Persistence of methane emission in OECD countries for 1750-2014: A fractional integration approach

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Abstract

The statistical structure of the methane emissions in a group of 36 OECD countries has been examined in this work for the time period 1750-2014 using techniques based on fractional integration. This allows us to determine the degree of persistence of the series and the potential presence of trends in the data. Our results indicate that all series are highly persistent, with orders of integration above 1 in the majority of the cases. Linear (positive) trends are observed in approximately half of the cases. One of the implications of these findings is that policies designed for decreasing methane emissions will have a long term impact in these countries.

JEL Classification: C25; Q50; Q53; Q58

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1. Introduction

Methane or CH_4 emission is a key component of greenhouse gases (GHGs). Anthropogenic methane emission accounts for almost 20% of radiative forcing of GHGs in the globe since pre-industrial periods, which makes it the second biggest contributor after CO_2 emission (Environmental Protection Agency, 2012; Fernández-Amador et al., 2020). Global methane emissions increased more than 10-fold between 1750 and 2000. It further increased by more than 23% between 2000-2014 (Hoesly et al., 2018).

The increase in methane emission is partly due to the multiple sources that generate it, which includes energy production, distribution and use; agriculture; and waste management. Agricultural activities are the biggest source of anthropogenic methane emissions as they account for almost 25% of total methane emissions (International Energy Agency, 2020). The second biggest emitter is the energy sector because biofuels, coal, oil and natural gas generate methane emissions (International Energy Agency, 2020). Therefore, if not effectively mitigated, methane releases and leaks could weaken the greenhouse gas benefit natural gas provides and portend a major setback for the climate. Decreasing methane emissions is essential to avoid the worst impacts of climate change. Moreover, methane is a major contributor to the detrimental air pollutant, tropospheric ozone. Although methane does not often cause direct harm to crop production or human health, ozone accounts for almost a million premature respiratory deaths in the globe, annually (Gil-Alana and Solarin, 2018).

Due to the significance of methane emission, several aspects of greenhouse gas have been examined in the literature including its economic impact (Crow et al., 2019); its environmental impact and responding strategies (Cheng et al., 2011). Wang et al. (2019) focussed on the evolution of methane emissions in global supply chains. Ma et al. (2018) explored the socioeconomic factors that determine the growth of methane emissions and showed that household income and consumption were the key determinants of methane growth, while changes in efficiency are the most important factor offsetting methane emissions. Park (2004) evaluated the use of biofilters to lessen atmospheric methane emissions. However, the persistence of methane emissions has been largely ignored in the extant literature. Persistence is a measure of the extent at which short term shocks (resulting from new government initiatives) is able to generate permanent future changes.

The importance of investigating the unit root process of methane emissions can hardly be overemphasized. If there is evidence for the persistence of methane emission, the consequence is that the shocks will have a permanent impact on methane emission and the methane emission level will not retreat back to its steady long-term growth path. Conversely, if methane emission is mean reverting, this means that methane emission levels move back to their long-run trend path after a shock, and the effect of the initiatives on methane emission will merely be momentary. Hence, changes in methane emission are related to transient fluctuations and a shock should have no impact on long-run methane emission (Zerbo and Darné, 2019). Moreover, a result indicating that methane emission is mean reverting might suggest efficiency gains during the sample period and that the ecological effect can be probably contained (Christidou et al., 2013). Whether methane emission is persistent is significant for validating numerous important theories in energy economics including The Stochastic Impacts by Regression on Population, Affluence, and Technology or (STIRPAT) model or the environmental Kuznets curve (EKC) hypothesis. The failure to appropriately test for the persistence of the relevant variables in the models might yield invalid inferences in favour of STIRPAT or EKC.

The main aim of this research is to contribute to the literature on the persistence of environmental indicators in two unique ways. The first contribution of this study is that it examines the persistence of methane emissions in 36 OECD countries for the 1750–

2014 period, which is likely to provide new information on a series that has been virtually overlooked in the extant literature. Most of the existing papers have focused on the persistence of CO_2 emissions (Christidou et al., 2013; Gil-Alana et al., 2017; Zerbo and Darné, 2019; Gil-Alana and Monge, 2020). There are several differences between CO_2 emissions and methane emissions, which might make the results of CO_2 emissions not applicable to methane emissions. The lifetime of carbon dioxide in the atmosphere is much longer than methane, but methane is more effective at trapping radiation than carbon dioxide (International Energy Agency, 2020). Unlike carbon dioxide, methane has commercial value because the extra methane captured can frequently be monetised directly (International Energy Agency, 2020). Secondly, we use the fractional integration technique that is more flexible than the common approaches based on integer differentiation, using ARMA-ARIMA models. Thirdly, we have also utilised a very long-time-series data in order to benefit from the larger sample size. For instance, a historical sample gives more reliable results with greater precision and power. Working with a long sample leads to large samples, which is important since it helps to remove the outliers in the sample. The use of a historical sample presents the unique advantage of covering different economic periods, which is crucial given the EKC literature demonstrates the importance of economic performance on the evolution of CO_2 emissions (Gil-Alana and Solarin, 2018)

We have focussed on OECD countries due to several reasons. With an aggregate gross domestic product of US$52 trillion (in 2010 prices), OECD countries accounted for 63% of the global GDP in 2018 (World Bank, 2020). Secondly, OECD countries have experienced growth in their methane emissions during most of the periods under investigation. The methane emissions increased by almost 800% between 1750 and 2014 in the OECD countries (Hoesly et al., 2018). Thirdly, OECD countries account for 19%

of the global methane emissions in 2014 (Hoesly et al., 2018). Fourthly, methane emission mitigation technologies in OECD countries are often more effective than those available in non-OECD countries. Many non-OECD countries look to OECD countries when formulating their methane emission mitigation policies.

The remainder of the paper is arranged as follows. The literature review in the next section. The methodology and data are discussed in the third section. The findings are provided in the fourth section. Conclusions are in the last section.

2. Literature review

Due to the roles that GHGs play in climate change as well as global warming, analysing the dynamic behaviour of these gases has become an important research area. Thus, an increasing number of papers have examined the persistence of pollutants, but the majority of these papers have focussed on the unit root process of CO_2 emissions. The papers on CO_2 emissions cut across different regions and have adopted different econometric techniques. For instance, Christidou et al. (2013) used nonlinear panel unit root tests to show the stationarity of CO_2 emissions in 33 countries during 1870–2006. Tiwari et al. (2016) used the nonlinear panel and time series unit root tests to examine the CO_2 emissions per capita in 35 African nations for 1960–2009 period. The panel tests showed that CO_2 emissions per capita follow a stationary process for all the countries, while the time series output showed mean reversion in 29% of the countries under investigation. Gil-Alana et al. (2017) examined the persistence of CO_2 emissions in BRICS and G7 countries. The results revealed that CO_2 emissions are not mean reverting in most cases. The only exceptions were the US, the UK and Germany where there was mean-reverting evidence for CO_2 emissions. Zerbo and Darné (2019) used a sequential testing procedure to show that non-stationarity existed in CO_2 emissions per capita of 29 countries including

OECD countries as well as Brazil, China, India and South Africa over the period 1960–2014. Gil-Alana and Trani (2019) examined the persistence of CO_2 emissions in 20 EU member countries, China and the US for the period of 1960-2013. Using fractional integration, the results provided evidence for mean reversion in the case of the UK, while an explosive behaviour characterizes the CO_2 emissions of such countries as Bulgaria, Greece, Italy Portugal and Spain. Gil-Alana and Monge (2020) analysed the persistence of global CO_2 emissions for the period, 1880-2015. The univariate results indicated that the CO_2 emissions were not mean reverting in all cases.

There are also studies on the persistence of other pollutants. For example, Lee and List (2004) investigated the persistence of NO_x (Nitrogen Oxide) emissions in the U.S. for the 1900–1994 period. The empirical findings provided mixed evidence for the series. McKitrick (2007) investigated the stationarity of different pollutants in the U.S. such as NOx and VOC (Volatile Organic Compounds). The results suggested that both series were nonstationary. Sidneva and Zivot (2014) examined the nonstationary of NO_x and VOC emissions. The results provided evidence for the stationarity of the series. Gil-Alana and Solarin (2018) examined the time series properties of the global and per capita NO_x and VOC emissions in the U.S. for the period, 1940-2014. Using fractional integration approaches, the results suggested that the two series were very persistent.

3. Data and Methodology

3.1 Data

The data for methane emissions (in kilotonnes or kt) in 36 OECD countries for the 1750-2014 period has been generated from the Joint Global Change Research Institute[1][2]. The

[1] The data is available in https://zenodo.org/record/3606753#.XqLGoMgzbIU.
[2] The list is based on OECD membership as at 1/4/2020. The list was generated from https://www.oecd.org/about/document/list-oecd-member-countries.htm.

only exception is Australia with a dataset running from 1781-2014. The data was generated within the Community Emissions Data System (CEDS) and has several advantages over the other sources of methane emissions data which often lack reproducibility and uncertainty estimates, have inadequate temporal resolution and do not have consistent procedures across emission types (Hoesly et al., 2018). CEDS utilises current emission inventories, emission factors, and driver/activity data to compute yearly country emissions over time and there are numerous phases involved in the calculation process.[3] The first phase involves data collection as well processing of data into a consistent timescale and format. In the second phase, default emissions from 1960 to 2014 for several OECD countries are computed using driver and emission factor data (where emission is equal to emission factors multiplied by the drivers). The drivers used in the computation include energy consumption (which is used as a driver for emissions from fuel combustion), population (which is used as a noncombustion emissions driver). In the third stage, default estimates are scaled so as to be consistent with current emission inventories where plausible, available and complete. In the fourth stage, scaled emission estimates are stretched back to 1750 to generate final emissions in each country. Finally, emissions are scrutinised and collated to generate data for release as well as analysis (Hoesly et al., 2018).

The descriptive statistics of the series are reported in Table 1 and it is shown that with an average of 10,089.76 kilotonnes of methane emission, the U.S. (which has the largest economy in the globe) was the biggest emitter among the OCED countries during the investigated period. The U.S. is also among the countries with largest coefficient of variation. The Skewness statistic show that most of the series are not skewed. The empirical analysis in Section 4 will be based on the logged transformed data.

[3] The country totals do not include international shipping or aircraft emissions.

Table 1: Descriptive statistics of the series

Series	Mean	Coeff. of variation	Skweness
AUSTRALIA	1,546.92	1.19	0.91
AUSTRIA	180.77	0.51	0.93
BELGIUM	318.23	0.31	0.47
CANADA	1,145.14	1.18	1.31
SWITZERLAND	102.72	0.55	0.80
CHILE	197.90	1.17	1.37
CZECH REP.	448.99	0.56	0.87
GERMANY	2,311.13	0.67	0.83
DENMARK	143.33	0.63	0.64
SPAIN	522.54	0.73	1.48
ESTONIA	38.88	0.55	1.08
FINLAND	115.19	0.76	0.84
FRANCE	2,598.54	0.38	0.24
U.K.	2,889.40	0.60	0.17
GREECE	184.46	0.75	1.03
HUNGARY	347.38	0.41	0.40
IRELAND	931.84	0.36	0.72
ICELAND	6.99	1.03	1.18
ISRAEL	26.44	1.30	2.44
ITALY	886.92	0.61	1.22
JAPAN	659.96	0.70	0.99
KOREA	480.85	0.74	1.28
LITHUANIA	135.05	0.39	0.64
LUXEMBOURG	10.77	0.49	0.62
LATVIA	87.64	0.32	0.53
MEXICO	1,151.05	1.30	1.65
NETHERLANDS	426.35	0.73	1.22
NORWAY	92.44	0.74	0.91
NEW ZEALAND	451.46	1.05	0.96
POLAND	1,110.54	0.59	0.80
PORTUGAL	195.27	0.77	1.40
SLOVAKIA	102.25	0.56	1.29

SLOVENIA	38.23	0.79	1.02
SWEDEN	144.00	0.62	0.99
TURKEY	736.10	0.90	1.89
USA	10,089.76	1.07	0.86

The series are in their original forms (in kilotonnes or kt)

3.2 Methodology

As earlier mentioned, we use fractional integration. This technique has the advantage of being more general than the standard methods that impose an integer degree of differentiation, usually 1, if the series is nonstationary. Thus, the order of integration of a series may be a fractional value and this allows for a large degree of flexibility in the dynamic specification of the model. In fact, by using this approach, we can consider a variety of modelling approaches, including short memory processes, if d = 0; stationary long memory models if 0 < d < 0.5; nonstationary mean reverting approaches if 0.5 ≤ d < 1; unit roots if d = 1, or even processes with d > 1. Moreover, the fact that we work with the logged transformed data implies that if the order of integration is higher than 1, the growth rates (which are the first differences of the logged series) will display a long memory (d > 0) pattern.

In the empirical work conducted in the following section, we also permit for deterministic terms such as an intercept and a time trend. Thus, the model under examination is:

$$y_t \; = \; \alpha \; + \; \beta t \; + \; x_t \, , \qquad (1 - L)^d x_t \; = \; u_t \, , \qquad t \; = \; 1, \, 2, \, \ldots, \qquad (1)$$

where y_t is each of the time series we observe (in logs); α and β are unknown coefficients referring respectively to a constant and a linear time trend, and d is the potentially fractional differencing parameter. We use a testing approach that is based on a simple version of the tests of Robinson (1994). These tests include a variety of approaches including for instance, those based on seasonal and cyclical integration. In this paper, however, and based on the annual nature of the data, we simply use the one that is characterized by the polynomial in (1). See Gil-Alana and Robinson (1997) for the description of the test statistic, which is based on the Whittle function expressed in the frequency domain.

4. Empirical results

We start this section by presenting the results under the assumption that u_t in (1) is uncorrelated. We display the results for three model specifications in Table 2. The first one, in column 2, reports the estimates of d (and the associated 95% confidence band) for the case where α and β are supposed to be zero, i.e., we do not include any deterministic term in the model; in the second case, in column 3, we suppose that only β is equal to zero, i.e., the model includes a constant but not a time trend; finally, in the last column, we report the estimates of d when α and β are jointly estimated with the differencing parameter d. The selected models according to these terms are reported in bold in the tables.

Table 2: Estimates of d: White noise disturbances

Series	No terms	An intercept	A time trend
AUSTRALIA	0.93 (0.86, 1.02)	1.79 (1.64, 1.98)	**2.03 (1.49, 2.19)**
AUSTRIA	0.98 (0.90, 1.09)	**1.53 (1.47, 1.60)**	1.53 (1.48, 1.60)
BELGIUM	0.98 (0.90, 1.09)	**1.08 (1.01, 1.18)**	1.08 (1.01, 1.18)
CANADA	0.97 (0.89, 1.08)	1.25 (1.19, 1.33)	**1.24 (1.18, 1.32)**
SWITZERLAND	0.98 (0.90, 1.02)	**1.54 (1.48, 1.62)**	1.53 (1.47, 1.61)
CHILE	0.98 (0.89, 1.08)	1.19 (1.12, 1.38)	**1.21 (1.14, 1.30)**
CZECH REP.	0.98 (0.90, 1.08)	**1.24 (1.16, 1.33)**	1.24 (1.16, 1.33)
GERMANY	0.98 (0.90, 1.08)	**1.11 (1.03, 1.23)**	1.11 (1.03, 1.23)
DENMARK	0.93 (0.86, 1.08)	**1.45 (1.39, 1.53)**	1.45 (1.39, 1.52)
SPAIN	0.98 (0.90, 1.08)	1.10 (1.06, 1.15)	**1.12 (1.07, 1.17)**

ESTONIA	0.98 (0.90, 1.08)	**1.30 (1.20, 1.44)**	1.30 (1.20, 1.44)
FINLAND	0.97 (0.89, 1.08)	1.53 (1.48, 1.60)	**1.52 (1.47, 1.59)**
FRANCE	0.98 (0.90, 1.08)	**1.08 (1.01, 1.18)**	1.08 (1.01, 1.18)
U.K.	0.97 (0.89, 1.07)	0.95 (0.91, 1.00)	**0.95 (0.91, 1.00)**
GREECE	0.97 (0.89, 1.07)	**1.39 (1.31, 1.50)**	1.41 (1.33, 1.51)
HUNGARY	0.98 (0.90, 1.08)	**1.24 (1.17, 1.32)**	1.23 (1.17, 1.31)
IRELAND	0.99 (0.91, 1.09)	**1.40 (1.34, 1.49)**	1.40 (1.34, 1.49)
ICELAND	1.30 (1.22, 1.39)	**1.52 (1.44, 1.61)**	1.52 (1.44, 1.61)
ISRAEL	1.03 (0.96, 1.12)	1.20 (1.15, 1.26)	**1.20 (1.15, 1.27)**
ITALY	0.98 (0.90, 1.08)	**1.37 (1.32, 1.43)**	1.37 (1.32, 1.43)
JAPAN	0.98 (0.90, 1.08)	**1.39 (1.29, 1.51)**	1.39 (1.29, 1.51)
KOREA	0.98 (0.90, 1.08)	**1.31 (1.23, 1.41)**	1.32 (1.24, 1.41)
LITHUANIA	0.98 (0.90, 1.07)	**1.47 (1.36, 1.62)**	1.47 (1.36, 1.62)
LUXEMBOURG	1.00 (0.92, 1.10)	1.32 (1.23, 1.43)	**1.31 (1.21, 1.42)**
LATVIA	0.98 (0.90, 1.08)	**1.52 (1.38, 1.70)**	1.52 (1.38, 1.70)
MEXICO	0.98 (0.90, 1.08)	1.32 (1.25, 1.41)	**1.33 (1.27, 1.42)**
NETHERLANDS	0.98 (0.89, 1.08)	**1.59 (1.52, 1.68)**	1.59 (1.52, 1.68)
NORWAY	0.97 (0.89, 1.07)	1.21 (1.15, 1.28)	**1.21 (1.16, 1.28)**
NEW ZEALAND	1.00 (0.89, 1.08)	1.29 (1.22, 1.35)	**1.25 (1.21, 1.31)**
POLAND	0.98 (0.90, 1.08)	0.98 (0.92, 1.06)	**0.98 (0.91, 1.06)**
PORTUGAL	0.98 (0.89, 1.08)	1.25 (1.18, 1.34)	**1.27 (1.21, 1.36)**
SLOVAKIA	0.98 (0.90, 1.08)	1.19 (1.12, 1.27)	**1.19 (1.13, 1.27)**
SLOVENIA	0.97 (0.88, 1.07)	1.18 (1.13, 1.24)	**1.19 (1.14, 1.25)**
SWEDEN	0.98 (0.90, 1.08)	**1.47 (1.41, 1.54)**	1.46 (1.41, 1.53)
TURKEY	0.99 (0.91, 1.09)	1.25 (1.19, 1.32)	**1.26 (1.21, 1.34)**
USA	0.98 (0.89, 1.08)	1.09 (1.04, 1.16)	**1.08 (1.03, 1.14)**

The values are the estimates of d using Robinson (1994). We report in parenthesis the 95% confidence band of the non-rejection values for d. In bold, the significant models for each series.

Table 3 focusses on the estimated coefficients of the selected models, reporting the estimate of the differencing parameter along with the coefficients of the deterministic terms, i.e., the constant and the time trend.

Table 3: Estimated coefficients of the selected models: White noise disturbances

Series	d	Intercept	Time trend
AUSTRALIA	2.03 (1.49, 2.19)	4.9522 (151.87)	0.7107 (15.19)
AUSTRIA	1.53 (1.47, 1.60)	4.4594 (759.19)	---
BELGIUM	1.08 (1.01, 1.18)	5.3000 (223.99)	---
CANADA	1.24 (1.18, 1.32)	3.5568 (225.51)	0.0172 (5.11)
SWITZERLAND	1.54 (1.48, 1.62)	3.7401 (841.63)	---
CHILE	1.21 (1.14, 1.30)	2.9565 (180.39)	0.0130 (4.33)
CZECH REP.	1.24 (1.16, 1.33)	5.1528 (184.34)	---
GERMANY	1.11 (1.03, 1.23)	6.6607 (146.34)	---
DENMARK	1.45 (1.39, 1.53)	3.8660 (743.80)	---
SPAIN	1.12 (1.07, 1.17)	5.3433 (462.88)	0.0071 (5.40)
ESTONIA	1.30 (1.20, 1.44)	2.8014 (137.47)	---
FINLAND	1.52 (1.47, 1.59)	3.1389 (480.45)	0.0078 (1.66)
FRANCE	1.08 (1.01, 1.18)	7.1744 (239.44)	---
U.K.	0.95 (0.91, 1.00)	6.5207 (146.57)	0.0052 (2.49)
GREECE	1.39 (1.31, 1.50)	4.0413 (606.21)	0.0054 (1.72)
HUNGARY	1.23 (1.17, 1.31)	4.8580 (301.78)	---
IRELAND	1.40 (1.34, 1.49)	6.6109 (589.15)	---
ICELAND	1.52 (1.44, 1.61)	0.2267 (21.85)	---
ISRAEL	1.20 (1.15, 1.27)	2.0535 (57.05)	0.0116 (1.85)
ITALY	1.37 (1.32, 1.43)	6.0012 (835.66)	---
JAPAN	1.39 (1.29, 1.51)	5.6489 (372.44)	---
KOREA	1.31 (1.23, 1.41)	5.5152 (353.60)	---
LITHUANIA	1.47 (1.36, 1.62)	4.1896 (291.02)	---
LUXEMBOURG	1.31 (1.21, 1.42)	1.4014 (185.87)	0.0064 (2.81)
LATVIA	1.52 (1.38, 1.70)	3.8820 (219.01)	---
MEXICO	1.33 (1.27, 1.42)	5.2704 (359.15)	0.0100 (2.05)
NETHERLANDS	1.59 (1.52, 1.68)	5.0474 (635.86)	---
NORWAY	1.21 (1.16, 1.28)	3.1312 (283.16)	0.0073 (3.63)
NEW ZEALAND	1.25 (1.21, 1.31)	3.2783 (414.64)	0.0144 (8.14)
POLAND	0.98 (0.91, 1.06)	5.9465 (182.06)	0.0061 (3.40)
PORTUGAL	1.27 (1.21, 1.36)	4.1644 (472.71)	0.0073 (3.37)
SLOVAKIA	1.19 (1.13, 1.27)	3.8333 (248.05)	0.0051 (2.02)
SLOVENIA	1.19 (1.14, 1.25)	2.4030 (215.01)	0.0070 (3.80)
SWEDEN	1.47 (1.41, 1.54)	3.9833 (536.40)	---

TURKEY	1.26 (1.21, 1.34)	5.7356 (406.40)	0.0080 (2.41)
USA	1.08 (1.03, 1.14)	5.4062 (221.01)	0.0181 (7.95)

The values in parenthesis in the third and fourth columns are the corresponding t-values.

Most of the estimated values of d are significantly higher than 1, implying long memory (d > 0) in the growth rate series (first differences). Evidence of unit roots is only found in the cases of the UK and Poland, and for the UK, this hypothesis is close to be rejected in favor of mean reversion. Time trends are observed in half of the countries examined (18), in all cases with positive coefficients running these values from 0.0051 (Slovakia) to 0.7107 (Australia).

In Table 4 we permit autocorrelated errors. That is, u_t in (1) may display some type of weak dependence structure. However, instead of imposing a parametric model (e.g., an ARMA(p, q) as is standard in the literature) we use here a non-parametric approach that was developed by Bloomfield (1973) and that accommodates very well in the context of the tests of Robinson (1994), (Gil-Alana, 2004). Bloomfield (1973) showed that the logarithm of the spectral density function of an AR process can be well approximated by the log of the following function,

$$f(\lambda)\ \frac{\sigma^2}{2\pi}\left(2\sum_{j=1}^{m}\tau_j\cos(\lambda j)\right). \qquad (2)$$

where σ^2 is the variance of the error term and m is an integer value. This method has also the advantage that it is stationary across all its values unlike what happens in the AR case. The results using this method are reported across Tables 4 and 5. The estimates of d are now slightly smaller, though, as in the previous table, most of the estimates are significantly higher than 1. Evidence of unit roots is obtained now for Belgium, Germany, Estonia, France and Latvia, this hypothesis being rejected in favour of d > 1 in the rest of the cases, and there are 14 countries with significant positive time trend coefficients (Table 5).

Table 4: Estimates of d: Autocorrelated disturbances

Series	No terms	An intercept	A time trend
AUSTRALIA	0.98 (0.86, 1.12)	1.40 (1.30, 1.54)	**1.32 (1.25, 1.42)**
AUSTRIA	0.94 (0.81, 1.10)	**1.78 (1.64, 1.96)**	1.76 (1.63, 1.96)
BELGIUM	0.94 (0.82, 1.11)	**1.02 (0.91, 1.14)**	1.02 (0.91, 1.14)
CANADA	0.93 (0.79, 1.12)	1.41 (1.29, 1.57)	**1.39 (1.29, 1.55)**
SWITZERLAND	0.94 (0.81, 1.11)	**1.63 (1.53, 1.79)**	1.63 (1.51, 1.76)
CHILE	0.94 (0.81, 1.13)	1.19 (1.09, 1.34)	**1.22 (1.11, 1.38)**
CZECH REP.	0.95 (0.80, 1.16)	**1.22 (1.10, 1.36)**	1.22 (1.10, 1.36)
GERMANY	0.93 (0.80, 1.09)	0.90 (0.81, 1.00)	**0.90 (0.80, 1.00)**
DENMARK	0.95 (0.81, 1.12)	**1.63 (1.51, 1.82)**	1.62 (1.50, 1.82)
SPAIN	0.96 (0.84, 1.12)	**1.38 (1.28, 1.57)**	1.44 (1.33, 1.58)
ESTONIA	0.93 (0.81, 1.13)	1.11 (0.99, 1.26)	**1.11 (0.99, 1.26)**
FINLAND	0.93 (0.80, 1.11)	**1.89 (1.73, 2.08)**	1.84 (1.69, 2.04)
FRANCE	0.95 (0.83, 1.11)	**1.03 (0.93, 1.17)**	1.03 (0.93, 1.16)
U.K.	0.93 (0.81, 1.12)	**1.21 (1.14, 1.29)**	1.21 (1.13, 1.30)
GREECE	0.94 (0.82, 1.11)	1.29 (1.19, 1.45)	**1.34 (1.23, 1.50)**
HUNGARY	0.95 (0.83, 1.12)	**1.27 (1.15, 1.40)**	1.27 (1.14, 1.40)
IRELAND	0.97 (0.86, 1.13)	**1.45 (1.34, 1.59)**	1.45 (1.34, 1.58)
ICELAND	1.34 (1.20, 1.52)	**1.56 (1.42, 1.79)**	1.56 (1.42, 1.77)
ISRAEL	1.06 (0.96, 1.22)	**1.53 (1.39, 1.73)**	1.54 (1.40, 1.73)
ITALY	0.96 (0.82, 1.12)	**1.65 (1.54, 1.80)**	1.65 (1.55, 1.78)
JAPAN	0.95 (0.81, 1.12)	**1.17 (1.09, 1.28)**	1.18 (1.09, 1.28)
KOREA	0.96 (0.82, 1.13)	**1.26 (1.16, 1.41)**	1.28 (1.18, 1.43)
LITHUANIA	0.95 (0.81, 1.12)	**1.27 (1.13, 1.46)**	1.26 (1.13, 1.45)
LUXEMBOURG	0.98 (0.83, 1.17)	1.30 (1.11, 1.52)	**1.25 (1.10, 1.49)**
LATVIA	0.95 (0.83, 1.12)	**1.05 (0.93, 1.19)**	1.04 (0.92, 1.19)
MEXICO	0.96 (0.83, 1.13)	1.27 (1.20, 1.39)	**1.31 (1.23, 1.42)**
NETHERLANDS	0.95 (0.80, 1.11)	**1.59 (1.50, 1.73)**	1.59 (1.50, 1.73)
NORWAY	0.93 (0.80, 1.10)	1.34 (1.23, 1.47)	**1.34 (1.26, 1.47)**

NEW ZEALAND	0.95 (0.81, 1.13)	1.56 (1.45, 1.69)	**1.49 (1.39, 1.61)**
POLAND	0.96 (0.83, 1.13)	1.11 (0.99, 1.28)	**1.11 (1.00, 1.29)**
PORTUGAL	0.96 (0.84, 1.13)	1.22 (1.13, 1.36)	**1.27 (1.17, 1.41)**
SLOVAKIA	0.95 (0.83, 1.13)	1.19 (1.09, 1.32)	**1.21 (1.09, 1.33)**
SLOVENIA	0.94 (0.80, 1.11)	**1.54 (1.40, 1.71)**	1.54 (1.42, 1.71)
SWEDEN	0.93 (0.80, 1.12)	**1.64 (1.52, 1.78)**	1.63 (1.53, 1.76)
TURKEY	0.97 (0.85, 1.13)	**1.35 (1.24, 1.54)**	1.37 (1.27, 1.55)
USA	0.95 (0.81, 1.12)	1.31 (1.22, 1.45)	**1.27 (1.18, 1.40)**

The values are the estimates of d using Robinson (1994). We report in parenthesis the 95% confidence band of the non-rejection values for d. In bold, the significant models for each series.

Table 5: Estimated coefficients of the selected models: Autocorrelated disturbances

Series	d	Intercept	Time trend
AUSTRALIA	1.32 (1.25, 1.42)	4.5269 (116.45)	0.1024 (8.12)
AUSTRIA	1.78 (1.64, 1.96)	4.4594 (856.99)	---
BELGIUM	1.02 (0.91, 1.14)	5.3005 (223.47)	---
CANADA	1.39 (1.29, 1.55)	3.5574 (234.22)	0.0164 (2.47)
SWITZERLAND	1.63 (1.53, 1.79)	3.7401 (879.56)	---
CHILE	1.22 (1.11, 1.38)	2.9566 (180.80)	0.0129 (4.10)
CZECH REP.	1.22 (1.10, 1.36)	5.1528 (183.42)	---
GERMANY	0.90 (0.80, 1.00)	6.6565 (149.76)	0.0046 (2.84)
DENMARK	1.63 (1.51, 1.82)	3.8660 (805.94)	---
SPAIN	1.38 (1.28, 1.57)	5.3476 (541.28)	---
ESTONIA	1.11 (0.99, 1.26)	2.7990 (132.96)	0.0044 (1.94)
FINLAND	1.89 (1.73, 2.08)	3.1427 (613.24)	---
FRANCE	1.03 (0.93, 1.17)	7.1748 (238.81)	---
U.K.	1.21 (1.14, 1.29)	6.5238 (160.59)	---
GREECE	1.34 (1.23, 1.50)	4.0409 (601.89)	0.0060 (2.58)
HUNGARY	1.27 (1.15, 1.40)	4.8579 (304.84)	---
IRELAND	1.45 (1.34, 1.59)	6.6110 (600.56)	---
ICELAND	1.56 (1.42, 1.79)	0.2267 (22.21)	---
ISRAEL	1.53 (1.39, 1.73)	2.0597 (67.76)	---
ITALY	1.65 (1.54, 1.80)	6.0012 (972.08)	---
JAPAN	1.17 (1.09, 1.28)	5.6488 (361.01)	---
KOREA	1.26 (1.16, 1.41)	5.5152 (350.20)	---

LITHUANIA	1.27 (1.13, 1.46)	4.1897 (275.24)	---
LUXEMBOURG	1.25 (1.10, 1.49)	1.4013 (183.72)	0.0064 (3.74)
LATVIA	1.05 (0.93, 1.19)	3.8832 (207.06)	---
MEXICO	1.31 (1.23, 1.42)	5.2702 (358.21)	0.0103 (2.32)
NETHERLANDS	1.59 (1.50, 1.73)	5.0474 (635.84)	---
NORWAY	1.34 (1.26, 1.47)	3.1317 (293.13)	0.0067 (1.80)
NEW ZEALAND	1.49 (1.39, 1.61)	3.2779 (444.22)	0.0148 (3.11)
POLAND	1.11 (1.00, 1.29)	5.9472 (185.01)	0.0058 (1.65)
PORTUGAL	1.27 (1.17, 1.41)	4.1644 (472.91)	0.0073 (3.37)
SLOVAKIA	1.21 (1.09, 1.33)	3.8333 (248.99)	0.0051 (1.81)
SLOVENIA	1.54 (1.40, 1.71)	2.4068 (257.95)	---
SWEDEN	1.64 (1.52, 1.78)	3.9832 (581.39)	---
TURKEY	1.35 (1.24, 1.54)	5.7398 (429.63)	---
USA	1.27 (1.18, 1.40)	5.4068 (234.37)	0.0175 (3.07)

The values in parenthesis in the third and fourth columns are the corresponding t-values.

Table 6 summarizes the results in terms of the time trends. We see that under the two assumptions for the error term, Australia occupies the first place, and Canada, the USA, New Zealand, Mexico, Chile are in the top places in the two cases. On the other hand, evidence of no trends in the two cases is found for Austria, Belgium, Switzerland, Czech Republic, Denmark, France, Hungary, Ireland, Iceland, Italy, Japan, Korea, Lithuania, Latvia, the Netherlands and Sweden.

Table 6: Summary table: Time trend coefficients

No autocorrelation		Autocorrelation	
No terms	Time trends	No terms	Time trends
Austria	Australia (0.7107)	Austria	Australia (0.1024)
Belgium	Canada (0.0172)	Belgium	U.S.A. (0.0175)
Switzerland	New Zealand (0.0144)	Switzerland	Canada (0.0164)
Czech Rep.	Chile (0.0130)	Czech Rep.	New Zealand (0.0148)
Germany	Israel (0.0116)	Denmark	Chile (0.0129)
Denmark	Mexico (0.0100)	Spain	Mexico (0.0103)

Estonia	U.S.A. (0.0081)	Finland	Portugal (0.0073)
France	Turkey (0.080)	France	Norway (0.0067)
Hungary	Finland (0.078)	U.K.	Luxembourg (0.0064)
Ireland	Norway (0.073)	Hungary	Greece (0.0060)
Iceland	Portugal (0.073)	Ireland	Poland (0.058)
Italy	Spain (0.071)	Iceland	Slovakia (0.0051)
Japan	Slovenia (0.070)	Israel	Germany (0.0046)
Korea	Luxembourg (0.064)	Italy	Estonia (0.0044)
Lithuania	Poland (0.061)	Japan	
Latvia	Greece (0.054)	Korea	
Netherlands	U.K. (0.052)	Lithuania	
Sweden	Slovakia (0.051)	Latvia	
		Netherlands	
		Sweden	
		Turkey	

In Table 7 the classification is made according to the level of persistence. We observe that Germany, France, Belgium and the UK display low degrees of persistence (lower 8) in the two cases of uncorrelated and autocorrelated errors, while Sweden, Iceland, Finland, Austria, Switzerland and the Netherlands are in the top 8 in the two cases.

Table 7: Summary table: Persistence

No autocorrelation		Autocorrelation	
Lower d	Upper d	Lower d	Upper d
U.K. (0.95)	Estonia (1.30)	Germany (0.90)	Australia (1.32)
Poland (0.98)	Korea (1.31)	Belgium (1.02)	Greece (1.34)
Belgium (1.08)	Luxembourg (1.31)	France (1.03)	Norway (1.34)
France (1.08)	Mexico (1.33)	Latvia (1.05)	Turkey (1.35)
U.S.A. (1.08)	Italy (1.37)	Estonia (1.11)	Spain (1.38)
Germany (1.11)	Greece (1.39)	Portugal (1.11)	Canada (1.39)
Spain (1.12)	Japan (1.39)	Japan (1.17)	Ireland (1.45)

Slovakia (1.19)	Ireland (1.40)	U.K. (1.21)	New Zealand (1.49)
Slovenia (1.19)	Denmark (1.45)	Slovakia (1.21)	Israel (1.53)
Israel (1.20)	Lithuania (1.47	Chile (1.22)	Slovenia (1.54)
Chile (1.21)	Sweden (1.47)	Czech Rep. (1.22)	Iceland (1.56)
Norway (1.21)	Iceland (1.52)	Luxembourg (1.25)	Netherlands (1.59)
Hungary (1.23)	Latvia (1.52)	Korea (1.26)	Switzerland (1.63)
Canada (1.24)	Finland (1.52)	Hungary (1.27)	Denmark (1.63)
Czeck Rep. (1.24)	Austria (1.53)	Lithuania (1.27)	Sweden (1.64)
New Zealand (1.25)	Switzerland (1.54)	Portugal (1.27)	Italy (1.65)
Turkey (1.26)	Netherlands (1.59)	U.S.A. (1.27)	Austria (1.78)
Portugal (1.27)	Australia (2.03)	Mexico (1.31)	Finland (1.89)

6. Concluding comments

In this article, we have examined the structure of the methane emissions in a group of 36 OECD countries by investigating the degree of persistence of the series and the potential presence of linear trends in the data. For this purpose, we have used fractional integration or I(d) techniques, which are more flexible than other approaches that simply use integer degrees of differentiation.

Our results indicate first that all series are highly persistent, with orders of integration which are statistically higher than 1 in the majority of the cases. In fact, the I(1) hypothesis is rejected in all cases except for the U.K. and Poland under the assumption of white noise errors. With autocorrelation, the exceptions are Germany, Belgium, France, Latvia and Estonia. Nevertheless, we cannot find a single case supporting the hypothesis of mean reversion ($d < 1$). The analysis was conducted on the log-transformed data. This implies that long memory ($d > 0$) is found in the majority of the growth rate series. Dealing with the presence of trends, a significant time trend coefficient is found in half of the series (16) under white noise errors, and allowing autocorrelation, the number of cases is slightly smaller.

The implication of the results is that the shocks to the methane emissions in the OECD countries are permanent. The methane emissions will not return back to their original mean after experiencing an economic or natural shock. Hence, policies aimed at reducing methane emissions will have permanent and lasting impacts. These policies include improvement in the equipment employed to generate, store, and convey natural gas and oil; changing manure management policies; modifications to animal feeding strategies; introduction of emission controls that capture landfill methane. Moreover, the national authorities and international environmental establishments need to focus on the long-run trends in methane emission as against concentrating on short-run targets.

Based on the long span of the data used in this work, back to 1750, an issue that is still to be considered is the possibility of structural breaks. This is a relevant point, taking into account, in particular, the fact that some authors argue that fractional integration may be a simple artefact generated by the presence of breaks in the data that has not been considered. This issue, along with others, such as the existence of smooth non-linear trends, still within this context of fractional integration will be examined in future papers.

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