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Beyond the Turning Point: The Environmental Kuznets Curve for CO_2 Emissions in Romania

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Abstract
This thesis studies the long-run relationship between environmental degradation measured by carbon dioxide (CO_2) emissions per capita and economic growth measured by gross domestic product per capita in Romania, between 1968 and 2018. The study is conducted using the environmental Kuznets curve (EKC) framework and the autoregressive distributed lag bounds test for cointegration. Furthermore, the direction of short-run and long-run causality is investigated using the Granger causality within the Vector Error Correction Model. The EKC hypothesizes the existence of an inverted U-shaped relationship between environmental degradation and the level of income. The analysis lends support to this hypothesis for Romania in the studied time frame and identifies a peak of CO_2 emissions at a real GDP per capita of approximately $5,700. This suggests that a green growth strategy is suitable for Romania at this point in time. However, the situation should be constantly monitored and efforts to reduce emissions ought to be maintained to reach the net zero targets.

Keywords: environmental Kuznets curve, autoregressive distributed lag, Granger causality, Romania
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Acronyms

ADF augmented Dikey-Fuller unit root test.

AIC Akaike information criterion.

ARDL autoregressive distributed lag.

CEE Central and Eastern Europe.

CO carbon monoxide.

$\mathbf{CO_2}$ carbon dioxide.

ECM error correction model.

ECT error correction term.

EKC environmental Kuznets curve.

ETS Emissions Trading Scheme.

EU European Union.

GDP gross domestic product.

GHGs greenhouse gases.

$\mathbf{SO_2}$ sulfur dioxide.

$\mathbf{SO_x}$ sulfur oxides.

SPM suspended particulate matter.

VECM vector error correction model.

1 Introduction

The stock of greenhouse gases (GHGs) in the atmosphere has been rising over time due to anthropogenic emissions, causing climate change with phenomena such as increasing average temperatures, extreme weather events, and rising sea levels (Nunez, 2021). Carbon dioxide (CO_2) is a heat-trapping gas (Fecht, 2021) that triggers irreversible change over centuries to millennia, its accumulation being attributable to the combustion of fossil fuels and deforestation (IPCC, 2022). It has long-run negative consequences and is a global pollutant (Cole et al., 1997). This creates an incentive for countries to free ride on the mitigation efforts of others, as the repercussions of CO_2 emissions transcend national borders and affect future generations (Roca et al., 2001; Kunnas and Myllyntaus, 2007). Nevertheless, economic growth could be part of the solution to environmental degradation (Rothman and de Bruyn, 1998) instead of the problem (Soytas et al., 2007), if the right policies are in place to safeguard the environment (Ng and Wang, 1993).

In this context, uncovering the dynamics between CO_2 emissions and income sheds light on whether Romania should adopt a green growth or a degrowth strategy, whilst keeping in mind the net zero targets of the European Green Deal. Therefore, this paper investigates whether there exists a long-run relationship between environmental degradation, measured by CO_2 emissions per capita and income, measured by real gross domestic product (GDP) per capita in Romania, between 1968 and 2018, as well as the nature of this relationship.

Romania makes an interesting case study given its shift from a centrally-planned to a market-based economy in in the early 1990s. Furthermore, the country's national average shares of renewable energy are below 2030 benchmarks, as it still relies heavily on coal and other fossil fuels (OECD, 2021). When considering all GHGs in post-communist Romania, the share of CO_2 emissions has been the largest and this has not changed over time (Romanian Ministry of Environment, Waters and Forests, 2019).

This analysis is centred around the environmental Kuznets curve (EKC) framework which depicts the dynamics between environmental degradation and income (Stern, 2018). The EKC hypothesises that initially, as income grows, the quality of the environment decreases with an increase in emissions. However, there is a tipping point and beyond that level of income, the trend is overturned, with more economic growth being associated with a higher environmental quality (MacDermott et al., 2019). Consequently, this implies that environmental degradation is a function of per capita income that is likely to display an inverse U-shaped pattern (Stern, 2018).

Using an autoregressive distributed lag (ARDL) approach, this study finds that CO_2 emissions per capita and GDP per capita showcase an inverted U-shaped relationship between 1968 and 2018 in Romania. The variables move together in the long-run, with 96.2 percent of the deviations from this equilibrium being corrected each year. Nevertheless, an extended EKC analysis that includes the cubic term of GDP per capita uncovers an inverse N-shaped pattern between emissions and income. This is consistent with and reinforced by dividing the main quadratic analysis into two distinct periods (1968-1988 and 1993-2018): communist Romania displays U-shaped dynamics, whereas democratic Romania exhibits an inverse U-shape.

With a turning point corresponding to an emissions peak at a GDP per capita level of $5,700 in 2011 prices in the main model and of $6,550 in the extended cubic model, the results suggest that green growth is a suitable strategy for Romania, as that value of income has long been surpassed. The level of per capita real GDP in Romania has exceeded the tipping points in the 1990s, thus further increases did not result in lower environmental quality as measured by CO_2 emissions. Furthermore, the vector error correction model (VECM) indicates that, in the short-run, there is feedback Granger causality between primary energy consumption and income. Emissions also Granger-cause both income and energy. The long-run tests reveal causality from real income per capita, square real income per capita, and primary energy consumption per capita to CO_2 emissions.

Given the revealed patterns, the present study finds that Romania has not reached a level of income per capita at which emissions start increasing again, when green technology can no longer compensate for the damage inflicted upon the environment (Lazăr et al., 2019). Thus, technological changes can still decrease the level of environmental degradation (Kaufmann et al., 1998; de Bruyn et al., 1998; Talukdar and Meisner, 2001; Bruvoll and Medin, 2003; Shi, 2003; Lantz and Feng, 2006). This effect is potentially driven by improvements in the efficiency of existing technologies or by changed production structures towards more environmentally-friendly activities (Kaufmann et al., 1998; Lantz and Feng, 2006).

The underlying dynamics leading to the aforementioned patterns may have to do with two important developments in the history of Romania. Firstly, the change in the political regime in December 1989, which was associated with a heavy deindustrialization process as for instance 1,256 factories shut down right after the Revolution (DCNews, 2013). A consequence of this could have therefore been the reduced CO_2 emissions.

Secondly, the European Union (EU) accession in 2007 is likely to back the EKC hypothesis, as it provided a framework to reduce emissions, especially through the Emissions Trading Scheme (ETS)

(European Commission, 2022). When the effects of GHGs occur over a long period of time and transcend national borders, incentives for immediate action are not likely to hold (Arrow et al., 1995; Perrings and Ansuategi, 2000; Roca et al., 2001). It is thus somewhat surprising to identify the EKC for CO_2, given the long-term negative consequences that affect future rather than present generations and alter incentives to curb these emissions (Arrow et al., 1995). Nevertheless, this hypothesis is likely to hold for global gases such as CO_2 when the country is subject to a multinational policy initiative (Arrow et al., 1995). Similarly, Jula et al. (2015) also note that emergent economies tend to display lower environmental standards before joining the EU.

Little research has so far been conducted on the EKC in Romania over time. Additionally, the existing studies of the EKC employ a multitude of methodologies resulting in a lack of consensus, so there is room for further research. This paper aims to add to the existing literature an in-depth analysis of the particularities of Romania and its EKC. The panel data studies conducted on Central and Eastern Europe (CEE) countries do not account for heterogeneities in the dynamics between income and environmental degradation, which occur due to different economic systems, geographical locations, and environmental regulations (Lazăr et al., 2019). Consequently, my paper takes the opportunity to fill this gap in the literature.

Compared to other EKC time series analyses of Romania (Jula et al., 2015; Shahbaz et al., 2013), the present study considers a more extended time frame, thus including recent developments. The time frame is then split around the democratic transition of Romania *i.e.* before and after the Romanian Revolution of December 1989. Furthermore, the extended (cubic) EKC is also examined in this paper. Alongside the ARDL analysis which is present in some of the Romanian EKC literature (Shahbaz et al., 2013), the VECM employed to establish the direction of the Granger causality between the variables is a novelty. Furthermore, this study estimates the turning point in the real GDP per capita levels for CO_2 emissions per capita which, to my knowledge, has not been researched until now.

The following section of this thesis describes the EKC framework and reviews the literature. Section three provides descriptive statistics, the fourth one sets out the methodology, whereas the fifth section conducts the empirical analysis and presents the results. The sixth section concludes the thesis.

2 The Environmental Kuznets Curve

2.1 Origins

The EKC indicates the relationship between environmental degradation measured by various indicators and per capita income (Stern, 2018). It stipulates that, as income increases, environmental deterioration also increases given the rise in emissions. Nonetheless, beyond the GDP turning point, the trend reverses (MacDermott et al., 2019; Katsoulakos et al., 2016). Therefore, this implies an inverted U-shaped relationship between environmental degradation and income (Stern, 2018). In this setup, green growth would thus be achievable and policies could focus on ensuring a knowledge-oriented and innovative economy (Katsoulakos et al., 2016; Leadbeater, 2000; Mol, 2001). Otherwise, if increases in GDP lead to lower environmental quality, a degrowth strategy would be more suitable to tackle climate change.

The name of the curve stems from Simon Kuznets, who first outlined the link between income inequality and per capita income as an inverted U pattern (MacDermott et al., 2019). This relationship was further adapted by Grossman and Krueger (1995), who studied the link between per capita income and four environmental indicators: urban air pollution, oxygen regime in river basins, fecal and heavy metal contamination of river basins. Similar to Kuznets, the authors found that there is some degree of environmental deterioration up to a certain point, followed by a subsequent phase of environmental improvement brought by economic growth (Grossman and Krueger, 1995). These results modified the original curve and led to the creation of what is known as the EKC (MacDermott et al., 2019), as depicted in Figure 1.

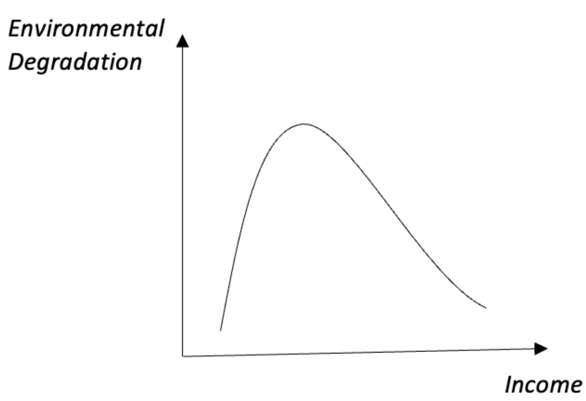

Figure 1: Environmental Kuznets Curve. Adapted from Panayotou (1993)

In the initial stage of development of a country, people are intuitively more keen on economic aspects such as jobs and income, rather than the quality of the environment because the society is not rich enough to be concerned about environmental deterioration or pollution abatement (Jula et al., 2015). As increased income levels imply higher demand for environmental quality, societies that are more well-off have incentives to use more resources for increasing environmental quality to the extent that the degradation impacts their lives (Roca et al., 2001). A high environmental quality requires 'immediate and continuous expenditures', while benefits are mostly felt in the future (Ng and Wang, 1993).

Furthermore, as outlined by MacDermott et al. (2019), the idea is that countries with lower levels of income often lack industrialisation and focus on agriculture. During their development process, the industrialization and the inherent dirty industries contribute to decreasing the environmental quality until a certain point when preferences change towards cleaner technologies (MacDermott et al., 2019). After that point, growth results in service-oriented economies since countries invest resources to 'provide cleaner water and air' such that the environmental degradation decreases and even regresses (MacDermott et al., 2019). This evolution can be summarised as follows:

$$\textit{Clean agriculture} -> \textit{Dirty industries} -> \textit{Cleaner industries} -> \textit{Services}$$

Along these lines, more recent studies extended the EKC theory to a cubic relationship, which depicts an N-shaped area between environmental degradation and per capita income (Figure 2). This effect was theorized by Grossman and Krueger (1995), occurring when the society moves away from industry towards knowledge and services. The extended relationship implies that, as per capita income grows, the environmental quality starts decreasing again after a certain time (Lazăr et al., 2019). These dynamics were explored in studies by Yang et al. (2015), Dogan and Seker (2016), Millimet et al. (2003), to name but a few, finding distinct results for different countries and pollutants. An explanation for this N-shaped link could be that the economic impact on the environment can no longer be compensated by the technological effect (Lazăr et al., 2019).

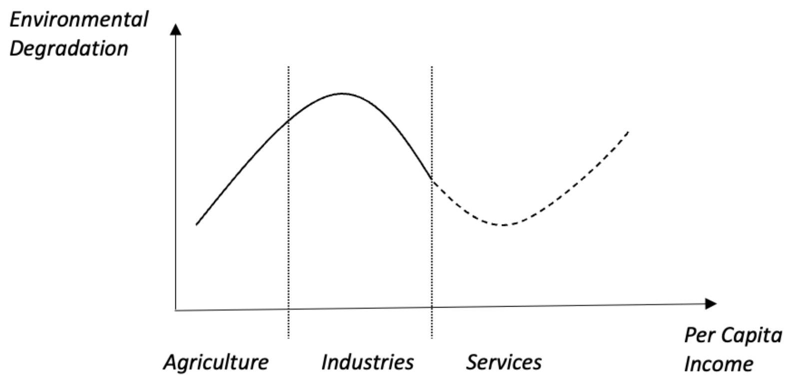

Figure 2: Extended Environmental Kuznets Curve. Adapted from Lazăr et al. (2019) and Grossman and Krueger (1995)

2.2 Environmental Degradation

GHGs are like a blanket which warms the Earth through slowing the rate at which the energy goes into space. They differ from each other due to their ability to absorb energy as well as through their lifetime in the atmosphere (United States Environmental Protection Agency, 2022a). In this regard, environmental deterioration depends on both the current flows of emissions and on prior environmental pressures which ‘affect the capacity of assimilation and the resilience of ecosystems’ (Roca et al., 2001). This aspect is even more relevant when imbalances and irreversible changes occur (Arrow et al., 1995).

Depending on what GHG is used for the study of environmental degradation, the literature finds different dynamics, such as an inverted U-shaped one as per the EKC. This kind of relationship is often uncovered for gases that have more short-term costs such as sulfur, particulates and fecal coliforms. However, CO_2 is a global gas, with more dispersed costs and negative effects that transcend borders and occur mostly in the long-run (Arrow et al., 1995). In this case, the EKC hypothesis is less likely to hold, with CO_2 emissions being rather an increasing function of income (Shafik and Bandyopadhyay, 1992; Socolow et al., 1997; Arrow et al., 1995).

It is often argued that prosperity is one of the primary drivers of CO_2 emissions, so governmental policies and technological choices play an important role (Ritchie and Roser, 2020). Therefore, investigating the nexus between environmental pollution and income is important, as it informs whether economic growth is detrimental to the environment. If that were the case, then a degrowth strategy might be preferred to tackle emissions and achieve neutrality by 2050. If however the EKC is applicable, then increases in income do not necessarily lead to more pollution following the tipping point, as environmental policies can achieve 'de-linking' income growth and environmental degradation (Roca et al., 2001).

Environmental deterioration can be measured in a number of ways. According to Bednar-Friedl and Getzner (2003), there are four indicators that help achieve that: emissions per capita, pollution intensity, ambient levels of pollution, and total emissions. This paper focuses on measuring environmental degradation through CO_2 emissions per capita, as this gas is the 'primary driver of climate change' (Ritchie and Roser, 2020). Being the 'most important man-made greenhouse gas', it has a lifetime of up to thousands of years (Clark, 2012).

Given the Industrial Revolution, the stock of CO_2 emissions in the atmosphere has been increasing due to the burning of fossil fuels and deforestation (Villarreal, 2011; IPCC, 2022). Once in the atmosphere, this GHG can influence the climate system for thousands of years (Clark, 2012). The abatement cost can also be more substantial compared to other pollutants such as sulfur dioxide (SO_2) which has a low cost (Roca et al., 2001).

However, the cost/benefit analysis for this gas has a hidden side, given its large lifetime (Arrow et al., 1995). This is in contrast to other GHGs such as methane which persists in the atmosphere less than 15 years and is eliminated by chemical reaction (Clark, 2012). Additionally, other gases have immediate negative consequences like acid rain due to SO_2 and nitrogen oxides released into the air (United States Environmental Protection Agency, 2022b).

Therefore, when considering the link between the economy and the environment, authors state that an inverse U-shaped relationship can occur for local GHGs with short-term consequences such as SO_2, due to their quick clear negative impacts upon the environment and human health (Roca et al., 2001). However, for global gases such as CO_2 with higher abatement costs, the environmental quality would not increase with income, given free riding and the lack of immediate incentives to curb emissions (Selden and Song, 1994; Arrow et al., 1995; Cole et al., 1997; Roca et al., 2001). These GHGs are likely to generate emissions that rise monotonically with income, or have tipping points occurring at high per capita incomes unless they are subject to a multilateral policy initiative (Arrow et al., 1995). This viewpoint is also shared by Bernard et al. (2015), who identify lower income turning points for SO_2 compared to CO_2.

2.3 Empirical Evidence: A Lack of Consensus

2.3.1 Specifications

Different specifications have been reported in the literature regarding the relation between GDP and CO_2 emissions: linear, quadratic and cubic (Bednar-Friedl and Getzner, 2003). The linear pattern is investigated in time series studies such as Akbostancı et al. (2009) (Turkey), He and Richard (2010) (Canada) or Kunnas and Myllyntaus (2007) (Finland) that find a monotonically increasing relationship.

A number of researchers examining the quadratic dynamics revealed an inverse U-shaped relationship (Shahbaz et al., 2013; Kasman and Duman, 2015; Lazăr et al., 2019), consistent with the EKC hypothesis. Others however discussed the existence of U-shaped links (Stern, 2018). In addition, including the cubic term for income in the analysis revealed both an N-shaped relationship between GDP and CO_2 emissions (Bednar-Friedl and Getzner, 2003; Simionescu, 2021; Lazăr et al., 2019), as well as an inverted N-shape (Jula et al., 2015; Lazăr et al., 2019).

Panel data research also uncovers new insights. Yang et al. (2015) studied the nexus between CO_2 emissions and economic growth for 67 countries, between 1971 and 2010. The authors highlight that there is no universal model that can fit all countries, with four models being the most widespread ones: inverted N-shaped, M-shaped, inverted U-shaped and monotonic, with developed countries typically following the first two ones and the developing ones displaying the inverted N and inverted U shapes, as well as the monotonically increasing relationships (Yang et al., 2015). Therefore, analysing the countries separately yields much diversity in terms of the nature of the dynamics (Lazăr et al., 2019). Similarly, two panel studies on CEE countries by Kasman and Duman (2015) and Lazăr et al. (2019) reveal different shapes of the relationship for different countries, with the EKC hypothesis being confirmed only in some.

The aforementioned lack of consensus in the literature stems from a variety of reasons, as identified by Hill and Magnani (2002). For instance, authors choose to study different environmental and economic variables such as SO_x, CO_2, volatile organic compounds, and CO (Yang et al., 2015), to name but a few. Furthermore, the different data collection strategies and the included variables could yield divergent results, as in the case of the EKC study in Malaysia (Lau et al., 2014; Azlina et al., 2014). The employed empirical methods are also distinguishable, such as ARDL, the Johansen approach, the generalized method of moments, or structural time series models (Yang et al., 2015). Lastly, it is also a case of individual country characteristics generating different patterns between environmental degradation and income (Lazăr et al., 2019), which cannot be generalized to other contexts (Socolow et al., 1997; Roca et al., 2001; Simionescu, 2021).

As per Bednar-Friedl and Getzner (2003), the general functional form that does not account for other explanatory variables is:

$$CO_{2,\mathrm{t}} = \beta_0 + \beta_1 Y_\mathrm{t} + \beta_2 {Y^2}_\mathrm{t} + \beta_3 {Y^3}_\mathrm{t} + \varepsilon_\mathrm{t} \tag{1}$$

The linear relationship implies that $\beta_1>0$ whereas $\beta_2=\beta_3=0$. The EKC hypothesis of an inverse U-shaped curve has $\beta_1>0$, $\beta_2<0$ and $\beta_3=0$. Lastly, the extended EKC which involves an N shape suggests $\beta_1>0$, $\beta_2<0$ and $\beta_3>0$ (Bednar-Friedl and Getzner, 2003).

The model can further be augmented with a vector of control variables, consistent with the existing literature such as the study by Egli (2004).

$$CO_{2,\mathrm{t}} = \beta_0 + \beta_1 Y_\mathrm{t} + \beta_2 {Y^2}_\mathrm{t} + \beta_3 {Y^3}_\mathrm{t} + \theta V_\mathrm{t} + \varepsilon_\mathrm{t} \tag{2}$$

To account for the communist history of the countries, Lazăr et al. (2019) include as control variables energy consumption and economic freedom, capturing the transition from a former centrally-planned economy to a market economy. For the same reason, my thesis also controls for primary energy consumption per capita, as well as for the change in political regime.

Other studies controlled for the share of nuclear power and coal within the entire energy system, as per Roca et al. (2001). However, this is not the case for Romania, as depicted in Figure 3, because no sector had a strong and sudden growth. Consequently, no such controls were included in this analysis. Since 1990, the shares of coal and natural gas have been decreasing in Romania, whereas renewable energy such as wind and solar were well below 3 percent in 2015 and had even slightly decreased by 2019. This proves that, from a policy perspective, there is still room for improvement in this regard.

Nevertheless, in the current context, Romania is looking for ways to achieve more energy independence, thus planning to reopen coal power plants (Barna, 2022), which is likely to increase the share of coal as a source of CO_2 emissions in the future.

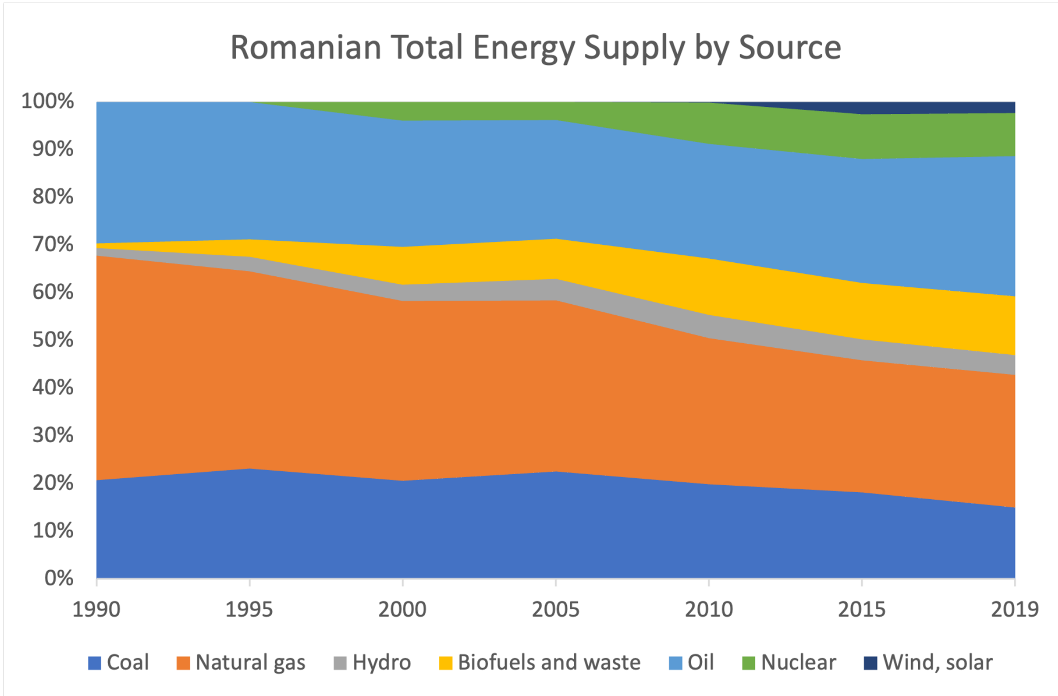

Figure 3: Romanian total energy supply. Author's calculations based on data from IEA (2022)

2.3.2 The Turning Point

The turning point denotes the level of income beyond which environmental degradation starts decreasing with economic growth, as depicted in Figure 1. Afterwards, the individuals' willingness to pay for environmental protection increases as they assign more value to it (Jula et al., 2015). This point is calculated in the literature for a great number of pollutants. For instance, estimations take GDP per capita values such as $3,670 (Shafik, 1994) or $4,053 (Grossman and Krueger, 1995) for SO_2, and $8,000 for SPM and SO_2 (Selden and Song, 1994). Lazăr et al. (2019) identify CO_2 turning points taking values between $9,9677 (Hungary) and $24,875 (Czechia) for peaks, as well as $9,253 (Bulgaria) and $21,668 (Estonia) for troughs.

It is highly relevant to note that the turning points have a wide variation (Hill and Magnani, 2002; López-Menéndez et al., 2014; Sulemana et al., 2017; Lazăr et al., 2019). The level of economic development, the variables and the model used are among the reasons why the turning point estimations differ in the literature (Lazăr et al., 2019). For example, the higher turning point by Selden and Song (1994) is theorized to occur given the use of aggregate emissions flow rather than the urban air pollution stocks from Shafik (1994), thus biasing downward the estimates of the turning point (Shafik, 1994).

The downturn in the relation between environmental degradation and income is hypothesized to be driven by factors like positive income elasticities for the quality of the environment, changes in production and consumption, as well as the levels of environmental awareness and openness in the political systems (Selden and Song, 1994). This is also consistent with the concept of dematerialization: in the 1990s, environmental economists theorized that developed capitalist economies tend to 'dematerialize', increasing energy efficiency and decoupling economic growth from the use of energy and materials and from waste materials, such that economic growth does not harm the environment (Katsoulakos et al., 2016).

All these differences imply that new contributions are needed regarding turning points. This is specifically relevant for the case of Romania and CO_2 emissions as, to the author's knowledge, no such point has been estimated so far. Therefore, the present thesis aims to fill this gap.

2.3.3 The Romanian Case Study

Shahbaz et al. (2013) investigated the relation between economic growth, CO_2 emissions and energy consumption between 1980 and 2010 in Romania. Using ARDL bounds testing, the authors lend support to the EKC hypothesis, while also revealing that energy consumption is a significant contributor to pollution. Furthermore, the study identifies how the democratic regime contributed to reducing CO_2 emissions through various economic policies that allocated more resources towards environmental projects (Shahbaz et al., 2013). While the approach of my study has the ARDL methodology in common with Shahbaz et al. (2013), this thesis augments the analysis through expanding the timeline of the analysis, dividing the time frame into two distinct periods, analysing the cubic EKC hypothesis, as well as through using Granger causality.

Moreover, Jula et al. (2015) also inspected the relation between economic growth measured by income per capita and environmental quality as CO_2 emissions per capita in Romania between 1960 and 2010. Using the Least Squares with Breakpoints approach and choosing 1990 as the breaking point, the authors find evidence of an inverse N-shaped relationship between economic growth and CO_2 emissions, revealing a long-run connection.

A more recent study covering a panel of countries examines GHG emissions in CEE countries including Romania between 1990 and 2019 (Simionescu, 2021). Using the EKC, the analysis reveals the inverted N-shape between both GDP and GHG emissions, and value added in agriculture and emissions.

Lazăr et al. (2019) also test the EKC hypothesis separately for each CEE country. For Romania, the authors find that the link between income and CO_2 emissions is not statistically significant.

2.4 Policy Implications

Early research outlines how economic growth generates needs and wants, making the society more complex and transient (Pezzey, 1992). In this context, people consume and demand more as they compare the standards of living against countries with higher ones (Pezzey, 1992). Nonetheless, Ng and Wang (1993) state that increases in income lead to higher environmental degradation. This view is not consistent with the EKC hypothesis and it thus suggests that income growth could be welfare-reducing despite the resulting higher per capita consumption in developed countries. In the authors' view, the welfare-reducing channel revolves around the social choice. This choice is centred on material production, given myopic voters and 'the tendency of politicians to respond more to the rat-race for material growth' which is non-welfare optimizing (Ng and Wang, 1993).

Therefore, there is scope for authorities to ensure that any negative environmental consequences of income growth are mitigated, as the welfare-enhancing measures such as environmental policies and scientific advances are more important than and enabled by GDP growth (Ng and Wang, 1993). These measures could thus lead to the occurrence of the EKC, as they would compensate the negative impacts of economic growth up to a certain point (Lazăr et al., 2019). Consequently, the implication of an increase in income can be higher environmental quality in the presence of enhanced environmental protection introduced by strong institutions and policies (Ng and Wang, 1993). This holds as pollution is a negative externality which can be internalized through policy interventions, with powerful institutions required to act upon this matter (Jula et al., 2015).

While such policies improve the quality of life, in some cases they are only adopted at the turning point when the society reaches a certain level of development (Jula et al., 2015), as the society shifts its focus from industry to services (Grossman and Krueger, 1995). Along these lines, income is used as a proxy for living standards (Jula et al., 2015). When the former overpasses a certain level, individuals start demanding better environmental quality, treating this as a normal good (Jula et al., 2015).

However, even if the EKC hypothesis holds, that would not imply a free pass for harming the environment up until the turning point and postponing environmental action through taking for granted a future decrease in emissions (Kunnas and Myllyntaus, 2007; Akbostancı et al., 2009). While some development paths manage to decouple CO_2 emissions and income, this is not a given and adequate institutions are still needed to safeguard environmental quality (Arrow et al., 1995; Socolow et al., 1997). For instance, Akbostancı et al. (2009) derive from an EKC analysis on Turkey that actions cannot wait until per capita income or the awareness about the environment increase, suggesting that pollution does not vanish with income. This implies a strong need for national and local policies to address the environmental degradation and combat pollution, regardless of the income levels (Akbostancı et al., 2009).

A highly important policy implication also stems from the ETS that Romania joined through its accession to the EU in 2007. By 2015, Romania had collected approximately 260 million euro in ETS revenues, with 71 percent of these revenues being invested in climate finance projects (World Bank, 2015). Nevertheless, policy briefs argue that Romania is lagging behind especially in sectors not included in the ETS, which might thus hinder the ability of the country to reach the net zero goals of 2050 (Catuti, 2022). In this regard, the revenues from the ETS auctioning are projected to reach four billion euro by 2030 and their efficient use could help reach climate targets and increase the ambition in the sectors that are still not covered by the regulation (World Bank, 2015), such as transport, building or agriculture (Catuti, 2022).

Lastly, while an inverted U-shaped relationship between income and emissions could be observed, it is recommended to take this analysis one step further and investigate the cubic relationship between the variables (Allard et al., 2018). This analysis should be conducted to rule out a second turning point in the level of GDP per capita, after which environmental degradation would start increasing again, as per the extended EKC in Figure 2. Should such an N-shaped relationship be present, that could serve as support in favour of a degrowth strategy, given that technological progress can no longer compensate and curb emissions after a certain income threshold (Lazăr et al., 2019).

3 Descriptive Statistics

3.1 Data

Table 1 describes the variables used in the present study. Data on carbon dioxide (CO_2) emissions and primary energy consumption (EN) originate from Ritchie et al. (2022), with gross domestic product (GDP) being added from the Maddison Project Database 2020 (Bolt and van Zanden, 2020) and aggregated by Ritchie et al. (2022). The variables are in per capita terms and I transformed them in natural logarithms to obtain more appropriate results compared to the linear form (Shahbaz et al., 2010), while enabling a direct interpretation of the elasticities (Shahbaz et al., 2010) and smoothing the data (Touitou, 2021). My thesis studies Romania between 1968 and 2018.

Table 1: Summary Statistics

	Mean	**Count**	**Variance**	**Min**	**Max**	**Unit**
lnCO_2	1.743	51	.094	1.312	2.208	tonnes/capita
lnGDP	8.894	51	.312	7.942	9.893	$/capita, 2011 prices
lnEN	10.073	51	.043	9.775	10.340	kilowatt-hours/capita/year

3.2 Evolution of CO_2 Emissions and GDP

Figure 4 depicts the time series of the natural logarithm of CO_2 emissions for Romania between 1968 and 2018. The data could be split into two different periods with different trajectories: in the first one, during the ruling of the Communist Party, emissions had been steadily increasing, reaching their peak right before the fall of the regime, at the end of the 1980s. At the beginning of the 1990s, the emissions started decreasing. This second period was marked by occasional increases, but they never reached pre-1989 levels again.

These periods coincide with significant events in Romania's history: until 1989, Romania had been a communist country, ruled by dictator Nicolae Ceaușescu. That time was marked by 'excessive industrialization', given the centrally-planned economy (Foarfă, 2019). With almost four million people working in the industrial sector (Foarfă, 2019), this was the primary occupation, albeit there were doubts regarding the way of allocating resources, as efficiency was not at the forefront of production processes before 1989 (Digi24, 2019). For instance, a thermal plant in Anina was built in 1976 and thousands were relocated there to work in what was known as a workers' colony. After 8,400 hours of operation, the plant was closed in 1988 (Moldovan, 2015).

As it moved from an industrial economy to a service-based one (Săgeată et al., 2021), Romania's transition towards the free market led to the collapse of some industries (Moldovan, 2015). Compared to 1989, total GHG emissions from the energy sector decreased by more than 65 percent in 2017, as the market economy transition was associated with a sudden and steep decline in the demand for heat and power by power plants (Romanian Ministry of Environment, Waters and Forests, 2019). In 2007, Romania joined the EU. Following accession, emissions started declining again, remaining

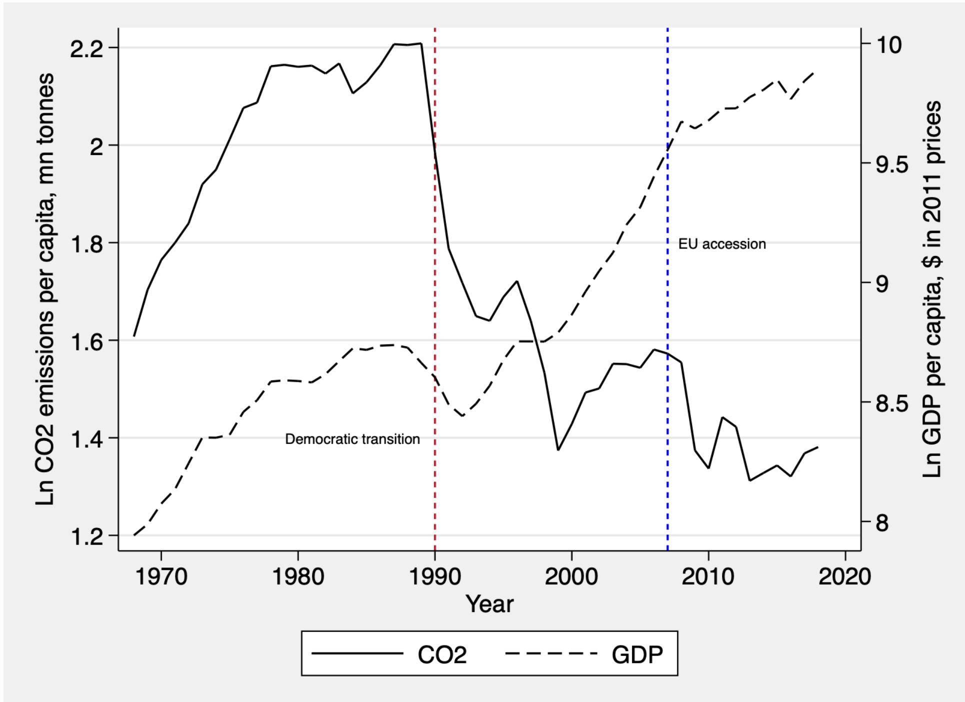

Figure 4: CO_2 emissions and real GDP per capita between 1968 and 2018

below accession levels at the end of 2018, as seen in Figure 4. Therefore, the past 50 years have been marked by complex processes of large-scale industrialization, followed by deindustrialization, and reindustrialization/tertiarization (Săgeată et al., 2021).

In terms of GDP, Figure 4 also outlines the evolution of real income per capita during the analyzed time frame. While income decreased right after the fall of communism, the decrease was not as dramatic as the emissions one. Following that period of political instability at the end of the 20$^{\text{th}}$ century, income per capita has been steadily increasing, almost reaching 20,000$ (in 2011 prices) per capita in 2018.

Figure 5 depicts a comparison between the sources of CO_2 emissions by sector from 1990 and 2018, using data from Hannah Ritchie and Rosado (2020). Electricity and heat account for most of the emissions both in 1990 and 2018, with 46 percent and 40 percent respectively. Manufacturing used to be the second largest share, but was overtaken by transport at the end of the time frame. Yet emissions from manufacturing still account for almost one fifth of the total CO_2 emissions.

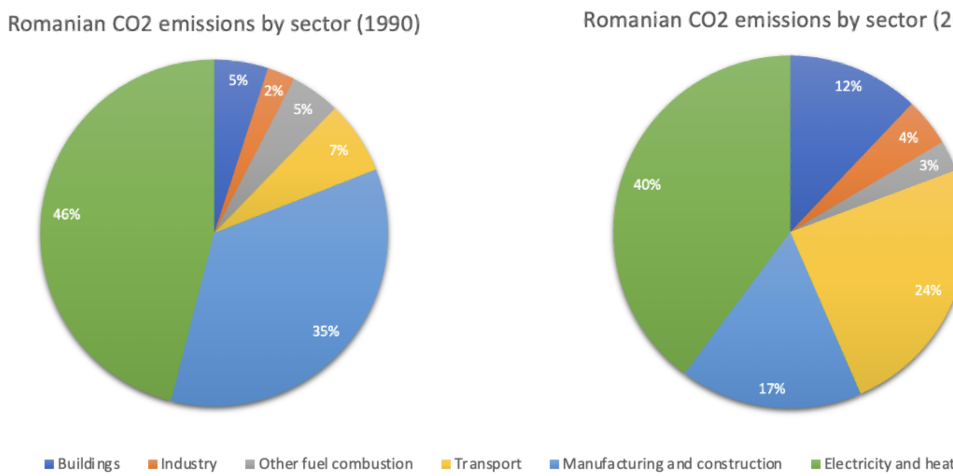

Figure 5: Comparison of emissions by sector. Author's calculations

4 Methodology

4.1 Specification

The log-log functional specification of the long-run relationship between per capita emissions and income is augmented with per capita primary energy consumption and a vector of exogenous variables, being expressed as follows:

$$lnCO_{2,\mathrm{t}} = \beta_0 + \beta_1 lnY_t + \beta_2 lnY^2{}_t + \beta_3 lnEN_t + \theta V_{\mathrm{t}} + \varepsilon_{\mathrm{t}} \tag{3}$$

$CO_{2,t}$ stands for carbon dioxide emissions per capita in year t, Y_{t} represents $\mathrm{GDP_t}$ per capita in year t, in \$ and 2011 prices, whereas EN_{t} indicates the primary energy consumption per capita in year t, measured in kilowatt-hours. $\mathrm{V_t}$ is a vector of exogenous variables, and ε_{t} is the error term. As per the EKC hypothesis, the quadratic relationship is expected to have $\beta_1 > 0$, whereas $\beta_2 < 0$ (Bednar-Friedl and Getzner, 2003). Furthermore, β_3 is expected to be positive, as more energy could lead to higher emissions, given that energy is the main source of CO_2 emissions in Romania (Romanian Ministry of Environment, Waters and Forests, 2019). The relationship between environmental degradation and income was first inspected without adding any other control variables to the model (apart from the democracy dummy for year 1990) and only including Y and Y^2. However, this model failed the Jarque-Bera normality test (2011) at all significance levels, while also being unstable as shown by the recursive residuals and namely the plot of their cumulative sum of squares (Baum, 2000). Therefore, more variables were included in the final model to obtain robust results.

Empirically, one of the methods of investigating the aforementioned relationship is based on time series data for a single region (Yang et al., 2015). Since correlation does not imply causation, a time series analysis is employed to avoid spurious regressions (Roca et al., 2001). The potential issues when using time series data are related to non-stationarity as well as serial correlation, given the limited sample size (Egli, 2004), which consists of 51 observations in the current study.

There are several methods for determining whether cointegration exists between the variables: the Granger (1981), the Engle and Granger (1987), and the Johansen and Juselius (1990) approaches as well as the autoregressive distributed lag (ARDL) bounds testing approach to cointegration (Pesaran and Shin, 1999; Pesaran et al., 2001). These became suitable solutions to non-stationarity and enable the assessment of the long-run relationship, as well as the reparameterization of the time series to the error correction model (ECM) (Nkoro and Uko, 2016). In this way, the short-run adjustments are integrated in the long-run equilibrium (Menegaki, 2019).

This paper employs ARDL bounds testing to study the long-run relationship between the variables, due to a number of advantages that will further be discussed. ARDL is a dynamic specification that uses the lagged dependent variable as well as lagged and contemporaneous values of the explanatory variables through which short-run effects and the long-run equilibrium can be estimated (Ghosh, 2010). The former are estimated directly through the coefficients of the regressors, whereas the latter indirectly, through the coefficient of the lagged error correction term (ECT) (Ghosh, 2010). The 'distributed lag' part of the name is given by the inclusion of the unrestricted lags of the explanatory variables in the regression (Nkoro and Uko, 2016).

An advantage of this approach over the conventional cointegration ones such as Engle and Granger (1987) and Gregory and Hansen (1996) is that the bounds tests results are robust when the sample size is small (Pesaran and Shin, 1999; Haug, 2002; Ghosh, 2010; Javaid and Zulfiqar, 2017; Shahbaz et al., 2010; Zhang, 2021) and less than 60 observations (Liew, 2004), as in this thesis. In addition, it is not necessary for the variables to have the same order of integration when using the ARDL approach, as it is possible for them to be a mix of I(0) and I(1), unlike the Granger (1981) and the Engle and Granger (1987) approaches which require the same order of integration (Nkoro and Uko, 2016). The last step in this analysis is to investigate the direction of causality using the Granger approach within the vector error correction model (VECM), similar to the methodology of Ozturk and Acaravci (2010).

Equation 3 is consequently expressed in the following ARDL form:

$$\begin{aligned}\Delta lnCO_{2,\mathrm{t}} =& \beta_0 + \beta_1 lnCO_{2,t-1} + \beta_2 lnY_{t-1} + \beta_3 lnY^2_{t-1} + \beta_4 lnEN_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta lnCO_{2,t-i} + \\ &+ \sum_{j=0}^{q} \gamma_j \Delta lnY_{t-j} + \sum_{k=0}^{r} \gamma_k \Delta lnY^2_{t-k} + \sum_{l=0}^{s} \gamma_l \Delta lnEN_{t-l} + \theta V_t + \mu_{\mathrm{t}}\end{aligned} \tag{4}$$

where Δ is the first difference operator, β_0 denotes the drift, β_1, β_2, β_3 and β_4 are the coefficients for the long-run dynamics. The lagged variables of ΔCO_2, ΔY, ΔY^2 and ΔEN are used to identify the short-run relationship. V_t is a vector of exogenous variables. These are a political dummy variable taking the value of 1 for the years after 1990, accounting for the regime change and 0 otherwise, as well as two dummies for years 1977 and 2006, due to important events that occurred at those times. To check this intuition, three Chow structural break tests (Chow, 1960) have been conducted to establish whether there was indeed a structural break in each of the selected three years. The statistic was calculated based on the following formula:

$$F_{Chow} = \frac{[RSS - (RSS_1 + RSS_2)]/k}{(RSS_1 + RSS_2)/(T - 2k)} \tag{5}$$

where RSS denotes the residual sum of squares of the regression covering the entire period, RSS_1 represents the residual sum of squares for the period up until the break point and RSS_2 denotes the time frame afterwards. k stands for the number of parameters and T is the number of observations.

Should the variables exhibit a long-run relationship, Equation (4) will further be fit into an ECM (Granger, 1986; Engle and Granger, 1987; Fawson and Chang, 1994), becoming the focus of this analysis.

4.2 Order of Integration

Before conducting the ARDL analysis, some prerequisites have to be fulfilled to be able to implement this method. In the case that variables are integrated of order two, the F-statistics of the bounds test are not valid, as they are based on the assumption that the variables are either stationary, are integrated of order one, or a mix between the two (Ouattara, 2006; Chigusiwa et al., 2011; Nkoro and Uko, 2016). Therefore, to ensure that all variables are either stationary in levels or in first differences, the present paper investigates the order of integration of the variables before conducting the ARDL analysis, following the framework of Nkoro and Uko (2016). Thus, unit root testing is carried out before the bound F-test to inspect the stationarity of the variables and ensure that the order of integration is maximum 1.

Consistent with the literature (Roca et al., 2001; Shahbaz et al., 2010; Javaid and Zulfiqar, 2017; Cañal-Fernández and Fernández, 2018; Simionescu, 2021), this paper uses the augmented Dikey-Fuller unit root test (ADF) (Dickey and Fuller, 1979) to examine the order of integration of the variables and rule out any variables being I(2). The test is conducted as follows:

$$\Delta lnCO_{2,t} = \alpha_0 + \alpha_1 t + \alpha_2 lnCO_{2,t-1} + \sum_{i=1}^{z} \lambda_i \Delta lnCO_{2,t-i} + u_t \tag{6}$$

In this equation, α_0 denotes the constant, t stands for the trend which is included when necessary, following visual inspection and reasoning, while u denotes the residuals. λ_i is the coefficient of the lagged difference. The number of lags is selected using the Akaike information criterion (AIC).

The coefficient of interest is α_2, as the null hypothesis of $\alpha_2 \geq 0$ implies a unit root process. This is tested against the alternative hypothesis of $\alpha_2 < 0$, implying it is a stationary process (Dickey and Fuller, 1979). For the ARDL method, it does not matter whether the variables are I(0) (*i.e.* stationary) or I(1) (*i.e.* with one unit root), or a combination of the two and the bounds testing for cointegration can be conducted as long as no variable is integrated of a higher order than one (Nkoro and Uko, 2016). The test is conducted for all the variables. Given that the data is annual, the number of lags is small, typically 1 or 2, so that not many degrees of freedom are lost (Wooldridge, 2009).

Additionally, the modified Dickey Fuller unit root test DF-GLS (Elliott et al., 1996) is also conducted. While similar to the ADF, this test is conducted on a transformed time series through a generalized least squares regression (StataCorp, 2022a) It has the same purpose as the previous test and its aim is to ensure the robustness of results.

4.3 Cointegration

In univariate models, a solution to the issue of non-stationarity is first differencing the data to obtain a stationary series (Enders, 2014). In the context of multivariate analyses, non-stationarity is less straight-forward. Nevertheless, it is possible to find a stationary linear combination of integrated variables, which are therefore defined as 'cointegrated', implying the possibility to simultaneously model both short-run and long-run dynamics (Enders, 2014).

As per Ozturk and Acaravci (2010), the null hypothesis of no cointegration is based on the joint

F-statistic or Wald statistic testing $\beta_1 = \beta_2 = \beta_3 = \beta_4 = 0$ in the ARDL model, against the alternative $\beta_1 \neq \beta_2 \neq \beta_3 \neq \beta_4 \neq 0$. Two sets of critical values were calculated by Pesaran et al. (2001): the upper bound I(1) and the lower bound I(0) ones. The resulting F-statistic is compared to the critical values in order to asses whether the variables exhibit a long-run relationship, as cointegration enables identifying the existence of a steady state equilibrium between the variables (Nkoro and Uko, 2016). This occurs as some series, despite each moving considerably across time, are bound together by equilibrium forces (Kripfganz and Schneider, 2018).

The discussion provided by Nkoro and Uko (2016) highlights the importance of cointegration. Time series data that diverge from their mean over time are non-stationary and the traditional estimations which assume that series are stationary or stationary around a deterministic trend can lead to misleading inferences and spurious regressions (Egli, 2004). Consequently, cointegration became an 'over-riding requirement for any economic model using non-stationary time series data' (Nkoro and Uko, 2016).

If the F-statistic is above the I(1) critical value, the null hypothesis is rejected, thus indicating cointegration. If it is below the lower bound, the null hypothesis cannot be rejected. Should the F-statistic fall between the critical bounds, the result is deemed to be inconclusive. Consistent with the literature, the optimal number of lags of the differenced regressors is chosen based on the AIC, given its power and suitability for small sample sizes (Ouattara, 2006; Shahbaz et al., 2013).

4.4 Error Correction Model

As mentioned, the lack of cointegration between non-stationary variables can lead to spurious regressions (Egli, 2004). According to Nkoro and Uko (2016), a solution for non-stationarity is first-differencing the data, but this method only provides short-run dynamics (through showing the difference between two consecutive periods) and does not inform of the long-run. Therefore, given cointegration, reparameterizing into the unrestricted ECM becomes 'imperative', as this method includes both short-run and long-run behaviours of the variables (Nkoro and Uko, 2016). The former show the effects of the regressors on the dependent variable, whereas the latter take into consideration the short-run fluctuations (Kripfganz and Schneider, 2018). For instance, if there is a change in an independent variable, the long-run relationship does no longer hold. As a consequence, the regressand reacts to this deviation and its short-run reaction occurs via the ECT, therefore being measured by the adjustment speed coefficient of it (Kripfganz and Schneider, 2018).

Consequently, the following ECM is the main specification of this paper:

$$\begin{aligned}\Delta lnCO_{2,\mathrm{t}} =&\beta_0 + \beta_1 lnCO_{2,t-1} + \beta_2 lnY_{t-1} + \beta_3 lnY_{t-1}^2 + \beta_4 lnEN_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta lnCO_{2,t-i} + \\ &+ \sum_{j=0}^{q} \gamma_j \Delta lnY_{t-j} + \sum_{k=0}^{r} \gamma_k \Delta lnY_{t-k}^2 + \sum_{l=0}^{s} \gamma_l \Delta lnEN_{t-l} + \zeta ECT_{t-1} + \theta V_t + \xi_{\mathrm{t}}\end{aligned} \tag{7}$$

β_0 is the constant, β_1, β_2, β_3 and β_4 correspond to the long-run dynamics, whereas γ_i, γ_j, γ_k, γ_l characterise the short-run ones. ζ is the coefficient of the lagged ECT. This indicates how fast the CO_2 emissions adjust to changes in the regressors before converging to the long-run equilibrium (Menegaki, 2019). The coefficient of the ECT is expected to be negative and statistically significant in order to have convergence over the long-run (Menegaki, 2019) and prove the stability of the long-run relationship (Banerjee et al., 1998; Shahbaz et al., 2013). The significance is established through a t-test (Shahbaz et al., 2010). Lastly, V_t is the vector of the three dummy variables corresponding to years 1977, 1990 and 2006.

4.5 Granger Causality

Knowing the direction of causality hinders erroneous conclusions regarding the nature of the interaction between emissions and income (Goulder and Schneider, 1999; Coondoo and Dinda, 2002). Therefore, this subsection presents the methodology of the Granger causality analysis. The estimated VECM equations are as follows:

$$\begin{aligned}\Delta lnCO_{2,t} =&\alpha_1 + \rho_1 ECT_{t-1} + \sum_{i=1}^{j} \tau_{1i} \Delta lnCO_{2,t-i} + \sum_{i=1}^{j} \phi_{1i} \Delta lnY_{t-i} + \sum_{i=1}^{j} \chi_{1i} \Delta lnY_{t-i}^2 + \\ &+ \sum_{i=1}^{j} \psi_{1i} \Delta lnEN_{t-i} + \omega_1 V_t + \varepsilon_{1t}\end{aligned} \tag{8}$$

$$\begin{aligned}\Delta lnY_t =&\alpha_2 + \rho_2 ECT_{t-1} + \sum_{i=1}^{j} \tau_{2i} \Delta lnCO_{2,t-i} + \sum_{i=1}^{j} \phi_{2i} \Delta lnY_{t-i} + \sum_{i=1}^{j} \chi_{2i} \Delta lnY_{t-i}^2 + \\ &+ \sum_{i=1}^{j} \psi_{2i} \Delta lnEN_{t-i} + \omega_2 V_t + \varepsilon_{2t}\end{aligned} \tag{9}$$

$$
\begin{aligned}
\Delta lnY_t^2 =& \alpha_3 + \rho_3 ECT_{t-1} + \sum_{i=1}^{j} \tau_{3i} \Delta lnCO_{2,t-i} + \sum_{i=1}^{j} \phi_{3i} \Delta lnY_{t-i} + \sum_{i=1}^{j} \chi_{3i} \Delta lnY_{t-i}^2 + \\
&+ \sum_{i=1}^{j} \psi_{3i} \Delta lnEN_{t-i} + \omega_3 V_t + \varepsilon_{3t}
\end{aligned} \tag{10}
$$

$$
\begin{aligned}
\Delta lnEN_t =& \alpha_4 + \rho_4 ECT_{t-1} + \sum_{i=1}^{j} \tau_{4i} \Delta lnCO_{2,t-i} + \sum_{i=1}^{j} \phi_{4i} \Delta lnY_{t-i} + \sum_{i=1}^{j} \chi_{4i} \Delta lnY_{t-i}^2 + \\
&+ \sum_{i=1}^{j} \psi_{4i} \Delta lnEN_{t-i} + \omega_4 V_t + \varepsilon_{4t}
\end{aligned} \tag{11}
$$

The lagged ECT is obtained from the long-run equation and thus the significance of its coefficient indicates long-run Granger causality among the co-integrated variables (Maddala and Kim, 1999; Lütkepohl, 2005; Tsegaye and Yoo, 2014; Usman et al., 2019). As per Johansen and Juselius (1990), the ECT informs of the speed of adjustment of the regressand from its long-run equilibrium (Tsegaye and Yoo, 2014). This 'speed of correction' thus indicates the adjustment towards the equilibrium from the short-run to the long-run (Schaffer, 2022). Furthermore, ε_{1t}, ε_{2t}, ε_{3t} and ε_{4t} are the residuals which are assumed to have a mean of zero (Usman et al., 2019).

The same three dummy variables for years 1977, 1990 and 2006 are included in the VECM (V_t) due to the aforementioned structural breaks identified by the Chow tests. This is consistent with the strategy suggested by Bianchi (1995) of including binary variables to control for the years when structural breaks occurred and obtain a correct inference based on the Granger causality. Another alternative mentioned by Bianchi (1995) would be to perform the tests within sub-samples without structural breaks. However, in the present case, the sample size would be too small.

The short-run directional causality is tested using the χ^2 statistic of the first differenced lag of the independent variables (Usman et al., 2019), thus having the null hypotheses of:

$\phi_{1i} = 0$, $\chi_{1i} = 0$, $\psi_{1i} = 0$;

$\tau_{2i} = 0$, $\chi_{2i} = 0$, $\psi_{2i} = 0$;

$\tau_{3i} = 0$, $\phi_{3i} = 0$, $\psi_{3i} = 0$;

$\tau_{4i} = 0$, $\phi_{4i} = 0$, $\chi_{4i} = 0$.

The tests above state that income does not Granger-cause emissions, energy does not Granger-cause emissions, emissions do not Granger cause income, being consistent with the approach of Ozturk and Acaravci (2010). Thus, the short-run Granger causality is indicated by a significant relationship between the variables in the first difference: for example, real GDP Granger-causes CO_2 emissions if 'the prediction error of current environmental degradation changes by using past values of real income per capita in addition to the past values of environmental degradation' (Usman et al., 2019).

Furthermore, the long-run study inspects whether the coefficients of the ECTs are statistically different form zero, testing: $\rho_1 = 0$, $\rho_2 = 0$, $\rho_3 = 0$ and $\rho_4 = 0$. Nevertheless, it is important to note that in multivariate analyses, a straight-forward interpretation of the error correction becomes more difficult given the increased number of variables to which Granger causality can be assigned (Tsegaye and Yoo, 2014). Along these lines, Y^2 was not reported in this analysis due to interpretation reasons.

Lastly, the strong Granger causality F test aims to analyse the short-run correction meant to rebuild the long-run equilibrium (Tursoy and Faisal, 2016). Thus, the joint significance tests indicate, as per Ozturk and Acaravci (2010) and Tursoy and Faisal (2016), the short-and-long run significance of the coefficients, with the null hypotheses of:

$\phi_{1i} = \rho_1 = 0$, $\chi_{1i} = \rho_1 = 0$, $\psi_{1i} = \rho_1 = 0$;

$\tau_{2i} = \rho_2 = 0$, $\chi_{2i} = \rho_2 = 0$, $\psi_{2i} = \rho_2 = 0$;

$\tau_{3i} = \rho_3 = 0$, $\phi_{3i} = \rho_3 = 0$, $\psi_{3i} = \rho_3 = 0$;

$\tau_{4i} = \rho_4 = 0$, $\phi_{4i} = \rho_4 = 0$, $\chi_{2i} = \rho_2 = 0$.

5 Empirical Analysis

5.1 Order of Integration

As per Table 2, both tests indicate that all variables are I(1), *i.e.* non-stationary in levels, but stationary in first difference. This occurs as the p-values from the third column are higher than 5 percent for the variables in levels, suggesting the presence of a unit root. However, in the first differences specifications, the null hypothesis of a unit root process is rejected for all variables, as the p-values are lower than 5 percent. Since no variable is integrated of order 2, this enables the pursuit of the ARDL analysis, consistent with the methodology of Ouattara (2006); Chigusiwa et al. (2011) and Nkoro and Uko (2016), to name but a few.

Table 2: Order of integration

Variable	**ADF stat**	**P-value**	**Lags**	**DF-GLS stat**	**Conclusion**
Levels					
lnCO_2	-2.689	0.241	2	-1.573	Unit root
lnY	-1.770	0.719	2	-2.047	Unit root
lnY^2	-1.661	0.768	2	-1.886	Unit root
lnEN	-2.760	0.212	2	-1.747	Unit root
First Difference					
ΔlnCO_2	-4.391	0.000	1	-3.195	Stationary
ΔlnY	-3.170	0.022	1	-2.620	Stationary
ΔlnY^2	-3.123	0.025	1	-2.651	Stationary
ΔlnEN	-3.530	0.007	1	-2.730	Stationary

5.2 Cointegration

As the variables are I(1), the next step is to investigate whether they are cointegrated *i.e.* if they exhibit a long-run relationship (Nkoro and Uko, 2016).

This can be performed using the Johansen cointegration test (1988): the null hypothesis of no cointegration is rejected if the log likelihood of the unconstrained model is statistically different from the one of the constrained model which does not include the cointegrating equations (StataCorp, 2022b). The number of cointegrating relations is revealed by the trace or Maximal eigenvalue. If a single long-run relation between the variables is established by either, then the ARDL analysis can be pursued (Nkoro and Uko, 2016).

Table 3: Johansen tests for cointegration

Maximum rank	**Statistic**	**5% CV**	**1% CV**
Trace Statistic			
0	132.664	109.99	119.80
1	84.411	82.49	90.45
2	53.96	59.46	66.52
3	34.002	39.89	45.58
Max Statistic			
0	48.253	41.51	47.15
1	30.451	36.36	41.00
2	19.957	30.04	35.17
3	14.64	23.80	28.82

Table 3 indicates that, at 1 percent significance level, both the trace and the max statistics display a single cointegrating relationship. As per the Johansen (1988) cointegration test, the null hypothesis states that there is no cointegration, against the alternative that there is at least one cointegrating relationship.

The first time when the value of the trace and max statistics is lower than the critical value (*i.e.* when the null hypothesis can be rejected) is on the second row for both statistics, which corresponds to a maximum rank of one, hence one single cointegrating relation. At 5 percent significance level, the max statistic indicates a single cointegrating relation so, as per Nkoro and Uko (2016), the ARDL analysis can be conducted and the ECM can be estimated, the next step being to conduct the ARDL bounds testing.

5.3 Main Results

The ARDL model from Equation 7 is estimated using the AIC method for the lag selection. As reported in Table 4, both the F and t statistics (23.889 and -7.936 respectively) are more extreme than the critical values for the upper bounds (6.777 and -4.584), so the null hypothesis of no level relationship is rejected at 1 percent significance level. Therefore, the test indicates that there is a long-run relationship between the variables, such that the variables are cointegrated, bound together in the long-run by equilibrium forces (Kripfganz and Schneider, 2018). As per Enders (2014), an important characteristic of cointegrated variables implies that their paths over time are influenced by the extent of deviations from their long-run equilibrium: if the system returns to its long-run equilibrium, this suggests that 'movements of at least some of the variables must respond to the magnitude of the disequilibrium' (Enders, 2014).

Table 4: ARDL cointegration test

ARDL(1,4,2,3)		
CO_2=$f(Y,\ Y^2,\ EN)$		
R^2	0.947	
Adj-R^2	0.919	
N	47	
Significance level	**Lower bound I(0)**	**Upper bound I(1)**
F = 23.889		
1%	4.834	6.777
5%	3.375	4.884
10%	2.761	4.081
t = -7.936		
1%	-3.554	-4.584
5%	-2.827	-3.768
10%	-2.469	-3.361

Accordingly, given the existence of such a long-run relationship in this paper, the ECM is estimated to inspect how the variables converge towards the long-run equilibrium and how fast they absorb the short-run shocks. Table 5 presents the results of the analysis.

The second row of Table 5 highlights that there is a positive and statistically significant relationship between GDP and CO_2 emissions in the long-run. Shahbaz et al. (2013) find an increase of 0.06 percent in emissions in Romania between 1980 and 2010 associated with a 1 percent increase in GDP; Jula et al. (2015) estimate a 0.47 percent increase with a time frame between 1961 and 1990, whereas Simionescu (2021) identify a decrease of 2.17 percent in an unbalanced panel between 1996 and 2019 of V4 countries, Bulgaria and Romania. The differences in magnitude could be explained by the different timelines. Furthermore, the panel data study of Simionescu (2021) does not account for individual country characteristics, thus leading to a different result in terms of the sign of the coefficient. Lastly, the real GDP in the various studies is calculated using different benchmark years.

The Y^2 term is significant and its sign is consistent with an inverted U-shaped relationship in the long-run, thus confirming the existence of the EKC in Romania. This suggests that emissions increase with income up to a certain turning point after which they start decreasing. A potential explanation for this is related to the fall of communism in December 1989. This event triggered the closure of thousands of factories (DCNews, 2013), which consequently brought about decreased emissions despite the steady increase in GDP ever since. Moreover, a flat tax rate of 16 percent was introduced in Romania at the end of 2005 which led to foreign companies locating in Romania and implementing new technologies that could have thus limited emissions (Shahbaz et al., 2013). Additionally, given the accession to the EU in 2007, Romania has been subject to the ETS as well as to the Integrated Pollution Prevention and Control Directive introduced in 2012 which aimed to curb emissions. This could have had an effect on technology, as the directive aimed at using up to date ecological technologies (Shahbaz et al., 2013).

However, it is interesting to note that, in the short-run, the signs of the Y and Y^2 coefficients flip. A reason for this could be related to the lack of environmental pressure in the short-run, since the damaging effects of CO_2 emissions are experienced mostly in the long-run (Roca et al., 2001). As expected, energy consumption and environmental degradation display a positive relationship: a 1 percent increase in consumption leads to a 0.996 percent increase in emissions in the long-run, holding everything else constant. Nevertheless, this variable is not statistically significant in the short-run.

Table 5: ARDL estimates.
CO_2 and GDP: inverse U-shape.

	Dependent variable: $\Delta \ln CO_2$	
	Coefficient	**t-Statistics**
ECT_{t-1}	-0.962***	-7.936
Long-Run		
lnY	2.327***	2.971
$\ln Y^2$	-0.135***	-3.085
lnEN	0.996***	12.350
Short-Run		
ΔlnY	-4.371***	-3.412
$\Delta \ln Y^2$	0.246***	3.532
ΔlnEN	0.102	0.594
br77	-0.063*	-2.032
br90	-0.172***	-8.810
br06	0.038	1.386
constant	-17.454***	-4.503
R^2	0.947	
Adj-R^2	0.919	
N	47	
Diagnostics Tests	**Statistics**	
Breusch-Pagan	0.200	
White heteroscedasticity	0.431	
Breusch–Godfrey LM	0.254	
Durbin alt	0.353	
Ramsey RESET	0.605	
J–B normality	0.918	
ARCH LM	0.689	

Notes: * p<0.10, ** p<0.05, *** p<0.01.
Results based on Equation (7).

The results have a high explanatory power, with an R-squared of 94.7 percent and an adjusted R-squared of 91.9 percent. The coefficient of the lagged ECT shows how quickly a disequilibrium in the regressor is adjusted from the short-run to the long-run (Shahbaz et al., 2013). It thus measures how strongly the outcome variable reacts to one period deviations from the equilibrium relationship, *i.e.* how fast such an equilibrium distortion is rectified (Kripfganz and Schneider, 2018). Should the coefficient be zero, this would imply the lack of a long-run relationship (Menegaki, 2019).

As seen in the first row of Table 5, the ECT coefficient is negative and statistically significant at 1 percent significance level, suggesting a stable long-run relationship (Banerjee et al., 1998). Consequently, each year, the speed at which CO_2 emissions per capita respond to changes in the independent variables and return to the long-run equilibrium is 96.2 percent. This is a slightly higher result compared to Shahbaz et al. (2013), which identified that 87.7 percent of the deviations from short-run to long-run are corrected every year.

The main results also inform of three events in the history of Romania. A large earthquake in 1977 destroyed around 763 industrial units (Georgescu and Pomonis, 2018), which could have led to decreased emissions afterward. The Chow test corresponding to this year is 13.51. The critical value for 4 parameters and 51 observations at 5 percent significance level is approximately 2.56. As the test is larger than the critical value, the null hypothesis of no structural break is rejected, hence the 1977 variable (*br77*) is included. Table 5 shows that this structural break variable is significant at 10 percent significance level, indicating a decrease in emissions following 1977.

Furthermore, the change in the democratic regime from the beginning on 1990 shifted the focus from the aforementioned 'excessive industrialization' to a market economy centered around services (Foarfă, 2019). This was a change compared to the previous period when state-owned companies utilised obsolete technology in order to minimise production costs (Shahbaz et al., 2013). As a consequence, another Chow test was conducted for this year. Having a value of 34.21 against the critical value of 2.56, the null hypothesis of no break is rejected, so the *br90* variable is included. The democracy dummy is significant at 1 percent significance level, showing a decrease in emissions linked with the democratic transition of Romania.

Lastly, the 2006 dummy (*br06*) is included for a number of reasons: firstly, the end of 2005 saw the introduction of a flat tax rate that attracted foreign companies and led to the installation of modern technologies that limit CO_2 emissions (Shahbaz et al., 2013). Secondly, this variable is also meant to control for the preparation of EU accession in 2007. The EU integration is of particular relevance because emergent economies tend to display lower environmental standards beforehand, so accession could lead to a number of changes that aim to increase the environmental quality (Jula et al., 2015). For instance, an increasing number of Romanian industries became subject to the ETS following accession. By 2014, around 200 installations and operations participated in the ETS, out of 10,000, emitting approximately 40 percent of the total GHG emissions (World Bank, 2015). The Chow test for year 2006 has a value of 2.95, which is higher than the 2.56 critical value, hence the variable is included in the main model. However, Table 5 indicates that *br06* is not significantly different from 0.

5.4 Goodness of Fit

The resulting ARDL model undergoes diagnostics and stability checks to ensure its goodness of fit. The normality, serial correlation, functional form and heteroscedasticity tests indicate that the model is suitable, as per the second half of Table 5. Furthermore, the cumulative sum of recursive residuals (CUSUM) in Figure 6 and the cumulative sum of squares of recursive residuals (CUSUMSQ) in Figure 7 show that the model is stable, within the confidence bands.

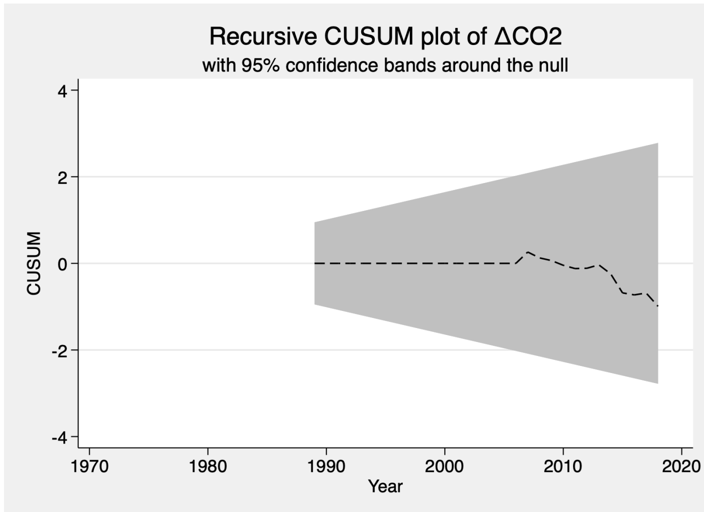

Figure 6: Cumulative sum of recursive residuals

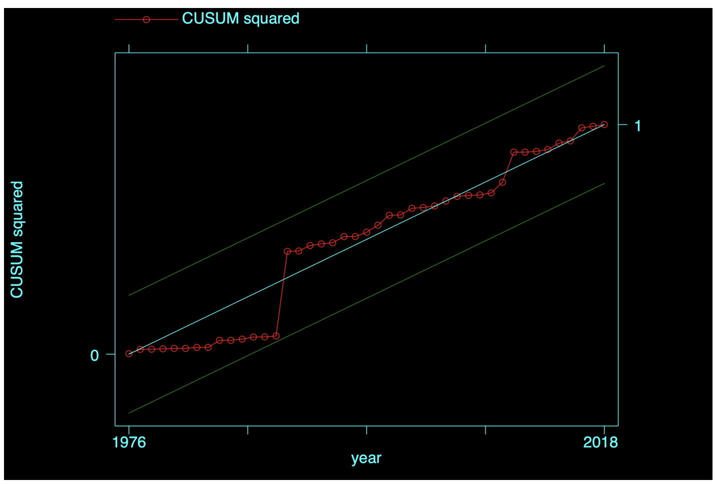

Figure 7: Cumulative sum of squares of recursive residuals

The following subsection of this paper proceeds to calculate the turning point after which environmental degradation decreases with increases in income.

5.5 The Turning Point

While the main model of the paper identifies an inverted U-shaped relationship between environmental damage and income per capita, the question revolves around which part of the curve the country is at the moment: if the economy is far from the turning point, that implies high environmental degradation over a period of time (Kunnas and Myllyntaus, 2007). This section thus proceeds to calculate the value of the turning point.

An initial plot of the emissions and income levels reinforces the existence of an inverse U-shaped relationship, as displayed by Figure 8.

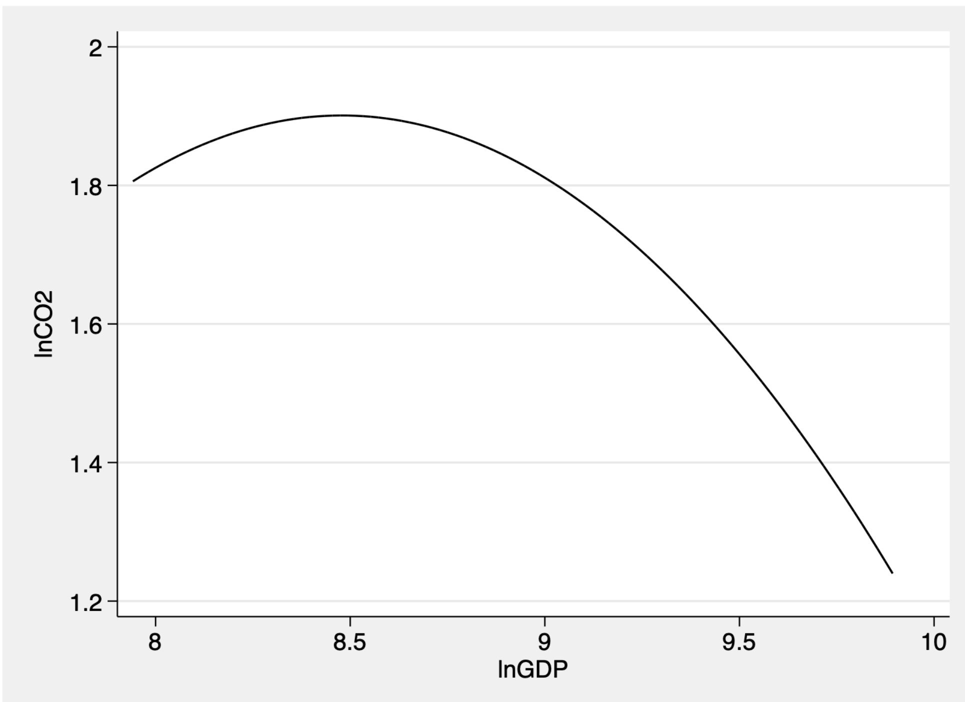

Figure 8: CO_2 emissions and GDP

Based on the original EKC specification, the ARDL ECM was estimated as previously discussed. The turning point of per capita income was then calculated by finding the maximum value of the function, *i.e.* taking the derivative of emissions with respect to income for the estimations of the long-run variables:

$$\frac{d(lnCO_{2,\mathrm{t}})}{dY_{\mathrm{t}}} = \beta_1 \frac{1}{Y_t} + \beta_2 2lnY_t \frac{1}{Y_t} \tag{12}$$

Setting Equation (12) equal to zero and rearranging yields:

$$lnY_t = -\frac{\beta_1}{2\beta_2} \tag{13}$$

Consequently, the turning point is calculated as $exp(-\frac{\beta_1}{2\beta_2})$. This methodology is consistent with the one used in the empirical literature, for instance by Zhang et al. (2019) or Bernard et al. (2015).

The estimated extreme point is at a value of per capita GDP of approximately \$5,700 measured in 2011 prices, which occurred in the beginning of the 1990s. This result is lower compared to the ones in literature such as the turning points calculated by Lazăr et al. (2019) which are higher than \$9,000 for all studied CEE countries. It is however acknowledged that there is a wide variation in the estimated turning points within the literature (Hill and Magnani, 2002; Lazăr et al., 2019). The timing of the Romanian turning point is associated with the change in the political regime.

5.6 Extended EKC Analysis

The extended EKC in Figure 2 states that, following the turning point, emissions will start increasing again in the long-run, as the damaging impact of economic growth upon the environment cannot be compensated anymore (Lazăr et al., 2019). The question of whether income growth 'has an automatic tendency to diminish environmental damage at high-income levels' (Kunnas and Myllyntaus, 2007) could be answered through employing an extended EKC analysis.

Therefore, a complete study requires the introduction of the third degree polynomial for the GDP term to augment the main model. Otherwise, identifying the inverse U-shape might be misleading, as it would suggest that emissions decrease with income and the strategy of 'green growth' is therefore working (Sinha et al., 2018). Instead, this paper is questioning whether, at higher income levels, the GDP increases are still associated with diminished pollution levels (Kunnas and Myllyntaus, 2007).

After establishing that the variables have one cointegrating relation (Table A.1 in Appendix), a new ARDL model is estimated.

The underlying extended EKC specification is as follows:

$$lnCO_{2,\mathrm{t}} = \iota_0 + \iota_1 lnY_t + \iota_2 lnY^2{}_t + \iota_3 lnY^3{}_t + \iota_4 lnEN_t + \theta V_t + \varepsilon_\mathrm{t} \tag{14}$$

As a consequence, the ARDL model becomes:

$$\Delta lnCO_{2,\mathrm{t}} = \iota_0 + \iota_1 lnCO_{2,t-1} + \iota_2 lnY_{t-1} + \iota_3 lnY_{t-1}^2 + \iota_4 lnY_{t-1}^3 + \iota_5 EN_{t-1} + \sum_{i=1}^{p} \kappa_i \Delta lnCO_{2,t-i} + \\ + \sum_{j=0}^{q} \kappa_j \Delta lnY_{t-j} + \sum_{k=0}^{r} \kappa_k \Delta lnY_{t-k}^2 + \sum_{l=0}^{s} \kappa_l \Delta lnY_{t-l}^3 + \sum_{m=0}^{u} \kappa_m \Delta lnEN_{t-m} + \theta V_t + \mu_\mathrm{t} \tag{15}$$

The bounds test in Table 6 indicates yet again that there is a long-run association between the variables, as both the F and t statistics are more extreme than the upper bounds at 1 percent significance level.

Table 6: ARDL cointegration test, extended model

ARDL(1,0,1,1,0)		
CO_2=f(Y, Y^2, Y^3, EN)		
R^2	0.911	
Adj-R^2	0.889	
N	50	
Significance level	**Lower bound I(0)**	**Upper bound I(1)**
F = 35.693		
1%	4.308	6.006
5%	3.093	4.468
10%	2.571	3.800
t = -12.587		
1%	-3.555	-4.806
5%	-2.866	-4.023
10%	-2.523	-3.629

Consequently, the ECM is estimated, following the same procedure as in the main model. The results of this subsection are therefore based on the following specification:

$$\Delta lnCO_{2,\mathrm{t}} = \iota_0 + \iota_1 lnCO_{2,t-1} + \iota_2 lnY_{t-1} + \iota_3 lnY_{t-1}^2 + \iota_4 lnY_{t-1}^3 + \iota_5 EN_{t-1} + \sum_{i=1}^{p} \kappa_i \Delta lnCO_{2,t-i} + \\ + \sum_{j=0}^{q} \kappa_j \Delta lnY_{t-j} + \sum_{k=0}^{r} \kappa_k \Delta lnY_{t-k}^2 + \sum_{l=0}^{s} \kappa_l \Delta lnY_{t-l}^3 + \sum_{m=0}^{u} \kappa_m \Delta lnEN_{t-m} + \zeta ECT_{t-1} + \theta V_t + \mu_\mathrm{t} \tag{16}$$

Table 7 presents the results of this augmented model, displaying the presence of an inverse N-shaped relationship between per capita emissions and real GDP in the long-run. The result is consistent with the finding of Jula et al. (2015) for Romania. On top of that, this kind of pattern was also identified by Lazăr et al. (2019) for countries such as Poland, Slovakia or Lithuania, albeit having higher values of income per capita at the peaks and troughs.

The ECM has a high explanatory power, with an R^2 of 91.1 percent and an adjusted R^2 of 88.9 percent. Furthermore, the ECT from the first row of Table 7 signifies that, within one year, approximately 98 percent of the deviations from the long-run equilibrium are corrected. It is negative and highly significant at 1 percent level. This result is similar to the main model one which identified a 96 percent adjustment.

Table 7: ARDL estimates for extended model.
CO_2 and GDP: inverse N-shape

	Dependent variable: ΔlnCO_2	
	Coefficient	**t-Statistics**
$ECT_{t\text{-}1}$	-0.979***	-12.587
Long-Run		
lnY	-18.187**	-2.089
lnY^2	2.089**	2.16
lnY^3	-0.08**	-2.236
lnEN	1.121***	20.763
Short-Run		
ΔlnY^2	-0.098	-1.344
ΔlnY^3	0.008	1.465
br77	-0.048**	-2.34
br90	-0.143***	-7.254
br06	0.005	.175
constant	42.439	1.663
R^2	0.911	
Adj-R^2	0.889	
N	50	
Diagnostics Tests	**Statistics**	
Breusch-Pagan	0.051	
White heteroscedasticity	0.292	
Breusch–Godfrey LM	0.843	
Durbin alt	0.862	
Ramsey RESET	0.577	
J–B normality	0.503	
ARCH LM	0.33	

Notes: * $p<0.10$, ** $p<0.05$, *** $p<0.01$.
Results based on Equation (16).

In terms of energy, the result is similar compared to the quadratic model: a one percent increase in primary energy consumption per capita is associated with a 1.12 percent increase in CO_2 emissions per capita in the long-run, *ceteris paribus*.

The exogenous variables are the same as in the main model, accounting for breaks in years 1977, 1990 and 2006, with the first two being statistically significant at 5 percent and 1 percent significance levels respectively.

Given the limited lag length and the employed AIC method for lag selection, the short-run dynamics cannot be identified. This is a limitation of the cubic model, and the results of this inverse N-shaped analysis should be taken with a grain of salt for a number of reasons. Given the small sample size, the maximum number of lags in the ARDL model had to be limited to one, as otherwise the model would not run and variables would be omitted due to collinearity. Furthermore, the Breusch Pagan test (1979) is much weaker: while we cannot reject the null hypothesis of constant variance among the residuals at 5 percent significance level, its test value is close to the rejection area, unlike the one in the main model. Nevertheless, the model passes the diagnostics tests, as shown by Table 7. In addition to this, the CUSUM (Figure 9) and CUSUMSQ (Figure 10) indicate the stability of the model.

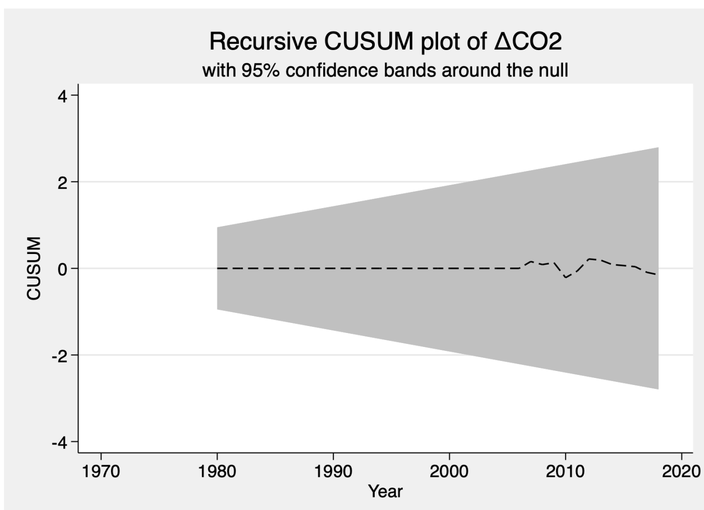

Figure 9: Cumulative sum of recursive residuals, extended model

Regarding the turning points, the trough is achieved at a level of GDP per capita of $5,569 in 2011 prices, whereas the peak is at a level of $6,552. GDP has been beyond that level since 1999. Consequently, the main takeaway of this cubic estimation is that Romania is not under the extended EKC scenario where emissions start increasing with the level of income, as described by an N-shaped relationship. In practice, an inverse N-shaped curve can be observed instead.

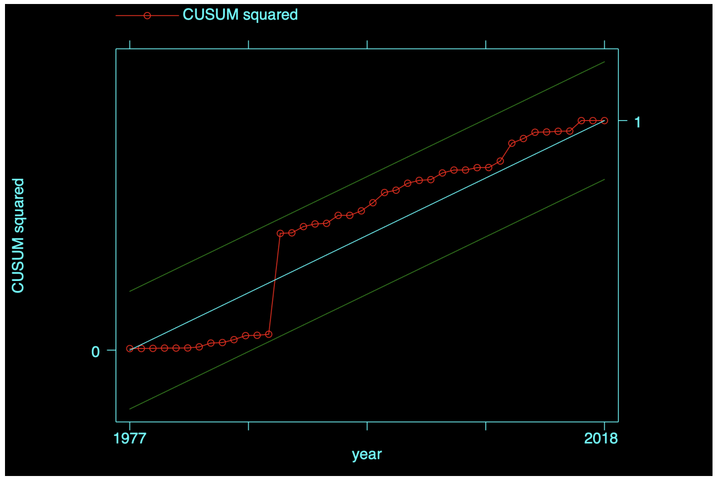

Figure 10: Cumulative sum of squares of recursive residuals, extended model

5.7 Two-Period Analysis

In this subsection, the main quadratic EKC analysis is further split into two distinct time frames to account for the significant change that Romania went through in 1989 from a centrally-planned economy to a free market. The number of observations is more limited compared to the initial analysis, so the power of the tests is likely to be lower.

5.7.1 1968-1988

As per the ADF and the DF-GLS tests (Table A.2 in the Appendix), the variables are a mix between I(0) and I(1). Like in the main model, a single cointegrating relationship is identified thourgh the Johansen test (Table A.3 in the Appendix). Furthermore, Table 8 confirms the existence of a long-run relationship between emissions and economic growth. Both the F and the t statistics (14.617 and -7.122) are more extreme than the upper bound critical values at 1 percent significance level (9.009 and -5.092 respectively), thus rejecting the null hypothesis of no cointegration

The ECM is consequently estimated and the results are presented in Table 9. The model passes the diagnostics and stability checks and has a high explanatory power. However, while the ECT is statistically significant at 1 percent significance level and negative, its absolute value is larger than 1, indicating some overcorrection. As per Narayan and Smyth (2006), a speed of adjustment of 114 percent could imply that instead of directly converging to the equilibrium, the process fluctuates around the long-run value in a dampening way instead of monotonically. When the process is complete, the convergence to the long-run equilibrium is fast (Narayan and Smyth, 2006).

Table 8: ARDL cointegration test for period 1968-1988

ARDL(1,1,0,0)		
CO_2=*f*(Y, Y^2, EN)		
R^2	0.895	
Adj-R^2	0.846	
N	20	
Significance level	**Lower bound I(0)**	**Upper bound I(1)**
F = 14.617		
1%	6.660	9.009
5%	4.173	5.813
10%	3.259	4.626
t = -7.122		
1%	-3.927	-5.092
5%	-3.003	-3.995
10%	-2.581	-3.495

Furthermore, the signs of the GDP terms indicate a U-shaped pattern in the long-run, given the negative sign of the Y coefficient in the second row of Table 10 and the positive sign of the Y^2 one in the third row. Interestingly, in this time period, the GDP coefficient has a negative sign both in the short-run and in the long-run, unlike in the main model where the long-run effect was positive. However, these coefficients are not statistically significant in this model.

Lastly, energy consumption and emissions have a positive relationship: a 1 percent increase in primary energy consumption per capita leads to a 1.01 percent increase in emissions per capita, holding everything else constant.

The model passes all diagnostics and stability checks, as confirmed by Table 9 and the CUSUM and CUSUMSQ plots (Figures A.1 and A.2 in the Appendix). It also has a good explanatory power, with an R^2 of 89.5 percent and an adjusted R^2 of 84.6 percent.

Table 9: ARDL estimates for period 1968-1988.
CO_2 and GDP: U-shape.

	Dependent variable: $\Delta \ln CO_2$	
	Coefficient	**t-Statistics**
ECT_{t-1}	-1.14***	-7.12
Long-Run		
lnY	-0.458	-0.26
$\ln Y^2$	0.031	0.30
lnEN	1.011***	9.01
Short-Run		
ΔlnY	-0.169	-1.47
br77	-0.055**	-2.68
constant	-7.538	-0.95
R^2	0.895	
Adj-R^2	0.846	
N	20	
Diagnostics Tests	**Statistics**	
Breusch-Pagan	0.052	
White heteroscedasticity	0.395	
Breusch–Godfrey LM	0.465	
Durbin alt	0.553	
Ramsey RESET	0.052	
J–B normality	0.965	
ARCH LM	0.137	

Notes: * p<0.10, ** p<0.05, *** p<0.01.
Results based on Equation (7).

5.7.2 1993-2018

The same analysis is conducted for the second period, following the fall of communism and the transition of Romania to a free market economy. The ADF and DF-GLS tests (Table A.4 in the Appendix) confirm that all variables are integrated of order 1 for this new period between 1993 and 2018.

The next step is to study whether there exists a cointegrating relationship between the variables using the Johansen cointegration test. Both the trace and max statistics indicate the existence of a single cointegrating relationship at 1 percent significance level (Table A.5 in the Appendix).

Furthermore, the ARDL cointegration test in Table 10 confirms the existence of a long-run connection between emissions and economic growth in the period between 1993 and 2018: both the F statistic of 21.989 and the t statistic of -8.696 are more extreme than the critical values for the upper bound. Thus, the null hypothesis of no cointegration is rejected, suggesting a long-run relationship.

The next step is to estimate the ECM and Table 11 reports the results. Unlike in the main period, the sign of the GDP coefficient does not flip and remains positive in the short-run and in the long-run.

Both the Y and Y^2 coefficients are significant at 10 percent significance level in the long-run, while the short-run GDP coefficient is significant at 1 percent level. As the squared income coefficient is negative, that implies, as seen in the main analysis, an inverse U-shape relationship between the variables.

Table 10: ARDL cointegration test for period 1993-2018

ARDL(1,1,0,0)		
$CO_2=f(Y,\ Y^2,\ EN)$		
R^2	0.872	
Adj-R^2	0.831	
N	26	
Significance level	**Lower bound I(0)**	**Upper bound I(1)**
F = 21.989		
1%	5.787	7.790
5%	3.843	5.324
10%	3.077	4.346
t = -8.696		
1%	-3.764	-4.855
5%	-2.953	3.915
10%	-2.569	-3.469

The fact that the short-run income variable has the same sign as the long-run one could be related to the EU accession. Given the previous discussion, CO_2 is a global gas (Cole et al., 1997), with incentives to free ride on others' mitigation efforts (Roca et al., 2001; Kunnas and Myllyntaus, 2007). There are also less incentives to tackle it in the short-run given that its negative consequences experienced more in the long-run (Arrow et al., 1995). However, given the EU accession, Romania became subject to the ETS and had more incentives to curb emissions, as well as more revenue from the ETS to be invested in climate finance (World Bank, 2015). This finding is also consistent with the view of Arrow et al. (1995) that global GHGs are more likely to be tackled in the short-run and to have lower turning points of per capita income when they are subject to multilateral policy initiative.

Table 11: ARDL estimates for period 1993-2018.
CO_2 and GDP: inverse U-shape.

	Dependent variable: ΔlnCO_2	
	Coefficient	**t-Statistics**
ECT_{t-1}	-0.881***	-8.7
Long-Run		
lnY	1.786*	1.80
lnY^2	-0.101*	-1.84
lnEN	1.16***	10.39
Short-Run		
ΔlnY	0.355**	2.51
br06	0.005	0.16
constant	-15.803***	-3.35
R^2	0.872	
Adj-R^2	0.831	
N	26	
Diagnostics Tests	**Statistics**	
Breusch-Pagan	0.633	
White heteroscedasticity	0.436	
Breusch–Godfrey LM	0.649	
Durbin alt	0.699	
Ramsey RESET	0.17	
J–B normality	0.994	
ARCH LM	0.459	

Notes: * $p<0.10$, ** $p<0.05$, *** $p<0.01$.
Results based on Equation (7).

In the long-run, a 1 percent increase in primary energy consumption per capita is associated with a 1.16 percent increase in emissions per capita, this being significant at 1 percent level. The result is similar to the one from the 1968-1990 period, as well as with the main analysis one, albeit being a slightly larger effect. The coefficient of the ECT is statistically significant and negative, suggesting that within a year, 88.1 percent of the deviations from the long-run equilibrium are corrected.

The model passes all diagnostics and stability checks, as confirmed by Table 11 and the CUSUM and CUSUMSQ plots (Figures A.3 and A.4 in the Appendix). It also has a good explanatory power, with an R^2 of 87.2 percent and an adjusted R^2 of 83.1 percent.

5.8 Granger Causality

As a cointegrating relationship between the variables exists in the main model, this implies that there ought to be Granger causality in no less than one direction (Ozturk and Acaravci, 2010). Therefore, this subsection of the present thesis will employ a VECM analysis to inspect the direction of the

Granger causality. An advantage of this procedure is that it identifies both short-run and long-run causality (Usman et al., 2019). Consequently, the purpose of the tests is to assist policymakers in the implementation of inclusive energy, economic and environmental policies that aim at supporting growth and safeguarding environmental quality in the long-run (Shahbaz et al., 2016).

Table 12 presents the results of the VECM Granger causality tests which were implemented in a similar manner to the ones by Shahbaz et al. (2016) and Ozturk and Acaravci (2010). Two lags were included. Following the methodology by Tsegaye and Yoo (2014), the reported values are the asymptotic Granger causality χ^2, whereas the *p-values* are displayed in parentheses.

Table 12: VECM Granger Causality Test

	Short Run			**Long Run**
Dep. Variable	$\Delta \ln CO_{2,t\text{-}1}$	$\Delta \ln GDP_{t\text{-}1}$	$\Delta \ln EN_{t\text{-}1}$	$ECT_{t\text{-}1}$
$\Delta \ln CO_{2,t}$	-	0.10 (0.749)	0.21 (0.649)	22.87 (0.000)
$\Delta \ln GDP_t$	4.93 (0.026)	-	3.02 (0.083)	4.17 (0.0412)
$\Delta \ln EN_t$	6.39 (0.012)	3.10 (0.078)	-	0.37 (0.541)
	Joint Significance			
	$\Delta \ln CO_{2,t\text{-}1}$, $ECT_{t\text{-}1}$	ΔlnGDP t-1, $ECT_{t\text{-}1}$	ΔlnEN t-1, $ECT_{t\text{-}1}$	
$\Delta \ln CO_{2,t}$	-	28.57 (0.000)	23.10 (0.000)	
$\Delta \ln GDP_t$	6.77 (0.034)	-	7.22 (0.027)	
$\Delta \ln EN_t$	6.47 (0.039)	5.10 (0.078)	-	

Notes: Asymptotic Granger causality χ^2.
p-values in parentheses.

The short-run section of Table 12 indicates bidirectional Granger causality between energy and GDP. Furthermore, CO_2 emissions Granger-cause GDP and energy. As depicted by Table 5 of the main analysis, emissions and income have a negative and significant relationship in the short-run. In this context and accounting for the direction of the Granger causality analysis, sudden decreases in environmental degradation could lead to increases in GDP, holding everything else constant. Hence, policy implications would also revolve around changing from fossil fuels to other alternatives such as renewable energy (Coondoo and Dinda, 2002) to diminish the CO_2 emissions.

The causation uncovered from emissions to income is consistent with the findings by Coondoo and Dinda (2002). The authors identify that for North America and Western Europe and, to some extent, Eastern Europe, the direction of causation is from environmental degradation to income. This direction from emissions to GDP could hence be attributed to the particularities of Romania. As Coondoo and Dinda (2002) hypothesise, manufacturing could have an important role in explaining this direction: given that manufacturing is more emission-intensive than other sectors such as the services one, countries with a high share of manufacturing in their income (of more than 20 percent) that depend on such activities for economic growth are likely to experience causality from emissions to GDP (Coondoo and Dinda, 2002). In the particular case of Romania, the country's industrial base is one of the largest ones in the CEE group, with manufacturing being the more substantial sector, making up over 20 percent of the value added in 2020 (Milatovic and Szczurek, 2020). The manufacturing's contribution to GDP used to be even larger: for instance, in 1990, the share of manufacturing output was over 30 percent (Macrotrends, 2022). Despite the decreases in the share throughout the studied time period, it still remains an important share, accounting for almost 20 percent of GDP (Macrotrends, 2022) and representing 17 percent of total CO_2 emissions (Figure 5) at the end of the studied timeline, which is therefore in line with the hypothesis of Coondoo and Dinda (2002).

In regard to the long-run, the ECT coefficient is significant at 1 percent level for emissions and 5 percent level for income, suggesting the existence of bidirectional Granger causality. The tests in Table 12 thus indicate long-run causality from real income per capita, square real income per capita, and primary energy consumption per capita to CO_2 emissions at 1 percent significance level. This is consistent with the results of Victor Bekun et al. (2019) and Usman et al. (2019) that pursuing economic growth exerts downward pressure on environmental quality. Similarly, at 5 percent significance level, there is also long-run Granger causality from environmental degradation, squared real GDP per capita and primary energy consumption per capita to real GDP per capita. This implies that efforts to reduce fossil-fuel based energy consumption could help reduce emissions, as per Usman et al. (2019).

The joint-significance test results inform of strong interdependence between CO_2 emissions and income, as well as between emissions and energy consumption. There is also two-way causality between income and energy consumption, albeit at 10 percent significance level, thus revealing complex interdependencies between the variables.

The policy recommendation associated with these findings is increasing the use of renewable resources, which could be linked with both decreased emissions and dependence on fossil fuels, leading to increased energy security (Usman et al., 2019). As suggested by Shahbaz et al. (2013), these results could also prompt the Romanian policymakers to increase efforts to acquire and develop new technologies with high energy efficiency and low pollution to further curb emissions. Consistent with the findings of Chontanawat (2020), policies decreasing energy consumption would also help increase environmental quality in the long-run, given the positive relationship between the two variables uncovered by Table 5. In terms of the implementation of such policies, Romania could follow the recommendation by the European Commission (2019) to use the EU ETS revenues for projects revolving around energy efficiency and increasing the share of renewable resources in the energy mix, as well as for the financing of decarbonisation technologies.

6 Conclusion and Potential Extensions

My thesis finds a long-run relationship between CO_2 emissions per capita and income per capita in Romania over the period between 1968 and 2018. The relationship is consistent with the EKC hypothesis and uncovering it is an interesting result given the particularities of the CO_2, its obscured short-term costs and its global nature (Cole et al., 1997).

This finding could be explained by a number of particularities of Romania, such as the change in the democratic regime associated with deindustrialization, and the accession to the EU together with the inherent participation in the ETS. When splitting the analysis into two parts, the results suggest that a U-shaped relationship between emissions and income can be observed between 1968 and 1988, whereas an inverse U-shape is present between 1993 and 2018. Over the entire period, the result of an inverse N-shape is consistent with the two-period analysis. These dynamics imply that a green growth strategy is currently applicable for Romania. However, the power of the two-period and extended analyses is weaker, given the very small sample size and the limited lag length.

Even though the results indicate a long-run relationship between the variables, it is also important to think about the external validity. Finding an inverse U-shaped or inverse N-shaped relation in Romania for a given time frame does not necessarily imply that this would be the case in other countries, given the heterogeneities in the dynamics between emissions and income (Lazăr et al., 2019). It is also no evidence that the same will also hold in the future, especially in a timely manner in order to avoid irreversible global consequences (Arrow et al., 1995). Subsequently, the findings - both in terms of the turning point, as well as the short-run and long-run dynamics between emissions and income - are not likely to be generalised to other settings. The reasons for that are related to the unique circumstances of Romania, the change in regime, its national legislation throughout the time period, as well as its own energy source mix (Simionescu, 2021; Roca et al., 2001). Therefore, the results lay out information that is mostly relevant for the Romanian policymakers, providing background on the long-run relationship between environmental degradation and economic growth. As a consequence, the situation ought to be closely monitored and the authorities should act at the first sign of an increase in environmental degradation following income growth. While the results of this thesis suggest that the tipping point has been surpassed in Romania before the turn of the 20$^{\text{th}}$ century, that does not necessarily imply that emissions will keep decreasing with increases in GDP over the long-run horizon (Kunnas and Myllyntaus, 2007). As long as emissions have not reached the net zero target, there is still room for improvement.

Since countries could be on different points of the EKC at the same time (Lazăr et al., 2019), a global emissions trading scheme where developed countries acquire permits from developing ones (Coondoo and Dinda, 2002) could be an area of further research. Even if my study uncovers an inverse N-shaped link between income and pollution in Romania, that may not be externally valid. Some developed countries might display an N-shaped pattern and find themselves on the rising end of the EKC (Bednar-Friedl and Getzner, 2003; Balsalobre et al., 2015; Lorente and Álvarez-Herranz, 2016; Allard et al., 2018). Simultaneously, developing countries are prone to be on the first part of the EKC (Allard et al., 2018), as their economic development and intensive use of abundant natural resources negatively impact the environment (Hasanov et al., 2019). This context together with the global nature of CO_2 (Cole et al., 1997) speak in favour of creating a global emissions trading scheme which would enable a transfer of funds from developed to developing countries (Coondoo and Dinda, 2002). This would be advantageous given the significant short-term costs associated with substituting away from fossil-fuel technologies by developing countries (Gillingham, 2019). These countries also tend to have stringent public budgets and raising revenue through taxation could harm resource allocation due to firms and households facing a distorted price regime (Eskeland and Jimenez, 1992). Thus, such transfer of funds would be an alternative policy to taxation that would help mitigate climate change despite the aforementioned tight public budget of developing countries. While the ETS already exists in the EU, it would be informative to study the consequences of extending the ETS globally, since for now the scheme only covers similar countries in terms of development levels. Such research should nonetheless be mindful of the potential backlash from developing countries wanting to industrialize, as well as of the inherent requirement for a global pollution monitoring agency (Coondoo and Dinda, 2002).

An extension to my thesis could be conducting this analysis on the separate regions of Romania using panel data. Such a study was conducted by Akbostancı et al. (2009) in the case of Turkey, thus shedding more light on the particularities of the individual country and accounting for heterogeneity in the different regions. As per the existing EKC literature, more variables could augment the present study, such as trade openness or investments in research and development projects in the area of renewable energy. The framework of this thesis could also be applied to research the nexus between income and other GHGs like in the literature, and the results could then be compared to the CO_2 ones.

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Appendix

Table A.1: Johansen tests for cointegration, extended model

Maximum rank	**Statistic**	**5% CV**	**1% CV**
Trace Statistic			
0	184.423	141.20	152.32
1	129.544	109.99	119.80
2	89.484	82.49	90.45
3	64.12	59.46	66.52
Max Statistic			
0	54.879	47.99	53.90
1	40.06	41.51	47.15
2	35.364	36.36	41.00
3	22.406	30.04	35.17

Table A.2: Order of integration for period 1968-1988

Variable	**ADF stat**	**P-value**	**Lags**	**DF-GLS stat**	**Conclusion**
Levels					
$\ln CO_2$	-1.410	0.858	1	-0.985	Unit root
lnY	-1.647	0.774	4	-0.863	Unit root
$\ln Y^2$	-1.661	0.768	4	-0.832	Unit root
lnEN	-2.165	0.023	3	-1.839	Inconclusive
First Difference					
$\Delta \ln CO_2$	-3.527	0.007	0	-4.625	Stationary
ΔlnY	-2.653	0.008	0	-4.068	Stationary
$\Delta \ln Y^2$	-2.708	0.008	0	-4.048	Stationary
ΔlnEN	-	-	3	-3.650	Stationary

Table A.3: Johansen tests for cointegration for period 1968-1988

Maximum rank	**Statistic**	**5% CV**	**1% CV**
Trace Statistic			
0	73.700	59.46	66.52
1	42.386	39.89	45.58
2	22.325	24.31	29.75
3	5.187	12.53	16.31
Max Statistic			
0	31.315	30.04	35.17
1	20.061	23.80	28.82
2	17.138	17.89	22.99
3	4.831	11.44	15.69

Table A.4: Order of integration for period 1993-2018

Variable	**ADF stat**	**P-value**	**Lags**	**DF-GLS stat**	**Conclusion**
Levels					
lnCO_2	-2.838	0.183	3	-2.748	Unit root
lnY	-1.065	0.935	1	-0.923	Unit root
lnY^2	-1.135	0.923	1	-1.008	Unit root
lnEN	-2.440	0.359	1	-2.593	Unit root
FD					
ΔlnCO_2	-4.081	0.001	0	-4.129	Stationary
ΔlnY	-4.264	0.001	0	-4.062	Stationary
ΔlnY^2	-4.344	0.000	0	-4.151	Stationary
ΔlnEN	-4.794	0.000	0	-4.530	Stationary

Table A.5: Johansen tests for cointegration for period 1993-2018

Maximum rank	**Statistic**	**5% CV**	**1% CV**
Trace Statistic			
0	101.032	59.46	66.52
1	38.041	39.89	45.58
2	18.725	24.31	29.75
3	8.820	12.53	16.31
Max Statistic			
0	62.991	30.04	35.17
1	19.316	23.80	28.82
2	10.905	17.89	22.99
3	5.878	11.44	15.69

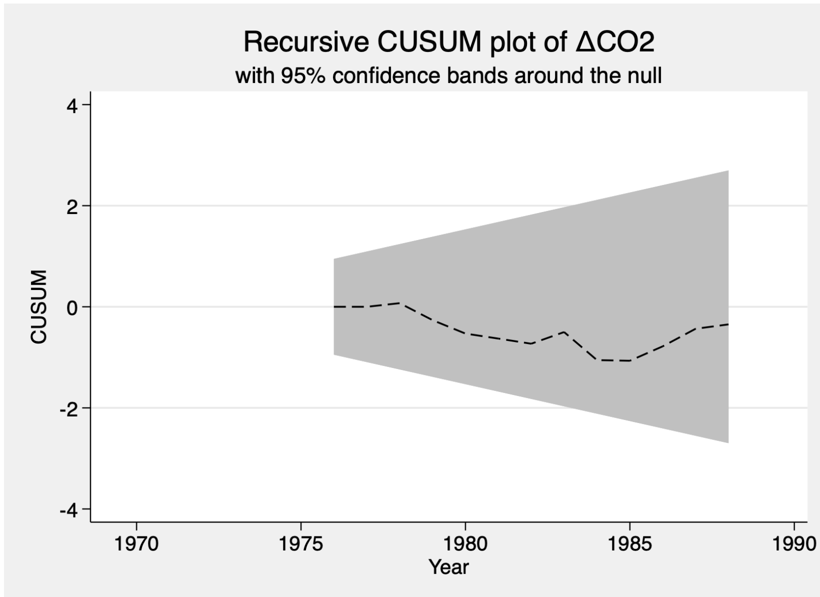

Figure A.1: Cumulative sum of recursive residuals for period 1968-1988

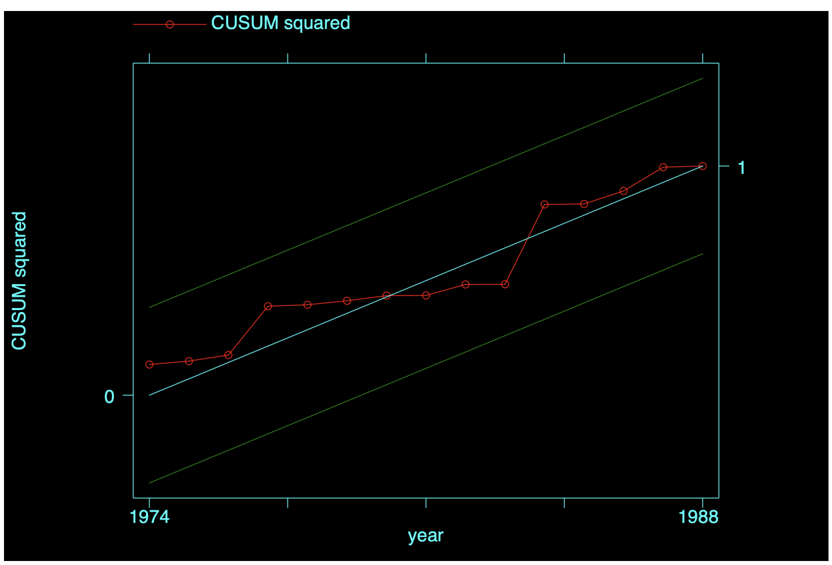

Figure A.2: Cumulative sum of squares of recursive residuals for period 1968-1988

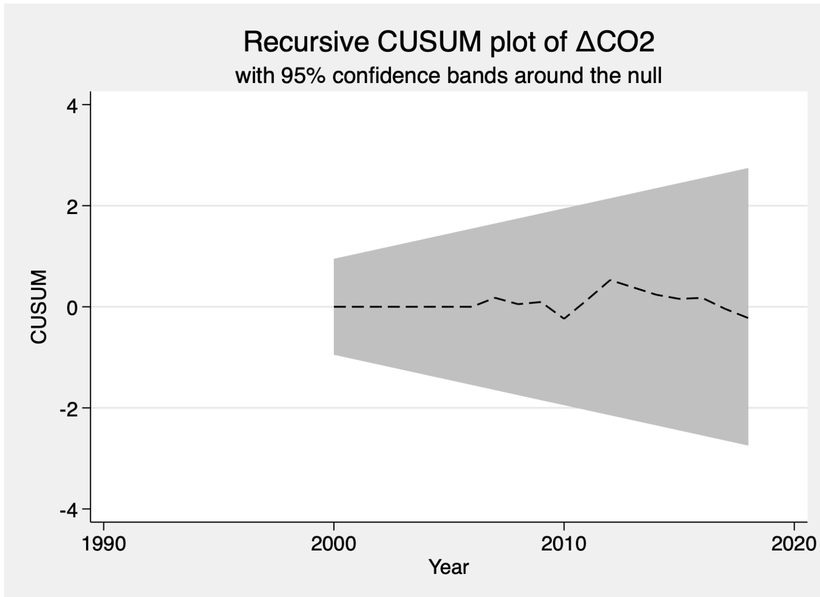

Figure A.3: Cumulative sum of recursive residuals for period 1993-2018

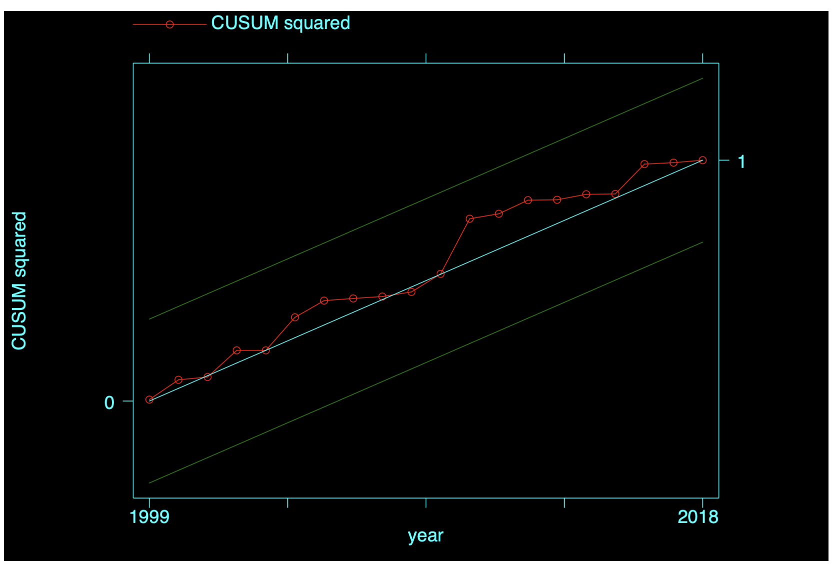

Figure A.4: Cumulative sum of squares of recursive residuals for period 1993-2018