Article

The Environmental Consequences of Growth: Empirical Evidence from the Republic of Kazakhstan

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Abstract: The main objective of this paper is to examine the effect growth has on CO_2 emissions in Kazakhstan, controlling for energy consumption, in the autoregressive distributed lag (ARDL) cointegration framework. We find that the environmental Kuznets curve (EKC) hypothesis seems to hold for Kazakhstan; this effect at a low level of income increases CO_2 but at a high level decreases it. We also find that energy consumption increases CO_2 emissions.

Keywords: ARDL; CO_2 emissions; environment; growth; Kazakhstan

JEL Classification: C2; Q4; Q5

1. Introduction

Examining the effect growth has on a country's environment has long been a popular subject of empirical research. Based on the modeling approach, the existing research can be classified into two groups. The first group generally includes early papers that have adopted cross-sectional or panel data to determine the effect of growth on the environment (e.g., Shafik and Bandyopadhyay 1992; Panayotou 1993; Holtz-Eakin and Selden 1995; Moomaw and Unruh 1997; Roberts and Grimes 1997; List and Gallet 1999; Heil and Selden 2001; Harbaugh et al. 2002; Perman and Stern 2003; Martinez-Zarzoso and Bengochea-Morancho 2004; Liu 2005; Frankel and Rose 2005). Holtz-Eakin and Selden (1995), for example, use the panel data of 130 countries for the years 1951 to 1986 when examining the growth-environment nexus; they find that growth improves the environment.

The second group claims that, since any beneficial growth effect on the environment in one country may be outweighed by an adverse growth impact in other countries, or vice versa, the results of the first group are likely to suffer from the so-called aggregation bias of data. To avoid the shortcoming, therefore, this group employs individual country level data and time series methods in tackling the issue (e.g., Soytas and Sari 2009; Jalil and Mahmud 2009; Halicioglu 2009; Iwata et al. 2010; Pao et al. 2011; Baek and Kim 2011, 2013; Akpan and Akpan 2012; Shahbaz et al. 2013; Baek 2015; Tutulmaz 2015; Ibrahiem 2016; Yang et al. 2017). Pao et al. (2011), for example, analyze the growth-environment nexus using a time series dataset of Russia; they find that growth indeed decreases pollution. Baek (2015) reports, in passing, a beneficial effect of growth on the environment only in some Arctic countries, after controlling for energy consumption.

Our primary interest in this paper is to contribute to the literature on the second group by assessing the effect growth has on Kazakhstan's environment using an autoregressive distributed lag (ARDL) model of Pesaran et al. (2001), which is one of the recently most widely used time series models. Since independence in 1991, Kazakhstan has been one of the fast-growing transitional economies in the world. For the years 2000 to 2014, for example, Kazakhstan's economy has recorded

an average economic growth rate of approximately 7.7%. As a result, Kazakhstan has been the largest economy in Central Asia. Kazakhstan's fast-growing economy, however, has resulted in environmental degradation because of the rapid growth of energy intensive and heavy industries. For the years 2000 to 2014, for example, carbon dioxide (CO_2) emissions, a major greenhouse gas, have increased by more than 80%. As a result, Kazakhstan currently is among the world's highest CO_2 emitter per unit of GDP. Up until now, however, attention of the literature on the second group has typically been on the growth-environment nexus for individual countries in Europe and East Asia with few studies considering the subject in Central Asian countries like Kazakhstan. Given the recent adoption of green growth policy in Kazakhstan that targets achieving growth while improving environmental quality, it is indeed timely to pursue this line of research. We hope that the current research contributes to a better understanding of the growth-environment nexus in Kazakhstan. It is worth mentioning that Bacon et al. (2007), Tamazian and Rao (2010), and Mitic et al. (2017) examine the effect of growth on CO_2 emissions using panel data set that includes Kazakhstan. Using panel data of 70 countries, for example, Bacon et al. (2007) conclude that growth has little effect on CO_2 emissions. Mitic et al. (2017), on the other hand, report the significant impact growth has on CO_2 emissions after analyzing 17 transitional economies. To the best of our knowledge, Mikayilov et al. (2017) is perhaps the only study that is tackling the issue in the individual country of Central Asia. Using a time series data for the 1990–2014 period, they find that growth does not affect CO_2 emissions from the transport sector in Azerbaijan.

The rest of the paper is presented as follows. In Section 2 we present our empirical model along with an application of an ARDL model. In Sections 3 and 4, we discuss our data and the empirical results. Section 5 provides our summary and conclusion.

2. The Models and Methods

When examining the impact growth has on Kazakhstan's environment, following the current literature (i.e., Baek 2015), we estimate the standard growth–environment model, with CO_2 emissions (c_t) proxied for a measure of environmental damage as the dependent variable and GDP per capita (y_t) proxied for growth and energy consumption (ec_t).

$$c_t = \beta_0 + \beta_1 y_t + \beta_2 y_t^2 + \beta_3 ec_t + u_t, \tag{1}$$

in which all variables are in the logarithmic form. In the current paper, we are particularly interested in β_1 and β_2. If β_1 is positive and β_2 is negative, the quadratic has a parabolic shape, meaning that CO_2 emissions are decreasing (increasing) with growth after (before) turnaround value of income, thereby confirming the Environmental Kuznets curve (EKC).[1] In addition, a growing body of literature provides empirical evidence that energy consumption and the level of economic development are among the most important causes of increased greenhouse gas (GHG) emissions like CO_2 emissions (e.g., Pao and Tsai 2011; Kivyiro and Arminen 2014; Baek 2016). By including energy consumption explicitly in the model, therefore, we are able to control for its effect on CO_2 emissions. We hypothesize that $\beta_3 > 0$; that is, an increase in total energy consumption results in more CO_2 emissions.

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[1]   It is worth discussing main theoretical explanations for the EKC hypothesis. One of the main theories explaining the ECK is that the shape of the EKC reflects changes in the demand for environment quality with growth, known as the *income effect* (Lopez 1994). That is, income growth driven by accumulation of production factors increases firms' demand for pollution inputs. At the same time, demand for environmental quality rises with growth as the willingness to pay for a clean environment increases. An alternatively widely cited explanation for the EKC is the *threshold effect* (John and Pecchenino 1994; Selden and Song 1995). That is, since pollution could be unregulated entirely at the early stage of development, pollution at first tends to rise with growth. After some threshold has been reached and regulation implemented, however, pollution tends to decline with growth. The *increasing returns to abatement effect* argues that as the scale of abatement increases, its efficiency tends to increase, which makes abatement more profitable and hence reduces pollution levels as more abatement is undertaken (Andreoni and Levinson 2001). Finally, the most recent explanation for the EKC is that growth tends to shift economic production system from high polluting industries to low polluting industries, known as the *structural change effect* (Marsiglio et al. 2016).

In order to conduct the ARDL approach, we first need to convert Equation (1) into an error-correction format.

$$\Delta c_t = \beta_0' + \sum_{k=1}^{p} \beta_{k1}' \Delta c_{t-k} + \sum_{k=0}^{p} \beta_{k2}' \Delta y_{t-k} + \sum_{k=0}^{p} \beta_{k3}' \Delta y_{t-k}^2 + \sum_{k=0}^{p} \beta_{k4}' \Delta ec_{t-k} ,$$
$$\sigma_0 c_{t-1} + \sigma_1 y_{t-1} + \sigma_2 y_{t-1}^2 + \sigma_3 ec_{t-1} + \xi_t \tag{2}$$

Equation (2) follows Pesaran et al. (2001) who include the linear combination of lagged level variables rather than lagged error term from Equation (1). In estimating Equation (2), a cointegration relationship among the four variables must first be established. For this, we need to conduct the *F*-test about the lagged four variables in Equation (2). However, the asymptotic distribution of this *F*-statistic is not standard under the null hypothesis, so that Pesaran et al. (2001) provide two new sets of critical values. Since the critical values account for integrating properties of all regressors, there is no need for pre unit-root testing under this approach, and regressors could be *I*(0) or *I*(1), which is the main advantage of this method.[2] Once cointegration is found, the short-run dynamics are captured in our coefficient series of β_{k1}–β_{k4}. The long-run relationship among the variables is identified by the estimates of σ_1–σ_3, which are normalized to σ_0.

3. Data

We use the data from 1991 through 2014 on CO_2 emissions, income, and energy consumption in Kazakhstan (Figure 1). Notably, the dataset can originally be traced back to 1985. Since Kazakhstan became independent in 1991, however, the six years of data between 1985 and 1991 are from the period of the Soviet Union, and the rest is from the period of independence. Given that the former period is characterized by the centralized, command-based economic system, while the latter period is characterized by the market-based economy, methodologies for calculating/constructing the data (i.e., GDP per capita) are likely to be very different from each other. For example, since GDP per capita has been calculated using different methodologies in the sub-periods of the whole period, their data generating processes are different from each other, and this is likely to lead to the so-called data measurement errors, thereby invalidating estimation results. For this reason, therefore, we decide to exclude the first six years from the empirical analysis.[3]

CO_2 emissions (measure in metric tons) are obtained from Statistical Yearbook published by Agency on Statistics of the Republic of Kazakhstan. To ensure comparability with income per capita in Equation (2), the CO_2 emissions per capita for Kazakhstan are calculated using its total population. The GDP per capita (measured in constant 2010 Kazakhstani Tenge) is taken from the World Development Indicator (WDI). The energy consumption per capita (measured in kg of oil equivalent per capita) is also collected from the WDI. Descriptive statistics of data are summarized in Table 1.

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[2] Another advantage of this method is that it has been proven to have superior small sample properties, which makes it a good choice for our sample of less than 30 annual observations compared to other cointegration methods (i.e., Johansen method).

[3] The authors thank an anonymous referee for raising this issue discussed here.

[4] It is worth mentioning that since Kazakhstan's independence from the USSR in 1991, the collapse of demand for Kazakhstan's heavy industry products has resulted in a sharp contraction of the economy during the 1990s. Since the beginning of economic reforms and opening up to the outside world in the early 2000s, however, Kazakhstan's economy has grown sharply (except for the global financial crisis in 2009). As illustrated in Figure 1, therefore, CO_2 emissions per capita, energy consumption per capita, and income per capita have persistently declined up to 2000 and have increased since then.

Table 1. Descriptive Statistics.

	Mean	**Standard Deviation**	**Min**	**Max**
CO_2 emissions	10.249	2.099	6.756	14.435
Income	452,622.2	165,430.3	259,194.0	738,066.4
Energy consumption	3626.116	855.193	2324.548	4796.144

Notes: CO_2 emissions are measured in metric tons of CO_2 emitted per capita. Income per capita is measured in 2010 Kazakhstani Tenge. Energy consumption is measured in kg of oil equivalent per capita.
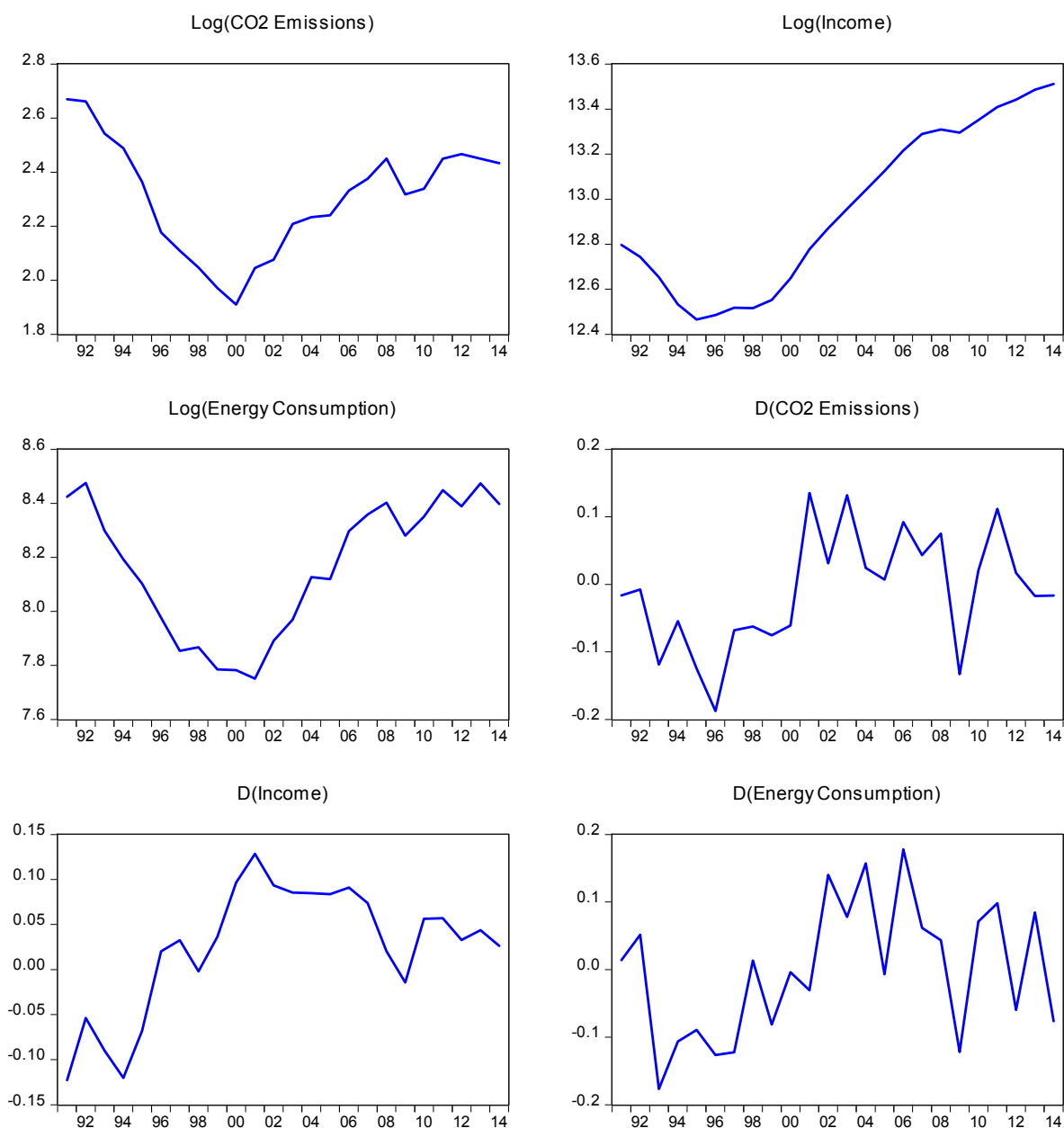

Figure 1. Plots of CO_2 Emissions, Income, and Energy Consumption[4]. *Notes*: D represent the first differences of the variables.

4. The Results

In this section, we estimate the ARDL model outlined by Equation (2) using annual data for the period 1991–2014. It should be noted that, although the ARDL can be employed regardless of

whether regressors are *I*(0) and/or *I*(1) or mixture of them, it cannot be applicable to *I*(2) or higher series. For this, we apply the ADF test developed by Dickey and Fuller (1979) and Perron-Vogelsang (PV) test developed by Perron and Vogelsang (1992) to the levels, as well as first-differenced variables, and report the results in Table 2. Note that, since the PV test is known as one of the most popular procedures to test for a unit root in an unknown structural break, this is employed to validate the results of the ADF test, as well as to identify a potential breakpoint in the series. From Table 2, it appears that CO_2 emissions and energy consumption are *I*(1), and the remaining variables are trend-stationary.[5] The PV test corroborates the findings of the ADF test, implying that an identified structural break does not seem to affect the unit root properties of the variables in the model. These results indeed justify adopting ARDL approach to estimate error-correction model in Equation (2).

With strong evidence that each of the four series is either *I*(0) or *I*(1), we impose two as the maximum lag length on each of the first-differenced variables, and using the Akaike Information Criterion (AIC), the autogressive distributed lag (ARDL) (1, 0, 0, 0) equation is identified as the optimal specification.[6] We then apply the *F*-test to determine if a long-run or cointegration relationship among the variables occurs. The obtained *F*-statistic for the null hypothesis of no cointegration (H_0: $\sigma_0 = \sigma_1 = \sigma_2 = \sigma_3 = 0$) is 8.45, which is well above the 5% upper critical value of 4.35[7], and hence a statistical rejection of the null. It is important to note that the asymptotic critical values provided by Pesaran et al. (2001) are not likely to be valid for small sample sizes like our sample of less than 30 annual observations. To address this concern, we also use the small sample critical values taken from Narayan (2005). The Narayan 5% critical value for 30 observation is 5.02 (taken from Case III: unrestricted intercept and no trend), which is higher than that of Pesaran et al. (2001), but we still can reject the null and support cointegration.[8] This implies that our *F*-tests seem robust, even in small samples. Therefore, we conclude that there is strong evidence of cointegration among the four variables, so we can pursue the ARDL on estimating the short- and long-run estimates in Equation (2).[9]

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[5]   It should be pointed out that when there is no trend in ADF and PV tests, the unit root hypothesis for the two income variables cannot be rejected. With the time trend included, however, we can strongly reject the null for both variables. Thus, the best characterization of the two income variables seems to be as a trend-stationary process; that is, a process that is stationary about its time trend.

[6]   The Schwarz Bayesian Criterion (SBC) generally used for low small size of studies like this paper also identifies ARDL (1, 0, 0, 0) as the optimal model.

[7]   Among five cases of testing for cointegration, case III (unrestricted intercept and no trend) is used for the analysis. The associated 5% and 10% critical value bounds are (3.23, 4.35) and (2.72, 3.77), respectively, which are taken from Table CI (iii) Case III: unrestricted intercept and no trend on p. 300 of Pesaran et al. (2001).

[8]   In order to capture the effects of technological progress or enhanced environmental awareness on CO_2 emissions, a time trend is included in estimating Equation (2). However, our findings are more conclusive when the *F*-test is applied to Equation (2) without a time trend. Further, a time trend is not statistically significant even at the 10% level. Hence, a time trend is excluded from the final model.

[9]   As a cross-check, we also perform the bounds *t*-test of H_0: $\sigma_0 = 0$ against H_1: $\sigma_0 < 0$. If the null is rejected using the upper critical value bounds tabulated by Pesaran et al. (2001, pp. 303–4), this would support cointegration relationship among the variables. In our case, the *t*-statistic on c_{t-1} in -4.58. When we look at Table CII(iii) (Case III: unrestricted intercept and no trend) on p. 303 of Pesaran et al. (2001), the associated 5% and 10% critical value bounds for the *t*-statistic are $(-2.86, -3.78)$ and $(-2.57, -3.46)$, respectively. Even at the 5% level, therefore, this result confirms our conclusion that there is a long-run relationship among the four variables.

Table 2. Results of Unit Root Tests.

Variable	ADF Test	
	Level	First Difference
CO_2 emissions	−1.855 (0)	−3.574 ** (0)
Income	−3.730 ** (0)	
(Income)^2	−3.777 ** (0)	
Energy consumption	−1.913 (0)	−4.252 ** (1)

Variable	Perron-Vogelsang Test			
	Level	Time Break	First Difference	Time Break
CO_2 emissions	−2.929 (1)	2003	−6.569 ** (1)	2000
Income	−6.299 ** (1)	2000		
(Income)^2	−6.259 ** (1)	2000		
Energy consumption	−2.195 (1)	2003	−5.237 ** (1)	2001

Notes: Numbers inside the parentheses are lag lengths, which are chosen by the Schwarz Information Criterion (SIC). ** demarcates rejection of the null hypothesis at the 5% level. The 5% and 10% critical values for the ADF (Perron-Vogelsang), including a constant and trend, are −3.60 and −3.24 (−4.86 and −4.61), respectively.

Panels A and B in Table 3 give the results of the short- and long-run results for the logarithm of CO_2 emissions. Each of the coefficients gives the estimated coefficient of CO_2 emissions with respect to the corresponding explanatory variables. The coefficients on income variables are our main interest in this paper. The variables y_t and y_t^2 have *t*-statistics of above 1.7 in both the short- and long-run, and so they are statistically significant at least at the 10% significance level. What about interpreting the effect of income on CO_2 emissions? Because the coefficient on y_t is positive and the coefficient on y_t^2 is negative, this equation literally suggests that, at low level of income, an additional income growth has a positive effect on CO_2 emissions; in other words, growth increases CO_2 emissions. After turnaround value of income, however, the effect becomes negative, and the quadratic shape means that CO_2 emissions are decreasing as income grows. Thus, this finding seems to support the EKC in Kazakhstan. Note that, since the estimated elasticity of CO_2 emissions with respect to income is not directly observed from the estimated equation, it is calculated using the formula: $\beta_1 + 2\beta_2 \log(\text{income})$, in which β_1 and β_2 are the estimated coefficients on income and income^2. In the long-run; for example, when we plug in the minimum and maximum values of income in the sample, the estimated elasticity of CO_2 emissions with income range from is 0.254 to −0.623. This implies that income at low levels increases CO_2 but reduces it at high levels, providing evidence of the EKC hypothesis. For completeness, we then calculate the turning point income of the EKC for Kazakhstan. In the long-run, for example, the absolute value of the coefficient on income, 10.779, divided by twice the coefficient on income^2, 0.422, gives the turning point income = exp(10.779/2(0.422)) ≈ 351,979.1 Kazakhstani Tenge, which coincides with GDP per capita around 2001.

Table 3. Results of Estimated Short- and Long-Run Estimates.

Panel A: Short-Run Results			
$\Delta(\text{income})_t$	$\Delta(\text{income})_t{}^2$	$\Delta(\text{energy consumption})_t$	ec_{t-1}
6.858	−0.268	0.619	−0.636
(2.208) **	(−2.256) **	(3.315) **	(−3.936) **
Panel B: Long-Run Results			
income_t	$\text{income}_t{}^2$	$\text{energy consumption}_t$	Constant
10.779	−0.422	0.973	−74.444
(1.689) *	(−1.725) *	(10.448) **	(−1.793) *

Panel C: Diagnostic Statistics					
F-statistic	*LM*	RESET	Normality	ARCH	Heteroskedasticity
8.453 **	0.561	0.829	0.607	0.236	1.069
	[0.454]	[0.363]	[0.738]	[0.889]	[0.397]

Notes: Numbers inside the parentheses and brackets are *t*-statistics and *p*-values, respectively. ** and * demarcate significance at the 5% and 10% levels, respectively. The upper critical values at the 5% and 10% level are 4.01 and 3.52, respectively. LM and RESET represent the Lagrange multiplier test of serial correlation and Ramsey's test for misspecification, respectively.

The estimated effect of the energy consumption on CO_2 emissions is positive and highly significant in both the short- and long-run, indicating that CO_2 emissions increase as energy consumption increases. For example, a 1% increase in energy consumption increases CO_2 emissions by approximately 0.973% (0.619%) in the long (short)-run, holding income fixed. This further suggests that as found in previous studies (e.g., Iwata et al. 2011; Baek 2015; Baek and Choi 2017), energy consumption is one of main determinants in determining CO_2 emissions in Kazakhstan and should be accounted for when estimating the income-environment nexus accurately.

It is worth mentioning that the error-correction term (ec_{t-1}) represented by the linear combination of lagged level variables in Equation (2) is negative and very significant (Panel A). If the error-correction term is significantly negative, it works to push the selected variables back toward the equilibrium with shocks and provides another evidence of cointegration. The results show that the *t*-statistic on ec_{t-1} of −3.936 in our model is highly significant, thereby confirming cointegration. The coefficient of ec_{t-1} is −0.636, implying that deviation from the long-run equilibrium is corrected by approximately 64% in a year.

We also report additional diagnostic statistics (Panel C). First, in order to test for serial correlation and functional form misspecification, we employ the Lagrange Multiplier (LM) test and Ramsey's regression specification error test (RESET), which have a χ^2 distribution with one degree of freedom. The LM (RESET) statistic of 0.093 (0.367) with a *p*-value of 0.760 (0.544) indicates that there is no evidence of serial correlation (functional misspecification) in the CO_2 emissions model. The autoregressive conditional heteroskedasticity (ARCH) statistic of 0.236 with a *p*-value of 0.889 affirms that there is little evidence of heteroskedasticity in the residuals. Second, in order to test for stability of short-run and long-run coefficient estimates, the cumulative sum (CUSUM), cumulative sum of squares (CUSUMSQ) tests, and recursive coefficient stability tests are applied to the residual of our optimum model (Figure 2). Apparently, all estimates are generally stable over the sample period.

Finally, a possible criticism of our efforts to analyze the growth-environment nexus in Kazakhstan is that, since empirical studies typically show that air pollution (i.e., CO_2 emissions) increases monotonically with growth in developing economies, our conclusion that the EKC holds for Kazakhstan seems unrealistic to some.[10] Further, given that a growing body of the EKC literature

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[10]　Using different data from different sources perhaps results in such finding. In this paper, for example, CO_2 emissions are taken from Statistical Yearbook published by Agency on Statistics of the Republic of Kazakhstan, whereas income and

recently claims that it may not correct to include energy consumption variable in examining the growth-CO_2 nexus (Itkonen 2012; Liddle 2015; Jaforullah and King 2017), it would be worthwhile to explore this possibility by excluding energy consumption in Equation (2). The results show that in the long-run, the estimated coefficients on income and income^2 are +13.143 and −0.514, respectively, and they are statistically significant at the 10% level. When we plug in the minimum and maximum values of income in the sample, the estimated elasticity of CO_2 emissions with respect to income ranges from 0.324 to −0.741, which is quite close to the elasticities obtained from Equation (2) with energy consumption. Thus, this finding also supports the EKC for Kazakhstan. It should be admitted, however, that the availability of data used for the analysis (only 24 annual observations) is too limited to draw a robust conclusion about the existence of the EKC in a developing economy like Kazakhstan; our findings should thus be viewed with caution.[11]
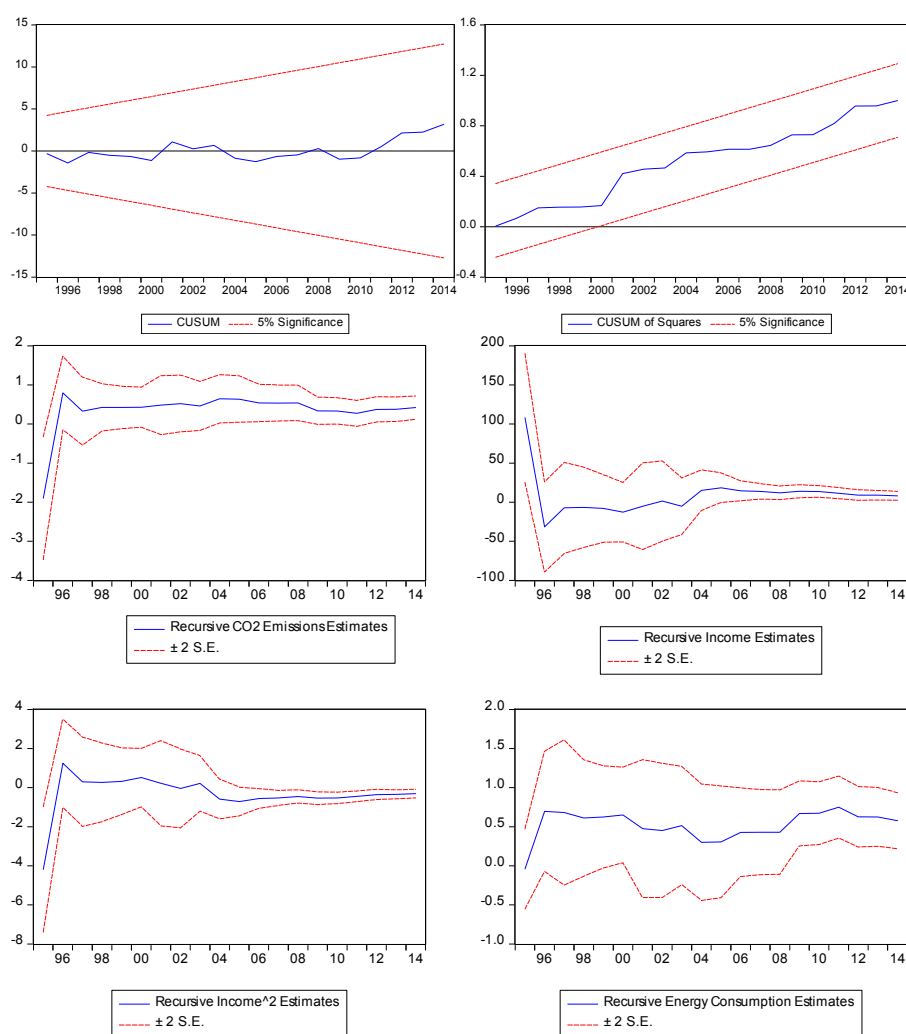
Figure 2. Plots of CUSUM, CUSUMSQ, and Recursive Coefficient Stability Tests.

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energy consumption are obtained from World Development Indicator (WDI) database. For this reason, we also re-estimate Equation (2) after collecting CO_2 emissions from WDI. However, we obtain almost the same results.

[11] It should be pointed out that, although we have a relatively small sample size, the regression fits reasonably well (*adj*. R^2 = 0.67) and passes all the necessary diagnostic tests (Panel C in Table 3). Further, we also adopt other alternative cointegration methods such as Fully Modified Least Squares (FMOLS), Dynamic Least Squares (DOLS), and Canonical Cointegration Regression (CCR) for robustness check, although those methods require all the variables to be *I*(1) processes. We also find almost the same results as those reported in Table 3. Combined with our diagnostic results, therefore, this should somehow mitigate our concern with the relatively short period of dataset and strengthen the credibility of our findings.

5. Concluding Remarks

In this short paper, the effect growth has on CO_2 emissions in Kazakhstan, controlling for total energy consumption, is examined in the autoregressive distributed lag (ARDL) cointegration framework. The contribution of this paper is to apply a country-specific time series method to the growth-environment nexus and to address the problem of aggregation bias drawn from the earlier cross-sectional and panel data studies. Our results show that growth increases CO_2 at a low level of income but at a high level decreases it, providing evidence in support of the EKC for CO_2 emissions in Kazakhstan. We also find that total energy consumption has an adverse effect on reducing CO_2 emissions.

A clear policy implication from our findings is that, given the fact that CO_2 emissions decrease with growth in Kazakhstan, any effort to promote economic growth may not cause a corresponding increase in CO_2 emissions. For this reason, government measures that need to be taken to reduce CO_2 emissions could be implemented without any sacrifices of growth in Kazakhstan. These measures include government policies directed more toward a low-fossil-fuel economy through an increase in the use of renewable energy and improved energy efficiency; the regulatory enforcement of reducing the greenhouse gas emitted from industry, transport, and heating; and the implementation of carbon sequestration technologies in power plants. Another policy implication is that, since energy consumption is mainly driven by continued growth in the industrial sector in Kazakhstan, any industrial policy implemented by the Kazakh government that aims to promote economic development could offset the positive income impacts on the environment, thereby leading to a rapid growth in CO_2 emissions. Thus, policy makers in Kazakhstan need to focus on attracting clean and energy efficient industries via foreign direct investment and reducing the carbon intensity of energy emitted per unit of energy consumption, thereby mitigating CO_2 emissions.

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