, GDP square, GDP cubic and energy consumption.	Yes
Baek [10]	1980–2009	Nuclear producing countries.	Pedroni cointegration, dynamic OLS and FMOLS	CO_2 emission, GDP, GDP square, nuclear energy consumption and energy consumption.	Yes
Begum et al. [11]	1970–2009	Malaysia	ARDL modelling approach, DOLS and the Sasabuchi–Lind–Mehlum tests.	CO_2 emission, GDP, GDP square, energy consumption and population.	Yes
Kasman and Duman [26]	1992–2010	European Union countries.	Pedroni and Kao cointegration, fully modified ordinary least square and VECM Granger causality.	CO_2 emission, GDP, GDP square, energy consumption, trade openness and urbanisation.	Yes
Ozturk and Al-mulali (2015)	1996–2012	Cambodia	Two-stage least square and GMM.	CO_2 emission, GDP, GDP square, energy consumption, urbanisation, the control of corruption and governess.	No
Shahbaz et al. [50]	1980–2012	African Countries.	Johansen cointegration test, Pedroni cointegration, vector error correction model (VECM) Granger causality.	CO_2 emission, GDP, GDP square and electricity intensity.	Yes
Tang and Tan [54]	1976–2006	Vietnam	Johansen cointegration and VECM Granger causality.	CO_2 emission, GDP, GDP square, energy consumption and FDI.	Yes
Yin et al. [61]		China	Panel random effects model.	CO_2 emission, GDP, GDP square, energy efficiency, population, trade openness, FDI, coal consumption, secondary industrial output and renewable energy consumption.	Yes

Table 2
Conventional unit root tests, China.

Variables	ADF unit root test T-statistic	PP unit root test T-statistic	DF-GLS unit root test T-statistic
$\ln C_t$	−1.886(1)	−1.665 (0.751)	−0.826 (0)
$\Delta \ln C_t$	−4.192*** (3)	5.851*** (3)	−2.896* (0)
$\ln Y_t$	−3.048(2)	−1.733 (4)	−1.167(2)
$\Delta \ln Y_t$	−7.307*** (1)	−5.227*** (4)	−4.513*** (0)
$\ln Y_t^2$	−1.928 (2)	−1.036 (4)	−1.204 (1)
$\Delta \ln Y_t^2$	−7.307*** (1)	−5.227*** (4)	−4.513*** (0)
$\ln H_t$	−2.251 (0)	−2.338 (0)	−1.723 (0)
$\Delta \ln H_t$	−7.515*** (0)	−7.515*** (0)	−7.552*** (0)
$\ln U_t$	−1.399 (4)	−1.383 (3)	−2.426 (3)
$\Delta \ln U_t$	−3.894**(3)	−4.926*** (4)	−4.030*** (0)

***, ** and * denote significance at the 1%, 5% and 10% levels, respectively. The Akaike Information Criterion is used to choose the optimal lag length, after imposing a maximum lag of 4 in the ADF and DF-GLS tests. The Bartlett with Newey-West bandwidth is employed in choosing the optimal lag for the PP test. A constant and time trend are added in each of the equations.

Table 3
Conventional unit root tests, India.

Variables	ADF unit root test T-statistic	PP unit root test T-statistic	DF-GLS unit root test T-statistic
$\ln C_t$	−2.142 (0)	−2.142 (0)	−1.436(0)
$\Delta \ln C_t$	−8.047***(0)	−8.064*** (1)	−8.113***(0)
$\ln Y_t$	−0.593(0)	−0.350(4)	−0.923 (3)
$\Delta \ln Y_t$	−5.408***(3)	−8.032*** (4)	−7.464***(0)
$\ln Y_t^2$	−0.162 (0)	0.128 (4)	−0.113 (1)
$\Delta \ln Y_t^2$	−7.457*** (0)	−7.588*** (4)	−7.302*** (0)
$\ln H_t$	−2.916 (0)	−2.916 (0)	−2.310 (0)
$\Delta \ln H_t$	−6.301*** (0)	−6.285*** (2)	−6.367*** (0)
$\ln U_t$	−2.829 (2)	−1.375 (3)	−2.257(2)
$\Delta \ln U_t$	−3.669**(1)	−4.822**(2)	−3.253** (1)

*, ** and *** indicate significance at 10%, 5% and 1% levels respectively. For other explanations, check the footnote of Table 2.

$$\begin{aligned}\Delta \ln H_t = \; & \alpha_{31} + \sum_{i=1}^{k} \alpha_{32}\Delta \ln C_{t-i} + \sum_{i=1}^{k} \alpha_{33}\Delta \ln Y_{t-i} + \sum_{i=1}^{k} \alpha_{34}\Delta \ln Y_{t-i}^2 \\ & + \sum_{i=1}^{k} \alpha_{35}\Delta \ln H_{t-i} + \sum_{i=1}^{k} \alpha_{36}\Delta \ln U_{t-i} + \\ & + \alpha_{37}T + \alpha_{38}\Delta D_1 + \alpha_{39}\Delta D_2 + \alpha_{310}ECT_{t-1} + \upsilon_t\end{aligned} \tag{6}$$

$$\begin{aligned}\Delta \ln U_t = \; & \alpha_{41} + \sum_{i=1}^{k} \alpha_{42}\Delta \ln C_{t-i} + \sum_{i=1}^{k} \alpha_{43}\Delta \ln Y_{t-i} + \sum_{i=1}^{k} \alpha_{44}\Delta \ln Y_{t-i}^2 \\ & + \sum_{i=1}^{k} \alpha_{45}\Delta \ln H_{t-i} + \sum_{i=1}^{k} \alpha_{46}\Delta \ln U_{t-i} + \\ & + \alpha_{47}T + \alpha_{48}\Delta D_1 + \alpha_{49}\Delta D_2 + \alpha_{410}ECT_{t-1} + \upsilon_t\end{aligned} \tag{7}$$

The *F*-test of joint significance of these lagged terms was utilized to examine the short-run Granger causality. According to the standard causality practice, if the F-statistic is not significant at 10% or better, there is no short run causality. On the other hand, there is evidence for short run causality, when the F-statistic is significant at 10% or better. For instance, causality runs from per capita real GDP, per capita square of real GDP to emission in the short term if the joint null hypothesis is rejected as $\alpha_{13} \neq \alpha_{14} \neq 0\forall$. The parameter associated with ECT signifies the speed of adjustment to the long-run equilibrium. The *t*-test for the coefficient of ECT provided the estimates of the long-run Granger causality. There is joint long run causality from the independent variables, when the t-statistic is significant at 10% or better and the coefficient is negative in the relevant equation. A significant ECT coefficient suggests that previous equilibrium errors play important roles in determining present values. If the coefficient α_{110} is significant in Eq. (4), then per capita real GDP, per capita square of real GDP, per capita hydroelectricity use, and urban population ratio Granger cause emission in the long term. The same analysis is applied to the remaining equations.

3. Results

We began the analyses of the data by testing the unit root properties of all the series in the current study. As a starting point, we initially use the traditional unit root tests, namely the Said and Dickey [47] or ADF test, Phillips and Perron [43] or PP test, Elliott et al. [18] or DF-GLS test to examine the unit root in the five series. The ADF test provides for serial correlation by allowing the lag terms of the dependent variable among the regressors in the unit root test. The PP test provides for autocorrelation by adopting a heteroskedasticity- and autocorrelation-consistent covariance matrix estimator. One benefit of the PP test relative to the ADF test is that the PP test is not susceptible to the general forms of heteroscedasticity in the error term. DF-GLS allows the time series to be transformed via a generalized least squares (GLS) regression before the traditional ADF is performed. According to Elliott et al. [18], the DF-GLS test performs better than the standard ADF test, when the sample size is small. The findings which are stated in Tables 2, 3 illustrate that we cannot reject the null of unit root when the variables are in level. Nevertheless, the null hypothesis can be rejected when the series are in their first differences.

The reliability of these tests tends to become questionable when the series encounter structural shifts. As a result, the estimates from the Lee and Strazicich [29,30] tests are subsequently stated. As displayed in Table 4, we are unable to reject the null of nonstationarity for all the series when their level forms are examined in China. When the variables are expressed in first difference, the null of nonstationarity can be rejected in all the variables for China. It is noted that two shifts are significant in all cases with the exclusion of when per capita real GDP and per capita hydroelectricity use are expressed in level form. Almost 33% of the structural breaks (or 6 of the 18 structural shifts) are within the latter part of 1970s, while 39% of the breaks (or 7 of the 18 structural shifts) are within the latter part of 1990s.

In the case of India, the output of the Lee and Strazicich [29,30], presented in Table 5, indicates that we cannot reject the null hypothesis when we express the series in their level forms. However, we are able to reject the null of nonstationarity when the series are expressed in their first differences. It is noted that two shifts are significant in all cases with the exception of when per capita real GDP as well as per capita hydroelectricity use are specified in their first differences and urban population ratio is expressed in level form. About 29% of the structural breaks (or 5 of the 17 structural breaks) are within the earlier part of 1980s while another 24% of the breaks (or 4 of the 17 structural breaks) are within the latter part of 1990s.

After observing that the series are integrated of order (1), we turn to the cointegration test. By applying the ARDL method, it is noted that there is long-run link in the variables for the two countries. The cointegration test, which is reported in Table 6, show that the F-statistics (7.005) is above the upper bounds critical values at 1% significance level (6.684), when the per capita emission is expressed as the dependent series, while the independent variables are real GDP, real GDP square, per capita hydroelectricity consumption and urban population ratio in China. The F-statistics (5.152) is above the upper bounds critical values at 1% significance level (5.064) in the case of India. The diagnostic tests suggest that there is no autocorrelation of the error term. Furthermore, the autoregressive conditional heteroskedasticity (ARCH) test (which is a test for heteroscedasticity of the model) denotes that the errors are homoscedastic and therefore, the estimates are efficient.[3] Using the Jarque-Bera normality tests, we cannot

[3] The test is based on the notion that heteroscedasticity exists if the variance of the current error term depends on the error terms of the previous period.

Table 4
LM Unit root test, China.

Variable	T-statistics	TB1	TB2	DU1	DT1	DU2	DT2
$\ln C_t$	−4.125[1]	1977	1995	0.110 (1.793)	−0.118*** (−4.512)	−0.079 (−1.267)	0.100*** (4.738)
$\Delta \ln C_t$	−6.060**[1]	1978	1997	0.018 (0.278)	−0.108*** (−3.865)	−0.052 (−0.802)	0.093*** (3.237)
$\ln Y_t$	−2.042[1]	1981	–	−0.055*** (−1.765)	0.049*** (4.998)	–	–
$\Delta \ln Y_t$	−10.228***[1]	1975	1997	0.100*** (3.658)	−0.021 (−1.544)	−0.053 (−2.039)	0.041*** (3.931)
$\ln Y_t^2$	−4.883 [1]	1980	2004	−0.062 (−0.191)	0.246** (2.011)	−0.291 (−0.860)	1.012*** (5.468)
$\Delta \ln Y_t^2$	−8.452***[1]	1976	1991	−2.334*** (−6.192)	0.882*** (5.326)	0.361 (1.062)	−0.646*** (−4.787)
$\ln H_t$	−4.381[0]	1980	1999	0.009(0.133)	0.129***(3.464)	−0.044(−0.631)	0.027** (2.086)
$\Delta \ln H_t$	−7.662***[1]	1999	–	0.059 (0.803)	−0.040* (−1.664)	–	–
$\ln U_t$	−5.052[1]	1979	1996	−0.028*** (−6.103)	0.036*** (17.468)	0.001 (0.095)	0.005 (2.039)
$\Delta \ln U_t$	−7.119*** [3]	1978	1997	−0.014*** (−4.679)	0.015*** (6.771)	0.002 (0.756)	−0.013*** (−8.154)

***, ** and * denote significance at the 1%, 5% and 10% levels, respectively. TB is the estimated break points. Critical values are in Lee and Strazicich [29,30]. *TB1 and TB2* are the structural break dates. DU1 and DU2 are the dummy variables for breaks in intercept, while DT1 and DT2 are the dummy variables for trend breaks. Critical values for the other coefficients are based on the standard t-distribution 1.65, 1.96, 2.58 We use the Akaike Information Criterion to choose the optimal lag length (which are presented in the brackets) after imposing a maximum lag of 4. The t-statistics are presented in the in parenthesis. The estimates do not show any evidence of serial correlation and heteroscedasticity.

Table 5
LM unit root test, India.

Variable	T-statistics	TB1	TB2	DU1	DT1	DU2	DT2
$\ln C_t$	−4.921[4]	1983	1999	−0.009 (−0.333)	0.080 (4.857)	0.047 (1.730)	−0.045*** (−3.090)
$\Delta \ln C_t$	−8.881***[0]	1988	1999	0.059** (2.283)	−0.042*** (−4.457)	−0.031 (−1.204)	0.064*** (4.929)
$\ln Y_t$	−4.491 [3]	1985	1998	−0.031 (−1.331)	0.052*** (5.058)	−0.008 (−0.344)	0.040*** (4.039)
$\Delta \ln Y_t$	−8.066***[0]	2002	–	0.027 (0.971)	−0.020 ** (−2.106)	–	–
$\ln Y_t^2$	−4.280 [3]	1986	2002	−0.341 (−1.231)	0.587*** (5.005)	−0.125 (−0.441)	0.577*** (4.778)
$\Delta \ln Y_t^2$	−6.443*** [1]	1982	2003	0.827*** (2.566)	−0.488*** (−3.952)	−0.303 (−0.903)	0.553*** (3.645)
$\ln H_t$	−4.637[1]	1980	2005	0.163* (1.921)	−0.191*** (−4.719)	0.103 (1.192)	0.095** (2.460)
$\Delta \ln H_t$	−6.654*** [0]	2003	–	0.300*** (2.985)	−0.117*** (−2.981)	–	–
$\ln U_t$	−1.048[0]	1983	–	0.004** (2.076)	−0.004*** (−5.963)	–	–
$\Delta \ln U_t$	−6.373**[4]	1983	1999	0.001 (0.651)	−0.004*** (−6.976)	−0.001 (−0.350)	0.003*** (6.35)

***, ** and * denote significance at the 1%, 5% and 10% levels, respectively. For other explanations, check the footnote of Table 4.

reject the null of normality of the model. Therefore, we conclude the errors are normally distributed.

Since we are able to ascertain the incidence of long-run relationship in the series, the next step is to analyse the impact of the variables on per capita emission and verify the existence of the EKC. The short and long-run results related to China and India are described in Table 7. The outcomes show that per capita GDP and urban population ratio positively affect the emission per capita in the long-run. However, the per capita square of GDP and per capita hydroelectricity use have negative influence on per capita emission in the long-run. The foregoing results provide evidence for EKC hypothesis or an inverted U-shaped relationship between economic growth and CO_2 emission in the country. While the break in 1977 has negative long-run influence on per capita emission, the break of 1995 has positive long-run effect on per capita emission. The results in the short run are not materially different from the foregoing outcomes.

The estimates further show that GDP and urbanisation long-run have positive influence on emission in India. Real GDP square and per capita hydroelectricity use have long-run negative impact on emission in the country. Therefore, there is evidence for EKC hypothesis in the country. The short run estimates are very similar to the long run coefficients. The diagnostics tests suggest the models are free from serial correlation and heteroscedasticity. There is no specification and normality concerns as the test statistics provide evidence for well specified model and normally distributed error. The cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) tests are two popular tests used to check the stability of the model. We observe that the tests largely support the stability of the coefficients of the regressions.

Table 6
The ARDL Cointegration Analysis.

Country	Model: $F_{\ln C_t}(\ln C_t / \ln Y_t,\ \ln Y_t^2,\ \ln H_t,\ \ln U_t)$			Diagnostic tests			
	Lag length	Structural Break	F-statistics	χ^2_{SERIAL}	χ^2_{ARCH}	χ^2_{NORMAL}	χ^2_{RESET}
China	(3,1,1,0,3)	1977, 1995	7.005***	0.686[1]	0.802[1]	0.160[1]	0.926[1]
India	(2,4,4,0,2)	1983, 1999	5.152**	0.113[1]	0.434[1]	0.605[2]	0.480[1]
Significant level	Critical values (T= 49)[#]						
	Lower bounds $I(0)$	Upper bounds $I(1)$					
1 per cent level	5.184	6.684					
5 per cent level	3.834	5.064					
10 per cent level	3.240	4.350					

*** imply 1% level of significance. We use the Akaike Information Criterion to choose the optimal lag length. The brackets contain the order of diagnostic tests. Narayan [33] provide the critical values and it is for case V: unrestricted intercept and unrestricted trend.

Table 7
Long run and short run analyses.

Panel A: Long-run elasticities			
Independent variable	China	Independent Variable	India
$\ln Y_t$	5.282*** (4.462)	$\ln Y_t$	1.620** (2.047)
$\ln Y_t^2$	−0.463*** (3.321)	$\ln Y_t^2$	−0.060*** (−5.639)
$\ln H_t$	−0.137*** (9.970)	$\ln H_t$	−0.124*** (−3.715)
$\ln U_t$	12.837* (1.775)	$\ln U_t$	9.478*** (7.263)
Constant	3.048*(1.751)	Constant	7.535 (1.556)
Trend	0.118*** (4.174)	Trend	−0.130*** (−6.189)
Dummy 1977	−0.209*** (−2.538)	Dummy 1983	−0.052*** (−6.158)
Dummy 1995	0.287*** (6.888)	Dummy 1999	0.032*** (7.553)
Panel B: Short run elasticities			
Independent Variable	China	Independent Variable	India
$\Delta \ln Y_t$	3.197*** (3.158)	$\Delta \ln Y_t$	1.121*** (2.616)
$\Delta \ln Y_t^2$	−0.315*** (−3.550)	$\Delta \ln Y_t^2$	−0.041*** (−3.283)
$\Delta \ln H_t$	−0.016** (−2.192)	$\Delta \ln H_t$	−0.152*** (−3.942)
$\Delta \ln U_t$	2.790* (1.790)	$\Delta \ln U_t$	15.079*** (3.793)
ΔConstant	−1.198*** (−3.876)	ΔConstant	9.194* (1.648)
ΔTrend	0.014 (−0.422)	ΔTrend	−0.159*** (−5.764)
ΔDummy1977	−0.025 (−0.525)	ΔDummy1977	0.064*** (7.068)
ΔDummy1995	0.034*** (3.755)	ΔDummy1995	−0.039*** (−4.606)
ECM (−1)	−0.119*** (−2.792)	ECM (−1)	−0.220*** (−5.713)
Adjusted R^2	0.935		0.716

Diagnostics test			
Test	China	Test	India
χ^2_{SERIAL}	0.632[1]	χ^2_{SERIAL}	0.887[1]
χ^2_{ARCH}	0.455[1]	χ^2_{ARCH}	0.457[1]
χ^2_{NORMAL}	0.106[2]	χ^2_{NORMAL}	0.172[2]
χ^2_{RESET}	0.358[1]	χ^2_{RESET}	0.159[1]
CUSUM	Stable	CUSUM	Stable
CUSUMSQ	Stable	CUSUMSQ	Stable

***, ** and * denote significance at the 1%, 5% and 10% levels, respectively. We employ the Akaike Information Criterion to choose the optimal lag length. The parentheses contain the probability values, while the bracket contains the order of diagnostic tests.

Table 8
The VECM granger causality analysis in China.

Dependent variable	Direction of causality				
	Short run				Long run
	$\Delta \ln C_{t-i}$	$\Delta \ln Y_{t-i}$, $\Delta \ln Y^2_{t-i}$	$\Delta \ln H_{t-i}$	$\Delta \ln U_{t-i}$	ECT_{t-1}
$\Delta \ln C_t$		3.972** (0.014)	2.250* (0.095)	2.257* (0.094)	−0.143*** [4.009]
$\Delta \ln Y_t \Delta \ln Y_t^2$	1.829 (0.158)		0.538 (0.707)	1.099 (0.381)	−1.234*** [−8.798]
$\Delta \ln H_t$	1.912 (0.143)	5.028*** (0.005)		3.824** (0.016)	−0.815*** [−7.721]
$\Delta \ln U_t$	2.186 (0.103)	2.506* (0.701)	0.215 (0.927)		−0.446*** [0.009]

***, ** and * denote significance at the 1%, 5% and 10% levels, respectively. The parentheses contain the probability values, while the brackets contains the t-statistics.

Then we examined the causal relationship between per capita real GDP (real GDP per capita square), per capita hydroelectricity consumption, per capita emissions, and urban population ratio in China and India. In China, as indicated in Table 8, we found strong evidence of dual Granger causality between per capita real GDP (real GDP per capita square) and CO_2 emissions per capita in the long-run. However, there is a unidirectional causality from per capita real GDP (real GDP per capita square) to emission per capita in the short-run. There is evidence for long-run bilateral causation between per capita hydroelectricity consumption and per capita emissions, while there is short-run causation from per capita hydroelectricity consumption to emissions per capita. In addition, we found strong evidence of dual Granger causation in per capita real GDP (real GDP per capita square) and per capita hydroelectricity use in the long-run. There is long-run dual causality between urbanisation and the other variables. Moreover, there is short-run unidirectional causation from urbanisation to per capita emission and per capita hydroelectricity consumption, but unidirectional causation from per capita real GDP (real GDP per capita square) to urbanisation.

The causality results of India are depicted in Table 9. We found strong indication of bidirectional Granger causality between real GDP per capita (real GDP per capita square) and CO_2 emissions per capita in the long-run. However, there is unidirectional causality from real GDP per capita (real GDP per capita square) to per capita emission in the short run. There is evidence for long-run bilateral causality between emissions per capita and per capita hydroelectricity consumption, while causality exists from per capita hydroelectricity consumption to emission per capita in the short run. We found strong evidence of bidirectional Granger causality between real GDP growth per capita and per capita hydroelectricity use in the long-run, but no causality in the short-run. There is long-run bidirectional causality between urbanisation and the other variables. Moreover, there is short-run unidirectional causality from urbanisation to economic growth per capita in the short-run.

Table 9
The VECM granger causality analysis in India.

Dependent variable	Direction of causality				
	Short run				Long run
	Δ ln C_{t-i}	Δ ln Y_{t-i}, Δ ln Y^2_{t-i}	Δ ln H_{t-i}	Δ ln U_{t-i}	ECT_{t-1}
Δ ln C_t		4.907*** (0.008)	4.438*** (0.008)	4.443*** (0.008)	−0.316*** [−2.630]
Δ ln $Y_t \Delta$ ln Y_t^2	0.507 (0.731)	–	1.101 (0.380)	4.539*** (0.008)	−0.353* [−1.891]
Δ ln H_t	1.057 (0.400)	1.741 (0.175)		1.549 (0.222)	−0.378** [−2.743]
Δ ln U_t	2.046 (0.121)	1.333 (0.288)	2.095 (0.114)	–	−0.048** [−2.069]

***, ** and * denote significance at the 1%, 5% and 10% levels, respectively. The parentheses contain the probability values, while the brackets contains the t-statistics.

4. Discussion

There is an indication for EKC in both countries and the long-run causality between emissions and income growth indicates that real GDP per capita (real GDP per capita square) influences CO_2 emission and vice-versa. Since there is evidence that causality flows from income to emission, this is corroborating evidence to the findings observed in the ARDL regression estimates that EKC exist in both countries. Related studies with similar results include Jalil and Mahmud [23] for China, Jayanthakumaran et al. [24] for China and India, Tiwari et al. [55], Kanjilal and Ghosh [25], as well as Govindaraju and Tang [22] for India. However, Wang et al. [56] was unable to observe EKC in China.

The positive relationship between income and emission is supported by the fact that these countries are populated by several industries that are fossil-fuel intensive. In China and India, coal and oil, which are significant sources of CO_2 emission, dominate the energy mix of the countries. In 2013, 67%, 18%, and 5% of the energy mix of China was constituted by coal, oil, and natural gas, respectively. In India, coal, oil, natural gas, and hydroelectricity constituted 55%, 29%, 8%, and 5%, respectively in 2013 [16,17] The evidence for positive relationship between income and emission is also supported by the fact that both emission and income have been growing simultaneously, over the years. The data show that emission has increased by more than 7700% for the period of 1965–2013, while the rise in real GDP per capita exceeded 3200% in China. In India, emission per capita rose by 170%, while the rise in real GDP per capita exceeded 477% for the period of 1965–2013. With continuous increase in hydroelectricity use and, possibly, focus on less-energy intensive industries, the rising pace of emission can be addressed.

Most of the breaks in China are within the latter part of 1970s and the latter part of 1990s. The shifts in the latter part of 1970s are associated with the first phase of the key economic transformation in China. From 1979, China has been progressing towards achieving a more open economy along with increasing international trades with several countries. The improvements also focussed on the rural areas and provided rural households rights to utilise jointly-owned land, and the right to dispose the marginal output thereof in the open market. There are improvements in quality of lives of many Chinese households and the economy has witnesses encouraging economic expansion [14]. Most of the shifts in the latter part of 1990s are due to the pattern of economic transformation that arose during that period in the country. The administrative and regulatory reform of rural-urban exodus policies, the tax system, the banking system, foreign trade, and foreign investment erased numerous binding restrictions on economic growth. Although Asian economic crisis occurred during this period, it did not have much impact on the country.

Several of the shifts in India are within the earlier part of 1980s and in the latter part of 1990s. The break dates of the earlier part of 1980s matched the oil crisis experienced within that time, which caused economic downturn in the country [24]. The break date of 1999 is related to the Asian financial crisis of that period. The crisis, which occurred in the Association of Southeast Asian Nations, resulted in considerable flows of capital from the ASEAN countries to economies including India. Generally**,** occurrences of structural breaks are practically expected in the periods of policy changes and economic turbulences. The occurrence of the crisis resulted in altering the relationship among the series [25,52].

The findings also revealed that there is long-run negative causality from hydroelectricity consumption to emission with feedback effect from emission in both countries. At the same time, there is bidirectional causality between hydroelectricity consumption and economic growth. The findings are consistent with the results of Wang et al. [56] for China as well as Govindaraju and Tang [22] and Tiwari et al. [55] for India. Zhang and Cheng [63] study for China contradicted the findings of the present study. The results obtained in this study came with no surprise as more intensive use of hydroelectricity might be associated with less use of fossil fuels. As it currently stands, hydroelectricity consumption accounts for a small share of the energy mix and there is still room for improvement. The share of hydroelectricity in the total energy mix was 7% and 5% in China and India, respectively in 2013 [16,17]. The degree of development is relatively small as only 27.3% of the total potentials are developed [31].

The Chinese government has rolled out programmes and incentives that will increase the use of hydroelectricity in the country. The Three Gorges Dam hydroelectric facility, which is the biggest hydroelectricity system in the globe, is located in China. It commenced operations in 2003 and the construction was completed in 2012 [16]. Moreover, China is stimulating investment in renewable energy and related transmission infrastructure via various economic and financial incentives. Chinese firms spent $65 billion in renewable energy schemes in 2012, which is 20% higher than investments in the previous year, and they arranged $473 billion on clean energy investments during the period of 2011–2015 [16].

India has enormous hydroelectricity resources as it is in the fifth place in terms of available hydroelectricity potentials in the globe. Several states in the country including Himachal Pradesh, Jammu, Kashmir, and Uttarakhand have significant river systems [17]. However, less than 25% has been established or earmarked for development [17,8]. Furthermore, the authorities have undertaken a series of measures to confront the concerns of prospective developers such as the provision of open access and trading, joint venture initiatives, devising of transparent bidding procedures, and notification of tariff determination procedures [8].

There is bidirectional causality between urbanisation and the remaining series. The fact that the urban centers in these countries hold the financial ace to the prosperity of the country makes this regression result not too surprising. Shanghai, Beijing, Hangzhou, Shenzhen, and Guangzhou are the commercial nerves of China, while Mumbai, Delhi, Banglore, and Hydderabad are the financial centers in India. Concurrently, they are the most urbanized cities in the two countries. Urban population has increased by more than 200 million in China since 2001 and it accounts for more than 80% of China's GDP [59]. Although India is relatively less urbanized than many countries, its urban population has increased by over 100 million since 2001. Cities are increasingly becoming the engine of the national economy, accounting for about 60% of India's GDP [57]. Thus, the role of urbanisation cannot be overemphasized. This result is not peculiar to these countries as global towns and cities are the hubs of prosperity—over 80% of world economic activities are generated in the urban areas that are just half of the world's population. Economic agglomeration promotes proficiency and it generates more income opportunities [58].

5. Conclusion and policy implications

The primary objective of this study is to test the link between hydroelectricity consumption, output, urbanisation, and CO_2 emissions in China and India during the period of 1965–2013. We augmented the conventional ARDL approach with structural breaks to investigate the long-run link between the variables. Our results indicated that all the variables are cointegrated for the long-run relationship when all the variables are expressed as the dependent variables. The study was able to establish EKC in these countries.

The findings further revealed that there is long-run negative causality from hydroelectricity consumption to emission with feedback effect from emission in both countries. With the use of more hydroelectricity in the energy mix, the utilisation of fossil fuels, which are responsible for most of the CO_2 emission, is likely to decrease. The policy implication emanating from these findings is that more hydroelectricity use is likely to decrease emission. Moreover, bidirectional causality between hydroelectricity consumption and economic growth does exist. This indicates that hydroelectricity use has boosted economic growth with economic activities also positively affecting the use of hydroelectricity. It implies that energy plays an essential part in stimulating economic growth, which means that decreasing energy use arbitrarily may have an adverse influence on the countries' economic development. Stimulated by continuous economic growth, increase in income levels, and rise in availability of goods and services, the incremental energy demands, including hydroelectricity use in both China and India, have been increasing. Thus, policies that reduce the use of hydroelectricity will have an adverse effect on economic growth in these countries. Any shortage of the hydroelectricity will also hinder economic growth. In addition, reduction in the output will adversely affect the demand for hydroelectricity. Shock to one of these variables will be passed to the other and the chain will persist via the feedback flow. Therefore, expansionary hydroelectricity policies are beneficial to both China and India.

There is bidirectional causality between urbanisation and all the remaining series. The bidirectional causality between urbanisation and emission suggests that urbanisation is also responsible for emission. Moreover, the feedback relationship between urbanisation and income suggests that urbanisation is a determinant of economic growth. As such, urbanisation is an instrument of economic development. Policymakers, who hope to increase the long-term economic growth by supporting urbanisation, are likely to achieve such objective.

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