Persistence of sulfur dioxide emissions in OECD countries between 1750-2014: A fractional integration approach

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Abstract

In this paper the degree of persistence of the sulfur dioxide emissions in a group of 37 OECD countries is examined by looking at the order of integration of the series. However, instead of using integer degrees of differentiation (i.e., 1 in case of unit roots and 0 for stationarity), fractional values are also considered. The results indicate high degrees of persistence and very little evidence of mean reversion. In fact, this property only holds for the three Latin American countries examined, namely Chile, Colombia and Mexico if the error follows a white noise process. If autocorrelation is permitted, however, the confidence intervals are wider and mean reversion is not found in any single case.

JEL Classification: C22; Q50; Q53; Q58

Keywords: Sulphur dioxide emissions; persistence; long range dependence

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1. Introduction

Combustion of sulfur-containing fuels including oil and coal, leads to the formation of sulfur dioxide (SO_2). It is produced from fuel combustion in the air, land and water transportation industries. However, power stations account for more SO_2 emissions than the remaining causes of the pollutant. SO_2 is one of the pollutants regarded as air contaminants by most developed countries (Srivastava, 2000). When SO_2 spreads into the atmosphere it forms fine particulate matter and sulfuric acid, secondary pollutants that have important negative impacts on people's health, the economy and the environment. (McLinden et al, 2016). Similar to other sulfur oxides, SO_2 emissions can cause acid rain which results in the damage of the ecosystems. At high concentrations, it can impair plants and trees by harming foliage and diminishing growth (Environmental Protection Agency, 2019).

Based on the importance of SO_2 emissions and the negative consequences that SO_2 emission buildup poses in the economy, especially on people's health, numerous features of the pollutant emissions have been examined in the extant literature such as the determining factors of SO_2 emissions (Managi et al, 2008; Zhou et al., 2017; Yang et al., 2017; Liu et al., 2019). There are also some articles that examine the convergence of SO_2 emissions (Solarin and Tiwari, 2020; Zhang et al., 2020). One of the areas that has received limited attention in the existing literature is the persistence of SO_2 emissions. The literature on the persistence of pollution indicators has been dominated by issues such as the persistence of CO_2 emissions (Belbute and Pereira, 2017). The trend observed for SO_2 emissions is quite different from the trend that CO_2 emissions has followed over the years (Solarin and Tiwari, 2020). The policies and methods aimed at reducing one or other type of pollutant emission differ considerably.

The important aspects of testing for persistence of SO_2 emissions are numerous. First, the lack of stationarity indicates that policy shocks to the SO_2 emissions brought about by the use of technologies or the imposition of standards or laws aimed at lowering SO_2 emissions will be permanent (Sidneva and Zivot, 2014). An example of such technologies is the wet system (which involves the mixing of crushed lime or limestone with water, which is then sprayed into the sulfur containing flue gases). Other examples are dry or semi-dry systems (which involve injecting a slurry of alkali sorbent, usually slaked lime, into the flue gases in a fine spray, following which the heat from the flue gases leads to evaporation of water as well as cooling the gases as it does so). Examples of current legislation and standards include the Clean Air Act of 1970 as well as the Energy Policy Conservation Act of 1975 in the United States; On-Road Vehicle and Engine Emission Regulations introduced through the Canadian Environmental Protection Act of 1999; and the various European Union Emission Standards for Light Commercial Vehicles (Timilsina and Dulal, 2009). Conversely, stationarity of SO_2 emissions implies that policy shocks to SO_2 emissions will have transient impacts.

An additional benefit is that from an econometrics viewpoint, nonstationarity in SO_2 emissions series has important consequences for the environmental Kuznets curve and (EKC) papers have employed SO_2 emissions as an indicator of pollutant emissions. A handful of EKC papers have relied on the assumption that there is trend stationarity in the pollutant emissions (Sidneva and Zivot, 2014). However, the EKC studies that use the level version of a nonstationary SO_2 emission (as the dependent variable) while the dependent variable such as real gross domestic product is also nonstationary, are likely to generate unreliable results (Newbold and Granger, 1974). In other words, econometric methods including the ordinary least squares (OLS) approach that are based on the premise that all the series in the analysis are stationary can yield spurious results, if the series are actually nonstationary (Hendry and Juselius, 2000).

Moreover, there is limited possibility of convergence among a set of series if the series are not individually stationary at level (Nieswiadomy and Strazicich., 2004). Therefore, any suggestion of convergence on the relative SO_2 emissions might not be valid if the SO_2 emissions series are not stationary at level.

The aim of the present paper is to extend the existing literature on pollutant emissions by examining the persistence of sulfur emissions in 37 OECD countries during the period 1750–2014. In order to do that, we adopt long range dependence techniques based on fractional integration, which is a more flexible method than the common approaches based on integer differentiation, using ARMA-ARIMA models corresponding to orders of integration of 0 and 1 respectively. This is relevant because we allow for a higher degree of flexibility in the dynamic specification of the models, and testing the order of integration of the series from a fractional viewpoint allows us to consider the possibility of nonstationary mean reverting processes with shock having transitory though with long lasting effects. Moreover, data from very long time series will be used to benefit from the larger sample size. We have focused on OECD countries for several reasons. With a gross domestic product worth US$52 trillion (at 2010 prices), OECD nations was responsible for 63% of world's GDP in 2018 (World Bank, 2020). Secondly, OECD nations have witnessed SO_2 emissions growth in majority of the sample period. SO_2 emissions increased by almost 200 times between 1750 and 2014 in the OECD countries (Hoesly et al., 2018). Thirdly, OECD nations was responsible for 15% SO_2 emissions in the globe in 2014 (Hoesly et al., 2018). Fourthly, SO_2 emission mitigation technologies in OECD nations are usually more efficient than those in existence in non-OECD nations. Several non-OECD nations refer to OECD nations when articulating their SO_2 emission mitigation blueprints.

The paper is presented as follows. First, the literature review will be examined in the following section. The methodology and data are presented in the Section 3. The fourth section discusses the main findings of the work. Finally, the conclusions of this research are presented in Section 5.

2. Literature review

Based on the existing correlation between economic development and environmental pollution, most of the current literature has focused on the Environmental Kuznets Curve (EKC) model as the main framework (Asumadu Sarkodie and Strezov, 2019). Thus, for example, Zhou et al. (2017) examined the relationship between SO_2 emissions and economic development by means of a spatial panel model and suggesting an inversely N-shaped environmental Kuznets curve. Other researchers suggested that air pollution is also related to a country's international trade which, in turn, promotes economic growth (Managi et al, 2008).

It is also worth highlighting studies aimed at identifying the convergence of SO_2 emissions in air pollution. Solarin and Tiwari (2019), using the panel stationarity test proposed in Nazlioglu and Karul (2017), examined sulfur dioxide (SO_2) emission convergence among 32 OECD countries. The results revealed the existence of convergence of SO_2 emissions among the OECD countries included in the study.

On the other hand, Zhang et al. (2020) employed the Fourier quantile unit-root test to survey SO_2 emissions per capita convergence in 74 cities of China for a time period from December 2014 to June 2019. Their results indicate that SO_2 emissions per capita in 72 out of 74 cities were convergent. They also conclude that the nonstationary behavior of SO_2 emissions per capita in these cities is asymmetrically persistent at different quantiles.

Other studies such as Yang et al. (2017) used data from 113 major cities in China in to investigate the impact of natural factors on SO_2 concentrations. The results indicated that precipitation exerts a significant effect on SO_2 reduction and temperature factors seems to aggravate SO_2 concentrations.

There are also studies on air pollution persistence. Meraz et al. (2015) used a long range dependence technique based on the rescaled range analysis (R/S) to examine the level of persistence of air pollutants in Mexico City. The air-pollution time series were hourly observations of nitrogen dioxide, ozone, sulfur dioxide and particulate matter obtained at the Mexico City downtown monitoring station from 1999 to 2014. The results showed that the long-range persistence property is not uniformly distributed over a wide range of time scales, from days to months.

Detrended fluctuation analysis (DFA) has also been employed in the analysis of persistence of air pollutants. Chelani (2013) examined time series of gaseous pollutants of CO, O_3 and NO_2 concentrations during the time period 2000–2009 at a traffic site in Delhi. Long range dependence (long memory) was observed in CO and NO_2, whereas O_3 showed anti-persistent behavior. Chen et al. (2016) examined similar issues in four major cities, Beijing, Guangzhou, Shenzhen and Shanghai, between 2013 and 2015. Their results revealed that pollution was persistent. In a similar way, Gil-Alana et al. (2020) analyzed the time series behavior of the air quality in the US states by looking at the degree of persistence of the particulate matter (PM_{10} and $PM_{2.5}$) datasets. They found heterogeneous results across the states with higher levels of persistence in the Western states with respect to those in the Eastern part, where a decreasing trend is also observed.

Based on the above review, it should be noted that studies on the persistence of sulfur dioxide emissions have been rarely discussed. Despite substantial overall downward trends in recent years, sulfur dioxide (SO_2) emitted from anthropogenic sources (e.g., coal-fired power plants) continues to have significant impacts on the environment (Li et al., 2020). Connecting the convergence hypothesis and the degree of persistence of sulfur dioxide (SO_2) emissions is important for environmental policymakers to effectively implement sulfur emission reduction strategies on a global scale.

3. Data and Methodology

3.1 Data

The SO_2 emissions dataset (in kilotonnes or kt) in 37 OECD nations for the period, 1750-2014, was obtained from the Joint Global Change Research Institute[1][2]. Australia is the only exception as it has dataset for only 1781-2014. The data was obtained within the Community Emissions Data System (CEDS) and has many advantages over the other sources of SO_2 emissions as they usually lack duplicability and uncertainty estimates (Hoesly et al., 2018). CEDS uses emission factors, current emission inventories and activity/driver data to generate annual national emissions across time and there are stages involved in the calculation phase. The first stage encompasses collection of data and the processing of data into a consistent format and timescale. In the second stage, default emissions for several OECD nations for 1960 to 2014 are calculated using driver and emission factor data. The drivers used in the computation include energy consumption (which is used as a driver for fuel combustion emissions), population (used as a driver for non-combustion emissions). In the third step, default estimates are scaled in order to be consistent with present emission inventories where available, plausible and complete. In the fourth step, scaled emission estimates are extended back to 1750 to obtain final

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[1] The data is available in https://zenodo.org/record/3606753#.XqLGoMgzbIU.

[2] The list is based on OECD membership as at 30/6/2020. The list was generated from https://www.oecd.org/about/document/list-oecd-member-countries.htm.

national emissions. Finally, emissions are scrutinized and collated to extract data for release and analysis (Hoesly et al., 2018).

Some relevant statistics of the series are presented in Table 1 and it is shown that with an average of 8576.478 kilotonnes of SO_2 emissions, the U.S. (which has the largest economy in the globe) was the biggest emitter among the OCED countries during the period under investigation. The U.S. is also among the countries with the largest standard deviation. The Skewness statistic suggests the majority of the series are not skewed. Hence, the logged transformed data is used for the empirical analysis.

Table 1: Descriptive statistics of the series

Series	Mean	Standard deviation	Minimum value	Maximum value
Australia	427.672	554.441	0.000	1816.861
Austria	79.022	99.663	1.884	377.456
Belgium	239.230	261.348	5.851	1028.546
Canada	1117.430	1526.571	0.237	5422.850
Chile	428.136	657.640	0.099	2396.017
Colombia	32.720	56.469	0.145	188.882
Czech Republic	490.288	605.728	2.767	2152.669
Denmark	58.707	120.327	0.911	608.662
Estonia	40.814	66.938	0.049	274.179
Finland	60.734	123.428	0.026	688.524
France	598.335	718.816	32.477	3204.397
Germany	2167.125	2577.280	11.967	8330.618
Greece	83.455	164.850	0.312	599.538
Hungary	302.672	421.279	3.985	1494.954
Iceland	2.225	4.275	0.003	20.012
Ireland	44.391	62.441	2.791	234.618
Israel	57.120	119.958	0.023	450.889
Italy	390.473	778.997	4.290	3386.328
Japan	531.588	883.988	2.017	4591.322
Korea	100.970	214.413	2.385	825.518
Latvia	19.920	38.652	0.443	165.179
Lithuania	32.767	61.593	0.981	256.506

Luxembourg	4.380	10.923	0.042	55.445
Mexico	380.199	650.110	0.776	2362.333
Netherlands	90.620	155.180	2.628	732.978
New Zealand	25.708	30.083	0.005	97.614
Norway	31.567	43.387	0.183	169.107
Poland	693.513	945.176	7.368	3628.186
Portugal	43.241	79.520	0.639	364.384
Slovakia	95.674	143.218	0.973	622.175
Slovenia	15.627	34.714	0.046	198.721
Spain	386.353	692.523	3.039	3020.669
Sweden	92.522	187.930	1.433	925.952
Switzerland	27.362	41.999	0.359	201.037
Turkey	302.826	669.287	1.145	2549.460
United Kingdom	2425.585	2070.001	101.381	6311.099
United States	8576.478	9573.654	2.130	29552.055

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

3.2 Methodology

Persistence is examined in this paper by using long range dependence techniques based on fractional integration that means that fractional differences are used in the differentiation of the series to render it stationary I(0).

We start by defining I(0) stationarity. A process $\{u_t, t = 0, \pm 1, \ldots\}$ is said to be I(0) or integrated of order 0 if its spectral density function (which is the Fourier transformation of the autocovariances) is finite and strictly positive at all frequencies. Then, a process is said to be I(d) or integrated of order d if it can be expressed as:

$$(1 - B)^d x_t = u_t, \quad t = 1, 2, \ldots, \qquad (1)$$

where B is the backshift operator ($Bx_t = x_{t-1}$), the parameter d can be a fractional value and u_t is stationary I(0) like a white noise or a stationary ARMA-type process.

The value of the differencing parameter d is relevant because it measures the degree of persistence in the data, noting that equation (1) can be expressed for any real value d as:

$$x_t = d\, x_{t-1} - \frac{d(d-1)}{2} x_{t-2} + \ldots + u_t. \qquad (2)$$

Thus, the higher the value of d is, the higher the level of association will be between the observations. Moreover, this specification is quite general since it allows us to consider different alternative modelling approaches such as short memory or I(0) processes (d = 0), stationarity with long memory (0 < d < 0.5); nonstationarity though mean reverting processes (0.5 ≤ d < 1), unit roots (d =1) and explosive processes (d > 1).

The estimation of d is carried out by using the frequency domain specification of the Whittle function, which is an approximation to the likelihood function. In more detail, we employ a testing procedure proposed in Robinson (1994) that has various advantages with respect to other approaches. Thus, it can be used for any value of d, including values above 0.5 in the nonstationary range; its limit distribution is a standard N(0,1) and this behavior holds regardless of whether deterministic terms such as an intercept and a linear time trend are included in the model.[3]

4. Empirical results

We examine the model given by the following equation,

$$y_t = \alpha + \beta t + x_t, \quad (1-B)^d x_t = u_t, \quad t = 1, 2, \ldots, \qquad (3)$$

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[3] The functional form of the version of the tests of Robinson (1994) used in this work can be found in Gil-Alana and Robinson (1997). The use of other approaches such as Sowell’s (1992) maximum likelihood method or the semiparametric method of Geweke and Porter-Hudak (1987) produced almost identical results to those reported in this work.

where y_t is the time series we observe, α and β are unknown parameters referring respectively to an intercept and a linear time trend, and the regression errors, x_t are integrated of order d, so that u_t is an integrated of order 0 process.

In Table 2 we suppose u_t is an uncorrelated process, so there is no other time dependence across the data than the one produced by the fractionally differenced parameter d. We report in the table the estimates of d in equation (3) under three different modelling specifications for the deterministic terms. Thus, in the second column, we present the results supposing that the two coefficients α and β are equal to 0 a priori, so no deterministic components are included in the model; in the third column, we report the estimated values of d under the assumption that α is unknown and β = 0 a priori, i.e., including a constant in the regression model. We have marked in bold the selected specification for each series. This selection is based on the significance of the coefficients in the d-differenced regression. We see that there are only three countries with a significant time trend. They are Colombia, Mexico and Turkey, and in the three of them the time trend is found to be positive (unreported). Focusing on the differencing parameter, we notice evidence of mean reversion only for Chile, Colombia and Mexico, which are the only three Latin American countries in the sample. Here, any shock affecting the series will have a transitory though long lasting effect. On the other hand, there are thirteen countries where the unit root null hypothesis (d = 1) cannot be rejected, while for the remaining 22 the estimated value of d is significantly higher than one, especially for Latvia (1.41), the Netherlands and Japan (1.49).

Table 2: Estimates of the differencing parameter: White noise errors

Country	No terms	With an intercept	With a linear trend
AUSTRALIA	**1.03 (0.94, 1.15)**	1.03 (0.94, 1.15)	1.03 (0.94, 1.15)
AUSTRIA	**1.05 (0.97, 1.14)**	1.05 (0.97, 1.14)	1.05 (0.97, 1.14)

BELGIUM	**1.18 (1.08, 1.31)**	1.18 (1.08, 1.31)	1.18 (1.08, 1.31)
CANADA	**1.16 (1.10, 1.25)**	1.16 (1.10, 1.25)	1.16 (1.10, 1.25)
SWITZERLAND	**1.16 (1.09, 1.25)**	1.16 (1.09, 1.25)	1.16 (1.09, 1.25)
CHILE	**0.92 (0.87, 0.98)**	0.92 (0.87, 0.98)	0.92 (0.87, 0.98)
COLOMBIA	0.83 (0.79, 0.89)	0.83 (0.79, 0.89)	**0.82 (0.78, 0.89)**
CZECK	**1.27 (1.19, 1.37)**	1.27 (1.19, 1.37)	1.27 (1.19, 1.37)
GERMANY	**1.14 (1.06, 1.26)**	1.14 (1.06, 1.26)	1.14 (1.06, 1.26)
DENMARK	**1.08 (1.01, 1.15)**	1.08 (1.01, 1.15)	1.08 (1.01, 1.15)
SPAIN	**1.12 (1.06, 1.20)**	1.12 (1.06, 1.20)	1.12 (1.06, 1.20)
ESTONIA	**1.06 (0.99, 1.16)**	1.06 (0.99, 1.16)	1.06 (0.99, 1.16)
FINLAND	**1.18 (1.09, 1.29)**	1.18 (1.09, 1.29)	1.18 (1.09, 1.29)
FRANCE	**1.17 (1.10, 1.26)**	1.17 (1.10, 1.26)	1.17 (1.10, 1.26)
GREAT BRITAIN	**0.99 (0.95, 1.04)**	0.99 (0.95, 1.04)	0.99 (0.95, 1.04)
GREECE	**1.46 (1.37, 1.55)**	1.46 (1.37, 1.55)	1.46 (1.37, 1.55)
HUNGARY	**1.23 (1.16, 1.32)**	1.23 (1.16, 1.32)	1.23 (1.16, 1.32)
IRELAND	**0.97 (0.90, 1.07)**	0.97 (0.90, 1.07)	0.97 (0.90, 1.07)
ICELAND	**1.11 (1.03, 1.20)**	1.11 (1.03, 1.20)	1.11 (1.03, 1.20)
ISRAEL	**1.02 (0.96, 1.09)**	1.02 (0.96, 1.09)	1.02 (0.96, 1.09)
ITALY	**1.35 (1.27, 1.45)**	1.35 (1.27, 1.45)	1.35 (1.27, 1.45)
JAPAN	**1.49 (1.39, 1.61)**	1.49 (1.39, 1.61)	1.49 (1.39, 1.61)
KOREA	**1.33 (1.28, 1.40)**	1.33 (1.28, 1.40)	1.33 (1.28, 1.40)
LITHUANIA	**1.07 (1.02, 1.14)**	1.07 (1.02, 1.14)	1.07 (1.02, 1.14)
LUXEMBOURG	**1.35 (1.26, 1.46)**	1.35 (1.26, 1.46)	1.35 (1.26, 1.46)
LATVIA	**1.41 (1.35, 1.49)**	1.41 (1.35, 1.49)	1.41 (1.35, 1.49)
MEXICO	0.83 (0.79, 0.89)	0.83 (0.79, 0.89)	**0.82 (0.79, 0.88)**
NETHERLANDS	**1.45 (1.32, 1.61)**	1.45 (1.32, 1.61)	1.45 (1.32, 1.61)
NORWAY	**1.02 (0.95, 1.13)**	1.02 (0.95, 1.13)	1.02 (0.95, 1.13)
NEW ZEALAND	**1.00 (0.93, 1.10)**	1.00 (0.93, 1.10)	1.00 (0.93, 1.10)
POLAND	**1.00 (0.94, 1.09)**	1.00 (0.94, 1.09)	1.00 (0.94, 1.09)
PORTUGAL	**1.00 (0.94, 1.08)**	1.00 (0.94, 1.08)	1.00 (0.94, 1.08)
SLOVAKIA	**1.25 (1.19, 1.32)**	1.25 (1.19, 1.32)	1.25 (1.19, 1.32)
SLOVENIA	**1.01 (0.93, 1.11)**	1.01 (0.93, 1.11)	1.01 (0.93, 1.11)
SWEDEN	**1.24 (1.17, 1.32)**	1.24 (1.17, 1.32)	1.24 (1.17, 1.32)
TURKEY	1.06 (1.00, 1.14)	1.06 (1.00, 1.14)	**1.06 (1.00, 1.14)**
UNITED STATES	**1.00 (0.94, 1.08)**	1.00 (0.94, 1.08)	1.00 (0.94, 1.08)

The values in parenthesis are the confidence intervals of the values of d where the null hypothesis cannot be rejected at the 5% level. In bold, the selected model for each series in relation with the deterministic terms.

In Table 3 we suppose that the disturbances term, u_t in equation (3) is autocorrelated. Here, we use the exponential spectral model of Blomfield (1973) which approximates autoregressive components in a very simple way. The time trend is now only required in the case of Australia. The estimated values of d are now much higher than in the previous case of white noise disturbances, and there is no evidence of mean reversion in any single case. The hypothesis of unit root (i.e., I(1) behavior) cannot be rejected in 15 countries, and in 22 the estimated value of d is found to be significantly higher than one.

Table 3: Estimates of the differencing parameter: Autocorrelated errors

Country	No terms	With an intercept	With a linear trend
AUSTRALIA	0.95 (0.85, 1.12)	0.95 (0.85, 1.12)	**0.94 (0.83, 1.12)**
AUSTRIA	**1.13 (0.95, 1.34)**	1.13 (0.95, 1.34)	1.13 (0.95, 1.34)
BELGIUM	**0.97 (0.86, 1.13)**	0.97 (0.86, 1.13)	0.97 (0.86, 1.13)
CANADA	**1.13 (1.03, 1.25)**	1.13 (1.03, 1.25)	1.13 (1.03, 1.25)
SWITZERLAND	**1.23 (1.09, 1.39)**	1.23 (1.09, 1.39)	1.23 (1.09, 1.39)
CHILE	**1.32 (1.17, 1.49)**	1.32 (1.17, 1.49)	1.32 (1.17, 1.49)
COLOMBIA	**1.08 (0.97, 1.23)**	1.08 (0.97, 1.23)	1.07 (0.97, 1.23)
CZECK	**1.21 (1.09, 1.36)**	1.21 (1.09, 1.36)	1.21 (1.09, 1.36)
GERMANY	**1.05 (0.92, 1.22)**	1.05 (0.92, 1.22)	1.05 (0.92, 1.22)
DENMARK	**1.26 (1.14, 1.40)**	1.26 (1.14, 1.41)	1.26 (1.14, 1.41)
SPAIN	**1.28 (1.16, 1.44)**	1.28 (1.16, 1.44)	1.28 (1.16, 1.44)
ESTONIA	**0.96 (0.86, 1.08)**	0.96 (0.86, 1.08)	0.96 (0.86, 1.08)
FINLAND	**1.02 (0.88, 1.21)**	1.02 (0.88, 1.21)	1.02 (0.88, 1.21)
FRANCE	**1.39 (1.22, 1.61)**	1.40 (1.22, 1.61)	1.40 (1.22, 1.61)
GREAT BRITAIN	**1.35 (1.24, 1.51)**	1.35 (1.24, 1.51)	1.35 (1.24, 1.51)
GREECE	**1.50 (1.36, 1.72)**	1.50 (1.36, 1.72)	1.50 (1.36, 1.72)
HUNGARY	**1.25 (1.14, 1.39)**	1.25 (1.14, 1.39)	1.25 (1.14, 1.39)
IRELAND	**0.96 (0.85, 1.11)**	0.97 (0.85, 1.11)	0.97 (0.85, 1.11)
ICELAND	**1.08 (0.96, 1.24)**	1.08 (0.96, 1.24)	1.09 (0.96, 1.25)
ISRAEL	**1.31 (1.18, 1.54)**	1.31 (1.18, 1.54)	1.32 (1.18, 1.54)
ITALY	**1.35 (1.24, 1.51)**	1.35 (1.24, 1.51)	1.35 (1.24, 1.51)
JAPAN	**1.29 (1.07, 1.57)**	1.29 (1.07, 1.57)	1.29 (1.07, 1.57)
KOREA	**1.73 (1.58, 1.90)**	1.72 (1.58, 1.90)	1.72 (1.58, 1.90)
LITHUANIA	**1.35 (1.22, 1.50)**	1.35 (1.22, 1.50)	1.35 (1.22, 1.50)
LUXEMBOURG	**1.29 (1.10, 1.52)**	1.29 (1.10, 1.52)	1.29 (1.10, 1.52)
LATVIA	**1.63 (1.49, 1.83)**	1.63 (1.49, 1.83)	1.63 (1.49, 1.83)
MEXICO	**1.01 (0.93, 1.11)**	1.01 (0.93, 1.11)	1.01 (0.92, 1.11)
NETHERLANDS	**1.05 (0.90, 1.28)**	1.05 (0.90, 1.28)	1.05 (0.90, 1.28)
NORWAY	**0.94 (0.83, 1.08)**	0.94 (0.83, 1.08)	0.94 (0.83, 1.08)
NEW ZEELAND	**1.06 (0.92, 1.26)**	1.06 (0.92, 1.26)	1.06 (0.92, 1.26)
POLAND	**1.14 (1.01, 1.30)**	1.14 (1.01, 1.30)	1.14 (1.01, 1.30)
PORTUGAL	**1.25 (1.10, 1.41)**	1.25 (1.10, 1.41)	1.25 (1.10, 1.41)
SLOVAKIA	**1.42 (1.30, 1.56)**	1.42 (1.30, 1.56)	1.42 (1.30, 1.56)

SLOVENIA	**1.06 (0.89, 1.29)**	1.06 (0.89, 1.29)	1.06 (0.89, 1.29)
SWEDEN	**1.33 (1.20, 1.49)**	1.33 (1.20, 1.49)	1.33 (1.20, 1.49)
TURKEY	**1.09 (1.01, 1.21)**	1.09 (1.01, 1.21)	1.10 (1.01, 1.20)
UNITED STATES	**1.04 (0.94, 1.17)**	1.04 (0.94, 1.17)	1.04 (0.94, 1.17)

The values in parenthesis are the confidence intervals of the values of d where the null hypothesis cannot be rejected at the 5% level. In bold, the selected model for each series in relation with the deterministic terms.

5. Concluding comments

In this paper we have examined the stochastic properties of the SO_2 emissions in 37 OECD countries by looking at their degree of differentiation from a fractional viewpoint. This methodology is more general than the classical one that is based exclusively on integer values for the order of integration. The results indicate that if the error term is uncorrelated, mean reversion occurs in the three Latin American countries, namely Chile, Colombia and Mexico, with orders of integration higher than 0.5 but strictly smaller than one and thus implying long lasting effects of shocks. For all the other countries in the sample, the order of integration is found to be equal to or higher than one implying permanency of shocks. If autocorrelation is permitted, the confidence intervals for the values of the differencing parameter are wider and mean reversion is not found in any single case.

The results presented in this work are of interest for the authorities since lack of mean reversion implies that shocks have a permanent nature; thus, if a negative shock occurs in relation to the SO2 emissions, strong policies should be adopted by the authorities if it is beneficial to recover the original levels, although the duration of SO2 emissions in the atmosphere might be not as long as what several other pollutants exhibit. In other words, efforts to reduce SO2 emissions are required when there is an increase in the pollutant as a result of various economic activities such as movements of automobiles and heavy equipment that burn fuel with a high sulfur content. On the other hand, the good news is that positive shocks reducing the level of emissions will also be permanent, and no strong actions should then be required.

Another implication of the foregoing results is that there is a need for adequate on investments on mitigation efforts at reducing SO2 emissions, such as wet systems which involve the use of sorbent materials to reduce SO2 emissions, especially when a rise in the pollutant is being experienced. This is because the lack of mean reversions in the results implies that the increase in the pollutant will be permanent if adequate investments for mitigation are not available.

The paper can be extended in several directions. More countries should be examined. In particular, the fact that the only three Latin American countries investigated in the paper display some degree of mean reversion is interesting and this might be a specific feature corresponding to this continent. Other potential extensions include the possibility of structural breaks and non-linear issues. Liddle and Messinis (2015) showed substantial breaks and nonlinearities in OECD sulfur emissions so this possibility should also be investigated in the context of fractional integration. In fact, authors such as Diebold and Inoue (2001) and Granger and Hyung (2004) demonstrated that these two issues (nonlinearities and fractional integration) were intimately related. Thus, a natural extension of this work might be to examine the potential presence of breaks in the data, using for instance the approach developed in Gil-Alana (2008) that extends Bai and Perron’s (2003) method to the fractional case. In addition, nonlinear deterministic terms can also be included in the model avoiding the abrupt changes produced by the classical structural breaks. Here, we can employ the Chebyshev polynomials in time proposed in Cuestas and Gil-Alana (2016), the Fourier functions (Gil-Alana and Yaya, 2021) or even neural networks (Yaya et al., 2021) still in the context of the fractional models used in this work. This line of research is now in progress.

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