# Effectiveness of Carbon Pricing Policy: The Case of Nordic Countries

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# Effectiveness of Carbon Pricing Policy: The Case of Nordic Countries

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## Abstract

Carbon taxation has become prominent as an effective policy tool in combating global climate change in today's world. This study aims to investigate the impact of carbon pricing on carbon emissions and carbon footprint (CF), focusing on the Nordic countries, which were the first to implement carbon taxation, using panel data analysis method between 1992 and 2021. The econometric tests applied in the study are as follows, in order: cross-section dependence tests, Delta homogeneity tests, second-generation panel unit root tests, Gegenbach et al. (2016) panel cointegration test, panel Dynamic Ordinary Least Squares Mean Group (DOLSMG) estimator, and Dumitrescu and Hurlin (2012) panel causality test. According to the findings of the coefficient estimation results, we determine that carbon pricing is effective in reducing carbon emissions and CF. In addition to the findings indicating cointegration among the variables, we also obtain evidence of a unidirectional causal relationship from carbon pricing to carbon emissions and CF.

**Key Words:** Carbon Tax, Carbon Footprint, Carbon Emissions, Environmental Pollution, Nordic Countries, DOLSMG, Panel Causality

## Introduction

In recent years, policymakers have been developing various policy tools to address the growing global warming and combat climate change. In this context, carbon taxation emerges as an effective policy tool in the fight against global climate change. There are numerous empirical studies in the literature to evaluate the effectiveness of environmental taxes. Although there are a significant number of studies on environmental taxes, particularly carbon taxation, the time periods covered in these studies are relatively narrow, and the number of empirical studies conducted specifically on carbon taxation is relatively small.

According to Bayer and Aklin (2020), policymakers, academics, and market participants in Europe have expressed doubts about the effectiveness of carbon pricing in tackling climate change. However, a significant portion of this doubts centers on prices that fall below anticipated levels, particularly in relation to the societal impact of carbon emissions. Another aim of the study is to contribute to the ongoing debates surrounding carbon pricing policies in Europe.

This study stands out from other studies in two aspects. Firstly, it covers a long time period from 1992 to 2021, allowing for a more comprehensive analysis. Secondly, the study not only examines the impact of carbon taxation on carbon emissions but also investigates its impact on CF. This provides a more holistic perspective on the issue. In this study, we employ the balanced panel data analysis method to investigate the effectiveness of carbon pricing policy in countries such as the Nordic countries, which were among the first to implement carbon taxation. In the study, we construct separate models for each dependent variable, and we select the carbon pricing rate as the independent variable. Additionally, we include per capita gross domestic product (GDP per capita) as a control variable in the models. We obtain the CF data from the Global Footprint Network database, while we obtain the data for carbon emissions (per capita in metric tons), gross domestic product (per capita), and carbon pricing from the World Bank database.

The overall aim of the study is to determine whether there exists a long-term relationship between carbon pricing and carbon emissions as well as CF, and if so, to identify the direction and coefficients of this relationship. Within this framework, we utilize panel data analysis method, which has been frequently employed in recent years. To determine whether there is any cross-sectional dependence among countries, we apply the Breusch-Pagan (1980), Pesaran (2004), and Pesaran et al. (2008) tests. In the next stage, we perform the Pesaran et al. (2008) delta homogeneity test to decide on the homogeneity status of the slope coefficients. Next, we employ the CIPS (Cross-sectional Augmented Im, Pesaran, and Shin) and CADF (Cross-sectional Augmented Dickey-Fuller) unit root tests, developed by Pesaran (2007), to determine the stationarity of the series. Subsequently, we apply the panel cointegration test developed by Gegenbach et al. (2016) to test for the existence of long-term relationships among the variables. We then estimate the long-term coefficients using the DOLS-MG (Dynamic Ordinary Least Squares-Mean Group) estimator proposed by Pedroni (2001). Finally, we conclude the empirical analysis by conducting the panel causality test using the Dumitrescu and Hurlin (2012) approach. The findings of the empirical analysis indicate similar results, particularly regarding GDP per capita, as observed in the studies conducted by Safi et al. (2021) and Xie and Jamaani (2022).

The study consists of four sections, including the introduction. The first section provides an assessment of the subject and significance of the study. The second section presents a detailed review of the theoretical background and relevant literature pertaining to the subject. In the third section, we describe the dataset and the econometric tests, and provide an explanation of why we chose these tests, along with their mathematical foundations. The final section presents policy recommendations regarding the effectiveness of carbon pricing based on the empirical findings.

## Theory and Literature

Global warming and climate change have emerged as one of the most urgent issues to be addressed on the world agenda in the past century. Due to the heavy reliance on fossil fuel-based energy production, countries have often overlooked environmentally friendly production methods. As a result of the unsustainable consumption of fossil fuels, greenhouse gas emissions have increased worldwide. Furthermore, the accumulation of greenhouse gases in the atmosphere has led to a significant increase in global average temperatures. During this process, the melting of glaciers has accelerated, sea levels have risen, and clean water sources have diminished. These changes are just a few examples of the serious consequences of climate change. Therefore, it is of great importance to control greenhouse gas emissions and adopt sustainable solutions. In this context, countries are making efforts to transition to a growth model that prioritizes environmental sustainability and to reduce greenhouse gas emissions.

Climate change and global warming are widely recognized as significant global issues, and the international community is taking necessary steps to address them. In this context, the Kyoto Protocol was first signed in 1997 with the aim of reducing greenhouse gas emissions (UNFCCC, 1998). Later, in 2015, the Paris Climate Agreement was signed, aiming to keep global warming below 2 degrees Celsius below previous level and, if possible, to reduce it below 1.5 degrees (UNFCC, 2015). The Paris Climate Agreement aims to reduce greenhouse gas emissions permanently through a more comprehensive and long-term approach, unlike the Kyoto Protocol. The fundamental difference between the two agreements lies in the focus of the Kyoto Protocol on emission reduction, whereas the Paris Climate Agreement aims to eliminate emissions. Therefore, the selection of effective measures for emission reduction is of great importance. Carbon pricing is considered a key policy instrument in addressing the climate issue and achieving the goals of the Paris Agreement (Lilliestam et al., 2021).

Carbon pricing has emerged as an effective policy instrument in combating climate change, particularly in Nordic countries, starting from 1990. This policy aims to internalize the environmental costs associated with greenhouse gas emissions by imposing a financial cost on carbon emissions. The main objective of carbon pricing policy is to reduce carbon emissions, promote the adoption of low-carbon technologies in industries, increase the use of renewable energy sources, and facilitate the transition to cleaner and more sustainable production processes (Stiglitz et al., 2017:9-10). Countries can choose from various instruments for carbon pricing. However, countries generally implement carbon pricing policy through emissions trading schemes (ETS) or carbon taxes. But in recent years, countries have been using both policy instruments.

Governments can implement carbon pricing in various ways by using it as a fiscal policy instrument. This policy can be applied both at the economy-wide level and specifically targeted to certain sectors, industries, or local regions. Local conditions, strategic priorities, and needs are key considerations in determining the scope or coverage of the policy implementation. The success of carbon pricing policy in reducing environmental pollution depends on the level and coverage of the targeted carbon price. It is generally accepted that high carbon pricing, when implemented on a broad scale, is more effective in reducing emissions. Furthermore, the implementation of carbon pricing at high levels and broad scope also promotes low-carbon development in the long term (World Bank, 2022:12).

Many governments today are implementing carbon pricing policies to combat climate change and reduce increasing greenhouse gas emissions. Many countries are also planning to implement carbon pricing policies soon. The number of countries implementing carbon tax policies has been increasing, particularly after the 2015 Paris Climate Agreement. In the countries covered in this study, carbon taxes were implemented prior to the 1997 Kyoto Protocol and the 2015 Paris Climate Agreement. Carbon tax was first implemented and Finland in 1990. Then in 1991, it was implemented in Sweden and Norway, and in 1992, it started in Denmark. It can be observed that Sweden, Norway, and Finland have implemented carbon pricing policies extensively and have generated significant tax revenue as a result (Table 1).

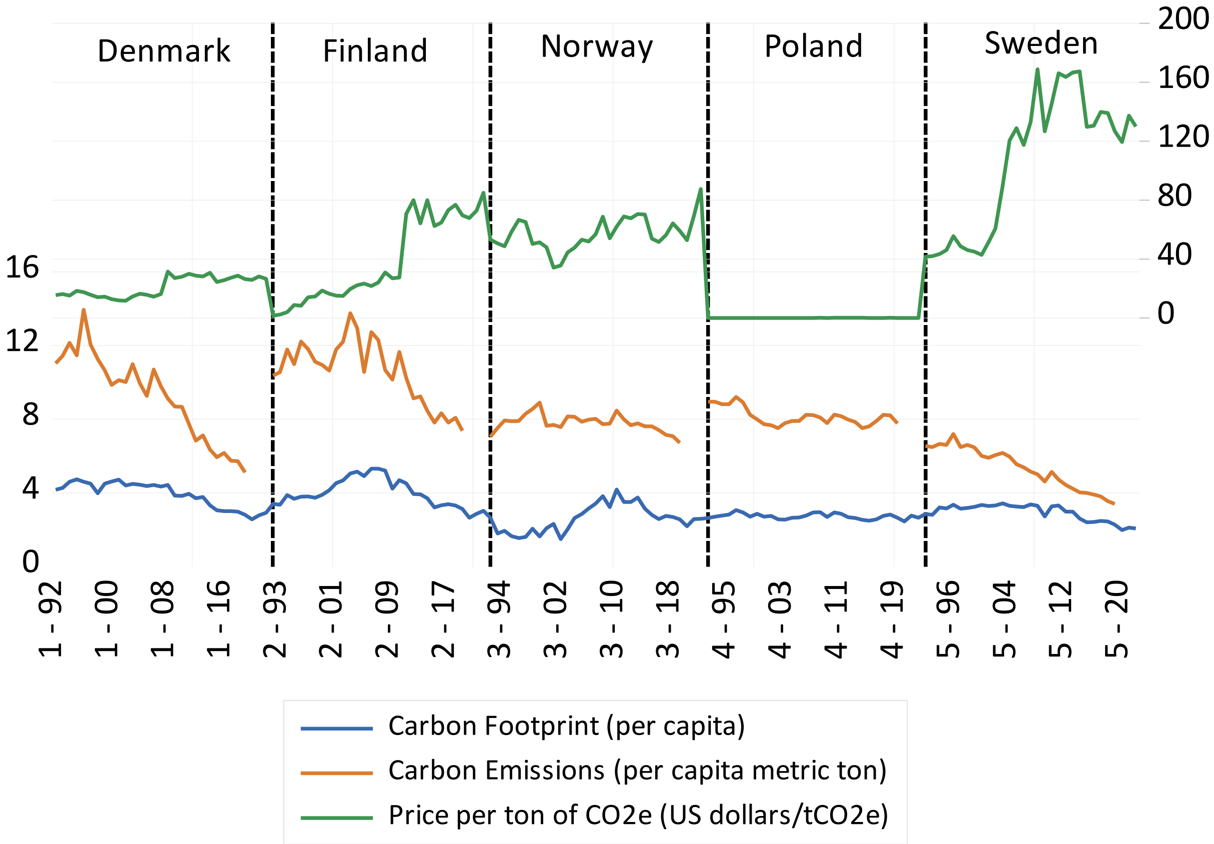
**Table 1.** Data on Carbon Tax

| **Countries** | **Implementation Year** | **Covered greenhouse gas emissions [MtCO2e][a]** | **Covered greenhouse gas emissions [%][b]** | **Covered national greenhouse gas emissions [%]** | **Price [US$/tCO2e][c]** | **Tax Revenue [Million US$][d]** |
|---|---|---|---|---|---|---|
| **Sweden** | 1991 | 24 | 0.05 | 86.7 | 125.55 | 2124.8 |
| **Denmark** | 1992 | 16 | 0.03 | 48.0 | 26.62/21.90 (fossil fuels and gases) | 492.8 |
| **Poland** | 1990 | 15 | 0.03 | 4.1 | 0.07/14.44 (co2/f-gases) | 6.48 |
| **Finland** | 1990 | 22 | 0.05 | 76.0 | 85.10/58.58 (transport/other fossil fuels) | 1706.7 |
| **Norway** | 1991 | 43 | 0.09 | 98.0 | 90.86/7.34 (general/reduced rate) | 1800.4 |

Source: Compiled using data from World Bank (2023). ([a] Based on official government sources and/or estimates, [b] Data taken from Köppl and Schratzenster (2022) for the year 2015, [c] represents 2023 carbon prices, [d] represents 2022 carbon tax revenues.

Graph 1 illustrates the trends over time in carbon emissions, carbon tax, and CF data of the countries. Graph 1 shows that carbon tax generally increases, while carbon emissions and CF decrease. Indeed, this situation suggests that carbon tax has a positive impact on environmental effects. Especially in Finland and Sweden, the successful implementation of carbon tax policies has contributed to the reduction of emissions and the decrease in CF. These data provide visual evidence that carbon tax implementations are important tools for sustainability and can be effective in achieving emission reduction targets. The increasing trend of carbon tax encourages the reduction of greenhouse gas emissions. A higher tax level increases the cost of activities that contribute to greenhouse gas emissions and encourages environmentally friendly practices such as reducing energy consumption from fossil fuels, transitioning to renewable energy sources, and improving energy efficiency.

**Figure 1** Trend of Carbon Price, Carbon Emissions, and CF




Source: Created based on data obtained from the World Bank (2023).

Table 2 summarizes multiple studies conducted about different countries on the impact of environmental taxes and carbon taxes on carbon emissions. The studies presented in Table 2 vary in their methodologies and findings, but most of them demonstrate the emission-reducing effect of environmental taxes. For example, Safi et al. (2021) provide evidence that an increase in environmental taxes reduces CO2 emissions both in the short and long term. In another study, Xie and Jamaani (2022) demonstrate the emission-reducing effect of environmental taxes in G-7 countries. However, some studies have obtained different results. For instance, Zaghdoudi and Maktouf (2017) find a positive relationship between environmental taxes and CO2 emissions. In another study, Saucedo et al. (2017) find a negative relationship between environmental taxes and CO2 emissions in a fixed-effects model. However, in the same study, they are unable to obtain a significant result based on the GMM estimator.

In the literature, we can see that researchers generally investigate the impact of environmental taxes on CO2 emissions. Despite this, we observe that there is a limited number of studies specifically conducted on carbon taxes. Examples of such studies include Zhang et al. (2020), Gugler et al. (2023), Gemechu et al. (2014), Bayer and Akalin (2012), Hajek et al. (2019), Lin and Li (2012), and Shmelev and Speck (2018). These studies demonstrate the mitigating effect of carbon taxation on CO2 emissions. In conclusion, the literature summary in Table 2 indicates that environmental taxes, including carbon taxation, generally have a mitigating effect on CO2 emissions. These findings reinforce the argument that policy instruments such as environmental taxes and carbon pricing are effective in reducing carbon emissions.

**Table 2.** Literature Review

| Authors | Period, Sample and Method | Dependent Variable | Findings |
|---|---|---|---|
| **Gugler et al. (2023)** | 2013-2015, England, Regression Discontinuity in Time | Co2 Emissions | The study has found a substantial decrease in CO2 emissions (26% or 38.6 MtCO2) in the British energy sector as a result of the carbon tax. |
| **Xie and Jamaani (2022)** | 1990-2020, G-7 Countries, Method of Moment Quantile Regression and Dumitrescu and Hurlin panel causality | Co2 Emissions | Environmental taxes reduce CO2 emissions by promoting renewable energy and green innovation. Additionally, an increase in GDP leads to an increase in CO2 emissions. |
| **Rufael and Weldemeskel (2022)** | 1994-2018, 18 Latin American and Caribbean countries, Panel Data Analysis | Co2 Emissions | Environmental taxes reduce CO2 emissions. |
| **Rafique et al. (2022)** | 1994-2016, 29 OECD Countries, Panel cointegration and FMOLS Method | Ecological Footprint (EF) | Environmental taxes reduce ecological footprint. |
| **Safi et al. (2021)** | 1990-2019, G-7, Panel Data Analysis | Co2 Emissions | An increase in environmental taxes reduces CO2 emissions both in the short and long term. |
| **Meireles et al. (2021)** | 2008-2018, EU Countries, Panel Data Analysis | Co2 Emissions | An increase in transportation taxes reduces CO2 emissions. |
| **Zhang et al. (2020)** | 2008-2016, China, Difference-in-difference (DID) Method | Co2 Emissions | They find that the Emission Trading System (ETS) policy has successfully reduced industrial CO2 emissions in China's seven pilot regions. |
| **Bayer and Aklin (2020)** | 2008-2016, EU Countries, Generalized Synthetic Control Method | Co2 Emissions | Even though carbon prices are low (less than 35 euros per ton), they are still effective in reducing CO2 emissions. |
| **Fernando (2019)** | 1990-2004, Denmark, Finland, Norway and Sweden, Synthetic Control Method | Co2 Emissions | In Norway and Sweden, CO2 taxes reduce CO2 emissions. However, in Denmark and Finland, the impact of CO2 taxes is not significant. |

| | | | |
|---|---|---|---|
| **Hajek et al (2019)** | 2005-2015, Denmark, Ireland, Finland, Sweden and Slovenia, Panel Data Analysis | Co2 Emissions | Based on the partial regression coefficient (-0.01158), an increase of one euro per ton in carbon tax would result in a reduction of annual per capita emissions by 11.58 kg. |
| **Shmelev and Speck (2018)** | 1960-2010, Sweden, Time Series Analysis | Co2 Emissions | Carbon tax, as well as energy taxes on coal and LPG, contribute to the reduction of carbon emissions. |
| **Zaghdoudi and Maktouf (2017)** | 1994-2014, OECD, Panel Threshold Regression | Co2 Emissions | There is a positive relationship between environmental taxes and CO2 emissions. |
| **Saucedo et al. (2017)** | 1994-2014, OECD, Static and Dynamic Panel Data Analysis | Co2 Emissions | According to the Fixed Effects Model, there is a negative relationship between environmental taxes and per capita CO2 emissions. However, in dynamic panel data analysis, the environmental tax does not have a significant effect on CO2 emissions. |
| **Kotnik et al. (2014)** | 1995-2010, EU-19 Countries, Panel Data Analysis | Co2 Emissions | Environmental taxes directly and indirectly impact greenhouse gas emissions negatively. |
| **Miller and Vela (2013)** | 1995-2010, Developed and developing countries, Panel Data Analysis | Co2 Emissions | An increase in environmental taxes reduces CO2 emissions. |
| **Morley (2012)** | 1995-2006, EU Countries and Norway, Panel Data Analysis | Co2 Emissions | There is a negative and significant relationship between environmental taxes and environmental pollution. |
| **Lin and Li (2011)** | 1990-2008, Denmark, Finland, Netherlands, Norway, Sweden, Difference in difference | Co2 Emissions | The results indicate that the carbon tax in Finland has led to a 1.7% decrease in per capita CO2 emissions. In Denmark, Sweden, and the Netherlands, negative effects on emissions were observed, but they were not statistically significant. |

## Data and Method

We construct two models to measure the impact of carbon tax on CF and CO2 emissions in the Baltic countries between 1992 and 2021. In the first model, the dependent variable is CO2 emissions, and in the second model, the dependent variable is CF. We include Denmark, Finland, Norway, Poland, and Sweden in the study. The reason for starting the analysis in 1992 is that it is the common year when the countries included in the study implemented the carbon tax. The reason for ending the first model in 2019 is the unavailability of data on CO2 emissions for subsequent years. Similarly, the reason for ending the second model in 2021 is the unavailability of data on per capita GDP for subsequent years at the time of conducting the analysis. Given the time frame of the study, we believe that the missing observation for the GDP per capita variable for one year would not significantly alter the analysis results. Equation 1 and 2 represent the first and second model respectively.

$$Co2_{it} = \alpha_0 + \alpha_1 GDPer_{it} + \alpha_2 CPrice_{it} + \in_{it} \quad (1)$$

$$CF_{it} = \alpha_0 + \alpha_1 GDPer_{it} + \alpha_2 CPrice_{it} + \in_{it} \quad (2)$$

In both equations, t, i and $\in_{it}$ represents the period, the cross-sections, the error term respectively. In equation 1, $Co2_{it}$ represents the dependent variable, which denotes per capita CO2 emissions. In equation 2, $CF_{it}$ represents the CF, symbolizing the dependent variable. In both equations, the independent variable, $CPrice_{it}$, represents the price paid per ton of carbon dioxide equivalent (CO2e) emissions in dollars. Lastly, the variable $GDPer_{it}$, included as a control variable in both equations, represents per capita gross domestic product (GDP). Table 3 shows detailed descriptions of the variables and the database from which they were obtained.

There are two carbon pricing rates in the World Bank (2023). The reason for this distinction is that countries apply carbon taxes to different sectors at different rates over time. In the study, we use a carbon price rate of 1 for each year between 1992 and 2021, considering complete data available for the countries' annual carbon pricing.

**Table 3.** Description of Variables

| Variables | Description | Unit | Source |
|---|---|---|---|
| $\boldsymbol{Co2_{it}}$ | Carbon emission | Per capita metric ton | World Bank |
| $\boldsymbol{CF_{it}}$ | Carbon footprint | Per capita | Global Footprint Data |
| $\boldsymbol{CPrice_{it}}$ | Price per ton of CO2e | US Dollars/tCO2e | World Bank |
| $\boldsymbol{GDPer_{it}}$ | Per capita GDP | US Dollars at 2015 prices | World Bank |

We use balanced panel data analysis, which allows for analysis with multiple cross-sectional units. To determine the impact of increasing globalization in recent years, we first conduct tests for cross-sectional dependence. These tests are Breush and Pagan (1980) LM, Pesaran (2004) $CD_{LM}$ and Pesaran et al. (2008) bias-adjusted LM. Panel data analyses conducted without considering cross-sectional dependence may lead to unreliable results (Destek et al., 2018; Hsiao, 2014:347).

The Breusch-Pagan LM test was developed to test the relationship between units based on the fixed effects model. The null hypothesis of the Breusch and Pagan (1980) LM test for cross-sectional dependence is that there is no cross-sectional dependence among the error terms of the regression equation. The LM statistic of the test is calculated through Equation 3.

$$CD_{lm} = T\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{p}_{iJ}^{2} \qquad (3)$$

Pesaran (2004) developed the CD (Cross-Sectional Dependence) test as an alternative to the Breusch-Pagan LM test. The CD test is designed to provide consistent results in both N>T and N<T situations, addressing the inconsistency issue of the LM test when N>T. In this test, the null hypothesis (Ho) is that there is no cross-sectional dependence. The test provides two test statistics based on pairwise correlation coefficients to be used in both balanced and unbalanced panels. Since this study is based on balanced panel data analysis, we only present the balanced panel CD test statistic in Equation 4.

$$CD = \sqrt{\frac{2T}{N(n-1)}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{p}_{iJ}\right) \qquad (4)$$

Pesaran, Ullah, and Yamagata (2008) developed a modified version of the Breusch and Pagan (1980) LM test by incorporating various adjustments and additions to address the issue of biased results when N>T. Equation 5 represents the calculation of the test statistic of this modified version (Pesaran et al. 2008).

$$LM_{adj} = \sqrt{\frac{2}{N(N-1)}}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\frac{(T-K)\hat{p}_{iJ}{}^{2}-\mu_{T_{ij}}}{v_{T_{ij}}} \qquad (5)$$

**Table 4.** Cross Section Dependency Test Results

| | **MODEL 1** | | **MODEL 2** | |
|---|---|---|---|---|
| **Test** | **İstatistik** | **p-değeri** | **İstatistik** | **p-değeri** |
| **LM** | 52.13 | 0.0000 | 51.12 | 0.0000 |
| **LM adj*** | 25.76 | 0.0000 | 26.21 | 0.0000 |
| **LM CD*** | 4.524 | 0.0000 | 5.705 | 0.0000 |

According to the results shown in Table 4, all three tests suggest the presence of cross-sectional dependence in the models. Determining the homogeneity of slope coefficients is necessary in order to identify the appropriate estimation methods in the subsequent stage.

Performing panel data analysis without considering slope heterogeneity can lead to biased results in panel cointegration and long-term coefficient estimation. We apply the delta homogeneity test developed by Pesaran and Yamagata (2008) to test for slope heterogeneity. In this test, there are two test statistics: $\hat{\Delta}$ test for large samples and $\hat{\Delta}_{adj}$ for small samples. The null hypothesis in this test is that the slope coefficient is homogeneous. Equation 6 shows the calculation of the test statistics.

$$\hat{\Delta} = \sqrt{N}\left(\frac{N^{-1}\hat{s}-k}{\sqrt{2k}}\right), \hat{\Delta}_{adj} = \sqrt{N}\frac{\left(N^{-1}\hat{S}-2k\right)}{\sqrt{var(\hat{z}_{iT}}} \qquad (6)$$

**Table 5.** Results of Pesaran and Yamagata (2008) Delta Test

| | MODEL 1 | | MODEL 2 | |
|---|---|---|---|---|
| **Test** | **İstatistik** | **p-değeri** | **İstatistik** | **p-değeri** |
| $\hat{\Delta}$ | 10.118 | 0.0000 | 11.402 | 0.0000 |
| $\hat{\Delta}_{adj}$ | 10.928 | 0.0000 | 12.247 | 0.0000 |

According to the results in Table 5, the slope coefficients of both models are heterogeneous. We consider this result in further tests.

According to Hsiao (2014), conducting statistical tests without considering the stationarity of the series leads to unreliable results and may cause the problem of spurious regression. In addition, to test for the presence of a long-term relationship between variables, it is necessary for the variables to be stationary in first differences. Due to all these reasons, we apply the cross-sectionally augmented ADF (CADF) and IPS (CIPS) tests, which are commonly used in the literature, developed by Pesaran (2007). These tests, which can be applied in the case of cross-sectional dependence, are referred to as second-generation unit root tests in the literature. The null hypothesis of the test is that the series is non-stationary. Equation 7 displays the calculation of the test statistic for the CADF unit root test.

$$\Delta Y_{it} = y_i + y_i Y_{i,t-1} + y_i \bar{X}_{t-1} + \sum_{i=0}^{p} Y_{it} \Delta \overline{Y_{t-\iota}} + \sum_{i=0}^{p} Y_{it} \Delta Y_{i,t-1} + \in_{it} \quad (7)$$

In equation 7, $Y_{i,t-1}$ represents the average lagged value of the parameter, and $\Delta Y_{i,t-1}$ represents the cross-sectional averages of the first-order differences. Additionally, equation 8 shows the calculation of the CIPS test statistic.

$$\widehat{CIPS} = \frac{1}{N} \sum_{i=1}^{n} CADF_i \quad (8)$$

In this equation, CIPS represents the cross-sectionally augmented IPS test, and CADF represents the cross-sectionally augmented Dickey-Fuller test.

Table 6. Results of the CADF and CIPS Unit Root Test

| | **Variables** | **CIPS** | | | | **CADF** | | | |
|---|---|---|---|---|---|---|---|---|---|
| | | Constant | | Constant + Trend | | Constant | | Constant + Trend | |
| | | I (0) | I (1) | I (0) | I (1) | I (0) | I (1) | I (0) | I (1) |
| **MODEL 1** | $Co2_{it}$ | -1.910 | -5.137* | -2.657 | -5.269* | -1.447 | -3.964* | -2.060 | -4.045* |
| | $CPrice_{it}$ | -1.944 | -4.835* | -1.929 | -4.860* | -1.870 | -3.420* | -1.812 | -3.404* |

| | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | $GDPer_{it}$ | -0.854 | -3.389* | -1.501 | -3.738* | -1.419 | -2.790* | -1.828 | -3.225* |
| **MODEL 2** | $CF_{it}$ | -1.096 | -5.341* | -2.806*** | -5.431* | -0.932 | -3.766* | -2.234 | -3.827* |
| | $CPrice_{it}$ | -2.030 | -5.068* | -1.866 | -5.084* | -1.968 | -3.504* | -1.660 | -3.498* |
| | $GDPer_{it}$ | -1.249 | -3.986* | -2.101 | -4.250* | -1.476 | -3.088* | -2.007 | -3.517* |

(*, **, and *** represent significance levels of 1%, 5%, and 10%, respectively. The CADF test statistic is the t-bar statistic used in balanced panel data.)

According to the CIPS and CADF test results in Table 6, all variables have unit roots at levels. However, after taking the first difference of the series, all variables become stationary. Therefore, all variables become stationary in first differences.

Finding the variables to be stationary after taking the first difference suggests the possibility of a long-term relationship among the series. For this purpose, we apply the panel cointegration test developed by Gegenbach, Urbain, and Westerlund (2016) to detect the long-term relationship among the series. This error correction-based test provides effective results against heterogeneity and cross-sectional dependence. Additionally, this test allows for unit-specific lag lengths. This test is based on the vector error correction model presented in equation 9, which utilizes the common factor structure (Gegenbach et al., 2016).

$$\Delta y_i = d\delta_{y.x_i} + \alpha_{y_i} y_{i,-1} + \omega_{i,-1} + \nu_i \pi_i + \varepsilon_{y.x_i} = \alpha_{y_i} y_{i,-1} + g_i^d + \varepsilon_{y.x_i} \qquad (9)$$

In the first stage of the test, the OLS estimation of the model is obtained for each unit, and the null hypothesis $H_0\colon \alpha_{y_i} = 0$ is tested using a t-test. The $(T-1-p) \times (T-1-p)$ dimensional projection matrix is defined as shown in equation 10.

$$M_A = I_{T-1-p} - A(A'A)A' \qquad (10)$$

Equation 11 and 12 represent the OLS estimator and variance of $\alpha y_i$, respectively.

$$\hat{\alpha}_{yi} = \frac{y'_{i,-1} M_{g_i^d} \Delta y_i}{y'_{i,-1} M_{g_i^d} y_{i,-1}} \qquad (11)$$

$$\sigma^2_{\hat{\alpha}_{y_i}} = \frac{\sigma^2_{y.xi}}{y'_{i,-1} M_{g_i^d} y_{i,-1}} \qquad (12)$$

The expansion of $\sigma^2_{y.xi}$ in the numerator of equation 12 is $T^{-1}\left(\Delta y_i - \hat{\alpha}_{yi} y_{i,-1}\right)' M_{g_i^d}(\Delta y_i - \hat{\alpha}_{yi} y_{i,-1})$. Equation 13 shows the calculation of the t-statistic.

$$t_{c_i} = t_{\alpha_{yi}} = \frac{\hat{\alpha} y_i}{\hat{\sigma}_{\hat{\alpha} y_i}} \tag{13}$$

The panel test statistic is obtained by taking the average of the unit-specific t-statistics. Equation 14 represents the calculation of the panel test statistic.

$$\bar{t}_c = \frac{1}{N} \sum_{i=1}^{N} t_{c_i} \tag{14}$$

The null hypothesis of the cointegration test is "$H_0$: there is no cointegration relationship among the series." Table 7 presents the results of the cointegration tests for both models.

Table 7. Results of Gegenbach et al. (2016) Panel Cointegration Test

| $\boldsymbol{d.y}$ | **Coefficients** | **T-bar Statistics** | **P-value** | |
|---|---|---|---|---|
| $\boldsymbol{Y(t-1)}$ | -0.787 | -3.741 | <=0.01 | **MODEL 1** |
| $\boldsymbol{Y(t-1)}$ | -0.602 | -3.372 | <=0.05 | **MODEL 2** |

(We determine the lag length heterogeneously.)

Table 7 shows that in both models, the null hypothesis $H_0$ is rejected at the 0.01 and 0.05 significance levels, respectively. In other words, there exists a cointegration relationship among the variables.

To estimate the coefficients of the long-term relationship among the variables obtained from the cointegration test, we use the DOLSMG estimator, which provides robust results in the presence of cross-sectional dependence and heterogeneity (Pedroni, 2001). In the DOLSMG estimator proposed by Pedroni (2001), it is based on the model presented in Equation 15.

$$Y_{it} = \mu_i = \beta_i X_{it} + u_{it} \quad i = 1, \ldots, N \quad t = 1, \ldots, T \tag{15}$$

The model in Equation 15 indicates that the cointegration model is heterogeneous across units. In the first stage, the cointegration model given in Equation 15 is estimated for each unit using the dynamic ordinary least squares (DOLS) approach with the inclusion of lagged values and leads. Subsequently, the results are combined for the entire panel using the Pesaran and Smith (1995) MG approach.

$$\hat{\beta}_{DOLSMG} = N^{-1} \left[ \sum_{i=1}^{N} \left( \sum_{t=1}^{T} (Z_{it} Z'_{it}) \right)^{-1} \right] \left( \sum_{t=1}^{T} (Z_{it} \overline{y_{it}}) \right) \tag{16}$$

In Equation 16, $Z_{it}$ represents the vector of explanatory variables, and $Y_i = Y_{it} - \bar{Y}_i$. Equation 17 illustrates the calculation of the DOLSMG test statistic.

$$t_{\hat{\beta}_{DOLS,i}} = N^{-1} \sum_{t=1}^{T} t_{\beta_{DOLS,i}} \quad (17)$$

**Table 8.** Results of the Long-Term Coefficient Estimation

| Country | Variables | MODEL 1 | | MODEL 2 | |
|---|---|---|---|---|---|
| | | **Beta** | **t-statistic** | **Beta** | **t-statistic** |
| **Denmark** | $CPrice_{it}$ | 0.05440 | 1.137 | 0.05122 | 5.661* |
| | $GDPer_{it}$ | 0. 0004265 | 0.9397 | 0.0000674 | -0.7776 |
| **Finland** | $CPrice_{it}$ | -0.02247 | -2.124* | -0.007417 | -2.535* |
| | $GDPer_{it}$ | 0.000436 | 3.942* | 0.0002532 | 7.586* |
| **Norway** | $CPrice_{it}$ | -0.05828 | -3.393* | -0.04008 | -2.691* |
| | $GDPer_{it}$ | -0.000254 | -2.094* | -0.0001252 | -1.224 |
| **Poland** | $CPrice_{it}$ | -0.06263 | -6.19* | -0.0131 | -2.048* |
| | $GDPer_{it}$ | 0.000135 | 2.323* | 0.0000982 | 2.63* |
| **Sweden** | $CPrice_{it}$ | -0.01531 | -8.454* | -0.008183 | -5.631* |
| | $GDPer_{it}$ | 0.000192 | 5.195* | 0.0000316 | 1.124 |
| **The entire panel** | $CPrice_{it}$ | -0.02086 | -8.507* | -0.003511 | -3.295* |
| | $GDPer_{it}$ | 0.0001872 | 4.609* | 0.0000381 | 4.176* |

(The critical value for α=0.05 is 1.96)

According to the findings in Table 8, the beta coefficients of the panel groups are statistically significant because their absolute t-statistic values are greater than the critical t-table value. In Model 1, a unit increase in the $CPrice_{it}$ decreases $Co2_{it}$ by 0.02086 units, while a unit increase in the $GDPer_{it}$ increases $Co2_{it}$ by 0.0001872 units. In Model 2, a unit increase in the $CPrice_{it}$ decreases $CF_{it}$ by 0.003511 units, while a unit increase in the $GDPer_{it}$ variable increases $CF_{it}$ by 0.0000381 units. According to the other findings in Table 8, for Model 1, the coefficient of the $CPrice_{it}$ is negative and statistically significant in Finland, Norway, Poland, and Sweden. In other words, an increase in the $CPrice_{it}$ leads to a decrease in $Co2_{it}$ in these countries. On the other hand, the findings of Model 2 are consistent with the findings in Model 1. These results imply that carbon pricing is an effective policy tool in reducing carbon emissions and carbon footprint. In the study, only in Model 2, carbon pricing increases carbon footprint in Denmark. Zaghdoudi and Maktouf (2017) also obtained similar results regarding the relationship between environmental taxes and CO2 emissions. Lin and Li (2011) detect the negative impact of carbon tax on carbon emissions, but this relationship is statistically insignificant. The reason of this could be Denmark's exemptions of the manufacturing industry and energy sectors in the implementation of carbon tax. Lastly, the findings for the overall panel support the existing literature.

For the detection of causality between series, we use the Dumitrescu and Hurlin (2012) test, which is an extension of the Granger causality test. This test is commonly used in the literature for panels with cross-sectional dependence and heterogeneity (Lopez and Weber, 2017). Equation 18 shows the basic regression of the Dumitrescu and Hurlin (2012) panel causality test.

$$y_{it} = \alpha_i + \sum_{k=1}^{k} \gamma_i^{(k)} Y_{it-k} + \sum_{k=1}^{K} \beta_i^{(k)} X_{it-k} + \varepsilon_{it} \tag{18}$$

In Equation 18, $X_{i,t}$ and $Y_{i,t}$ represent two stationary variables. The parameters in the equation may vary across cross sections, but they are assumed to be time-invariant. K represents the lag length. The hypotheses of the Dumitrescu and Hurlin (2012) panel causality test are as follows:

- $H_0$: There is no causality.
- $H_a$: There is causality.

To test the null hypothesis, the Wald Test statistics calculated for each cross section are summed and averaged. Equation 19 shows the calculation of the Wald test statistics.

$$\bar{W}_{N,T} = \frac{1}{N} \sum_{i=1}^{N} W_{i,T} \tag{19}$$

In the Dumitrescu and Hurlin (2012) panel causality test, when T is large, the test statistic specified in equation 20 is used.

$$\bar{Z}_{N,T} = \sqrt{\frac{N}{2K}} \left( \bar{w}_{N,T} - K \right) = \overrightarrow{T, N \to \infty}\ N(0,1) \tag{20}$$

Table 9. Results of the Dumitrescu and Hurlin (2012) Panel Causality Test

| **MODEL 1** | | | **MODEL 2** | | |
|---|---|---|---|---|---|
| **Direction** | **Test Statistics** | **P-value** | **Direction** | **Test Statistics** | **P-value** |
| $\boldsymbol{CPrice_{it}}$=>$\boldsymbol{Co2_{it}}$ | 2.7682 | 0.0263 | $\boldsymbol{CPrice_{it}}$=>$\boldsymbol{CF_{it}}$ | 4.3356 | 0.0000 |
| $\boldsymbol{Co2_{it}}$=>$\boldsymbol{CPrice_{it}}$ | 1.4640 | 0.1432 | $\boldsymbol{CF_{it}}$=>$\boldsymbol{CPrice_{it}}$ | 1.2288 | 0.2191 |
| $\boldsymbol{GDPer_{it}}$=> $\boldsymbol{Co2_{it}}$ | 0.2316 | 0.8168 | $\boldsymbol{GDPer_{it}}$=>$\boldsymbol{CF_{it}}$ | -0.7238 | 0.4692 |
| $\boldsymbol{Co2_{it}}$=>$\boldsymbol{GDPer_{it}}$ | 2.2314 | 0.0257 | $\boldsymbol{CF_{it}}$=>$\boldsymbol{GDPer_{it}}$ | 0.4855 | 0.6273 |
| $\boldsymbol{CPrice_{it}}$=>$\boldsymbol{GDPer_{it}}$ | 5.6878 | 0.0000 | $\boldsymbol{CPrice_{it}}$=>$\boldsymbol{GDPer_{it}}$ | 2.5374 | 0.0112 |
| $\boldsymbol{GDPer_{it}}$=>$\boldsymbol{CPrice_{it}}$ | -0.8786 | 0.3796 | $\boldsymbol{GDPer_{it}}$=>$\boldsymbol{CPrice_{it}}$ | -0.9636 | 0.3434 |

(In the Dumitrescu and Hurlin panel causality test, we make the series stationary by taking their first differences. We set the lag value to 1. Since T>N, we use the $\bar{Z}_{N,T}$ statistic.)

In model 1, we detect the causality from the $CPrice_{it}$ to the $Co2_{it}$, from the $Co2_{it}$ to the $GDPer_{it}$, and from the $CPrice_{it}$ to the $GDPer_{it}$. In model 2, there are casualties from the $CPrice_{it}$ to $CF_{it}$ and from the $CPrice_{it}$ to the $GDPer_{it}$. These results confirm the findings of the DOLSMG estimator for the overall panel. In another words, there is a unidirectional causal relationship between carbon pricing and the dependent variables.

**Conclusion**

This study investigates the impact of carbon pricing policy on carbon emissions and carbon footprint. We can say that while the impact of environmental taxes on carbon emissions is widely studied in the literature, its effect on carbon footprint is generally overlooked. However, we observe that studies focusing on carbon taxes are not quantitatively sufficient. In these limited studies, we also observe that the time period is kept very narrow. In this context, the study stands out from its counterparts by investigating the impact of carbon pricing policy on both carbon emissions and carbon footprint over a wide time period. Furthermore, the use of balanced panel data methodology enables the investigation of long-term relationships among variables. In these contexts, we believe that the study makes a significant contribution to the literature. We summarize the limitations of the study as follows:

• Including only five countries: Many countries have implemented carbon pricing policies over time. However, due to the selection of balanced panel data analysis method to investigate the long-term relationship between variables, we are compelled to include countries that adopted this policy in the early stages and have similarities in terms of geography and culture.

• Period: We aim for the study to cover the years 1992-2022. However, due to the unavailability of CO2 emissions data beyond 2019 and per capita GDP data beyond 2021 in the relevant databases, we cannot include the recent years in the time range.

The findings of the analysis are particularly important in shaping governments' carbon tax policies. Currently, many countries such as Italy, Hungary, Greece, and Turkey are considering implementing carbon taxes. This study provides a projection for governments that are considering carbon taxation. The reason why these countries have not implemented carbon taxes yet could be their concerns about the negative impact on production costs. At this point, these countries can incentivize companies to transition to more environmentally friendly production methods using the revenue generated from carbon taxes. Furthermore, the revenue from carbon taxes can significantly increase public income and assist in financing government expenditures. Additionally, when it comes to carbon taxes, carrot-based approaches may be more recommended than stick-based approaches for taxpayers. In this context, in new implementations, a more inclusive, incentivizing, and internalizing approach can be offered, considering the perspectives of taxpayers. This way, resistance to carbon taxes can be reduced.

Carbon taxation can play a crucial role in ensuring justice in collecting taxes from different social classes, beyond its other functions. This can be achieved by taxing luxury goods and services that contribute to carbon emissions. Carbon emissions resulting from personal use of private jets, carbon emissions from large fuel-consuming vehicles, and carbon emissions from energy-intensive homes are examples of micro-scale areas that can be subject to taxation. In this way, higher taxes can be imposed on carbon emissions originating from luxury consumption favored by the capitalist class. This policy proposal offers an approach that aims to reconcile environmental sustainability and social justice goals.

Consequently, it is important for such a policy design to consider factors such as feasibility, effectiveness, and social acceptance. Through detailed analysis and studies, the effects and feasibility of the policy proposal should be evaluated in collaboration with stakeholders. Fair determination of tax rates is crucial for achieving the intended environmental impacts and ensuring social acceptance. Additionally, factors such as the economic effects of the policy design, industrial transformation processes, and social protection mechanisms should be considered. Therefore, future research should focus on the detailed examination and improvement of the policy proposal, considering the opinions of relevant stakeholders. The more consensus and agreement are reached, the greater the inclusiveness of environmental taxes becomes.

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## Statements & Declarations

**Ethics approval**

All authors approved to consider ethical standards.

**Consent to participate**

All authors involved in this paper have confirmed their participation in the preparation of the paper.

**Consent for publish**

All authors approved the paper for publication.

**Author contributions**

All authors have contributed equally to the study. All authors read and approved the final manuscript.

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