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The Environmental Kuznets Curve in Poland – Evidence from Threshold Cointegration Analysis

A b s t r a c t. The article aims to look at the long-run equilibrium relationship between per capita greenhouse gas emissions and per capita real GDP (EKC hypothesis) in an asymmetric framework using the non-linear threshold cointegration and error correction methodology for Polish economy during the period 2000 to 2012 (quarterly data). To test the robustness of the results the additional explanatory variable (per capita energy consumption) is added to the EKC model. The EKC hypothesis is tested using threshold autoregressive (TAR) and momentum threshold autoregressive (MTAR) cointegration method. Moreover, the threshold error correction model (TECM) is implemented in order to examine both the short-run and the long-run Granger-causal relationship between per capita greenhouse gas emissions and per capita income. We found strong evidence in favour of the EKC hypothesis for the Polish case and additionally we confirmed that adjustment of deviations toward the long-run equilibrium is asymmetric.

K e y w o r d s: Environmental Kuznets Curve, greenhouse gas emission, energy consumption, growth, threshold cointegration, Granger causality.

J E L Classification: C24, Q43, Q50.

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Introduction

The harmful effects of climate change made policy makers become increasingly interested in reducing greenhouse gas (GHG) emissions using different policy tools such as environmental taxation and regulation imposing the increased use of renewable energy. At the international level various steps are taken to motivate countries to reduce the emissions of GHG, e.g. the Kyoto Protocol or the EU energy and climate obligations for member countries. The problem of how the pollutants relate to the economy has been the subject of intense research in the last decades. One of the main developments in understanding the link between the environment and economy was the environmental Kuznets curve (EKC). The term 'environmental Kuznets curve' was coined almost simultaneously by Shafik and Bandyopadhyay (1992), Grossman and Krueger (1995) and Panayotou (1993). It refers by analogy to the inverted U-shaped relationship between the level of economic development and the degree of income inequality formulated by Kuznets (1955). The EKC hypothesis says that environmental degradation increases with per capita income during the early stages of economic growth, and then declines with per capita income after arriving at a threshold[1]. Hence, the relationship between income per capita and some types of pollutants is approximately an inverted U-shaped.

There is a wide stream of researches that has employed cointegration techniques to examine the relationship between per capita income and some types of pollutants, among others, Aspergis and Payne (2009), Halicioglu (2009), Soytas and Sari (2009), Ang (2009), Soytas et al. (2007). However, empirical results are mixed and not conclusive to give policy recommendations that can be applied across countries. The common feature of these studies was the linear approach and symmetric cointegration which may be a possible reason for ambiguous results. It has been suggested more recently (Balke and Fomby, 1997; Enders and Granger, 1998; Enders and Siklos, 2001) that the adjustment of deviations toward the long-run equilibrium need not be symmetric and reverting each period. To our knowledge, there are a very few studies that use non-linear (threshold) cointegration techniques for testing the EKC hypothesis, e.g. Fosten et. al. (2012), Esteve and Tamarit (2012).

This paper aims to look at testing the EKC hypothesis for the presence of threshold cointegration between per capita greenhouse gas emissions and per

[1] See Stern (2004), Coondoo and Dinda (2002), Dinda (2004), Luzzati and Orisni (2009), Halicioglu (2009) for extensive review surveys of studies which tested the EKC hypothesis.

capita income for the Polish economy during the period 2000 to 2012 (quarterly data). This approach will allow for a different speed of adjustment to the long-run equilibrium depending on whether emissions of greenhouse gas are above or below the EKC. Additionally, another explanatory variable is added, i.e. energy consumption, to test the robustness of the results. The standard EKC hypothesis and its extended version including energy consumption will be tested using threshold autoregressive (TAR) and momentum threshold autoregressive (MTAR) cointegration method of Enders and Granger (1998) and Enders and Siklos (2001). We will also concentrate on the short-run and long-run causal relationship between per capita greenhouse gas emissions and per capita income using threshold error correction models (T-ECM) and momentum threshold error correction model (M-TECM). To our knowledge, there is no such study that uses this approach for the case of Poland as if the EKC hypothesis is concerned.

The remainder of this paper is organized as follows. Section 2 presents the environmental Kuznets curve. Section 3 describes the methodology employed in the analysis. Section 4 describes the data and reports the empirical results. Section 5 concludes.

1. The Environmental Kuznets Curve

In classical approach to modelling the relationship between environmental degradation and income level the following quadratic function with the turning point occurring at a maximum pollutant level is used (Agras and Chapman, 1999):

$$EP_t = \beta_0 + \alpha_1 GDP_t + \alpha_2 GDP_t^2 + \mu_t, \tag{1}$$

where EP_t – emissions of some pollutant (per capita), GDP_t – real income (per capita), both variables are in logarithms, α_1, α_2, β_0 – estimated parameters, μ_t error term that may be serially correlated.

Based on the parameter values we may conclude about the shape of environmental pollution and income linkage. If $\alpha_1 > 0$ (or $\alpha_1 < 0$) while $\alpha_2 = 0$, then the function (1) is a monotonically increasing (decreasing) according to linear function behavior. Otherwise, if $\alpha_1 > 0$ and $\alpha_2 < 0$, an inverted U-shape describes the situation when the pollution level increases as a country develops until this development reaches a turning point and after that the rising incomes are accompanied by decreasing environmental degradation. The turning point value is approximated by following relation (Stern, 2004):

$$GDP_{TP} = \exp(-\alpha_1 / 2\alpha_2). \tag{2}$$

In the EKC literature the more sophisticated functional form is also taken into consideration with the third order polynomial for income factor (Dinda, 2004):

$$EP_t = \beta_0 + \alpha_1 GDP_t + \alpha_2 GDP_t^2 + \alpha_3 GDP_t^3 + \mu_t. \tag{3}$$

Similarly to the EKC model (2), the parameter estimates in equation (3) indicate the direction and character of the relationship between environmental pollutant and income. In the case of $\alpha_3 > 0$, $\alpha_2 < 0$, $\alpha_1 > 0$ and $\beta_0 < 0$ the N-shaped function is monotonically increasing with two possible turning points. In the case of opposite signs of cubic polynomial parameters, namely $\alpha_3 < 0$, $\alpha_2 > 0$, $\alpha_1 < 0$ and $\beta_0 > 0$, an inverse-N shape is more accurate for describing analyzed relationship.

In order to capture the effect of technological progress on pollution emission level the deterministic time trend (squared time trend) and some additional variables X_t that may affect EP_t can be included in equation (1) or (3).

In our empirical research of the long-run relationship between greenhouse gas emissions and economic growth first the standard EKC model (1) is assumed and further the EKC model with energy consumption (E_t) as additional variable is considered to test the robustness of results. In the latter case the model takes the form:

$$EP_t = \beta_0 + \alpha_1 GDP_t + \alpha_2 GDP_t^2 + \gamma E_t + \mu_t, \tag{4}$$

where γ − estimated parameter.

2. Methodology

The concept of cointegration implicitly assumes linearity and symmetry, what means that the adjustment of the deviations towards the long-run equilibrium is made instantaneously at each period and increases or decreases of the deviations are corrected in the same way. However, the cointegration tests and their extensions are misspecified if adjustment is asymmetric. To take the property of asymmetry into account, Enders and Siklos (2001) developed the concept of threshold cointegration. This is indeed an extension of residual-based two-stage estimation as developed by Engle and Granger (1987). The differences between them consist in the formulation of linearity and non-linearity from their second stage of unit root test.

Extracting from the long-run relationship (1) or (4) the disturbance term μ_t (first stage), in the second stage we focus on the coefficient estimates of ρ_1 and ρ_2 (adjustment parameters) in the following regression:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1-I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{r}\theta_i\Delta\mu_{t-i} + \varepsilon_t, \tag{5}$$

where ε_t is a white noise disturbance.
The Heaviside indicator function I_t is defined to depend on the lagged values of the residuals μ_t:

$$I_t = \begin{cases} 1 & if\ \mu_{t-1} \geq \tau \\ 0 & if\ \mu_{t-1} < \tau \end{cases} \tag{6}$$

or on the lagged changes in μ_t:

$$I_t = \begin{cases} 1 & if\ \Delta\mu_{t-1} \geq \tau \\ 0 & if\ \Delta\mu_{t-1} < \tau \end{cases} \tag{7}$$

where τ is a threshold value.
Equations (5)-(6) are referred to as the TAR model (threshold autoregressive model, Enders and Sikolos, 2001), while equations (5) and (7) are named the MTAR model (momentum-threshold autoregressive model; see Enders and Granger, 1998).

Petruccelli and Woolford (1984) proved that the necessary and sufficient conditions for the stationarity of residuals $\{\mu_t\}$ from the EKC model are $\rho_1 < 0$, $\rho_2 < 0$ and $(1+\rho_1)(1+\rho_2) < 1$ for any threshold value τ (Enders and Siklos, 2001). If these conditions are satisfied, $\mu_t = 0$ can be considered as the long run equilibrium value of the sequence. If μ_t is higher than the long-run equilibrium, the adjustment is $\rho_1\mu_{t-1}$, but if μ_t is lower than the long-run equilibrium, the adjustment is $\rho_2\mu_{t-1}$. Therefore, the equilibrium error behaves like a threshold autoregressive process (TAR). The MTAR model – according to Enders and Granger (1998) – is especially valuable when adjustment is asymmetric as the deviations μ_t exhibit more 'momentum' in one direction than in the other. Hence, the TAR model allows to examine whether the positive deviations $(\mu_t > 0)$ from the long-run equilibrium have different effects on the behavior of emissions than do the negative deviations $(\mu_t < 0)$, while the MTAR model allows to display various amounts of autoregressive decay depending on whether the series is increasing or decreasing (Fosten et al., 2012). There is no prescribed rule as to whether to use the TAR or MTAR model, but it is recommended to select the best adjustment mechanism (TAR or MTAR) using the AIC (Akaike Information Criterion) or SBC (Schwarz Bayesian Criterion) information criteria (Enders, Chumrusphonlert, 2004).

In general, the threshold value τ governing the asymmetric behavior is unknown and has to be estimated along with the values of adjustment parameters ρ_1 and ρ_2. In our studies we follow Enders and Siklos (2001) and Yau and Nieh (2009) by employing Chan's (1993) methodology[2] of searching the consistent estimates of threshold value. However, in many economic applications this value is set to zero, $\tau = 0$, and then the cointegrating vector coincides with the attractor ($\mu_t = 0$).

Once the threshold value τ is obtained and the TAR or MTAR models are estimated, then testing for threshold cointegration can be performed. First, the null hypothesis of no cointegration $H_0 : \rho_1 = \rho_2 = 0$ is tested, and when it is rejected, then the null hypothesis of symmetric adjustment $H_0 : \rho_1 = \rho_2$ is verified. To test the null hypothesis of no threshold cointegration, Enders and Siklos (2001) proposed the Φ_μ statistics which is the F statistics. As the distribution of Φ_μ is non-standard, appropriate critical values were tabulated by Enders and Siklos (2001) and later modified by Wane et al. (2004). In the presence of cointegration (rejection of $H_0 : \rho_1 = \rho_2 = 0$), the null hypothesis $H_0 : \rho_1 = \rho_2$ of symmetric adjustment can be tested using the standard F statistics. When the adjustments coefficients are equal (symmetric adjustment), equation (5) converges the standard ADF test. Rejecting both the null hypotheses of $\rho_1 = \rho_2 = 0$ and $\rho_1 = \rho_2$ implies the existence of threshold cointegration with asymmetric adjustment. For instance, if the null of symmetric adjustment is rejected and $|\rho_1| > |\rho_2|$, then it implies that positive deviations (above threshold) of emissions from the long-run EKC tend to revert quickly towards equilibrium, whereas negative deviations of emissions (below threshold) tend to persist. Hence, it suggests faster convergence for positive deviations of emissions (above thresh-

[2] The Chan's method to find the consistent estimate of the threshold value arranges the values μ_{t-1} or $\Delta\mu_{t-1}$ in ascending order, excludes the smallest and largest 15 percent, and the parameter that yields the smallest sum of squared residuals over the remaining 70 percent is the consistent estimate of the threshold.

old) than negative deviations (below threshold) from the long-run EKC relationship.

Given the threshold cointegration is found, the next step proceeds with the Granger-causality test using the advanced threshold error correction model (TECM) or momentum-threshold error correction model (M-TECM) by Enders and Granger (1998) and Enders and Siklos (2001). The threshold ECM is expressed as the following:

$$\Delta Y_{jt} = \beta + \lambda_1 Z_{t-1}^{+} + \lambda_2 Z_{t-1}^{-} + \sum_{i=1}^{q_1} \delta_i \Delta EP_{t-i} + \sum_{i=1}^{q_2} \theta_i \Delta GDP_{t-i} + \sum_{i=1}^{q_3} \gamma_i \Delta GDP_{t-i}^2 + v_t, \quad (8)$$

or when additional variable E_t is included

$$\begin{aligned}\Delta Y_{jt} = \beta + \lambda_1 Z_{t-1}^{+} + \lambda_2 Z_{t-1}^{-} + \sum_{i=1}^{q_1} \delta_i \Delta EP_{t-i} + \sum_{i=1}^{q_2} \theta_i \Delta GDP_{t-i} + \sum_{i=1}^{q_3} \gamma_i \Delta GDP_{t-i}^2 + \\ + \sum_{i=1}^{q_4} \alpha_i \cdot \Delta E_{t-i} + v_t,\end{aligned} \quad (9)$$

where in equation (8) and (9): $\Delta Y_{jt} = \left(\Delta EP_t, \Delta GDP_t, \Delta GDP_t^2\right)$, and $Z_{t-1}^{+} = I_t \hat{\mu}_{t-1}$ and $Z_{t-1}^{-} = (1 - I_t)\hat{\mu}_{t-1}$, I_t – Heaviside indicator function determined by (6) or (7), $\hat{\mu}_{t-1}$ is obtained from the estimated long-run relationship (1) or (4), and v_t is a white noise disturbance.

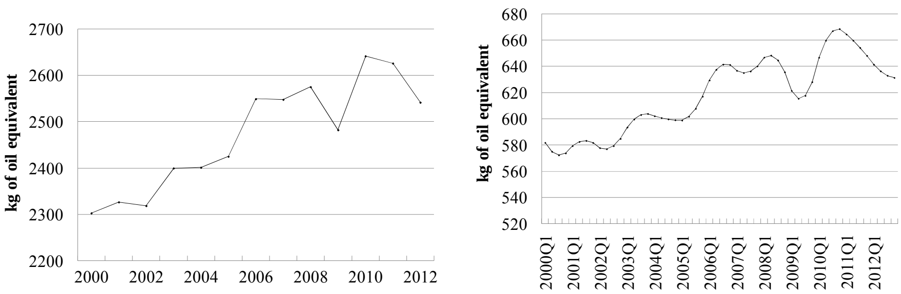
Based on the equation (8) and (9) the Granger-causality tests are employed. The long-run causality is determined by the significance of adjustment parameters λ_1 and λ_2. To test short-run causality (or weak causality) the joint significance of all the coefficients δ_i of ΔEP_{t-i}, θ_i of ΔGDP_{t-i}, γ_i of ΔGDP_{t-i}^2 or α_i of ΔE_{t-i} is examined using the Wald F test. It is also desirable to check whether this two sources of causation (short and long-run) are jointly significant. This can be done by providing the Wald F statistics for the interactive terms, i.e. the λ terms and the explanatory variables (e.g. $H_0 : \theta_i = \lambda_1 = 0$, $H_0 : \theta_i = \lambda_2 = 0$). The joint test indicates which variables bear the burden of short-run adjustment to re-establish long-run equilibrium, given a shock to the system (Asafu-Adjaye, 2000; Mehrara et al., 2012). This is referred to as the strong Granger causality test.

3. Empirical Results

3.1. Data Source

The data used in this study consist of greenhouse gas emissions (EP_t) (in tons of CO_2 equivalent per capita), real gross domestic product per capita (GDP_t) and energy consumption[3] (E_t) in kilo of oil equivalent per capita in Poland. The quarterly data for GDP are obtained from the Central Statistical Office (www.stat.gov.pl) of Poland, while annual data describing greenhouse gas emissions and energy consumption in Poland were obtained from the Eurostat database.

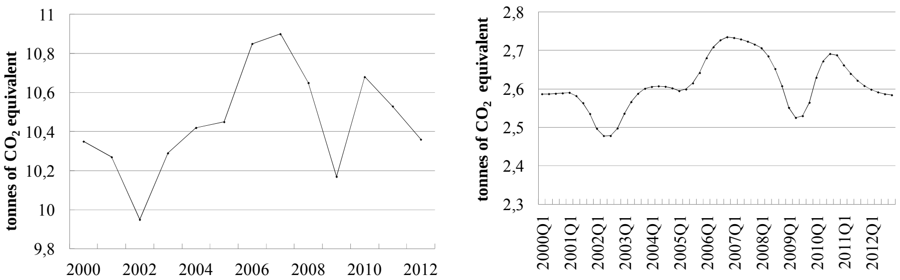
To obtain the real GDP some transformations were made, i.e. quarterly nominal GDP data were transformed by the authors into real GDP in 2005 prices using GDP deflator. Since the GDP data were characterized by significant quarterly seasonality, the TRAMO procedure of Gretl software was applied to adjust it seasonally. The frequency of real GDP series is quarterly, however, the greenhouse gas emissions and energy consumption data are only available at annual frequency. Therefore, we interpolated annual data to quarterly frequency by employing the Denton-Cholette method (Sax, Steiner, 2013) in the R software. The sample period is from 2000:Q1 to 2012:Q4. All variables are employed with their natural logarithms form to reduce heteroscedasticity and to obtain the growth rate of the relevant variables by their differenced logarithms.



a) annual data b) disaggregated data to quarterly frequency

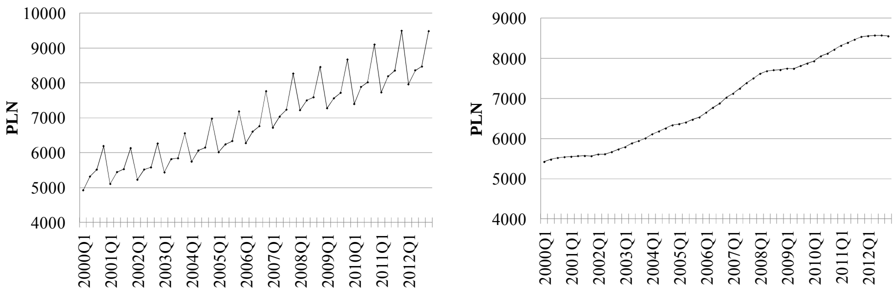
Figure 1. Energy use (in kg of oil equivalent per capita) in Poland in the period 2000–2012 – for annual data and disaggregated data to quarterly frequency

[3] Energy consumption covers: consumption by the energy sector itself; distribution and transformation losses; final energy consumption by end users (see epp.eurostat.ec.europa.eu, Eurostat glossary).



a) annual data b) disaggregated data to quarterly frequency

Figure 2. Greenhouse gas emissions (in tons of CO_2 equivalent per capita) in Poland in the period 2000–2012 – for annual data and disaggregated data to quarterly frequency



a) seasonally unadjusted b) seasonally adjusted

Figure 3. GDP (in PLN per capita, in 2005 prices) in Poland, 2000:Q1–2012:Q4

The evolution of time series data (original and disaggregated) employed in our analysis is shown in Figure 1–3.

3.2. Cointegration Analysis with Asymmetric Adjustment

Before performing cointegration analysis, we use the Augmented Dickey-Fuller Generalized Least Squares (ADF-GLS; Elliot et al., 1996) and Kwiatkowski-Phillips-Schmidt-Shinn (KPSS; Kwiatkowski et al., 1992) tests to identify the order of integration for each variable. In Table 1, the ADF-GLS tests show that the unit root hypothesis cannot be rejected at any significant level for each variable in levels. Further investigations of the unit root hypothesis indicate that the first differenced variables are stationary at least at the 10% level of significance. We also apply the KPSS unit root test based on the null hypothesis of stationarity (or no unit root). The results show that the null hypothesis of stationarity is rejected at least at the 10% significance level. Hence, all series are found to be integrated of order $I(1)$.

Table 1. The results of unit root tests

Variables	Levels		Differences	
	ADF-GLS	KPSS	ADF-GLS	KPSS
EP_t	–1.522 [3]	0.463 [2]*	–5.246 [3]***	0.073 [2]
GDP_t	0.041 [4]	1.218 [4]***	–2.307 [4]**	0.139 [4]
GDP^2_t	–0.144 [5]	1.217 [4]***	–1.676 [5]*	0.132 [4]
E_t	–0.957 [2]	1.548 [2]***	–5.368 [2]***	0.110 [2]

Note: (***), (**), (*) in ADF-GLS tests respectively indicate the rejection of the null hypothesis that series has a unit root at 1%, 5% and 10% levels of significance, while in KPSS tests indicate the rejection of the null hypothesis that series is stationary. The numbers inside the brackets are the optimum lag lengths determined using AIC in ADF-GLS tests and the bandwidth is used using the Newey-West method in KPSS tests.

Table 2 contains cointegration test results of the standard long-run EKC in the form (1) and Table 3 – long-run EKC including energy consumption in the form (4), when considering threshold (TAR) and momentum adjustment (MTAR). The tables report values of the adjustment coefficients ρ_1 and ρ_2, the Φ_μ statistics for the null hypothesis of no cointegration (a unit root in μ_t) against the alternative of cointegration with asymmetric adjustment. The *F*-test is used to test whether the adjustment back to long-run equilibrium is symmetric $\rho_1 = \rho_2$.

It can be seen that all coefficients ρ_1, ρ_2 have negative signs and are significant at least at 10% significance level. The necessary and sufficient conditions for cointegration hold in the case of all TAR and M-TAR models because the null hypothesis $H_0 : \rho_1 = \rho_2 = 0$ is rejected (at 5% significance level)[4] – see Table 2. Using the standard F-statistics for the restriction $H_0 : \rho_1 = \rho_2$ it is shown that asymmetric cointegration is strongly significant only in the TAR model with $\tau = -0.0169$, whereas in the TAR model with $\tau = 0$ the support for asymmetric cointegration is only at 10% significance level. This evidence favors that the adjustment back to equilibrium between greenhouse gas emissions (EP_t) and gross domestic product (GDP_t) is nonlinear. For choosing the more appropriate adjustment process (TAR or M-TAR) we will follow Enders and Chumrusphonlert's (2004) advice in using AIC or SBC to select the best adjustment mechanism. The reported SBC for each model shows that the TAR model with threshold value $\tau = -0.0169$ is more appropriate adjustment mechanism (the minimum SBC is in bold in Table 2).

[4] It is worth noting that this null hypothesis is also rejected when critical values from Wane et al. (2004) are taken.

Table 2. Results of TAR and M-TAR Enders-Siklos (E-S) test for cointegration on the standard EKC model

	ρ_1	ρ_2	Φ_μ ($\rho_1=\rho_2=0$)	F ($\rho_1=\rho_2$)	Lag	SBC	LB(4)
TAR	–0.124	–0.199	12.76***	3.075	5	–387.6	5.15
τ=0	(–2.372)**	(–5.036)***		[0.087]*			[0.273]
M-TAR	–0.125	–0.191	11.722***	1.722	5	–386.1	6,5
τ=0	(–2.489)**	(–4.804)***		[0.197]			[0.165]
TAR	–0.119	–0.222	15.184***	6.236	5	–390.9	4.74
τ=–0.0169	(–2.939)**	(–5.500)***		[0.017]**			[0.315]
M-TAR	–0.148	–0.179	10.722***	0.416	5	–384.6	5.35
τ=0.0023	(–2.925)**	(–4.536)***		[0.522]			[0.253]

Note: *** (**) (*) indicate significance at 1% (5%) (10%) level. Critical values for Φ_μ statistics from Enders, Siklos (2001). T-statistics for in ρ parentheses. *p*-values in brackets. LB(4) for Ljung-Box statistics. The lag length is selected such that the AIC is minimized.

Table 3. Results of TAR and M-TAR Enders-Siklos (E-S) test for cointegration on the EKC model including energy consumption

	ρ_1	ρ_2	Φ_μ ($\rho_1=\rho_2=0$)	F ($\rho_1=\rho_2$)	Lag	SBC	LB(4)
TAR	–0.104	–0.056	5.83*	2.155	2	–488.06	6.91
τ=0	(–3.249)**	(–2.20)*		[0.149]			[0.141]
M-TAR	–0.071	–0.072	4.537	0.002	2	–485.77	7.26
τ=0	(–2.683)*	(–2.325)*		[0.961]			[0.123]
TAR	–0.115	–0.055	6.616**	3.462*	2	–489.40	6.67
τ=0.0097	(–3.516)**	(–2.237)*		[0.069]			[0.155]
M-TAR	–0.064	–0.082	4.74	0.34	2	–486.13	0.03
τ=8.8E-05	(–2.36)*	(–2.74)**		[0.563]			[0.999]

Note: ** (*) indicate significance at 5% (10%) level. Critical values for Φ_μ statistics from Enders, Siklos (2001). T-statistics for ρ in parentheses. *p*-values in brackets. LB(4) for Ljung-Box statistics. The lag length is selected such that the AIC is minimized.

We can see in Table 2 that the point estimates ρ_1 and ρ_2 suggest faster convergence for the deviations from long-run EKC equilibrium when they are below the threshold ($\mu_t < \tau = -0.0169$) than when they are above the threshold because $|\rho_1| < |\rho_2|$. We see that about 22% of the deviation from equilibrium is corrected in the next period when emissions are falling, compared to about 12% when they are rising. This means that short-run adjustment towards the EKC equilibrium reverts more quickly when the greenhouse gas emissions are decreasing (below the threshold) and tends to persist more when the greenhouse gas emissions are increasing (above the threshold). While the opposite result might be expected, this evidence should not be surprising when we look at the Poland's energy profile. Heavy reliance on coal makes Poland a relatively carbon-intensive economy, compared to the IEA Europe average. The large-scale transition of the Polish energy sector, which is characterised by ageing infrastructure[5], to a low-carbon economy requires huge long-term investments and an adequate policy and regulatory framework. Therefore, from a short run perspective the potential reduction of greenhouse gas emissions should be rather combined with the energy efficiency improvements.

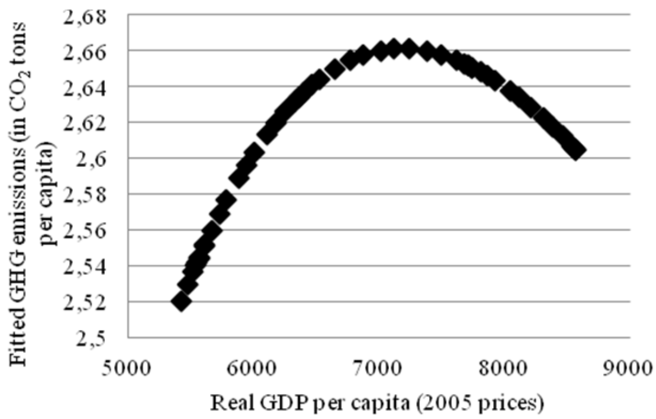


Figure 4. The fitted values of the estimated EKC results for greenhouse gas emissions

Table 4. Estimated parameters of long-run EKC equation

Variable	Estimates of parameters (t-statistics)	
Constant	–53.298 (–4.228)***	–23.041 (–2.444)**
GDP_t	12.225 (4.280)***	4.345 (1.978)*
GDP^2_t	–0.688 (–4.255)***	–0.255 (–2.065)**
E_t		0.863 (7.639)***

Note: *** (**) (*) indicate significance at 1% (5%) (10%) level. T-statistics for ρ in parentheses. Estimated parameters for equation (1) and (4).

[5] Nearly half of today's operating capacity of the Polish energy sector is older than 30 years. See: Energy Policies of IEA Countries. Poland (2011).

Having established the TAR cointegration in the long-run relationship for greenhouse gas emissions, it is now justified to analyse the estimation results (Table 4) and the implications for the EKC hypothesis. We can see that in terms of the coefficients described in the EKC relation (1), we have $\beta_0 < 0,\ \alpha_1 > 0$ and $\alpha_2 < 0,$ which implies the inverted U-shaped function.

The fitted values of greenhouse gas emissions for the observed values of real GDP are displayed in Figure 4. This results show the strong evidence in favour of the EKC hypothesis. The turning point in the observed range of real GDP for greenhouse gas occurs at PLN7218.8 per capita which corresponds to around 2007Q2.

The addition of energy consumption (E_t) to the standard EKC model has not affected the results in terms of the presence of asymmetric cointegration (Table 3), i.e. in the TAR framework (with $\tau = 0.0097$) the asymmetric cointegration is observed but only at 10% significance level when critical values from Enders and Siklos (2001) are used. Moreover, the non-linear relationship remains significant and correctly signed (ρ_1 and ρ_2 are negative), suggesting the relationship is reasonably robust (see Table 3). However, when critical values from Wane et al. (2004) are taken, then the null hypothesis of non-cointegration $(H_0 : \rho_1 = \rho_2 = 0)$ cannot be rejected. Hence, the results behind the threshold cointegration for the EKC model with energy consumption are rather weak and should be treated with caution.

Also the point estimates ρ_1 and ρ_2 differ with regard to the speed of adjustment process and direction of convergence for deviations above and below the threshold value when compared to the standard EKC model (Table 2). Now, faster convergence for deviations (from the long-run EKC) above the threshold than for those below the threshold is observed. We can see (Table 3) that about 12% of the deviation from equilibrium is corrected when emissions are above the threshold $(\mu_t \geq \tau = 0.0097)$, compared to only 5.5% when they are below the threshold $(\mu_t < \tau = 0.0097)$. Therefore, the short-run adjustment towards the long-run equilibrium reverts more quickly when emissions are increasing (above the threshold) and tends to persist when emissions are decreasing (below the threshold). This result is quite the opposite to that obtained for the standard EKC model what may indicate that energy consumption (E_t) is an important determinant of greenhouse gas emissions. The faster correction for deviations in emissions, if they are too high, indicates that in the presence of environmental regulation[6] the pressure to reduce them into their long-run levels was occurred. However, these emissions reductions were to a large extent achieved through the restructuring of Polish industry and energy efficiency improvements.

3.3. Threshold Error Correction Models

The positive finding of cointegration with TAR adjustment justifies estimation of threshold error correction model (8) and (9) and testing the Granger causality[7]. The T-ECM models are estimated for changes in greenhouse gas emissions (ΔEP_t) and gross domestic product (ΔGDP_t), assuming the estimates of threshold values obtained in previous step, i.e. $\tau = -0.0169$ − in model (8) and $\tau = 0.0097$ − in model (9) (see also Table 2 and 3). The T-ECM for the ΔE_t is not estimated since the energy consumption is added into the standard EKC model to test the robustness of the results.

Based on equations (8) and (9), the Granger causality tests are employed to verify whether all the coefficients of ΔEP_{t-i}, ΔGDP_{t-i} or ΔGDP_{t-i}^2 (or ΔE_{t-i} in eq. (9)) are jointly statistically different from zero based on a standard F-test (Wald test) and/or whether the coefficients (λ_1, λ_2) of the error correction term are significant. To determine the appropriate lag lengths we apply the SBC criterion, and empirically find that the lag lengths are equal: $q_1 = q_2 = 2$, $q_2 = q_3 = 3$. Table 5 presents estimates of the error correction parameters along with Wald F test statistics regarding Granger causality.

We will first interpret the results for the T-ECMs and then for T-ECMs including ΔE_t (Table 5).

While the adjustment speed on the exceeding or underlying threshold level in the T-ECM model for ΔEP_t has the 'right' direction (λ_1 and λ_2 have a negative sign and are significant) by acting to eliminate deviations from the long-run equilibrium, the T-ECM model for ΔGDP_t adjusts to the

[6] Regulation on air pollution has become increasingly stringent, including international protocols such as the Oslo Protocol, the Kyoto Protocol and the EU energy and climate obligation for member countries.

[7] The T-ECM for ΔGDP_t^2 is run, but not reported because it has little useful economic interpretation (results are available from authors on request).

'wrong' direction (λ_1 has a positive sign) for one regime, and additionally parameters λ_1 and λ_2 are insignificant. In the ΔEP_t model the adjustment speed responds faster in the lower regime (emissions below the threshold value) than in the higher regime with increasing deviations from the long-run equilibrium (emissions above the threshold value). This is consistent with the results in Table 2. Now, the greenhouse gas emissions converge to their long-run equilibrium at the rate of 13.2% with a deviation below the threshold and at a lower rate of 9.3% with a deviation above the threshold.

Table 5. Estimates of threshold error correction model (T-ECM) and the Wald *F* statistics for Granger causality

	T-ECM		T-ECM including ΔE_t	
	ΔEP_t	ΔGDP_t	ΔEP_t	ΔGDP_t
λ_1	–0.093* (–2.021)	0.046 (0.693)	–0.081 (–1.012)	–0.015 (–0.141)
λ_2	–0.132** (–2.610)	–0.073 (–1.008)	–0.022 (–0.357)	–0.163* (–1.961)
H_0: θ_i=0	1.328 [0.280]		1.519 [0.227]	
H_0: θ_i=λ_1=0	2.028 [0.111]		1.308 [0.286]	
H_0: θ_i=λ_2=0	2.952** [0.032]		1.209 [0.324]	
H_0: γ_i=0	1.386 [0.262]	0.562 [0.643]	1.622 [0.202]	1.000 [0.404]
H_0: γ_i=λ_1=0	2.059 [0.106]	0.489 [0.744]	1.387 [0.259]	0.750 [0.565]
H_0: γ_i=λ_2=0	3.039** [0.029]	0.786 [0.542]	1.293 [0.292]	1.894 [0.133]
H_0: δ_i=0		5.626** [0.007]		1.162 [0.325]
H_0: δ_i=λ_1=0		3.796** [0.018]		1.056 [0.380]
H_0: δ_i=λ_2=0		4.487** [0.009]		3.367** [0.029]
H_0: α_i=0			3.857** [0.031]	0.129 [0.879]
H_0: α_i =λ_1=0			2.674* [0.062]	0.187 [0.904]
H_0: α_i =λ_2=0			2.572* [0.069]	2.012 [0.130]
LB(3)	2.60 [0.457]	1.60 [0.659]	2.66 [0.448]	1.95 [0.583]
BIC	–395.72	–362.00	–387.32	–360.22

Note: *** (**) (*) indicate significance at 1% (5%) (10%) level. *t*-statistics for λ_1, λ_2 in parentheses. *p*-values for Wald statistics in brackets. LB(3) for Ljung-Box statistics along with p-values.

The results of the Granger causality tests show that there is no short-run causality running from real GDP_t (and square of GDP_t) to EP_t, but there is an unidirectional short-run causality from greenhouse gas emissions to per capita real GDP_t. This can be interpreted as the non-rejection of $H_0 : \theta_i = 0$ and $H_0 : \gamma_i = 0$ in the ΔEP_t equation and rejection of $H_0 : \delta_i = 0$ in the ΔGDP_t equation (see Table 5, 2nd and 3rd column). Besides, the strong long-run causality running from per capita real GDP (and square of real GDP) to greenhouse gas emissions is found in the regime below the threshold value of −0.00169 (rejection of $H_0: \theta_i = \lambda_2 = 0,\ \ H_0: \gamma_i = \lambda_2 = 0$). The results for the ΔGDP_t equation suggest reverse long-run causality running from greenhouse gas emissions to per capita real GDP (rejection of $H_0: \delta_i = \lambda_1 = 0$ and $H_0: \delta_i = \lambda_2 = 0$). This implies that deviations from the long-run EKC are corrected not only by movements in greenhouse gas emissions but also by movements in per capita real GDP.

In the T-ECM model for ΔEP_t, including energy consumption (Table 5, 4th column), the parameters λ_1, λ_2 are correctly signed, but statistically insignificant[8]. Nonetheless, it is worth noting that the correction back to equilibrium is faster in regime above the threshold value of 0.0097 (since $|\lambda_1| > |\lambda_2|$) unlike the results obtained in the T-ECM not including energy consumption where faster convergence to equilibrium occurred in the regime below the threshold. This means that the addition of energy consumption into the T-ECM changed the response of greenhouse gas emissions to error correction. The manifested influence of energy consumption to reduce greenhouse gas emissions should be rather attached to the promotion of technological progress and the energy efficiency improvements, but not the large-scale transition to low-carbon economy, as has already been mentioned.

Further, there is neither short-run Granger causality (or weak causality) nor long-run causality from GDP_t to EP_t in the T-ECM including energy consumption (the non-rejection of null hypotheses $\theta_i = 0,\ \ \theta_i = \lambda_1 = 0,$ $\theta_i = \lambda_2 = 0,\ \ \gamma_i = 0,\ \ \gamma_i = \lambda_1 = 0,\ \gamma_i = \lambda_2 = 0$ in Table 5, 4th column). The non-significance of the F-statistics for GDP indicates that it is exogenous in the system. However, there exists the short-run and long-run causality from energy consumption to greenhouse gas emissions (the rejection of null hypotheses: $\alpha_i = 0,\ \ \alpha_i = \lambda_1 = 0$ and $\alpha_i = \lambda_2 = 0$ in the ΔEP_t model including ΔE_t). This evidence suggests that energy consumption has an effect on greenhouse gas emissions and bears the burden of short-run adjustment to restore long-run equilibrium after a shock to the system. In the T-ECM for

[8] This finding should be rather attached to the effect of small sample and reduction of degrees of freedom in the estimated threshold error correction models due to bigger number of variables and lagged terms in comparison to previous TAR estimations (see Table 2 and 3). Therefore the results in Table 5 should be interpreted with caution and validated on the larger sample.

ΔGDP_t (including energy consumption) the rejection of $H_0 : \delta_i = \lambda_2 = 0$ suggests the unidirectional long-run causality from greenhouse gas emissions to real GDP in regime below the threshold value of 0.0097. The results from causality analysis based on the T-ECM (including energy consumption) indicate that re-establishing of the long-run equilibrium is carried out by the interaction of greenhouse gas emissions and energy consumption, but the long run effect of GDP is rather weak.

Conclusions

The paper has aimed to investigate the EKC hypothesis for the case of the Polish economy during the period 2000-2012. We tested for the presence of threshold cointegration between per capita greenhouse gas emissions and per capita income (real GDP). Moreover, to test the robustness of the results, we considered the standard EKC hypothesis with the addition of per capita energy consumption to the model. To address such an issue, we applied the threshold cointegration model which allows the nonlinear adjustment to the long-run equilibrium.

In the case of the standard EKC relationship, the results of threshold cointegration indicate that per capita greenhouse gas emissions and per capita real GDP are cointegrated with an asymmetric adjustment process. Adjustments towards the long-run equilibrium revert more quickly when emissions are below the threshold value and tend to persist more when emissions are above the threshold value. The evidence of long persistence of adjustments to the equilibrium in higher regime (emissions above the threshold) may be explained by the specific of Polish energy sector, namely heavy reliance on coal. As a consequence, the transition to a low-carbon economy will require huge long-term investments and an adequate policy and regulatory framework. Therefore, from a short run perspective the reduction of emissions should be rather combined with the energy efficiency improvements. With regard to the Granger causality tests the results are following. In the lower regime the bidirectional long-run causality between per capita real GDP and per capita greenhouse gas emissions is found. The short-run dynamics suggests that there is no causal relationship from real GDP to greenhouse gas emissions but there is an unidirectional Granger causality from greenhouse gas emissions to real GDP.

Our results find strong evidence in favour of the EKC hypothesis with per capita greenhouse gas emissions having an inverse U-relation with real GDP per capita. The evidence suggests that the turning point in the observed range of real GDP for greenhouse gas emissions occurred at PLN7218.8. As a consequence, a decoupling between the two variable appears, i.e. the growth of emissions of some pollutant is slower than the economic growth.

The addition of energy consumption to the standard EKC model has not affected the results in terms of the presence of asymmetric cointegration but has changed the direction of convergence for deviations above and below the threshold value. Namely, adjustment towards long-run equilibrium is faster in the higher regime (emissions above the threshold) than in the lower regime (emissions below the threshold). This mean that in the presence of environmental regulation the pressure to reduce emissions into their long-run levels is observed. However, it should be emphasized that the abatement of emissions is rather achieved through the restructuring of Polish industry, promoting technological progress and energy efficiency improvements, but not the large-scale transition to low-carbon economy. The results of Granger causality tests indicate that there is no short-run causality from real GDP to greenhouse gas emissions. However, there is short and long-run causality from energy consumption to greenhouse gas emissions. Besides, the unidirectional long-run causality from greenhouse gas emissions to real GDP in regime below the threshold value is observed. The results from causality analysis indicate that re-establishing of the long-run EKC is carried out by the interaction of greenhouse gas emissions and energy consumption, but the long run effect of GDP is rather weak.

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Ekologiczna krzywa Kuznetsa dla Polski − analiza progowej kointegracji

Z a r y s t r e ś c i. Artykuł przedstawia analizę relacji długookresowej między emisją gazów cieplarnianych w przeliczeniu na mieszkańca a realnym PKB w przeliczeniu na mieszkańca (hipoteza EKC, ekologiczna krzywa Kuznetsa) z wykorzystaniem podejścia progowej (asymetrycznej) kointegracji i modelu korekty błędem dla polskiej gospodarki w okresie 2000-2012 (dane kwartalne). Standardowy model EKC został rozszerzony o zużycie energii w celu zbadania wpływu dodatkowych zmiennych na wyniki. Hipoteza EKC była sprawdzana z wykorzystaniem progowych modeli autoregresyjnych (TAR i MTAR). Ponadto, dla zbadania krótkookresowej i długookresowej przyczynowości Grangera między emisją gazów cieplarnianych w przeliczeniu na mieszkańca a realnym PKB w przeliczeniu na mieszkańca zastosowano progowy model korekty błędem. Otrzymane wyniki dostarczają istotnych dowodów na rzecz hipotezy EKC dla przypadku Polski oraz wskazują, że tymczasowe odchylenia od ścieżki długookresowej równowagi EKC są korygowane w asymetryczny sposób.

S ł o w a k l u c z o w e: ekologiczna krzywa Kuznetsa, emisja gazów cieplarnianych, zużycie energii, wzrost gospodarczy, kointegracja progowa, przyczynowość w sensie Grangera.