Environmental Science and Pollution Research (2019) 26:14562–14569
https://doi.org/10.1007/s11356-019-04835-6

RESEARCH ARTICLE

Testing environmental Kuznets curve for the USA under a regime shift: the role of renewable energy

Umit Bulut[1]

Received: 4 January 2019 / Accepted: 8 March 2019 / Published online: 14 March 2019

Abstract

The goals of this paper are to examine whether the environmental Kuznets curve (EKC) holds and to investigate whether renewable energy consumption can decrease CO_2 emissions in the USA using monthly data spanning the period 2000:M01–2018:M07. For these purposes, the paper employs a cointegration test with a regime shift and observes the long-run coefficients before and after the regime shift. The findings support the presence of the EKC. The findings also indicate that renewable energy consumption has negative effects on CO_2 emissions, while these effects are greater when the share of renewable energy consumption in total energy consumption is higher in the USA. Theoretical and practical implications for these findings are discussed.

Keywords CO_2 emissions · Environmental Kuznets curve · The US economy · Renewable energy consumption · Cointegration test with a regime shift · DOLS estimator

Introduction

In today's world, one of the most discussed problems has been environmental degradation within the scope of global warming (Yavuz 2014). Global warming has serious influences on economies and the ecological system, namely higher temperatures, longer frost-free seasons, changes in precipitation patters, more droughts and heat waves, and stronger hurricanes (National Aeronautics and Space Administration 2018). The main cause of global warming is CO_2 emissions stemming from the use of fossil sources, such as coal, oil, and natural gas (Guris 2016).

As a result of the environmental degradation caused by fossil sources, the empirical energy literature has focused on two considerable research fields. The first group of the studies examines whether renewable energy, as a clean energy source, can decrease environmental problems arising from the utilization of fossil energy sources. The second group of the studies investigates whether environmental destruction may begin to decrease after economies reach a threshold value for income. This phenomenon is defined as the environmental Kuznets curve (EKC) hypothesis in the energy literature. Many papers have searched for the validity of the EKC since the seminal paper of Grossman and Krueger (1995), who adapt the original work of Kuznets (1955) about income inequality and economic growth for environment. The EKC hypothesis implies that the level of environmental degradation first increases as a result of economic growth and then begins to decrease after income reaches a threshold value/turning point (Stern 2004; Du et al. 2018). Accordingly, at the onset of the pathway to economic growth, more energy sources are employed for production activities and so more waste and pollutant gas emissions emerge (Bilgili et al. 2016; Pata 2018; Sun and Fang 2018). During this period, economic development process is described by high energy consumption that is mainly met from fossil energy sources as fossil energy sources are cheaper compared to renewable energy sources (Sarkodie and Strezov 2018a). The reason why the environmental quality increases after the threshold value of income is that a high-income economy can substitute dirty and old technologies with clean and new technologies (Copeland and Taylor 2003). Put differently, the high-income economy tends to spend more on research and development and to

Responsible editor: Philippe Garrigues

✉ Umit Bulut
ubulut@ahievran.edu.tr

[1] Faculty of Economics and Administrative Sciences, Department of Economics, Kirsehir Ahi Evran University, Kirsehir, Turkey

replace fossil energy sources with renewable energy sources (Sarkodie and Strezov 2018a). Besides, when an economy grows, the production structure of this economy shifts from industry sector to services sector, which is not an energy-intensive sector (Ulucak and Bilgili 2018). Hence, the EKC hypothesis implies an inverted U-shaped relationship between income and environmental destruction. The existence of the EKC implies that environmental degradation can be controlled by economic growth (Su and Chen 2018).

In the energy economics literature, there are lots of papers examining the validity of the EKC and/or the influence of renewable energy consumption on CO_2 emissions. Some of these papers focus on the USA, having the greatest economy and one of the greatest energy-consuming countries in the world. When one examines the estimation methodologies of these papers, he/she will observe that all these papers assume the long-run coefficients do not change over time. Put differently, they suppose a fixed coefficient for independent variables throughout the observed period. Starting from this point of view, this paper examines whether the EKC dominates in the USA and investigates the influence of renewables consumption on CO_2 emissions for the USA using monthly data from January 2000 to July 2018 under the presence of a regime shift. In other words, the empirical analysis in the paper employs the cointegration test propounded by Gregory and Hansen (1996) and considers a change in the coefficients of the independent variables after the regime shift. In doing so, the paper is supposed to make a significant contribution to the energy economics literature as it is the first paper investigating the presence of the EKC and the influence of renewables consumption on CO_2 emissions for the USA under a regime shift. More clearly, the distinctive characteristic of this paper from the previous papers is that it is the first paper in the energy literature that examines whether or not the EKC prevails for the USA and investigates whether or not renewable energy consumption can decrease CO_2 emissions in the USA regarding a regime shift.

The rest of the paper is structured as follows: Literature review is given in the second section. The third section exhibits model and data. Estimation methodology and findings are reported in the fourth section. Finally, fifth section concludes the paper.

Literature review

Based on the empirical model that is employed, this paper presents the empirical literature estimating whether the EKC is valid for the USA and the empirical literature examining whether renewable energy consumption can decrease CO_2 emissions in the USA.

One can detect that there has been an extending empirical literature testing for the validity of the EKC in the energy economics literature (see, e.g., Shahbaz and Sinha 2018 for the empirical literature on the EKC, among others). One can also notice that some papers in the empirical literature examine the validity of the EKC for the USA. For instance, Flores et al. (2014), who use state-level data over the period 1929–1994 and employ quantile regression fixed effects model, find evidence in favor of the validity of the EKC. Baek (2015) utilizes data spanning the period 1960–2010 and performs the autoregressive distributed lag (ARDL) methodology. He yields that the EKC is not valid for the USA. Bilgili et al. (2016), who use data for the period 1977–2010 and employ the dynamic ordinary least squares (DOLS) and the fully modified ordinary least squares (FMOLS) estimators, find that CO_2 emissions are not related to income level and so the EKC is not valid in the USA. Dogan and Turkekul (2016) use data from 1960 to 2010 and employ the ARDL approach. They find out that the EKC is not valid for the USA. Atasoy (2017) examines the EKC for 50 states in the USA over the period 1960–2010 through the augmented mean group (AMG) and the common correlated effects mean group (CCEMG) estimators. While the AMG estimator indicates that the EKC is valid for 30 states, the CCEMG estimator implies that it is valid for 10 states. Apergis et al. (2017), who use data over the period 1960–2010 and carry out the common correlated effects (CCE) estimator, investigate whether the EKC dominates in the USA within a state-level framework. Their findings show that the EKC prevails in 10 states in the USA. Finally, Sarkodie and Strezov (2018b) explore that the EKC does not dominate for the USA by using data over the period 1971–2013 and employing the ARDL approach. As is seen, the papers in the energy literature do not exhibit clear-cut evidence about the EKC for the USA.

One can observe throughout the energy economics literature that most of the papers that focus on the influence of renewables consumption on CO_2 emissions examine this relationship using a panel data framework and report the parameter of renewables consumption for the entire panel (see, e.g., Apergis et al. 2010, Zoundi 2017, and Wang et al. 2018 for the empirical literature, among others). Besides, there are several papers which follow time series analysis for the USA or present the parameter of renewable energy consumption for individual countries including the USA. For instance, Menyah and Wolde-Rufael (2010) investigate the causal relationships between renewable energy consumption and CO_2 emissions using data over the period 1960–2007. They find one-way causal relationship running from CO_2 emissions to renewable energy consumption. Ozbugday and Erbas (2015), using data for the period 1971–2009 and performing the CCE estimator, examine the effect of renewables consumption on CO_2 emissions and find that CO_2 emissions are not associated with renewables consumption. Bilgili et al. (2016) analyze the effect of renewable energy consumption on CO_2 emissions over the period 1977–2010 through the FMOLS and DOLS estimators. They yield mixed findings. That is, while the FMOLS estimator implies CO_2 emissions are positively associated with renewables consumption, the DOLS estimator implies CO_2 emissions are negatively associated with renewables consumption. As is seen, the papers focusing on the impact of renewable energy consumption on CO_2 emissions yield mixed findings.

Model and data set

To examine the validity of the EKC and the effect of renewables consumption on CO_2 emissions in the US, following Bilgili et al. (2016), the paper uses the empirical model below:

$$\ln CO_{2t} = \delta_0 + \delta_1 \ln IP_t + \delta_2 (\ln IP_t)^2 + \delta_3 \ln REC_t + \varepsilon_t \qquad (1)$$

where ln, CO_2, IP, $(IP)^2$, REC, and ε stand for natural logarithm, CO_2 emissions (million metric tons), industrial production index (2012 = 100), the square of industrial production index, renewables consumption (quadrillion Btu), and the error term, respectively. This paper utilizes monthly data for the period 2000:M01–2018:M07. Data for CO_2 emissions and renewable energy consumption are sourced from Energy Information Administration (2018, hereafter EIA), while data for industrial production index are obtained from Federal Reserve Bank of St. Louis (2018). All variables are seasonally adjusted through Census X-13 method.

With the estimation of the empirical model in Eq. (1), researchers may explore several results represented as follows:

(i) If $\delta_1 = \delta_2 = 0$, no relationship between income and CO_2 emissions
(ii) If $\delta_1 > 0$ and $\delta_2 = 0$, a monotonically increasing relationship between income and CO_2 emissions
(iii) If $\delta_1 < 0$ and $\delta_2 = 0$, a monotonically decreasing relationship between income and CO_2 emissions
(iv) If $\delta_1 < 0$ and $\delta_2 > 0$, U-shaped relationship between income and CO_2 emissions
(v) If $\delta_1 > 0$ and $\delta_2 < 0$, the EKC holds

Finally, δ_3 is expected to be negative and statistically significant as renewable energy sources are cleaner and lead to fewer CO_2 emissions compared with fossil energy sources.

Descriptive statistics along with correlation matrix for the variables in the empirical model are reported in Table 1. Accordingly, all descriptive statistics of $(\ln IP)^2$ are higher than those of other variables in the model. Besides, the output of the correlation matrix indicates that (i) $\ln CO_2$ is negatively correlated with lnIP, $(\ln IP)^2$, and lnREC; (ii) there is a very high and positive correlation between lnIP and $(\ln IP)^2$ as the latter is the square of the former; and (iii) lnREC appears to be positively correlated with lnIP and $(\ln IP)^2$.

Table 1 Descriptive statistics and correlation matrix for the variables

	$\ln CO_2$	lnIP	$(\ln IP)^2$	lnREC
Descriptive statistics				
Mean	6.144	4.592	21.091	2.058
Median	6.145	4.604	21.197	2.080
Maximum	6.252	4.680	21.907	2.466
Minimum	5.970	4.467	19.951	1.625
Std. deviation	0.058	0.051	0.470	0.242
Observations	223	223	223	223
Correlation matrix				
$\ln CO_2$	1.000			
lnIP	−0.241	1.000		
$(\ln IP)^2$	−0.242	0.999	1.000	
lnREC	−0.868	0.531	0.532	1.000

Figure 1 plots of the variables in the empirical model. Accordingly, industrial production in the USA sharply decreased during the global financial crisis in 2007–2008 and then began to recover over time. Besides, $\ln CO_2$ has a tendency to decrease over time in the USA while lnREC has a tendency to increase over time. Hence, time plots of the variables indicate that (i) the variables in the empirical model may contain a unit root and so may not be stationary and (ii) the increase in lnREC may have a role in decreases in $\ln CO_2$.

Descriptive statistics, correlation matrix, and graphical observations provide some preliminary inspection for researchers. However, to be able to obtain efficient output about the relationships among the variables in the empirical model, researchers need to obtain some econometric methodologies, namely unit root and cointegration tests. Hence, the next section presents the estimation methodology and the empirical findings in the paper.

Estimation methodology and findings

Unit root tests

The first step is to examine the order of integration of variables in an empirical model to avert possible spurious regression problem in a time series analysis. The unit root tests developed by Dickey and Fuller (1981, hereafter ADF) and Phillips and Perron (1988, hereafter PP) are largely performed in econometric analyses. In his seminal paper, Perron (1989) remarks that these tests do not take structural breaks, stemming from wars, natural disasters, economic crises, radical changes in economic policies, etc., in series into account. To be able to obtain efficient output about stationarity of the variables in the empirical model, this paper utilizes Zivot and Andrews (1992, hereafter ZA) unit

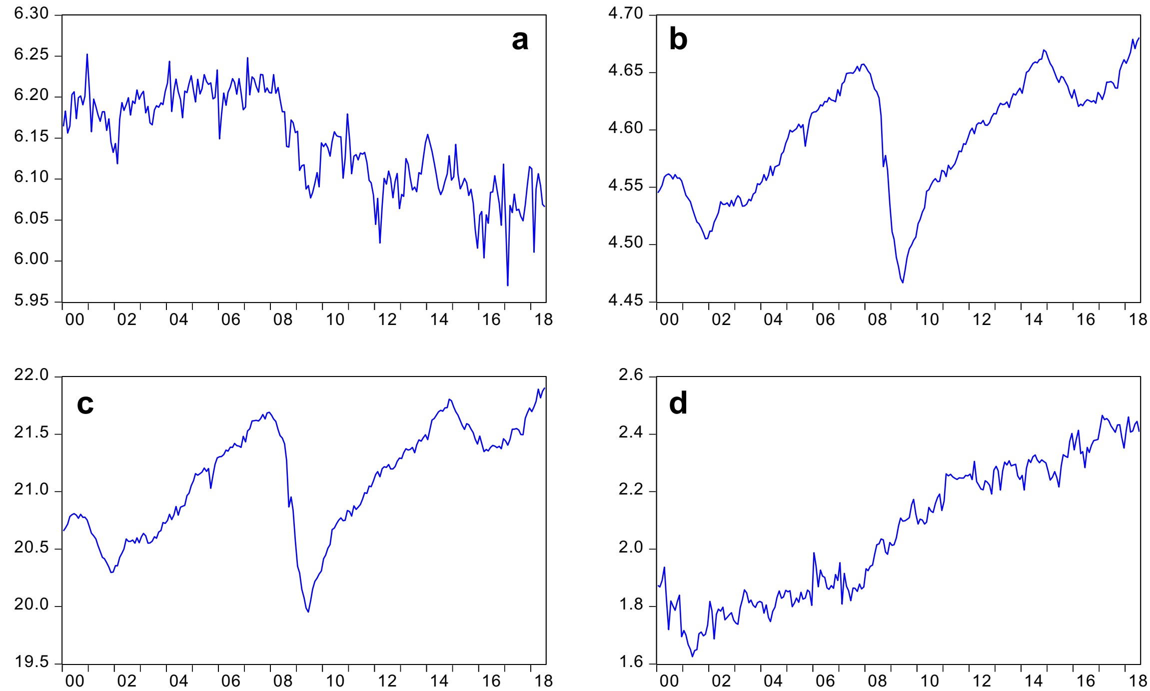


Fig. 1 Time plots of the variables. **a** $\ln CO_2$. **b** lnIP. **c** $(\ln IP)^2$. **d** lnREC

root test with one structural break along with ADF and PP unit root tests.

The results of these unit root tests are demonstrated in Table 2. Accordingly, the null hypothesis of a unit root is rejected for all variables at first differences. In other words, unit root tests imply that all variables become stationary at their first difference forms. Besides, the global financial crisis can account for all the breaks detected for the variables.

Cointegration test with a regime shift

In a time series analysis, researchers need to determine whether series are cointegrated prior to estimating long-run parameters if series are $I(d)$, where $d \neq 0$, because they may face biased and inefficient output about t, F, and/or Wald statistics. Seminal papers of Engle and Granger (1987), Johansen (1988), and Johansen and Juselius (1990) have been greatly employed in the literature to investigate the cointegration relationship between variables. These cointegration methods suppose that long-run parameters do not change over time. Put differently, they assume that there are no regime shifts (structural breaks in intercept and slope coefficient(s)) throughout the observed period. That is, they estimate a coefficient for each independent variable in the empirical model. However, long-run parameters may change over time due to some considerable developments in an economy, such as wars, natural disasters, economic crises, and radical changes in economic policies. One may therefore obtain inefficient output about the cointegration relationship if he/she employs cointegration methods without structural breaks. Hence, this paper employs the cointegration test with a structural break propounded by Gregory and Hansen (1996, hereafter GH). GH produce three models, namely structural break in intercept (level shift), structural break in intercept with trend, and regime shift. These models are exhibited as follows:

Level shift model:

$$y_t = \beta_0 + \beta_0' \mathrm{D} + \beta_1 x_t + \varepsilon_t \tag{2}$$

Level shift with trend model:

$$y_t = \beta_0 + \beta_0' \mathrm{D} + \theta \mathrm{t} + \beta_1 x_t + \varepsilon_t \tag{3}$$

Regime shift model:

$$y_t = \beta_0 + \beta_0' D + \beta_1 x_t + \beta_1' D x_t + \varepsilon_t \tag{4}$$

where y, x, β_0, β_0', D, β_1, β_1', and ε stand for the dependent variable, the independent variable, intercept before the regime shift, the change in intercept after the regime shift, dummy variable, the coefficient of the independent variable before the regime shift, the change in the coefficient of the independent variable after the regime shift, and the error term, respectively.

Table 2 Results for ADF, PP, and ZA unit root tests

Variable	Level			1st difference		
	ADF	PP	ZA[a]	ADF	PP	ZA
$\ln CO_2$	− 1.610	− 2.568	− 4.614 (August 2008)	− 16.831[b]	− 26.487[b]	− 16.989[b]
lnIP	− 2.175	− 1.483	− 4.549 (August 2008)	− 3.844[b]	− 13.317[b]	− 5.889[b]
$(\ln IP)^2$	− 2.154	− 1.471	− 4.549 (August 2008)	− 3.847[b]	− 13.368[b]	− 5.893[b]
lnREC	− 0.134	− 0.710	− 4.426 (March 2009)	− 15.179[b]	− 18.851[b]	− 11.611[b]
5% critical value	− 2.874	− 2.874	− 4.93	− 2.874	− 2.874	− 4.93

[a] Break dates are illustrated in parentheses

[b] Illustrates statistical significance and indicates the rejection of the null hypothesis of a unit root

In order to model the break, GH define the dummy variable as

$$D = \begin{Bmatrix} 0\,\text{if}\,t \le [n\tau] \\ 1\,\text{if}\,t > [n\tau] \end{Bmatrix}$$

with the unknown parameter $\tau \in (0,1)$ indicating the relative timing of the change point. Finally, the bracket signifies the integer part.

GH use the ADF test produced by Engle and Granger (1987) along with Z_α and Z_t tests suggested by Phillips (1987) to investigate the existence of the cointegration relationship in the empirical model. If test statistics are higher than the critical values, then the null hypothesis of no cointegration can be rejected. To be able to follow the possible changes in slope coefficients, this paper considers the regime shift model. Therefore, Eq. (1) can be rewritten under the regime shift as the following:

$$\ln CO_{2t} = \delta_0 + \delta_1 \ln IP_t + \delta_2 (\ln IP_t)^2 + \delta_3 \ln REC_t + \delta_0' D + \delta_1' D \ln IP_t + \delta_2' D (\ln IP_t)^2 + \delta_3' D \ln REC_t + \varepsilon_t \quad (5)$$

where D is the dummy variable. According to Eq. (5), δ_0 is the intercept before the regime shift while δ_0' is the change in the intercept after the regime shift. Besides, δ_1, δ_2, δ_3 respectively stand for the coefficients of lnIP, $(\ln IP)^2$, and lnREC before the regime shift while δ_1', δ_2', and δ_3' respectively denote the changes in δ_1, δ_2, and δ_3 after the regime shift. Finally, ε is the error term.

Table 3 reports the results of the GH cointegration test for the regime shift model. As is seen, the null hypothesis of no cointegration can be rejected as per all test statistics. In other words, test statistics signify that there occurs a cointegration relationship among the variables in the empirical model described in Eq. (1) and that long-term coefficients can be estimated. Besides, while ADF test indicates the regime shift occurs in September 2011, Z_t and Z_α tests imply it occurs in October 2011. In this paper, the structural break is assumed to occur in October 2011 as two out of three test statistics point out this date. Therefore, the dummy variable, namely D, takes the value of 0 from January 2000 to October 2011 while it takes the value of 1 from November 2011 to July 2018, which is the end of the used data in the paper. Hence, the paper observes the coefficients of the independent variables between January 2000 and October 2011 along with the changes in the coefficients of the independent variables after the regime shift in October 2011.

After exploring the existence of the cointegration relationship among the variables in the empirical model, the next stage is to estimate the empirical model with a regime shift. To estimate Eq. (5), the DOLS estimator suggested by Stock and Watson (1993) can be performed. This method, which is widely employed in the economics literature to estimate long-run parameters, can correct possible endogeneity and serial correlation problems (Esteve and Requeana 2006).

The results for the DOLS estimator are depicted in Table 4. Accordingly, before the regime shift in October 2011, lnIP,

Table 3 GH cointegration test

Statistics	Regime shift model
ADF	− 6.455[a] (September 2011)
Phillips (Z_t)	− 11.22[b] (October 2011)
Phillips (Z_α)	− 161.6[b] (October 2011)

Values in parentheses show break dates

[a] Illustrates 5% statistical significance

[b] Illustrates 1% statistical significance

Table 4 DOLS estimator

Variable	Coefficient	Std. error	t statistic
Intercept	− 49.023[a]	18.038	− 2.717
lnIP	23.793[a]	7.887	3.016
$(\ln IP)^2$	− 2.548[a]	0.862	− 2.995
lnREC	− 0.155[a]	0.013	− 11.980
D	− 60.545	110.905	− 0.546
D*lnIP	25.635	47.864	0.535
$D*(\ln IP)^2$	− 2.693	5.163	− 0.521
D*lnREC	− 0.196[a]	0.043	− 4.585
$R^2 = 0.90$, adj. $R^2 = 0.89$			

[a] Illustrates 1% statistical significance

$(\ln IP)^2$, and lnREC have the estimations of 23.793, − 2.548, and − 0.155, respectively. As is seen, all these coefficients are statistically significant at 1% level. These findings imply that not only the EKC prevails but also renewable energy consumption has negative impacts on CO_2 emissions in the USA before the regime shift. After the regime shift, D*lnIP, $D*(\ln IP)^2$, and D*lnREC have the estimations of 25.635, − 2.693, and − 0.196, respectively. Among these coefficients, the coefficient of D*lnREC appears to be significant, whereas other coefficients are insignificant. Therefore, the findings indicate that the coefficients of lnIP and $(\ln IP)^2$ did not change, while the coefficient of lnREC took the final value of − 0.351 (= − 0.155–0.196) after the regime shift. Therefore, the paper explores that the EKC dominates after the regime shift, too, and the negative effect of renewables consumption on CO_2 emissions increased after the regime shift.

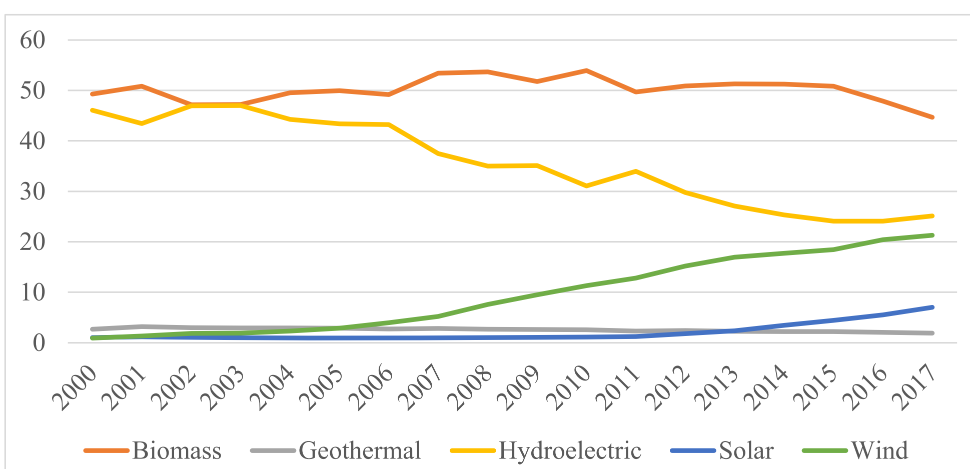
The findings of this paper for the EKC hypothesis concur with those of Flores et al. (2014) and contradict with those of Baek (2015), Bilgili et al. (2016), and Dogan and Turkekul (2016). Additionally, the findings of the paper in terms of the effect of renewable energy consumption on CO_2 emissions conflict with those of Menyah and Wolde-Rufael (2010) and Ozbugday and Erbas (2015).

Conclusion

This paper has examined the impact of renewables consumption on CO_2 emissions in the USA within the scope of the EKC hypothesis using monthly data from January 2000 to July 2018. After performing unit root tests, the paper performed the GH cointegration test with a regime shift to detect whether or not there existed a cointegration relationship among the series in the empirical model and then employed the DOLS estimator to estimate the long-run parameters. The findings signified that the EKC prevailed in the USA throughout the observed period and the regime shift did not affect the coefficients associated with the EKC hypothesis. Put differently, the findings implied that the USA, as a developed country, experienced the EKC. The findings also indicated that renewable energy consumption had statistically significant and negative impacts on renewable energy consumption in the USA, while this influence was higher after the regime shift.

The findings towards the effect of renewable energy consumption on CO_2 emissions have important implications for the US economy. Accordingly, while the portion of renewable energy consumption in total energy consumption was 9.3% in October 2011 in the USA (the date when the regime shift occurred), it reached 12.93% in May 2018 due to investment tax credits, production tax credits, and other state energy incentives (EIA 2018). Therefore, this ratio has a tendency to increase after the regime shift in the USA. The finding for renewable energy consumption therefore implies that when the share of renewable energy consumption in total energy

Fig. 2 The shares of renewable energy sources in total renewable energy consumption (%). Source: EIA (2018)



Source: EIA (2018)

consumption increases, the influence of renewable energy consumption on CO_2 emissions augments. On the other hand, one can notice from Fig. 2 that the energy mix of the USA considerably changed with regard to renewable energy sources from 2000 to 2017. Accordingly, the US economy has replaced hydroelectric with solar and wind sources during 2000s. From 2004 to 2017, the shares of wind and solar energy consumption in total energy consumption increased from 2.33% to 21.27% and from 0.96% to 7.02%, respectively. Awareness for solar and wind energy in the USA and the declines in costs of wind and solar energy appear to have critical roles in these remarkable increases. The US governments have encouraged production of wind and solar energy through production tax credit and investment tax credit over the last decades. As a result of these supports along with technological developments and investments in wind and solar energy industries, the costs of wind and solar energy technologies have notably decreased in the USA over the last years. For example, average capital costs for wind energy projects declined by 65% during the period 1980–2014 and some studies explored that this decrease would continue in the future (Lantz et al. 2012). Additionally, the cost of solar photovoltaic cell per watt decreased from 76.67 USD to 0.74 USD over the period 1977–2013 (Economist 2013).

Based on the empirical findings and the developments in the renewable energy sector of the USA, this paper argues that the US governments should go on supporting renewable energy technologies. By doing so, they not only can decrease the environmental problems arising from the use of fossil sources but also can make a contribution to economic growth as energy is a crucial input for economic activities.

References

Apergis N, Payne JE, Menyah K, Wolde-Rufael Y (2010) On the causal dynamics between emissions, nuclear energy, renewable energy, and economic growth. Ecol Econ 69:2255–2260

Apergis N, Christou C, Gupta R (2017) Are there environmental Kuznets curves for US state-level CO_2 emissions? Renew Sust Energ Rev 69: 551–558

Atasoy BS (2017) Testing the environmental Kuznets curve hypothesis across the US: evidence from panel mean group estimators. Renew Sust Energ Rev 77:731–747

Baek J (2015) Environmental Kuznets curve for CO_2 emissions: the case of Arctic countries. Energy Econ 50:13–17

Bilgili F, Kocak E, Bulut U (2016) The dynamic impact of renewable energy consumption on CO_2 emissions: a revisited environmental Kuznets curve approach. Renew Sust Energ Rev 54:838–845

Copeland BR, Taylor MS (2003) Trade and the environment: theory and evidence. Princeton University Press, New Jersey

Dickey DA, Fuller WA (1981) Likelihood ratio statistics for autoregressive time series with a unit root. Econometrica 49:1057–1072

Dogan E, Turkekul B (2016) CO_2 emissions, real output, energy consumption, trade, urbanization and financial development: testing the EKC hypothesis for the USA. Environ Sci Pollut Res 23: 1203–1213

Du G, Liu S, Lei N, Huang Y (2018) A test of environmental Kuznets curve for haze pollution in China: evidence from the penal data of 27 capital cities. J Clean Prod 205:821–827

Economist (2013) Sunny uplands. http://www.economist.com/news/21566414-alternative-energy-will-no-longer-be-alternative-sunny-uplands. Accessed 24 Nov 2018

Energy Information Administration (2018) October 2018 monthly energy review. https://www.eia.gov/totalenergy/data/monthly/. Accessed 11 Nov 2018

Engle RF, Granger CW (1987) Co-integration and error correction: representation, estimation, and testing. Econometrica 55:251–276

Esteve V, Requeana FA (2006) Cointegration analysis of car advertising and sales data in the presence of structural change. Int J Econ Bus 13:111–128

Federal Reserve Bank of St. Louis (2018) Industrial production index. https://fred.stlouisfed.org/series/INDPRO. Accessed 11 Nov 2018

Flores CA, Flores-Lagunes A, Kapetanakis D (2014) Lessons from quantile panel estimation of the environmental Kuznets curve. Econ Rev 33:815–853

Gregory AW, Hansen BE (1996) Residual-based tests for cointegration in models with regime shifts. J Econ 70:99–126

Grossman GM, Krueger AB (1995) Economic growth and the environment. Q J Econ 110:353–377

Guris B (2016) CO_2 emissions and energy consumption in the US: results from a threshold error correction model. Energ Source Part B 11: 314–320

Johansen S (1988) Statistical analysis of cointegration vectors. J Econ Dyn Control 12:231–254

Johansen S, Juselius K (1990) Maximum likelihood estimation and inference on cointegration-with applications to the demand for money. Oxford B Econ Stat 52:169–210

Kuznets S (1955) Economic growth and income inequality. Am Econ Rev 45:1–28

Lantz E, Hand M, Wiser R (2012) The past and future cost of wind energy. https://www.nrel.gov/docs/fy12osti/54526.pdf. Accessed 20 Aug 2018

Menyah K, Wolde-Rufael Y (2010) CO_2 emissions, nuclear energy, renewable energy and economic growth in the US. Energy Policy 38: 2911–2915

National Aeronautics and Space Administration (2018) How climate is changing? https://climate.nasa.gov/effects/. Accessed 1 July 2018

Ozbugday FC, Erbas BC (2015) How effective are energy efficiency and renewable energy in curbing CO_2 emissions in the long run? A heterogeneous panel data analysis. Energy 82:734–745

Pata UK (2018) The influence of coal and noncarbohydrate energy consumption on CO_2 emissions: revisiting the environmental Kuznets curve hypothesis for Turkey. Energy 160:1115–1123

Perron P (1989) The great crash, the oil price shock, and the unit root hypothesis. Econometrica 57:1361–1401

Phillips PC (1987) Time series regression with a unit root. Econometrica 55:277–301

Phillips PCB, Perron P (1988) Testing for a unit root in time series regression. Biometrika 75:335–346

Sarkodie SA, Strezov V (2018a) A review on environmental Kuznets curve hypothesis using bibliometric and meta-analysis. Sci Total Environ 649:128–145

Sarkodie SA, Strezov V (2018b) Empirical study of the environmental Kuznets curve and environmental sustainability curve hypothesis for Australia, China, Ghana and USA. J Clean Prod 201:98–110

Shahbaz M, Sinha A (2018) Environmental Kuznets curve for CO_2 emission: a literature survey. J Econ Stud 46:106–168

Stern DI (2004) The rise and fall of the environmental Kuznets curve. World Dev 32:1419–1439

Stock JH, Watson M (1993) A simple estimator of cointegrating vectors in higher order integrated systems. Econometrica 61:783–820

Su ECY, Chen YT (2018) Policy or income to affect the generation of medical wastes: an application of environmental Kuznets curve by using Taiwan as an example. J Clean Prod 188:489–496

Sun S, Fang C (2018) Water use trend analysis: a non-parametric method for the environmental Kuznets curve detection. J Clean Prod 172: 497–507

Ulucak R, Bilgili F (2018) A reinvestigation of EKC model by ecological footprint measurement for high, middle and low income countries. J Clean Prod 188:144–157

Wang S, Li G, Fang C (2018) Urbanization, economic growth, energy consumption, and CO2 emissions: empirical evidence from countries with different income levels. Renew Sust Energ Rev 81:2144–2159

Yavuz NC (2014) CO_2 emission, energy consumption, and economic growth for Turkey: evidence from a cointegration test with a structural break. Energ Source Part B 9:229–235

Zivot E, Andrews DWK (1992) Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis. J Bus Econ Stat 10: 251–270

Zoundi Z (2017) CO_2 emissions, renewable energy and the environmental Kuznets curve, a panel cointegration approach. Renew Sust Energ Rev 72:1067–1075