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Are per capita CO_2 emissions increasing among OECD countries? A test of trends and breaks

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Abstract

We empirically analyze the trend characteristics of per capita CO_2 emissions in OECD countries from 1971 to 2009. We use a statistically robust procedure, which is valid regardless of whether per capita CO_2 emissions are trend stationary or contain a stochastic trend, to test for the presence of a deterministic trend and a structural break in the trend. Our results suggest that the trend in per capita CO_2 emissions shifts downward or is reversed for a number of OECD countries either after the 1970s oil shocks or during the early- to mid-2000s.

JEL classification: C12, Q53
Keywords: carbon dioxide, trend, structural break, robust test

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1 Introduction

Carbon dioxide (CO_2) is a primary greenhouse gas emitted by human activities and there is an increasing concern about its influence on the climate system. Despite the introduction of climate policies at both domestic and international levels, whether emission targets can be achieved within a reasonable time frame depends on whether a negative non-stochastic emission-time relationship (i.e., deterministic trend) exists. Understanding the trend properties of per capita CO_2 emissions is therefore critical to policy makers.

In this paper, we use a statistically robust procedure to empirically explore: one, whether a deterministic trend is present in the per capita CO_2 emissions in OECD countries; two, whether this relationship has been subject to a discrete structural shift; and three, whether a similar trend pattern can be observed among OECD countries. Hypothesis testing for the presence of a deterministic trend and for the stability of the trend in emission series is complicated. The validity of such tests depends on whether the existence of a deterministic trend can be correctly discriminated from a stochastic trend process. The differentiation, however, is empirically difficult for two reasons. First, whether the deterministic trend is embedded in a stationary $I(0)$ series or non-stationary $I(1)$ series is *a priori* unknown. Second, a permanent shock to a time series is observed in a similar way as a structural shift in the deterministic trend (Perron, 1989). Perron and Yabu (2009) have recently developed a methodology that tackles these two difficulties and in this paper we adopt this statistically robust procedure to evaluate the trend characteristics of per capita CO_2 emissions.

We focus on the OECD countries because they are responsible for more than

40% of world emissions and their per capita emissions are, on average, 129% higher than world average (IEA, 2012). If per capita CO_2 emissions have successfully responded to any emission control policies implemented by a country or through multinational cooperation, we expect to observe evidence of a downward trend of per capita CO_2 emissions or a similar trend shift among the countries participating in the cooperation.

Correctly identifying the nature and extent of the deterministic trend, if any exists, is also crucial in the estimation of the carbon Kuznets curve, which postulates an inverted U-shaped relationship between per capita CO_2 emissions and per capita income (Wagner, 2008). Evidence for the inverted U-shaped relationship between the two variables can be falsely identified if the estimated model fails to differentiate the emission-income relationship and the emission-time relationship (Vollebergh et al., 2009). To this end, our study also provides a first, but critical, step towards the correct estimation of the carbon Kuznets curve.

2 Econometric methods and data

2.1 Robust test for deterministic linear trend

Let $\{y_t;\, t = 1, \ldots, T\}$ denote annual time series of per capita CO_2 emissions. We start with a model that fits a linear trend line to the per capita CO_2 emissions data, that is:

$$y_t = \mu + \beta t + u_t,\; u_t = \rho u_{t-1} + e_t \tag{1}$$

where e_t satisfy all assumptions in Perron and Yabu (2009) and $-1 < \rho \leq 1$ so that u_t can be $I(0)$ or $I(1)$. Our aim is to test the null hypothesis $H_0 : \beta = 0$ against the alternative $H_1 : \beta \neq 0$. Given u_t may be $I(1)$, the standard t-test based on OLS estimation of (1) is inappropriate for assessing H_0.

Perron and Yabu (2009) propose a modified t-test using quasi-feasible GLS (FGLS) estimation of the deterministic linear trend which allows u_t to be either $I(0)$ or $I(1)$. Their test-statistic takes the form $t_\beta^F = \hat{\beta}^F / se(\hat{\beta}^F)$, where $\hat{\beta}^F$ and $se(\hat{\beta}^F)$ are obtained from the OLS estimation of:

$$y_t - \hat{\rho} y_{t-1} = (1 - \hat{\rho})\mu + \beta[t - \hat{\rho}(t-1)] + (u_t - \hat{\rho} u_{t-1}) \tag{2}$$

in which $\hat{\rho}$ is the estimate of ρ in the regression $\hat{u}_t = \rho \hat{u}_{t-1} + \sum_{i=1}^{k} \zeta_i \Delta \hat{u}_{t-i} + e_{tk}$ using a truncated weighted symmetric least-squares with a lag length k selected by a modified AIC criterion and $\hat{u}_t$ are the OLS residuals in (1). The limiting distribution of t_β^F is standard normal under both $I(0)$ and $I(1)$ errors. Hence, the test rejects H_0 at level α if $|t_\beta^F|$ is greater than the desired standard normal critical value, regardless of whether u_t are $I(0)$ or $I(1)$.

2.2 Structural shift in the deterministic linear trend

The tests based on (1) may fail to identify a linear trend if a structural break, including the reversal of the time trend, occurs during the sample period. We therefore extend model (1) to allow for the presence of a break in the deterministic trend, such that:

$$y_t = \mu + \beta t + \gamma DT_t + u_t, \; u_t = \rho u_{t-1} + e_t \tag{3}$$

where $DT_t = 1(t > T_1)(t - T_1)$ with $1(C) = 1$ if condition C holds and 0 otherwise. The break date $T_1 = \lfloor \tau^* T \rfloor$ is unknown, $\tau^* \in (0, 1)$ and $\lfloor x \rfloor$ is the integer part of the number x. Our goal is to test the hypotheses: (i) $H_\beta : \beta = 0$ (no deterministic pre-break trend), (ii) $H_\gamma : \gamma = 0$ (no structural shift in trend) and (iii) $H_{\beta+\gamma} : \beta + \gamma = 0$ (no deterministic post-break trend).

To assess the hypotheses in (i)-(iii), we follow a two-step procedure. In the first step, we identify the break date by assuming that the true break fraction, τ^*, lies in $[0.1, 0.9]$. We regress (3) using OLS for all $\tau \in [0.1, 0.9]$ and select $\hat{\tau}$ that minimizes the sum of squared residuals. We then compute the estimated break date as $\hat{T}_1 = \lfloor \hat{\tau} T \rfloor$. In the second step, we use Perron and Yabu's (2009) FGLS procedure for assessing hypotheses (i)-(iii). The estimated equation takes the form:

$$y_t - \tilde{\rho} y_{t-1} = (1 - \tilde{\rho})\mu + \beta[t - \tilde{\rho}(t-1)] + \gamma[DT_t - \tilde{\rho} DT_{t-1}] + (u_t - \tilde{\rho} u_{t-1}) \qquad (4)$$

where $\tilde{\rho}$ is estimated in the regression $\tilde{u}_t = \rho \tilde{u}_{t-1} + \sum_{i=1}^{k} \zeta_i \Delta \tilde{u}_{t-i} + e_{tk}$ using the same truncation methodology as in (2). Here, $\tilde{u}_t$ are the OLS residuals from (3). The test-statistics for assessing hypotheses (i)-(iii) are expressed as:

$$\begin{aligned} t_\beta^{FB} &= \hat{\beta}^{FB}/se(\hat{\beta}^{FB}), \quad t_\gamma^{FB} = \hat{\gamma}^{FB}/se(\hat{\gamma}^{FB}), \\ t_{\beta+\gamma}^{FB} &= (\hat{\beta}^{FB} + \hat{\gamma}^{FB})/se(\hat{\beta}^{FB} + \hat{\gamma}^{FB}) \qquad (5) \end{aligned}$$

where $\hat{\beta}^{FB}$ and $\hat{\gamma}^{FB}$ are the OLS estimators of β and γ in (4). The corresponding tests reject H_0 at level α when the absolute value of the test-statistic is greater than the desired standard normal critical value, regardless of whether u_t are $I(0)$ or $I(1)$.

2.3 Data

The annual per capita CO_2 emissions from fuel combustion in OECD countries are retrieved from International Energy Agency (IEA, 2012). The sample spans from 1971 to 2009. Among the 34 OECD member countries, Estonia and Slovenia are excluded due to lack of data.

3 Results and discussion

Table 1 presents the estimates of the trend coefficient β in equation (1), which disregards a possible structural shift in the deterministic trend of per capita CO_2 emissions. We report the test-statistics t_β^F proposed by Perron and Yabu (2009) and t_{HAC}, which is calculated using the heteroskedasticity autocorrelation consistent (HAC) standard errors. The t_{HAC}-test suggests that, for most OECD countries, per capita CO_2 emissions are either positively or negatively time-related at a statistically significant level. By contrast, the t_β^F-test, which is robust to either $I(0)$ or $I(1)$ errors, suggests that about half of the OECD countries exhibit neither a positive nor a negative trend. This significant discrepancy between the two test results highlights the fact that the t_{HAC}-test, which is often referred to as a *robust* test procedure, suffers from severe size distortions when the errors are strongly correlated or are $I(1)$.

In general, we cannot observe a similar trend pattern among the OECD countries. For instance, all OECD countries in the Asia and Oceania region exhibit a positive estimate of the trend coefficient. In particular, the Republic of Korea, Australia and Israel have the three most positive trends among the OECD countries. By contrast, for many European countries, the estimated trend coefficients are negative. Nevertheless, the t^F_β-test suggests that, for most countries, the negative trend coefficients are statistically insignificant, except for Germany, Sweden, Switzerland and the United Kingdom.

Table 2 reports the estimates of the parameters β, γ and $\beta + \gamma$ in equation (3), which incorporates a structural shift in the trend. For many OECD countries, we find that the trend in per capita CO_2 emissions either shifts downward ($\hat{\gamma} < 0$) or is reversed ($\hat{\beta} > 0$ and $\hat{\beta}+\hat{\gamma} < 0$) during the period 1971 to 2009. Only three countries, Chile, the Republic of Korea and Turkey, exhibit a statistically significant steeper positive trend in the post-break period ($\hat{\beta} + \hat{\gamma} > 0$). Further, all of the statistically significant trend shifts occur either in the post-oil shock period of 1979 to 1984 or the early to mid 2000s.

For the downward trend shift, these two periods may correspond to the times when the countries have switched the source of electricity production, which is a major contributor of CO_2 emissions, from the carbon-intensive fossil fuels to the less carbon intensive sources, such as nuclear power and renewables. For instance, Czech Republic, Germany, Hungary and Slovak Republic exhibit a statistically significant negative trend shift between 1979 to 1982. For all four countries, the share of electricity production from nuclear power has increased from the end of the 1970s to the beginning of the 1990s, whereas the coal share of the total electricity generation has decreased constantly during the same period.

We, however, note that not all countries that have replaced the fossil fuels with low-carbon electricity sources exhibit a statistically significant downward shift in the trend in per capita CO_2 emissions. For instance, France has increased its share of electricity production from nuclear power extensively from 6% in 1971 to 77.2% in 2009. Similarly, Finland has increased the share of nuclear power and renewables from 0% to 47.1%. Nevertheless, the downward shifts in the trend in per capita CO_2 emissions for both countries are statistically insignificant.

4 Conclusions

Our estimation shows that most of the statistically significant structural shifts in the trend of per capita CO_2 emissions in OECD countries are downward. As a result, the significant post-break trends are mainly negative. We postulate that the downward shift in the trend of per capita CO_2 emissions would be associated with changes in the type of fuels used for generating electricity. Burke (2010) recently finds that countries increase the share of low-carbon fuels in the electricity production as their income increases. However, further empirical research is needed to formally test for the relationship between changes in the energy mix and per capita CO_2 emissions.

Our empirical results also demonstrate that both the size and sign of the deterministic trend coefficient vary significantly across the countries and time periods. This suggests that a simple assumption about the emissions-time relationship in the panel estimation of the carbon Kuznets curve can be problematic.

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Table 1: OLS estimates of the trend coefficient and test statistics

	$\hat{\beta}$	t_{HAC}	t_{β}^{F}
Americas			
Canada	13.00	0.82	-0.01
Chile	65.61	7.57***	1.56
Mexico	35.76	6.51***	2.85***
United States	-62.85	-3.70***	-1.07
Asia and Oceania			
Australia	193.41	16.49***	3.03***
Israel	140.43	16.99***	2.73***
Japan	61.10	5.54***	0.72
New Zealand	84.18	9.32***	1.41
Republic of Korea	262.34	21.37***	4.25***
Europe			
Austria	51.07	6.24 ***	0.52***
Belgium	-58.19	-3.29***	-0.81
Czech Republic	-159.48	-5.52***	-1.21
Denmark	-47.27	-2.74***	-1.74*
Finland	60.97	3.85***	0.29***
France	-80.17	-5.85***	-1.43
Germany	-113.62	-9.41***	-2.24**
Greece	152.20	24.03***	3.70***
Hungary	-60.67	-3.84***	-0.28
Iceland	-0.60	-0.08	-0.27
Ireland	109.21	9.16***	0.67
Italy	54.44	9.41***	0.61
Luxembourg	-576.15	-6.14***	-1.56
Netherlands	4.26	0.50	0.26
Norway	58.83	7.48***	0.79***
Poland	-91.35	-4.05***	-0.46
Portugal	125.66	12.33***	2.10**
Slovak Republic	-99.79	-4.40***	-1.01
Spain	98.60	10.54***	1.51
Sweden	-136.45	-8.58***	-2.34**
Switzerland	-10.19	-2.84***	-2.87***
Turkey	62.69	21.83***	3.05***
United Kingdom	-68.57	-11.54***	-1.86***

Note: *, **, *** denote rejection at the 10%, 5% and 1% levels, respectively.

Table 2: Estimation results of the structural break model

	Break year	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\beta}+\hat{\gamma}$
Americas				
Canada	1991	-46.20	128.20	82.05
Chile	1984	-28.30	129.80**	101.48***
Mexico	1981	129.80***	-114.40***	15.37
United States	1983	-193.50	172.40	-21.12
Asia and Oceania				
Australia	2005	202.90***	-259.90***	-57.07
Israel	2002	154.90***	-150.60*	4.38
Japan	1983	-29.30	119.30	90.00
New Zealand	2005	94.80**	-290.50**	-195.77*
Republic of Korea	1984	170.60*	126.80	297.37***
Europe				
Austria	1989	16.00	65.20	81.17
Belgium	1984	-181.20	170.00	-11.21
Czech Republic	1982	152.80	-394.90*	-242.14**
Denmark	1996	10.20	-208.00	-197.79
Finland	2004	78.00	-317.60	-239.56
France	1988	-176.20*	166.50	-9.68
Germany	1979	112.70*	-257.60***	-144.91***
Greece	2005	163.20***	-302.50***	-139.31
Hungary	1979	259.80***	-364.80***	-104.97***
Iceland	2005	8.30	-243.90*	-235.57*
Ireland	2005	126.80	-483.00***	-356.16**
Italy	2005	66.10**	-321.10***	-254.99***
Luxembourg	1983	-1628.80***	1388.90**	-239.88
Netherlands	1978	92.10	-97.20	-5.10
Norway	2004	64.40	-104.30	-39.87
Poland	1979	363.20**	-517.40***	-154.19*
Portugal	2004	144.70***	-354.30***	-209.67*
Slovak Republic	1981	257.50**	-434.70***	-177.23**
Spain	1988	61.60	64.20	125.79
Sweden	1984	-286.40***	207.20	-79.18
Switzerland	2005	-6.40	-102.70	-109.19
Turkey	1983	36.60	34.40	71.01***
United Kingdom	1983	-106.30	49.80	-56.51

Note: *, **, *** denote rejection at the 10%, 5% and 1% levels, respectively. The significance of the parameters are based on t_{β}^{FB}-, t_{γ}^{FB}- and $t_{\beta+\gamma}^{FB}$-statistics, respectively.

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