ISCTE IUL
Instituto Universitário de Lisboa

Department of Economics

CO_2 emissions and economic growth: an assessment of the power and transport sectors

A Thesis presented in partial fulfilment of the Requirements for the Degree of Doctor in Economics

By

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Resumo

Esta tese é composta por três capítulos subordinados ao tema da Economia do Ambiente. Nos capítulos 1 e 2 investiga-se empiricamente a existência da hipótese da Curva Ambiental de Kuznets (CAK) para descrever a relação entre as emissões de CO_2 e o crescimento económico para Portugal, entre 1960 e 2010. É um estudo inovador, pois além da vertente de âmbito nacional, analisa setorialmente a CAK para os dois principais emissores de dióxido de carbono: o setor da produção de eletricidade e o setor dos transportes. Adicionalmente, são incorporadas no modelo outras variáveis que potencialmente podem ser determinantes para as emissões de CO_2, nomeadamente o preço do crude, o preço médio dos combustíveis, a taxa de motorização, a temperatura média e a precipitação média. No capítulo 1 recorre-se à metodologia da cointegração não linear para as formas funcionais quadrática e cúbica, quer em níveis quer em logaritmos naturais. No capítulo 2, utiliza-se a técnica da cointegração com quebras de estrutura desconhecidas para um modelo linear, também em níveis e logaritmos. As duas abordagens econométricas evidenciam resultados distintos.

No capítulo 3, desenvolve-se um modelo de Cournot de dois períodos para se aferir o comportamento estratégico dos produtores de eletricidade térmica e de eletricidade produzida a partir de fontes renováveis (E-FER), quando este último, para além da quantidade, tem de decidir o nível de investimento em capacidade instalada para estar disponível no período seguinte. Os resultados são comparados com as soluções do ótimo social.

Palavras-chave: Crescimento económico, setor de produção de eletricidade, setor dos transportes, curva ambiental de Kuznets, cointegração não linear, cointegração com quebras de estrutura, E-FER, modelo de Cournot, tarifas feed-in, emissões de CO_2, políticas ambientais

Classificação JEL: C32, L13, Q48, Q58

Abstract

This thesis consists of three chapters in the field of Environmental Economics. Chapters 1 and 2 investigate the empirical existence of the hypothesis of the Environmental Kuznets Curve (EKC) to describe the relationship between CO_2 emissions and economic growth for Portugal between 1960 and 2010. It is an innovative study because it goes beyond the aggregate national data to provide a sectoral analysis of the EKC hypothesis for the two major carbon emitters: the power generation sector and the transport sector. Additionally, we add into the model other variables that can potentially act as determinants of CO_2 emissions, namely the price of crude oil, the average price of fuel, the rate of motorization, the average temperature and average precipitation. Chapter 1 uses the non-linear cointegration methodology to analyse both the quadratic and cubic functional forms, in levels or in natural logarithms. Chapter 2 employs the approach of cointegration with unknown structural breaks in a linear model, also at levels and logarithms. The two econometric approaches reveal different results.

In Chapter 3, a two-period Cournot model is developed to assess the strategic behaviour of both thermal and renewable electricity (RES-E) producers when the latter, besides quantity, has to make a decision about capacity investment to be available in the next period. The results are then compared with socially optimal solutions.

Keywords: Economic growth, power generation sector, transport sector, environmental Kuznets curve, nonlinear cointegration, cointegration with structural breaks, RES-E, Cournot model, feed-in tariffs, CO_2 emissions, environmental policies

JEL codes: JEL: C32, L13, Q48, Q58

Resumo alargado

Os gases com efeito de estufa (GEE), sobretudo o dióxido de carbono (CO_2), associados ao paradigma económico assente nos combustíveis fósseis que vigora desde a revolução industrial, têm imposto custos externos crescentes à sociedade. A principal externalidade negativa relacionada com a emissão de GEE são as alterações climáticas.

Segundo o quinto relatório elaborado pelo Painel Intergovernamental para Alterações Climáticas das Nações Unidas, as manifestações das alterações climáticas já estão a ocorrer através de eventos como secas, ondas de calor, e cheias. Os danos causados por fenómenos climáticos extremos vão acentuar-se no futuro, levando à redução do rendimento de culturas agrícolas, incêndios ou extinção de espécies, com consequentes impactos económicos.

A produção de eletricidade e os transportes são setores intensivos em carbono, e por isso, os principais emissores, pelo que assumem especial relevância no contexto atual cujo objetivo é a mitigação e adaptação às consequências das alterações climáticas.

A presente tese, subordinada ao tema da Economia do Ambiente, visa ser um contributo para a discussão das alterações climáticas incidindo nos dois principais setores emissores anteriormente mencionados. Composto por três capítulos, este trabalho de investigação, resulta de uma abordagem empírica e teórica ao tema central das emissões de GEE.

Nos capítulos 1 e 2 investiga-se empiricamente a existência da hipótese da Curva Ambiental de Kuznets (CAK) para descrever a relação entre as emissões de CO_2 e o crescimento económico para Portugal, entre 1960 e 2010. É um estudo inovador, pois incide simultaneamente numa análise de âmbito nacional e setorial – produção de eletricidade e setor dos transportes – para examinar a hipótese da CAK, com recurso a metodologias de cointegração não linear e de cointegração com quebras de estrutura. A hipótese da CAK é alargada a outras variáveis explicativas, para além do produto interno bruto (PIB), nomeadamente, o preço do crude, o preço médio dos combustíveis, a taxa de motorização, a temperatura média e a precipitação média. Que saibamos, esta tese é a primeira a incluir como regressores a taxa de motorização e a precipitação

A hipótese da Curva Ambiental de Kuznets relaciona a qualidade ambiental com o crescimento económico. No início dos anos 90, vários investigadores como por exemplo Grossman e Krueger (1991), Shafik e Bandyopadhyay (1992), e Panayotou (1993), identificaram uma configuração de U invertido entre a degradação ambiental e o rendimento per capita. A esta relação, Panayotou (1993) designou de Curva Ambiental de Kuznets.

A hipótese da CAK é, essencialmente, empírica, mas diversos autores têm procurado fundamentações teóricas que a sustentem (e.g. Grossman e Krueger, 1991; Shafik e Bandyopadhyay, 1992; Panayotou, 1993; Panayotou, 1997; De Bruyn *et al*., 1998; Dinda *et al.*, 2000; Hettige *et al.*, 2000; Panayotou, 2003; Friedl e Getzner, 2003). É amplamente aceite que a relação em U invertido entre o crescimento económico e a degradação ambiental tem subjacente a ação de três efeitos: escala, composição e tecnológicos. A CAK resulta da interação destes três efeitos à medida que ocorre crescimento económico.

As mudanças estruturais na atividade económica são um processo dinâmico associado ao crescimento económico. Numa primeira fase, este crescimento traduz-se num aumento da atividade económica, e a passagem do setor primário para o setor secundário. A consequência desta transição é uma maior utilização dos recursos naturais como inputs dos processos produtivos, bem como um aumento dos níveis de poluição. Este mecanismo, em que se verifica uma degradação da qualidade ambiental, é identificado como efeito de escala. Com a continuação do crescimento económico, surgem os efeitos de composição e tecnológicos que conduzem a melhorias na qualidade ambiental. Estes efeitos aparecem com a transição para uma economia pós-industrial, em que o setor terciário é predominante. O peso das indústrias poluentes no PIB vai sendo reduzido, havendo uma substituição por indústrias assentes no conhecimento e na informação, que são menos poluentes e requerem uma menor necessidade de extração de recursos naturais (Panayotou, 1997; Panayotou, 2003). Esta transição consiste no efeito de composição. Os efeitos tecnológicos referem-se às melhorias dos processos de produção. O crescimento económico a par com a liberalização do comércio promove os investimentos em investigação e desenvolvimento levando ao progresso e inovação tecnológicos e ao aparecimento de tecnologias mais eficientes, contribuindo para uma diminuição da poluição e da extração dos recursos naturais.

O ponto de viragem da CAK acontece quando a magnitude dos efeitos de composição e dos efeitos tecnológicos superam a magnitude do efeito de escala (Dinda, 2004).

Para além dos efeitos de escala, composição e tecnológicos, a elasticidade rendimento da procura de qualidade ambiental é amplamente reconhecida como um dos fatores que também explicam a CAK. Para maiores rendimentos per capita, as pessoas tornam-se mais conscientes da importância da qualidade ambiental, estando dispostos a pagar mais para terem melhorias na qualidade ambiental (Roca, 2003), e exigem da parte das instituições um enquadramento regulatório mais rigoroso na proteção ambiental (Panayotou, 1993; Panayotou, 2003; Dinda, 2004).

No capítulo 1 recorre-se à metodologia da cointegração não linear, para aferir da possibilidade de existência da CAK, através de modelos quadráticos e cúbicos, quer em níveis quer em logaritmos naturais. As metodologias de cointegração são frequentemente utilizadas para analisar a hipótese da CAK para séries temporais. Como o modelo da CAK na sua forma paramétrica reduzida tem como regressores o PIB, o seu quadrado e, muitas vezes também, o seu cubo, surgem questões metodológicas que podem comprometer a fiabilidade dos resultados. A aplicação de técnicas de cointegração linear não é apropriada para modelos não lineares. Assim sendo, neste capítulo recorremos às metodologias de cointegração não linear desenvolvidas por Breitung (2001), e Choi e Saikkonen (2004; 2010), e aplicamos o procedimento de Hong e Phillips (2010) para testar a especificação dos modelos em análise. De acordo com este estudo, genericamente, a especificação cúbica em níveis é a que melhor descreve a relação entre as emissões de CO_2 e o PIB, quer a nível setorial quer a nível nacional.

A hipótese da CAK é validada para o setor de produção de eletricidade e para as emissões totais. Não obstante, as conclusões para as emissões provenientes da produção de eletricidade e as emissões totais são distintas. No caso da produção de eletricidade, os resultados indiciam que o ponto de viragem foi superado e que, portanto, o nível de emissões está posicionado na parte descendente da curva. Como tal, para este setor, as emissões de CO_2 já não estão ligadas ao crescimento económico. Para as emissões totais de CO_2, com base nos resultados obtidos, o ponto de viragem está muito próximo de ser alcançado, o que evidencia uma estabilização do crescimento das emissões.

Para os transportes, não existe um padrão de CAK, verificando-se que a relação entre as emissões de carbono e o crescimento é monotónica positiva.

Relativamente às restantes variáveis testadas como possíveis regressores, o preço do crude e a precipitação ajudam a explicar as emissões de CO_2 no setor de produção de eletricidade. No que concerne o setor dos transportes, para além do PIB, também o preço dos combustíveis contribuem para justificar as emissões de carbono. As emissões de CO_2 a nível nacional são explicadas pelos preços do crude e dos combustíveis e pela precipitação. A inclusão de variáveis explicativas adicionais não provoca alterações expressivas nos coeficientes estimados associados ao PIB.

No capítulo 2, utiliza-se a cointegração com quebras de estrutura desconhecidas para um modelo linear, também em níveis e logaritmos naturais. A hipótese de CAK descreve a relação de longo prazo entre a qualidade ambiental e o crescimento económico. Neste enquadramento, é provável que a relação entre as emissões de CO_2 e o PIB sofra alterações ao longo do tempo. Fatores legislativos, institucionais, políticos, desenvolvimento tecnológico, ou choques petrolíferos, podem induzir a quebras estruturais que afetam o padrão da ligação CO_2-PIB.

Negligenciar a ocorrência de quebras estruturais levanta problemas econométricos que podem induzir a conclusões erróneas. Conforme Perron (1989) demonstra, os testes de raiz unitária usuais tendem a rejeitar a existência de raiz unitária se a série em análise seguir um processo estacionário linear com quebras de estrutura. Kejriwal and Perron (2008, 2010) defendem que os testes de estabilidade podem rejeitar a hipótese nula de estabilidade dos coeficientes quando na realidade o que acontece é que se está na presença de uma relação espúria.

Perante estes factos, é pertinente analisar a relação CO_2-PIB à luz da metodologia da cointegração com quebras de estrutura. Para o efeito seguimos os trabalhos de Esteve e Tamarit (2012) e de Liddle e Messinis (2014), para avaliar a existência da CAK para os setores da eletricidade e transportes e também a nível nacional. Recorremos a modelos lineares para analisar a possibilidade de cointegração com múltiplas quebras estruturais desconhecidas. Para tal, utilizamos os testes de cointegração desenvolvidos por Gregory and Hansen (1996), e Arai e Kurozumi (2007). O número de quebras estruturais na relação de cointegração e a estimação das respetivas datas são obtidos com recurso à metodologia de Kejriwal e Perron (2010).

Os resultados obtidos devem ser encarados com prudência, pois a qualidade da análise de cointegração com quebras de estrutura depende da dimensão da série temporal, e na presente tese dispomos somente de 51 observações anuais o que pode condicionar a fiabilidade dos resultados. Ainda assim, foi possível alcançar resultados que contribuem para uma melhor compreensão da relação entre os níveis de emissão de CO_2 e o crescimento económico.

Para o setor de produção de eletricidade, para o modelo em níveis, é estimada uma alteração de regime em 1999. Esta quebra pode ter origem em vários fatores. Por um lado, a reduzida precipitação em 1998, o que levou a uma maior produção de eletricidade produzida a partir de combustíveis fósseis para fazer face à procura. Por outro lado, a inversão do ciclo económico em 1999. Acresce ainda o facto da data estimada ter capturado o efeito da criação do Programa Energia pelo Decreto-Lei 195/94 de 19 de julho e pela Resolução do Conselho de Ministros nº 68/94, que vêm materializar o apoio do Quadro Comunitário de Apoio (QCA) para o período de 1994 a 1999.

Com base na especificação em logaritmos, 1995 é a data estimada para a quebra estrutural. Este resultado é inconclusivo, pois as possíveis fundamentações são contrastantes. Esta quebra pode ser explicada pelo crescimento económico acentuado que se verificou com a adesão de Portugal à Comunidade Económica Europeia (CEE) em 1986. Todavia, uma outra explicação plausível é o abrandamento da economia em 1992 ou até mesmo o crescimento negativo do PIB em 1993.

O setor dos transportes aparenta ter três regimes para a especificação em níveis, com datas estimadas em 1993 e 1999. Para a especificação em logaritmos, apenas se identificam dois regimes com a data estimada em 1998. Apesar da diferença no número de quebras, as datas estimadas são coincidentes. Os QCA I (1986-1993) e QCAII (1994-1999) parecem estar na origem das alterações estruturais, devido ao forte investimento nas acessibilidades, sobretudo infraestruturas rodoviárias.

A avaliação a nível nacional, para ambas as especificações, as datas estimadas são as mesmas: 1991 e 1999. A primeira quebra estrutural pode estar relacionada com a quebra de 1993 no setor dos transportes; a segunda quebra pode dever-se à alteração do ciclo económico em Portugal registado em 1999. Sendo os setores de produção de eletricidade e dos transportes, as principais fontes de emissão de CO_2, é expectável que o CO_2 total tenha quebras próximas das destes setores. No entanto, é importante referir que para a contabilização das emissões totais de CO_2 estão incluídos outros setores que estão fora do âmbito desta tese.

O capítulo 3 consiste num estudo teórico em que é desenvolvido um modelo de Cournot de dois períodos para setor de produção de eletricidade. São considerados dois produtores, um que produz eletricidade térmica e outro que produz eletricidade a partir de fontes de energia renovável (E-FER). O objetivo deste modelo é analisar o comportamento estratégico dos produtores com e sem tarifas feed-in. Para além da quantidade no primeiro período, o produtor de E-FER tem de decidir o investimento em capacidade instalada para estar disponível no período seguinte. Seguidamente são calculadas as soluções do ótimo social para comparar com as soluções de Cournot-Nash.

Os principais resultados sugerem que devido ao aumento da procura no segundo período, as quantidades de ambos os produtores são superiores às do primeiro período. Relativamente aos níveis de investimento, este é mais elevado quando são aplicadas tarifas feed-in para promover a E-FER, o que leva a que o produtor de eletricidade térmica produza menos eletricidade quando comparado com a situação em que não existe este mecanismo de apoio à E-FER.

A comparação entre as soluções de Cournot-Nash e as do ótimo social revela, sobretudo, a importância da magnitude dos custos externos. Elevados custos externos levam a um excesso de produção de eletricidade térmica no contexto do modelo de Cournot. Quando os custos externos são moderados, por outro lado, o efeito da concorrência imperfeita prevalece e a quantidade de eletricidade socialmente ótima é superior à solução de Cournot-Nash. O nível de investimento e, consequentemente a produção de E-FER, no ótimo social é superior à verificada no modelo de Cournot.

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Contents

INTRODUCTION **1**

1. Economic growth and other determinants of Portuguese CO_2 emissions – A nonlinear cointegration approach **17**

1.1. Introduction 17

1.2. Data and unit root tests 20

1.2.1. Data 20

1.2.2. Unit root tests 22

1.3. Methodology under nonlinearities 28

1.3.1. Econometric specification 28

1.3.2. Cointegration analysis 30

1.4. Empirical results 40

1.4.1. Cointegration tests 40

1.4.2. GDP-CO_2 emissions' long-run relationship pattern 42

1.4.3. Extended model with other explanatory variables 47

1.5. Discussion 51

1.6. Concluding remarks 54

2. Structural breaks in the Economic growth - Portuguese CO_2 emissions relationship **57**

2.1. Introduction 57

2.2. Methodology under Breaks 59

2.2.1. Econometric specification 60

2.2.2. Cointegration tests 61

2.2.3. Structural break tests and correspondent estimated regression 64

2.3. Empirical results 65

2.3.1. Cointegration tests 66

2.3.2. Stability tests and structural break dates 67

2.3.3. GDP-CO_2 emissions' long-run relationship pattern 68

2.3.4. Extended model with other explanatory variables 70

2.4. Discussion 72

2.5. Concluding remarks 75

3. RES-E capacity expansion decisions in a duopolistic electricity market **77**

3.1. Introduction 77

3.2. The model 80

3.3. Results and discussion 81

3.3.1. Constrained RES-E generation 82

3.3.2. Constrained RES-E generation and feed-in tariff support scheme 84

3.4. Social optimum 86

3.5. A discussion on the potential role of electric vehicles 90

3.5.1. GHG emissions 90

3.5.2. Intermittent RES-E 91
3.5.3. Some insights from the literature for Portugal 92
3.6. Concluding remarks 93
FINAL CONSIDERATIONS AND FUTURE RESEARCH 97
Bibliography 99
APPENDIX A – DATA PLOTS AND UNIT ROOT TESTS 109
APPENDIX B – NONLINEAR COINTEGRATION TESTS 119
APPENDIX C – NONLINEAR COINTEGRATION REGRESSIONS (REDUCED FORM MODELS). 127
APPENDIX D – NONLINEAR COINTEGRATION REGRESSIONS (EXTENDED MODELS) 129
APPENDIX E – COINTEGRATION TESTS WITH STRUCTURAL BREAKS 131
APPENDIX F – STRUCTURAL BREAKS TESTS 135
APPENDIX G – COINTEGRATING REGRESSIONS WITH STRUCTURAL BREAKS (REDUCED-FORM MODELS) 139
APPENDIX H – COINTEGRATING REGRESSIONS WITH STRUCTURAL BREAKS (EXTENDED MODELS) 143
APPENDIX I – PUBLISHED PAPER 149

List of figures

Figure 1.1: Per capita real GDP (base year=2006) and per capita CO_2 emissions, from 1960 to 2010. 21

Figure 1.2: Per capita real GDP versus per capita CO_2 emissions from electricity generation sector for Portugal (1960 to 2010). 44

Figure 1.3: Per capita real GDP versus per capita CO_2 emissions from transport sector for Portugal (1960 to 2010). 45

Figure 1.4: Per capita real GDP versus per capita total CO_2 emissions for Portugal (1960 to 2010)... 47

Figure 1.5: Per capita real GDP versus per capita CO_2 emissions from electricity sector for Portugal with additional control variables (1960 to 2010). 48

Figure 1.6: Per capita real GDP versus per capita CO_2 emissions from transport sector for Portugal with additional control variables (1960 to 2010). 50

Figure 1.7: Per capita real GDP versus per capita total CO_2 emissions for Portugal with additional control variables (1960 to 2010). 51

Figure 1.8: Long-run relationship between per capita real GDP and per capita CO_2 emissions for Portugal (1960 to 2010). 53

List of tables

Table 1.1: Descriptive statistics of data 21

INTRODUCTION

Greenhouse gas (GHG) emissions, among which carbon dioxide (CO_2) is paramount, are a by-product of fossil fuel-driven economy, and have been growing steadily since the onset of the industrial revolution. The most important external cost imposed by carbon intensive activities, particularly electricity generation and transport sector, is climate change.

Climate change stems from the fact that both producers and consumers have taken into account only private costs or benefits of their fossil fuel based activities, respectively, which leads to an overproduction of GHG. Thus, as the market equilibrium does not meet the social optimum, significant external costs borne by society arise.

According to the Fifth Assessment Report of the United Nation's Intergovernmental Panel on Climate Change (IPCC) — Climate Change 2014: Impacts, Adaptation and Vulnerability — climate change is already occurring globally, increasing the frequency of extreme weather events, such as heavy rainfall, floods, heat waves and droughts. Higher or lower agricultural yields, species extinctions and increases in the ranges of disease vectors are among further consequences of climate change. Damages from weather-related catastrophes are on the rise and will, very likely, continue to do so. These weather events are expected to become ever more frequent and severe, although there is great uncertainty regarding future impacts (IPCC, 2013). As global climate change may threaten social well-being, intra and intergenerational equity, economic growth and ecosystems, tackling it is an urgent priority for the 21st century.

Following the Kyoto Protocol, the 2009 Copenhagen Accord set a maximum limit of 2 degrees Celsius, when compared with pre-industrial values, for the increase in mean global temperature. Although the accord itself is not legally binding, the European Union (EU) is committed to the implementation of measures that aim at the stabilization of GHG in the atmosphere. Through the implementation of various policies, including an Emissions Trading System for large industrial emitters, and within the context of the European 2020 strategy, the 28 member states of the European Union (EU 28) managed to reduce GHG emissions (excluding land use, land Use change and forestry, (LULUCF) by 19.2% from 1990 to 2012 (EEA, 2014). This downward trend was felt in every sector except transport (EEA, 2014).

The contrasting behaviour of transport GHG emissions begs for a sectoral analysis, especially of the two main emitting sectors: the Electricity and Heat Production sector and the Transport sector. In 2012, these two sectors were responsible for about 53.7% of all GHG emissions, in the EU-28 (IEAa, 2014). The Electricity and Heat Production sector was responsible for 34%, and transport sector responsible for 19.7% (EEA, 2014). In the same year, 82% of GHG emissions, excluding LULUCF, inside the EU-28, were CO_2 (EEA, 2014).

In the period between 2001 and 2011, the installed capacity of electricity generation plants went up by 31% in the EU-28. The profile of the power stations was altered in the first decade of this century. In 2001, thermal power stations, hydropower and nuclear power represented about 58%, 20% e 19%, of the total installed capacity, respectively. At this time, other renewable energy sources had a share of only 3%. Ten years later, thermal plants have a 53% share of the total installed capacity, and the other renewable energy sources (RES) fare second, representing 17%; after these comes hydro with 16% and, last, nuclear power, with a weight of 14% of the total capacity (Eurostat, 2013). Throughout this time period, electricity generated from hydro and other renewables increased by about 50%, while the installed capacity of thermal plants grew only 1%, and the power generation produced from nuclear energy decreased 7% (Eurostat, 2013). As a consequence, in 2011, the share of hydro and other RES to total EU-28 electricity generation reached 21%, while in 2001 that value was only 14% (Eurostat, 2013).

Regardless of the growing installed capacity of renewable energy power stations, and the decrease in CO_2 emissions, in 2011 the heat and electricity generation sector remained the greatest emitter of GHG in the EU**.**

In 2011, the transport sector was the second highest emitter of GHG, after power and heat production. In 2011 GHG emissions in the EU-28 transport sector increased 19%, when compared with 1990 levels. Although the growth trend in emissions in this sector was interrupted in 2008, because of a decrease in passenger and freight transport traffic due to the economic slowdown, growth is expected to resume as the economy recovers. Road transport was the mode of transportation that most stood out, as in previous years, since it was responsible for 71.9%[1] of the total final energy consumption of transports, most of it from fossil fuels (Eurostat, 2013). Road transport therefore remains the main source of GHG emissions within the transport sector, accounting, in 2012, for 95% of the total transport emissions within the EU-28 Member-States (Eurostat, 2013). In fact, there has been an increase of the motorization rate in almost all Member States, and the use of a private car continues to account for a significant share of the total passenger transport (Eurostat, 2013).

Within this framework, the main concern is to establish adaptation and mitigation policies, in order to manage, reduce and control the impacts and risks associated with climate change and global warming. Since climate change has distinctive negative consequences in the varied regions of the globe, and GHG emission trends are also different for the two major emitting sectors, climate change policies cannot be reproduced.

In order to have effective policies of climate-change mitigation, it is crucial that they are specific, not only for each region, but also for each sector. Additionally, such policies should take into account the relation between the GHG emissions and economic activity, lest their impacts be misunderstood.

In this context, the current thesis contributes to a better understanding of the relationship between economic growth and CO_2 emissions, focusing on electricity generation and transport sectors since they are the major contributors. The general objectives of this thesis are:

[1] http://ec.europa.eu/clima/policies/transport/index_en.htm

- To empirically investigate and compare the relationship between economic growth and CO_2 emissions for both the electricity generation and transport sectors, using Portugal as a case study. The underlying rationale is to better understand the behaviour of the economic growth-CO2 emissions link for each sector in order to help policy makers to design and implement more effective mitigation policies.
- To test additional variables that may act as determinants of CO_2 emissions besides the EKC hypothesis to tackle one of the major criticisms made to the reduced-form of the EKC hypothesis.
- To explore the strategic behaviour of thermal and RES-E producers and the impact on investment decisions in renewable generation capacity, with and without feed-in tariffs, taking into account environmental externalities.

To accomplish these goals, we split them into a set of sub-objectives to be achieved in chapters 1, 2 and 3.

Chapters 1 and 2

Chapters 1 and 2 carry out an empirical analysis of the Environmental Kuznets Curve (EKC) hypothesis for Portugal, from 1960 to 2010.

From 1990 to 2012, the total GHG emissions without LULUCF increased 13.1%. Similar to its European counterparts, in Portugal energy-related activities are the main source of GHG. In 2012 energy accounted for about 70% of total GHG emissions, which reflects an increment of 15% over the 1990-2012 period (APA, 2013).

In 2012, almost 73% of total GHG correspond to CO2 emissions and of these around 93% of came from energy use. The relevance of energy is explained by the fact that from 1990 to 2012, 83% of the primary energy consumed in the country was generated by fossil-fuel combustion (coal, oil and natural gas) while renewable energy sources represented 17% on average over the same period (APA, 2013).

Additionally, between 2001 and 2011, Portuguese energy intensity decreased by 8.5% while the EU28 average reduction was about 16%, and in 2012, Portugal dependency on energy imports in was near 80%, again, a value higher than the EU28 average (Eurostat, 2013).

Electricity and heat production and transport are the main emitters in the country. In 2012, Electricity and Heat and Transport were responsible for 39.2% and 34% of CO2, respectively (IEA, 2014).

Portugal has adopted a set of energy and environmental policies within the EU guidelines to achieve the commitments under both the Kyoto Protocol and the EU Climate and Energy Package 20-20-20.

These policies are mainly visible from the beginning of the twenty-first century, and aim at the reduction of energy dependence, the promotion of renewable energy sources, including the use of biofuels, increased energy efficiency and security of supply. In what follows we highlight a few measures implemented by Portugal since 2000.

In 2001, the E4 Program, Energy Efficiency and Endogenous Energies was approved. The energy objectives are reinforced within the single energy market through the adoption of the first National Strategy for Energy in 2005.

In 2010 the National Energy Strategy 2020 came into force with the following established goals for 2020: 1) reduce energy dependence of the country to the exterior to 74%, 2) obtain 60% of electricity and 31% of final energy consumption from renewable sources, 3) reduce final energy consumption by 20%, 4) reduceof final energy consumption in the transport sector by 10%, and 5) maintain the contribution of biofuels to meet the renewable energy targets in the transport sector.

Since 2001, programs and plans have been developed (reviewed periodically) that contribute to attain specific goals:

- National Programme for Climate Change (NPCC)
- National Plan for Emissions Allocation Plan (NAP)
- National Action Plan for Energy Efficiency (NAPEE)
- National Action Plan for Renewable Energy (NAPRE)
- Energy Efficiency Program for Public Administration (ECO.AP)

These measures make clear the strong commitment of Portugal towards a more sustainable energy sector, including electricity generation and transport sector. Despite energy and environmental policies take into account both sectors, measures targeting the electricity sector have been much more demanding, with better instruments, when compared to transport.

The two chapters use different methodologies to assess the relationship between carbon dioxide and economic growth and to test if the EKC hypothesis holds. The data is common to both. Thus, we decided to review the EKC literature and to present the time series variables in this section.

Since the 1990's, EKC have been widely used to empirically probe the link between economic growth and environmental quality. The initial literature was mostly empirical and it presented the tantalizing suggestion that economic growth, albeit environmentally damaging at first, actually contributes to better environmental outcomes as countries keep growing. A large set of EKC studies, considering various pollutants in a number of countries, studied individually or in groups, have been undertaken and are seen to be fairly inconclusive, with contradicting results. Theoretical reasons and econometric issues have been pointed out, by several authors, as the main justifications behind this result diversity.

We employ two distinct econometric methodologies. In the first chapter, we apply a nonlinear cointegration approach for the quadratic and cubic specifications, both in levels and natural logarithms. The goal is to assess which specification is more appropriate to describe the income-CO_2 emissions relationship. Chapter 2 focus on the possible regime shifts of the long-run relationship. To do so, we do a cointegration analysis with unknown structural breaks, applied to a linear specification, both in levels and natural logarithms.

The reduced-form models are extended to add economic and climate regressors, specifically:

- Average crude oil prices and average fuel prices – The use of fossil fuels is the major cause for GHG emissions, including CO_2, thus it is of major importance to study the impact of price changes in emissions levels. It is expected that a price increase will enhance the substitution of fossil fuels by renewable energy sources to electricity generation. As for the transport sector, higher fuel prices will lead to alternative mobility patterns concerning, for instance, the choice of mode of transportation and the frequency of long distance trips. Other authors have already considered crude oil prices (e.g. He and Richard, 2010; Saboori *et al.*, 2014) and fuel prices (e.g. Kumar and Viswanathan, 2004; Liddle, 2015) as additional explanatory variables of the EKC hypothesis for GHG.
- Rate of motorization – This concept is defined as the number of passenger cars per 1,000 inhabitants and it is an indicator of economic development and environmental matters. As far as one can tell from the EKC literature, this is the first study that includes this variable. Due to the lack of data on distance traveled per person, we use motorization rate as a proxy. We are assuming that higher rates of motorization translate into more travels, which may be a fragile assumption. Still, by including this variable we want to test how CO_2 emissions react to variations in the rate of motorizations and, consequently, to travel behaviour. Higher levels of income ought to lead to more kilometers travelled, thereby increasing emissions.
- Average temperature – The use of temperature as an explanatory variable is justified by the use of cooling and heating systems for buildings powered by electricity. In order to meet the increase in electricity demand due to temperature, we need more electricity supply which may have effects on emissions levels. However, it should be point out that the impact of temperature on carbon emissions depends on the electricity supply mix and it is more likely to happen in regions with extreme temperatures which is not the Portuguese case that has a mild climate. This variable has already been tested in the EKC context for GHG (e.g. Friedl and Getzner, 2003; Mota and Dias, 2006). We address how changes in temperature affect the CO_2 from electricity generation, through increasing electricity demand.
- Average precipitation – To the best of our knowledge, the present work is pioneer in using precipitation as a regressor. Hydroelectric power is the largest source of emissions-free in the world. However, its availability depends on the precipitation levels. In dry years it is greater the amount of electricity generated by thermal power plants, resulting in an increase of GHG emissions. Precipitation is particularly relevant for countries such as Portugal, where hydroelectricity represents a significant share of the total installed capacity and there are no nuclear power plants. We examine how average precipitation impacts the CO_2 emissions from electricity generation.

Environmental Kuznets Curve theory

The Environmental Kuznets Curve (EKC) studies the income – environmental quality relationship. In the early 1990s, several researchers, such as Grossman and Krueger (1991), Shafik and Bandyopadhyay (1992), and Panayotou (1993) identified an inverted U-shaped pattern between environmental degradation and per-capita income. Panayotou (1993) was the first to designate this specific relation as the Environmental Kuznets Curve, due to the previous work of Simon Kuznets, who postulated an identical relation between income inequality and per-capita income in the 1950s.

Since the 1990s and for a large number of countries, much empirical research has been carried out on different types of pollution, using time-series, panel data and cross-section methodologies. The EKC theory states that environmental quality changes over time, together with economic growth. The early stages of economic growth encompass the deterioration of environmental quality; however, beyond a certain level of income, the environmental degradation starts to decline. Thus, there is an EKC-turning point, after which economic growth has a positive impact on environmental quality.

Several authors tried to identify the underlying economic factors of the inverted U-shape for EKC. Grossman and Krueger (1991) were the first ones, followed by others (Shafik and Bandyopadhyay, 1992; Panayotou, 1993; Panayotou, 1997; De Bruyn *et al.*, 1998; Dinda *et al.*, 2000; Hettige *et al.*, 2000; Panayotou, 2003; Friedl and Getzner, 2003). Most of these authors have come to the conclusion that the role of economic development on the environmental quality may be explained by three effects: scale, composition (or structure) and technical effects. Economic growth takes place in stages and the interplay of the three effects, at each moment, is crucial for the inverted U-shape.

The first phase of economic growth implies an increase in economic activity, and a shift from the primary to the secondary sector. All else being equal, the consequence is the use of more natural resources as inputs in the production process, as well as an increase in pollution. This mechanism, harmful for the environment, is called the scale effect, and it explains the upward trend of the EKC. As economic growth continues, the composition and technical effects emerge, leading to positive impacts on the improvement of environmental quality. Economic structural change is a dynamic process linked to economic development (Panayotou, 1993). In the first phase, the composition effect accelerates environmental degradation, as the structure of the economic activity shifts from the primary sector to the secondary sector (which include heavy polluting industries). In the course of economic growth, structural changes take place, and the weight of industrial activities in gross domestic product (GDP) starts to decline, as the pre-industrial economy is replaced by a post-industrial economy, characterized by knowledge-based industries and services which are less polluting (Panayotou, 1997; Panayotou, 2003).

The technical effect relates to improvements in production processes. Wealthier countries and trade liberalization promote R&D investments leading to innovation, technological progress and more efficient technologies. Obsolete technologies are replaced by cleaner and more efficient ones, reducing both the use of resources and the levels of pollution in goods production.

As Panayotou (1993) stresses: *"The faster economic growth is, the faster the structural change that propels industry from a minor to a dominant sector of the economy, and from light, through heavy, to technologically sophisticated industry."* (Panayotou, 1993, pg. 3)

Together, the composition and technical effects overcome the scale effect and reverse the slope of the EKC (Dinda, 2004).

Additionally to scale, composition and technical effects, income elasticity of environmental quality demand is broadly recognized as one of the factors that also explains the shape of the EKC. At higher per-capita income levels, people become increasingly aware of environmental quality. When a certain level of income is attained, the willingness to pay for cleaner environment rises by a greater proportion than income (Roca, 2003). Moreover, consumers with higher incomes enforce environmental protection and regulatory frameworks (Panayotou, 1993; Panayotou, 2003; Dinda, 2004).

Andreoni and Levinson (2001), derive an explanation for the EKC directly from the technological link between consumption and abatement of pollution. They postulate that in richer economies the optimal scale of operation is larger which give rise to efficiencies in abatement. These efficiencies allow abatement at average lower costs even though pollution policy is maintained. Thus, the fact that abatement efficiency increases with increases in the scale of abatement, is a sufficient condition for an EKC.

Some authors include other explanatory variables besides income – e.g. international trade, demography, energy consumption, nuclear energy for electricity generation, share of service sector – to capture the relation between environmental quality and economic growth. International trade has been extensively discussed. In the long-run international trade contributes to environmental quality improvement because of its link with scale, composition and technique effects. In short and medium-run, the scale effect of the EKC, international trade, through exports, drives economic growth and pollution rise. Notwithstanding, trade has also an important role to play in fostering environmental quality by means of composition and technique effects, and regulatory framework. Tighter environmental regulation and protection claimed by wealthier consumers push forward pollution-reducing innovation, efficient production processes and changes in trade patterns. There is a transition from pollution-intensive industries to cleaner technology industries, knowledge-based industries, and service-oriented economy. Heavy polluting industries relocate away from high-income countries with strict environmental regulations to less developed economies with weaker environmental protection laws. This international migration of heavy polluting industries is known as the pollution haven hypothesis (PHH). The PHH states that loose environmental regulation provides a comparative advantage in itself (Dinda, 2004).

Brock and Taylor (2010) design a green Solow model which is the extension of the Solow's economic growth model by including the assumptions that pollution is a by-product of the production process and that the levels of pollution can drop because a constant share of economic output is spent on pollution abatement. They conclude that the ratio of pollution to per capita income in the optimal path first increases and after attaining a certain threshold it begins to decrease. By incorporating technological development in abatement, the EKC becomes an inevitable by-product of convergence towards sustainable growth.

Smulders et al. (2011) develop an endogenous growth model in which the EKC is caused by modifications in the scale, technology and composition of production. In their model, as new production technology emerges, the levels of a certain pollutant increase. This pollution increase raises consumers' awareness and leads to the imposition of an emission tax. The tax implementation leads to a decrease of pollution which then remains constant prior to the invention of a cleaner technology. In this situation, one may say that the EKC shape is generated by the replacement of one pollutant by another. The downward slope for the regulated pollutant occurs because its source is substituted by an alternative technology that emits another pollutant.

Note that, for some pollutants all of the causes described above may serve to justify the EKC. Notwithstanding, the EKC for certain pollutants can be originated by only a single or a few of the reasons stated above. Moreover, the relationship between economic growth and environmental degradation has been intensively studied through the use of several indicators for pollutants and the depletion of natural resources, in different regions.

Before reviewing some EKC literature for carbon dioxide emissions, we present the groundbreaking studies of Grossman and Krueger (1991), Shafik and Bandyopadhyay (1992), and Panayotou (1993), who were responsible for the creation of the EKC theory, and triggered the interest of the scientific community in empirically assessing the existence of the inverted U-shaped for other environmental variables and regions.

Grossman and Krueger (1991) modelled the economic growth and air pollution relationship for urban areas of 42 different countries, as part of a study to evaluate the consequences of the North American Free Trade Agreement on Mexico's environment.

The estimated regression included, as regressors, the per-capita GDP, time trend, city and site-specific characteristics, and trade intensity. The pollutants were sulphur dioxide (SO_2), dark matter and suspended particles. The main discovery by these authors was that it is the cubic function that best described the relationship between pollution and GDP.

The estimation results for both sulphur dioxide and dark matter indicated an N-shaped relationship between these and GDP. Initially, at low levels of national income, the concentration of these pollutants goes hand in hand with the per-capita GDP growth; however, when higher levels of income are attained, the concentration of these pollutants starts to fall. The turning-point for both pollutants is when income levels rise to about $4,000-5,000.

Subsequently, at income levels over $10,000-$15,000, the concentrations of SO_2 and dark matter become stationary or even rise again. The N-shaped relationship does not hold for mass suspended particles in the air. The concentration of this pollutant decreases as the per-capita GDP increases, at early stages of economic development; this relation is maintained until the per-capita GDP reaches $9,000, after which economic growth no longer has any impact on the concentration of mass suspended particles. In general, Grossman and Krueger's results suggest that economic growth tends to improve environmental quality.

Shafik and Bandyopadhyay (1992) carried out several regressions using log-linear, log-quadratic, and log-cubic polynomial in per-capita GDP, and a time trend. Their work is part of the 1992 World Development Report. They used observations from up to 149 countries for the 1960-90 period, where the environmental variables were: lack of clean water, lack of urban sanitation, ambient levels of suspended particulate matter, ambient sulphur oxides, change in forestation from 1961-86, annual rate of deforestation from 1961-86, dissolved oxygen in rivers, faecal coliforms in rivers, per-capita municipal waste, and per-capita carbon emissions. The authors also included other regressors such as trade orientation, electricity prices and site-specific characteristics. The estimation results were contradictory, depending on the type of environmental indicator. Only the ambient levels of suspended particulate matter and ambient sulphur oxides were in accordance with the EKC shape. Concerning the lack of clean water and lack of urban sanitation, the results showed that there was a uniform decline, over time, as income rised. The results for both deforestation indicators were not statistically significant.

Regarding river quality indicators, the results suggested that higher incomes led to an increase in environmental degradation. As for per-capita municipal waste and per-capita carbon emissions, Shafik and Bandyopadhyay found a positive monotonic relationship between these two indicators and income increase.

In contrast to the previously mentioned studies of Grossman and Krueger (1991), and Shafik and Bandyopadhyay (1992), Panayotou (1993) employed only cross-sectional data to estimate EKCs for both developed and developing countries. The environmental indicators were sulphur dioxide, nitrogenous oxides (NO_x) and solid particulate matter (SPM) – data from 54 countries – and annual rate of deforestation in the mid-1980s – data from 68 countries. The author used a log-quadratic functional form in per-capita GDP for pollutants, and a translog function in population density and per-capita GDP and a dummy variable for tropical countries. For all the environmental indicators, results suggested an inverted U-shape. The turning-points were $3,000; $5,500; $4,500; and $823 for SO_2, NO_x, SPM, and deforestation, respectively.

The substantive findings of these three pioneer studies remain critical for the current EKC knowledge. On the one hand, they were able to identify, for different environmental indicators, the EKC pattern. On the other, the results allow us to verify the ambiguity in the relation between economic growth and environmental quality. The divergence of results for the same environmental indicator remains visible in more recent analyses, regardless of the econometric methodology and samples chosen.

EKC literature for CO_2 emissions

Even though the first empirical EKC-studies on CO_2 emissions go back to the beginning of the 1990s, the growing concern with climate change has pushed for the application of the EKC theory, in order to better understand the connection between economic growth and CO_2 emissions. In the present Section, we mention but a few illustrative references of the extensive literature on this subject. The conclusion we come to is that, regardless of the econometric methodology and data sources used, the empirical findings are inconclusive as to the EKC-pattern for this pollutant. In the literature, we may find three types of results: 1) results that totally refute the EKC shape for CO_2 emissions, 2) ambivalent results, and 3) results that confirm the EKC theory.

Since the 1990s, several studies have been published whose empirical outcomes refute the EKC for CO_2 emissions. Holtz-Eakin and Selden (1995) were some of the first authors to empirically test the EKC for CO_2 emissions. They used a panel data with country and year fixed effects, for 130 countries; however, complete data was only available for 108 countries, for the period between 1951 and 1986. To estimate the per-capita CO_2 emissions and per-capita GDP (constant 1986 US dollars) relationship, they used two models: one that is quadratic in levels and another that is quadratic in natural logarithms (log). For both models, they were able to identify an EKC pattern for the per-capita CO_2 emissions – per-capita GDP nexus. However, the turning point assumed very different values for both models. The quadratic model in levels revealed a turning point at per-capita US$35,428, while for the natural logarithmic specification, the turning point occurred at a per-capita income above US$8 million. Since the turning point values were higher than the sample per-capita GDP values, the authors concluded that, despite the existence of the inverted U-shape, in fact, and for the sample period, the CO_2 emissions accompanied economic growth, which denotes a monotonic relationship.

Claiming that an EKC pattern using panel data does not have to hold for single countries, over time, De Bruyn *et al.* (1998) chose the time-series approach, which included energy prices, as an additional explanatory variable to account for the intensity of use of raw materials. This alternative growth model was estimated for CO_2, NO_x and SO_2 pollutants in four countries – the Netherlands, United Kingdom (UK), United States and West Germany – between 1961 and 1990. Despite using a different methodology from the one used by Holtz-Eakin and Selden (1995), the authors arrived at similar conclusions for each country, i.e., they were able to identify a linear positive relationship between per-capita emissions of the three pollutants and per-capita GDP. The authors also stressed that emission reductions might have been reached as a result of structural and technological changes to the economy.

Roca *et al.* (2001) also chose to do a single-country empirical study. They contributed to the EKC debate by looking into the annual emission flux of six atmospheric pollutants – CO_2, methane (CH_4), nitrous oxide (N_2O), SO_2, NO_x and non-methane volatile organic compounds (NMVOC) – in Spain. The time period for CO_2 went from 1973 to 1996, and for the remaining pollutants from 1980 to 1996. The authors used a cubic regression taking variables in natural logarithms, to test the evidence from the EKC. Except for SO_2, for which a U-shaped relationship with the per-capita GDP (thousands of 1986 Spanish peseta) was identified, for the remaining pollutants, the results showed an absence of correlation between higher income levels and reduced emissions. The authors were peremptory in claiming that the relationship between income levels and pollution emissions hinges on many variables, and that it is reckless to think that economic growth, on its own, is the solution for all environmental problems. They confirmed the monotonic increase of CO_2 emissions with income increase in Spain, and these results are in line with those obtained by Bruyn *et al.* (1998) for the Netherlands, UK, United States and West Germany.

The references presented so far represent a large part of the literature that suggests that CO_2 emissions tend to increase monotonically with economic growth. However, some studies present mixed results. Examples are the works of Tucker (1995), Richmond and Kaufmann (2006), and Galeotti *et al.* (2006). Tucker (1995) looked at the per-capita CO_2 emissions – per-capita GDP relationship for a sample of 137 countries, from 1971 to 1991. The author used a cross-section quadratic model at first differences for each year. The results revealed several interesting findings. There was a positive relationship between CO_2 emissions and per-capita GDP, for the linear term of the regression, for all years analysed. This positive relationship, as years go by, became increasingly more pronounced. During the first few years of the samples – early 1970s – the relationship was temporarily weakened, most likely due to the increase in oil prices. As for the quadratic term of the regression, results differed, depending on the year being analysed. The quadratic term had a negative coefficient for 11 out of the 13 years where it was statistically significant, sustaining the EKC hypothesis for these 11 years. The author attributed the negative coefficient to an increase in the demand for tighter environmental regulation and protection. Despite reporting an inverted-U curve that rises in statistical significance over time - essentially during the course of 1980s - the author stressed that such results do not mean that CO_2 emissions decreased; rather, it indicates that there was a slowing down of the growth of CO_2 emissions for higher per-capita income levels. Richmond and Kaufmann (2006) used a panel data model, which included other explanatory variables besides income (constant 1996 US dollars), for 36 nations – 20 OECD countries and 16 non-OECD countries – to investigate the existence of an EKC for CO_2 emissions. They concluded that fuel shares, the specification for income, and the degree of economic development influence results regarding the existence of an EKC between CO_2 emissions and income. More, the results obtained clearly separate developed countries (OECD) and developing ones (non-OECD). For OECD countries they were able to prove, with limited support, the inverted U-shaped relationship, with a turning point of per-capita \$25,500. For developing countries, the existence of an EKC relationship was rejected; alternatively, a positive CO_2 emissions-income relationship was proved. Due to the disparity of the results, Richmond and Kaufmann (2006) stated that policy makers should not make CO_2 emissions savings policies depend on reaching of a turning point in the relationship. Galeotti *et al.* (2006), decided to assess how robust the mixed results were, regarding the existence of an EKC for CO_2 emissions. For this purpose, the authors, on the one hand, used the alternative emissions data source (International Energy Agency), and on the other hand, proposed a non-linear functional form, other than the normally used for the EKC analysis, which consists of a generalization of the three-parameter Weibull function. The model was run from 1960 to 1998 for the Annex II countries of the United Nations Framework Convention on Climate Change, and from 1971 to 1998 for all other countries. The empirical study was carried out independently for two samples, one including the high-income countries and another that encompasses low-income nations. These were the two main conclusions of this research: first, the evidence on the EKC for CO_2 emissions didn't seem to depend on the data source and second, for high-income countries, the relationship between per-capita CO_2 emissions and per-capita GDP revealed an inverted-U pattern, with a reasonable turning point. For low income countries, the EKC shape wasn't so pronounced, showing only a slight concavity.

There are also numerous empirical analyses where researchers argue that there is evidence supporting the EKC relationship. Schmalensee *et al.* (1998), in a paper where they tried to assess the existence of an EKC for CO_2 emissions, adopted a spline regression with ten piece-wise segments with fixed year- and country-specific effects, for a dataset of 141 countries, including non-OECD countries, for the period between 1950 and 1990. They reported an inverted-U shaped relationship between per-capita CO_2 emissions and per-capita GDP (1985 US dollars), with a turning point that takes place in an interval of values of per-capita income between US$ 10,000 to US$ 17,000. For most countries in the sample, the peak of per-capita CO_2 emissions took place in 1985.

This brief EKC literature review allows us to confirm conflicting results. Even empirical works that study the EKC hypothesis for CO_2 emissions, both on individual country-level and on group of countries, show contradicting outcomes. The literature has yet to provide a definitive conclusion to this relationship.

Nonetheless, in most empirical analysis on CO_2 it predominates a monotonically increasing relation. Lieb (2004) gives an explanation for the monotonic pattern. According to this author, because CO_2 has a long lifetime, it is considered a stock pollutant. Other stock pollutants, such as sulphur dioxide or carbon monoxide, are considered flow pollutants from a long-run perspective since they all have short lifetimes (Lieb, 2004). Therefore, the negative effects on the environment will be more evident in the future as for this types of pollution the environmental damage occurs mainly due to accumulation. Thus, governments do not have the incentive to reduce stock pollutants emissions because most of the benefits will be felt by future generations (Lieb, 2004).

Electricity generation sector and the EKC hypothesis

Since the energy sector is the main emitter of GHG, several EKC-works have included energy consumption as an additional variable. However, there are very few studies that assess the production side of electricity in the behaviour of CO_2 emissions.

Iwata *et al.* (2010) examined the EKC for CO_2 emissions in France, between 1960 and 2003. The authors included other determinants of CO_2 emissions besides income, namely nuclear energy (percentage of the total electricity produced), trade, and energy consumption. They adopted a quadratic log-log model and applied the autoregressive distributed lag (ARDL) methodology for cointegration analysis. The estimation results confirmed the EKC pattern for the relationship between per-capita CO_2 emissions and per-capita GDP (constant local currency). As for the consequences of having electricity generated from nuclear energy, these were significantly negative, both in the short-run and the long-run. The estimated long-run elasticity of CO_2 emissions on nuclear energy was between -0.27 to -0.31, which means that every 1% increase in nuclear-sourced electricity leads to a decrease in per-capita CO_2 emissions of about 0.27 to 0.31%.

As for the impact of energy consumption on CO_2 emissions, the authors found evidence of statistical significance in the short-run, but not in the long-run. These results lead to two main conclusions. First, nuclear energy for power generation plays a relevant role in decreasing CO_2 emissions. Second, a possible explanation for the fact that the energy consumption coefficient is not statistically significant in the long-run, is that while energy consumption leads to an increase in CO_2 emissions, nuclear energy emits fewer CO_2, and the effect of the nuclear energy on CO_2 emissions seems to be greater than energy consumption. Lastly, the results of the trade impact on CO_2 emissions proved not to be statistically significant in both the short-run and the long-run.

At a later date, Iwata *et al.* (2011) used a dataset of 28 countries – 17 OECD and 11 non-OECD countries – to validate the EKC for CO_2 emissions and the role of nuclear power. They used a log-log quadratic model, and employed the pooled mean group estimation method. The authors' findings suggested a monotonic increase of the per-capita CO_2 emissions as the per-capita GDP increases. This relationship was valid for all three cases – the full sample, OECD countries, and non-OECD countries. Although the EKC pattern was verified, the turning point was outside the sample range, and therefore the authors interpreted this result as a positive monotonic relationship. As expected, there was evidence of the negative impact of nuclear energy for electricity production, on CO_2 emissions. The reduction in CO_2 emissions as a reflection of the increase in electricity produced from nuclear energy is greater for non-OECD countries. However, that does not mean that the increase in the share of nuclear energy in electricity generation contributes to the existence of an EKC. The rate of economic growth in non-OECD countries is very fast, which translates into an increase in the secondary sector, in both economic activity and increased consumption. Given that more efficient technologies are not adopted by these countries, CO_2 emissions will increase, despite the CO_2 emissions savings derived from the nuclear power. These results highlight the importance of cooperation between OECD and non-OECD countries in the resolution of climate change, since the estimation outcomes show that nuclear power, on its own, is insufficient in decreasing CO_2 emissions, despite playing an important role in doing so (Iwata *et al.*, 2011).

Burke (2012) resorted to a binomial dependent variable model to assess the weight of decarbonisation of electricity generation in the appearance of an EKC for CO_2 emissions for a sample of 105 countries. For this author, the fact that the EKC pattern for CO_2 emission-income nexus relationship is valid for some countries, but not others, has to do, primarily, with primary energy sources used to produce electricity. In this paper, Burke puts forward the hypothesis that the extent to which the electric power supply shifts from carbon-intensive power production (e.g. coal) towards low-carbon power production technologies (nuclear and renewable, excluding hydro) is the key determinant of whether that country has experienced an EKC-shape for CO_2 emissions is verified. In his previous work, Burke (2010) stressed that the shift to the use of nuclear power and renewable sources of energy is directly related to an increase in income. The evidence found supports the formulated hypothesis, confirming that the increase in the electricity generation quota, produced from nuclear power and renewable energy, is the main determinant for the downward slope of the EKC for CO_2 emissions, in the long-run.

Sulaiman, Azman and Saboori (2013), apart from testing the existence of an EKC for Malaysia, looked at the impact of renewable energy sources for electricity generation (RES-E) in CO_2 emissions, and trade openness, during the period between 1980 and 2009. The authors employed a linear logarithmic quadratic model and the ARDL approach. The estimation results concerning the EKC hypothesis support the existence of an inverted U-shaped relationship between per-capita CO_2 emissions and per-capita real GDP (constant 2000 US$). One must interpret these results cautiously, as the turning point was 8.77 US$, which is above the highest value of per-capita real GDP - 8.49 US$ - of the sample. When it comes to the impact of RES-E on CO_2 emissions, both the short-run and the long-run elasticity of CO_2 emissions with respect to electricity production from renewable sources are negative and statistically significant. According to the estimation results for long-run elasticity, the coefficient value is 0.11, i.e., a 1% increase in per-capita RES-E translates into a decrease of 0.11% in CO_2 emissions. These findings are consistent with the results of Iwata *et al.* (2010) and Iwata *et al.* (2011) for nuclear power generation. Trade openness has a significant negative impact on CO_2 emissions, in the long-run.

To sum up, the central conclusion we may draw from the few existing empirical studies is that, in effect, the type of primary energy resources for electricity generation has an impact not only on the levels of CO_2 emissions, but it also contributes to the descending part of the EKC shape.

Transport sector and the EKC hypothesis

Together with the electricity power generation, the transportation sector is one of the main emitters of GHG, and yet, the EKC literature focusing on the relationship between transport-related gas emissions and economic growth is still at an embrionic stage.

Liddle (2004) performed an ordinary least squares (OLS) regression, fixed and time effect regressions, with time dummies, on a data panel of 23 OECD countries, with observations being taken for five time periods – 10-year intervals – from 1960 to 2000. The goal was to investigate the EKC relationship between per-capita road energy use and per-capita GDP. The author also included geographic and demographic variables. Three models were estimated: a quadratic model in levels, a quadratic log-log model, and a lin-log model. For all the models, the EKC hypothesis was rejected, because the parameters on the GDP squared terms were not statistically significant, and the turning point values were well above the sample range. Therefore, the relationship between per-capita road energy use and per-capita GDP was monotonic.

Tanishita (2006) studied the evidence of an EKC for energy intensity from passenger transportation, using city-based data, for the period between 1980 and 1995. The results supported the EKC relationship between energy intensity of private and public transportation and per-capita Gross Regional Product (GRP), with a turning point that ranged from US$22,000 to $26,000 (PPP, 1995).

Ubaidillah (2011) explored this same relationship for the United Kingdom, from 1970 to 2008. Instead of using CO_2 as an environmental indicator, he took carbon monoxide (CO), as this pollutant is a more road transport specific indicator. A quadratic model in levels was employed, as well as the Johansen maximum likelihood methodology for cointegration analysis. Just like Tanishita (2006), the author found evidence of the EKC pattern. In this case, the per-capita CO – per-capita GDP (2000 constant price) relationship has a turning point of $21,402. For this author, the EKC pattern for the UK's road transport sector is explained by the increased usage of private or passenger vehicles, which follows the growth trajectory of income. In turn, the use of private or passenger vehicles translates into an increase in fuel combustion and, as a consequence, into an increase in CO emissions. However, from a certain level of per-capita income, an improvement in technology, rules and behaviours may be explained by the realisation that clean air is more important, and thus CO emissions start to decline.

Cox *et al.* (2012) resorted to a 2006 survey, undertaken in six case study areas in Scotland, to inquire about the existence of an EKC for household transport CO_2 emissions. The authors used a simple OLS regression of log household CO_2 emissions, and a household annual income dummy. The most important result in this study is that households with annual incomes equal to or greater than £52,000 produce 92% more CO_2 emissions than lower income households. On average, richer households have more than one vehicle, newer but not less polluting, and used more often. Therefore, richer families never truly become aware of the social costs of the pollution emitted by their vehicles. To conclude, the results obtained do not support the existence of the EKC for private road vehicles.

Abdallah *et al.* (2013) examined the relationship between transport value added, road transport-related energy consumption, road infrastructures, fuel prices and the CO_2 emissions for the transport sector, in Tunisia, from 1980 to 2010. Regarding the analysis of the relationship between per-capita CO_2 emissions from the transport sector and the per-capita transport value added, the authors selected a log-log cubic model and the Johansen cointegration methodology. The parameter of the cubic term is negative and statistically significant, suggesting that the relationship between transport value added and transport CO_2 emissions is described by an inverse N-shape, refuting the existence of the EKC. Despite these results, the authors believe that, in practice, the relationship is described by a monotonically increasing curve as, on the one hand, the first turning point is equal to 74.88 Tunisian national dinars (constant 2000 TND), which is a very low value and on the other, the second turning point – 578.82 TND – exceeds the dataset values. This result follows in the same direction as those of Liddle (2004).

Unlike the electricity generation sector, these results are different, as to the confirmation of a transportation-EKC for CO_2. We need to remember that while this sectoral analysis is still at an early stage, more robust conclusions demand more empirical analyses.

Sub-objectives of chapters 1 and 2

Throughout these empirical EKC studies, the sub-objectives for each sector and at national level are:

- to examine if the EKC hypothesis holds (chapter 1);

- to calculate the income turning point (when applied);
- to estimate the income elasticities;
- to test crude oil price, fuel price, rate of motorization, temperature and precipitation as explanatory variables of CO_2 emissions;
- to evaluate the stability of the CO_2-GDP relationship overtime and the dating of structural breaks (chapter 2);

The findings of the two chapters are complementary rather than substitutes. With both methodological approaches, we intend to shed light on the CO_2-GDP relationship.

Chapter 3

RES-E capacity investment is key to electricity generation decarbonisation and security of electricity supply in the context of imperfect competition. In chapter 3 a two-period Cournot model is developed building on the work of Genc and Thille (2011), to examine the strategic behaviour of both thermal and RES-E producers when the latter faces generation capacity investment decisions in deregulated markets with and without the adoption of feed-in tariffs. Besides the generation costs our model takes in consideration the environmental damages caused by the thermal power plant to compute the socially optimum solutions and compare them with the Cournot-Nash ones. A review of the literature on this type of electricity market models is provided within the chapter, which also includes some considerations on potential role of electric vehicles.

Sub-objectives of chapter 3

This game theoretical model allows attaining the following sub-objectives:

- to determine the Cournot-Nash solutions for both producers without feed-in tariffs;
- to calculate the Cournot-Nash solutions for both producers under feed-in tariffs scheme;
- to compare the Cournot-Nash solutions with the social optimum;

1. Economic growth and other determinants of Portuguese CO_2 emissions – A nonlinear cointegration approach[2]

1.1. Introduction

One of the major econometric critiques concerning the EKC hypothesis has been the prevalence of panel-data applications over the time-series approach, mostly justified by the inexistence of time-series data observed over a long period of time. With panel data the sample size increases as a result of the compilation of data from several countries at the expense of assuming homogeneity in the model. Within this framework, we have a single relationship between income and environmental quality which is valid for the whole set of countries. Nevertheless, this relationship does not mean that each individual country follows exactly that pattern (De Bruyn *et al.*, 1998). Hence, empirical EKC studies should be better circumscribed to a given country with longer time span data (De Bruyn *et al.*, 1998; List and Gallet, 1999; Dijkgraaf and Vollebergh, 2001, 2005; Jalil and Mahmud, 2009).

Within the time series approach, econometric methodology, should pay attention to the statistical properties of the series. In particular, whenever the series are found to have unit roots, cointegration analysis must be done. The EKC hypothesis uses on a second- or third- order polynomial function to proxy the income – environmental quality nexus, in which GDP and its square and/or cube are the explanatory variables. Since GDP and environmental variables are widely accepted as nonstationary and integrated of order one (Liu *et al.*, 2009), the majority of empirical EKC studies for a single country is based on cointegration techniques.

For instance, Friedl and Getzner (2003) did a time-series analysis of the EKC hypothesis for Austria, for the time period of 1960 to 1999. Besides per-capita GDP (constant 1995 euros), the share of imports, as a proxy for the pollution haven hypothesis, and the share of the service sector in GDP were both added to the model, to reflect possible structural changes in economic activity. The cubic parametric model was the one that best suited the relationship between the annual level of total CO_2 emissions and the per-capita GDP, and the Engle-Granger two-step cointegration analysis was carried out. According to the results, the relationship between Austrian CO_2 emissions and per-capita GDP follows an N-shaped pattern. Fodha & Zaghdoud (2010) also used a time-series approach to investigate the link between income and two pollutants – CO_2 and SO_2 – in Tunisia, for the time period 1961-2004, using a logarithmic cubic function and the Johansen maximum likelihood cointegration analysis. According to this study, the relationship between per-capita SO_2 emissions and per-capita GDP constant 2000 US dollars) may be described through an inverse N-shape, while the relationship between per-capita CO_2 emissions and per-capita GDP turned out to be monotonically increasing, once again. The authors justified these results, highlighting the fact that unlike what happens with SO_2, whose effects are felt mainly at the local and regional levels, CO_2 has a global impact. Nasir and Rehman (2011) conducted

[2] A partial version of this work was presented at the Global Conference on Environmental Taxation (Copenhagen, 2014) and published in *Natural Resources Forum* (Sousa *et al.*, 2015).

a time-series investigation on the existence of an EKC for CO_2 emissions in Pakistan, for the period between 1972 and 2008. Besides income (2000 US dollars), energy consumption and foreign trade were used as explanatory variables. The authors adopted the log-log quadratic model, and employed the Johansen maximum likelihood cointegration approach. The long-run results supported the EKC for CO_2 emissions in Pakistan, and the authors reported the positive effect of both energy consumption and trade openness on CO_2 emissions. On the contrary, short-run results failed to confirm the existence of the EKC. The authors alert to the fact that the contradiction between short-run and long-run results should be taken into consideration by policy makers. Shahbaz *et al.* (2013) assessed the impact of economic growth, financial development, coal consumption and trade openness on environmental quality, in South Africa, for the period between 1965 and 2008. The authors used a log-log quadratic model and the autoregressive distributed lag (ARDL) bounds testing approach to cointegration. According to the results, the EKC holds for the relationship between per-capita CO_2 emissions and per-capita GDP, with a turning point of per-capita US$3463. As to the relationship between CO_2 emissions and the other variables, the correlation between financial development and CO_2 emissions, and trade openness and CO_2 emissions, proved to be negative.

More recently, Lau *et al.* (2014) tested the validity of the EKC pattern in Malaysia for the period between 1970 and 2008. They used a log-log quadratic specification and the ARDL methodology for cointegration. Along with the per-capita GDP (2000 US dollars) and its square, foreign direct investment and trade openness were part of this model. The empirical findings, which are consistent for both short-run and long-run analysis, showed the presence of EKC-like behaviour between per-capita CO_2 emissions and per-capita GDP. Moreover, the authors also concluded that foreign direct investment and trade openness contributed to environmental degradation.

Portugal has occasionally been covered in the EKC literature, with contradictory results. Mota and Dias (2006) applied a time-series approach to look for an EKC for per-capita CO_2 emissions, between 1970 and 2000. They test for stationarity and cointegration, using the ADF and the Engle and Granger tests, respectively. The relation between per-capita CO_2 emissions and per-capita real GDP proved inconclusive, as both linear and cubic models present reliable econometric results, with the authors choosing the linear model. Regardless of which model best describes the CO_2-GDP link, an EKC does not appear to exist. They also identified a positive contribution from the service sector to CO_2 emissions. In order to explain this positive relation, the authors stress that the services sector encompasses the transport subsector. However, a sectoral analysis was not performed.

Acaravci and Ozturk (2010) perform an autoregressive distributed lag (ARDL) bounds cointegration analysis, using 1960 – 2005 data for Portugal, among others, and they also find that the EKC doesn't hold. Shahbaz *et al.* (2015) use the same ARDL methodology for Portugal, during the period of 1971-2008. Their results are in sharp contrast to those of Mota and Dias (2006) and Acaravci and Ozturk (2010), since they find that the EKC hypothesis holds.

This brief literature review shows that empirical evidence of an EKC for CO_2 emissions in Portugal is mixed. These contradictory results raise an important issue, since the key objective of an empirical analysis of the EKC for CO_2 is to bring new insights to climate policy design and implementation. Policy makers should formulate policies considering the pattern of the relationship between income and CO_2 emissions in order to establish appropriate targets and implement the most suitable policies to pursue then. Testing the EKC hypothesis goes beyond the simplistic interpretation that environmental problems are solved with economic growth.

Despite their widespread use, conventional cointegration methodologies applied to the parametric reduced-form model of the EKC raise econometric problems as it includes per-capita GDP and its squares and cubes as explanatory variables (Müller-Fürstenberger & Wagner, 2007). It is not that straightforward to assume that the properties of linear models with nonstationary variables are hold similarly in nonlinear contexts (Granger and Hallman, 1991; Ermini and Granger, 1993; Corradi, 1995; Müller-Fürstenberger & Wagner, 2007).

It is widely accepted that GDP is nonstationary. Thus, since per-capita GDP is an integrated variable, the powers of GDP are a nonlinear transformation of an integrated I(1) process, which in turn are not usually integrated (Müller-Fürstenberger & Wagner, 2007; Hong and Wagner, 2008). Therefore, EKC regressions cannot be analysed within the usual linear unit root and cointegration methodologies, as they call for an alternative type of asymptotic theory, and lead to different properties of estimators (Park and Phillips, 1999 and 2001; Chang, Park and Phillips, 2001). Nonetheless, most EKC models up to date neglect the methodological implications of this issue and investigate cointegration using traditional tests. In this context, a significant part of the empirical evidence of the EKC is doubtful.

Henceforth, in order to obtain reliable EKC outcomes and possibly confirm the existing ones, GDP and its integer powers should be considered and nonlinear cointegration methodologies applied. The absence of an appropriate nonlinear cointegration analysis can lead to misleading results. If the EKC regressions do not fulfil the cointegration property, then the estimates may be spurious (Aslanidis, 2009).

To overcome the econometric limitations of linear cointegration techniques such as Engle and Granger's two-step technique (1987) and ARDL or Johansen's (1995) maximum likelihood approach, this chapter will rely on nonlinear cointegration methodologies based on the procedures developed by Breitung (2001) and Choi and Saikkonen (2004, 2010) and Hong and Phillips (2010). Additionally, we include other independent variables – economic and climate variables – as we believe that they may act as important determinants of CO_2 emissions.

1.2. Data and unit root tests

1.2.1. Data

To conduct our empirical analysis we employ annual time series for Portuguese population, CO_2 emissions and real GDP at 2006 prices, from 1960 to 2010. Data for population and for CO_2 emissions were downloaded from the World Bank website. CO_2 emissions from electricity and heat generation (CO2elect) comprise the emissions from main activity producer electricity generation, combined heat and power generation and heat plants; from the generation of electricity and/or heat by autoproducers; and from fuel combusted in petroleum refineries, for the manufacture of solid fuels, coal mining, oil and gas extraction and other energy-producing industries. CO_2 emissions from transportation (CO2transp) encompass the emissions from fossil fuel combustion for all transportation activities, irrespective of the activity sector and excluding the International Marine Bunkers and International Aviation. Total CO_2 emissions (CO2total), in metric tons per capita, include those stemming from the burning of fossil fuels and the manufacture of cement, which includes CO_2 produced during consumption of solid, liquid, and gas fuels and gas flaring. Originally, CO2elect and CO2transp were in million metric tons. In order to have the emissions variables in the same units, they were converted to metric tons per capita. The time series for real GDP in euros is available at Pordata.

The crude oil annual average prices (crude) in real 2010 USD per barrel were also obtained from the World Bank website. The weighted annual average fuel prices (diesel plus gasoline) for Portugal are calculated using data from different sources. The annual average prices for diesel and gasoline, in euros at 2006 prices, were downloaded from Pordata and the road sector diesel and gasoline fuel consumption per capita in kg of oil equivalent were acquired from the World Bank website, and converted to liters.

As for climate variables, the average annual temperature (temp) in degrees Celsius and precipitation (precip) in millimetres were obtained from the Agência Portuguesa do Ambiente website (original data from Instituto Português do Mar e da Atmosfera).

The rate of motorization (number of vehicles/1000 habitants) is calculated considering the vehicle stock in Portugal provided by the Associação Automóvel de Portugal.

Table 1.1 shows descriptive statistics for both dependent and independent variables. The lowest level of per capita real GDP was observed in 1960, while the highest level was in 2007. Similarly, the minimum value of per capita CO_2 emissions was in 1961 for the electricity sector and 1960 for the transport sector and countrywide. However, the maximum values for CO_2 emissions and per capita GDP value do not match in time. One may interpret these disparities as a change in economic growth – CO_2 emissions nexus.

Table 1.1: Descriptive statistics of data.

Variable	Obs	Mean	Std. Dev.	Min	Max
GDP (€ per capita)	51	9587.528	4079.716	3135.867	15521.780
CO_2 emissions from electricity generation (metric tons per capita)	51	1.138621	0.866575	0.060475	2.6247880
CO_2 emissions from transport sector (metric tons per capita)	51	0.944879	0.568982	0.199826	1.8891690
Total CO_2 emissions (metric tons per capita)	51	3.533636	1.782134	0.928578	6.4445870
Crude oil price (USD/bbl)	51	29.57257	22.02735	5.171954	94.312320
Fuel price (€/L)	51	1.123167	0.194861	0.833869	1.6304340
Rate of motorization (number of vehicles/1000 habitants)	51	234.9306	181.9400	26.69988	548.31350
Temperature (ºC)	51	15.86920	0.573164	14.54100	17.211000
Precipitation (mm)	51	902.3633	220.6849	503.1000	1509.8000

Note: All data is in levels.

Figure 1.1 represents the evolution over time of GDP and CO_2 variables. Two comments can be drawn. First, we perceive a more similar pattern of fluctuations between CO2elect and CO2total than between CO2transp and CO2total, possible because the electricity generation sector is the greater contributor to total carbon emissions in Portugal. It is possible to see that around 1999, both sectoral and total emissions tend to stabilize and, around 2005 start to decline. Since the electricity generation sector has a large weight on national carbon emissions and it is highly dependent on precipitation patterns due to hydropower, the CO2elect and CO2total stabilization is seen as a fluctuation around a certain value.

Second, up to 2001 there is a steady growth of the per capita GDP, with the notable exception of the 1974 revolution period. After 2001, and especially since 2005, GDP growth is much more subdued. Moreover, real GDP and CO_2 emissions seem to co-move, in a nonlinear fashion, especially until 2005.

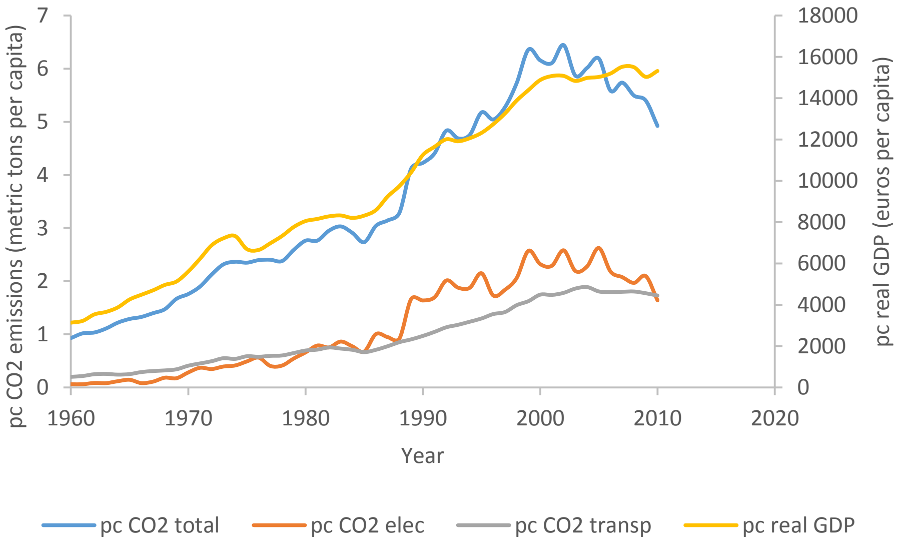

Figure 1.1: Per capita real GDP (base year=2006) and per capita CO_2 emissions, from 1960 to 2010.

The plots of the other explanatory variables are depicted in appendix A.

1.2.2. *Unit root tests*

We begin our empirical analysis by investigating the stationarity of the data, that is, by testing for the presence of unit roots for all the dependent and independent variables in levels. To do so, we consider six distinct tests in order to obtain more solid results – Augmented Dickey and Fuller (1979) (ADF), Philips and Perron (1988) (PP), Kwiatkowski, Phillips, Schmidt and Shin (1992) (KPSS), Dickey-Fuller Generalised Least Squares (DFGLS) test proposed by Elliott, Rothenberg, and Stock (1996), point-optimal Elliott, Rothenberg, and Stock (1996) (ERS) and Ng and Perron (2001) (Ng-PP) – with intercept only, and with intercept plus trend. The ADF, DFGLS, ERS and Ng-Perron tests are computed with the automatic lag length selection based on the Schwarz information criterion (maximum of 10 lags), and the PP and KPSS tests are run with automatic bandwidth selection. These tests enable to test as to whether the time series have unit roots or not.

Description of the unit root tests

Augmented Dickey and Fuller test

The Augmented Dickey and Fuller test (1979) corrects for serial autocorrelation in the errors of the auxiliary regression by adding lagged difference terms of the dependent variable – Δy_{t-i} – to the right-hand side of the test regression. The test equation is one of the following:

$$\Delta y_t = \alpha y_{t-1} + \sum_{i=1}^{p} \delta_i \, \Delta y_{t-i} + \varepsilon_t \tag{1.1}$$

$$\Delta y_t = c + \alpha y_{t-1} + \sum_{i=1}^{p} \delta_i \, \Delta y_{t-i} + \varepsilon_t \tag{1.2}$$

$$\Delta y_t = c + \delta t + \alpha y_{t-1} + \sum_{i=1}^{p} \delta_i \, \Delta y_{t-i} + \varepsilon_t \tag{1.3}$$

Where y_t is the observed time series, t is the linear time trend term and ε_t is the white noise error term with zero mean and constant variance. Equation (1.2) has an intercept term – c – representing a I(1) process under the null hypothesis and equation (1.3) is designated by an ADF test with drift and linear trend. The Dickey-Fuller test statistic does not follow a standard t-distribution and is left-tailed. The null hypothesis of a unit root, $\alpha = 0$, is tested against the alternative of $\alpha < 0$, which corresponds to a mean-stationary process when using equation (1.2) or to a trend stationary process when considering equation (1.3).

Philips and Perron test

Phillips and Perron (1988) nonparametric tests for unit roots are a modification and generalization of Dickey-Fuller's procedures which allows for general forms of serial autocorrelation.

The PP test equations can be written as follows:

$$\Delta y_t = \alpha y_{t-1} + \varepsilon_t \tag{1.4}$$

$$\Delta y_t = c + \alpha y_{t-1} + u_t \tag{1.5}$$

$$\Delta y_t = c + \delta t + \alpha y_{t-1} + u_t \tag{1.6}$$

The null hypothesis of unit root is again $\alpha = 0$ tested against the alternative of $\alpha < 0$.

The statistics are designated by Z_t and Z_α:

$$Z_\alpha = T(\hat{\alpha} - 1) - (s^2 - s_\varepsilon^2)(2T^{-2} \sum_{t=1}^{T} y_{t-1}^2)^{-1} \tag{1.7}$$

$$Z_t = \left({}^{s_\varepsilon}/_{s}\right) t_\alpha - \left({}^{1}/_{2}\right)(s^2 - s_\varepsilon^2)(s^2 T^{-2} \sum_{t=1}^{T} y_{t-1}^2)^{-1/2} \tag{1.8}$$

Where $\hat{\alpha}$ is the OLS estimate of α, T is the sample size, and s^2 and s_ε^2 are consistent estimators of the long and short run variances.

Kwiatkowski, Phillips, Schmidt and Shin test

Kwiatkowski, Phillips, Schmidt and Shin (1992) propose a test for the null hypothesis of level or trend-stationarity against the alternative of a unit root process. The null hypothesis of stationarity is particularly useful when the previous unit root tests fail to reject the null of unit root as it happens for many economic time series data (KPSS, 1992).

The process is expressed as the sum of a deterministic trend, random walk, and stationary error:

$$y_t = \xi_t + r_t + \varepsilon_t \tag{1.9}$$

$$r_t = r_{t-1} + u_t \text{, } with\, u_t \approx iid\ N(0, \sigma_u^2) \tag{1.10}$$

The LM test statistic is defined as:

$$LM = \sum_{t=1}^{T} {}^{S_t^2}\Big/_{\hat{\sigma}_\varepsilon^2} \tag{1.11}$$

with

$$S_t = \sum_{i=1}^{t} e_i, \qquad t = 1,2, \dots, T \tag{1.12}$$

Where S_t is the partial sum of the OLS residuals and $\hat{\sigma}_\varepsilon^2$ is the point estimate of the error variance. The non-rejection of the null hypothesis arises whenever the LM statistic is smaller than the critical value.

The standard ADF and PP tests have in some cases low power of rejecting the null, thus the KPSS test is a valuable complement to verify the stationary properties of a series (Fodha and Zaghdoud, 2010).

Dickey-Fuller Generalised Least Squares test

Elliot, Rothenberg and Stock (1996) developed the DFGLS test using a generalized least squares procedure. This test consists of a modification to the ADF approach in which the time series are detrended prior to running the unit root test. This detrending procedure is done by removing the deterministic terms out of the data. The DFGLS test involves the application of the ADF method to the GLS detrended data, according to the equation:

$$\Delta y_t^d = \alpha y_{t-1}^d + \sum_{i=1}^{p} \beta_i \Delta y_{t-i}^d + v_t \qquad (1.13)$$

Where y_t^d is the generalised least squares detrended data and v_t is the independently and identically distributed (iid) error term. Similar to ADF test the null hypothesis is $\alpha = 0$ against the alternative of $\alpha < 0$. The null hypothesis is not rejected when the test statistic is greater than the critical value, tabulated in Elliott, Rothenberg and Stock (1996).

The number of lags is determined using either the Akaike Information Criteria or the Schwarz Bayesian Information Criteria.

Elliott, Rothenberg and Stock test (Point optimal test)

ERS (1996) propose a second unit root test entitled point-optimal test. It is assumed that the data is generated as follows:

$$y_t = d_t + u_t, \qquad t = 1, 2, \ldots, T \qquad (1.14)$$

With

$$u_t = \alpha u_{t-1} + v_t \qquad (1.15)$$

Where d_t refers to the deterministic term and the unobserved stationary zero-mean error process is denoted by v_t.

The null hypothesis of unit root is $\alpha = 1$ against the alternative $\alpha = \bar{\alpha}$, smaller than one in absolute value, and the test statistics is:

$$P_T = \frac{[S(\bar{\alpha}) - \bar{\alpha} S(1)]}{\hat{\omega}^2} \qquad (1.16)$$

Where $S(\alpha)$ is the OLS sum squared residuals from the regression of y_α (T- dimensional column vector y_α) on Z_α ($T\ x\ q$ matrix):

$$y_\alpha = (y_1, y_2 - ay_1, \dots, y_T - ay_{T-1})' \quad (1.17)$$

$$Z_\alpha = (z_1, z_2 - az_1, \dots, z_T - az_{T-1})' \quad (1.18)$$

z_t is the deterministic q-dimensional vector.

And $\widehat{\omega}^2$ is the consistent estimator of $\omega^2 = \sum_{k=-\infty}^{\infty} E(v_t v_{t-k})$.

The ERS test rejects the null hypothesis of unit root for a smaller value of the test statistic than the critical value.

Ng and Perron test

Ng and Perron (2001) use the GLS detrending procedure of ERS (1996) to develop efficient versions of the modified PP tests of Perron and Ng (1996). These efficient modified PP tests do not reveal the severe size distortions of the PP tests for errors with large negative MA or AR roots.

Let the series $\{y_t\}_{t=0}^{T}$ be generated by:

$$y_t = d_t + u_t \quad (1.19)$$

$$u_t = \alpha u_{t-1} + v_t \quad (1.20)$$

Where $E(u_0^2) < \infty$, $v_t = \delta(L)e_t = \sum_{j=0}^{\infty} \delta_j e_{t-j}$ with $\sum_{j=0}^{\infty} j\left|\delta_j\right| < \infty$ and $\{e_t\} \sim iid\ (0, \sigma_e^2)$. The non-normalized spectral density at frequency zero of v_t at frequency zero is given by $\sigma^2 = \sigma_e^2 \delta(1)^2$. Also, $d_t = \psi'\, z_t$ where z_t is a set of deterministic components.

The authors consider $d_t = \sum_{i=0}^{p} \psi_i t^i$, with special focus on $p = 0, 1$. The null hypothesis of unit root – $\alpha = 1$ – is tested against $|\alpha| < 1$. Let the ADF test be the t statistic for $\beta_0 = 0$ in the following auto-regression:

$$\Delta y_t = d_t + \beta_0 y_{t-1} + \sum_{j=1}^{k} \beta_j \Delta y_{t-j} + e_{tk} \quad (1.21)$$

In Perron and Ng (1996), the properties of three tests, MZ_α, MZ_t, and MSB, collectively referred to as the M tests, were derived:

$$MZ_\alpha = (T^{-1} y_T^2 - s_{AR}^2)(2T^{-2} \sum_{t=1}^{T} y_{t-1}^2)^{-1} \quad (1.22)$$

$$MSB = \left(T^{-2} \sum_{t=1}^{T} y_{t-1}^2 \Big/ s_{AR}^2\right)^{1/2} \quad (1.23)$$

$$MZ_t = MZ_\alpha + MSB \tag{1.24}$$

The M tests are based on an autoregressive estimate of the spectral density at frequency zero of v_t at zero frequency, s_{AR}^2, given by:

$$s_{AR}^2 = \hat{\sigma}_k^2 / \left(1 - \hat{\beta}(1)\right)^2 \tag{1.25}$$

Where $\hat{\beta}(1) = \sum_{i=1}^{k} \hat{\beta}_i$ and $\hat{\sigma}_k^2 = (T-k)^{-1} \sum_{t=k+1}^{T} \hat{e}_t^2$, with $\hat{\beta}_i$ and $\{\hat{e}_t\}$ obtained from equation (1.21).

MZ_α and MZ_t, as modified versions of the Phillips (1987) and Phillips-Perron (1988) Z_α and Z_t tests, are referred as to the Z tests. The MSB test statistic is associated to Bhargava's (1986) R-statistic. The Z tests suffer from severe size distortion when v_t has a negative moving average root. The M tests demonstrate having much smaller size distortions than most of the existing unit root tests, including the Z tests, if k is appropriately selected.

Ng and Perron (2001) adapted the local to unity GLS detrending procedure proposed previously in ERS (1996). For any series $\{x_t\}_{t=0}^{T}$, define $\left(x_0^{\bar{\alpha}}, x_t^{\bar{\alpha}}\right) \equiv (x_0, (1-\bar{\alpha}L)x_t)$, $t = 1, \ldots, T$ for some chosen $\bar{\alpha} = 1 - \bar{c}/T$. The GLS detrended series is defined as:

$$\tilde{y}_t \equiv y_t - \hat{\psi}' z_t \tag{1.26}$$

Where $\hat{\psi}$ minimizes $s(\bar{\alpha}, \psi) = (y^{\bar{\alpha}} - \psi' z^{\bar{\alpha}})'(y^{\bar{\alpha}} - \psi' z^{\bar{\alpha}})$. If v_t is iid normal, the point optimal test of the null hypothesis $\alpha = 1$ against the alternative $\alpha = \bar{\alpha}$ is the likelihood ratio statistic, $L = S(\bar{\alpha}) - S(1)$, where $S(\bar{\alpha}) = min_\psi S(\bar{\alpha}, \psi)$. ERS (1996) considered a feasible point optimal test that takes into consideration the possibility that v_t may be serially correlated. The statistic is:

$$P_T = [S(\bar{\alpha}) - \bar{\alpha} S(1)] / s_{AR}^2 \tag{1.27}$$

The value of $\bar{c}$ is chosen such that the asymptotic local power function of the test is tangent to the power envelope at 50% power. For $p = 0$ this is -7.0 and for $p = 1$, it is -13.5. ERS (1996) also suggested the DF^{GLS} as the t-statistic for testing $\beta_0 = 0$ from the following regression estimated by OLS:

$$\Delta\tilde{y}_t = \beta_0 \tilde{y}_{t-1} + \sum_{j=1}^{k} \beta_j \Delta\tilde{y}_{t-1} + e_{tk} \tag{1.28}$$

Based on ERS' (1996) previous work, Ng and Perron (2001) also use the GLS detrending approach to the M tests, which are designated as the M^{GLS} tests. Ng and Perron examined the asymptotic properties of the M^{GLS} tests and calculate the respective critical values. Furthermore, the authors considered two modified feasible point optimal tests which are given as follows:

$$p = 0: \quad MP_T^{GLS} = [c^{-2} T^{-2} \sum_{t=1}^{T} \tilde{y}_{t-1}^2 - \bar{c} T^{-1} \tilde{y}_T^2] / s_{AR}^2$$

$p = 1$: $\quad MP_T^{GLS} = [c^{-2}T^{-2}\sum_{t=1}^{T}\tilde{y}_{t-1}^2 + (1-\bar{c})T^{-1}\tilde{y}_T^2]/s_{AR}^2$ (1.29)

The test statistics are named $\bar{M}^{GLS}$ and $\bar{Z}^{GLS}$ when s_{AR}^2 is derived from equation (1.28). When the estimation of s_{AR}^2 uses the OLS detrended data of equation (1.21), the test statistics are called M^{GLS} and Z^{GLS}. The null hypothesis is not rejected if the test statistics are greater than the critical values. Ng and Perron (2001) suggest a class of modified information criteria (MIC) in order to select k.

Properties of the data

Tables A.1-A.6 of Appendix A display the results for the unit root tests conducted in this study. Our analysis relies on the results for intercept and an intercept and trend only, due to the nature of the series. All the tests give the same results for the real GDP, CO2transp and temperature series. Except for temperature, evidence suggests that is I(0), the hypothesis of a unit root in levels is not rejected for the other variables.

For the remaining variables under study, the results are contradictory and inconclusive. Concerning the CO2total, according to PP, KPSS and Ng-Perron (intercept), the series is integrated of order one. However, the ADF (intercept), Ng-Perron (intercept and trend) and ERS tests give inconclusive results, and ADF (intercept and trend) DFGLS (both intercept and intercept and trend) indicate that the series is stationary in levels, which is rather implausible. Regarding CO2elec, the series is considered integrated of order one, although the Ng-Perron test indicates that the series is integrated of order two (only the MSB statistic, for intercept, suggests that the series is integrated of order one), and KPSS test (intercept and trend) does not reject the null of stationarity. In the case of the rate of motorization, based on the ADF, PP, DFGLS and ERS tests, the second differences is stationary. Based on KPSS results the rate of motorization is integrated of order one, while Ng-Perron suggests the rate of motorization is stationary in levels (intercept) and integrated of order one (intercept and trend). As for crude and fuel series, all tests indicate that these series are stationary at first differences, except KPSS test that evidences both crude (intercept and trend) and fuel (intercept, intercept and trend) are stationary in levels which is rather unlikely.

In respect to precipitation, apart from ERS test (intercept), the series is considered stationary in levels as expected.

In spite of some contradictory results, most tests indicate the following statistical properties of the variables: GDP, CO2total, CO2elect, CO2transp, crude and fuel are integrated of order one, temperature and precipitation are stationary, and the rate of motorization is integrated of order two, for a 5% significance level.

Having determined the order of integration, we proceed with the possibly long-run cointegration relationship between the variables. We assume that the motorization rate is I(2), thus to conduct the cointegration analysis we take the first differences of the motorization rate series. In our regressions, the carbon emissions variables are integrated of order one and, thus, they should not be explained by regressors with different order of integration such as I(2).

1.3. Methodology under nonlinearities

In this section we briefly outline the econometric specifications and cointegration procedures applied to the data for this empirical study.

1.3.1. Econometric specification

Considering the available data and noting the strong link between the crude and fuel prices, the more general expressions for the dependent variables under study could be specified as follows:

$CO2elect = f(y, y^2, y^3, crude, temp, precip)$

$CO2transp = f(y, y^2, y^3, crude, motor)$ or $CO2transp = f(y, y^2, y^3, fuel, motor)$

$CO2total = f(y, y^2, y^3, crude, temp, precip, motor)$

or

$CO2total = f(y, y^2, y^3, fuel, temp, precip, motor)$

where $CO2elect$, $CO2transp$ and $CO2total$, are emissions from the Portuguese electricity generation sector, the Portuguese transportation sector and the total emissions in Portugal, respectively. y represents the per capita real GDP, $crude$ refers to the crude oil price, $fuel$ is the weighted average fuel price, $temp$ is the temperature, $precip$ is the precipitation, and $motor$ denotes the rate of motorization.

A time trend is not included since univariate test statistics indicate that the CO_2 time series enclose a stochastic trend. Hence, the inclusion of a deterministic trend may cause spurious outcomes (Richmond and Kaufmann, 2006).

However, we will begin by analysing only the relationship between per capita GDP and per capita carbon dioxide emissions, which can be specified through several candidate functional forms. All of them will be tested for both sectors under analysis as well as for total CO_2 emissions. First we focus on the quadratic and cubic reduced forms in levels and natural logarithms:

$$CO2_t = \beta_0 + \beta_1 y_t + \beta_2 y_t^2 + \varepsilon_t \tag{1.30}$$

$$CO2_t = \beta_0 + \beta_1 y_t + \beta_2 y_t^2 + \beta_3 y_t^3 + \varepsilon_t \tag{1.31}$$

$$lnCO2_t = \beta_0 + \beta_1 ln(y_t) + \beta_2 ln(y_t)^2 + \varepsilon_t \tag{1.32}$$

$$lnCO2_t = \beta_0 + \beta_1 ln(y_t) + \beta_2 ln(y_t)^2 + \beta_3 ln(y_t)^3 + \varepsilon_t \tag{1.33}$$

Where $CO2_t$ and y_t are the per capita CO_2 emissions and per capita real GDP in year t, respectively, and ε_t is the stochastic error term.

Then the set of other explanatory variables will be added, as follows:

$$CO2_t = \beta_0 + \beta_1 y_t + \beta_2 y_t^2 + \beta_k V_t + \varepsilon_t \tag{1.34}$$

$$CO2_t = \beta_0 + \beta_1 y_t + \beta_2 y_t^2 + \beta_3 y_t^3 + \beta_k V_t + \varepsilon_t \tag{1.35}$$

$$lnCO2_t = \beta_0 + \beta_1 ln(y_t) + \beta_2 ln(y_t)^2 + \beta_k ln(V_t) + \varepsilon_t \tag{1.36}$$

$$lnCO2_t = \beta_0 + \beta_1 ln(y_t) + \beta_2 ln(y_t)^2 + \beta_3 ln(y_t)^3 + \beta_k ln(V_t) + \varepsilon_t \tag{1.37}$$

V_t is a vector of k additional variables – crude oil prices, average fuel prices, change rate of motorization, temperature and precipitation – that together with per capita real GDP and its squares and cubes, may influence the GDP-CO_2 emissions nexus.

The above equations encompass the following possible relationships between economic growth and CO_2 emissions:

(i) $\beta_1 = \beta_2 = \beta_3 = 0$: Flat pattern or no relationship;

(ii) $\beta_1 > 0$ and $\beta_2 = \beta_3 = 0$: Monotonic increasing linear relationship;

(iii) $\beta_1 < 0$ and $\beta_2 = \beta_3 = 0$: Monotonic decreasing linear relationship;

(iv) $\beta_1 > 0$, $\beta_2 < 0$ and $\beta_3 = 0$: Inverted-U-shaped relationship;

(v) $\beta_1 < 0$, $\beta_2 > 0$ and $\beta_3 = 0$: U-shaped relationship;

(vi) $\beta_1 > 0$, $\beta_2 < 0$ and $\beta_3 > 0$: N-shaped relationship;

(vii) $\beta_1 < 0$, $\beta_2 > 0$ and $\beta_3 < 0$: Inverted N- shaped relationship;

EKC may hold only for forms iv and vii, depending on the estimated value of the turning point of the per capita real GDP.

The log-log models allows us to assess if the EKC holds, and also to measure the income elasticity respect to GDP, which is measured by the slope coefficient β_1.

1.3.2. Cointegration analysis

Rank tests for cointegration

Breitung's (2001) rank test for cointegration is a nonparametric methodology which has the advantage of not requiring any particular functional form of the nonlinear cointegrating relationship. Moreover, it can detect either linear or nonlinear cointegration relations. The cointegration test is based on the rank transformation of the time series, where each observation of x_t is substituted by its rank in the observed series. The rank test relies on the fact that if x_t is a random walk, then the ranked series of x_t resembles a random walk as well. Likewise, if two series are (nonlinear) cointegrated, the ranked series are cointegrated as well. The null hypothesis of no (nonlinear) cointegration is not rejected if the test statistic is smaller than the respective critical value, available in table 1 of Breitung (2001).

Consider the existing nonlinear cointegration relationship, given by:

$$u_t = g(y_t) - f(x_t) \qquad (1.38)$$

For $t = 1, \dots, T (T$ is the sample size), where $f(x_t) \sim I(1)$, $g(y_t) \sim I(1)$ and $u_t \sim I(0)$. The functions $f(x)$ and $g(y)$ are monotonically increasing[3]. If it is unknown whether these functions are monotonically increasing or decreasing, a two-sided test is accessible.

The rank series is defined as:

$$R_T(x_t) = Rank\ [of\ x_t\ among\ x_1, \dots, x_T]$$

And $R_T(y_t)$ is constructed accordingly.

The rank statistic is built by replacing $f(x_t)$ and $g(y_t)$ by $R_T(x_t)$ and $R_T(y_t)$, respectively.

The advantage of a statistic based on the sequence of ranks is that the functions $f(.)$ and $g(.)$ need not be known.

[3] The resulting test is valid for more generic cases (Breitung, 2001) and thus it can be applied to our analysis.

Two “distance measures” are considered between the sequences $R_T(x_t)$ and $R_T(y_t)$:

$$\kappa_T = T^{-1}\sup_t |d_t| \tag{1.39}$$

and

$$\xi_T = T^{-3}\sum_{t=1}^{T} d_t^2 \tag{1.40}$$

where $d_t = R_T(y_t) - R_T(x_t)$. κ_T is a Kolmogorov-Smirnov type of statistic and ξ_T is a Cramer-von-Mises type of statistic. Breitung (2001) proposes the corrected versions of the original test statistics κ_T^* and ξ_T^*, if $f(x_t)$ and $g(y_t)$ are mutually correlated as in this case the random walk sequences are more complex.

$$\kappa_T^* = \frac{\kappa_T}{\hat{\sigma}_{\Delta d}} \tag{1.41}$$

$$\xi_T^* = \frac{\xi_T}{\hat{\sigma}_{\Delta d}^2} \tag{1.42}$$

Where: $\hat{\sigma}_{\Delta d}^2 = T^{-2}\sum_{t=2}^{T}(d_t - d_{t-1})^2$

The critical values for the test statistics κ_T, ξ_T, κ_T^* and κ_T^* are computed for T=500 (Breitung, 2001, table 1).

The simulation results in Breitung’s (2001) study demonstrate that the κ_T^* and ξ_T^* tests can be applied when the correlation between $f(x_t)$ and $g(y_t)$ is small to moderate. Additionally, for correlations close to one, Breitung (2001) defines:

$$\kappa_T^{**} = \frac{\kappa_T^*}{\lambda_\kappa^\alpha(E\rho_T^R)} \tag{1.43}$$

and

$$\xi_T^{**} = \frac{\xi_T^*}{\lambda_\xi^\alpha(E\rho_T^R)} \tag{1.44}$$

Where $\lambda(.)$ is a correction term that depends on α (test significance level) and ρ_T^R , the expected correlation coefficient of the rank differences.

For a 5% level, Breitung shows that $\lambda_\kappa^\alpha(E\rho_T^R)$ is approximately $1 - 0.174(\rho_T^R)^2$ and $\lambda_\xi^\alpha(E\rho_T^R)$ is approximately $1 - 0.462\rho_T^R$, where ρ_T^R is the correlation coefficient of the rank differences:

$$\rho_T^R = \frac{\sum_{t=2}^{T} \Delta R_T(x_t) \Delta R_T(y_t)}{\sqrt{\left(\sum_{t=2}^{T} \Delta R_T(x_t)^2\right)\left(\sum_{t=2}^{T} \Delta R_T(y_t)^2\right)}} \tag{1.45}$$

The statistics κ_T^{**} and ξ_T^{**} have the same limiting distribution than κ_T^{*} and ξ_T^{*}.

The rank test can be generalized to cointegration among $k+1$ variables,y_t, x_{1t},... x_{kt}, (see Breitung, 2001, for details[4]).

Neglected nonlinearity

In the same paper, Breitung develops a rank test for neglected nonlinearity, useful when it is assumed that there is a stable long-run relationship between the time series of a linear or nonlinear form. The test statistic follows a standard asymptotic χ^2 distribution under the null hypothesis of linear cointegration.

Consider the nonlinear relationship:

$$y_t = \gamma_0 + \gamma_1 x_t + f^*(x_t) + u_t \tag{1.46}$$

where $\gamma_0 + \gamma_1 x_t$ is the linear part of the cointegration model. Under the null hypothesis, $f^*(x_t) = 0$ for all t and u_t is $I(0)$, so there is linear cointegration. Under the alternative hypothesis, for $f^*(x_t) \neq 0$ and u_t $I(0)$, there is nonlinear cointegration.

A score-type test statistic is given by the $T.R^2$statistic from the least squares regression:

$$\tilde{u}_t = c_0 + c_1 x_t + c_2 R_T(x_t) + e_t \tag{1.47}$$

where $\tilde{u}_t = y_t - \tilde{\gamma}_0 - \tilde{\gamma}_1 x_t$ and $\tilde{\gamma}_0$ and $\tilde{\gamma}_1$ are the least squares point estimates.

The test statistic is distributed as χ^2 with one degree of freedom. The extension of this test to more than two nonstationary variables is straightforward.

Usually, the errors u_t are found to be serially correlated and the regressor x_t may be endogenous. In these cases, under the null hypothesis of linear cointegration, the following representation holds[5]:

$$y_t = \gamma_0^* + \sum_{j=1}^{\infty} \alpha_j y_{t-j} + \gamma_1^* x_t + \sum_{j=-\infty}^{\infty} \pi_j^* \Delta x_{t-j} + \varepsilon_t \tag{1.48}$$

[4] For the EKC hypothesis k=1.
[5] see Stock and Watson 1993; Inder 1995

Hence, a test for nonlinear cointegration can be obtained by truncating the infinite sums appropriately and forming $T.R^2$ for the regression of the residuals $\tilde{\varepsilon}_t$ on the regressors of (1.48) and $R_t(x_t)$. Once again, the resulting score statistic is asymptotically χ^2 distributed under the null hypothesis.

Linearity tests

Choi and Saikkonen (2004, hereafter CS (2004)) propose statistical tests for detecting linearity in cointegrated smooth transition regression (STR) models, considering $I(1)$ variables. The tests allow for correlation between the regressors and errors of the model, by correcting for endogeneity on the basis of the lead-and-lags (LL) approach (DOLS, dynamic OLS estimation).

Consider the cointegrating STR model:

$$y_t = \mu + vg(z_{st}) + \alpha' x_t + \beta' x_t g(z_{st}) + u_t, \qquad t = 1, 2, \dots, T \tag{1.49}$$

Where $x_t = [x_{1t}, \dots, x_{pt}]'$ is a p-dimensional $I(1)$ process, u_t is a zero-mean stationary error term, and

$$z_{st} = \gamma(x_{st} - c), \quad \gamma \neq 0, \;\; s \in \{1, \dots, p\} \tag{1.50}$$

Furthermore, $g(z_{st})$ is a smooth, real-valued transition function of the process x_{st} and the scalars γ and c. Here, μ and v are scalars and α and β are $p \times 1$ vectors. In model (1.49), there is a single $g(z_{st})$ that is affected by the single transition variable x_{st}. It is straightforward to extent the linearity tests for this model to multiple transition functions and transition variables.

The nonlinear nature of the model is determined by the transition function $g(z_{st})$.

Based on model (1.49) the null hypothesis of interest is:

$$H_0 : v = 0 \; and \; \beta = 0 \tag{1.51}$$

In model (1.49), the transition function $g(z_{st})$ allows the relationship between x_t and y_t to change depending on where x_{st} are located relative to the parameter c. The smoothness of the change is characterized by the parameter γ.

The test procedure for the linearity restrictions (1.51) against the cointegrating STR model (1.49) is developed allowing for the cases of both no drift and drift in regressors. This testing problem is non-standard as the nuisance parameters γ and c are not identified under the null hypothesis. This follows from equation (1.49) because, under the null hypothesis, the transition functions $g(z_{st})$ and, consequently, the parameters γ and c can take any values without any effect on the model specification.

Instead of explicitly using the given transition function $g(z_{st})$, the authors CS (2004) use its Taylor's series approximations: the first-order and third-order tests. Both tests are Lagrange Multiplier (LM)-type tests in that they require estimating the model only under the null hypothesis of linearity. The simplicity of these tests stems from the fact that OLS can be used in this context.

- **First-order test:**

The first-order Taylor series approximation of the function $g(z_{st})$ around the origin is given by:

$$g(z_{st}) \approx b\beta(x_{st} - c) \tag{1.52}$$

Substituting this approximation for $g(z_{st})$ in model (1.49), the auxiliary regression model becomes:

$$y_t = \phi + \rho' x_t + \sum_{k=1}^{p} \theta_k x_{kt} x_{st} + \eta_t \tag{1.53}$$

where the parameters ϕ (scalar), $\rho = [\rho_1, \dots, \rho_p]'$ and θ_k (scalar) are implicitly defined as a function of the original parameters. The error term η_t is the sum of u_t in model (1.49) and the approximation error. Under the null hypothesis, the approximation error vanishes and, consequently, $\eta_t = u_t$. The idea is to test the original null hypothesis (1.51) by:

$$H'_0: \theta_k = 0 \ (k = 1, \dots, p) \tag{1.54}$$

in the auxiliary regression model (1.53).

However, as the tests allow for correlation between the regressors and the error term, an endogeneity correction is employed on the leads-and-lags methodology[6].

$$y_t = \phi + \rho' x_t + \zeta' n_t + \sum_{j=-K}^{K} \pi'_j \Delta x_{t-j} + \eta_{Kt}, \quad t = K+1, \dots, T-K \tag{1.55}$$

For fixed K, where Δ is the difference operator, $\zeta = [\theta_1, \dots, \theta_p]'$, $n_t = [x_{1t}x_{st}, \dots, x_{pt}x_{st}]'$, and η_{Kt} is the error term.

Let M be the moment matrix for the auxiliary regression model (1.55) and $(M^{-1})_{nn}$ the block of the matrix M^{-1} corresponding to n_t. Then the LM test for the null hypothesis (1.54) is defined as:

$$\mathcal{T}_1 = \hat{\zeta}'[\tilde{\omega}_e^2 \, (M^{-1})_{nn}]^{-1} \, \hat{\zeta} \tag{1.56}$$

[6] See Saikkonen and Choi (2004).

where $\hat{\zeta}$ is the OLS estimator of ζ in (1.55) and $\widetilde{\omega}_e^2$ is the standard long-run variance estimator based on the residuals of the corresponding restricted estimation. This is the test statistic once there are no drifts in the regressors. The standard χ^2 distribution can be used to test for the linearity hypothesis.

Presume that the transition variables contain the unknown drift parameter μ_x. Repeating the same procedure to obtain (1.55) but now substituting x_{st} by $x_{st} - t\hat{\mu}_{sx}$, where $\hat{\mu}_x$ is the corresponding estimator and $\hat{\mu}_{ix}$ is its ith component, yields:

$$y_t = \phi + \rho' x_t + \zeta' \hat{n}_t + \sum_{j=-K}^{K} \pi'_j \Delta x_{t-j} + \eta_{Kt}, \quad t = K+1, \ldots, T-K \tag{1.57}$$

Where $\hat{n}_t = \left[(x_{1t} - t\hat{\mu}_{1x})(x_{st} - t\hat{\mu}_{sx}), \ldots, \left(x_{pt} - t\hat{\mu}_{px}\right)(x_{st} - t\hat{\mu}_{sx})\right]'$

In (1.57), x_t contains both deterministic and stochastic trends that need to be separated for our test. This can be done by using the model augmented by a time trend:

$$y_t = \phi + \tau t + \rho' x_t + \zeta' \hat{n}_t + \sum_{j=-K}^{K} \pi'_j \Delta x_{t-j} + \eta_{Kt}, \quad t = K+1, \ldots, T-K \tag{1.58}$$

Instead of $\mathcal{T}_1$ we now have the test statistic:

$$\mathcal{T}_{1\mu} = \hat{\zeta}'_\mu \left[\widetilde{\omega}_{e\mu}^2 \, (M^{-1})_{\hat{n}\hat{n}}\right]^{-1} \hat{\zeta}_\mu \tag{1.59}$$

Where $\hat{\zeta}_\mu$ is the OLS estimator of ζ in (1.58), $(M^{-1})_{\hat{n}\hat{n}}$ is defined in the same way as $(M^{-1})_{nn}$ and $\widetilde{\omega}_{e\mu}^2$ is a long-run variance estimator obtained from the OLS residuals of (1.58). The limiting null distribution of the test statistic $\mathcal{T}_{1\mu}$is the same as that of $\mathcal{T}_1$.

- **Third-order test:**

In equations (1.53) and (1.55), the parameters θ_k are not functions of v, which only affects the parameters ϕ and ρ. Because the parameter v cannot be detached from ϕ and ρ, the test statistic $\mathcal{T}_1$ may have low power when the value of v is large and the elements of β take small absolute values. To deal with this difficulty, CS (2004) consider a test that relies on the third-order Taylor series approximation of $g(z_{st})$.

The third-order Taylor series approximation of the function $g(z_{st})$ around the origin can be written as:

$$g(z_{st}) \approx b\gamma(x_{st} - c) + d\gamma^2(x_{st} - c)^2 + h\gamma^3(x_{st} - c)^3 \tag{1.60}$$

where b, d and h are constants determined by partial derivatives of $g(z_{st})$ at the origin. Because the motivation for using the third-order approximation is to improve the power of test statistic $\mathcal{T}_1$ under $v \neq 0$, we use the approximate relation (1.60) at the transition of the intercept term, $vg(z_{st})$, and keep the relation (1.52) for the transition involving the regressors, $\beta' x_t g(z_{st})$. The resulting auxiliary regression model (cf (1.55)) is:

$$y_t = \psi + \xi' x_t + \varsigma' h_t + \sum_{j=-K}^{K} \pi'_j \Delta x_{t-j} + \eta_{Kt}^*, \quad t = K+1, \ldots, T-K \tag{1.61}$$

Here $\varsigma = [\varphi_1, \ldots, \varphi_p, \lambda]'$ and $h_t = [x_{1t}x_{st}, \ldots, x_{pt}x_{st}, x_{st}^3]'$. Moreover, the $p \times 1$ parameter vector ξ and the scalar parameters φ_k, λ and ψ are defined appropriately.

Instead of the original null hypothesis (1.51), the idea is now to test the null hypothesis:

$$H_0''': \varsigma = 0 \tag{1.62}$$

in the auxiliary regression model (1.61) using the standard LM test. The test statistic is expressed as:

$$\mathcal{T}_2 = \hat{\varsigma}' \left[\widetilde{\omega}_e^2 \left(\overline{M}^{-1} \right)_{hh} \right]^{-1} \hat{\varsigma} \tag{1.63}$$

which follows a χ^2 distribution and whose terms are defined as previously. The number of degrees of freedom has increased by the number of additional restrictions included in the auxiliary null hypothesis.

The above test can be modified to allow for drifts in the regressors. Instead of (1.61), the test is based on:

$$y_t = \psi + \tau t + \xi' x_t + \varsigma' \hat{h}_t + \sum_{j=-K}^{K} \pi'_j \Delta x_{t-j} + \eta_{Kt}^*, \quad t = K+1, \ldots, T-K \tag{1.64}$$

Where $\hat{h}_t$ is defined by replacing x_{st} in the definition of the h_t by $x_{st} - t\hat{\mu}_{sx}$. This leads to test statistic:

$$\mathcal{T}_{2\mu} = \hat{\varsigma}_\mu' \left[\widetilde{\omega}_{e\mu}^2 \left(\overline{M}^{-1} \right)_{hh} \right]^{-1} \hat{\varsigma}_\mu \tag{1.65}$$

Where $\hat{\varsigma}_\mu$ is the OLS estimator of ς in (1.63), $\left(\overline{M}^{-1} \right)_{hh}$ is an analogue of $\left(\overline{M}^{-1} \right)_{hh}$ based on $\hat{h}_t$ instead of h_t, and $\widetilde{\omega}_{e\mu}^2$ is a long-run variance estimator obtained from the OLS residuals of (1.64) with the constraint $\zeta = 0$. The limiting null distribution of test statistic $\mathcal{T}_{2\mu}$ is the same as that of $\mathcal{T}_2$.

When it is known that the coefficients associated to the regressors are not subject to smooth transition (i.e. $\beta = 0$), the test is built on the regression equation:

$$y_t = \psi + \xi' x_t + \varsigma' h_t + \sum_{j=-K}^{K} \pi'_j \Delta x_{t-j} + \eta_{Kt}^*, \quad t = K+1, \ldots, T-K \tag{1.66}$$

where $\varsigma = [\varphi, \lambda]'$ and $h_t = [x_{st}^2, x_{st}^3]'$. The test for the null hypothesis $\varsigma = 0$ is $\mathcal{T}_3$ and it follows the standard χ^2 distribution with two degrees of freedom.

This test may have higher power than $\mathcal{T}_2$ when $\beta = 0$. In some transition models, $d = 0$ (cf (1.60)) and the auxiliary regression model reduces to:

$$y_t = \psi + \xi' x_t + \lambda x_{st}^3 + \sum_{j=-K}^{K} \pi'_j \Delta x_{t-j} + \eta_{Kt}^*, \quad t = K+1, \dots, T-K \tag{1.67}$$

In this case, $\mathcal{T}_3$ tests the null hypothesis $\lambda = 0$, and the test statistic follows the standard χ^2 distribution with one degree of freedom.

When the regressors contain drifts, the resulting test is designated by $\mathcal{T}_{3\mu}$.

Tests for nonlinear cointegration

Choi and Saikkonen's (2010, hereafter CS (2010)) methodology to detect nonlinear cointegrating relationships is an alternative to Breitung's rank tests. Contrary to Breitung, CS (2010) requires the precise specification of the nonlinear model before running the test, and uses the same test statistics as the KPSS test.

Let the nonlinear cointegrating regression be:

$$y_t = g(x_t, \theta) + u_t, \quad t = 1,2, \dots, T \tag{1.68}$$

where $x_t (p \times 1)$ is an $I(1)$ regressor vector, u_t is a zero-mean stationary error term, $g(x_t, \theta)$ a known, smooth function of the process x_t and the parameter vector $\theta(k \times 1)$.

In order to test for cointegration, one has to infer the properties of the error process u_t. CS (2010) test the null hypothesis of stationarity of u_t against the alternative that it is an $I(1)$ process. CS (2010) first study the cointegration tests using full-sample residuals. For a NLLS estimation of equation (1.68) the residuals are given by:

$$\tilde{u}_t = y_{tT} - g\left(x_{tT}, \tilde{\theta}_T\right) \tag{1.69}$$

which are used to define the test statistic:

$$C_{NLLS} = T^{-2} \tilde{\omega}_u^{-2} \sum_{t=1}^{T} \left(\sum_{j=1}^{t} \tilde{u}_j\right)^2 \tag{1.70}$$

where $\tilde{\omega}_u^2$ is a consistent estimator of the long-run variance ω_u^2 based on the full-sample residuals.

Correspondingly, considering the case of leads-and-lags (LL) estimation (DOLS) for equation (1.68):

$$y_{tT} = g(x_{tT}, \theta) + V_t' \pi + e_{Kt,} \quad t = K+2, \dots, T-K \tag{1.71}$$

where $x_{tT} = \left({T_0}/{T}\right)^{1/2} x_t$, $V_t = [\Delta x'_{t-K} \dots \Delta x'_{t+K}]'$, $\pi = [\pi'_{-K} \dots \pi'_{K}]'$, and $e_{Kt} = e_t + \sum_{|j|>K} \pi'_j v_{t-j}$, it is possible to compute the test on the residuals from (1.71):

$$\hat{e}_{Kt} = y_{tT} - g\left(x_{tT}, \hat{\theta}_T^{(1)}\right) - V'_t \hat{\pi}_T^{(1)}, \quad t = K+2, \dots, T-K \tag{1.72}$$

Again $\{\hat{e}_{Kt}\}_{t=K+2}^{T-K}$ are the full-sample residuals and the test statistic is:

$$C_{LL} = N^{-2} \widetilde{\omega}_e^{-2} \sum_{t=K+2}^{T-K} \left(\sum_{j=K+2}^{t} \hat{e}_{Kj}\right)^2 \tag{1.73}$$

where $\widetilde{\omega}_e^{-2}$ is a consistent estimator of the long-run variance ω_e^2 based on $\{\hat{e}_{Kt}\}_{t=K+2}^{T-K}$.

Nonetheless, the limiting null distributions of the test statistics C_{NLLS} and C_{LL} are impractical, i.e., depend on unknown parameters, for nonlinear models. Hence, CS (2010) introduce a new test for nonlinear cointegration using sub-sample residuals and the Bonferroni procedure. For this purpose consider the following test statistics:

$$C_{NLLS}^{b,i} = b^{-2} \widetilde{\omega}_{i,u}^{-2} \sum_{t=i}^{i+b-1} \left(\sum_{j=i}^{t} \tilde{u}_j\right)^2 \tag{1.74}$$

and

$$C_{LL}^{b,i} = (b - 2K - 1)^{-2} \widetilde{\omega}_{i,e}^{-2} \sum_{t=i+K+2}^{i+b-K} \left(\sum_{j=i+K+2}^{t} \hat{e}_{Kj}\right)^2 \tag{1.75}$$

which have the same functional forms as (1.70) and (1.73), correspondingly, however the subresiduals $\{\tilde{u}_t\}_{t=i}^{i+b-1}$ and $\{\hat{e}_{Kt}\}_{t=i}^{i+b-1}$ are used instead of the full residuals. The index i represents the starting point of the subresiduals and the size of subresiduals – block size – is denoted by b.

The test statistics $C_{NLLS}^{b,i}$ and $C_{LL}^{b,i}$ are expected to have low power compared to those applying full residuals. To overcome this issue, CS (2010) use the Bonferroni procedure, by selecting M tests $C_{NLLS}^{b,i_1}, \dots, C_{NLLS}^{b,i_M}$ and define:

$$C_{NLLS}^{b,max} = max\left(C_{NLLS}^{b,i_1}, \dots, C_{NLLS}^{b,i_M}\right) \tag{1.76}$$

and

$$C_{LL}^{b,max} = max\left(C_{LL}^{b,i_1}, \dots, C_{LL}^{b,i_M}\right) \tag{1.77}$$

Despite having the same block size, the M tests use distinct starting points $i_1, \dots, i_M$. This testing procedure entails that the α-level critical values for test statistics $C_{NLLS}^{b,max}$ and $C_{LL}^{b,max}$ are taken from the distribution of $\int_0^1 W^2(s)ds$ using the level $\frac{\alpha}{M}$, where $W(s)$ is a standard Brownian motion.

CS (2010) suggest that for a specific block size b, the choice of M, the number of subresidual-based tests used in the Bonferroni procedure, and $i_1, \dots, i_M$, the starting points of the subresiduals required for test statistic $C_{NLLS}^{b,max}$, should be done as follows:

Step 1: Let $M = \left[{}^{T}/_{b}\right]^*$, where $[x]^*$ denotes the smallest integer greater than or equal to x.

Step 2: Let $i_1 = 1$, $i_2 = T - b + 1$, $i_3 = b + 1$, $i_4 = T - 2b + 1, \dots$

These steps assure that the whole sample is used to calculate $C_{NLLS}^{b,max}$, while trying to minimize M, which is important because if M is too large, the test procedure will have quite low power. For the $C_{LL}^{b,max}$, substitute T with the actual sample size $T - 2K - 1$ in the above steps.

The authors use the minimum volatility rule to decide on the block size b (see CS, 2000, for details).

Modified Regression Error Specification Test

Hong and Phillips (2010; hereafter HP (2010)) propose a modified Regression Error Specification Test (RESET) in order to remove the size distortions of the conventional RESET test when applied to nonstationary time series. The modified RESET test has the advantage of testing the null hypothesis of linear cointegration against both alternatives of no cointegration and a specific nonlinear cointegration relationship.

Let the regression residuals $\hat{u}_t$ in:

$Y_t = \theta X_t + u_t$ be modelled as $\hat{u}_t = \sum_{j=1}^{k} \beta_j F_j(X_t) + e_t$ (1.78)

where $F_j(X_t) = {X_t}^{j+1}$ form the polynomial basis functions. The nonstationarity of X_t includes bias terms in the limit distribution of the sample covariance between X_t^m – polynomials of X_t – and $\hat{u}_t$ – regression residuals – leading to the noncentral chi-squared limit distribution of the conventional RESET statistic.

To correct the biases, when $\{X_t, Y_t\}$ are linearly cointegrated, the following modified RESET statistic (MR_n) has a limiting central $\chi^2(k)$ distribution:

$$MR_n = \{\hat{u}FD_n - E'_n - S'_n\}\left(\widehat{\Omega}_{uu.v} D'_n \widetilde{F}' \tilde{F} D_n\right)^{-1} \times \{D_n F' \hat{u} - E_n - S_n\} \quad \sim \chi^2(k) \tag{1.79}$$

Where $\hat{u}$ is an $n \times 1$ vector of residuals from the linear cointegration regression (1.78), $F = [F_1, \dots F_n]'$, and $\tilde{F}_t = F_t - X_t(\sum_t X_t X'_t)^{-1} \sum_t X_t F_t$. The matrix $\widehat{\Omega}_{uu.v}$, the $k \times k$ normalization matrix D_n and the $(m-1)th$ elements of the two $k \times 1$ correction vectors $E_n = [E_n(1), \dots, E_n(k)]'$ and $S_n = [S_n(1), \dots, S_n(k)]'$ are defined in Hong and Phillips (2010).

The null hypothesis sets all coefficients equal to zero, $\beta_j = 0$ for $j = 1, \dots, k$. The rejection of the null hypothesis occurs when at least one coefficient deviates enough from zero, in other words, if at the minimum one polynomial basis function – $F_j(X_t) = X_t^{j+1}$ – is capable to capture some "part" of the nonlinearity.

The modified RESET test can be interpreted as a Lagrange multiplier test, where the basis functions are potential alternative nonlinear specifications. By construction, the modified RESET test has highest power against such alternatives. Moreover, if the test rejects the null of linearity, the estimated nonlinear cointegration relationship provides a plausible alternative nonlinear model. In situation when the linear model is rejected and the alternative polynomial model is estimated, the procedure to correct the biases in the least squares coefficient estimators is also employed. Consider the following nonlinear cointegration model:

$$Y_t = \theta X_t^m + u_t \qquad t = 1, \dots, n \tag{1.80}$$

Then the corresponding fully modified estimator of θ is:

$$\tilde{\theta}_m = \left(\sum X_t^{2m}\right)^{-1}\left\{\sum X_t^m Y_t - n^{(m+1)/2}[E_m + S_m]\right\} \tag{1.81}$$

with the respective correction terms:

$$E_m \equiv \hat{\Lambda}_{vu} \frac{m}{n} \sum_{t=1}^{n} \left(\frac{X_t}{\sqrt{n}}\right)^{m-1} \tag{1.82}$$

and

$$S_m \equiv \hat{\Omega}_{uv} \hat{\Omega}_{vv}^{-1} \left\{\sum_{t=1}^{n} \left(\frac{X_t}{\sqrt{n}}\right)^m \frac{v_t}{\sqrt{n}} - \hat{\Delta}_{vv} \frac{m}{n} \sum_{t=1}^{n} \left(\frac{X_t}{\sqrt{n}}\right)^{m-1}\right\} \tag{1.83}$$

(see HP, 2010, for details) The authors also show that the modified RESET test also has power against lack of cointegration as well.

1.4. Empirical results

1.4.1. Cointegration tests

The rank tests for cointegration (Breitung, 2001) are summarized in the table B.1 of appendix B. For total CO_2 emissions, CO_2 from electricity generation, and CO_2 from transport sector, and based on the results of the $k^*_{T=500}$ statistic, the null hypothesis cannot be rejected thus suggesting that none of these variables are cointegrated with GDP. The $\xi^*_{T=500}$ statistic at a 10% significance level reveals the existence of (nonlinear) cointegration between CO_2 from electricity generation and CO_2 from transport sector, and GDP.

The existence of a long-run relationship between the CO_2 variables and GDP is analysed performing the alternative tests of CS (2010) for nonlinear cointegration. The results depicted in tables B.2-B.7 of appendix B suggest a cointegrating relationship between CO_2 emissions – total, electricity and transport – and GDP, at a 5% significance level, for both quadratic and cubic specifications (in levels). These results are valid for both fixed block size, with automatic or fixed bandwidth, and using the minimum volatility rule, and by either NLLS or DOLS K=1, 2 and 3 estimation methods. Similar to CS (2010), we use the $b_{small} = [T^{0.7}]$ and $b_{big} = [T^{0.9}]$.

Although the rank tests fail to clearly confirm the cointegration hypothesis, the most recent tests for nonlinear cointegration and the well-understood economic nexus between CO_2 emissions and GDP provided by the extensive literature underpin the likely existence of cointegration among variables. Therefore we assume cointegration, and proceed with the tests that identify whether the variables are linearly or nonlinearly cointegrated.

In accordance with the results of the rank test for neglected nonlinearity and the test of linearity of the cointegrating relation, reported in tables B.8 and B.9 of appendix B, nonlinear cointegration is predominantly accepted.

The score statistic of the rank test for neglected nonlinearity indicates that cointegration is of a nonlinear form for the relationship between GDP and each of the three CO_2 emissions variables. In this case, the null hypothesis of linear cointegration is rejected for a 1% level of significance.

Moreover, the test of linearity proposed by CS (2004) is also conducted. This test is applied to both quadratic and cubic polynomial specifications in levels for each CO_2 emissions variables, under the alternative hypothesis. As for the CO2total-GDP quadratic model, the linear cointegration hypothesis is rejected at a 10% significance level, for 1 and 2 leads and lags. For K=3, the null hypothesis is accepted. Overall, the outcomes suggest a quadratic long-run relationship between the two variables. When applied to the cubic model, for all K, the $\mathcal{T}_{2\mu}$ test statistic is larger than the critical value at a 1% significance level and, therefore, the hypothesis of a cubic long-run relationship between CO2total and GDP is clearly accepted.

Concerning the link between CO2elect and GDP, the $\mathcal{T}_{1\mu}$ statistic is smaller than the critical value, for all considered leads and lags, suggesting a linear long-run relationship between the variables instead of a quadratic one. Opposite results are obtained when the alternative hypothesis is a cubic cointegration form. In this case, the $\mathcal{T}_{2\mu}$ statistic is larger than the critical values, for K=1, 2 and 3, at a significance level of 1%.

The results for CO2transp and GDP relationship in the long-run point out to a nonlinear cointegration for either alternative hypothesis – quadratic or cubic specifications. For a 5% significance level, $\mathcal{T}_{1\mu}$ and $\mathcal{T}_{2\mu}$ statistics take values above the critical ones.

Because the rank tests for cointegration and the tests for nonlinear cointegration produce ambiguous results, the Hong and Phillips modified RESET test (HP, 2010) is also employed. The interpretation of the results is complementary to the previous ones. The results shown in table B.10 of appendix B are slightly different depending on the alternative hypothesis. Under the alternative of no cointegration or quadratic cointegration, the null hypothesis is rejected with a p-value smaller than 5%, except for the CO2elect case for which the null hypothesis is rejected at a 10% statistical significance. When no cointegration or cubic cointegration is the alternative, the results for all the cases are unanimous in rejecting the null hypothesis with a p-value of zero. Two conclusions can be drawn from these results. First, given the observed p-values, the alternative hypothesis of no cointegration or cubic cointegration is "stronger" than the no cointegration or quadratic cointegration one. Second, the cubic cointegration form prevails over the no cointegration scenario, relying on the previous results of the rank tests for cointegration and the tests for nonlinear cointegration.

Upon the aforementioned tests' results, we find evidence of a nonlinear cointegration relationship between CO2total, CO2elect and CO2transp, and GDP, most probably through a cubic form. As our series are I(1), the validation of cointegration avoids spurious results.

1.4.2. GDP-CO_2 emissions' long-run relationship pattern

Four reduced-form regressions are estimated for the emissions variables, allowing for quadratic and cubic polynomial specifications in levels and logs. The purpose is to ascertain which specification conforms better to the dataset. However, it is worth noting that due to the nonlinearity of the models with nonstationary variables it isn't possible to apply the traditional econometric techniques in order to evaluate if the quadratic or the cubic terms are statistically significant, or which specification produces better diagnostic results. We begin our analysis by comparing the results for quadratic and cubic specifications in levels, and then we address the quadratic and cubic log-log functions. Since the conclusion of previous section indicates that the cubic specification in levels is the one that better describes the relationship between the CO_2 emissions and GDP, the results of the following subsections are analysed in light of that.

The estimated coefficients using the NLLS and DOLS estimation approaches stay roughly the same indicating that these coefficients are "stable" irrespectively of the chosen estimator. Similar to the results of the empirical example of CS (2010), both the NLLS and leads-and-lags regressions provide qualitatively identical outcomes. Thus, we do not indicate all turning points values as they are very similar.

CO_2 emissions from electricity generation

Table C.1, appendix C, reports the estimation results for CO_2 emissions from the electricity generation sector. The coefficients of the linear and quadratic terms of model (1.1), using only GDP as an explanatory variable, are positive for all estimation procedures, indicating that the emissions of CO2elect rise monotonically with GDP. This monotonic pattern is unexpected if we consider the evolution of CO2elect emissions plotted in figure 1.1.

The evidence based on all the estimations undertaken for the cubic model in levels – model (1.2) – establishes an inverted N-shape for the CO2elect-GDP relationship. The minimum and maximum turning points are about €4,227.9 per capita and €14,775.8 per capita, respectively for the NLLS estimation. The EKC is confirmed for CO2elect once the minimum turning point is reached. The turning point values are within the observed sample range. This means that the emissions from the electricity sector have already surpassed the turning point and show a downward-sloping trend. Thus, at present, the CO2elect emissions are delinked from economic growth. The validity of the EKC for this sector is in line with the main finding of Moutinho *et al.* (2014).

Estimated models (1.1) and (1.2) present contrasting results. The plot of the evolution of the emissions of CO2elect between 1960-2010 (Figure 1.1) helps us to select the cubic level specification as the best fitting model.

As for the logarithmic forms – models (1.3) and (1.4) – the results are quite distinct. The positive and negative signs of the coefficients of the linear and the quadratic terms of model (1.3), respectively, validate the EKC.

Apparently, the inverted U-shape is consistent with the cubic model in levels, because it supports the existence of an EKC relationship between emissions from electricity generation and GDP. However, this is a merely statistical result as the turning point – €325. 01 per capita for the NLLS regression – is well below the smallest observed value of the dataset for all regressions, which makes it not plausible at all. Thus, model (1.3) is not appropriate to illustrate the CO2elect—GDP nexus. The cubic regressions in logs give identical results as the cubic level ones, but with slightly higher turning points. Apart from the NLLS estimation, the minimum and maximum turning points are within the sample range. Taking DOLS K=1 as an example, the minimum turning point is €3,176 per capita and the maximum turning point is €15,166.6 per capita, corroborating the emissions decoupling from economic growth.

Looking at the estimated income elasticities, mixed results are obtained. For the log-log quadratic model, for all regressions, the income elasticity is higher than one. For NLLS estimation, a 1% increase in GDP will increase CO2elect emissions by 10%. Visibly, this income elasticity cannot fit the reality. Once again, model (1.3) proves to be unsuitable. On the contrary, the income elasticity for the cubic log-log functional form is less than zero for all regressions. Based on the DOLS K=1 regression, -399.6 is the income elasticity of CO2elect. The negative sign of the elasticity makes sense since the general results indicate a decrease in emissions with economic growth, though the absolute value of the elasticity seems unrealistic.

Given the previously mentioned results, it appears that the cubic reduced-form in levels is the most adequate model for the CO2elect-GDP relationship. Figure 1.2 displays the regressions for model (1.2).

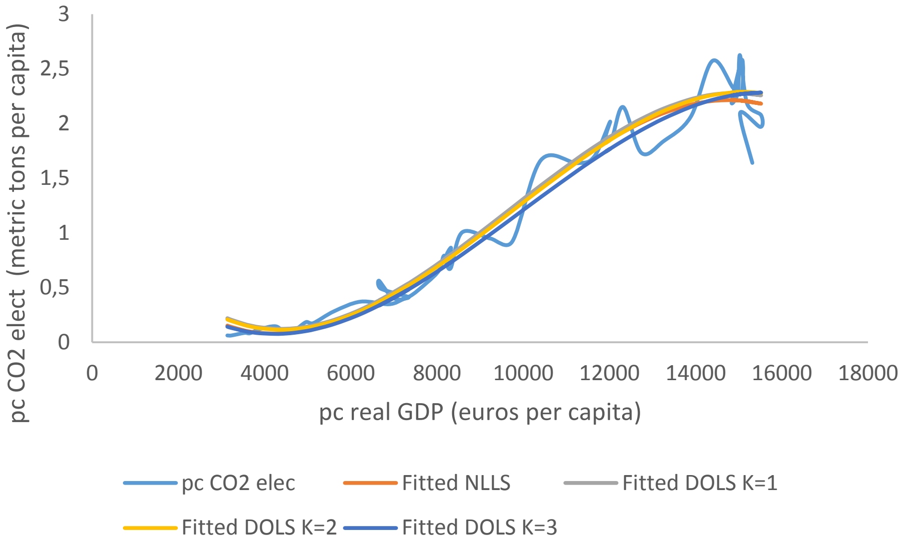

Figure 1.2: Per capita real GDP versus per capita CO_2 emissions from electricity generation sector for Portugal (1960 to 2010).

CO_2 emissions from transport

The estimation results are summarized in table C.2, appendix C. Again, we first estimate the models in a non-logarithmic form. All the estimated coefficients of model (1.5) have positive signs. These results, for NLLS and DOLS estimations, entail a positive monotonic association between the CO2transp and GDP. The estimated models (1.6) yield an N-shaped curve, implying the existence of two turning points. Yet, in practice, the N-shaped curve represents a monotonic increasing relationship between the variables because the roots of the equation are imaginary, which implies that the ascendant part of the N curve is the only one relevant. Therefore, the cubic reduced-form reinforces that CO2transp emissions are growing with GDP, not supporting the EKC hypothesis (Sousa *et al.*, 2015).

The signs of the logarithmic models lead to distinct shapes, but identical conclusions. Model (1.7) – quadratic log – fails to corroborate the EKC hypothesis for transport sector because of the negative sign of the linear term and the positive sign of the quadratic term. These results support a U-shape relation but only in theory because of the very small value of per capita income estimated turning point, for all four regressions. For instance, for the NLLS estimation, the turning point occurs at €13.8[7] per capita, a much smaller value than the observed minimum over time.

Considering model (1.8), an inverted N-shape curve is found for all the regressions, but as in model (1.6) this is only a mathematical result. For both NLLS and DOLS K=1 procedures, it is possible to calculate the per capita GDP values for the two local extremes of the curve. The NLLS turning points come about at €405.2 and over €6 billion, and for DOLS K=1 the minimum and maximum local extreme are about €961.8 and €845.124.7, respectively. These estimated turning points are far from being economically relevant, thus supporting the monotonic relation. As for DOLS K=2 and K=3, the turning points involve imaginary numbers.

For the quadratic-log form, the estimated income elasticities of emissions are negative for all regressions, which is incompatible with the monotonic pattern of CO2transp-GDP. The income elasticities for the cubic-log model have quite mixed results: The negative elasticities for NLLS and DOLS K=1 estimations do not fit reality; For DOLS K=2 and 3, the income elasticities are positive and elastic.

Based on the evolution of the emissions from CO2 from transport sector and on the above results, model (1.6) seems to predict more accurately the long-run CO2transp-GDP pattern. Figure 1.3 plots the estimated CO2transp-GDP path considering the cubic level model.

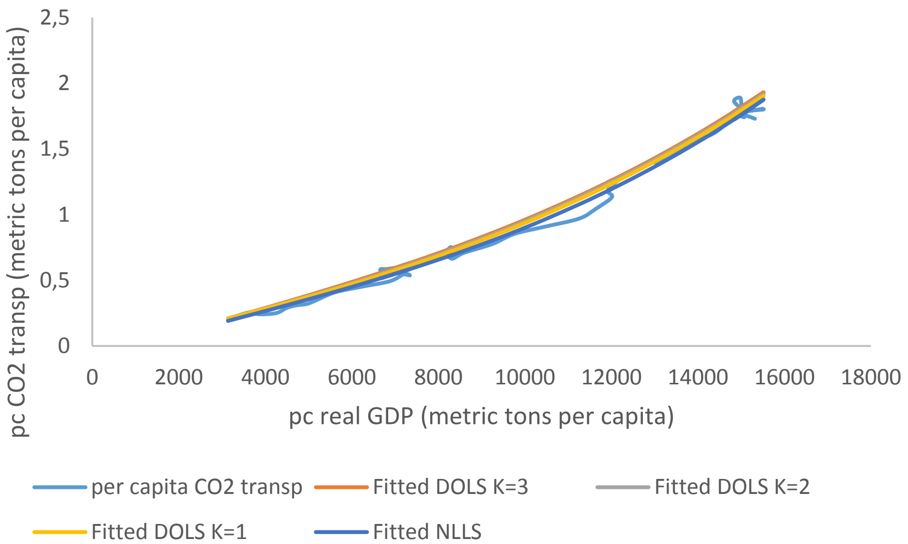

Figure 1.3: Per capita real GDP versus per capita CO_2 emissions from transport sector for Portugal (1960 to 2010).

[7] Similar results are found for DOLS K=1, 2 and 3 estimations.

Total CO_2 emissions

The estimations of the quadratic and cubic polynomial specifications, both in levels and logs, are presented in table C.3, appendix C, for total CO_2 emissions as the dependent variable.

Considering model (1.9), with the exception of the constant, the estimated coefficients for the linear and quadratic income terms are positive, pointing to a monotonically rising relationship between CO2total and GDP. Conversely, based on model (1.10), the relationship follows an inverted N-shaped curve with a minimum estimated turning point of €2,447.1 per capita after which the emissions start to rise with per capita GDP, until reaching the maximum turning point of around €16,421.3 per capita. Thus, the EKC hypothesis holds from the moment the minimum turning point is reached.

The estimated values for both turning points call for particular attention because they are outside the observed sample range, yet very close to it. This could indicate that, at the present, the CO2total emissions are in the upward part of the EKC but the rate of emissions seems to be slowing down as the maximum turning point is slightly above the largest observed per capita GDP value.

Instead of being contradictory, the results of models (1.9) and (1.10) are in line with each other. The quadratic form is more restrictive than the cubic one; consequently it only partially captures the relation CO2total-GDP. The cubic specification goes beyond indicating that, in fact, CO2total emissions are still increasing but they are close to stabilizing.

Models (1.11) and (1.12) represent quadratic and cubic specifications in logs, respectively. Concerning the shape of the relationship, there are no significant differences when compared with the corresponding models in levels. However, it should be noted that the maximum estimated turning point of the cubic log model is €25,351.3 per capita, a considerably higher value than cubic model in levels. Because of this, the cubic log model indicates a monotonic increasing link between CO2total and GDP.

The log-log specification also allows examining the estimated income elasticity of CO2total. According to the quadratic log estimation, the income elasticity is positive but less than one: The total emissions increase by around 0.59% when GDP increases by 1%. This result is consistent with the conclusions drawn from the models in levels, indicating the gradual decoupling of emissions from economic growth.

The elasticity of -72% from the cubic log NLLS regression is not consistent with any of the previous results, as it would imply that total emissions have already attained the downward slope of the EKC, which seems not to be the case.

The above results suggest the cubic level model as the most appropriate one to define the long-run relationship between CO2total emissions and GDP again.

The results for the estimated cubic-levels model are plotted in figure 1.4. It is possible to see that regardless of the estimator, the outcomes are very similar.

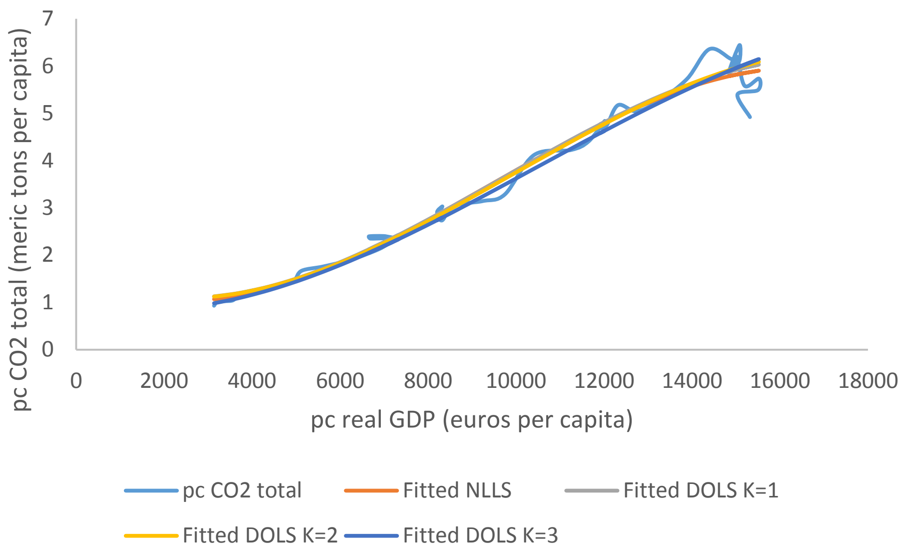

Figure 1.4: Per capita real GDP versus per capita total CO_2 emissions for Portugal (1960 to 2010).

1.4.3. Extended model with other explanatory variables

This section includes the estimation of regressions, in levels and logs for both quadratic and cubic specifications, with further explanatory variables that may influence CO_2 emissions besides income. Along with additional economic variables – crude oil price, average fuel price and rate of motorization – temperature and precipitation are also explored as possible determinants of CO_2 emissions.

Due to the statistical limitations of the nonlinear cointegration methodology and in particular the inability to test for the statistical significance of the regressors, we make use of the linear error correction model (ECM)[8] to have a more concrete idea about the significance of the new independent variables.

CO_2 emissions from electricity generation

The extended models for the electricity generation sector encompass, besides GDP, the crude oil price, precipitation and temperature.

Considering the results depicted in table D.1, appendix D, the inclusion of additional variables, hardly affects the size of the income-related estimated coefficients for electricity emissions for both for level (models (1.13) and (1.14)) and log (models (1.15) and (1.16)) specifications. As expected, crude oil price and precipitation lead to carbon emissions reduction and based on the error correction model (ECM) outcomes the three variables are significant at a 5% level. Electricity is a secondary energy carrier that can be generated using distinct technologies and a diversity of primary energy sources. The costs of

[8] Table with the ECM results available upon request.

power generation are influenced by the price of the primary energy used, and thus the crude oil price fluctuations and precipitation levels play a crucial role in the development of CO_2 emissions from power generation. On one hand, crude oil prices have indirect impacts on energy sources used to produce electricity such as fuel oil, natural gas and even coal. This significant relationship may be because some primary energy sources have a substantial amount of crude oil in their production, or due to oil-indexed contracts which can have major impacts for natural gas, coal or fuel oil fired power plants. So, when the price of crude oil rises it impacts the energy mix of power generation, which has a direct effect on CO_2 emissions. On the other hand, the installed hydroelectric capacity in Portugal corresponded to 26%[9] of the total installed capacity in 2010. Moreover, about 30%[10] of the electricity produced from renewable energy sources (RES-E) was hydropower. This substantial weight of hydropower on the Portuguese electricity sector endows the level of precipitation great importance to CO_2 emissions development. Besides being free of charge, precipitation turns into river flows and, in dams, becomes a major energy source without directly emitting carbon dioxide.

As for temperature, the expected negative sign is only verified for model (1.14). This result is rather questionable if we take into consideration that the temperature-related electricity consumption in Portugal is mostly for heating. These results are not particularly relevant to our work as the ECM provide evidence that temperature is not statistical significant. Moreover, this is reinforced by Mota and Dias (2006) results. The authors also tested the importance of temperature for Portugal and found negative estimated coefficient, though it is not statistically significant.

Figure 1.5 displays the estimated regression plots.

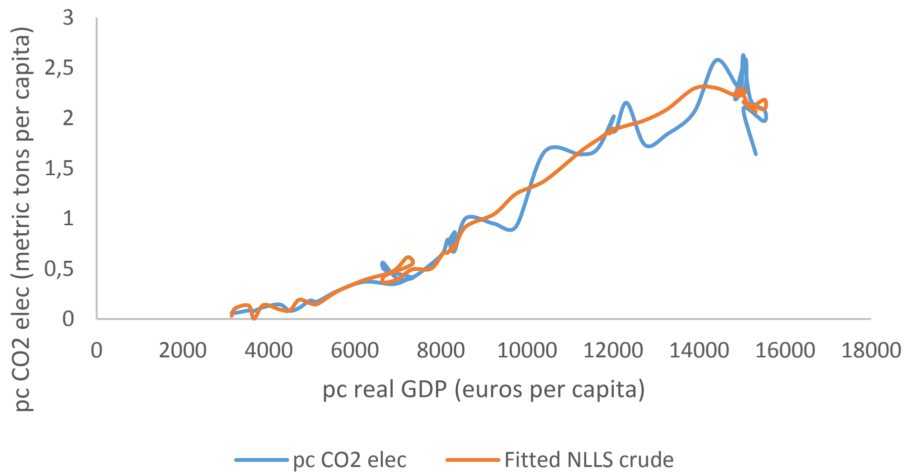

Figure 1.5: Per capita real GDP versus per capita CO_2 emissions from electricity sector for Portugal with additional control variables (1960 to 2010).

[9] www.ren.pt
[10] www.dgeg.pt

CO_2 emissions from transport

In the case of transport, the additional variables tested as possible regressors are crude oil prices, fuel oil prices and rate of motorization. Similar to electricity sector results, the insertion of such variables, does not produce considerable changes in the size and signs of the main regression coefficients, as can be noted by the results provided in table D.2, appendix D.

For both quadratic (models (1.17) and (1.18)) and cubic (models (1.19) and (1.20)) in levels specification, crude oil and fuel price coefficients have the expected negative signs. An increase in either crude oil or average fuel prices has a direct impact on kilometre (km) travelled, reducing these, since a rise in prices not only translates into a decrease in travel demand but also in people choosing alternative mobility options such as public transportation, cycling, electric vehicles or even walking. Fewer km travelled and mobility choices less dependent on fossil fuels have direct effects on CO_2 emissions from the transport sector.

According to the ECM, based on a significance level of 5%, together with GDP, the average fuel price is the only control variable that explains CO_2 emissions from transport sector.

The rate of motorization, which would be predictable a positive impact on carbon emissions, presents negative estimated coefficients for all models. These results are not only contradictory to those obtained for total CO_2 emissions (below), but also very doubtful. Most probable, a greater number of vehicles per 1000 inhabitants is positively related with the distance travelled, and thus with CO_2 emissions. However this result can be neglected because according to the ECM procedure, the rate of motorization it not statistically significant.

From the quadratic (models (1.21) and (1.22)) and cubic (models (1.23) and (1.24)) log specifications a 1% increase in crude oil price or in average fuel price leads to an increase of less than 1% in CO_2 emissions. These results are not in line with those in levels. Presumably, a 1% increase in prices should lead to a decrease in CO_2 emissions from transport. However, according to these estimated elasticities, there is a small CO2transp growth response to a rise in prices of crude or fuel. The negative elasticities associated to the rate of motorization are obviously implausible. Note that the results for the crude oil prices and rate of motorization should be ignored relying on the ECM outcomes.

Again, the unrealistic result for average fuel prices, prove that the specification in levels should prevail upon logs.

Our results for CO2transp are consistent with those obtained by Mota and Dias (2006). In their regression they include the service sector that encompasses, among other activities, the transport sector. They argue that the positive sign of services' coefficient is possibly due to the inclusion of transportation.

Figure 1.6 shows the estimated regression plots.

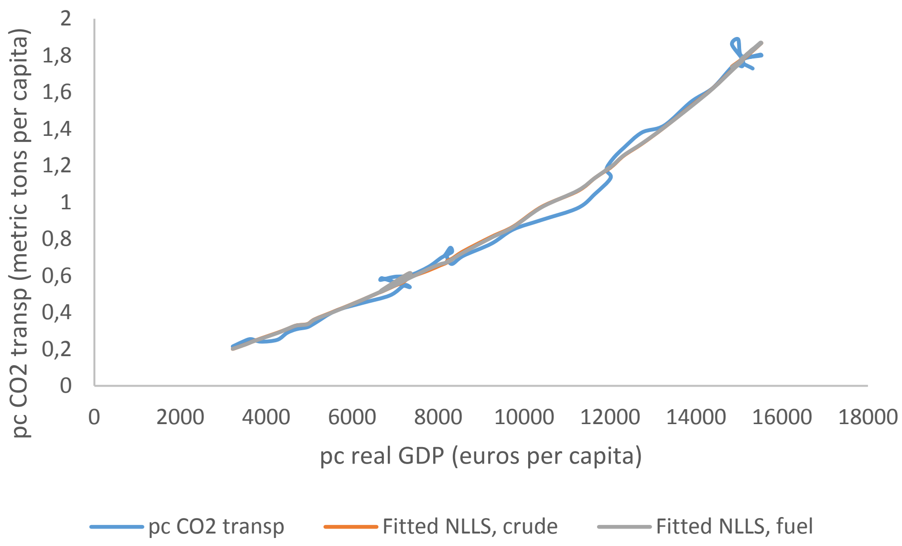

Figure 1.6: Per capita real GDP versus per capita CO_2 emissions from transport sector for Portugal with additional control variables (1960 to 2010).

Total CO_2 emissions

The estimation results for total CO_2 emissions' models are displayed in table D.3, appendix D. While the size of the GDP-related coefficients is relatively similar when compared to the reduced-form specifications, the other estimated coefficients appear to have the expected signs even though, at this point, we cannot fully assess their statistical significance. The crude oil price, average fuel price and precipitation have negative coefficients, as expected, for all models. The negative signs of crude and fuel prices can be explained by the demand impact of price changes: whenever prices increase, the quantity of energy demanded decreases, leading to CO_2 emission reductions. As noted above, a possible reason for the negative impact of precipitation is hydropower capacity, which accounts for an important share of the total Portuguese installed capacity. High precipitation levels allow for more hydropower generation, and subsequently, the reduction of CO_2 emissions released by fossil fuel powered generation facilities.

As for the rate of motorization, all regressions' coefficients are positive, possibly because greater private vehicle ownership rates might be associated to an increase of vehicle travelled.

Temperature is the only variable which presents distinct and more controversy results. Except for model (1.32), the coefficients for temperature are positive. We believe that these results should be put in perspective due to the reason mentioned above for the CO2elect sector.

The ECM offers a conjecture about the statistical significance of the independent variables. Only GDP, crude oil price and average fuel price are significant at 5%, and precipitation is significant at the 10%

significance level, with the respective expected signals. The rate of motorization and temperature do not explain the trajectory of CO2total, thus the estimated coefficients should be ignored.

For both NLLS and ECM estimations, the size of the average fuel price coefficient is higher than those of the average crude oil prices. This might indicate that the fuel price has a more substantial impact on the total CO_2 emissions than the crude price, which is consistent with the plots of figure 1.7, but calls for further analysis. National carbon dioxide emissions aggregate emissions from distinct economic activities, among which electricity generation and transport stand out as being the largest emitters. The role crude oil price and average fuel price play in the CO_2 trend is better understood within the sectoral investigation of the CO_2 emissions.

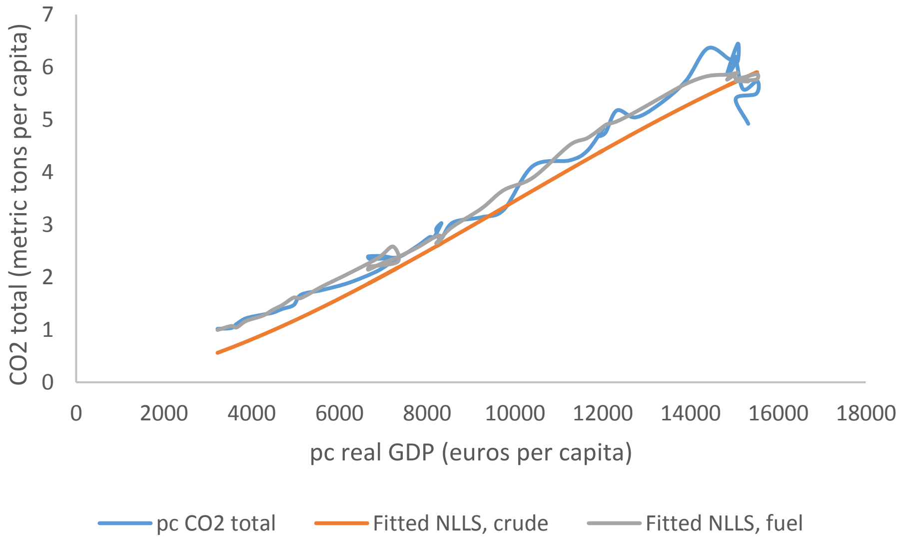

Figure 1.7: Per capita real GDP versus per capita total CO_2 emissions for Portugal with additional control variables (1960 to 2010).

1.5. Discussion

In this chapter we have performed a more traditional EKC analysis of total CO_2 emissions but also looked at the EKC from a disaggregated, sectoral, position. Concerning the electricity generation sector, although only Moutinho *et al.* (2014) have examined the evidence for the EKC hypothesis for this sector (also using Portugal as a case study), the relevance of electricity for CO_2 emissions is clear from the work of several authors who test independent variables such as energy consumption and the share of nuclear power and renewable energy for electricity production (e.g. Iwata *et al.*, 2011; Burke, 2011; Sulaiman *et al.*, 2013). A general conclusion that can be drawn from previous works is that the primary energy source for power generation is crucial for the evolution of CO_2 emissions. These results are coherent with our findings for precipitation, which is also a primary energy source used to generate electricity through its role in hydropower.

Focusing on the CO2elect-GDP long-run relationship, GDP is indeed a key determinant for electricity demand, and our results validating the existence of the EKC are similar to those obtained by Moutinho *et al.* (2014).

From our point of view there are two main reasons why economic growth is no longer synonym of an increase of CO_2 emissions from electricity. The first one is, as Burke (2011) states, because the economic development enables the emergence of more environmentally friendly power generation technologies such as the efficiency improvements of the thermoelectric power plants with the introduction of gas-fired combined-cycle power plants, which are more efficient and less polluting. Between 1990 and 2010, the efficiency[11] of the electricity generation sector rose from 48% to 59%, correspondingly. Efficiency gains were particularly significant in thermal generation, due to the commissioning of new gas-fired combined cycles plants. Along with the efficiency increases, there is also the deployment of renewable technologies, not only hydro but also wind power.

The second reason concerns the more demanding regulatory framework as a consequence of the economic growth. The electricity generation sector is covered by the European Union Emissions Trading Scheme (EU ETS). Similar to what happens in the other Member-States, in Portugal the electricity generation sector has the highest weight of all installations covered by the Portuguese National Allocation Plan[12] (NAP) I (2005-2007) and II (2008-2012). From all the installations considered, the thermoelectric power plants were responsible for more than a half of CO_2 emissions in the period comprehended between 2005 and 2010. Thus, it is evident the effort to achieve significant GHG reductions from the electricity generation sector, corresponding to a decrease of 30% in NAP II when compared to NAP I (ERSE, 2012). From 2005 to 2010, GHG emissions from the thermal power plants decreased from 678 to 579g CO_2 / kWh. This reduction was mainly due to less use of fossil fuels as primary energy source and to a greater contribution of hydro and wind power production source (ERSE, 2012). Additionally to the EU ETS, the country has a legal context capable of promoting and encouraging the investment in renewable energy sources for power generation at early stages of development, such as the mandatory purchase and subsidies.

The present work does not include renewable energy sources (RES) as an independent variable because it is assumed that it is encompassed in income growth (technological effect). Despite the absence in our models, the RES is part of the technological effect of the EKC hypothesis and thus it probably explains part of our results. However, probably, regressions with better quality would be obtained if we included RES-E installed capacity as regressors. Notwithstanding, it can be inferred that with the economic growth the effect of renewable electricity production, mainly hydro and wind, overlapped the effect of electricity demand on CO_2 emissions, which justifies the descendent trajectory.

[11] Data source: http://www.wec-indicators.enerdata.eu/power-generation-efficiency.html
[12] NAP establishes what are the installations covered by the EU ETS and defines the respective number of licenses.

Our empirical analysis demonstrates a completely different scenario for the transport sector. The monotonically increase of CO_2 emissions from transport with per capita GDP is, nonetheless, consistent with most previous work for other countries (Liddle, 2004; Abdallah *et al.* 2013; Cox *et al.*, 2012). As in the electricity generation sector, transport also benefits from the technological effect of economic development, namely more efficient engines and fuels with better quality. Nevertheless, the technological effects are in some measure absorbed by other effects responsible for driving up transport sector CO_2 emissions. Probably, some of the factors pointed out by Meunie and Pouyanne (2009) also help to explain the increase of carbon emissions along with GDP growth for Portugal. More specifically, the combination between higher income and behaviours like car ownership and distance-travelled per car per year or the purchase of more potent vehicles with higher energy consumption, tend to overlap the advantages of more efficient technologies and fuels of better quality.

The contrasting findings for the sectors we have studied shed some light on the relationship between CO_2 emissions and income, with policy implications. Our results may be used as a tool for public policy design and decision-making. If Portugal wants to continue to fulfil international commitments in terms of greenhouse gases, it has to be aware of the different drivers of emissions in these two sectors.

The impact of each sector on the performance of total CO_2 emissions can be drawn from the fitted curves plotted in figure 1.8. The split of the total CO_2 emissions' curve enlightens the slowing trend of emissions' increase. The decrease of CO_2 from electricity generation is lowering emissions growth, while the transport sector is pushing up CO_2 emissions. The combination of these two counteracting forces is a plausible justification for the validation of an EKC shape for total emissions. This finding contradicts the results of Mota and Dias (2006) and Acaravci and Ozturk (2010), yet is in line with Shahbaz *et al.* (2015).

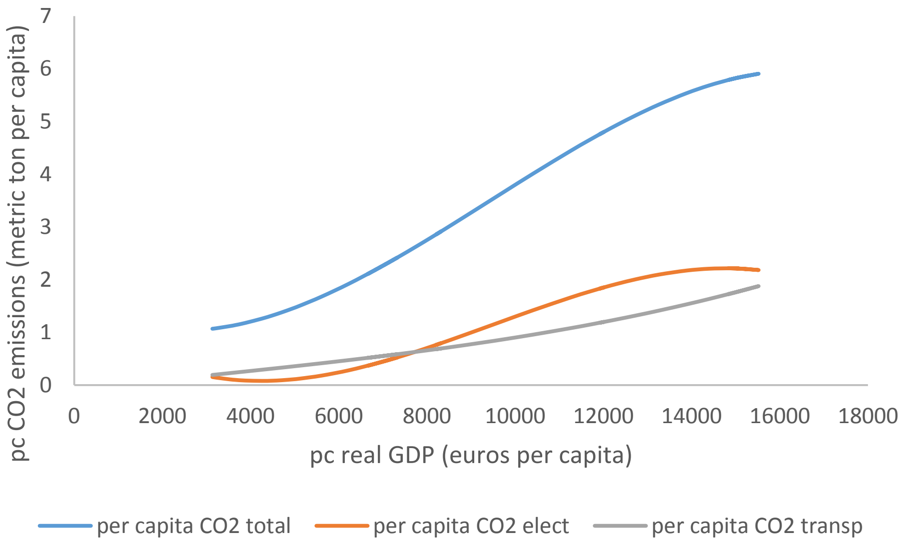

Figure 1.8: Long-run relationship between per capita real GDP and per capita CO_2 emissions for Portugal (1960 to 2010).

Based on these new insights, policy-makers may take note on the way these strategic sectors have been handled, with differentiated measures and goals. Focusing on the electricity generation sector, the support policies applied so far have increased production from RES and thus effectively diminished CO_2 emissions. Now, Portugal ought to take full advantage of existing RES-E installed capacity and continue to promote its diffusion. As for the transport sector, to achieve sustainability it will inevitably go through several changes, as it has been insufficiently targeted in existing policies. The regulatory framework should be improved to be capable of reversing the progression of CO_2 emissions.

As noted in Sousa *et al.* (2015): *"Travel behaviour, irrespective of the mode of transport selected, is influenced by several factors, such as fuel prices (including taxes), infrastructure, demographics, changes in lifestyle, and the price and quality of public transport. Such public policies should also be able to incorporate and balance any conflicting impacts of higher income levels on transport, mainly road transport, which is by far the most critical source of CO2 emissions. Indeed, richer families can buy more than one car and choose more powerful engines, while at the same time become more demanding, when it comes to environmental quality. Sustainable transport policies will allow for a reduction in the number of travels, a decrease in distances travelled, and the adoption of alternative modes of transport such as public transport, cycling, walking, and the substitution from both car and aviation by rail for intercity travel. (…) As incomes rise, governments have at their disposal several measures that may contribute to an offsetting reduction in CO2 emissions from transport, such as economic incentives (price increases through fuel taxes), campaigns to raise awareness on climate change, and policies to support alternative modes of transport like carsharing, cycling, walking, and the use of public transport. However, it is essential to know which measures best fit each national reality, and how they should be implemented, so that they may lead to a reduction in CO2 levels, and at the same time promote inclusion and social equality, without compromising economic growth."* (Sousa *et al.*, 2015, pg.12 and 13)

1.6. Concluding remarks

Time-series data and cointegration analysis are often applied in empirical EKC studies. The polynomial reduced-form specifications of the EKC demand for nonlinear cointegration techniques to achieve reliable results. In this chapter we assess the relationship between economic growth and CO_2 emissions for Portugal, for a period of 51 years, by undertaking a nonlinear cointegration analysis on time-series data, at both aggregate and sectoral scales for the two most important sectors: electricity generation and transport.

The work developed in this chapter not only allows us to corroborate the presence of a connection among economic growth and CO_2 emissions, but also provides evidence on the nature of that connection. While most EKC works focus on either regression in levels only or in logarithms, it is truly an empirical question which specification fits the data better (Lieb, 2003). To help clarifying this question we test models in levels and natural logarithms for both quadratic and cubic reduced functional forms.

In general, the cubic model in levels proves to be the best one for the three cases under study. Our findings are of further relevance because this is the first empirical study that makes use of proper nonlinear cointegration techniques.

Contrary to the bulk of the EKC literature for Portugal, we find evidence of an EKC for both the electricity sector and aggregate emissions. With respect to the transport sector, there is an indication of a positive monotonic relationship between economic growth and CO_2 emissions. While early EKC studies were interpreted, optimistically, as an indication that growing economies would eventually solve environmental problems, it has since become clear that policy choices are a crucial tool to achieve EKC-type turning points. Our sectoral analysis shows that differentiated measures and goals undoubtedly have an impact on the relationship between emissions and GDP. In the electricity generation sector, participation in the EU ETS and generous support policies for renewable energy sources have significantly increased production from such sources and thus effectively diminished CO_2 emissions. No equivalent policies were applied in the transport sector, on the contrary – most investment was directed to high-emission road transport. Our work confirms that in order to break the link between emissions and economic activity, environmental policies must cover all relevant sectors, albeit taking into account sector-specific characteristics.

Apart from GDP and its powers, we also test for other economic and climate independent variables. To our knowledge, this is the first empirical work of the EKC hypothesis that includes rate of motorization and precipitation as regressors.

The results indicate that crude oil prices and precipitation help to explain the CO_2 emissions from electricity generation sector, and that average fuel price is a determinant of transport emissions. As for total CO_2 emissions in Portugal, crude oil prices, fuel prices and precipitation are found to be statistically significant explanatory variables.

Even though this work seeks to contribute to the statistical robustness of the EKC, further theoretical and empirical investigation is necessary prior to any indubitable conclusion can be drawn on the existence and validity of the EKC. Thus, in the subsequent chapter we will continue to examine the relationship between carbon dioxide emissions and economic growth focusing on another common econometric caveat in the empirical applications, which is the possibly neglected structural break(s) in the model that may induce to misleading conclusions. Although in the recent years this matter has been addressed by some authors (see for example Esteve and Tamarit, 2012), the majority of the literature pays little attention to it. As one of our aims is to endow to deepen the knowledge about the link between economic growth and emissions, the next chapter evaluates the existence of the EKC's underlying economic rationale employing unknown structural break techniques in cointegration models.

2. Structural breaks in the Economic growth - Portuguese CO_2 emissions relationship

2.1. Introduction

The nonlinear cointegration methodology performed in chapter 1 aimed to sort out one of many critiques that the Environmental Kuznets Curve (EKC) has been subject to. Still, there are many other criticisms that are made to EKC theory and that were not addressed in the previous chapter. The imposition of quadratic or cubic polynomials is not only seen by several authors as too restrictive to illustrate the relationship between pollution levels and income growth (Lindmark, 2004; Liddle and Messinis, 2014), but also as the cause of econometric caveats such as multicollinearity because of the squared and cubic terms of GDP (Narayan and Narayan, 2010). An alternative is to use a linear reduced form, because as Liddle and Messinis (2014) stress, we can achieve the same conclusions to those we would if we had used a nonlinear reduced form without fixing a priori a specific pattern for the CO_2 emissions and GDP relationship. Yet, even with the linear reduced form, econometric problems may arise because it is relevant to assess if the cointegrating relationship is stable or if it changes over time.

The EKC hypothesis is a long-run empirical phenomenon, so it is very likely that the pollution-income nexus suffers changes over time due to structural transformations. The world is permanently subject to structural changes (breaks) originated by several national motives such as legislative, institutional, political, and technological changes, or more global events like oil price shocks. Time-series regressions should take into account the possible structural changes that induce modifications in the estimated parameters over time.

Perron (1989) demonstrates that standard tests of the unit-root hypothesis against trend-stationary alternatives fail to reject the unit-root hypothesis if the true data is a stationary linear trend process with structural break. Additionally, Kejriwal and Perron (2008, 2010) show that the stability tests can reject the null hypothesis of coefficient stability when in fact there is no cointegration (i.e. the regression is spurious).

Thus, assuming that the long-run relationship between carbon emissions and economic growth is stable can lead to erroneous conclusions. Since nonlinearity and instability are in general difficult to differentiate and are both compatible (Esteve and Tamarit, 2012), we can take a different perspective of the results of chapter one. Perhaps the EKC shape has its explanation on the CO_2-GDP relationship's instability along the years. Macroeconomic events, technological progress or environmental regulation are incorporated in the CO_2-GDP nexus, changing the relationship's pattern. If the presence of structural breaks is confirmed, the relationship CO_2-GDP remains linear but different when a regime shift occurs, which can alter the relationship's slope in sign and/or absolute values.

Within this framework, the magnitude of scale, composition and technological effects associated to an increased demand for environmental quality and a better regulatory framework, are insufficient to explain the existence of the inverted-U shape of the EKC.

Although it is highly probable that structural breaks occur, few EKC studies take them into account. For all we know, Moonaw and Unruh (1997) conduct the first analysis with structural breaks under the EKC hypothesis for CO_2. The sample includes 16 OECD countries, from 1950 to 1992. To introduce exogenous shocks, a piece-wise linear spline function is considered for the periods 1950 to 1973 and 1974 to 1992, for the exogenous structural break in 1973, which was the year of the first oil price shock. The Chow test is performed for structural change. The regression is applied to each nation individually and also to the full panel. The break in 1973 is confirmed, and the results are similar for both approaches (country or full panel). In the first regime, there is a high positive correlation between carbon emissions and GDP for all countries. As for the second regime, only half of the countries have statistically significant coefficients at 95% level; of these, six are negative and two are positive, and based on the adjusted R-squared values there is only strong evidence of negative correlation for France and Sweden. The outcomes for the panel data are similar. For the period before 1973, the coefficients are positive and highly significant. After 1973, there is evidence of a negative relation between CO_2 and GDP. Overall, the main conclusion of this empirical work is that the declining of CO_2 emissions is not related with income levels. Instead, is a consequence of an historical exogenous shock (Moonaw and Unruh, 1997).

Lanne and Liski (2004) explore the properties of the time series of per capita CO_2 emissions in 16 OECD countries over the period 1870–1998 to detect possible endogenous structural breaks in the slope of the trend. The econometric technique is based on univariate unit-root tests which endogenously select multiple structural breaks using a sequential procedure. Only six countries show evidence of a regime-wise stationary emissions trend with a downturn in slope. In most cases, the endogenous structural breaks are identified in the early 20th century, which refutes the Moomaw and Unruh (1997) findings of a break during the oil crises in the 1970s. Only for the United Kingdom and Sweden do downturns in trends entail downward sloping stable trends.

Huntington (2005) focuses his analysis on the United States. He runs a single-break procedure to address the endogenous break in the per capita CO_2 emissions and per capita real GDP relationship for the country for the period 1870-1998, taking in consideration the impact of technological progress on emissions' level. He uses several Chow tests and the Andrews-Quandt approach to estimate the break date in 1913. Before and after the regime shift date the results suggest a stable income elasticity of 0.9, which means that 1% increase in per capita real GDP has a positive impact of 0.9% in per capita CO_2 emissions for constant time trends that control for technological progress, proxied by time trend. However, in association with technological trend effects, the emissions decrease only if the growth of GDP is under 1.8% per year. Otherwise, CO_2 emissions increase.

The empirical research of Esteve and Tamarit (2012), which provides the basis of our work, examines the relationship between per capita CO_2 emissions and GDP for Spain over 1857-2007. They resort to a cointegration model accounting for endogenous structural breaks, using the instability test approach developed by Kejriwal and Perron and the cointegration tests proposed by Arai and Kurozumi. According to the Kejriwal and Perron tests, there are two estimated break dates at 1971 and 1967 and, thus, three regimes. To compare the results with and without structural breaks, the authors estimate the model for the full sample and for the three subsamples. The long-run income elasticity over the three regimes is declining but positive (2.67, 1.10 and 0.56). Hence, albeit per-capita CO_2 increases with income growth, the income elasticity is less than one. This indicates that although the shape of the EKC does not follow an inverted U, the decreasing income-elasticity pattern suggests a forthcoming turning point. Concerning the full sample estimation, the long-run income elasticity is 1.37, which is considerably greater than the value for the third regime. Based on these findings, Esteve and Tamarit (2012) conclude that neglecting structural changes in the long-run cointegration relationship between CO_2 and income may overemphasize its magnitude.

Based on the same methodology as Esteve and Tamarit (2012), Liddle and Messinis (2014) extend the country-specific analysis of per capita carbon dioxide emissions and per capita real GDP to 23 OECD countries, including Portugal, over 1950-2010. The EKC hypothesis is confirmed for four countries. Specifically for Portugal – to our knowledge, this is the first and only study in the literature – there is one estimated break date at 1988. The income elasticity declines from regime one (1.086) to regime two (0.993), but still the elasticity value over the last regime remains high and near proportional. Given these results, the authors question the existence of a regime shift for Portugal. Similar conclusions are obtained for Spain, which goes in the opposite direction of Esteve and Tamarit (2012). Nevertheless, Liddle and Messinis (2014) stress that the time span has influence on the outcomes.

Following the works of Esteve and Tamarit (2012) and Liddle and Messinis (2014), in this chapter we test the EKC hypothesis through a linear reduced form. We will examine the EKC hypothesis using cointegration with multiple unknown structural breaks for both the electricity generation sector, the transport sector and nationwide. Furthermore, the same control variables as in chapter 1 will be included.

2.2. Methodology under Breaks

This section provides a concise overview of the methodological approach applied to conduct the empirical study focusing on the possible existence of cointegration with structural breaks.

We assume that the CO_2 emissions, GDP and control variables have no structural breaks, the results of the unit root tests conducted in chapter 1 remain valid, and thus we carry on the cointegration analysis with unknown multiple structural breaks.

2.2.1. Econometric specification

In this Chapter, the EKC hypothesis and the importance of additional control variables are tested using linear models, in levels and natural logarithms. Thus, the general model excludes the square and cube of GDP and is represented as follows:

$CO2elect = f(y, crude, temp, precip)$

$CO2transp = f(y, crude, motor)$ or $CO2transp = f(y, fuel, motor)$

$CO2total = f(y, crude, temp, precip, motor)$ or $CO2total = f(y, fuel, temp, precip, motor)$

Our procedure starts by examining the long-run relationship between carbon emissions variables and GDP, labelled by "y". To do so, the following reduced-form equations are estimated for both electricity generation and transport carbon emissions in addition to total CO_2 emissions.

$$CO2_t = \alpha_j + \beta_j y_t + \varepsilon_t \tag{2.1}$$

$$lnCO2_t = \alpha_j + \beta_j ln(y_t) + \varepsilon_t \tag{2.2}$$

Where $CO2_t$ is the per capita CO_2 emissions and y_t the per capita real GDP in year t, the coefficients are associated to regime j, and ε_t is the stochastic error term.

Then we include the control variables in the equations, as follows:

$$CO2_t = \alpha_j + \beta_j y_t + \beta_k V_t + \varepsilon_t \tag{2.3}$$

$$lnCO2_t = \alpha_j + \beta_j ln(y_t) + \beta_k ln(V_t) + \varepsilon_t \tag{2.4}$$

Where V_t is a vector of k additional variables – crude oil prices, average fuel prices, rate of motorization, temperature and precipitation.

These equation specifications allow us to investigate the potential existence of regime shifts in the long-run relationship between carbon emissions and GDP. If, for each regime, the impact of GDP on CO_2 emissions remains positive then the relationship between the variables is monotonic, but the emissions growth pace may be declining or increasing from one regime to another. If the growth pace falls along the years we can draw a parallel with the upward part of the EKC near the turning point value, which means that income development has contributed less to carbon dioxide emissions. A negative impact of GDP on emissions in a following regime that occurs after the turning point moment is coherent with the EKC hypothesis. As Esteve and Tamarit (2012) state, from a dynamic perception, the EKC hypothesis entails that along with the beginning of economic growth is the environmental degradation and pollution increase, after reaching a certain level of per capita income this pattern changes so that beyond the turning point level of income there is environmental improvements or, at the minimum the degradation is smoother which means an environmental improvement in relative terms. This indicates that even though the environmental impact indicator is not an inverted U-shaped function of per capita income, it reveals a declining growth trend over time (Esteve and Tamarit, 2012).

2.2.2. Cointegration tests

Since the variables of the models are I(1) processes, we now test the existence of a cointegrating relationship with structural breaks. Thus, we consider both the Gregory and Hansen (1996, hereafter GH (1996)) and Arai and Kurozumi (2007, hereafter AK (2007)) tests.

Gregory and Hansen (1996)

The GH (1996) method is an extension of the Engle-Granger's residual-based test which allows for a single unknown endogenous structural break in the cointegration vector, and it tests the null hypothesis of no cointegration against the alternative of cointegration with one regime shift.

GH (1996) propose ADF, Z_α, and Z_t tests for cointegration, defining three distinct models for the analysis of structural change in the parameters of the cointegrating vector. Let the observed data be $y_t = (y_{1t}, y_{2t})$, where y_{1t} is real-valued, y_{2t} is a $I(1)$ m-vector, and e_t is the error term.

The structural change is captured by the dummy variable denoted by $\varphi_{t\tau}$ which is indicated as follows:

$$\varphi_{t\tau} = \begin{cases} 0 & if \quad t \leq [n\tau] \\ 1 & if \quad t > [n\tau] \end{cases}$$

where $\tau \in (0,1)$ is the unknown break fraction parameter which denotes the (relative) timing of the change point for a sample of size n, and [ ] indicates the integer part.

Model C – Level shift:

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha^T y_{2t} + e_t, \quad t = 1, \dots, T \tag{2.5}$$

where μ_1 represents the intercept before the shift and μ_2 is the change in the intercept at the moment of the shift, $[T\tau]$.

Model C/T – Level shift with a time trend:

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \beta t + \alpha^T y_{2t} + e_t, \quad t = 1, \dots, T \tag{2.6}$$

Model C/S – Regime shift:

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha_1^T y_{2t} + \alpha_2^T y_{2t}\varphi_{tT} + e_t, \quad t = 1, \dots, T \tag{2.7}$$

Here μ_1 and μ_2 are as in model (2.5), α_1 indicates the slope coefficients prior to the regime shift and the change in the slope coefficients is denoted by α_2.

To trace the cointegrating relationship with the possible existence of structural break, for each of the above models, estimate its OLS residuals $\hat{e}_{t\tau}$ for fixed τ. Based on these residuals, the bias-correct first order serial correlation coefficient is:

$$\hat{\rho}_\tau^* = \sum_{t-1}^{n-1}\left(\hat{e}_{t\tau}\hat{e}_{t+1\tau} - \hat{\lambda}_\tau\right)/\sum_{t=1}^{n-1}\hat{e}_{t\tau}^2 \tag{2.8}$$

where $\hat{\lambda}_\tau$ is the estimate of a weighted sum of autocovariances, $M = M(T)$ is the selected bandwidth, such that $M \to \infty$ and $M/T^5 = O(1)$.

Thus the Phillips test statistics can be written as follows:

$$Z_\alpha(\tau) = T(\hat{\rho}_\tau^* - 1) \tag{2.9}$$

$$Z_t(\tau) = (\hat{\rho}_\tau^* - 1)/\hat{s}_\tau \tag{2.10}$$

with $\hat{s}_{\tau=} \hat{\sigma}_\tau^2/\sum_1^{n-1}\hat{e}_{t\tau}^2$ and $\hat{\sigma}_\tau^2$ is a long-run variance and

$$ADF(\tau) = tstat(\hat{e}_{t-1\tau}), \tag{2.11}$$

i.e., from the ADF regression applied to the residuals $\hat{e}_{t\tau}$. According to GH (1996), the test statistics (2.9), (2.10) and (2.11), are suitable for cointegration analysis with no structural break. Thus, they develop test statistics accounting for cointegration tests with one regime shift. These statistics are:

$$Z_\alpha^* = \inf_{\tau\in\boldsymbol{T}} Z_\alpha(\tau) \tag{2.12}$$

$$Z_t^* = \inf_{\tau\in\boldsymbol{T}} Z_t(\tau) \tag{2.13}$$

$$ADF^* = \inf_{\tau\in\boldsymbol{T}} ADF(\tau), \tag{2.14}$$

Where, typically, $\boldsymbol{T} = (0.15, 0.85)$. The null hypothesis of no cointegration is rejected when the observed statistic of interest is smaller than the critical value. In our work, we apply the GH (1996) methodology for both model C (level shift) and model C/S (regime shift). An estimate of the break point follows from the corresponding test statistic at its smallest value.

Arai and Kurozumi (2007) and Kejriwal (2008)

Additionally to the GH (1996) cointegration tests, we conduct the procedure proposed by AK (2007) which is a residual-based test for the null hypothesis of cointegration with a single structural break against the alternative of no cointegration. The authors modify the Lagrange Multiplier test of Shin (1994) based on partial sums of residuals where the break point is obtained by minimizing the sum of squared residuals, in order to allow for breaks under the null, and considered the same type of structural breaks as in GH (1996) work.

AK (2007) employ the dynamic OLS estimator by adding the leads and lags of the first differences of the regressors, thus modifying the GH (1996) as follows:

Model C' – level shift:

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \beta' y_{2t} + \sum_{i=-K}^{K} \pi'_i \Delta y_{2,t-i} + \varepsilon_t^*, \quad t = 1, \dots, T \tag{2.15}$$

Model C/T' – Level shift with a time trend:

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha t + \beta' y_{2t} + \sum_{i=-K}^{K} \pi'_i \Delta y_{2,t-i} + \varepsilon_t^*, \quad t = 1, \dots, T \tag{2.16}$$

Model C/S' – Regime shift:

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \beta'_1 y_{2t} + \beta'_2 y_{2t}\varphi_{t\tau} + \sum_{i=-K}^{K} \pi'_i \Delta y_{2,t-i} + \varepsilon_t^*, \quad t = 1, \dots, T \tag{2.17}$$

Where, for a fixed chosen K, π_i is a $(m \times 1)$ parameter vector for all $-K \leq i \leq K$. These are regression models where the leads and lags of Δy_{2t} are added to the original GH (1996) models. The estimator for the break point is obtained by minimizing the sum of the squared residuals over all admissible values defined by:

$$\hat{\tau} = arg \inf_{\tau \in \mathcal{T}} SSR_n(\tau) \tag{2.18}$$

where $\mathcal{T} = [\underline{\tau}, \overline{\tau}]$, $0 < \underline{\tau} < \overline{\tau} < 1$, (usually, the trimming is $\underline{\tau} = 0.15$ and $\overline{\tau} = 0.85$),and SSR is the sum of squared residuals based on the break fraction τ.

As Kejriwal (2008) points out, the AK (2007) testing procedure is some way restrictive as it only encompasses one structural break under the null hypothesis. Consequently, the null of cointegration may possible be rejected when the true data generating process exhibits cointegration with multiple breaks. To solve this caveat, Kejriwal (2008) extend the AK (2007) test through the inclusion of multiple breaks under the null hypothesis. Thus, the test statistic is given by:

$$\tilde{V}_{n\hat{\tau}}(\hat{\tau}) = n^{-2} \sum_{t=1}^{n} \tilde{S}_{t\hat{\tau}}^2 / \widehat{\omega}_{1.2\hat{\tau}} \tag{2.19}$$

where $\tilde{S}_{t\hat{\tau}} = \sum_{s=1}^{t} \tilde{e}_{s\hat{\tau}}$ and $\widehat{\omega}_{1.2\hat{\tau}}$ is a consistent estimator of the long run variance of ε_t^*. See Kejriwal (2008) for the definition of the residual and error terms $\tilde{e}_{s\hat{\tau}}$ and ε_t^*, respectively.

The test statistics are compared with the critical values obtained by simulation (Kejriwal, 2008). The null hypothesis of cointegration with one (or multiple breaks) is rejected when the observed test statistic is larger than the critical value. We run the test for model C/S', for one, two and three breaks.

2.2.3. Structural break tests and correspondent estimated regression

After testing for unit roots and cointegration with structural breaks, the subsequent step is to estimate the cointegrating relationship between per capita CO_2 variables and per capita real GDP, using the Dynamic Ordinary Least Squares[13] (DOLS) by Saikkonen (1991) and Stock and Watson (1993).

DOLS estimation, in our case with 2 leads and lags of the GDP, takes into consideration the possible endogeneity of the regressors, and also it corrects serial correlation in the error terms of the Ordinary Least Squares estimation.

We test the number of structural breaks in our cointegration relationship and the estimation of the break dates and the linear regression models for each regime. We employ the Kejriwal and Perron (2010, henceforth KP (2010) methodology.

KP (2010) consider the linear regression with m breaks, thus $m + 1$ regimes[14]:

$$y_t = c_j + z'_{ft}\delta_f + z'_{bt}\delta_{bj} + x'_{ft}\beta_f + x'_{ft}\beta_{bj} + u_t \qquad (2.20)$$

for $t = T_{j-1} + 1, \dots, T_j$, $j = 1, \dots, m + 1$, where $T_0 = 0$, $T_{m+1} = T$, and T is the sample size. y_t is a scalar-dependent $I(1)$ variable, $x_{ft}\left(p_f \times 1\right)$ and $x_{bt}\,(p_b \times 1)$ are vectors of $I(0)$ variables, $z_{ft}\left(q_f \times 1\right)$ and $z_{bt}\,(q_b \times 1)$ are vectors of $I(1)$ variables. The subscripts b and f denote "break" and "fixed" (across regimes) meaning that the cointegration vector changes through δ_{bj}. The break points $(T_1, \dots, T_m)$ are unknown.

A diversity of particular models can be derived from the data generating process (2.20) which is the most general one. The authors classify these particular models into two categories: a) models with only $I(1)$ regressors, and b) models with both $I(1)$ and $I(0)$ regressors.

[13] For more details see Saikkonen (1991) and Stock and Watson (1993).
[14] Authors' original notation.

KP (2010) consider three types of tests: The first type applies when the null of interest is no structural breaks against the alternative of fixed (arbitrary), $m = k$ number of breaks. They consider the supWald test, scaled by the number of regressors, as follows:

$$\sup F_T^*(k) = \sup_{\lambda \in \Lambda_\epsilon} \frac{SSR_0 - SSR_k}{\hat{\sigma}^2} \tag{2.21}$$

where SSR_0 indicates the sum of the squared residuals under the null hypothesis and SSR_k designates the sum squared residuals under the alternative hypothesis of k breaks. $\lambda = \{\lambda_1, \dots, \lambda_m\}$ is the vector of break fractions defined by $\lambda_i = T_i/T$ for $i = 1, \dots, m,$, $\hat{\sigma}^2$ is the estimated long-run variance of the residuals, $\Lambda_\epsilon = \{\lambda \; : \; |\lambda_{i+1} - \lambda_i| \geq \epsilon, \lambda_1 \geq \epsilon, \lambda_k \leq 1 - \epsilon\}$ and ϵ is some arbitrary small positive number (usually equal to 0.15).

The rejection of the null hypothesis occurs when the test statistic is greater than the critical value. The second type of test concerns the null of no structural breaks against the alternative of an unknown number of structural breaks between 1 and some upper bound M:

$$UDmax\ F_T^*(M) = \max_{1 \leq k \leq M} F_T^*(k) \tag{2.22}$$

According to KP (2010), this test is the most useful for detecting the presence of structural breaks. The rejection of the null hypothesis occurs when the test statistic is greater than the critical value. The third type of test is a sequential procedure for the null hypothesis of k breaks against the alternative of $k + 1$ breaks, given by the expression (see KP (2010) for details):

$$SEQ_T(k+1|k) = \max_{1 \leq j \leq k+1} \sup_{\tau \in \Lambda_{j,\varepsilon}} T\left\{SSR_T\left(\hat{T}_1, \dots, \hat{T}_k\right) - SSR_T\left(\hat{T}_1, \dots, \hat{T}_{j-1}, \tau, \hat{T}_j, \dots, \hat{T}_k\right)/SSR_{k+1}\right\} \tag{2.23}$$

Additionally to the sequential test procedure, we also apply the Bayesian Information Criterion (BIC) and the modified Schwarz' Criterion (LWZ) proposed by Liu, Wu and Zidek (1997) to select the number of breaks and use a trimming value of $\epsilon = 15\%$.

2.3. Empirical results

In this Section, we empirically study the EKC hypothesis by first assessing whether the cointegration with structural breaks exists and, if it does, then by determining the number and break dates of such relationships and finally by estimating the models subject to regime switching.

2.3.1. Cointegration tests

We start our structural break cointegration analysis of CO_2 variables and GDP by running the GH (1996) tests (results in table E.1 of appendix E). The null of no cointegration is not rejected for the ADF^*, Z_α^* and Z_t^* tests, either for the level or for the regime shifts models, for the CO2total-GDP and CO2transp-GDP relationships. These results are valid for both specifications in levels and natural logarithms, at 1%, 5% and 10% levels of significance.

In the case of the CO_2 emissions from electricity generation, the results indicate the existence of cointegration with one structural break. Concerning the specification in levels, for the level shift model, the null is rejected at the 5% and 10% significance levels for the ADF^* and Z_t^*, whereas for the Z_α^* test the null is rejected at the 10% level of significance. The tests conducted for the regime shift model result in some differences. The $ADF^*$and Z_t^* tests rejects the null hypothesis at the 10%, and at the 5% and 10% significance levels, respectively. Based on the Z_α^* test, the null of no cointegration cannot be rejected for all levels of significance.

Regarding the logarithmic specification of the CO2elect-GDP relationship, and for the level shift model, all the test results indicate the existence of cointegration with a structural break. As for the regime shift model case, the ADF^* rejects the null at the 5% and 10% significance levels. The Z_t^* also rejects the null for all relevant significance levels, whereas the Z_α^* test result suggests the inexistence of cointegration.

The results obtained for the AK (2007) and Kejriwal (2008) tests (tables E.2-E.4, appendix E) using the regime shift model (which is more general than the level model), considering the null hypothesis cases of 1) cointegration with one break, 2) cointegration with two, and 3) cointegration with three breaks, confirm the cointegrating relationship with breaks between each of the CO_2 emissions variables and GDP, at all significance levels, for both level and logarithmic specifications (except for the level specification of the CO2total-GDP relationship, where the null of cointegration with one structural break is only accepted for a 1% and 5% of significance levels).

Apart from for the electricity case, the GH (1996) and AK (2007) tests produce contradictory findings. Even so, we assume that CO2total, CO2elect and CO2transp series and GDP are cointegrated under structural breaks. Our decision relies on the fact that as AK (2007) and Kejriwal (2008) stress, the GH (1996) tests may have reduced power when the alternative hypothesis contains more than one break. Moreover, since we are interested in testing for cointegration (with breaks), the null hypothesis of cointegration appears to be a more appropriate choice as it are the cases of AK, 2007 and Kejriwal, 2008.

2.3.2. *Stability tests and structural break dates*

Once the cointegrating relationship between CO_2 emissions and GDP is confirmed including one, two or three structural breaks, it is necessary to identify the exact number of breaks and the corresponding estimated dates. For this purpose, we employ the KP (2010) approach for structural changes. As the value of the trimming is equal to 15%, the maximum number of breaks allowed is of 5, and both the intercept and slope are allowed to change. The results of the stability tests – supWald and DUmax tests – and the number of breaks selected are depicted in tables F.1-F.3 of appendix F. Similar to what Kejriwal (2008) states, we assume that the occurrence of four or five structural breaks are unreasonable considering the time span of the data. The choice of the estimated number of structural breaks is based on the SEQ test and on the LWZ criterion.

For the level specification between CO_2 emissions from electricity generation and GDP, the null hypothesis is rejected for one (5% and 10% significance levels), two (5% and 10% significance levels) and three breaks (10% significance levels), suggesting coefficient's instability. The SEQ test and the information criteria suggest distinct break dates. Based on the SEQ test, there is a single break at 1999 and, consequently, the two regimes at 1960-1998 and 1999-2010 periods. The BIC and LWZ suggest 2 breaks estimated at 1992 and 1995, which gives rise to three regimes for the electricity sector: 1960-1991, 1992-1994, and 1995-2010. Without certainties, the one break scenario seems more reasonable than two breaks, due to the proximity of the two last break dates. The existence of two endogenous structural breaks with estimated dates so next to each other, may not be realistic and statistically undistinguishable.

Looking at the logarithmic specification, the results lead to distinct conclusions when compared to the previous situation. There is evidence of coefficient instability for one and for two (5% and 10% significance levels) breaks. The null of stability is not rejected when the tests are run for three breaks. As for the number of structural breaks, the SEQ test, BIC and LWZ criteria all select the existence of one structural break at 1995 and, thus, two regimes: 1960-1994 and 1995-2010.

Regarding the CO2transp-GDP relationship, the results are in some way inconsistent with those previously obtained for the cointegration tests allowing for structural breaks. The long-run cointegrating relationship between the two variables is put into question by the stability tests for the specification in levels, regardless the number of breaks under test. The non-rejection of the null of no structural breaks for the supWald and the DUmax tests, is coherent with the zero-breaks result from the SEQ test. Conversely, according to the LWZ criterion, the number of estimated breaks is two, at 1993 and 1999, giving rise to three regimes: 1960-1992, 1993-1998 and 1999-2010. The BIC indicates 4 breaks, which seems highly implausible given the dimension of the sample, thus it is not subject to further examination.

Contrasting results are obtained when taking the natural logarithm of the CO2transp and GDP variables. In this instance, the supWald and the DUmax indicate coefficient instability for one break only, at the 5% and 10% significance levels. As for the number of breaks, the results diverge. The SEQ test defines one break at 1998 and, thus, two regimes: the first one from 1960 to 1997 and the second one between 1998 and 2010. The LWZ criterion, as well as for the BIC case, choose two breaks at 1988 and 1990. In this concrete case, one estimated break better depicts the behaviour of the relationship between CO2transp and GDP because having two breaks very close to each other in time is not convincing from the economic and statistical point of view.

Focusing on the specification of the total CO_2 emissions and GDP relationship, the stability tests reject the null of coefficient stability for one, two and three breaks (5% and 10% significance levels). These results indicate the existence of a cointegrating relationship with breaks. The SEQ test as well as the BIC and LWZ criteria, select 2 breaks estimated at 1991 and 1999, and, thus, three regimes: 1960-1990, 1991-1998 and 1999-2010.

With respect to the logarithmic specification, there are mixed results. The null hypothesis of coefficient stability is not rejected for one and three breaks, and rejected at the 1%, 5% and 10% of significance levels for two breaks. The SEQ test determines zero breaks, whereas the information criteria identify 2 breaks. The results obtained for the stability and the cointegration tests support the idea of two breaks and three regimes, the same as those of the specification in levels.

2.3.3. GDP-CO_2 emissions' long-run relationship pattern

Making use of the previously estimated break dates[15], regime-specific linear regressions are estimated for CO2elect-GDP, CO2transp-GDP and CO2total-GDP, which provide twofold information. For each regime, while the coefficients of the equations in levels allow us to find the marginal effect of GDP on the respective CO_2 emissions dependent variable, the logarithmic specifications give us the point estimates of the income-CO_2 emissions elasticities. Because the limiting distribution of the t-statistic is not standard, it won't be possible to check if GDP is statistically significant or not to explain CO_2 emissions. The whole existing theory suggests that it indeed is significant.

CO_2 emissions from electricity generation sector

The DOLS (K=2) regression estimates for the electricity generation sector are shown in table G.1, appendix G. For the specification in levels (model (2.1)), the regressions are estimated for both one and two breaks based on the break dates selected in the previous section. When a single break is considered at period 1999, the estimated slope coefficients for the two regimes have positive signs meaning that an increase in per capita GDP lead to an increase of carbon emissions. Despite the estimated coefficient for the first regime (0.000162) is greater than the one in the second regime (0.000148), the values do not differ significantly from one regime to the next.

[15] We run regressions for at most three breaks. The results are shown in tables G.1-G.3 of appendix G.

If the regression equation contemplates two breaks, the estimated slope coefficients are positive for the first two regimes – 0.000160 and 0.000087 –, with a decline from the first to the second regime, and negative for the last regime – -0.000485. Again, relying on the value of the R-squares, the fit is tighter in the first regime (88% as in the case of one break) and the two other regimes have R-squares below 50%. These results point to rather debatable conclusions as to whether the long-run relationship between CO2elect and GDP has one or two breaks. Concerning the one break case, the fact that there is not much difference between the estimated slope coefficients might indicate that it does not make much sense to have a break in the CO2elec-GDP link. On the contrary, the slope coefficient values for the two breaks regression are in line with the pattern of the Portuguese CO_2 emissions from electricity generation. It is possible to identify a progressive delinking of CO2elect and GDP along the three regimes. Moreover, the negative slope for the regime 1995-2010 supports the existence of the EKC hypothesis, that is, after 1995, economic growth induces an emissions reduction. However, as mentioned earlier, the presence of two consecutive breaks at the close dates 1992 and 1995 is somewhat questionable.

As for the logarithmic specification – model (2.2) – the regression of interest is with one break at 1995. Because of the regime shift, the equilibrium income elasticity of pollution changes when moving from the first to the second regime. The effect of income growth on emissions is strongly positive, higher than unity (2.68) in the first regime. In the second regime, the elasticity remains positive though less than unity (0.78). None of the elasticities is negative, and therefore there is no evidence of EKC and in the second regime the elasticity remains high and positive. Nonetheless, there is a decrease in the long-run effect among CO2elec and GDP. If our time span were longer, we could have verified if the relationship CO2elect-GDP was maintained and, in this case, it was monotonically increasing, or if the emissions-income elasticity continued to decline and, in this case, we were in the presence of the EKC for the electricity generation CO_2 emissions.

CO_2 emissions from transport sector

The estimates concerning the long-run relationship between CO2transp and GDP are reported in table G.2 of appendix G. The results for the specification in levels – model (2.3) – raise some interrogations about the existence of structural breaks in this cointegrating relationship. The negative sign of the estimated slope for the third regime (1999-2010) doesn't quite illustrate the long-run equilibrium CO2transp-GDP, because we know from the data that carbon emissions rise monotonically with GDP. The negative sign could point to misleading conclusions, because it would validate the EKC hypothesis. The R-squared for regimes one and two are about 98%, while for the third regime the coefficient of determination is slightly below 50%, which indicates that only a smaller part of the CO2transp variations are explained by GDP between 1999 and 2010. These results are even more doubtful once we recall that the SEQ test selected zero breaks.

As for the specification in logarithms – model (2.4) – our analysis focuses on regressions with one and two breaks. The same conclusions can be drawn for both one and two breaks. There is evidence that income growth has a strong impact on emissions, because the equilibrium emissions-income elasticities for all regimes are positive and greater than one. Also, the elasticity increases from the first regime to the second and, in the case of two breaks, it increases again in the third regime. These results corroborate the monotonically increasing nexus between CO2transp and GDP and are in line with the time pattern of CO_2 from transport sector along with income growth. Overall, this specification produces more trustworthy findings when compared to the regression in levels, irrespectively of the number of regime shifts considered. Clearly, these results rule out the EKC pattern to describe the relationship between CO2transp and GDP.

Total CO_2 emissions

From the results for the CO2total and GDP long-run relationship, in table G.3, appendix G, it is possible to verify that the conclusions for both specifications in levels (model (2.5)) and logarithms (model (2.6)) are compatible with each other in terms of the signs of the estimated coefficients and also for the R-squared values. Both regressions of interest contemplate two breaks. For the regression in levels, the first (1960-1990) and second (1991-1998) regimes have estimated slope coefficients that are positive with a small increase in the second regime. With regard to the third regime (1999-2010), there is a steep change of the estimated coefficient associated with GDP as it becomes negative. This may support the existence of an EKC for total CO_2 emissions for Portugal, after reaching a certain income level. However, this change may be excessive when looking at the evolution of national emissions over time. It is interesting to observe that the R-squared for regimes one and two are higher than 95% against the value of 44% for the third regime. Apparently, the explanatory power of the model for the last regime is smaller.

The estimated emissions-income elasticity for regime one is positive and higher than unitary. A similar result is obtained for the second regime in which the equilibrium elasticity is greater when compared to the first regime. Thus, during 1960-1990 and 1991-1998 periods, there is a strict link between total CO_2 emissions and GDP. In the third regime (1999-2010), the elasticity is negative (-1.13). The transition to the last regime, where the elasticity changes signs from positive to negative, in theory coincides with a turning time point of the EKC. Thus, in this situation the EKC is considered valid. Similar to what happens for the model in levels, the R-squared values are about 98% and 95% for the first and second regimes, respectively, while the regression of the third one produces a small value which puts into question the reliability of the correspondent equilibrium emissions-income elasticity.

2.3.4. Extended model with other explanatory variables

The regressions including the control variables serve two purposes. The first one is to assess whether the fact that we add other variables to the reduced-form model influences the estimated coefficients of per capita real GDP. The second purpose is to examine the signs of the estimated coefficients of the additional variables and to compare them with those obtained in chapter 1.

Like it was previously mentioned, we should keep in mind that the technique of cointegration with unknown structural breaks does not allow us to properly evaluate the statistical significance of the explanatory variables. In order to tackle this theoretical drawback, we make use of the error correction model employed in chapter 1 as a guide.

The results are shown in tables H.1 to H.5 of appendix H. For all regime shifts considered, the introduction of control variables in the regressions for CO2elec, CO2transp and CO2total produces little effects on the estimated coefficients for GDP, the signals remaining the same, and the R-squared are about equal. Thus, in respect to the GDP variable, the results are almost the same as without the control variables (the same happened for the nonlinear cointegration analysis).

Focusing on the CO2elect regressions – models (2.7) and (2.8) –, the crude oil price and precipitation have the expected negative signs and similar values as those obtained in chapter one. As for the average temperature, the estimated coefficients are positive and negative for the lin-lin and log-log specifications, respectively. Recall that with the nonlinear cointegration methodology, the signs of temperature's coefficients were already the same way and the ECM approach provided evidence for no statistical significance of this variable.

The results for the CO2transp regressions including crude oil price and rate of motorization – models (2.9) and (2.10) – give mixed and doubtful results. For the model in levels, the sign of the estimated coefficient of crude oil price is positive which is highly unlikely. Contrary to the findings in chapter one, now the estimated coefficient of the rate of motorization is positive which is more realistic when compared to the negative coefficients found in the previous chapter for both quadratic and cubic specifications. Regarding the logarithmic specification, the estimated coefficients of both crude oil price and rate of motorization are negative. Comparing these results with those under nonlinear cointegration, it is possible to see that with cointegration with unknown structural breaks we obtain the correct (negative) sign for crude oil price, for one or two breaks regressions, meaning that a 1% increase in this variable lead to a decrease in carbon emissions from transport sector. Like in the nonlinear cointegration analysis, the negative elasticities of the rate of motorization do not seem to fit reality.

In models (2.11) and (2.12), the average fuel price substitutes the crude oil price. Regarding model (2.11), in levels, with two structural breaks, contrary to the results found using nonlinear cointegration techniques, the average fuel price has a positive coefficient which is unexpected because it is more realistic that an increase in fuels price cause demand to stretch as people tend to change their mobility behaviour such as reducing the use of private car, which contributes to CO_2 emissions reduction. Looking at the natural logarithmic specification, the choice of one or two breaks gives distinct outcomes for the control variables. The one break regression indicates a positive elasticity for fuel prices and negative for the rate of motorization. Because of the arguments mentioned above and in chapter 1, these results do not conform to reality. For the two breaks regression, the elasticity associated to fuel prices is negative which is rather plausible and contradicts the results obtained in chapter 1 for the quadratic and cubic logarithmic specifications.

It is important to recall that, at a 5% significance level, the ECM indicates that, besides GDP, the average fuel price is the only regressor that explains CO_2 emissions from transport sector.

Regarding the CO2total, for the two breaks regressions including crude oil prices – models (2.13) and (2.14) – or average fuel prices – models (2.15) and (2.16) – the estimated coefficients have similar signs for each additional variable. Thus, both crude oil price and average fuel price have a positive coefficient when considering the model in levels and a positive elasticity when taking into consideration the logarithmic specification. These results are the opposite of those from nonlinear cointegration regressions, and according to what has been explained earlier and in chapter 1, are not to be expected.

Temperature has also a negative estimated coefficient for all models, as expected. As for precipitation, the values of the estimated coefficients are always negative, as expected, and considerably higher than those of chapter 1.

It is worth to mention that through the results of ECM conducted in chapter one, apart from GDP and at a 5% significance level, crude oil price and average fuel price are statistically significant and precipitation appears to be significant at a 10% significance level.

Finally, it is worth underlining that even when the independent variables have the same plausible sign when using both methodologies[16], the magnitude of the coefficients does not necessarily follow a pattern. In some cases, the results of this chapter are higher (e.g. precipitation for CO2total regressions) and in other the coefficients are lower (e.g. crude oil price for CO2elec). These differences can be explained by the fact that in chapter 1 we use nonlinear models and in the present chapter the models are linear with structural breaks.

2.4. Discussion

The economic interpretation is critical to validate the econometric results, namely, it is relevant to assess if the estimated break dates can match the Portuguese reality along the years or if they are merely statistical outcomes.

There are two general observations concerning the results of the present study. First, the choice of the specification in levels or in natural logarithms may influence the cointegrating relationship as we can see by the different break dates when using the same data. This issue, which has been addressed in

[16] Nonlinear cointegration (chapter 1) and cointegration with unknown structural breaks.

more detail in chapter 1, is of particular importance when these results are used to design environmental policies. However, as we are going to see, the dates do not differ considerably, they are even coherent, yet the specification decision may have effects on the results, thus leading to confusing conclusions. The results therefore require careful analysis. The second general observation relates to the lagged behaviour. As Esteve and Tamarit (2012) stress, the main focus of the cointegration analysis with unknown multiple structural breaks is on the development of the relationship between the per capita CO_2 emissions and per capita real GDP, not on the evolution of each variable individually.

To understand how the regime shifts affect the long-run relationship between CO_2 emissions and GDP, we have to consider the main changes in the Portuguese economy. During the time span of our sample, it is possible to identify the scale, composition and technological effects that underlie the EKC hypothesis, although these effects are not sufficient conditions for an inverted U shape to be verified because it also depends on the magnitude of each effect. Still, Portugal did move from an agrarian economic structure to a more industrial and then, to service and knowledge-based economy.

Over the years the country has passed through important changes of which it is worth highlighting the entrance into the European Free Trade Area in 1960, which consequently led to a higher openness of the economy, thus promoting industrialization and increased exports; the implementation of democracy in 1974; and, in 1986, the accession to the European Economic Community (EEC) with the ensuing Community Support Frameworks (CSF).

Banco de Portugal (BdP, 2009) identifies, from 1986 to 2008, two business cycles for Portugal: one between 1986 and 1997 and another from 1998 to 2008. By the early nineties, the country witnessed sharp economic growth and a high speed of convergence to the EU average in per capita income. In the second half of the nineties, however, structural problems arose, particularly concerning productivity growth and fiscal and external imbalances. From 1998 onwards, and although Portugal became a member of the European Monetary Union (EMU), the Portuguese economy stagnated (Pereira and Lains, 2010).

Aguiar-Conraria *et al.* (2010) also mention similar transformations in Portuguese economic growth. These authors reveal that the period of major convergence between Portugal and the EU's wealthier countries occurred from 1986 to 1991. Since 1992 economic growth started to slow down, reflecting the necessary adjustments made by Portugal in order to meet the required conditions for entry to the euro zone. In 1993 the annual average growth rate of GDP was even negative, largely due to the international economic environment, while in 1994 the economy recovered to positive values. Overall, from 1986 to 1995, Portugal had the second highest growth rate of cohesion countries, using per capita GDP as indicator (Lima, 2000).

As for the electricity generation sector, one regime shift at 1999 is the more reasonable choice for the specification in levels. The change of the long-run equilibrium between CO2elec and GDP can be caused by the average rainfall for the year 1998 (the fourth year of lower precipitation of our sample), implying that to meet the electricity demand Portugal had to make use of fossil fuel power plants. This year also coincided with the reversal of the economic cycle. Additionally, it is possible that the estimated date might capture the effect of the implementation of the Energy Programme on electricity generated from renewable sources (RES-E), other than large hydro, created by decree-Law 195/94 of 19 July and Council of Ministers Resolution n. 68/94, supported by CSF II (1994-1999).

For the logarithmic specification, however, it is not plausible that the estimated date at 1995 can be explained by the foregoing reasons. This date is somewhat inconclusive, because, hypothetically, it could be explained by two contrasting motives. One motive relates with the sharp economic growth due to the EEC accession in 1986, the other motive concerns the slowing down of the economy in 1992 or even the negative growth rate of GDP in 1993.

Focusing on the transport sector, as mentioned in previous section concerning the specification in levels, the existence of structural breaks is highly questionable. Notwithstanding, because the results are ambiguous, we still analyse the possible causes for the two-break results. Despite the different number of regime shifts depending on the regression specification, in levels or logarithms, the estimated dates are coincident. We believe that the main explanation for the regime shifts is the same for both specifications: the CSF I (1986-1993) and CSF II (1994-1999).

Given the peripheral location of Portugal, improved accessibility has always been considered to be strategic to tackle internal accessibility shortcomings, improve international connectivity, and achieve economic development (EU, 2000). So, the basis of the CSF I and II was the modernization and expansion of transport infrastructure so Portugal could converge with its European counterparts.

Public investment under the both CSFs was mainly channelled to road infrastructures. National roads and highways were the greatest beneficiaries: their share in total public investment passed from 27.4% and 10.1%, respectively, in the 1980s, to 33.7% and 16.5% in the 1990s. The CSFs also promoted investment in railways, but at a much lower scale: from 1989 to 1998 this increased to 19.3% of total public investment after having accounted for about 17% until 1988. Not all transport infrastructures benefited from EU support. In fact, during the period covered by the CSF I and II, the share of public investment on municipal roads, ports and airports declined (all data from Pereira and Andraz, 2001).

The estimated regime shifts at 1993 and 1999 for the specification in levels, and at 1998 for the specification in logarithms matches with the investment actions under the CSFs and the economic growth and convergence of the nineties. The combination of these issues might be determinant to the shift in the long-run equilibrium of CO2transp-GDP.

In the matter of the long-run relationship between total CO_2 emissions and GDP for Portugal as a whole, both specifications give the same estimated break dates at 1991 and 1999. Since power generation and transport are the major sources of Portuguese carbon emissions, the regime shifts for CO2total-GDP should incorporate the sectoral shifts. The regime shift in 1999 is consensual for the two sectors and the underlying reasons were mentioned above. As for the first regime shift at 1991, this could be tied to the 1993 regime shift for the transport sector, as the dates are close enough to be considered within the same time interval. As we previously assumed one estimated break date for the electricity generation sector, we can say that, probably, the bulk of the first regime change in CO2total-GDP is due to the investment on transport network infrastructures.

The estimated break dates do not match completely, not only because total CO_2 emissions encompass more sectors other than electricity and transport, but also due to limitations of the econometric procedures such as the limited number of observations or the proximity of break dates. The latter occurs for the estimated break dates with three regime shifts for the specification in levels for the electricity sector and the logarithmic specification for the transport sector. Moreover, the mixed results concerning the stability tests, SEQ tests and cointegration tests suggest weak evidence in favour of breaks for the three cases under analysis. Our findings are even more questionable when compared to Liddle and Messinis' (2014) empirical work for the Portuguese case considering the total CO_2 emissions. Despite the similar methodological approach and an identical time span (1950-2010), using natural logarithms, they estimate a break date at 1988 which is quite different from our results. Notwithstanding, the authors argue that it is rather possible that Portugal, between 1950-2010, does not present a regime shift. Therefore, although we find economic events that can frame our regime shifts, we have mixed econometric results that not allow us to draw a solid conclusion about the existence of changes in the equilibrium relationship between carbon emissions and income.

2.5. Concluding remarks

The widespread criticisms about the econometric issues of the EKC studies (e.g. Stern *et al.*, 1996), the mixed empirical results, and the lack of sound theoretical foundations, have led researchers to seek alternatives that better describe the relationship between pollution and economic growth.

In this framework, the present study is the culmination of the research started in chapter 1, and aims to examine the CO_2 emissions-GDP nexus without recurring to the standard quadratic or cubic reduced-form of the EKC. Instead, we employ a linear reduced-form, in both levels and natural logarithms, and a cointegration analysis allowing for unknown multiple structural breaks in time.

Our work can be seen as pioneer because is the first study that allows for endogenous structural breaks generated by the data at a sectoral level. We go one step further than Esteve and Tamarit (2012) and Liddle and Messinis (2014), because besides the country-specific analysis, we also decompose the structural breaks that may have effects on a particular sector. To do so, we use the same data as in chapter 1.

This research permit us to compare not only the existence and number of structural breaks for the cointegrating relationship CO_2-GDP for both electricity generation and transport, but also to assess whether structural breaks occur on the same dates. This is of major relevance because it allows us to discuss whether a structural change affects the two main carbon emitters in the same way. Detecting sectoral regime shifts is a pertinent matter, as it helps improve policy formulation for climate change mitigation.

Despite the evident structural changes that Portugal has gone through, our results show weak evidence of regime shifts, particularly for the transport sector. Notwithstanding, the results indicate that the cointegrating relationship for electricity generation sector, transport sector and Portugal as a whole has changed over time. Based on the estimated dates we claim that, in different manners, the Community Support Frameworks I and II were the major force behind structural changes in both sectors as well as nationwide overall emissions. In the future it will be interesting to repeat this empirical study for a longer period of time, since the reliability of the results depends on the time length of the sample, with larger samples increasing the likelihood of break identifications. The importance of time span is evident in the findings for Spain: Esteve and Tamarit (2012) use 151 year observations whereas Liddle and Messinis (2014) use 61, and differences in the results are clear. Esteve and Tamarit (2012) found three regime shifts while Liddle and Messinis (2014) detected two, at different dates. Moreover, Liddle and Messinis (2014) conjecture, based on the income-elasticity values, that in the period under study there is the possibility that Spain hadn't experienced a structural break at all, and mention the importance of a longer time span.

For now, the overall conclusion that can be drawn is that CO_2 mitigation at national level requires the implementation of measures targeted to each sector. After the empirical work presented in chapters 1 and 2, no certainty regarding the relationship between carbon emissions and economic growth at both sectoral and national level has been achieved, yet new questions were identified.

3. RES-E capacity expansion decisions in a duopolistic electricity market[17]

3.1. Introduction

The economic and social costs of climate change, along with security of energy supply and fossil fuel price fluctuations and exhaustibility, are strong reasons to foster electricity generation from renewable energy sources (RES-E) to meet electricity demand. Electricity demand can be restrained through the adoption of energy-efficiency improvements in buildings, industry and transport. Nonetheless, despite the efforts directed at improving energy-efficiency, electricity demand is likely to continue to rise in the coming years. The International Energy Agency (IEAb, 2014) estimates an increase of about 80% in world electricity demand between 2012 and 2040. More than half of the growth in global primary energy use is due to electricity generation. Furthermore, it is expected that over the same period, the share of fossil fuels for electricity generation will drop from 68% to 55%, and the use of renewables will increase from 21% to 33% (IEAb, 2014). Still fossil fuels will remain predominant (IEAb, 2014).

As the expansion of RES-E is one of the top priorities of the long-term energy policy of many economies, it is imperative to ensure that investment in new RES-E generation capacity is made. This investment will be key in the pursuit of decarbonisation of the electricity supply mix while guaranteeing that supply covers future electricity. Simultaneously, there are many countries around the world where electricity generation markets have been deregulated from monopolies, or near monopolies, towards increased competition. Nevertheless, there are some economic barriers that limit the implementation of a perfectly competitive market structure such as economies of scale, absolute cost advantage of the established firms, high capital requirements and funding constraints (Soares *et al.*, 2012). Therefore, instead of the perfect competition hypothesis, an oligopolistic market is a more realistic description of the electricity market (Murphy and Smeers, 2005).

Oligopoly is a market structure where there are few firms producing the total supply of a given good, and thus there is strategic interdependence between firms. The Cournot model is one the most popular oligopoly models, wherein firms decide simultaneously the output levels, and it is the preferred market structure to represent deregulated electricity markets worldwide (Reichenbach and Requate, 2012), since its assumption of quantity competition is adequate to model decisions in electricity markets, where price competition does not usually occur.

The Cournot model can be explained using the simplest model of duopoly competition. Both firm 1 and firm 2 adjust their output level in order to maximize profit given market demand and its competitor' decisions (Tirole, 1988). If firm 1 expects firm 2 to produce the output q_2, then firm 1 will choose its profit maximizing output given by $q_1^*(q_2)$, which is denominated the reaction function, or best-response

[17] A previous version of this work was presented at the Global Conference on Environmental Taxation (Madrid, 2011) and at the Spanish-Portuguese Association of Resource and Environmental Economics (Faro, 2012).

function, of firm 1. Firm 2 behaves in an identical way, and therefore firm 2's reaction function is $q_2^*(q_1)$. Each reaction function corresponds to a firm's best response to each possible choice of another firm. The Cournot-Nash equilibrium is reached where the two reaction curves intersect, and, since at this point none of the firms wishes to change its decisions, market demand balances market supply.

In a dynamic setting, however, generation capacity is not fixed and market equilibrium will also depend on firms' investment decisions. In the context of deregulated electricity markets, generation planning is undertaken by private investors whose goal is long-term profit maximization. Thus, it is of key importance to understand how the strategic behaviour of the producers influences the investment levels. The electricity market has been extensively modeled in the literature, as described in Ventosa *et al.* (2005). Closer to our work are papers that analyse investment incentives on electricity. Murphy and Smeers (2005) compare investment decisions of one and two stage games under uncertainty. They propose a two-period model of generation capacity investment in competitive electricity markets to evaluate how imperfectly competitive markets affects generation investment capacities. To do so they build three models. The first model represents perfect competition. The second model is an open-loop[18] Cournot game in which capacities are built and sold in long-term contracts, at the same time. The third model is a closed-loop Cournot game with investment decisions being made in the first stage while spot market operations take place in the second stage. The findings show that a more complex electricity market structure increases the difficulty of the investment decision process. Moreover, both investments and outputs in the closed loop game fall between the open loop game and the efficient solutions.

Bushnell and Ishii (2007) develop a Cournot model to assess capacity expansion decisions on restructured electricity markets. To this study, firms' investment decisions are a consequence of their market position, their obligations in the market, and uncertainty which can delay the investment. Garcia and Schen (2010), through a dynamic Cournot model with stochastic demand growth, conclude that, compared with the social optimum, Cournot firms underinvest. Genc and Thille (2011) investigate how hydro and thermal electricity generators compete under demand uncertainty. Through a two-period Cournot game they analyse thermal producer's incentives for capacity investments, both for S-adapted open-loop equilibrium and Markov perfect equilibrium. The conclusion is that the equilibrium for investment can be higher or lower than the efficient level.

Léautier (2013) evaluates the reasons of underinvestment in electric power generation, and also the many corrective market designs that have been suggested and executed. From this work stem four main findings. First, market power seems to be a more relevant cause of underinvestment when compared to the imposition of a price cap. Second, physical capacity certificates markets adopted by the United States restore optimal investment levels, however producers' profits rise more than on imperfect competition context. The third conclusion is that financial reliability options are unable to promote

[18] The terms "open-loop" and "closed-loop" denote distinct information structures for multiple-period games. In the open-loop model, all decisions are taken simultaneously by all players, thus they cannot observe the play of the others. On the contrary, in the closed-loop model, agents are aware of all past plays before deciding their actions at the beginning of each period (Fudenberg and Levine, 1988).

investment. Fourth, a single market for energy and operating reserves subject to a price cap is isomorphic to a simple energy market. Léautier (2013) argues that together these findings indicate that, in order to guarantee electric power generation adequacy, the main focus of policy makers should be to control a reduce the power market. Wogrin *et al.* (2013) is another example of investment models that account for imperfect competition. The authors develop two game-theoretic models of the capacity expansion in liberalized markets. The first model consists of an open-loop equilibrium model, and the second one is a closed-loop model. In both models, a conjectural variation of the price response function is assumed to define the strength for competition among producers in the spot market. When considering one load period, for any conjectural variations ranging from perfect competition to Cournot, the closed-loop equilibrium yields the same as for Cournot open-loop equilibrium. The results for multiple load periods reveal that for different conjectural variations, the closed-loop equilibria may differ from each other as well as from open-loop equilibria. Furthermore, alternative conjectured price response models with switching conjectures are examined. This analysis suggests that the rank ordering of closed loop equilibria with regard to consumer surplus and total social welfare is ambiguous. Therefore, regulatory frameworks that drive marginal cost-based bidding in spot markets might reduce market efficiency and consumer surplus by diminishing incentives for investment.

Di Cosmo and Valeri (2014) examine the impacts of increasing wind power generation on the incentives to invest in thermal plants and on the wholesale cost of electricity, in the context of the Irish Single Market. In deregulated markets, is up to private investors the decision to invest in a specific power generation plant with a certain installed capacity. Because this decision hinges on expected profits, the authors address how such profits are influenced by the steadily growing wind power generation. Revenue and cost streams of three types of thermal generation plants – coal plant, combined cycle gas turbine plant and an open cycle gas turbine plant – are compared to observe changes in plants' expected profits with wind generation capacity growth. The major finding of this work is the evidence of a negative correlation between electricity shadow prices and wind installed capacity, while an extension from the present 2000MW of wind installed capacity to around 3000MW has distinct impacts considering the type of plant. The reduction in profits is more expressive for baseload gas plants, and minor for less flexible coal-fuelled plants. In the scenario without wind power generation, the power plants' expected profits are higher. Filomena *et al.* (2014) focus on the process of technology selection and capacity investment for electricity generation in a competitive market under uncertainty concerning marginal costs. They use both open and closed loop Cournot models, in order to assess the link between various technologies' cost structures and the portfolio of generation technologies adopted by each firm in the equilibrium. The authors find that even with risk-neutral firms and technologies with distinct cost presumptions, diverse portfolio of technology emerges.

Another strand of literature relevant for this paper concerns RES-E diffusion in the presence of support mechanisms such as feed-in tariffs (FIT). For instance, Toke *et al.* (2008) investigate the main factors influencing wind power deployment, comparing six different countries. One of the variables examined was financial support mechanisms, namely the existence of FIT. They conclude that wind power capacity was greatly expanded in the countries under a feed-in regime, such as Germany or Spain.

Similar results were obtained by Mulder (2008), who studied how governments in the EU (15) countries have succeeded in stimulating investments in wind turbines between 1985 and 2005. The author emphasizes economic attractiveness as being a necessary condition for wind power investment increase, highlighting the role of feed-in tariffs in this process and again underlining Germany and Spain as success cases.

These two previous examples based their conclusions on installed capacity to assess the effectiveness of renewable support mechanisms. Other authors adopt a different approach by focusing directly on investment. Bürer and Wüstenhagen (2009) perform a survey of investment professionals in venture capital and private equity funds, in order to assess what energy policies encouraged them to invest in clean technologies. The results corroborate previous findings of feed-in tariff effectiveness. More recently, Wüstenhagen and Menichetti (2011) suggest the use of variables like risk-return perceptions and portfolio effects to explain renewable energy investment options.

This chapter builds on the two-period Cournot game formalized in Genc and Thille (2011) to examine the strategic behaviour of both thermal and RES-E producers when the latter faces generation capacity investment decisions in deregulated markets with and without the adoption of feed-in tariffs. Besides generation costs, our model takes into account the environmental damages caused by the thermal power plant. Finally, we compute the socially optimum solutions in order to assess the relative importance of both market failures (market power and pollution externality). The final section of the chapter provides some reflections on the potential impact of electric vehicles on the electricity market.

3.2. The model

As in Genc and Thille (2011), we use a two-period Cournot duopoly game for the electricity generation market with asymmetric technologies. In our case, we design a model with thermal and renewable (it could be hydro, solar or other) generating plants for two scenarios, acknowledging their different contributions to emissions damages. The difference between the two scenarios is the adoption of feed-in tariffs in the second period. A basic feed-in[19] support scheme is where a guaranteed fixed price is paid to RES-E producers per unit of electricity fed into the electricity grid (Klein *et al.*, 2008).

Both producers are assumed to maximize their profit by setting quantities. Additionally, the RES-E producer has to decide in period one how much to invest in generation capacity to be available in the following period, with and without a feed-in support scheme, given its capacity constraints. We use the same linear inverse demand function, quadratic thermal generation cost function, and quadratic investment capacity cost function as Genc and Thille (2011). Contrary to the latter, which supposed that electricity demand could increase or decrease, we assume that there is an increase of electricity demand

[19] Usually feed-in tariffs are differentiated by technology, installed capacity or location (for details about feed-in design see Klein *et al.*, 2008).

from period one to period two. This assumption is justified by the expected evolution in actual markets, as noted in the Introduction.

The inverse demand function is given by $P_t(Q_t) = D_t - Q_t$, with $t = 0, 1$. D_t is a constant and Q_t is the total electricity generated. As for generation technologies, the cost function for the thermal producer is $C(q_t) = c_\alpha q_t + \frac{c_\beta}{2} q_t^2$, where the parameters c_α and c_β are positive, and RES-E plants have zero marginal costs, since renewable sources are free so variable operation and maintenance costs are negligible (see EIA, 2014).

For both periods thermal and RES-E producers have to choose their output levels, denoted as $(q_0\, , q_1)$ and (z_0, z_1), respectively. For simplicity thermal capacity is assumed to be not binding, because even if it were binding, it could not be easily scaled. Therefore, the thermal producer will only select how much quantity to produce in each period's Cournot-Nash market equilibrium. For RES-E production, however, the capacity constraint is binding, thus, the producer has to choose not only the output levels but also make a decision on the investment in new capacity – I_0. The investment is irreversible, it becomes available in the following period, and old capacity doesn't depreciate. The investment capacity costs are modeled as a quadratic function of investment given by $e\frac{I_0^2}{2}$.

Our contributions relative to the Genc and Thille (2011) model are two-fold: we consider the impact of possible feed-in implementation for RES-E producers, given the popularity of such schemes in many countries; and we explicitly take into account the external costs of GHG emissions produced by the thermal generator, as this externality is crucial for the comparison between market outcomes and the social optimum.

3.3. Results and discussion

If both thermal and RES-E producers had unconstrained generation capacities, then there would be no need to invest in capacity expansion from the first to the second period. Because it is assumed that only thermal generating plants have production costs and electricity is a homogeneous commodity, the Cournot-Nash solution would in this case be a larger share of RES-E when compared to thermal. This equilibrium would be unstable because if there were no capacity constraints the zero-cost producer could engage in price competition to drive out thermal production. However, such a possibility is merely a theoretical exercise. The RES-E firm cannot drive its rival out of the market because only a share of RES-E capacity can be viewed as securely available capacity (due to intermittency and unpredictability). Therefore, conventional dispatchable generators, such as thermal plants, are needed as backup capacity in order to ensure electricity supply security and balance the electric system (Fürsch *et al.*, 2010). Therefore we only present here the model where both sources co-exist in the market but RES-E generation has a binding capacity constraint.

3.3.1. Constrained RES-E generation

We begin by focusing on the thermal producer's profit-maximization problem, when taking the quantity of the RES-E producer as given. Suppose that this producer is always below full capacity, so it does not have to make a decision on investment. Thus the Cournot model is as follows, for $t = 0{,}1$:

$$max_{q_t}\,(D_t - z_t - q_t)q_t - \left(c_\alpha q_t + \frac{c_\beta}{2}q_t^2\right) \tag{3.1}$$

The profit-maximization first-order condition is $\frac{\partial \pi}{\partial q_t} = 0$, and the reaction function is:

$$q_t = \frac{D_t - z_t - c_\alpha}{2 + c_\beta} \tag{3.2}$$

Equation (3.2.) gives the expected results for a typical Cournot solution. The thermal quantity will be smaller the lower demand, the higher its production costs and the higher the RES-E output.

Assuming that the RES capacity constraint is binding, it is possible for the RES-E producer to decide to invest in extra capacity even without a support policy. Note that the decisions for z_0 and I_0 can be modeled independently, as capacity will only become available in the following period and we assume there are no financing constraints. To analyse the investment decision, we assume that the RES-E producer will produce at capacity in the initial period. Since there is no uncertainty and investment is costly, moreover, it is clear that it will only be done to the extent that the resulting output will be sold in the second period, so that no idle capacity will exist in that period, either. Therefore the crucial decision for this player is investment, while quantities are given by the capacity constraints: for $t = 0$, $z_0 = k_0$, and for $t = 1$, $z_1 = k_0 + I_0$, where k_0 is the initial fixed capacity of the renewable generating plant.

The RES-E producer will choose how much to expand capacity by solving an intertemporal profit-maximization problem:

$$max_{I_0} - \frac{e}{2}I_0^2 + \frac{1}{1+r}(D_1 - z_1 - q_1)z_1 \tag{3.3}$$

$$s.t.\ I_0 \geq 0$$

$$z_1 = k_0 + I_0$$

Where r is the discount rate.

The problem is to maximize the following Lagrangean[20]:

$$L = -\frac{e}{2} I_0^2 + \frac{1}{1+r} (D_1 - k_0 - I_0 - q_1)(k_0 + I_0) \tag{3.4}$$

$$\frac{\partial L}{\partial I_0} = -eI_0 + \frac{1}{1+r} [-2(k_0 + I_0) + D_1 - q_1] - \lambda_0 \leq 0; \qquad I_0 \geq 0;\ I_0 \frac{\partial L}{\partial I_0} = 0 \tag{3.5}$$

The Cournot-Nash equilibrium investment is calculated by substituting q_1 in equation (3.5) by the reaction function (3.2.) for $t = 1$:

$$q_1 = \frac{D_1 - k_0 - I_0 - c_\alpha}{2 + c_\beta} \tag{3.6}$$

The investment function is given by:

$$I_0^c = \begin{cases} I_0 = 0 \quad if \quad D_1 < \frac{(3+2c_\beta)k_0 - c_\alpha}{(1+c_\beta)} \\ I_0 = \frac{D_1(1+c_\beta) - k_0(3+2c_\beta) + c_\alpha}{3+2c_\beta + (2+c_\beta)e(1+r)} \quad if\ otherwise \end{cases} \tag{3.7}$$

The interior solution in (3.7) is characterized by a level of chosen investment that depends positively on next-period demand (D_1) and negatively on initial capacity (k_0), marginal investment costs (e) and interest rate (r). Using this solution for investment allows us to rewrite RES-E output for $t = 1$, $z_1^c = k_0 + I_0^c$ as follows:

$$z_1^c = \frac{D_1(1+c_\beta) + k_0(2+c_\beta)e(1+r) + c_\alpha}{3+2c_\beta + (2+c_\beta)e(1+r)} \tag{3.8}$$

Substituting (3.8) in the reaction function (3.2.), we calculate the Cournot-Nash solution for thermal generator at $t = 1$:

$$q_1^c = \frac{D_1 + (D_1 - k_0 - c_\alpha)e(1+r) - 2c_\alpha}{3+2c_\beta + (2+c_\beta)e(1+r)} \tag{3.9}$$

Knowing that $Q_0^c = q_0^c + z_0$, and $z_0 = k_0$, allows us to determine total quantity equilibrium for $t = 0$:

$$Q_0^c = \frac{D_0 + k_0(1+c_\beta) - c_\alpha}{2+c_\beta} \tag{3.10}$$

Replacing the Q_0^c in the inverse demand function we obtain the equilibrium price:

[20] Since our functions are linear and quadratic the first order necessary conditions are also sufficient for global optimality.

$$P_0^c = \frac{(D_0 - k_0)(1 + c_\beta) + c_\alpha}{2 + c_\beta} \tag{3.11}$$

For $t = 1$, we determine Q_1^c through the sum of $q_1^c + z_1$, where $z_1 = k_0 + I_0$, thus for the interior solution of I_0, yields:

$$Q_1^c = \frac{D_1(2 + c_\beta) + [D_1 + k_0(1 + c_\beta) - c_\alpha]e(1 + r) - c_\alpha}{3 + 2c_\beta + (2 + c_\beta)e(1 + r)} \tag{3.12}$$

$$P_1^c = \frac{D_1(1 + c_\beta) + [(D_1 - k_0)(1 + c_\beta) + c_\alpha]e(1 + r) + c_\alpha}{3 + 2c_\beta + (2 + c_\beta)e(1 + r)} \tag{3.13}$$

Whether Q_1^c is larger or smaller then Q_0^c depends on the increment in electricity demand relative to the investment level. If $I_0 > D_1 - D_0$ then $q_1^c < q_0^c$, because the expansion in capacity generation will meet not only the increase of electricity demand but also eat into the market share of thermal generation. Then, $Q_1^c < Q_0^c$ and as $D_1 > D_0$, through the inverse demand function we know that $P_1^c > P_0^c$. Conversely, $I_0 < D_1 - D_0$ implies that $q_0^c < q_1^c$; however, in this case it is not possible to reach a conclusion regarding Q_0^c and Q_1^c and, consequently, P_0^c and P_1^c.

We assume that RES-E is GHG-free, hence thermal production results are used as an indicator of emissions from electricity production. GHG emissions will be higher in the second period when $I_0 < D_1 - D_0$ because the increase in demand will be met by both producers; conversely, emissions will decrease if $I_0 > D_1 - D_0$. If the demand in $t = 1$ was equal or less than in $t = 0$, on the other hand, the increase of RES-E capacity would always induce a reduction in the thermal producer's operation and therefore a drop in emissions.

In order to examine the isolated impact of investment in GHG emissions levels if there was no change in demand, we determine the equilibrium quantity under the assumption that $D_1 = D_0$. By substituting the D_1 for D_0 and $z_1 = k_0 + I_0$ in equation (3.2.) for $t = 1$, we find that for any $I_0 > 0$, because the demand remains the same in both levels, $q_1^c < q_0^c$, resulting in GHG emission savings from $t = 0$ to $t = 1$. Moreover, we can infer that $Q_1^c > Q_0^c$ and $P_1^c < P_0^c$.

3.3.2. Constrained RES-E generation and feed-in tariff support scheme

Now we assume that RES technologies are less mature than thermal ones, and thus, in order to promote RES-E massification, feed-in tariffs are applied in period 1. To have an effect, the feed-in tariff must be higher than the market price that would occur in the market equilibrium in $t = 1$, and given that there are no production costs for the renewable generating plant and investment is costly, RES-E producer will always operate at full capacity. For $t = 0$ the Cournot-Nash solutions are the same as section 1.3.1., whereas the intertemporal profit-maximization problem associated with the investment decision becomes:

$$max_{I_0} -\frac{e}{2}I_0^2 + \frac{1}{1+r} f z_1 \tag{3.14}$$

$$s.t.\ I_0 \geq 0$$

$$z_1 = k_0 + I_0$$

Where f denotes the feed-in tariff.

Thus the chosen level of investment will simply be given by:

$$I_0^c = \frac{f}{(1+r)e} \tag{3.15}$$

That means investment will increase with the feed-in tariff and decrease with the interest rate and with investment cost, as expected. Moreover, note that the value for investment will now be independent of all electricity market parameters (such as demand or cost for the other producer). Since the feed-in tariff should be larger than the electricity market price given by equation (3.13), investment in this case is higher than the investment level obtained in the absence of feed-in tariff, as expected.

Assuming the thermal producer acts according to its best-response function, Cournot-Nash equilibrium quantities are:

$$z_1^c = k_0 + \frac{f}{(1+r)e} \tag{3.16}$$

$$q_1^c = \frac{(D_1 - k_0 - c_\alpha)e(1+r) - f}{(2+c_\beta)e(1+r)} \tag{3.17}$$

For $f > P_1^c$, investment is higher when feed-in tariffs are introduced in $t = 1$. The higher the feed-in tariff, the more capacity is built and the lower the quantity that is available to the thermal producer in the second period. GHG will be lower in the second period as long as $I_0 > D_1 - D_0$, as before, because in this situation $q_1^c < q_0^c$, and GHG emissions are a by-product of thermal electricity generation.

For this equilibrium total quantity is given by:

$$Q_1^c = \frac{[D_1 + k_0(1+c_\beta) - c_\alpha]e(1+r) + f(1+c_\beta)}{(2+c_\beta)e(1+r)} \tag{3.18}$$

and equilibrium price is:

$$P_1^c = \frac{[(D_1 - k_0)(1+c_\beta) + c_\alpha](1+r)e - f(1+c_\beta)}{(2+c_\beta)e(1+r)} \tag{3.19}$$

From our results we can draw some conclusions about the feed-in tariff's effect on market equilibrium. When comparing market total quantities for $t = 0$ and $t = 1$, and given the equilibrium expressions, it is possible to see that the total quantity is higher in the second period when compared to the first one.

Total quantity in the second period rises relative to the first period due to both the additional investment made by the RES-E producer and the increase in electricity demand. Conversely, equilibrium price decreases. A possible explanation for this is that the thermal electricity producer generates electricity depending on demand level. So the greater the share of total demand supplied by the RES-E producer, the lower the residual demand faced by thermal electricity producer, and the market price goes down.

Moreover, if $f > \frac{D_1(2+c_\beta) - c_\alpha}{1+c_\beta}$ holds, then the total quantity under feed-in support is also higher than in the simple Cournot model for $t = 1$. It is expected that this is the case in which RES-E producer tend to invest more. This can be explained by two factors: (1) feed-in tariffs with mandatory purchase, and (2) the total electricity demand increase. Given the stable framework provided by RES-E support mechanism, the producer will invest more at $t = 0$ in order to have more installed capacity at $t = 1$.

3.4. Social optimum

Even without environmental externalities, the market equilibrium described in the previous sections would not be Pareto-efficient since firms have some market power. The Cournot solution generally results in lower output than would occur under a perfectly competitive market, therefore it is inefficient. However, the generation of GHG emissions from fossil fuels is a key aspect in the evaluation of electricity markets and it is therefore imperative that such external effects be taken into account in the calculation of the social optimum.

In this section we compare social welfare at the Cournot-Nash equilibrium and at the optimum, with respect to q_0, q_1, z_0, z_1 and I_0. In particular, we maximize the two-period expected welfare subject to non-negativity and generation capacity constraints, but contrary to Genc and Thille (2011) we take into consideration the external damages borne by society because of thermal generation, by adopting a convex and strictly positive damage cost function, defined as $\frac{s}{2}q_t^2$ as in Nordhaus (1994) and Reichenbach and Requate, (2012), where s is a non-negative parameter.

The social planner formulates the welfare maximization problem adding consumer and producer surplus in both periods, as follows[21]:

$$max_{q_0,z_0 I_0,q_1,z_1} \int_0^{Q_0} P(Q_0)dQ_0 - c_\alpha q_0 - \frac{c_\beta}{2} q_0^2 - \frac{s}{2} q_0^2 - \frac{e}{2} I_0^2 + \frac{1}{1+r} \left[\int_0^{Q_1} P(Q_1)dQ_1 - c_\alpha q_1 - \frac{c_\beta}{2} q_1^2 - \frac{s}{2} q_1^2\right]$$

(3.20)

$$s.t.\, q_0, z_0, I_0, q_1, z_1 \geq 0$$

$$z_0 = k_0$$

$$z_1 = k_0 + I_0$$

When RES capacity exists it's costless to use it and there are no external damages, so the capacity constraints in the welfare maximization problem are also binding.

The maximization problem can then be determined by maximizing the correspondent Lagrangian function. Assuming an interior solution:

$$L = D_0(q_0 + z_0) - \frac{(q_0+z_0)^2}{2} - \frac{s}{2} q_0^2 - \frac{e}{2} I_0^2 + \frac{1}{1+r}\left[D_1(q_1 + k_0 + I_0) - \frac{(q_1+k_0+I_0)^2}{2} - c_\alpha q_1 - \frac{c_\beta}{2} q_1^2 - \frac{s}{2} q_1^2\right]$$

(3.21)

The first order conditions are:

$$\frac{\partial L}{\partial q_0} = D_0 - q_0 - z_0 - c_\alpha - c_\beta q_0 - s q_0 = 0; \qquad (3.22)$$

$$\frac{\partial L}{\partial q_1} = \frac{D_1 - q_1 - z_1 - c_\alpha - c_\beta q_1 - s q_1}{1+r} = 0; \qquad (3.23)$$

$$\frac{\partial L}{\partial I_0} = -eI_0 + \frac{1}{1+r}(D_1 - q_1 - k_0 - I_0) = 0; \qquad (3.24)$$

To determine the socially efficient outcomes and investment we use the identical procedure used to identify the Cournot-Nash solutions. The first order conditions allow us to determine socially efficient outcomes and investment.

For $t = 0$, social welfare is maximized when:

$$z_0^* = k_0 \qquad (3.25)$$

[21] Although several authors have pointed out that social discount rates should be lower than private discount rates (Evans and Sezer, 2005), this complication would not bring significant added value in a two-period framework.

$$q_0^* = \frac{D_0 - k_0 - c_\alpha}{1 + c_\beta + s} \tag{3.26}$$

$$I_0^* = \frac{(D_1 - k_0)(c_\beta + s) + c_\alpha}{(1 + c_\beta + s)e(1+r) + c_\beta + s} \tag{3.27}$$

$$Q_0^* = \frac{D_0 - c_\alpha + k_0(c_\beta + s)}{1 + c_\beta + s} \tag{3.28}$$

$$P_0^* = \frac{(D_0 - k_0)(c_\beta + s) + c_\alpha}{1 + c_\beta + s} \tag{3.29}$$

As for $t = 1$, the optimum conditions are as follows:

$$z_1^* = \frac{k_0(1 + c_\beta + s)e(1+r) + D_1(c_\beta + s) + c_\alpha}{(1 + c_\beta + s)e(1+r) + c_\beta + s} \tag{3.30}$$

$$q_1^* = \frac{(D_1 - k_0 + c_\alpha)e(1+r) - c_\alpha}{(1 + c_\beta + s)e(1+r) + c_\beta + s} \tag{3.31}$$

$$Q_1^* = \frac{[D_1 - c_\alpha + k_0(c_\beta + s)]e(1+r) + D_1(c_\beta + s)}{(1 + c_\beta + s)e(1+r) + c_\beta + s} \tag{3.32}$$

$$P_1^* = \frac{(D_1 - k_0)e(1+r)(c_\beta + s)}{(1 + c_\beta + s)e(1+r) + c_\beta + s} \tag{3.33}$$

The above conditions can be compared to the Cournot-Nash solutions for the case of RES-E constrained capacity without feed-in tariffs. This analysis takes into consideration the fact that in imperfect markets, such as Cournot oligopoly, firms maximize their profits by producing a suboptimal quantity. Consequently, it is possible that although they ignore the external environmental costs created by their production choices, such firms might produce closer to the socially efficient level when compared to perfect competitive firms. It all depends on the magnitude of external and private cost distortions (Baumol and Oates, 1988, cap. 6). It isn't straightforward where the Cournot-Nash equilibrium will be relative to the social optimum, because the social planner considers the damage caused by thermal generation. Thus our analysis pays special attention to the value of the damage cost function parameter – s.

Let us begin with the evaluation for thermal generation quantities, and market quantity and price, for the first period. Because $z_0 = k_0$, for both Cournot-Nash equilibrium and social optimum, variations in the total quantity and price only depend on q_0^c and q_0^*. Therefore through the comparison between equation (3.2) for $t = 0$ and equation (3.26) we are able to reach some conclusions depending on the values attributed to the external damage parameter s. Knowing which thermal output is higher for $t = 0$, and using the inverse demand function, allows us to conclude whether the market quantity and price is higher or smaller at the optimum or at Cournot-Nash solutions.

$$s > 1 \begin{cases} q_0^* < q_0^c \\ Q_0^* < Q_0^c \\ P_0^* > P_0^c \end{cases} \tag{3.34}$$

$$s = 1 \begin{cases} q_0^* = q_0^c \\ Q_0^* = Q_0^c \\ P_0^* = P_0^c \end{cases} \tag{3.35}$$

$$1 > s \geq 0 \begin{cases} q_0^* > q_0^c \\ Q_0^* > Q_0^c \\ P_0^* < P_0^c \end{cases} \tag{3.36}$$

Higher external costs lead to smaller thermal outcomes at the optimum, as expected. They are also decisive for optimal market size, since higher damages will lead to smaller overall quantities and higher market prices (eq. 3.34). For the particular case of $s = 1$, the two externalities exactly cancel out and the Cournot-Nash market solution turns out to be Pareto-efficient (eq. 3.35). However, if the value of the marginal external costs is relatively small ($s < 1$, eq. 3.36), the market failure associated with imperfect competition dominates and corrective measures to approach the social optimum would actually lead to an increase in emissions. Thus the simple use of a Pigouvian tax on thermal producers or a subsidy on renewable power, such as that provided by a feed-in tariff, is not necessarily welfare-improving.

The complexity of our analytical solutions for both Cournot-Nash equilibrium and Social optimum limits our analysis for $t = 1$. Although the two previously mentioned effects will still exist and have countervailing effects (market power vs. external damages), it is necessary to also consider investment costs and benefits, where the latter will come from the lower production costs of RES-E. We restrict our evaluation to thermal (eq. 3.9 and 3.31) and RES-E quantities, and investment (eq. 3.7 and 3.27), under some specific conditions particularly regarding parameter s. We look at the case where the external costs are moderate or even inexistent, and at the case where $s = 1$.

Regarding the thermal producer outputs, focusing on the denominators of equations (3.9) and (3.31) we can see that for $1 \geq s \geq 0$, in the Cournot-Nash solution has the highest value, thus when $c_\alpha > D_1$ holds, $q_1^* > q_1^c$ because with this condition the numerator for q_1^* is higher than for q_1^c. The explanation is similar to the one mentioned for the first period regarding the magnitude of the external costs. As $s \to 0$, the greater is q_1^* in comparison with q_1^c.

Concerning the investment level, a similar procedure as was used for the comparison of thermal quantities allows us to draw some conclusions. $I_0^* > I_0^c$ and $z_1^* > z_1^c$ holds for $s = 1$, since $c_\beta > 0$ which is one of the assumptions of our model. The signal of the partial derivative $\frac{\partial I_0^*}{\partial s}$ does not depend on s, so we can extend the previous conclusion for $s \geq 1$. If there are no external costs ($s = 0$), then the condition $D_1 > k_0(c_\beta + 3)$ must additionally be verified for $I_0^* > I_0^c$ and $z_1^* > z_1^c$. Even for moderate and zero external costs, therefore, the inefficiency of the duopoly decisions leads to under investment and thus to a level of RES-E generation that is too low.

3.5. A discussion on the potential role of electric vehicles

In chapter 2 it was shown that power generation and transport, which are the main sources of GHG emissions in most countries, do not necessarily respond to changes in economic activity in the same way. Hitherto these two markets have evolved fairly separately, with power generation depending less and less on the oil-based products that supply transport. For instance, in 1973 oil accounted for 24.8% of fuel used in power generation in the world, while in 2012 the share was 5%. Concurrently, 45.4% of total final oil consumption in 1973 was for transport, while in 2012 this sector accounted for a much larger proportion of 63.7% (IEAb, 2014).

In the future, the spread of electric vehicles (EV) will bring about a closer linkage of the electricity and transport markets. This is particularly relevant because RES-E and EV policies share common goals. Therefore this chapter would not be complete without a description of two branches in the literature that discuss the possible interactions between EV and the electricity market, especially RES-E. The first one handles the potential environmental advantages of EV, particularly GHG emissions savings. The second one focuses on EV storage capacity, which could facilitate the integration of intermittent renewable energy in power grids and, probably, decrease market power.

3.5.1. GHG emissions

Compared with internal combustion vehicles, electric ones have more efficient engines and produce zero tailpipe emissions. Nonetheless, it is by no means certain that EV use is GHG emission free, as this will depend on how the electricity to charge batteries is generated. The determination of GHG emissions must thus be well-to-wheel, which means that the calculation includes the electricity generation process. There are several studies in the literature related with this matter.

WWF (2008) compares diesel, gasoline and electric vehicles in different countries. The results were significantly diverse, emphasizing the importance of primary energy sources for electricity production on vehicles' CO_2 emissions. For instance, in Greece power plants are mainly coal-fired, which leads to a result where EV CO_2 emissions are similar to those emitted by conventional vehicles. Austria is the opposite example, because it uses low-carbon electricity sources, so that the introduction of EV translates into an effective emission reduction of more than 70%.

Hadley and Tsvetkova (2008) examine the possible impacts of plug-in hybrid electric vehicles (PHEV) on the production of electricity in several U.S. regions. One of the variables is the GHG emissions caused by the introduction of the PHEV. The authors concluded that the adoption of these vehicles results in increased emissions because, although the electricity mix varies regionally, in most cases power plants use fossil fuels, particularly coal, as primary energy.

Göransson *et al.* (2009) analyze the effects of introducing PHEV in an electrical system where 25% of electricity is generated from wind power and the remaining 75% from thermal power plants. The conclusion was that these vehicles can lead to CO_2 savings if the integration occurs with load management. In the absence of load management, it is likely that the use of these vehicles results in increased emissions. The findings of Sioshansi and Denholm (2009) reinforce the possible role of PHEV in GHG reduction when recharge flexibility is provided. Emissions decrease is accentuated when the PHEV provide Vehicle-to-Grid (V2G) services.

3.5.2. Intermittent RES-E

The growth of renewable energy for electricity generation should intensify in the future, especially for wind and solar energy; since the costs of these technologies have declined enough to make them competitive with fossil-fuel based alternatives and thus mature for new deployment. However, despite their advantages, both wind and solar energy are intermittent, which has consequences for the balance and reliability of electric grids.

Of intermittent RES, wind is the one that requires more attention due to its unpredictability, which can cause two types of problems. First, the production of electricity from wind energy may be insufficient to meet demand at each time. For low levels of penetration, the fluctuation can be managed through existing mechanisms to balance supply and demand. But if the share of intermittent RES-E exceeds 10 to 30% of electricity supply, incremental backups are needed (Kelly and Weinberg, 1993, quoted by Kempton and Tomic, 2005). Second, production may exceed demand, as occurs during off-peak periods when wind power is higher and, at the same time, baseload electricity plants are operating. Given the inflexibility of the latter, the solutions are to export the wind-generated power, use it to pump water for hydropower storage, or just turn off the wind turbines. Solving these problems increases grid management costs, with negative consequences for consumers.

Given their technical characteristics, EV appear as an alternative solution for RES-E absorption into the system. The capacity of EV to store and inject electricity into the grid is called Vehicle-to-Grid (V2G) technology. V2G allows the regulation of grid stability in real time, making use of parked EV. This technology has a fast response with a relatively low capital cost (compared with other solutions), yet it entails a large initial network investment to implement the transition to smart grids.

The work of Kempton and Tomic (2005) was one of the first to suggest that V2G technology can play an important role for intermittent RES-E storage, accommodating solutions where the latter accounts for 50% or more of total electricity produced. The authors point to V2G technology as the missing critical factor for a major renewable energy deployment without incurring high storage costs, while assuring electric grid balance and reliability.

Lund and Kempton (2008) analyse the effects of the electrification of road transport on the electric system with large-scale wind power. The results indicate that EV diffusion prevents wind turbines from being turned off.

Although V2G technology is at an early stage, in the future it has the potential to provide optimal management of electric systems, by allowing the optimal charging or discharging of electricity. This will hasten the dynamic incorporation of intermittent energy sources.

Taking into consideration both the additional demand and V2G services associated to electric vehicles, Schill (2010) develop a game-theoretic model to assess the impacts of a hypothetic fleet of one million plug-in electric vehicles (PIEV) on the imperfectly competitive German electricity market. The analysis is made for different scenarios in relation to who is in charge of vehicle recharging management and storage capacity use, which can be either controlled by individual vehicle owners or by service providers (electricity generating firms or players which do not own electricity generation capacity). Electricity could be used for vehicle recharging, then stored electricity would be sold back to the market. If excess batteries are allowed for storage, electricity producers suffer from the price-smoothing effect of additional storage, while consumer surplus increase. These results differ depending on who is responsible for storage operations, and on battery degradation costs. Moreover, storage operations mitigate the market power of strategic electricity generators. This mitigation is more significant when storage operations are managed by a single player that is also an electricity generation firm. Looking at the potential market power exerted by electric vehicle fleets, Schill (2010) argues that it is an unlikely situation in Germany, regardless who controls them.

3.5.3. Some insights from the literature for Portugal

Lopes *et al.* (2009) examine the impacts of different shares of EVs on the Portuguese Low Voltage grid and also on the global generation pattern. The analysis of the maximum share of EV that can be integrated into the grid without violating the system´s technical restrictions and act in accordance with drivers' demands regarding the predicted use of vehicles is based on two distinct charging strategies: dumb charging and smart charging. The results suggest that when dumb charging is applied, an 11% share of EVs can be perfectly integrated without requiring modifications in either electricity generation or distribution networks for the residential area under study. This finding implies that EV diffusion is hobbled if no charging control is applied. If smart charging is implemented, however, it is possible to increase EV share up to 61%, without the need to reinforce the grid, if only half of those EVs are subject to control charge. In this work the authors also focus on the benefits that could be achieved from the better use of excess renewable energy, currently forfeited, and the environmental benefits of EV in terms of emissions, discussed in the previous section. Once again there are differences depending on the adopted charging strategy. Smart charging allows for higher penetration and thus for more energy storage by EV, to be discharged later into the system, and it also has more environmental benefits.

Camus *et al.* (2011) perform simulations for several scenarios for 2020 EV diffusion and charging patterns integrated with different alternatives for the electricity generation mix using Portugal as a case study, in order to identify the effects in load profiles, spot electricity prices and emissions levels. For the year 2020 and a hypothetical 2 million EV, a scenario of low hydroelectricity production and high prices translates into energy costs for EV recharge of 20 cents/kWh, with 2 million EV charging mostly at evening peak hours. For the opposite case – high hydroelectricity production, low prices and off-peak charging – the energy costs for EV recharging decrease to 5.6 cents/kWh. The impacts on energy, fossil fuels use and emissions are more significant when recharge occurs at off-peak periods. For example, in another scenario where a dry year is assumed along with off-peak recharge, savings in primary energy, fossil-fuel use and carbon dioxide emissions of up to 3%, 14% and 10%, correspondingly, are shown to occur for electricity and transport, taken together, when compared with a scenario without EV.

Pina *et al.* (2014) choose the Azores, namely the Flores Island, to investigate the impacts on primary energy consumption and CO_2 emissions of EV penetration in the energy system, based on four different scenarios regarding different EV share and recharging strategies (fixed and flexible). The results indicate that only for a small amount of EV share – 10% to 40% - is RES recharging guaranteed, when there is a substantial weight of RES-E production, ranging from 60% to 62%.The additional electricity demand has to be generated mostly from diesel generators. Flexible recharging strategies double the share of RES-E for EVs recharging, when compared to fixed recharge, and as a results induce a decrease in total primary energy consumption between 02% and 1.1%. As for CO_2 emissions, these could be reduced by between 0.3% and 1.7%. Relying on these findings, Pina *et al.* (2014) emphasize that the sustainability of energy systems can be improved by EV introduction.

These few examples of the literature that focus on Portugal suggest that EVs massification may in fact contribute to GHG savings and RES-E deployment, but the final outcomes will depend on the share of EVs and their recharging scheme.

3.6. Concluding remarks

Electricity generated from renewable energy sources is one of the main strategies to reduce dependence on fossil fuels, improve energy supply security and achieve GHG emissions savings. Deregulated electricity markets are frequently modeled as Cournot oligopolies because there are some economic barriers that hinder perfect competition.

Within this framework, it is important to examine the long-run investment decisions of private RES-E producer in capacity expansion. To do so, we design a two-period Cournot model based on the work developed by Genc and Thile (2011), with a thermal generation producer and a RES-E producer. We assess the investment decision of a RES-E producer with capacity constraints with and without feed.in tariffs. This work enabled us to reach some conclusions. The increase of electricity demand in the second period *per se* is not sufficient to guarantee the increase of total market quantity because it depends on whether the RES-E capacity expansion is higher or lower than $D_1 - D_0$. As the positive or negative balance of GHG emissions depends directly on the thermal producer's output, only when $I_0 > D_1 - D_0$ are GHG savings verified. The role of this relation between investment and the increase of electricity demand is similar with and without feed-in tariffs. The assumption that the feed-in value is higher than the market price for $t = 1$ ensures that the investment level is superior when this RES-E support mechanism is applied.

When confronting the Cournot-Nash solution with the social optimum, the environmental damages caused by the thermal producer must be considered. The main conclusion is that the underproduction of duopoly firms is not always verified, because it depends on the magnitude of the external costs.

The model highlights that deregulated electricity markets raise additional challenges for regulators, because of the concurrent existence of two market failures. The first market failure is market power, which is related to the strategic underproduction by oligopolistic firms and where prices do not reflect the marginal costs of production. The second market failure is the negative externality caused by the fossil-fuel based electricity producers during the generation process, which leads social costs to be higher than private production costs. If treated separately by distinct regulators, these linked externalities can yield some problematic results. In perfectly competitive markets with negative externalities, the socially optimal quantities are less and the price higher when compared to the case where social costs imposed to society are not considered. Hence market power may decrease the undesirable consequences on welfare of a negative externality (Lipsey and Lancaster 1956; Hammer 2000). The adoption of regulation measures must take in consideration the balance between the welfare losses as consequence of underproduction due to the market power and the welfare gains that subsequently arise from that same underproduction of fossil-fuel electricity generation plants. As the market failures are linked, joint, or at least coordinated, actions are required from the involved regulators in order to choose the most suitable instruments to correct them.

Regulators have at their disposal a range of regulatory instruments, such as taxes, subsidies, tradable permits, or price regulation for addressing different types of market failures. The choice depends on the problems to be addressed and in which circumstances the instrument is adopted (Bennear and Stavins, 2007). Market power and environmental negative externalities need multiple policy instruments decided by both economic and environmental regulators. The inexistence of coordination of both economic and environmental regulators may lead to potential overlaps and negative interactions among regulatory instruments (Goulder and Parry, 2008). Therefore, regulators should work together in the selection of one or more policy instruments to increase the Pareto-efficiency of oligopolistic electricity generation markets.

This chapter 3 should be seen a partial analysis of future electricity markets due to the present status of electric vehicles worldwide. The interaction between these vehicles and RES-E deserves special attention, because of two main reasons. On the one hand it is desirable that the additional electricity demand for battery charging is from renewable sources; on the other hand, the potential V2G services provided by EV is aimed to instigate of intermittent RES-E penetration. The strategic behaviour of electricity generation firms could incorporate the role of EV not only as drivers of increased demand but also as storage devices and potential electricity suppliers. The absence of a road transport model with electric vehicles and associated services thus constitutes a fragility of our model. Future research may address this limitation by focusing on the interplay between electricity and road transport sector. As we could see from section 3.5.3., EVs penetration may in fact contribute to achieve GHG emissions and RES-E targets, combined with the recharging management.

Apart from the Cournot model, alternative models might be designed in order to illustrate the steady growth of RES-E in the market and capacity expansion decisions. Ventosa *et al.* (2002) developed both a Cournot and a Stackelberg model for capacity expansion planning considering future demand as certain and where only one competitor can invest. The two models have in common the competition on quantity, which suits the electricity generation markets, but contrary to Cournot, the Stackelberg game illustrates a leader-follower reality in which the leader producer moves first and then the followers decide their quantities based on the strategy chosen by the leader. Generally, the largest producer is the leader (Chen *et al.*, 2006), thus it would be interesting to apply the Stackelberg game to our model assuming the thermal producer as the leader and RES-E producers as the followers, to compare the equilibrium results.

Alternatively, as in Reichenbach and Requate (2012), we could incorporate RES-E in an imperfect market as several small producers that act as a competitive fringe instead of one RES-E producer. In this case, fringe competitors could, for example, be price-takers or receive a feed-in tariff. It would be interesting to examine how the dominant firms would behave in the presence of their fringe competitors.

Finally, given that renewable technologies are in different stages of maturity and also differ regarding the intermittency (e.g. wind) and unpredictability (e.g. precipitation levels), another alternative approach to ours would be to distinguish between renewable sources. For instance, hydropower producer could be represented as one of the dominant firms as in Genc and Thille (2011) and other more immature renewable technologies could act as competitive fringe.

For future work, instead of linear inverse demand function and quadratic costs functions which are broadly used in related literature, other functional forms should be explored in order to compare the equilibrium solutions, and try to understand which one best suits the electricity generation market under an oligopoly. Moreover, future research should be conducted providing numerical simulations to evaluate our theoretical findings and also to overcome some of the model shortcomings.

FINAL CONSIDERATIONS AND FUTURE RESEARCH

The present and future consequences of greenhouse gas (GHG) emissions in terms of climate change have boosted the debate over the dynamics between the economy and emissions. The debate is still full of uncertainties and, thus, urges us to strive for a lasting effort to deepen our knowledge about climate change from an economic perspective.

With this thesis we tried to contribute to the clarification of some of the issues salient in the economics of climate change, focusing on the case of Portugal. Along with national aggregate data, sectoral analyses regarding the two main CO_2 emitters – electricity generation and transport – allowed us to better characterize the relationship between economic growth and emissions.

To proceed with the empirical analysis based on the EKC hypothesis, we used two complementary econometric procedures which have seldom been used in the literature in spite of their relevance to the methodological debates. The outcomes of nonlinear cointegration indicate that electricity generation and transport sectors are at different stages regarding the relationship between economic growth and carbon emissions, and that the relationship at national level is, as expected, strongly influenced by the reality of these two sectors. Thus, it is interesting to note that the electricity generation sector is on the descending part of an EKC, unlike the transport sector for which the results suggest a positive monotonic relationship instead of an inverted U-shaped link. The national results validate the EKC hypothesis; however, the turning point, albeit close, has not yet been reached.

As for the empirical study using a linear functional formula and considering the possibility of structural breaks between 1960 and 2010, we realized that the estimated break dates stress the effect that the Community Support Framework I and II had in the long-term relationship of CO_2 emissions and economy, both at sectoral and national levels. The validation of EKC is not consistent with the results obtained with previous nonlinear cointegration methodology. However, it is worth to stress that the two econometric procedures are complementary rather than substitutes. Still, it is relevant to continue to study this long-term relationship, for example without imposing a functional formula *a priori*, drawing on nonparametric methodologies.

In a further contribution of this thesis, for both empirical analyses we extend the EKC model to include control variables that also possibly explain CO_2 emissions. From this analysis, we highlight the importance of the economic and climate variables. In particular, crude and fuel prices and precipitation contribute to a fuller explanation of carbon emissions. While crude and fuel prices affect both of the main sectors, precipitation helps explain CO_2 emission reductions for the electricity generation sector and total CO_2, due to the significant weight of hydropower in the national electric generation system. Future work should explore other control variables such as RES-E installed capacity, the average km-travelled or the evolution in the annual number of public transport trips per capita.

Alongside the empirical studies of chapters 1 and 2, we also sought to examine from a theoretical approach the strategic behaviour of thermal electricity and RES-E producers in a deregulated market, in chapter 3. With a two-period Cournot model we were able to conclude that the increase or decrease of thermal generation, market quantity and GHG emissions from $t = 0$ to $t = 1$ is not straightforward, as it depends on the relation between investment level and electricity demand. Thermal generation and GHG emissions are directly related, hence, when RES-E investment capacity is higher than electricity demand, there is a reduction of the quantity of electricity produced by the thermal generator in the second period as well as GHG emissions and market quantity. However, nothing can be said about market equilibrium when the investment is lower than the increase on electricity demand. These conclusions are valid with and without feed-in tariffs. We showed that feed-in tariffs can play an important role in promoting RES-E expansion capacity, as long as they are higher than the market price.

Through the comparison of Cournot-Nash solutions and welfare-maximizing solutions sought by the social planner, we found that the socially optimal quantities and investment outcomes are not always higher than the oligopoly ones. Fundamentally, it will depend on the dimension of the environmental damages caused by the thermal generator.

Despite these findings, our Cournot model may be seen as an isolated analysis of a much wider research topic. This model should be developed further to overcome its present limitations. Given the potential role of electric vehicles (EVs), a road transport model should be designed capable of capture the interplay between the two sectors linked by EV. Special emphasis should be given to the impacts that the massification of EV might have in promoting intermittent RES-E, and to the potential net GHG emissions savings considering the two sectors together. The relevance of this future research relates to the fact that EVs induce an increase in electricity demand, but are also capable of bidirectional charging (V2G), storing electricity from intermittent RES-E that may subsequently be injected into the network to cope with periods of increased demand for electricity.

This thesis left many unanswered questions and was driven by the appearance of many others. Nonetheless, there are two general conclusions that can, and should, contribute to the discussion of climate-change mitigation policies. First, the reduction of GHG emissions at national level may require measures that take into account the specificities of each emitting sector, since their evolution has different drivers and similar behaviour cannot be taken for granted. Second, measures aimed at emission reduction must then be analysed in models that consider possible interactions between the different policies, because measures that are too narrowly focused on sectoral emissions could be counterproductive at a national level, resulting in loss of social welfare.

We believe that this thesis, despite the results already achieved, is, above all, the beginning of an investigation aimed at contributing to the debate on climate change and the implementation of cost-effective GHG mitigation policies.

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APPENDIX A – DATA PLOTS AND UNIT ROOT TESTS

A.1. Data plots

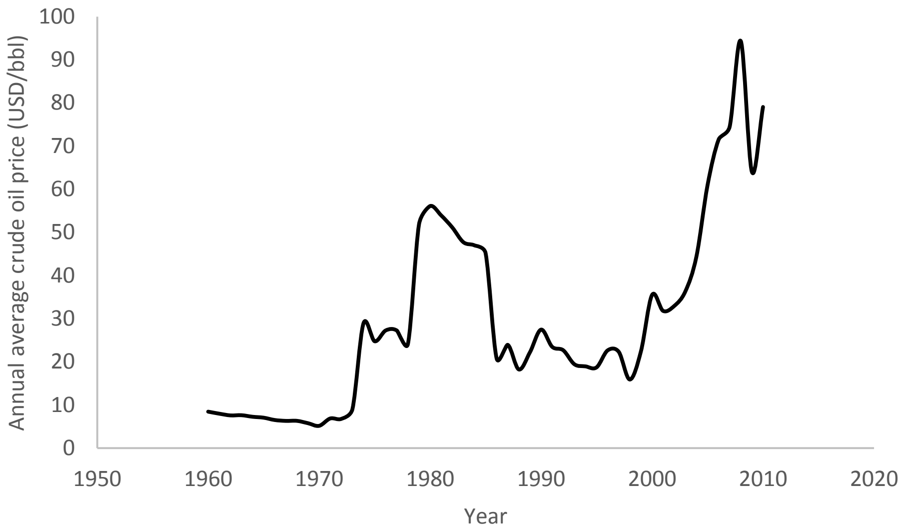

Figure A.1: Annual average crude oil price from1960 to 2010 (USD, base year=2010).

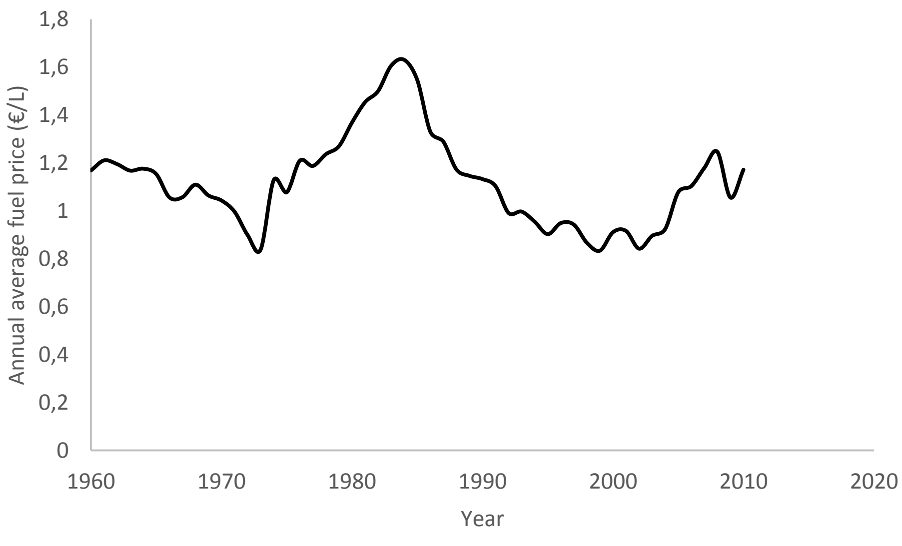

Figure A.2: Annual average fuel price for Portugal from1960 to 2010 (euros, base year=2006).

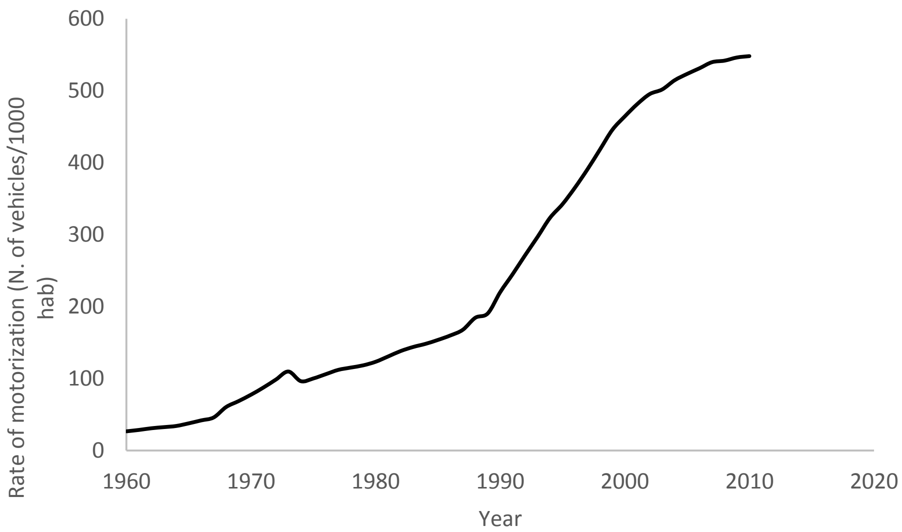

Figure A.3: Rate of motorization for Portugal from1960 to 2010.

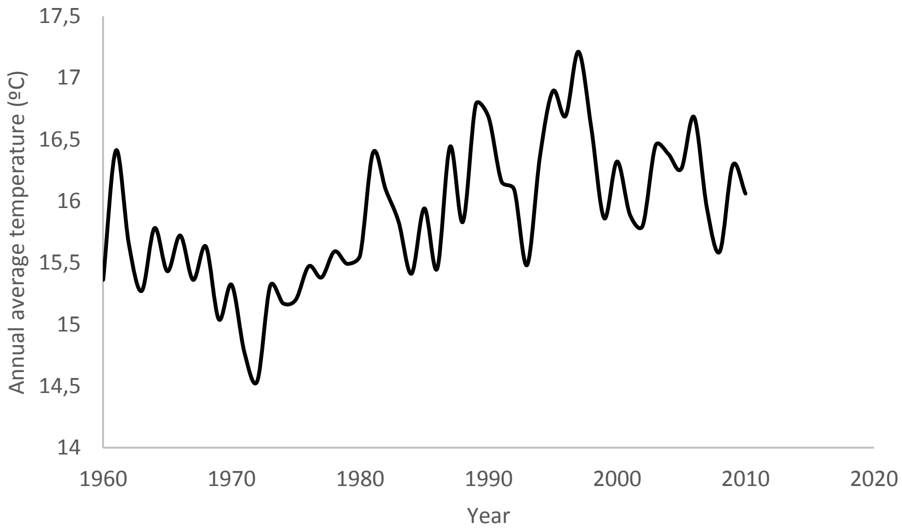

Figure A.4: Annual average temperature for Portugal from1960 to 2010.

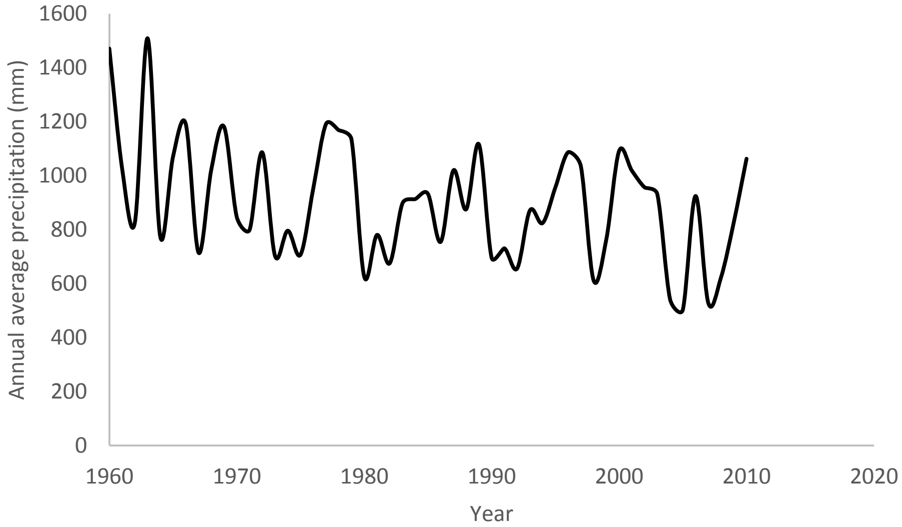

Figure A.5: Annual average precipitation for Portugal from1960 to 2010.

A.2. Unit root tests

Table A.1: Results of the ADF test.

	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
ADF test (Intercept)									
Test statistic	-0.890022	-1.282640	-1.229925	-0.557709	-0.774849	-1.563627	-0.819488	-3.519397	-6.506973
Critical value 1%	-3.571310	-3.568308	-3.568308	-3.571310	-3.568308	-3.568308	-3.574446	-3.568308	-3.568308
Critical value 5%	-2.922449	-2.921175	-2.921175	-2.922449	-2.921175	-2.921175	-2.923780	-2.921175	-2.921175
Critical value 10%	-2.599224	-2.598551	-2.598551	-2.599224	-2.598551	-2.598551	-2.599925	-2.598551	-2.598551
Lag length	1	0	0	1	0	0	2	0	0
Order of integration	I(1)	#	I(1)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)
ADF test (Intercept and trend)									
	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
Test statistic	-2.914987	-3.884886	-1.973997	-1.681723	-1.738204	-1.529839	-2.264935	-4.466807	-6.997206
Critical value 1%	-4.156734	-4.205004	-4.152511	-4.156734	-4.152511	-4.152511	-4.161144	-4.152511	-4.152511
Critical value 5%	-3.504330	-3.526609	-3.502373	-3.504330	-3.502373	-3.502373	-3.506374	-3.502373	-3.502373
Critical value 10%	-3.181826	-3.194611	-3.180699	-3.181826	-3.180699	-3.180699	-3.183002	-3.180699	-3.180699
Lag length	1	10	0	1	0	0	2	0	0
Order of integration	I(1)	I(0)	I(1)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)

#: Inconclusive results

Table A.2: Results of the PP test.

	PP test (Intercept)								
	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
Test statistic	-0.745922	-1.256098	-1.163386	-0.277320	-0.640911	-1.798850	0.982113	-3.550850	-6.506973
Critical value 1%	-3.568308	-3.568308	-3.568308	-3.568308	-3.568308	-3.568308	-3.568308	-3.568308	-3.568308
Critical value 5%	-2.921175	-2.921175	-2.921175	-2.921175	-2.921175	-2.921175	-2.921175	-2.921175	-2.921175
Critical value 10%	-2.598551	-2.598551	-2.598551	-2.598551	-2.598551	-2.598551	-2.598551	-2.598551	-2.598551
Bandwidth	2	5	4	4	2	3	5	4	0
Order of integration	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)
	PP test (Intercept and trend)								
	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
Test statistic	-1.880131	-0.989194	-2.006420	-1.777776	-1.773051	-1.789003	-1.700636	-4.556133	-7.278897
Critical value 1%	-4.152511	-4.152511	-4.152511	-4.152511	-4.152511	-4.152511	-4.152511	-4.152511	-4.152511
Critical value 5%	-3.502373	-3.502373	-3.502373	-3.502373	-3.502373	-3.502373	-3.502373	-3.502373	-3.502373
Critical value 10%	-3.180699	-3.180699	-3.180699	-3.180699	-3.180699	-3.180699	-3.180699	-3.180699	-3.180699
Bandwidth	2	5	5	4	3	3	5	3	5
Order of integration	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)

Table A.3: Results of the KPSS test.

	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
KPSS test (Intercept)									
Test statistic	0.938959	0.897507	0.880295	0.904211	0.514722	0.175866	0.891407	0.622271	0.554778
Critical value 1%	0.739000	0.739000	0.739000	0.739000	0.739000	0.739000	0.739000	0.739000	0.739000
Critical value 5%	0.463000	0.463000	0.463000	0.463000	0.463000	0.463000	0.463000	0.463000	0.463000
Critical value 10%	0.347000	0.347000	0.347000	0.347000	0.347000	0.347000	0.347000	0.347000	0.347000
Bandwidth	5	5	5	5	5	5	5	5	3
Order of integration	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)	I(0)	I(0)
KPSS test (Intercept and trend)									
	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
Test statistic	0.094519	0.094159	0.113314	0.166354	0.094867	0.105280	0.219779	0.115378	0.066317
Critical value 1%	0.216000	0.216000	0.216000	0.216000	0.216000	0.216000	0.216000	0.216000	0.216000
Critical value 5%	0.146000	0.146000	0.146000	0.146000	0.146000	0.146000	0.146000	0.146000	0.146000
Critical value 10%	0.119000	0.119000	0.119000	0.119000	0.119000	0.119000	0.119000	0.119000	0.119000
Bandwidth	5	5	5	5	5	5	5	4	2
Order of integration	I(0)	I(0)	I(0)	I(1)	I(0)	I(0)	I(1)	I(0)	I(0)

Table A.4: Results of the DFGLS test.

	DFGLS test (Intercept)								
	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
Test statistic	0.396439	-1.995227	-0.623716	-0.327898	-0.478170	-1.568027	-0.981380	-3.197692	-3.174491
Critical value 1%	-2.613010	-2.619851	-2.612033	-2.614029	-2.612033	-2.612033	-2.614029	-2.612033	-2.612033
Critical value 5%	-1.947665	-1.948686	-1.947520	-1.947816	-1.947520	-1.947520	-1.947816	-1.947520	-1.947520
Critical value 10%	-1.612573	-1.612036	-1.612650	-1.612492	-1.612650	-1.612650	-1.612492	-1.612650	-1.612650
Lag length	1	7	0	2	0	0	2	0	0
Order of integration	I(1)	I(0)	I(1)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)
	DFGLS test (Intercept and trend)								
	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
Test statistic	-2.993701	-4.959469	-2.137790	-1.594826	-1.863140	-1.586180	-2.139147	-4.551154	-5.649491
Critical value 1%	-3.770000	-3.770000	-3.770000	-3.770000	-3.770000	-3.770000	-3.770000	-3.770000	-3.770000
Critical value 5%	-3.190000	-3.190000	-3.190000	-3.190000	-3.190000	-3.190000	-3.190000	-3.190000	-3.190000
Critical value 10%	-2.890000	-2.890000	-2.890000	-2.890000	-2.890000	-2.890000	-2.890000	-2.890000	-2.890000
Lag length	1	10	0	1	0	0	2	0	0
Order of integration	I(1)	I(0)	I(1)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)

Table A.5: Results of the Ng-PP test.

	Ng-PP test (Intercept)								
Test statistic	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
MZ_{α}	0.62102	-6.82160	-0.77844	-0.31789	-1.20309	-4.54517	-49.8113	-14.2362	-13.5152
Critical value 1%	-13.8000	-13.8000	-13.8000	-13.8000	-13.8000	-13.8000	-13.8000	-13.8000	-13.8000
Critical value 5%	-8.10000	-8.10000	-8.10000	-8.10000	-8.10000	-8.10000	-8.10000	-8.10000	-8.10000
Critical value 10%	-5.70000	-5.70000	-5.70000	-5.70000	-5.70000	-5.70000	-5.70000	-5.70000	-5.70000
Order of integration	I(1)	I(1)	I(2)	I(1)	I(1)	I(1)	I(0)	I(0)	I(0)
MZ_t	0.45210	-1.75688	-0.49063	-0.16716	-0.39833	-1.50611	-4.88853	-2.64747	-2.59534
Critical value 1%	-2.58000	-2.58000	-2.58000	-2.58000	-2.58000	-2.58000	-2.58000	-2.58000	-2.58000
Critical value 5%	-1.98000	-1.98000	-1.98000	-1.98000	-1.98000	-1.98000	-1.98000	-1.98000	-1.98000
Critical value 10%	-1.62000	-1.62000	-1.62000	-1.62000	-1.62000	-1.62000	-1.62000	-1.62000	-1.62000
Order of integration	I(1)	I(1)	I(2)	I(1)	I(1)	I(1)	I(0)	I(0)	I(0)
MSB	0.72800	0.25755	0.63027	0.52585	0.33109	0.33137	0.09814	0.18597	0.19203
Critical value 1%	0.17400	0.17400	0.17400	0.17400	0.17400	0.17400	0.17400	0.17400	0.17400
Critical value 5%	0.23300	0.23300	0.23300	0.23300	0.23300	0.23300	0.23300	0.23300	0.23300
Critical value 10%	0.27500	0.27500	0.27500	0.27500	0.27500	0.27500	0.27500	0.27500	0.27500
Order of integration	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(0)	I(0)
MPT	37.3493	3.90243	22.0831	19.2058	10.4240	5.39307	0.74968	1.79972	1.82905
Critical value 1%	1.78000	1.78000	1.78000	1.78000	1.78000	1.78000	1.78000	1.78000	1.78000
Critical value 5%	3.17000	3.17000	3.17000	3.17000	3.17000	3.17000	3.17000	3.17000	3.17000
Critical value 10%	4.45000	4.45000	4.45000	4.45000	4.45000	4.45000	4.45000	4.45000	4.45000
Order of integration	I(1)	I(1)	I(2)	I(1)	I(1)	I(1)	I(0)	I(0)	I(0)
Lag length	1	7	0	2	0	0	2	0	0

Table A.5: Results of the Ng-PP test (continued).

Ng-PP test (Intercept and trend)									
Test statistic	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
MZ_{α}	-20.8410	2.79190	-9.28572	-6.13975	-6.78699	-4.76501	-11.8505	-20.9662	-23.2853
Critical value 1%	-23.8000	-23.8000	-23.8000	-23.8000	-23.8000	-23.8000	-23.8000	-23.8000	-23.8000
Critical value 5%	-17.3000	-17.3000	-17.3000	-17.3000	-17.3000	-17.3000	-17.3000	-17.3000	-17.3000
Critical value 10%	-14.2000	-14.2000	-14.2000	-14.2000	-14.2000	-14.2000	-14.2000	-14.2000	-14.2000
Order of integration	I(1)	#	I(2)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)
MZ_t	-3.16126	4.29384	-1.91912	-1.72662	-1.73658	-1.52118	-2.41629	-3.22197	-3.33217
Critical value 1%	-3.42000	-3.42000	-3.42000	-3.42000	-3.42000	-3.42000	-3.42000	-3.42000	-3.42000
Critical value 5%	-2.91000	-2.91000	-2.91000	-2.91000	-2.91000	-2.91000	-2.91000	-2.91000	-2.91000
Critical value 10%	-2.62000	-2.62000	-2.62000	-2.62000	-2.62000	-2.62000	-2.62000	-2.62000	-2.62000
Order of integration	I(1)	#	I(2)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)
MSB	0.15168	1.53796	0.20667	0.28122	0.25587	0.31924	0.20390	0.15367	0.14310
Critical value 1%	0.14300	0.14300	0.14300	0.14300	0.14300	0.14300	0.14300	0.14300	0.14300
Critical value 5%	0.16800	0.16800	0.16800	0.16800	0.16800	0.16800	0.16800	0.16800	0.16800
Critical value 10%	0.18500	0.18500	0.18500	0.18500	0.18500	0.18500	0.18500	0.18500	0.18500
Order of integration	I(1)	#	I(2)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)
MPT	4.77566	633.050	10.7235	14.8237	13.5143	18.9826	7.78515	4.44233	4.39059
Critical value 1%	4.03000	4.03000	4.03000	4.03000	4.03000	4.03000	4.03000	4.03000	4.03000
Critical value 5%	5.48000	5.48000	5.48000	5.48000	5.48000	5.48000	5.48000	5.48000	5.48000
Critical value 10%	6.67000	6.67000	6.67000	6.67000	6.67000	6.67000	6.67000	6.67000	6.67000
Order of integration	I(1)	#	I(2)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)
Lag length	1	10	0	1	0	0	2	0	0

#: Inconclusive results

Table A.6: Results of the ERS test.

	ERS test (Intercept)								
	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
Test statistic	183.8637	110.7184	29.85947	143.4646	12.83095	5.248583	36.32615	2.167476	4.712498
Critical value 1%	1.871600	1.871600	1.871600	1.871600	1.871600	1.871600	1.871600	1.871600	1.871600
Critical value 5%	2.972800	2.972800	2.972800	2.972800	2.972800	2.972800	2.972800	2.972800	2.972800
Critical value 10%	3.915200	3.915200	3.915200	3.915200	3.915200	3.915200	3.915200	3.915200	3.915200
Lag length	1	0	0	1	0	0	2	0	0
Order of integration	I(1)	#	I(1)	I(1)	I(1)	I(1)	I(2)	I(0)	I(1)
	ERS (Intercept and trend)								
	GDP	CO2total	CO2elect	CO2transp	Crude	Fuel	Motor	Temperature	Precipitation
Test statistic	4.400892	725.7649	10.70058	16.26952	13.22694	18.28890	12.94773	4.281823	6.536774
Critical value 1%	4.220800	4.220800	4.220800	4.220800	4.220800	4.220800	4.220800	4.220800	4.220800
Critical value 5%	5.718400	5.718400	5.718400	5.718400	5.718400	5.718400	5.718400	5.718400	5.718400
Critical value 10%	6.770400	6.770400	6.770400	6.770400	6.770400	6.770400	6.770400	6.770400	6.770400
Lag length	1	10	0	1	0	0	2	0	0
Order of integration	I(1)	#	I(1)	I(1)	I(1)	I(1)	I(2)	I(0)	I(0)*

#: Inconclusive results

APPENDIX B – NONLINEAR COINTEGRATION TESTS

Table B.1: Results of the rank test for cointegration (Breitung, 2001).

	Rank test for cointegration					
	CO2total-GDP		CO2elect-GDP		CO2transp-GDP	
Test statistics	$K^*_{T=500}$	$\xi^*_{T=500}$	$K^*_{T=500}$	$\xi^*_{T=500}$	$K^*_{T=500}$	$\xi^*_{T=500}$
	0.6262	0.0301	0.5505	0.0195	0.4523	0.0189
	Critical values					
	Critical value 1%		Critical value 5%		Critical value 10%	
$K^*_{T=500}$	0.3165		0.3635		0.3941	
$\xi^*_{T=500}$	0.0130		0.0188		0.0232	

Table B.2: Results of tests for nonlinear cointegration (Choi and Saikkonen, 2010):

CO2total-GDP (quadratic) – NLLS residuals						
Rule	Bandwidth	Block size	M	$C_{NLLS}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.2661	0.5299	0.0250
	Fixed[1]	35	2	0.2722	0.5221	0.0250
	Fixed[2]	35	2	0.2310	0.5783	0.0250
Minimum volatility	Automatic	28	2	0.4386	0.3633	0.0250
	Fixed[1]	29	2	0.4225	0.3753	0.0250
	Fixed[2]	28	2	0.2030	0.6223	0.0250
CO2total-GDP (quadratic) – DOLS, K=1 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1229	0.7825	0.0250
	Fixed[1]	35	2	0.1605	0.7002	0.0250
	Fixed[2]	35	2	0.1504	0.7210	0.0250
Minimum volatility	Automatic	27	2	0.1651	0.6910	0.0250
	Fixed[1]	27	2	0.2933	0.4967	0.0250
	Fixed[2]	27	2	0.1648	0.6917	0.0250
CO2total-GDP (quadratic) – DOLS, K=2 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1127	0.8071	0.0250
	Fixed[1]	35	2	0.1395	0.7445	0.0250
	Fixed[2]	35	2	0.1583	0.7047	0.0250
Minimum volatility	Automatic	32	2	0.1329	0.7593	0.0250
	Fixed[1]	32	2	0.1482	0.7255	0.0250
	Fixed[2]	27	2	0.1330	0.7591	0.0250
CO2total-GDP (quadratic) – DOLS, K=3 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1551	0.7112	0.0250
	Fixed[1]	35	2	0.1616	0.6979	0.0250
	Fixed[2]	35	2	0.1595	0.7021	0.0250
Minimum volatility	Automatic	27	2	0.0866	0.8738	0.0250
	Fixed[1]	27	2	0.1290	0.7681	0.0250
	Fixed[2]	27	2	0.0958	0.8498	0.0250

(1) Lag length for the long-run variance estimation: $[4(b/100)^{0.25}]$
(2) Lag length for the long-run variance estimation: $[12(b/100)^{0.25}]$

Table B.3: Results of tests for nonlinear cointegration (Choi and Saikkonen, 2010):

CO2elect-GDP (quadratic) – NLLS residuals						
Rule	Bandwidth	Block size	M	$C_{NLLS}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.3113	0.4765	0.0250
	Fixed[1]	35	2	0.2342	0.5736	0.0250
	Fixed[2]	35	2	0.1419	0.7392	0.0250
Minimum volatility	Automatic	28	2	0.8662	0.1692	0.0250
	Fixed[1]	28	2	0.8662	0.1692	0.0250
	Fixed[2]	28	2	0.4250	0.3735	0.0250
CO2elect-GDP (quadratic)– DOLS, K=1 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1484	0.7252	0.0250
	Fixed[1]	35	2	0.1639	0.6933	0.0250
	Fixed[2]	35	2	0.1297	0.7666	0.0250
Minimum volatility	Automatic	27	2	0.4198	0.3774	0.0250
	Fixed[1]	27	2	0.5929	0.2709	0.0250
	Fixed[2]	27	2	0.4123	0.3832	0.0250
CO2elect-GDP (quadratic)– DOLS, K=2 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1307	0.7644	0.0250
	Fixed[1]	35	2	0.1386	0.7466	0.0250
	Fixed[2]	35	2	0.1405	0.7424	0.0250
Minimum volatility	Automatic	32	2	0.1733	0.6751	0.0250
	Fixed[1]	32	2	0.1809	0.6609	0.0250
	Fixed[2]	32	2	0.1960	0.6341	0.0250
CO2elect-GDP (quadratic)– DOLS, K=3 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1365	0.7512	0.0250
	Fixed[1]	35	2	0.1380	0.7479	0.0250
	Fixed[2]	35	2	0.1868	0.6502	0.0250
Minimum volatility	Automatic	27	2	0.3837	0.4067	0.0250
	Fixed[1]	30	2	0.3586	0.4291	0.0250
	Fixed[2]	30	2	0.5476	0.2945	0.0250

(1) Lag length for the long-run variance estimation: $[4(b/100)^{0.25}]$
(2) Lag length for the long-run variance estimation: $[12(b/100)^{0.25}]$

Table B.4: Results of tests for nonlinear cointegration (Choi and Saikkonen, 2010):

CO2transp-GDP (quadratic) – NLLS residuals						
Rule	Bandwidth	Block size	M	$C_{NLLS}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1657	0.6898	0.0250
	Fixed[1]	35	2	0.1657	0.6898	0.0250
	Fixed[2]	35	2	0.1482	0.7255	0.0250
Minimum volatility	Automatic	23	3	0.3441	0.4428	0.0167
	Fixed[1]	25	3	0.4642	0.3453	0.0167
	Fixed[2]	32	2	0.0968	0.8473	0.0250
CO2transp-GDP (quadratic)– DOLS, K=1 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1004	0.8379	0.0250
	Fixed[1]	35	2	0.1153	0.8008	0.0250
	Fixed[2]	35	2	0.1027	0.8321	0.0250
Minimum volatility	Automatic	32	2	0.0858	0.8758	0.0250
	Fixed[1]	32	2	0.1133	0.8057	0.0250
	Fixed[2]	32	2	0.0955	0.8507	0.0250
CO2transp-GDP (quadratic)– DOLS, K=2 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.0932	0.8566	0.0250
	Fixed[1]	35	2	0.1174	0.7956	0.0250
	Fixed[2]	35	2	0.0937	0.8553	0.0250
Minimum volatility	Automatic	31	2	0.1406	0.7420	0.0250
	Fixed[1]	32	2	0.1142	0.8034	0.0250
	Fixed[2]	32	2	0.0954	0.8508	0.0250
CO2transp-GDP (quadratic)– DOLS, K=3 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1283	0.7699	0.0250
	Fixed[1]	35	2	0.0997	0.8399	0.0250
	Fixed[2]	35	2	0.1039	0.8291	0.0250
Minimum volatility	Automatic	30	2	0.0837	0.8813	0.0250
	Fixed[1]	31	2	0.1339	0.7570	0.0250
	Fixed[2]	31	2	0.0869	0.8730	0.0250

(1) Lag length for the long-run variance estimation: $[4(b/100)^{0.25}]$
(2) Lag length for the long-run variance estimation: $[12(b/100)^{0.25}]$

Table B.5: Results of tests for nonlinear cointegration (Choi and Saikkonen, 2010):

CO2total-GDP (cubic) – NLLS residuals						
Rule	Bandwidth	Block size	M	$C_{NLLS}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.0987	0.8424	0.0250
	Fixed[1]	35	2	0.1143	0.8032	0.0250
	Fixed[2]	35	2	0.0929	0.8572	0.0250
Minimum volatility	Automatic	32	2	0.0838	0.8810	0.0250
	Fixed[1]	32	2	0.1036	0.8299	0.0250
	Fixed[2]	31	2	0.0992	0.8411	0.0250
CO2total-GDP (cubic) – DOLS, K=1 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1133	0.8058	0.0250
	Fixed[1]	35	2	0.1272	0.7723	0.0250
	Fixed[2]	35	2	0.1221	0.7845	0.0250
Minimum volatility	Automatic	32	2	0.1055	0.8251	0.0250
	Fixed[1]	29	2	0.1365	0.7511	0.0250
	Fixed[2]	32	2	0.0902	0.8644	0.0250
CO2total-GDP (cubic) – DOLS, K=2 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1044	0.8277	0.0250
	Fixed[1]	35	2	0.1383	0.7473	0.0250
	Fixed[2]	35	2	0.1253	0.7768	0.0250
Minimum volatility	Automatic	32	2	0.0956	0.8503	0.0250
	Fixed[1]	29	2	0.1512	0.7193	0.0250
	Fixed[2]	31	2	0.0833	0.8823	0.0250
CO2total-GDP (cubic) – DOLS, K=3 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1568	0.7076	0.0250
	Fixed[1]	35	2	0.1765	0.6691	0.0250
	Fixed[2]	35	2	0.1760	0.6700	0.0250
Minimum volatility	Automatic	29	2	0.1123	0.8082	0.0250
	Fixed[1]	27	2	0.1441	0.7345	0.0250
	Fixed[2]	29	2	0.0851	0.8777	0.0250

(1) Lag length for the long-run variance estimation: $[4(b/100)^{0.25}]$
(2) Lag length for the long-run variance estimation: $[12(b/100)^{0.25}]$

Table B.6: Results of tests for nonlinear cointegration (Choi and Saikkonen, 2010):

CO2elect-GDP (cubic) – NLLS residuals						
Rule	Bandwidth	Block size	M	$C_{NLLS}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1071	0.8210	0.0250
	Fixed[1]	35	2	0.0780	0.8962	0.0250
	Fixed[2]	35	2	0.1351	0.7542	0.0250
Minimum volatility	Automatic	32	2	0.1318	0.7618	0.0250
	Fixed[1]	28	2	0.1358	0.7527	0.0250
	Fixed[2]	28	2	0.1738	0.6742	0.0250
CO2elect-GDP (cubic)– DOLS, K=1 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1003	0.8382	0.0250
	Fixed[1]	35	2	0.0669	0.9246	0.0250
	Fixed[2]	35	2	0.2222	0.5915	0.0250
Minimum volatility	Automatic	28	2	0.1475	0.7272	0.0250
	Fixed[1]	27	2	0.0868	0.8732	0.0250
	Fixed[2]	32	2	0.2089	0.6125	0.0250
CO2elect-GDP (cubic)– DOLS, K=2 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.0758	0.9018	0.0250
	Fixed[1]	35	2	0.0529	0.9581	0.0250
	Fixed[2]	35	2	0.4245	0.3738	0.0250
Minimum volatility	Automatic	32	2	0.1277	0.7713	0.0250
	Fixed[1]	27	2	0.0646	0.9305	0.0250
	Fixed[2]	27	2	0.4696	0.3417	0.0250
CO2elect-GDP (cubic)– DOLS, K=3 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1250	0.7776	0.0250
	Fixed[1]	35	2	0.0888	0.8679	0.0250
	Fixed[2]	35	2	0.2760	0.5174	0.0250
Minimum volatility	Automatic	27	2	0.3024	0.4863	0.0250
	Fixed[1]	26	2	0.1385	0.7467	0.0250
	Fixed[2]	25	3	0.8859	0.1638	0.0167

(1) Lag length for the long-run variance estimation: $[4(b/100)^{0.25}]$
(2) Lag length for the long-run variance estimation: $[12(b/100)^{0.25}]$

Table B.7: Results of tests for nonlinear cointegration (Choi and Saikkonen, 2010):

CO2transp-GDP (cubic) – NLLS residuals						
Rule	Bandwidth	Block size	M	$C_{NLLS}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1775	0.6672	0.0250
	Fixed[1]	35	2	0.1791	0.6642	0.0250
	Fixed[2]	35	2	0.2007	0.6261	0.0250
Minimum volatility	Automatic	31	2	0.1260	0.7751	0.0250
	Fixed[1]	24	3	0.2785	0.5143	0.0167
	Fixed[2]	31	2	0.0968	0.8473	0.0250
CO2transp-GDP (cubic)– DOLS, K=1 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.1006	0.8374	0.0250
	Fixed[1]	35	2	0.1067	0.8219	0.0250
	Fixed[2]	35	2	0.1083	0.8181	0.0250
Minimum volatility	Automatic	32	2	0.0947	0.8528	0.0250
	Fixed[1]	32	2	0.1008	0.8370	0.0250
	Fixed[2]	32	2	0.1052	0.8259	0.0250
CO2transp-GDP (cubic)– DOLS, K=2 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.0948	0.8524	0.0250
	Fixed[1]	35	2	0.0998	0.8396	0.0250
	Fixed[2]	35	2	0.0967	0.8474	0.0250
Minimum volatility	Automatic	31	2	0.0842	0.8800	0.0250
	Fixed[1]	32	2	0.1020	0.8339	0.0250
	Fixed[2]	31	2	0.0874	0.8716	0.0250
CO2transp-GDP (cubic)– DOLS, K=3 residuals						
Rule	Bandwidth	Block size	M	$C_{LL}^{b,max}$	$p-value$	5% level $\left(\frac{\alpha}{M}\right)$
Fixed block	Automatic	35	2	0.0872	0.8721	0.0250
	Fixed[1]	35	2	0.0838	0.8811	0.0250
	Fixed[2]	35	2	0.1004	0.8379	0.0250
Minimum volatility	Automatic	30	2	0.0850	0.8780	0.0250
	Fixed[1]	30	2	0.1125	0.8077	0.0250
	Fixed[2]	30	2	0.0907	0.8630	0.0250

(1) Lag length for the long-run variance estimation: $[4(b/100)^{0.25}]$
(2) Lag length for the long-run variance estimation: $[12(b/100)^{0.25}]$

Table B.8: Results of the rank test for neglected nonlinearity (Breitung, 2001).

	Rank test for neglected nonlinearity		
	CO2total-GDP	CO2elect-GDP	CO2transp-GDP
Score statistic	12.5373	8.5436	7.0949
	Critical values		
	Critical value 1%	Critical value 5%	Critical value 10%
χ^2 distribution with one degree of freedom	6.63	3.84	2.71

Table B.9: Results of Linearity test (Choi and Saikkonen, 2004):

	Linearity tests					
	T1μ			T2μ		
	CO2total-GDP	CO2elect-GDP	CO2transp-GDP	CO2total-GDP	CO2elect-GDP	CO2transp-GDP
K=1	3.2694 (0.0706)	0.2789 (0.5974)	7.3734 (0.0066)	13.3600 (0.0013)	12.2484 (0.0022)	7.4172 (0.0245)
K=2	3.3701 (0.0664)	0.0096 (0.9220)	8.1576 (0.0043)	18.2296 (0.0001)	15.5770 (0.0004)	8.1725 (0.0168)
K=3	0.6973 (0.4037)	1.9288 (0.1649)	6.4587 (0.0110)	13.0816 (0.0014)	14.1564 (0.0008)	6.5347 (0.0381)
	Critical values					
	Critical value 1%	Critical value 5%	Critical value 10%	Critical value 1%	Critical value 5%	Critical value 10%
χ^2 distribution	6.63	3.84	2.71	9.21	5.99	4.61

p-values in brackets.

Table B.10: Results of the Modified RESET test (Hong and Phillips, 2010).

	Modified RESET test					
	CO2total-GDP		CO2elect-GDP		CO2transp-GDP	
	Quadratic	Cubic	Quadratic	Cubic	Quadratic	Cubic
Test statistic	11.4290	12166.2176	2.7239	3371.6160	7.7541	8478.8054
$p - value$	0.0007	0.0000	0.0989	0.0000	0.0054	0.0000

APPENDIX C – NONLINEAR COINTEGRATION REGRESSIONS (REDUCED FORM MODELS)

Table C.1: Estimation of the long-run relationship between CO_2 emissions from electricity sector and per capita real GDP.

CO_2 emissions from electricity sector (levels)							
Estimator	Model 1.1 (quadratic)			Model 1.2 (cubic)			
	Constant	GDP$_t$	GDP$_t^2$	Constant	GDP$_t$	GDP$_t^2$	GDP $_t^3$
NLLS	-6.7713x10^{-1}	1.6779x10^{-4}	1.9129x10^{-9}	1.3826	-6.8244x10^{-4}	1.0380x10^{-7}	-3.6414x10^{-12}
DOLS, K=1	-7.5824x10^{-1}	1.6493x10^{-4}	2.5821x10^{-9}	1.5496	-7.2301x10^{-4}	1.0680x10^{-7}	-3.6905x10^{-12}
DOLS, K=2	-8.2887x10^{-1}	1.5544x10^{-4}	3.3418x10^{-9}	1.4889	-6.9585x10^{-4}	1.0252x10^{-7}	-3.5056x10^{-12}
DOLS, K=3	-8.7649x10^{-1}	1.4677x10^{-4}	3.9623x10^{-9}	1.2260	-5.9834x10^{-4}	8.9928x10^{-8}	-3.0277x10^{-12}
CO_2 emissions from electricity sector (logarithms)							
Estimator	Model 1.3 (quadratic – EKC hypothesis)			Model 1.4 (cubic)			
	Constant	GDP$_t$	GDP$_t^2$	Constant	GDP$_t$	GDP$_t^2$	GDP $_t^3$
NLLS	-5.7893x10^{+1}	1.0457x10^{+1}	-4.5199x10^{-1}	9.0263x10^{+2}	-3.1457x10^{+2}	3.6138x10^{+1}	-1.3704
DOLS, K=1	-6.5077x10^{+1}	1.2008x10^{+1}	-5.3520x10^{-1}	1.1589x10^{+3}	-3.9959x10^{+2}	4.5532x10^{+1}	-1.7159
DOLS, K=2	-7.3047x10^{+1}	1.3655x10^{+1}	-6.2055x10^{-1}	1.1740x10^{+3}	-4.0440x10^{+2}	4.6033x10^{+1}	-1.7331
DOLS, K=3	-8.6616x10^{+1}	1.6524x10^{+1}	-7.7266x10^{-1}	1.2303x10^{+3}	-4.2344x10^{+2}	4.8172x10^{+1}	-1.8129

Table C.2: Estimation of the long-run relationship between CO_2 emissions from transport sector and per capita real GDP.

CO_2 emissions from transport sector (levels)							
Estimator	Model 1.5 (quadratic)			Model 1.6 (cubic)			
	Constant	GDP$_t$	GDP$_t^2$	Constant	GDP$_t$	GDP$_t^2$	GDP $_t^3$
NLLS	7.5174x10^{-2}	2.6049x10^{-5}	5.7277x10^{-9}	-8.4559x10^{-2}	9.1985x10^{-5}	-2.1738x10^{-9}	2.8239x10^{-13}
DOLS, K=1	8.1074x10^{-2}	3.1415x10^{-5}	5.5018x10^{-9}	-4.6848x10^{-2}	8.0631x10^{-5}	-2.7490x10^{-10}	2.0456x10^{-13}
DOLS, K=2	7.8648x10^{-2}	3.4656x10^{-5}	5.3699x10^{-9}	-5.2657x10^{-2}	8.2883x10^{-5}	-2.4865x10^{-10}	1.9860x10-^{13}
DOLS, K=3	8.2531x10^{-2}	3.6795x10^{-5}	5.2796x10^{-9}	-8.3883x10^{-2}	9.5772x10^{-5}	-1.5248x10^{-9}	2.3965x10^{-13}
CO_2 emissions from transport sector (logarithms)							
Estimator	Model 1.7 (quadratic)			Model 1.8 (cubic)			
	Constant	GDP$_t$	GDP$_t^2$	Constant	GDP$_t$	GDP$_t^2$	GDP $_t^3$
NLLS	2.1661	-2.0164	1.9231x10^{-1}	1.0484x10^{+1}	-4.8311	5.0918x10^{-1}	-1.1867x10^{-2}
DOLS, K=1	2.2379	-1.9905	1.8911x10^{-1}	3.6583x10^{+1}	-1.3541x10^{+1}	1.4818	-4.8151x10^{-2}
DOLS, K=2	-1.5922x10^{-1}	-1.4437	1.5830x10^{-1}	-4.9836x10^{+1}	1.5209x10^{+1}	-1.7001	6.9037x10^{-2}
DOLS, K=3	-2.6625	-8.7670x10^{-1}	1.2649x10^{-1}	-1.8896x10^{+2}	6.1366x10^{+1}	-6.7977	2.5647x10^{-1}

Table C.3: Estimation of the long-run relationship between total CO_2 emissions and per capita real GDP.

Total CO_2 emissions (levels)							
Estimator	Model 1.9 (quadratic)			Model 1.10 (cubic)			
	Constant	GDP_t	GDP_t^2	Constant	GDP_t	GDP_t^2	GDP_t^3
NLLS	-5.0513×10^{-1}	4.0813×10^{-4}	1.1627×10^{-9}	1.5390	-4.3566×10^{-4}	1.0228×10^{-7}	-3.6138×10^{-12}
DOLS, K=1	-5.2558×10^{-1}	3.8929×10^{-4}	2.7814×10^{-9}	1.6233	-4.3748×10^{-4}	9.9822×10^{-8}	-3.4363×10^{-12}
DOLS, K=2	-5.2380×10^{-1}	3.6287×10^{-4}	4.4972×10^{-9}	1.5609	-4.0280×10^{-4}	9.3700×10^{-8}	-3.1530×10^{-12}
DOLS, K=3	-4.7289×10^{-1}	3.2967×10^{-4}	6.4582×10^{-9}	9.7654×10^{-1}	-1.8401×10^{-4}	6.5723×10^{-8}	-2.0873×10^{-12}
Total CO_2 emissions (logarithms)							
Estimator	Model 1.11 (quadratic)			Model 1.12 (cubic)			
	Constant	GDP_t	GDP_t^2	Constant	GDP_t	GDP_t^2	GDP_t^3
NLLS	-7.0455	5.8935×10^{-1}	3.4182×10^{-2}	$2.0926 \times 10^{+2}$	$-7.2606 \times 10^{+1}$	8.2742	-3.0861×10^{-1}
DOLS, K=1	-7.6253	7.0394×10^{-1}	2.8736×10^{-2}	$2.4336 \times 10^{+2}$	$-8.3702 \times 10^{+1}$	9.4755	-3.5187×10^{-1}
DOLS, K=2	-7.1986	5.8524×10^{-1}	3.6517×10^{-2}	$2.6838 \times 10^{+2}$	$-9.1796 \times 10^{+1}$	$1.0346 \times 10^{+1}$	-3.8298×10^{-1}
DOLS, K=3	-7.7120	6.6839×10^{-1}	3.3427×10^{-2}	$2.7124 \times 10^{+2}$	$-9.2528 \times 10^{+1}$	$1.0401 \times 10^{+1}$	-3.8401×10^{-1}

APPENDIX D – NONLINEAR COINTEGRATION REGRESSIONS (EXTENDED MODELS)

Table D.1: Estimation of the extended models for CO_2 emissions from electricity sector with additional control variables.

CO_2 emissions from electricity sector (levels)		
Dependent variable	Model 1.13	Model 1.14
Constant	-2.4678	1.2745
GDP_t	$1.5877\text{x}10^{-4}$	$-5.1862\text{x}10^{-4}$
GDP_t^2	$2.8636\text{x}10^{-9}$	$8.4392\text{x}10^{-8}$
GDP_t^3		$-2.9218\text{x}10^{-12}$
$Crude_t$	$-7.0482\text{x}10^{-3}$	$-4.0301\text{x}10^{-3}$
$Temp_t$	$1.3403\text{x}10^{-1}$	$-3.1133\text{x}10^{-3}$
$Precip_t$	$-1.5990\text{x}10^{-4}$	$-1.8693\text{x}10^{-4}$
CO_2 emissions from electricity sector (logarithms)		
Dependent variable	Model 1.15	Model 1.16
Constant	$-6.3109\text{x}10^{+1}$	$8.5481\text{x}10^{+2}$
GDP_t	$1.1262\text{x}10^{+1}$	$-2.9820\text{x}10^{+2}$
GDP_t^2	$-4.9985\text{x}10^{-1}$	$3.4269\text{x}10^{+1}$
GDP_t^3		-1.2994
$Crude_t$	$-2.3793\text{x}10^{-3}$	$-1.8067\text{x}10^{-3}$
$Temp_t$	$1.3148\text{x}10^{-1}$	$2.5494\text{x}10^{-2}$
$Precip_t$	$-1.7288\text{x}10^{-4}$	$-2.1436\text{x}10^{-4}$

Table D.2: Estimation of the extended models for CO2 emissions from transport sector with additional control variables.

CO_2 emissions from transport sector (levels)				
Dependent variable	Model 1.17	Model 1.18	Model 1.19	Model 1.20
Constant	$4.7027\text{x}10^{-2}$	$8.0895\text{x}10^{-2}$	$-5.0861\text{x}10^{-2}$	$3.5200\text{x}10^{-3}$
GDP_t	$3.6108\text{x}10^{-5}$	$3.6928\text{x}10^{-5}$	$7.7315\text{x}10^{-5}$	$6.7325\text{x}10^{-5}$
GDP_t^2	$5.3820\text{x}10^{-9}$	$5.2477\text{x}10^{-9}$	$2.9032\text{x}10^{-10}$	$1.4007\text{x}10^{-9}$
GDP_t^3			$1.8516\text{x}10^{-13}$	$1.4142\text{x}10^{-13}$
$Crude_t$	$-3.2650\text{x}10^{-4}$		$-3.9788\text{x}10^{-4}$	
$Fuel_t$		$-3.3885\text{x}10^{-2}$		$-2.9172\text{x}10^{-2}$
$d(motor)_t$	$-2.0666\text{x}10^{-3}$	$-1.9488\text{x}10^{-3}$	$-1.5758\text{x}10^{-3}$	$-1.4301\text{x}10^{-3}$
CO_2 emissions from transport sector (logarithms)				
Dependent variable	Model 1.21	Model 1.22	Model 1.23	Model 1.24
Constant	1.8780	2.7069	$4.4975\text{x}10^{+1}$	$3.9714\text{x}10^{+1}$
GDP_t	-1.9412	-2.1593	$-1.6529\text{x}10^{+1}$	$-1.4712\text{x}10^{+1}$
GDP_t^2	$1.8736\text{x}10^{-1}$	$2.0120\text{x}10^{-1}$	1.8298	1.6180
GDP_t^3			$-6.1506\text{x}10^{-2}$	$-5.3195\text{x}10^{-2}$
$Crude_t$	$6.7111\text{x}10^{-4}$		$6.3274\text{x}10^{-4}$	
$Fuel_t$		$3.1565\text{x}10^{-2}$		$1.9875\text{x}10^{-2}$
$d(motor)_t$	$-5.8334\text{x}10^{-4}$	$-1.3154\text{x}10^{-3}$	$-8.4825\text{x}10^{-4}$	$-1.6100\text{x}10^{-3}$

Table D.3: Estimation of the extended models for total CO_2 emissions with additional control variables.

Total CO_2 emissions (levels)				
Dependent variable	Model 1.25	Model 1.26	Model 1.27	Model 1.28
Constant	-1.6906	-1.6076	-1.5508×10^{-1}	8.0667×10^{-1}
GDP_t	3.8810×10^{-4}	3.5510×10^{-4}	1.0080×10^{-4}	-9.4569×10^{-5}
GDP_t^2	2.9050×10^{-9}	2.2625×10^{-9}	3.8562×10^{-8}	5.8791×10^{-8}
GDP_t^3			-1.2912×10^{-12}	-2.0535×10^{-12}
$Crude_t$	-8.7489×10^{-3}		-8.2348×10^{-3}	
$Fuel_t$		-3.8653×10^{-1}		-3.8950×10^{-1}
$d(motor)_t$	5.1898×10^{-3}	1.4257×10^{-2}	2.5986×10^{-3}	8.7687×10^{-3}
$Temp_t$	9.4039×10^{-2}	1.1348×10^{-1}	4.0000×10^{-2}	2.9385×10^{-2}
$Precip_t$	-1.0858×10^{-4}	-2.3482×10^{-5}	-1.3190×10^{-4}	-7.1934×10^{-5}
Total CO_2 emissions (logarithms)				
Dependent variable	Model 1.29	Model 1.30	Model 1.31	Model 1.32
Constant	-8.2693	-9.0766	$1.3230 \times 10^{+2}$	$1.4630 \times 10^{+2}$
GDP_t	7.6131×10^{-1}	9.8614×10^{-1}	$-4.6652 \times 10^{+1}$	$-5.1549 \times 10^{+1}$
GDP_t^2	2.4447×10^{-2}	9.0718×10^{-3}	5.3521	5.9266
GDP_t^3			-1.9910×10^{-1}	-2.2171×10^{-1}
$Crude_t$	-1.2345×10^{-3}		-1.3534×10^{-3}	
$Fuel_t$		-1.8759×10^{-2}		-5.7801×10^{-2}
$d(motor)_t$	1.3272×10^{-3}	3.0802×10^{-3}	6.9128×10^{-4}	2.1221×10^{-3}
$Temp_t$	3.1964×10^{-2}	3.1284×10^{-2}	1.7583×10^{-2}	-5.7801×10^{-2}
$Precip_t$	-1.9077×10^{-5}	-4.1218×10^{-6}	-2.8413×10^{-5}	-1.5884×10^{-5}

APPENDIX E – COINTEGRATION TESTS WITH STRUCTURAL BREAKS

Table E.1: Results of the Gregory and Hansen cointegration tests (Gregory and Hansen, 1996).

Test statistics	**CO2total-GDP (levels)**		**CO2elect-GDP (levels)**		**CO2transp-GDP (levels)**	
	Level shift	**Regime shift**	**Level shift**	**Regime shift**	**Level shift**	**Regime shift**
ADF*	-3.9712$^{\#}$	-4.4901$^{\#}$	-4.8786^{b,c}	-4.8641^{c}	-4.2509$^{\#}$	-4.4000$^{\#}$
Breakpoint	0.5686	0.5882	0.5490	0.5490	0.7059	0.5294
Break date	1988	1989	1987	1987	1995	1986
Zt*	-2.8566$^{\#}$	-4.2067$^{\#}$	-4.9281^{b,c}	-5.1571^{b,c}	-3.8622$^{\#}$	-3.8446$^{\#}$
Breakpoint	0.8235	0.8235	0.5490	0.5882	0.7255	0.6667
Break date	2001	2001	1987	1989	1996	1993
$Z\alpha^{*}$	-18.4256$^{\#}$	-29.9860$^{\#}$	-39.0571^{c}	-40.3964$^{\#}$	-23.8740$^{\#}$	-26.1466$^{\#}$
Breakpoint	0.8235	0.8235	0.5490	0.5882	0.7255	0.5490
Break date	2001	2001	1987	1989	1996	1987
Test statistics	**CO2total-GDP (logarithms)**		**CO2elect-GDP (logarithms)**		**CO2transp-GDP (logarithms)**	
	Level shift	**Regime shift**	**Level shift**	**Regime shift**	**Level shift**	**Regime shift**
ADF*	-3.7089$^{\#}$	-4.4336$^{\#}$	-5.2215	-4.9365^{b,c}	-3.5037$^{\#}$	-3.5121$^{\#}$
Breakpoint	0.8431	0.8235	0.7647	0.4706	0.6667	0.2745
Break date	2002	2001	1998	1983	1993	1973
Zt*	-3.7466$^{\#}$	-4.4786$^{\#}$	-5.3905	-5.6618	-3.6497$^{\#}$	-3.6388$^{\#}$
Breakpoint	0.8431	0.8235	0.7255	0.5882	0.6863	0.1569
Break date	2002	2001	1996	1989	1994	1967
$Z\alpha^{*}$	-25.5202$^{\#}$	-31.6574$^{\#}$	-39.8886^{c}	-41.7612$^{\#}$	-22.2705$^{\#}$	-22.4678$^{\#}$
Breakpoint	0.8431	0.8235	0.7255	0.5882	0.7059	0.5490
Break date	2002	2001	1996	1989	1995	1987

The critical values are taken from Gregory and Hansen (1996), table 1, m=1.
a, b and c denote the rejection of the null hypothesis of no cointegration at 1%, 5% and 10% levels of significance, respectively.
# denotes the acceptance of the null hypothesis of no cointegration.

Table E.2: Results of the Arai and Kurozumi cointegration tests (Arai and Kurozumi, 2007) with Kejriwal (2008) extension for CO2elect-GDP relationship.

CO2elect-GDP (levels)			
m=1			
$\tilde{V}_1(\hat{\tau}_1)$	$\hat{\tau}_1$	-	-
0.0650	0.5652	-	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.2040	0.1340	0.1030	
m=2			
$\tilde{V}_2(\hat{\tau}_2)$	$\hat{\tau}_1$	$\hat{\tau}_2$	-
0.0698	0.5652	0.7826	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.1780	0.1100	0.0860	
m=3			
$\tilde{V}_3(\hat{\tau}_3)$	$\hat{\tau}_1$	$\hat{\tau}_2$	$\hat{\tau}_3$
0.0424	0.3696	0.5652	0.7826
Critical value 1%	Critical value 5%	Critical value 10%	
0.0890	0.0610	0.0520	
CO2elect-GDP (logarithms)			
m=1			
$\tilde{V}_1(\hat{\tau}_1)$	$\hat{\tau}_1$	-	-
0.0360	0.5652	-	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.2040	0.1340	0.1030	
m=2			
$\tilde{V}_2(\hat{\tau}_2)$	$\hat{\tau}_1$	$\hat{\tau}_2$	-
0.0299	0.1522	0.5652	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.1430	0.0950	0.0740	
m=3			
$\tilde{V}_3(\hat{\tau}_3)$	$\hat{\tau}_1$	$\hat{\tau}_2$	$\hat{\tau}_3$
0.0269	0.1522	0.3696	0.5652
Critical value 1%	Critical value 5%	Critical value 10%	
0.1120	0.0720	0.0570	

m: number of breaks

Table E.3: Results of the Arai and Kurozumi cointegration tests (Arai and Kurozumi, 2007) with Kejriwal (2008) extension for CO2transp-GDP relationship.

CO2transp-GDP (levels)			
m=1			
$\tilde{V}_1(\hat{\tau}_1)$	$\hat{\tau}_1$	-	-
0.0497	0.6304	-	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.2380	0.1420	0.1150	
m=2			
$\tilde{V}_2(\hat{\tau}_2)$	$\hat{\tau}_1$	$\hat{\tau}_2$	-
0.0394	0.6304	0.8696	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.2330	0.1290	0.1030	
m=3			
$\tilde{V}_3(\hat{\tau}_3)$	$\hat{\tau}_1$	$\hat{\tau}_2$	$\hat{\tau}_3$
0.0551	0.4783	0.6739	0.8696
Critical value 1%	Critical value 5%	Critical value 10%	
0.1420	0.0840	0.0630	
CO2transp-GDP (logarithms)			
m=1			
$\tilde{V}_1(\hat{\tau}_1)$	$\hat{\tau}_1$	-	-
0.0520	0.4783	-	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.1900	0.1340	0.1040	
m=2			
$\tilde{V}_2(\hat{\tau}_2)$	$\hat{\tau}_1$	$\hat{\tau}_2$	-
0.0542	0.1522	0.4565	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.1800	0.1120	0.0840	
m=3			
$\tilde{V}_3(\hat{\tau}_3)$	$\hat{\tau}_1$	$\hat{\tau}_2$	$\hat{\tau}_3$
0.0567	0.1522	0.2826	0.6739
Critical value 1%	Critical value 5%	Critical value 10%	
0.1130	0.0740	0.0600	

m: number of breaks

Table E.4: Results of the Arai and Kurozumi cointegration tests (Arai and Kurozumi, 2007) with Kejriwal (2008) extension for CO2total-GDP relationship.

CO2total-GDP (levels)			
m=1			
$\tilde{V}_1(\hat{\tau}_1)$	$\hat{\tau}_1$	-	-
0.1925[c]	0.7826	-	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.3520	0.2040	0.1490	
m=2			
$\tilde{V}_2(\hat{\tau}_2)$	$\hat{\tau}_1$	$\hat{\tau}_2$	-
0.0365	0.5652	0.8696	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.1770	0.1150	0.0890	
m=3			
$\tilde{V}_3(\hat{\tau}_3)$	$\hat{\tau}_1$	$\hat{\tau}_2$	$\hat{\tau}_3$
0.0214	0.5652	0.6957	0.8696
Critical value 1%	Critical value 5%	Critical value 10%	
0.1750	0.1060	0.0820	
CO2total-GDP (logarithms)			
m=1			
$\tilde{V}_1(\hat{\tau}_1)$	$\hat{\tau}_1$	-	-
0.1439	0.7826	-	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.3520	0.2040	0.1490	
m=2			
$\tilde{V}_2(\hat{\tau}_2)$	$\hat{\tau}_1$	$\hat{\tau}_2$	-
0.0298	0.5652	0.8696	-
Critical value 1%	Critical value 5%	Critical value 10%	
0.1770	0.1150	0.0890	
m=3			
$\tilde{V}_3(\hat{\tau}_3)$	$\hat{\tau}_1$	$\hat{\tau}_2$	$\hat{\tau}_3$
0.0243	0.1304	0.5652	0.8696
Critical value 1%	Critical value 5%	Critical value 10%	
0.1200	0.0760	0.0620	

a, b and c denote the rejection of the null hypothesis of no cointegration at 1%, 5% and 10% levels of significance, respectively.
m: number of breaks

APPENDIX F – STRUCTURAL BREAKS TESTS

Table F.1: Results of the Kejriwal-Perron tests for multiple structural breaks (Kejriwal and Perron; 2008, 2010) for CO2elect-GDP relationship.

	CO2elect-GDP (levels)		
	m=1	m=2	m=3
Test statistics			
$SupF^{*}(1)$	14.267366^{b,c}	14.267366^{b,c}	14.267366 ^{b,c}
$SupF^{*}(2)$	11.562215^{b,c}	11.562215^{b,c}	11.562215 ^{b,c}
$SupF^{*}(3)$	8.457661^{c}	8.457661^{c}	8.457661^{c}
$SupF^{*}(4)$	6.452927$^{\#}$	6.452927$^{\#}$	6.452927$^{\#}$
$SupF^{*}(5)$	5.269597$^{\#}$	5.269597$^{\#}$	5.269597$^{\#}$
$UDmax$	14.267366^{b,c}	14.267366^{b,c}	14.267366^{b,c}
Number of breaks selected	SEQ 1	BIC 2	LWZ 2
Break dates			
$\hat{T}_1$	1999	1992	1985
$\hat{T}_2$	-	1995	1985
$\hat{T}_3$	-	-	1995
	CO2elect-GDP (logarithms)		
	m=1	m=2	m=3
Test statistics			
$SupF^{*}(1)$	19.022583	19.022583	15.502329^{b,c}
$SupF^{*}(2)$	11.097258^{b,c}	11.097258^{b,c}	11.097258^{b,c}
$SupF^{*}(3)$	6.956278$^{\#}$	6.956278$^{\#}$	6.956278$^{\#}$
$SupF^{*}(4)$	5.147810$^{\#}$	5.147810$^{\#}$	5.147810$^{\#}$
$SupF^{*}(5)$	4.394657$^{\#}$	4.394657$^{\#}$	4.394657$^{\#}$
$UDmax$	19.022583	19.022583	15.502329
Number of breaks selected	SEQ 1	BIC 1	LWZ 1
Break dates			
$\hat{T}_1$	1995	1985	1979
$\hat{T}_2$	-	1995	1985
$\hat{T}_3$	-	-	1985

The critical values are taken from Kejriwal and Perron (2010), table 1, non-trending case, q_b=1.
a, b and c denote the rejection of the null hypothesis of no structural breaks at 1%, 5% and 10% levels of significance, respectively.
# denotes the acceptance of the null hypothesis of no structural breaks.
m: number of breaks

Table F.2: Results of the Kejriwal-Perron tests for multiple structural breaks (Kejriwal and Perron; 2008, 2010) for CO2transp-GDP relationship.

	CO2transp-GDP (levels)		
	m=1	m=2	m=3
Test statistics			
$SupF^{*}(1)$	9.640715[#]	9.640715[#]	8.791025[#]
$SupF^{*}(2)$	6.646609[#]	6.646609[#]	6.646609[#]
$SupF^{*}(3)$	7.581012[#]	7.581012[#]	7.581012[#]
$SupF^{*}(4)$	6.733361[c]	6.733361[c]	6.733361[c]
$SupF^{*}(5)$	4.703901[#]	4.703901[#]	4.703901[#]
$UDmax$	9.640715[#]	9.640715[#]	8.791025[#]
Number of breaks selected	SEQ 0	BIC 4	LWZ 2
Break dates			
$\hat{T}_1$	1999	1993	1987
$\hat{T}_2$	-	1999	1988
$\hat{T}_3$	-	-	1999
	CO2transp-GDP (logarithms)		
	m=1	m=2	m=3
Test statistics			
$SupF^{*}(1)$	12.934071[b,c]	12.934071[b,c]	11.782188[c]
$SupF^{*}(2)$	8.446639[#]	8.446639[#]	8.446639[#]
$SupF^{*}(3)$	8.307063[#]	8.307063[#]	8.307063[#]
$SupF^{*}(4)$	5.638052[#]	5.638052[#]	5.638052[#]
$SupF^{*}(5)$	4.569566[#]	4.569566[#]	4.569566[#]
$UDmax$	12.934071[b,c]	12.934071[b,c]	11.782188[c]
Number of breaks selected	SEQ 1	BIC 2	LWZ 2
Break dates			
$\hat{T}_1$	1998	1988	1980
$\hat{T}_2$	-	1990	1981
$\hat{T}_3$	-	-	1990

The critical values are taken from Kejriwal and Perron (2010), table 1, non-trending case, q_b=1.
a, b and c denote the rejection of the null hypothesis of no structural breaks at 1%, 5% and 10% levels of significance, respectively.
# denotes the acceptance of the null hypothesis of no structural breaks.
m: number of breaks

Table F.3: Results of the Kejriwal-Perron tests for multiple structural breaks (Kejriwal and Perron; 2008, 2010) for CO2total-GDP relationship.

	CO2total-GDP (levels)		
	m=1	m=2	m=3
Test statistics			
$SupF^{*}(1)$	14.496050[b,c]	14.496050[b,c]	14.496050[b,c]
$SupF^{*}(2)$	12.003037[b,c]	12.003037[b,c]	12.003037[b,c]
$SupF^{*}(3)$	8.340191[#]	8.340191[#]	8.340191[#]
$SupF^{*}(4)$	6.003618[#]	6.003618[#]	6.003618[#]
$SupF^{*}(5)$	4.986297[#]	4.986297[#]	4.986297[#]
$UDmax$	14.496050[b,c]	14.496050[b,c]	14.496050[b,c]
Number of breaks selected	SEQ 2	BIC 2	LWZ 2
Break dates			
$\hat{T}_1$	1999	1991	1985
$\hat{T}_2$	-	1999	1989
$\hat{T}_3$	-	-	1999
	CO2total-GDP (logarithms)		
	m=1	m=2	m=3
Test statistics			
$SupF^{*}(1)$	10.290950[#]	10.290950[#]	9.135424[#]
$SupF^{*}(2)$	13.446389	13.446389	13.446389
$SupF^{*}(3)$	8.087853[#]	8.087853[#]	8.087853[#]
$SupF^{*}(4)$	6.297442[#]	6.297442[#]	6.297442[#]
$SupF^{*}(5)$	5.051530[#]	5.051530[#]	5.051530[#]
$UDmax$	13.446389[b,c]	13.446389[b,c]	13.446389[b,c]
Number of breaks selected	SEQ 0	BIC 2	LWZ 2
Break dates			
$\hat{T}_1$	1999	1991	1985
$\hat{T}_2$	-	1999	1985
$\hat{T}_3$	-	-	1999

The critical values are taken from Kejriwal and Perron (2010), table 1, non-trending case, q_b=1.
a, b and c denote the rejection of the null hypothesis of no structural breaks at 1%, 5% and 10% levels of significance, respectively.
# denotes the acceptance of the null hypothesis of no structural breaks.
m: number of breaks

APPENDIX G – COINTEGRATING REGRESSIONS WITH STRUCTURAL BREAKS (REDUCED-FORM MODELS)

Table G.1: Estimation of the long-run relationship between CO_2 emissions from electricity sector and per capita real GDP under breaks.

	CO2elect-GDP (levels): Model 2.1		
	m=1	m=2	m=3
Parameter estimates			
α_1	-0.645384 (0.037799)	-0.582032 (0.037909)	-0.299564 (0.047282)
β_1	0.000162 (0.043098)	0.000160 (0.061128)	0.000112 (0.064983)
α_2	0.061975 (0.000022)	0.855048 (0.061128)	-0.766919 (0.061648)
β_2	0.000148 (0.000027)	0.000087 (0.000022)	0.000192 (0.061648)
α_3	-	9.640866 (0.000065)	0.861130 (0.000036)
β_3	-	-0.000485 (0.000198)	0.000089 (0.000121)
α_4	-	-	10.230584 (0.000066)
β_4	-	-	-0.000524 (0.000199)
R^2	1: 0.882017 2: 0.546320 - -	1: 0.882017 2: 0.325457 3: 0.432760 -	1: 0.844976 2: 0.437650 3: 0.325457 4: 0.432760
Break dates			
$\hat{T}_1$	1999	1992	1985
$\hat{T}_2$	-	1995	1985
$\hat{T}_3$	-	-	1995
	CO2elect-GDP (logarithms): Model 2.2		
	m=1	m=2	m=3
Parameter estimates			
α_1	-24.515779 (0.047191)	-15.418090 (0.081273)	-15.615839 (0.087108)
β_1	2.684756 (0.053806)	1.589827 (0.049331)	1.616747 (0.072880)
α_2	-6.667224 (0.172282)	-24.001737 (0.048082)	-16.681429 (0.076822)
β_2	0.781567 (0.440075)	2.625497 (0.701200)	1.793898 (0.051534)
α_3	-	-7.525888 (0.368016)	-16.987726 (0.751547)
β_3	-	0.869861 (0.393258)	1.856486 (0.822798)
α_4	-	-	-7.053330 (1.264428)
β_4	-	-	0.820874 (0.421495)
R^2	1: 0.882017 2: 0.546320 - -	1: 0.429663 2: 0.795655 3: 0.584673 -	1: 0.429663 2: 0.359215 3: 0.449321 4: 0.584673
Break dates			
$\hat{T}_1$	1995	1985	1979
$\hat{T}_2$	-	1995	1985
$\hat{T}_3$	-	-	1985

α_i:intercept; β_i: estimated slope
Standard errors in brackets.
1, 2, 3 and 4 denote the R^2 of the first, second, third and fourth regime, respectively.
m: number of breaks

Table G.2: Estimation of the long-run relationship between CO_2 emissions from transport sector and per capita real GDP under breaks.

	CO2transp-GDP (levels): Model 2.3		
	m=1	m=2	m=3
Parameter estimates			
α_1	-0.174495 (0.026386)	-0.168739 (0.022519)	-0.181729 (0.020145)
β_1	0.000109 (0.034464)	0.000108 (0.036565)	0.000111 (0.031496)
α_2	-1.028969 (0.000013)	-1.040976 (0.049509)	-0.480038 (0.031496)
β_2	0.000189 (0.000026)	0.000189 (0.000011)	0.000139 (0.038575)
α_3	-	2.507316 (0.000030)	-0.970436 (0.000013)
β_3	-	-0.000044 (0.000192)	0.000185 (0.000023)
α_4	-	-	2.321124 (0.000028)
β_4	-	-	-0.000032 (0.000150)
R^2	1: 0.978783	1: 0.978783	1: 0.962596
	2: 0.953321	2: 0.981599	2: 0.963013
	-	3: 0.454645	3: 0.991118
	-	-	4: 0.454645
Break dates			
$\hat{T}_1$	1999	1993	1987
$\hat{T}_2$	-	1999	1988
$\hat{T}_3$	-	-	1999
	CO2transp-GDP (logarithms): Model 2.4		
	m=1	m=2	m=3
Parameter estimates			
α_1	-12.973091 (0.014923)	-9.448969 (0.025080)	-9.591185 (0.021437)
β_1	1.406282 (0.013677)	0.981282 (0.017734)	1.000662 (0.023154)
α_2	-15.356032 (0.057894)	-14.116602 (0.013271)	-7.833657 (0.013368)
β_2	1.658818 (0.066626)	1.533915 (0.216381)	0.817425 (0.014644)
α_3	-	-15.687883 (0.159578)	-11.088416 (0.184949)
β_3	-	1.692663 (0.064647)	1.195196 (0.252280)
α_4	-	-	-13.779608 (0.074725)
β_4	-	-	1.494025 (0.178518)
R^2	1: 0.968264	1: 0.848931	1: 0.848931
	2: 0.980974	2: 0.864206	2: 0.691379
	-	3: 0.980974	3: 0.965982
	-	-	4: 0.949419
Break dates			
$\hat{T}_1$	1998	1987	1980
$\hat{T}_2$	-	1990	1981
$\hat{T}_3$	-	-	1990

α_i:intercept; β_i: estimated slope
Standard errors in brackets.
1, 2, 3 and 4 denote the R^2 of the first, second, third and fourth regime, respectively.
m: number of breaks

Table G.3: Estimation of the long-run relationship between total CO_2 emissions and per capita real GDP under breaks.

	m=1	m=2	m=3
CO2total-GDP (levels): Model 2.5			
Parameter estimates			
α_1	-0.804008 (0.041963)	-0.271961 (0.050472)	-0.255162 (0.050169)
β_1	0.000456 (0.079621)	0.000385 (0.068783)	0.000385 (0.104437)
α_2	15.737692 (0.000015)	-1.161929 (0.105068)	0.721010 (0.090445)
β_2	-0.000647 (0.000257)	0.000507 (0.000030)	0.000344 (0.104437)
α_3	-	13.524797 (0.000047)	-1.079606 (0.000030)
β_3	-	-0.000506 (0.000408)	0.000503 (0.000180)
α_4	-	-	13.720393 (0.000089)
β_4	-	-	-0.000519 (0.000405)
R^2	1: 0.984513	1: 0.976839	1: 0.976839
	2: 0.463323	2: 0.953155	2: 0.837248
	-	3: 0.444511	3: 0.869702
	-	-	4: 0.444511
Break dates			
$\hat{T}_1$	1999	1991	1985
$\hat{T}_2$	-	1999	1989
$\hat{T}_3$	-	-	1999
CO2total-GDP (logarithms): Model 2.6			
	m=1	m=2	m=3
Parameter estimates			
α_1	-10.308348 (0.009885)	-9.082210 (0.009930)	-6.882471 (0.022482)
β_1	1.262903 (0.025522)	1.125695 (0.013533)	0.861669 (0.012314)
α_2	11.280413 (0.025034)	-9.589564 (0.020671)	-9.111514 (0.014718)
β_2	-0.988605 (1.505968)	1.191616 (0.036253)	1.128876 (0.022482)
α_3	-	12.691886 (0.119303)	-9.627439 (0.212128)
β_3	-	-1.134942 (1.219730)	1.195352 (0.078813)
α_4	-	-	13.107528 (0.129752)
β_4	-	-	-1.178157 (1.326560)
R^2	1: 0.991378	1: 0.982188	1: 0.975179
	2: 0.449007	2: 0.955342	2: 0.942558
	-	3: 0.449007	3: 0.955342
	-	-	4: 0.449007
Break dates			
$\hat{T}_1$	1999	1991	1985
$\hat{T}_2$	-	1999	1985
$\hat{T}_3$	-	-	1999

α_i:intercept; β_i: estimated slope
Standard errors in brackets.
1, 2, 3 and 4 denote the R^2 of the first, second, third and fourth regime, respectively.
m: number of breaks

APPENDIX H – COINTEGRATING REGRESSIONS WITH STRUCTURAL BREAKS (EXTENDED MODELS)

Table H.1: Estimation of the extended models for CO_2 emissions from electricity sector with crude, temperature and precipitation as control variables.

CO2elect-GDP (levels): Model 2.7			
Parameter estimates	m=1	m=2	m=3
α_1	-1.003307	-1.010746	0.172890
β_1	0.000180	0.000164	0.000126
α_2	-0.349213	0.625667	0.316466
β_2	0.000161	0.000069	0.000137
α_3	-	8.224410	1.593420
β_3	-	-0.000417	0.000078
α_4	-	-	9.114695
β_4	-	-	-0.000406
$Crude_t$	-0.003417	-0.001987	-0.002654
$Temp_t$	0.028987	0.038757	-0.027121
$Precip_t$	-0.000087	-0.000147	-0.000094
R^2	1: 0.882017	1: 0.882017	1: 0.844976
	2: 0.546320	2: 0.325457	2: 0.437650
	-	3: 0.432760	3: 0.325457
	-	-	4: 0.432760
Break dates			
$\hat{T}_1$	1999	1992	1985
$\hat{T}_2$	-	1995	1985
$\hat{T}_3$	-	-	1994
CO2elect-GDP (logarithms) Model 2.8			
Parameter estimates	m=1	m=2	m=3
α_1	-23.844171	-13.563613	-13.474190
β_1	2.673911	1.467540	1.491463
α_2	-5.608643	-24.349922	-20.138207
β_2	0.732140	2.755864	2.309544
α_3	-	-6.312362	-12.355239
β_3	-	0.830830	1.473160
α_4	-	-	-7.087918
β_4	-	-	0.949132
$Crude_t$	-0.000612	-0.001081	-0.002243
$Temp_t$	-0.026866	-0.039973	-0.062064
$Precip_t$	-0.000148	-0.000171	-0.000107
R^2	1: 0.932088	1: 0.429663	1: 0.429663
	2: 0.584673	2: 0.795655	2: 0.359215
	-	3: 0.584673	3: 0.449321
	-	-	4: 0.584673
Break dates			
$\hat{T}_1$	1995	1985	1979
$\hat{T}_2$	-	1995	1985
$\hat{T}_3$	-	-	1985

α_i:intercept; β_i: estimated slope
1, 2, 3 and 4 denote the R^2 of the first, second, third and fourth regime, respectively.
m: number of breaks

Table H.2: Estimation of the extended models for CO_2 emissions from transport sector with crude and d(motor) as control variables.

CO2transp-GDP (levels): Model 2.9			
Parameter estimates	m=1	m=2	m=3
α_1	-0.173932	-0.160995	-0.163568
β_1	0.000111	0.000104	0.000106
α_2	-1.071148	-1.069835	-0.481563
β_2	0.000193	0.000189	0.000137
α_3	-	2.999428	-0.959639
β_3	-	-0.000080	0.000183
α_4	-	-	2.782267
β_4	-	-	-0.000065
$Crude_t$	-0.000433	0.000592	0.000477
$d(motor)_t$	0.000242	0.000821	0.000331
R^2	1: 0.978783	1: 0.978783	1: 0.962596
	2: 0.953321	2: 0.981599	2: 0.963013
	-	3: 0.454645	3: 0.991118
	-	-	4: 0.454645
Break dates			
$\hat{T}_1$	1999	1993	1987
$\hat{T}_2$	-	1999	1988
$\hat{T}_3$	-	-	1999
CO2transp-GDP (logarithms): Model 2.10			
Parameter estimates	m=1	m=2	m=3
α_1	-13.042709	-9.528358	-9.272169
β_1	1.414292	0.991377	0.963920
α_2	-15.431602	-14.480930	-7.466335
β_2	1.667527	1.576353	0.776678
α_3	-	-15.818410	-10.509128
β_3	-	1.708476	1.133042
α_4	-	-	-15.104724
β_4	-	-	1.633428
$Crude_t$	-0.000045	-0.000316	-0.000528
$d(motor)_t$	-0.000426	-0.000445	0.000979
R^2	1: 0.968264	1: 0.848931	1: 0.848931
	2: 0.980974	2: 0.864206	2: 0.691379
	-	3: 0.980974	3: 0.965982
	-	-	4: 0.949419
Break dates			
$\hat{T}_1$	1998	1988	1980
$\hat{T}_2$	-	1990	1981
$\hat{T}_3$	-	-	1990

α_i:intercept; β_i: estimated slope
1, 2, 3 and 4 denote the R^2 of the first, second, third and fourth regime, respectively.
m: number of breaks

Table H.3: Estimation of the extended models for CO_2 emissions from transport sector with fuel and d(motor) as control variables.

CO2transp-GDP (levels): Model 2.11			
Parameter estimates	m=1	m=2	m=3
α_1	-0.120934	-0.174747	-0.199102
β_1	0.000111	0.000107	0.000109
α_2	-0.993106	-1.078712	-0.511381
β_2	0.000190	0.000191	0.000139
α_3	-	2.535337	-0.995630
β_3	-	-0.000047	0.000185
α_4	-	-	2.466917
β_4	-	-	-0.000043
$Fuel_t$	-0.051025	0.009455	0.021986
$d(motor)_t$	0.000344	0.000554	0.000224
R^2	1: 0.978783 2: 0.953321 - -	1: 0.978783 2: 0.981599 3: 0.454645 -	1: 0.962596 2: 0.963013 3: 0.991118 4: 0.454645
Break dates			
$\hat{T}_1$	1999	1993	1987
$\hat{T}_2$	-	1999	1988
$\hat{T}_3$	-	-	1999
CO2transp-GDP (logarithms): Model 2.12			
Parameter estimates	m=1	m=2	m=3
α_1	-13.027554	-9.332892	-8.909272
β_1	1.411407	0.973796	0.939079
α_2	-15.472703	-14.694478	-6.087808
β_2	1.670774	1.605483	0.635207
α_3	-	-15.377279	-10.068204
β_3	-	1.665598	1.102899
α_4	-	-	-14.038813
β_4	-	-	1.534907
$Fuel_t$	0.006257	-0.045363	-0.134780
$d(motor)_t$	-0.000331	-0.000399	0.000190
R^2	1: 0.968264 2: 0.980974 - -	1: 0.848931 2: 0.864206 3: 0.980974 -	1: 0.848931 2: 0.691379 3: 0.965982 4: 0.949419
Break dates			
$\hat{T}_1$	1998	1988	1980
$\hat{T}_2$	-	1990	1981
$\hat{T}_3$	-	-	1990

α_i:intercept; β_i: estimated slope
1, 2, 3 and 4 denote the R^2 of the first, second, third and fourth regime, respectively.
m: number of breaks

Table H.4: Estimation of the extended models for total CO_2 emissions with crude, d(motor), temperature and precipitation as control variables.

CO2total-GDP (levels): Model 2.13			
Parameter estimates	m=1	m=2	m=3
α_1	-2.013702	0.491623	0.857456
β_1	0.000445	0.000367	0.000369
α_2	10.095299	-0.519660	2.613734
β_2	-0.000354	0.000515	0.000269
α_3	-	16.805172	0.028508
β_3	-	-0.000685	0.000499
α_4	-	-	17.111209
β_4	-	-	-0.000681
$Crude_t$	-0.005767	0.001272	0.000943
$d(motor)_t$	0.005567	-0.005701	0.000789
$Temp_t$	0.091641	-0.025294	-0.049665
$Precip_t$	0.000056	-0.000292	-0.000279
R^2	1: 0.984513	1: 0.976839	1: 0.976839
	2: 0.463323	2: 0.953155	2: 0.837248
	-	3: 0.444511	3: 0.869702
	-	-	4: 0.444511
Break dates			
$\hat{T}_1$	1999	1991	1971
$\hat{T}_2$	-	1999	1981
$\hat{T}_3$	-	-	1991

CO2total-GDP (logarithms): Model 2.14			
Parameter estimates	m=1	m=2	m=3
α_1	-10.212948	-8.835887	-6.765544
β_1	1.220950	1.105752	0.851249
α_2	0.673722	-9.665264	-8.826011
β_2	0.084543	1.209822	1.099020
α_3	-	20.215112	-9.763089
β_3	-	-1.911122	1.213398
α_4	-	-	18.313010
β_4	-	-	-1.718818
$Crude_t$	-0.000767	0.000434	0.000263
$d(motor)_t$	0.001067	-0.001377	-0.000899
$Temp_t$	0.019253	-0.001769	0.001878
$Precip_t$	0.000008	-0.000059	-0.000061
R^2	1: 0.991378	1: 0.982188	1: 0.975179
	2: 0.449007	2: 0.955342	2: 0.942558
	-	3: 0.449007	3: 0.955342
	-	-	4: 0.449007
Break dates			
$\hat{T}_1$	1999	1991	1985
$\hat{T}_2$	-	1999	1985
$\hat{T}_3$	-	-	1999

α_i:intercept; β_i: estimated slope
1, 2, 3 and 4 denote the R^2 of the first, second, third and fourth regime, respectively.
m: number of breaks

Table H.5: Estimation of the extended models for total CO_2 emissions with fuel, d(motor), temperature and precipitation as control variables.

	CO2total-GDP (levels): 2.15		
Parameter estimates	m=1	m=2	m=3
α_1	-1.667799	0.443240	0.954680
β_1	0.000426	0.000370	0.000369
α_2	12.130266	-0.618386	2.637751
β_2	-0.000490	0.000523	0.000277
α_3	-	16.559083	0.242902
β_3	-	-0.000666	0.000495
α_4	-	-	17.584121
β_4	-	-	-0.000701
$Fuel_t$	-0.327585	0.117402	0.000876
$d(motor)_t$	0.007917	-0.005812	0.133897
$Temp_t$	0.089923	-0.030579	-0.065660
$Precip_t$	0.000097	-0.000290	-0.000274
R^2	1: 0.984513 2: 0.463323 - -	1: 0.976839 2: 0.953155 3: 0.444511 -	1: 0.976839 2: 0.837248 3: 0.869702 4: 0.444511
Break dates			
$\hat{T}_1$	1999	1991	
$\hat{T}_2$	-	1999	
$\hat{T}_3$	-	-	
	CO2total-GDP (logarithms): 2.16		
Parameter estimates	m=1	m=2	m=3
α_1	-10.002163	-8.923107	-6.834292
β_1	1.198711	1.113676	0.856128
α_2	2.603713	-9.940663	-8.824794
β_2	-0.119394	1.237936	1.096106
α_3	-	18.093607	-9.964749
β_3	-	-1.690394	1.232494
α_4	-	-	17.086771
β_4	-	-	-1.592809
$Fuel_t$	-0.081478	0.029579	0.019737
$d(motor)_t$	0.000648	-0.001441	-0.000927
$Temp_t$	0.024626	-0.002291	0.002382
$Precip_t$	-0.000010	-0.000057	-0.000060
R^2	1: 0.991378 2: 0.449007 - -	1: 0.982188 2: 0.955342 3: 0.449007 -	1: 0.975179 2: 0.942558 3: 0.955342 4: 0.449007
Break dates			
$\hat{T}_1$	1999	1991	1985
$\hat{T}_2$	-	1999	1985
$\hat{T}_3$	-	-	1999

α_i:intercept; β_i: estimated slope
1, 2, 3 and 4 denote the R^2 of the first, second, third and fourth regime, respectively.
m: number of breaks

APPENDIX I – PUBLISHED PAPER

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Economic growth and transport: On the road to sustainability

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Abstract

Transport sustainability is an essential driving force towards achieving sustainable development. In particular, greenhouse gas (GHG) reduction policies cannot overlook the growing importance of the transport sector as economies expand. In this context, it is important to assess the relationship between carbon dioxide (CO_2) emissions from the transport sector and economic growth, in order to design adequate transport policies. This paper tests the stated relationship for the Portuguese transport sector, using a non-linear cointegration methodology for the first time in this field. The main conclusion is that gross domestic product (GDP) and CO_2 emissions from transport exhibit an increasing monotonic relationship, indicating that economic growth per se will be insufficient to mitigate the emissions. Therefore, it is necessary to search for concrete public policies that can influence both passenger travel behaviour and freight mobility plans, so as to reduce CO_2 emissions without compromising economic development.

Keywords: Sustainability; economic growth; transport sector; environmental Kuznets curve; public policies; non-linear cointegration.

1. Introduction

Transport is a key issue in our lives. Families, governments and companies make transport-related decisions on a daily basis, and with increasing flows of goods and people throughout the globe such decisions will become increasingly relevant. Yet, the undeniable relationship between transport and economic growth is far from being well understood; this poses additional challenges to policymakers when it comes to designing and planning sustainable transport policies. Transport contributes to economic activity by providing the infrastructure that supports the movement of people and goods. While it promotes human and economic development by providing greater mobility, it is also a reflection of the economy, since economic growth drives transport demand through higher rates of motorization, intensification of commutes, and an increase in the quantity of goods being transported, using different modes of transport. Moreover, transport also plays a social role, since it allows people to reach education and health services, as well as promoting accessibility to employment. Therefore, the provision of transport is central to achieving a more inclusive and sustainable society.

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The relevance of transport and mobility in sustainable development is recognized worldwide. To meet sustainable development goals, one must achieve sustainable transport. The European Union Council of Ministers defines sustainable transport as one that (EU, 2001):

- allows the basic access and development needs of individuals, companies and societies to be met safely and in a manner consistent with human and ecosystem health, and promotes equity within and between successive generations;
- is affordable, operates fairly and efficiently, offers choice of transport mode, and supports a competitive economy, as well as balanced regional development; and
- limits emissions and waste within the planet's ability to absorb them, uses renewable resources at or below their rates of generation, and, uses non-renewable resources at or below the rates of development of renewable substitutes while minimizing the impact on the use of land and the generation of noise;

The environmental dimension of sustainable transport is also mentioned in the Report of the United Nations Conference on Sustainable Development, Rio+20, according to which "Sustainable transport achieves better integration of the economy while respecting the environment", (UN, 2012). A key aspect relates to the global environmental impact of transport as a source of greenhouse gas (GHG) emissions. These are a by-product of fossil fuel-intensive economic activity, and have been growing steadily since the onset of the industrial revolution. Moreover, anthropogenic GHG emissions, among which carbon dioxide (CO_2) is paramount, are very likely to be leading contributors to climate change. According to the Fifth Assessment Report of the United Nation's Intergovernmental Panel on Climate Change (IPCC) — Climate Change 2014: Impacts, Adaptation and Vulnerability — climate change is already occurring globally, increasing the frequency of extreme weather events, such as heavy rainfall, floods, heat waves and droughts. Higher or lower agricultural yields, species extinctions and increases in the ranges of disease vectors are among further consequences of climate change. Damages from weather-related catastrophes are on the rise and will, very likely, continue to do so. These weather events are expected to become ever more frequent and severe, although there is great uncertainty regarding future impacts (IPCC, 2013).

As global climate change may threaten social well-being, intra and intergenerational equity, economic growth and ecosystems, tackling it is an urgent priority for the 21st century. Following the Kyoto Protocol, the 2009 Copenhagen Accord set a maximum limit of 2 degrees Celsius, when compared with pre-industrial values, for the increase in mean global temperature. Although the accord itself is not legally binding, the European Union (EU) is committed to the implementation of measures that aim at the stabilization of GHG in the atmosphere. Through the implementation of various policies, including an Emissions Trading System for large industrial emitters, and within the context of the European 2020 strategy, the 27 member states of the European Union (EU 27) managed to reduce GHG emissions (excluding land use, land Use change and forestry) by 18.4% from 1990 to 2011. This downward trend was felt in every sector except transport, which in 2011 accounted for 20.2% of emissions (Eurostat, 2013). The contrasting behaviour of transport GHG emissions begs for a sectoral analysis.

The transport sector, which includes the aviation, road, navigation and railway subsectors, is currently the second highest emitter of GHG, after power and heat production. Although the growth trend in emissions in this sector was interrupted in 2008, because the economic slowdown led to a decrease in passenger and freight transport traffic, growth is expected to resume as the EU economy recovers. By 2010, GHG emissions in the EU 27 transport sector had increased 19%, when compared with 1990. Focusing on CO_2, which is the main GHG, weighing 82.4% of total GHG emission in the EU 27 (Eurostat, 2013), it is clear that the contribution of each mode of transport was highly unequal. In that year, domestic civil aviation emitted 11.6% of the 12.4% share of total civil aviation in CO_2 emissions from transport; road transport and railway contributed 72.1% and 0.6%, respectively; and domestic navigation's emissions accounted for 11.2% of a total navigation share of 14.1% (EU, 2013a). From 1990 to 2010, CO_2 emissions from road transport grew by 23% in spite of the economic crisis (EC, 2012). This subsector was responsible for about 82% of the total final energy consumption of transport, most of it from fossil fuels (Eurostat, 2013). Road transport is, therefore, by far, the main source of GHG emissions within the transport sector, in spite of EU measures aiming for more efficient engines in new passenger cars, which led to a 16.7% decrease in the CO_2 emissions of the average car between 2007 and 2012. Moreover, there has been an increase of the motorization rate in almost all Member States, and the use of a private car continues to account for a significant share of the total passenger transport (Eurostat, 2013).

Thus, to meet the European Union's Sustainable Development Strategy's overall objective on transport — "to ensure that our transport systems meet society's economic, social and environmental needs, whilst minimizing their undesirable impacts on the economy, society and the environment" — a robust assessment of the complex relation between road transport and economic growth is required.

Effective policies of climate-change mitigation should take into account the relation between GHG emissions and economic activity, as well as sectoral and regional aspects, lest their impacts be misunderstood. Since the 1990's, the environmental Kuznets curve (EKC) hypothesis has been widely used to probe the link between economic growth and environmental quality. The initial literature was mostly empirical and it presented the tantalizing suggestion that economic growth, albeit environmentally damaging at first, can actually contribute to better environmental outcomes as countries keep growing. In the early 1990s, several researchers (Grossman and Krueger, 1991; Shafik and Bandyopadhyay, 1992; Panayotou, 1993) identified an inverted U-shaped pattern between environmental degradation and per capita income. Panayotou (1993) was the first to designate this specific relation as the environmental Kuznets curve, due to the previous work of Simon Kuznets, who postulated a similar link between income inequality and per capita income in the 1950s.

Much empirical research has since been carried out on different types of pollution, using time-series, panel data, and cross-section methodologies. If EKCs exist, they should display a turning point, after which economic growth has a positive impact on environmental quality. Underlying economic factors are often organized in three effects: scale, composition (or structure), and technical effects. The first phase of economic growth implies an increase in economic activity, and a shift from the primary to the secondary sector. With all other factors being equal, the consequence is the use of more natural resources as inputs in the production process, as well as an increase in pollution. This mechanism, while harmful for the environment, is called the scale effect, and it explains the initial upward trend of the EKC. As economic growth continues, the composition and technical effects emerge, leading to positive impacts on the improvement of environmental quality. Economic structural change is a dynamic process linked to economic development (Panayotou, 1993). In the first phase, the composition effect accelerates environmental degradation, as the secondary sector includes heavy polluting industries. As the weight of industrial activities in gross domestic product (GDP) starts to decline the pre-industrial economy is replaced by a post-industrial economy, characterized by knowledge-based industries and services which are less polluting (Panayotou, 1997; 2003). The technical effect relates to improvements in production processes. Wealthier countries and trade liberalization promote research and development (R&D) investments leading to innovation, technological progress, and more efficient technologies. Obsolete technologies are replaced by cleaner and more efficient ones, reducing both the use of resources and the levels of pollution in goods production. Together, the composition and technical effects overcome the scale effect and reverse the slope of the EKC (Dinda, 2004).

Additionally to scale, composition and technical effects, income elasticity of environmental quality demand is broadly recognized as one of the factors that also explains the shape of the EKC. At higher per capita income levels, people become increasingly aware of environmental quality. When a certain level of income is attained, the willingness to pay for cleaner environment rises by a greater proportion than income (Roca, 2003). Moreover, consumers with higher incomes enforce environmental protection and regulatory frameworks (Panayotou, 1993; Dinda, 2004).

A large set of EKC empirical works, considering various pollutants in a number of countries, studied individually or in groups, have been undertaken and are seen to be fairly inconclusive, with contradicting results. Theoretical reasons and econometric issues have both been pointed out by several authors to explain this result diversity (e.g., Dinda, 2004; Stern, 2004; Van Alstine and Neumayer, 2008; Carson, 2010).

Portugal has occasionally been covered in the EKC literature, also with conflicting results. Mota and Dias (2006) applied a time-series approach to look for an EKC for per capita CO_2 emissions, between 1970 and 2000, in the country. This approach tests for stationarity and cointegration, using the Augmented Dickey-Fuller test (ADF) and the traditional Engle and Granger procedures, respectively. The relation between per capita CO_2 emissions and per capita real GDP proved to be inconclusive, as both the linear and cubic models present reliable results, although the authors believe the linear model to be the better one. Regardless of which model best describes the CO_2-GDP link, an EKC does not appear to exist. They also identified a positive contribution from the service sector to CO_2 emissions. In order to explain this positive relation, the authors stress that the service sector encompasses the transport subsector. However, a sectoral analysis was not performed.

Acaravci and Ozturk (2010) execute an autoregressive distributed lag (ARDL) bounds cointegration analysis, using 1960-2005 data for Portugal, among other countries, and they also find that the EKC does not hold. Shahbaz *et al*. (2010) use the same ARDL methodology for Portugal, during the period of 1971-2008. Their results are in sharp contrast to those of Mota and Dias (2006) and Acaravci and Ozturk (2010), since they find that the EKC hypothesis does hold. Similar results were obtained in Jaunky (2011), where the EKC was tested for Portugal and 35 other high-income countries, using data from 1980 to 2005.

This brief literature review shows that empirical evidence of an EKC for CO_2 emissions in Portugal is mixed, hindering the attainment of useful insights for climate policy design and implementation. Nevertheless, a transport sector analysis has not yet been performed. In this context, the current paper contributes to the understanding of the relationship between economic growth and CO_2 emissions by highlighting the role of transport, since it is a major contributor and has been insufficiently targeted in existing policies. In particular, and contrary to the traditional approach, we apply the non-linear cointegration methodology proposed by Breitung (2001) and Choi and Saikkonen (2004) to the Portuguese transport sector. This is, furthermore, the first time that a non-linear cointegration approach is used to assess the EKC for transport. We find evidence of non-linear cointegration and an increasing monotonic relationship between income and CO_2 emissions from transport.

The remainder of the paper is organized as follows: Section 2 briefly reviews existing literature on transport emissions and income, the data and unit root tests are described in Section 3, the EKC model and cointegration methodologies employed are introduced in Section 4, the results and discussion are presented in Section 5, and the concluding remarks are provided in Section 6.

2. Transport emissions and economic growth: empirical findings

Although the transport sector is one of the main emitters of GHG, second only to power generation, the literature focusing on the relationship between transport-related emissions and economic growth is scarce. Liddle (2004) performed an ordinary least squares (OLS) regression, fixed and time effect regressions, with time dummies, on a data panel of 23 Organisation for Economic Co-operation and Development (OECD) countries, with observations being taken over 10-year intervals from 1960 to 2000. The goal was to investigate the EKC relationship between per capita road energy use and per capita GDP. The author also included geographic and demographic variables. Three models were estimated: a quadratic model in levels, a quadratic log-log model, and a lin-log model. For all the models, the EKC hypothesis was rejected, because the parameters on the GDP-squared terms were not statistically significant and the turning point values were well above the sample range. Therefore, the relationship between per capita road energy use and per capita GDP was found to be monotonic.

Tanishita (2006) studied the evidence of an EKC for energy intensity from passenger transport, using city-based data, for the period between 1980 and 1995. The results supported the EKC relationship between energy intensity of private and public transport and per capita gross regional product (GRP), with a turning point that ranged from US$22,000 to US$26,000 (PPP, 1995).

The validity of the EKC shape for transport was also analysed by Meunie and Pouyanne (2009), who stress that the income effect on transport emissions is a combination of three factors. The first acts through behaviour changes, encompassing a direct link — higher incomes lead to more car ownership and use — as well as an indirect influence, since higher-income families move to detached houses, which are more prevalent in the suburbs than in city centres. The second factor acts through technology choice and it is ambiguous, since in one situation a growing level of income makes it easier to purchase more efficient cars (incorporating the most recent environmental innovations) and in another situation cars are linked to social position, so richer households may tend to buy potent vehicles with higher energy consumption. Lastly, the third aspect is political, i.e., related to the influence of government policy. At higher incomes, governments find it easier not only to promote environmentally-friendly technologies, but also to implement an environmentally-rich regulatory framework. The empirical work of Meunie and Pouyanne (2009) considers the possible theoretical mechanisms supporting an EKC for daily mobility. The quadratic functional form was applied using the Millennium Cities Database, for 88 cities, for the year 1995, to estimate cross-section models for carbon monoxide (CO), nitrogen oxides (NO_x), volatile hydrocarbons (VHC), SO_2, and also for density of pollution. The EKC hypothesis was tested for pollution emissions from transports and for per capita energy consumption. There was evidence of an EKC for CO and NO_x emissions as well as per capita energy consumption. As for VHC and SO_2 emissions, the EKC could not be applied. For these authors, more than the level of income, travel behaviour is essential to explain pollution from transport, and the EKC hypothesis failed to explain the environmental consequences of transport behaviours, which are most tightly connected with public policies. They also argue that higher concentrations of pollutants drive citizens to demand tighter ecological regulation, and that higher income levels allow regulators to implement environmental policies.

Ubaidillah (2011) explored this relationship for the United Kingdom, from 1970 to 2008. Again, CO was used, as this pollutant is a more road-transport specific indicator. A quadratic model in levels was employed, as well as the Johansen maximum likelihood methodology for cointegration analysis. Just like Tanishita (2006), the author found evidence of the EKC pattern. In this case, the per capita CO — per capita GDP (2000 constant price) relationship has a turning point at US$21,402. The EKC pattern for the UK's road transport sector is explained by the increased usage of private or passenger vehicles, which follows the growth trajectory of income. In turn, the use of private or passenger vehicles translates into an increase in fuel combustion and, as a consequence, into an increase in CO emissions. However, from a certain level of per capita income, an improvement in technology, rules, and behaviours occurs, possibly explained by the increasing value attributed to clean air, and thus CO emissions start to decline.

Cox *et al*. (2012) resorted to a 2006 survey, undertaken in six case study areas in Scotland, to inquire about the existence of an EKC for household transport CO_2 emissions. The authors used a simple ordinary OLS regression of log household CO_2 emissions, and a household annual income dummy. This study highlights that households with annual incomes equal to or greater than £52,000 produce 92% more CO_2 emissions than lower-income households. On average, richer households have more than one vehicle, newer but not less polluting, and drive more often, therefore the existence of the EKC for private road vehicles is not supported.

Ben Abdallah *et al*. (2013) examined the relationship between transport value added, road transport-related energy consumption, road infrastructures, fuel prices, and CO_2 emissions for the transport sector, in Tunisia, from 1980 to 2010. Regarding the analysis of the relationship between per capita CO_2 emissions from the transport sector and the per capita transport value added, the authors selected a log-log cubic model and the Johansen cointegration methodology. The parameter of the cubic term is negative and statistically significant, suggesting that the relationship between value added and CO_2 emissions in transport is described by an inverse N-shape, refuting the existence of the EKC. Despite these results, the authors believe that, in practice, the relationship is described by a monotonically increasing curve as the first turning point is equal to 74.88 Tunisian national dinars (constant 2000 TND) (corresponds to US$39.64, exchange rate at 13 January 2015), which is a very low value and the second turning point — 578.82 TND (corresponds to US$306.40, exchange rate at 13 January 2015) — exceeds the dataset values.

Liddle (2013) looked at the inverted-U relationship between per capita GDP and three air pollutants — carbon monoxide (CO), nitrogen oxide (NOx), and volatile hydrocarbons (VHC) — emitted from urban transport, and analyzed, further, the relationship between per capita GDP and urban transport energy consumption. The paper starts by underlining the importance of an influence zone for each transport pollutant, since the EKC are more likely for pollutants with local impact, whose negative consequences can be immediately sensed and are controlled locally, than for pollutants with a global impact. Since the environmental consequences of CO and VHC are local, unlike those of NO_x, it is to be expected that, if an EKC exist, it is far more likely for the first two pollutants. To test the EKC hypothesis, the Millennium Cities Database for Sustainable Transport was used for 84 cities, in both developed and developing countries, for the year 1995, and cross-section quadratic log-log models were estimated for each pollutant and also for energy consumed in transport. The variables 'urban density' and 'fuel prices' were included in the regressions, since, in previous studies, a negative relation between vehicle miles traveled or energy consumed in private transport, and both urban density and fuel prices was identified. Results confirm the EKC shape for all the dependent variables under study. However, two distinct interpretations can be made, justified by the values of the turning points. The turning points per capita GDP levels for CO, VHC and NOx were US$7,322; US$9,124; and US$15,939, respectively. All these values were within the sample range, but of the three pollutants under study, NO_x was the one that presented the highest turning point. This result was as expected, since the impact zone for this pollutant is higher than for the other two, and CO and VHC emissions can be reduced more easily through improvements in basic technologies, such as more efficient catalytic converters.

Regarding energy consumption, the turning point was found at US$137,698, an amount that exceeds those from the sample. The author concluded, then, that the EKC holds for the three pollutants, but not for energy consumed in transport, whose relationship with income is monotonic. Regarding other explanatory variables, results indicate that, the higher the urban density, the more people make use of non-motorized and public transport alternatives, thus leading to a reduction in per capita energy consumption in transport and, consequently, a reduction in per capita emissions. Still, the literature described in this section is fairly limited, and stronger conclusions demand more empirical work.

3. Data and unit root tests

3.1. Data

The Portuguese transport sector underwent several changes, mostly since joining the European Economic Community (EEC), later European Union (EU), in 1986. A part of EU funds was directed towards developing the national transport infrastructures. Despite the investments, each mode of transport has evolved distinctly, with quite different impacts on the mobility of both people and goods. Based on the scope of this work and on the available data, we give particular emphasis to road transport because it is still the dominant mode of transport in Portugal, although we provide some information below on all modes of transport that are included in the CO_2 emissions data, namely: domestic aviation, inland waterway transport, railway, and road transport.

3.1.1. Domestic civil aviation

The main Portuguese airports are located in Lisbon, Porto, and Faro. The aviation sector grew steadily over the last decades, in both number of yearly flights and passengers, but the domestic part is diminishing. For instance, throughout 2012, there were 5.7% fewer offered seats and also a 6.3% decrease in the number of passengers transported. For the same year, the regular domestic traffic flight operations represented 22.2%, 15.5%, and 7% of the total number of flights, km-traveled and number of flight hours, respectively. These figures indicate that domestic flights is a small share of the total aviation sector (INE, 2013).

3.1.2. Inland waterway transport

Historically, the inland waterway mode has had no significant expression in the Portuguese transport sector. The region of Lisbon, due to the boat crossings over the rivers Tejo and Sado, has most of the demand for this mode, and in spite of annual fluctuations, the overall trend there is stable to decreasing. In 1998, the opening of the Vasco da Gama bridge lead to a severe drop in both passengers and vehicles in the Tejo: from 1998 to 2004, a 40% and 71% decrease in passengers and vehicles, respectively, notwithstanding the investment made since 1995 in the modernization of the ferry fleet. As for the Sado, crossings have also declined in both the number of passengers and vehicles (ECORYS, 2006; INE, 2013).

3.1.3. Railways

The EU structural and cohesion funds made the modernization of the Portuguese railway track possible. In 1993, around 85% of all railway lines in operation was single track. In 2004, this number had declined to 79%, because of the improvement of some rail sections to double or quadruple track, which translated into an increase of 35% from 449 to 607 km (ECORYS, 2006). The total length of the national rail network, however, has been decreasing. At the end of 2012, the total length of the railway track in operation was about 2,541.2 km, a 9.0% reduction from 2011 (INE, 2013). In fact, according to Eurostat the railway density in Portugal is considerably lower than the EU average. The number of passengers in railway mode has also been declining. In 2012, the number of passengers was 11.3% less than in 2011. For the same period, the transport of goods by rail registered a 2.7% reduction when compared with 2011 (INE, 2013).

3.1.4. Road transport

Where the EU structural and cohesion funds had a massive impact was in road infrastructures. In the period between 1986 and 2006 alone, 2,700 km of main itineraries and complementary itineraries were built; of those, 2,300 km had a motorway profile (INE, 2007). While in 1990, the total length of motorways in Portugal was 316 km, in 2010 this number rose to 2,737 km. Improvement in road infrastructures and an increase in family income spearheaded the growth of vehicle fleets, especially private car ownership, and road travel. The evolution of the Portuguese rate of motorization (number of passenger cars per 1,000 inhabitants) reflects these changes. In 1990, the rate of motorization was 185 cars per 1,000 inhabitants, significantly lower than the EU 27 average of 345 cars per 1,000 inhabitants, while in 2011 it was 447 cars per 1,000 inhabitants, a value still below, but closer, to the EU 27 average of 483 (EU, 2013b).

The growth of mobility, in particular road transport, was accompanied by a marked increase in GHG emissions. As in the EU 27, the transport sector is the second main emitter of GHG in Portugal, representing 24.7% of total emissions, only surpassed by energy production (25.3%) (Pereira *et al.*, 2014). The higher energy demand associated with the increase in the number of trips, more powerful engines and increase in kilometres traveled, translated into an increase of GHG emissions, especially of CO_2. Road transport stands out in Portuguese CO_2 emissions (international bunkers included), as expected. In 2010, domestic aviation was responsible for 13.2% of CO_2 emissions, from a total of 13.1% for the aviation sector. Inland waterway transport emitted 12.3% of the total CO_2 emissions from navigation (8%). Railway was the smallest CO_2 emissions contributor — 0.2% — and road transport represented 78.7% of the CO_2 emitted by the entire transport sector (EU, 2013a).

Yet road transport has been changing and that, once again, is reflected in its GHG emissions. Until the beginning of the 2000s there was a steep increase in emissions; from 2002, GHG emissions from transport started to stabilize and, in 2005, began to decline. From 1990-2002 emissions from transport sources increased 97%, due to the steady growth of vehicle fleets (in particular those with more powerful engines); those emissions also grew due to road travel, from 1990 to the early 2000s, reflecting an increase in family income and the strong investment in road infrastructures, in the country, in previous decades. However, from 2005 to 2012, transport emissions were reduced by around 15% (Pereira *et al.*, 2014).

To conduct our empirical analysis we employed annual time series for population, CO_2 emissions and real GDP at 2006 prices, from 1960 to 2010. Data for population and for CO_2 emissions from transport were downloaded from the World Bank website (original data from the International Energy Agency). CO_2 emissions from transport encompass the emissions from fossil fuel combustion for all transport activities, irrespective of the activity sector and excluding the International Marine Bunkers and International Aviation. The data for real GDP is from Pordata (original data from INE and Banco de Portugal). Table 1 shows descriptive statistics for per capita CO_2 emissions from transport and per capita GDP (both in euro and US$). The highest level of per capita real GDP was observed in 2007 — 15,521 euros constant 2006 (corresponds to US$18,343.64, exchange rate at 13 January 2015) — while the lowest level was in 1960. The maximum of per capita CO_2 emissions was in 2004 — 1.889 — and the minimum was also in 1960.

Figure 1 represents the evolution of both variables, and the similarity between them is notable until 2004. However, from 2005 onwards, the per capita GDP seems to stabilize, while we can see a decrease in per capita CO_2 emissions from the transport sector.

3.2. Unit root tests

We begin our empirical analysis by investigating the stationarity of the data, testing for the presence of unit roots for both per capita real GDP and for per capita CO_2 from transport series, to avoid the problem of spurious regressions. To do so, we use three distinct tests to draw conclusions in more solid grounds — Augmented Dickey and Fuller (1979) (ADF), Phillips and Perron (1988) (PP) and Elliott *et al.* (1996) (ERS) — with an intercept. The ADF and ERS tests are run with automatic lag length selection on the basis of the Schwarz information criterion (maximum of 10 lags), and the PP test is run with automatic bandwidth selection according to Newey and West. For all tests (ADF, PP and ERS) the null hypothesis is that the series is non-stationary against the alternative that it is stationary. The results reported in Table 2 show that the null hypothesis of unit root is clearly not rejected at the usual level, and found to be stationary after taking first differences. Hence, we

Table 1. Descriptive statistics of data

Variable	Sample size	Mean	Std. Dev.	Min	Max
Per capita real GDP	51 years	9,587.528€ (US$11,330.54)	4,079.716€ (US$4,821.41)	3,135.867€ (US$3,705.97)	15,521.78€ (US$18,343.64)
Per capita CO_2 emissions from transport	51 Years	0.9448798	0.568981	0.199826	1.889169

Notes: All data is in levels. The values in brackets indicate the per capita GDP in US$(Rate of the day: 2015-01-13).
Source: Author's elaboration.

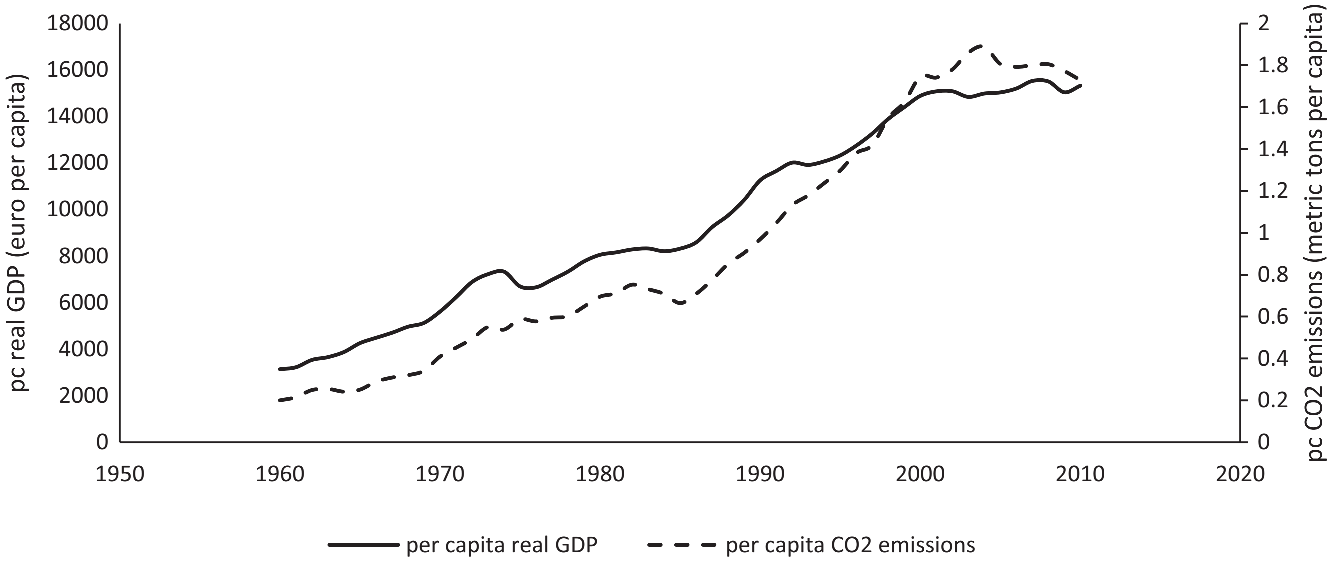

Figure 1. Per capita real GDP (base year = 2006) and per capita CO_2 emissions from transport sector, from 1960 to 2010.
Source: Author's elaboration.

Table 2. Results of the unit root tests

Variable	ADF		PP		ERS		Order of integration
	Level	1st difference	Level	1st difference	Level	1st difference	
Per capita real GDP	−0.890022	−3.880122*	−0.745922	−3.942680*	183.8637	1.549672*	I(1)
Per capita CO_2 emissions from transport	−0.557709	−4.061222*	−0.277320	−3.966196*	143.4646	1.611751*	I(1)

Note: The regressions include intercept for both variables.
*Indicates the rejection of hypothesis of non-stationarity 1% significance level.
Source: Author's elaboration.

conclude that per capita real GDP and per capita CO_2 emissions from transport are integrated of order one – I(1) – at a 1% significance level.

4. Methodology

4.1. EKC model

The bulk of the EKC literature assumes that the relationship between environmental degradation and economic growth can be adequately described by a polynomial function of income. The model applied empirically varies in two dimensions: (1) it can be linear, log-linear, linear-log or log-log, and (2) it can be quadratic or cubic. Although in the original paper of Grossman and Krueger (1991), the relationship was specified as a cubic functional form, which is more flexible, the second-order polynomial functional form is more often applied in empirical EKC analysis.

The conventional reduced-form model, for both cross and single-country studies, to test the EKC theory consists of relating the per capita environmental indicator to the per capita income (proxied by real per capita GDP). Extra explanatory variables of the environmental degradation may or may not be included.

For a parametric time-series analysis of the EKC hypothesis, the reduced-form model can be generally specified as:

$$E = \beta_0 + \beta_1 GDP_t + \beta_2 GDP_t^2 + \beta_3 GDP_t^3 + u_t \qquad (1)$$

where: E is the environmental indicator; GDP is per capita real GDP; β_k is the coefficient of the kth explanatory variables; u is the error term; t is the time period.

Model (1) allows for testing the possible patterns of the environmental degradation-income relationship, namely:

(i) $\beta_1 = \beta_2 = \beta_3 = 0$: Flat pattern or no relationship between environmental indicator and GDP;
(ii) $\beta_1 > 0$ and $\beta_2 = \beta_3 = 0$: Monotonic increasing linear relationship between environmental indicator and GDP;
(iii) $\beta_1 < 0$ and $\beta_2 = \beta_3 = 0$: A monotonic decreasing linear relationship between environmental indicator and GDP;

(iv) $\beta_1 > 0$, $\beta_2 < 0$ and $\beta_3 = 0$: Inverted-U-shaped relationship between environmental indicator and GDP, i.e., EKC holds;
(v) $\beta_1 < 0$, $\beta_2 > 0$ and $\beta_3 = 0$: U-shaped relationship between environmental indicator and GDP;
(vi) $\beta_1 > 0$, $\beta_2 < 0$ and $\beta_3 > 0$: N-shaped relationship between environmental indicator and GDP; and
(vii) $\beta_1 < 0$, $\beta_2 > 0$ and $\beta_3 < 0$: Inverted N- shaped curve relationship between environmental indicator and GDP;

Note that of all the possibilities listed above, only (iv) suggests the existence of an EKC.

4.2. Cointegration analysis

A vast existing EKC literature makes use of cointegration analysis. This raises important statistical problems in parametric reduced-form models such as the EKC, as it includes not only the per capita GDP but also its squares and cubes as explanatory variables (Müller-Fürstenberger & Wagner, 2007).

It is widely accepted that GDP is non-stationary. Thus, since per capita GDP is an integrated variable, the powers of GDP are a non-linear transformation of an integrated I(1) process, which in turn are not necessarily integrated (Müller-Fürstenberger & Wagner, 2007; Hong and Wagner, 2008). Therefore, EKC regressions cannot be analysed within the usual linear cointegration methodologies, as they call for an alternative type of asymptotic theory, and lead to different properties of the estimators (Hong and Wagner, 2008). Most up to date EKC models neglect the econometric implications of this issue. In this context, a significant part of the empirical evidence of the EKC is doubtful.

Hence, in order to obtain reliable EKC outcomes and confirm the existing ones, GDP and its integer powers should be considered by applying non-linear cointegration methodologies. The absence of relevant non-linear cointegration analysis can lead to misleading results. If the EKC regressions do not fulfill the cointegration property, then the estimates may be spurious (Aslanidis, 2009).

To overcome the econometric constraints of linear cointegration techniques such as the Engle and Granger (1987) two-step technique or Johansen and Juselius' (1990) maximum likelihood approach, the present research will rely on non-linear cointegration methodologies based on the procedures developed by Breitung (2001) and Choi and Saikkonen (2004). This is the first time that a non-linear cointegration approach will be used to assess the EKC for transport.

Since the unit root tests' results support the evidence for non-stationarity of per capita GDP and per capita CO_2 emissions from road transport, we proceed by examining the long-run cointegrating relation between the variables.

The first step is to perform the Rank Tests for Cointegration proposed by Breitung (2001). Breitung's rank test for cointegration is a nonparametric methodology which can be applied no matter what the nature of the series is, as it can detect both linear and non-linear cointegration relations. The rank test is based on rank transformed series and this monotonic transformation is a required condition in finding cointegration. Two rank statistics — K_T is a Kolmogorov-Smirnov type of statistic and ξ_T is a Cramer-von-Mises type of statistic — are considered to assess the rejection of the null hypothesis of no (non-linear) cointegration. Breitung proposes the correction of the original test statistics to K^*_T and ξ^*_T, if the time series are mutually correlated. In either cases, there is no (non-linear) cointegration if the observed test statistics are smaller than the respective critical values, available in table 1 of Breitung (2001).

If cointegration is found, the second step of the non-linear cointegration methodology is to investigate if the cointegration is linear or non-linear. We employ two different tests: the rank test for neglected non-linearity by Breitung (2001), and the linearity test developed by Choi and Saikkonen (2004).

The rank test for neglected non-linearity (Breitung, 2001) identifies the linearity or non-linearity of the cointegration relation. Under the null hypothesis of linear cointegration, the associated score test statistic has an asymptotic chi-square distribution with one degree of freedom. The null hypothesis of linear cointegration is rejected if the observed value exceeds the corresponding critical value.

The linearity test (Choi and Saikkonen, 2004) opposes the null hypothesis of linear cointegration against the alternative of a specific non-linear form of cointegration, namely a smooth-transition process. In this sense, Breitung's test is more general. Choi and Saikkonen's test is based on first and third-order Taylor approximations of the functional form under the alternative hypothesis and requires auxiliary regressions to run the actual tests. The potential existence of endogeneity is corrected based on a leads and lags estimation approach. The test statistics — $\tau 1$ and $\tau 2$ (first-order test); and $\tau 1\mu$ and $\tau 2\mu$ (for third-order tests) — follow the standard chi-square distribution. The null hypothesis is rejected if the test statistics values are larger than the critical ones.

We conduct a dynamic ordinary least squares (DOLS) estimation approach to investigate if the EKC hypothesis holds for the Portuguese road transport sector.

5. Results

5.1. Cointegration results

5.1.1. Breitung rank tests for cointegration

The results of the rank tests for cointegration are summarized in the first column of Table 3. Although for the $K^*_{T=100}$ statistics the null hypothesis of no (non-linear) cointegration

Table 3. Results of the rank test for cointegration, rank test for neglected non-linearity, and tests of linearity

Rank test for cointegration		**Rank test for neglected non-linearity**	**Tests of linearity**		
$K^*_{T=100}$	$\xi^*_{T=100}$	**Linearity Test Statistic**	$\tau 2\mu$ (K = 1)	$\tau 2\mu$ (K = 2)	$\tau 2\mu$ (K = 3)
0.4523	0.0189	7.0949	7.4172	8.1725	6.5347

Source: Author's elaboration.

cannot be rejected, based on $\xi_{T=100}$ at a 5% significance level, the series are cointegrated. Despite the mixed results, we assume that there is cointegration (i.e. the series co-move) not only because of the $\xi_{T=100}$ statistics but also because it is consistent with previous empirical EKC studies. As our series are I(1), this avoids having spurious results.

5.1.2. Non-linear cointegration tests

The second and third columns of Table 3 report the results for the rank test for neglected non-linearity and for tests of linearity of the cointegrating relationship. The null hypothesis of linear cointegration is rejected by both procedures at a 1% significance level which clearly indicates the existence of a non-linear cointegration relationship in the long-run between per capita GDP and per capita CO_2 emissions from road transport.

5.2. Estimation of the EKC model

In order to check if the EKC hypothesis holds for the Portuguese case, we have to estimate the cubic reduced-form model. The signs of the estimated coefficients by DOLS (equation 2) confirm the existence of an N- shaped relationship, which exhibits the same pattern as the inverted-U curve only at the initial phase of the CO_2-GDP relationship, starting with increasing CO_2 emissions and followed by a downward trend after the first turning point. Beyond a certain income level, however, there is a second turning point and the relationship between the two variables is ascending again.

Apparently, these results suggest that the CO_2 emissions would eventually come back with economic growth, after a period of reduction. However, this would be a hasty conclusion, because both the estimated per capita income turning points are out of the observed sample range. The turning points are well below the lowest observed per capita GDP value, making them irrelevant for the analysis. The appropriate interpretation suggests the existence of an increasing monotonic relationship between per capita GDP and per capita CO_2 emissions from transport. Figure 2 shows that our regression model fits the data well, as the actual and fitted curves are very similar.

$$CO_{2\,Transp} = 8.3883 * 10^{-2} + 9.5772 * 10^{-5} GDP_t - 1.5248 * 10^{-9} GDP_t^2 + 2.3965 * 10^{-13} GDP_t^3 \quad (2)$$

To attain better and more complete conclusions based on the results obtained, we should confront them not only with the previous EKC empirical findings for total CO_2 emission for Portugal, but also with earlier and scarce EKC studies focused on transport. This is the goal of the following subsection.

5.3. Discussion

When comparing our results with outcomes of the EKC studies for total CO_2 emissions in Portugal mentioned in Section 1, two relevant comments arise. First, the outcomes of the present paper are in line with those obtained by Mota and Dias (2006). They found a linear positive relationship between economic growth and total CO_2 emissions for Portugal. Even though they do not examine the EKC hypothesis for transport only, it is included in the service sector weight they use as an additional explanatory variable in the model. Second, the behaviour pattern between economic growth and CO_2 emissions (or another pollutant under study) is a consequence of the behaviour patterns in different sectors. Thus, our results for transport do not necessarily contradict those obtained by Shahbaz *et al.* (2010) and Jaunky (2011): we only focus on transport, yet there are other important sectors such as power that influence the overall shape of the economic growth-CO_2 emissions nexus.

Considering the EKC studies for transport, our results strengthen the idea that transport will keep growing with income, as in Liddle (2004), Cox *et al*. (2012) and Ben Abdallah *et al.* (2013). Despite the contradicting results noted in Section 2, with other authors verifying the EKC for this sector, it is known that the existence of the EKC is partially dependent on the type of pollutant studied. The EKC hypothesis is more likely to occur for local pollutants with direct effects (Arrow *et al*., 1995; Stern, 2004; Liddle, 2013), which does not prove to be the case for CO_2.

The increasing monotonic relationship found in the Portuguese case suggests that transport is a normal good, since an increase in CO_2 emissions follows an increase in income. As income rises, people tend to consume more mobility, as they are willing to pay more for it and can also afford it. Therefore, the typical measures that have been applied in several countries, including Portugal, such as restrictions on the motor industry that cap maximum emissions per vehicle, increase in biofuels, more efficient

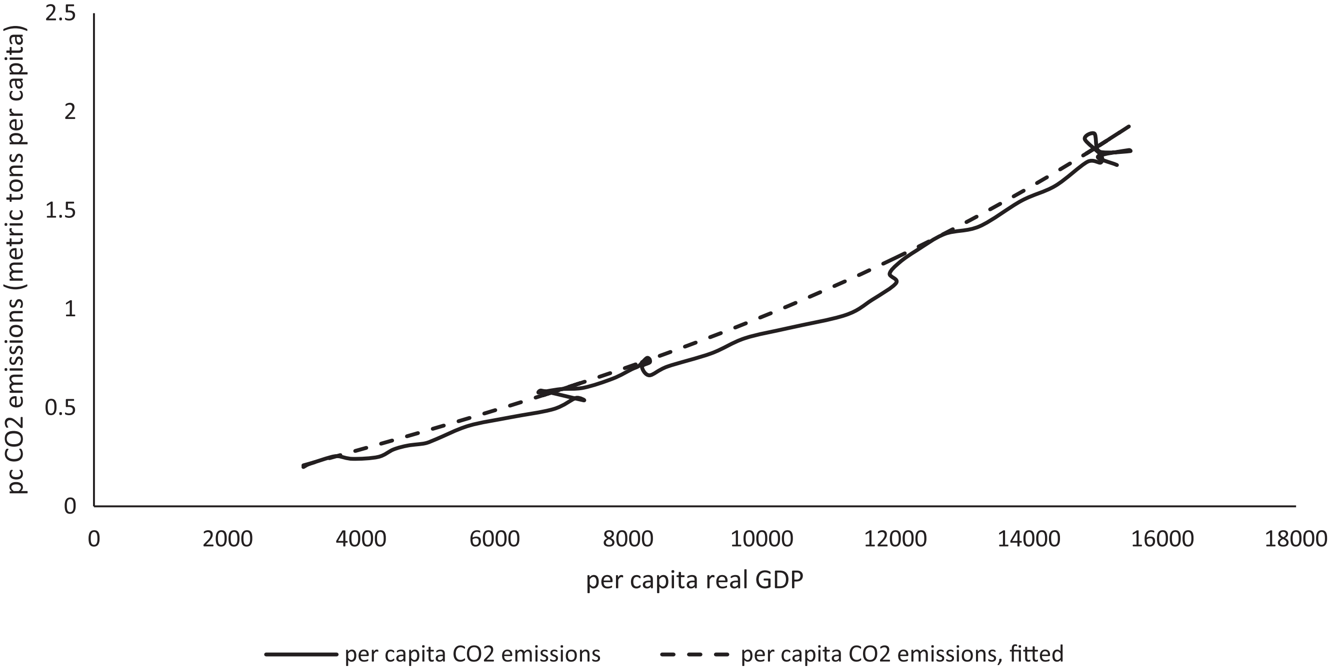

Figure 2. Per capita real GDP versus per capita CO_2 emissions from Portuguese transport sector from 1960 to 2010.
Source: Author's elaboration.

engines, or conventional fuels of better quality, have not been sufficient to decrease CO_2 emissions since they did not completely overcome the income effect. Note that the non-verification of the EKC does not invalidate the scale, composition and technological effects on the transport sector. What happens is that the magnitude of these effects is influenced by the economic, political, cultural, and environmental specificities of each country (Cole, 2007), which may help to explain the opposite findings of previous empirical works such as those mentioned in section 2.

6. Concluding remarks

The understanding of the CO_2 emissions profile of a country is necessary but not sufficient to design and implement efficient mitigation policies. It is imperative to undertake an individual analysis of each of the main emitting sectors to assess sectoral patterns in the long-run relationship between economic growth and emissions. The present paper provides a contribution in this direction, by examining the second largest source of anthropogenic carbon dioxide emissions in Portugal, the transport sector. Our work not only allows us to corroborate the link between economic growth and CO_2 emissions, but also the behaviour of that link. According to our findings, there is a monotonically increasing relationship between economic growth and CO_2 emissions from transport, which allows us to draw several important conclusions, not only for Portugal, but also in a more general perspective.

As Cole (2007) points out, an enhancement in environmental quality does not directly follow economic growth, by itself. What happens is that economic growth stimulates the implementation of legislation on environmental protection and facilitates capital investment, both of which can contribute to the reduction in per capita emissions of some pollutants. It is, therefore, necessary to conduct an analysis of the relationship between growth and environmental quality by country, and this analysis must include the role of governance and the environmental regulation framework.

The inability of economic growth *per se* in reducing CO_2 emissions has policy implications. Public policies geared towards the transport sector are needed. The design of such policies requires a deep knowledge of other variables that explain the rising emissions of CO_2, namely travel behaviour and its related conditioning factors (Meunie and Pouyanne, 2009). Travel behaviour, irrespective of the mode of transport selected, is influenced by several factors, such as fuel prices (including taxes), infrastructure, demographics, changes in lifestyle, and the price and quality of public transport. Such public policies should also be able to incorporate and balance any conflicting impacts of higher income levels on transport, mainly road transport, which is by far the most critical source of CO_2 emissions. Indeed, richer families can buy more than one car and choose more powerful engines, while at the same time become more demanding, when it comes to environmental quality. Sustainable transport policies will allow for a reduction in the number of travels, a decrease in distances traveled, and the adoption of alternative modes of transport such as public transport, cycling, walking, and the substitution from both car and aviation by rail for intercity travel.

Though transport sustainability is a worldwide goal, its attainability demands customized policies, based on the characteristics of each country. For instance, in 2010 (the latest year of our sample), Portugal had a per capita GDP (in purchasing power standards) 20% below the EU 28 average, while the Netherlands was 30% above (Eurostat, 2013). For that same year, according to the 'Future of Transport' analytical report (GO, 2011), these countries displayed mobility patterns which can be contrasted with the EU average: 52.9% of EU citizens used their car as the principal mode of transport on a daily basis, 22% opted for public transport, 12.6% chose to walk, and 7.4% selected cycling. Portugal's numbers were similar for car usage and public transport, while 17% of the Portuguese population chose to walk and only 1.6% indicated cycling. In the Netherlands, although the private car is also the main mode of transport for 48.5% of the citizens, cycling comes second, for 31.2% of people; public transport was elected by 11% of the population as their main mode of transport and, lastly, 3% of the citizens preferred to walk.

This disparity in the choice of the main mode of transport for daily activities — work and leisure — illustrates the importance of factors other than income, such as infrastructure, demographics or culture, on the efficacy of public policy directed towards transport. Sustainable transport policies cannot be applied wholesale by all countries, since it is imperative to take into consideration national specificities.

As incomes rise, governments have at their disposal several measures that may contribute to an offsetting reduction in CO_2 emissions from transport, such as economic incentives (price increases through fuel taxes), campaigns to raise awareness on climate change, and policies to support alternative modes of transport like car-sharing, cycling, walking, and the use of public transport. However, it is essential to know which measures best fit each national reality, and how they should be implemented, so that they may lead to a reduction in CO_2 levels, and at the same time promote inclusion and social equality, without compromising economic growth.

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