Exploring the effect of economic growth and government expenditure on the environment

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ABSTRACT
This paper examines the effect of economic growth and government spending on the environment using a panel of 71 countries for the time period 1970-2008. In particular, we test the hypothesis of the existence of an inverted U-shaped relationship between economic performance and pollution, as well as the hypothesis of a negative direct relationship between fiscal spending and pollution. To take into account that environmental degradation may respond to changes in income and government spending with a time lag, due to technological and institutional reasons, we apply appropriate dynamic econometric methods. We report the estimates for both the short-run and long-run effects on two different air pollutants, namely SO_2 and CO_2, distinguishing the results for different levels of economic development. Policy implications range depending on the level of income of the considered countries.

Keywords: Government expenditure; economic growth; environment; dynamics.

JEL Classifications: H50; E60; Q53; Q54; Q56.

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1. Introduction

The purpose of this paper is twofold, namely to examine the effect of economic growth as well as that of government expenditure on environmental degradation, taking into account the dynamic nature of these relationships. The environmental Kuznets curve (hereafter EKC) hypothesis posits that in the early stages of economic development environmental degradation will increase until a certain level of income is reached and then environmental improvement will occur (Gross and Krugman, 1995). On the other hand, government expenditure has recently expanded in many countries to alleviate the adverse effects of the recent economic crisis, with a large fraction of GDP spent by governments affecting a variety of economic variables and prosperity in general. A recent strand of literature suggests that government spending is an important determinant of environmental quality (Lopez et al., 2011, Halkos and Paizanos, 2013; Galinato and Islam, 2014).

According to Halkos (2003) EKC studies identify several factors as the most important in determining the inverted-U shape of the curve. In particular these factors include, improvements in environmental quality occurring from advances in production technology, the exportation of ‘dirty industry’ to less developed countries, the role of preferences and regulation on the emissions profile of polluters, the better institutional set up in the form of credible property rights, regulations and good governance which may create public awareness against environmental degradation and finally, the technological link between the consumption of a desired good and the abatement of its undesirable by-products in the form of pollution.

On the other hand, the mechanisms through which government expenditure and environment interact with each other are investigated in theoretical papers by Heyes (2000), Lawn (2003) and Sim (2006). Higher government expenditure is more likely to include redistributive transfers, which result to increased income equality and thus to higher demand for environmental quality. Moreover, if the environment is a luxury public good, it is likely that it will only be demanded when the demand for other public goods has been satisfied, i.e. at large levels of government size (Frederik and Lundstrom, 2001). Lopez et al. (2011) identify four mechanisms by which the level and composition of fiscal spending may affect pollution levels[1], namely the scale (increased environmental pressures due to more economic growth), composition (increased human capital intensive activities instead of physical capital intensive industries that harm the environment more), technique (due to higher labor efficiency) and income (where increased income raises the demand for improved environmental quality) effects.

However, in examining the aforementioned relationships their dynamic nature should be taken into account. In particular, it is highly unlikely that the above effects of income and government spending on the environment occur instantaneously (Halkos, 2003; Lopez et. al., 2011) which may be the case for several reasons. For example, technological advances that usually accompany economic development may take several years until fully implemented and employed by industries. In addition, for psychological reasons and as a result of the force of habit (inertia), industries and consumers may not alter their production methods and behaviour immediately following a technological advance or a distributional effect from a change in public spending, a result that may also be augmented by imperfect knowledge. Finally, one may also expect institutional reasons to contribute to the existence of lags in the examined relationships.

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[1] In particular, they examine the effect of the share of public goods in total government expenditure on pollution.

Given this background, our purpose is to investigate firstly how increases in income and government spending affect pollution at given income levels in the short-run and then to estimate how this changes influence environmental quality in the long-term.

To the best of our knowledge the present paper is the first that explicitly studies the short-run as well as the long-run effects of both economic development and government expenditure on the environment. For that reason, we estimate an augmented EKC equation, employing a sample of 71 countries covering the period 1970-2008 for two air pollutants (sulfur dioxide, SO_2 and carbon dioxide, CO_2). In estimating the proposed model we take into account the dynamic nature of the relationships examined, by employing appropriate econometric methods for the estimation of dynamic panels.

The remainder of the paper is organized as follows: Section 2 presents the data used in the analysis and section 3 discusses the proposed econometric models. The empirical results are reported in section 4 while the final section concludes the paper.

2. Data

Our sample consists of 71 countries[2] with a full set of SO_2, CO_2, GDP/c and share of government expenditure, for the period 1970-2008. The analysis for SO_2 takes place up to the year 2003 because of limited availability of data on this

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[2]Albania, Angola, Argentina, Australia, Austria, Belgium, Bolivia, Brazil, Bulgaria, Canada, Chile, China, Colombia, Denmark, Djibouti, Equador, Egypt, El Salvador, Ethiopia, Finland, France, Ghana, Greece, Guatemala, Guinea Bissau, Honduras, Hungary, India, Ireland, Israel, Italy, Jamaica, Japan, Jordan, Kenya, Lebanon, Liberia, Madagascar, Mauritius, Mexico, Mongolia, Morocco, Mozambique, Netherlands, New Zealand, Nigeria, Norway, Panama, Paraguay, Peru, Philippines, Poland, Portugal, Qatar, Romania, Sierra Leone, South Africa, Spain, Sri Lanka, Sudan, Sweden, Switzerland, Syria, Thailand, Trinidad, Tunisia, Turkey, Uganda, United Kingdom, United States, Uruguay

pollutant after that period. The data for SO_2 and CO_2 are from Stern (2005) and Boden et. al. (2011) respectively, the data on national income from Maddison (2010) and finally the data on government share of income were collected from the Penn World Table (2009). The database consists of up to 2,698 observations per variable.

Data on emissions of the two pollutants were used rather than their concentrations, to avoid dependence of results on geographic location characteristics and atmospheric conditions. An important distinction between the two pollutants that has to do with their atmospheric life characteristics is their geographical range of effect (Cole, 2007). Considering that two-thirds of SO_2 moves away from the atmosphere within 10 days after its emission, its impact is mainly local or regional and thus, historically, sulfur dioxide has been subject to regulation. In contrast, CO_2 has not been regulated by governments, since its atmospheric life varies from 50 to 200 years and hence its impact is global.

Moreover, the sources of pollution vary by pollutant. The main sources of SO_2 emissions are electricity generation and industrial processes. On the other hand, apart from energy transformation and industry, an important source of CO_2 emissions is transport. Apparently SO_2 pollution is characterized as production-generated, while CO_2 emissions are a mix between production and consumption–generated pollution. This distinction is important since the mechanism by which government expenditure size affects consumption pollution is likely to differ compared to production pollution. SO_2 emissions can be decreased by reducing consumption of fossil fuels (especially high-sulfur content coal), by using smoke-scrubbing equipment in power plants and by increasing energy efficiency. However, in consumption related pollutants the use and influence of environmental policies is more difficult, since the main tool to reduce these is the implementation of environmental taxes, which are often avoided as they are not politically popular.

3. Methodology

To establish the specification between air pollution and GDP/c, Box-Cox tests have been performed to test linearity against logarithmic functional forms. In addition, implementation of the Akaike and Bayesian information criteria indicated that the appropriate use of powers of the income variable is three, thus we use a cubic specification. In addition, employing greater powers of the income variable leads to multicollinearity among the explanatory variables. Specifically, findings of the tests suggest the following model which represents a conventional cubic formulation of the EKC, augmented by the lagged share of government expenditure over income:

$$\ln(P/c)_{it} = \mu_i + \zeta_t + \beta_1 \ln Govshare_{it-1} + \beta_2 \ln(GDP/c)_{it} + \beta_3 (\ln(GDP/c))^2_{it} + \\ + \beta_4 (\ln(GDP/c))^3_{it} + \varepsilon_{it} \qquad (1)$$

where subscripts i and t represent country and time respectively and all variables are expressed in natural logarithms, unless otherwise stated.

The income variable and its powers in (1) control for scale effects. The term μ_i is a country effect which can be fixed or random, ζ_t is a time effect common to all countries and ε_{it} is a disturbance term with the usual desirable properties. Following the terminology used to classify the pollution effects in the trade literature, the coefficient on the government expenditure variable captures the composition, income and part of the technique effect.

3.1 Econometric issues and estimation

In estimating equation (1) we must take into account the unobserved heterogeneity across countries. The standard approach is to use fixed and random effects, hereafter FE and RE respectively, model formulations with the choice depending on the assumption adopted about the correlation between the cross-section specific error-component and the explanatory variables. When such correlation is present, then RE estimators are not consistent and efficient and the use of FE is more appropriate. For instance, in the pollutants equations these country-specific characteristics may include differences in climate, geography and fossil fuels endowments, all of them potentially correlated with emissions (Leitao, 2010). Additionally, it is very likely that country unobserved characteristics are correlated with income and the other explanatory variables, implying that FE estimation is preferred. This assumption is supported by the use of Hausman test, in which the RE model was rejected in favor of the FE model, for equation (1) in all cases.

Since the balanced panel data used in this paper consists of large N and T dimensions, non-stationarity is important, hence in estimating the models we are particularly concerned about the dynamic misspecification of the pollutants equations. In particular, if we rely on a static model, then all adjustments to any shock occur within the same time period in which they occur, but this could be justified only in equilibrium or if the adjustment mechanism is rapid. However, according to Perman and Stern (1999) this is extremely unlikely and on the contrary, it is expected that the return to long-run equilibrium emission levels is a rather slow process.

To estimate a non-stationary dynamic panel we employ the dynamic fixed effects (DFE) estimator developed by Pesaran and Smith (1995) and Pesaran et al. (1997, 2004). In DFE estimation we assume that intercepts differ across countries but that the long-run coefficients are equal across countries. However, if equality of the slope coefficients does not hold in practice, this technique yields inconsistent estimators. This assumption is tested using a Hausman test.

For equation (1), adopting the formalization by Blackburne III and Frank (2007), we set-up an initial general autoregressive-distributed lag model AD (p,q1,…,qk) of the form:

$$\ln(P/c)_{it} = \mu_i + \sum_{j=1}^{p} \lambda_{ij} \ln(P/c)_{i,t-j} + \sum_{j=0}^{q} \beta_{ij}^{'} K_{i,t-j} + \varepsilon_{it} \tag{2}$$

where number of countries $i = 1,2,...,N$; number of periods $t = 1,2,...,T$, for sufficiently large T; K_{it} a $k \times 1$ vector of explanatory variables including government expenditure and income variables; and μ_i a country-specific effect.

If the variables in equation (2) are integrated of order one (that is I(1)) and cointegrated, then the error term is an I(0) process for all i. A principle feature of cointegrated variables is their responsiveness to any deviation from the long-run equilibrium. Hence, it is possible to specify an error correction model in which deviations from the long-run equilibrium affect the short-run dynamics of the variables. The error correction equation is formed as:

$$\Delta \ln(P/c)_{it} = \phi_i [\ln(P/c)_{i,t-1} - \zeta_i^{'} K_{it}) + \sum_{j=1}^{p-1} \lambda^{*}{}_{ij} \Delta \ln(P/c)_{i,t-1} + \sum_{j=0}^{q-1} \beta_{ij}^{'\,*} \Delta K_{i,t-j} + \mu_i + \varepsilon_{it} \tag{3}$$

where $\phi_i = -(1 - \sum_{j=1}^{p} \lambda_{ij}), \zeta_i = \sum_{j=0}^{q} \beta_{ij} / (1 - \sum_k \lambda_{ik}), \lambda_{ij}^{*} = -\sum_{m=j+1}^{p} \lambda_{im} \;\; j = 1,2,..., p-1$

and $\beta_{ij}^{*} = -\sum_{m=j+1}^{q} \beta_{im} \;\; j = 1,2,..., q-1$.

Nonlinearity in the parameters requires that the models are estimated using maximum likelihood.

Another econometric concern for equation (1) is the bias occurring from the potential endogeneity between government spending and pollution, since government spending often increases with pollution because governments implement ecological taxes. Moreover, as already mentioned, the impact of government expenditure may not occur instantaneously. For this reason, we use the lagged share of government expenditure which also may mitigate bias from reverse causality.

3.2 Identifying the short- and long-run effects

Including more than one lags of the government expenditure and income variables in (1) to capture dynamics may result in multicollinearity. Thus, we employ the Koyck transformation of estimating distributed lag models. In particular, we assume that the subsequent effects of government expenditure and income are all of the same sign as their short-run counterparts and that they decline geometrically each year following:

$$\beta_{\kappa\tau} = \beta_{i0}\lambda^{t} \tag{4}$$

In addition we assume that, after a change in government expenditure and income, the speed of adjustment rates of the pollutants' emissions to their long-run equilibrium are similar and thus we propose the model:

$$\begin{aligned} \ln(P/c)_{it} &= (\mu_i + \zeta_t)^{*} + \beta_1 \ln Govshare_{it-1} + \beta_2 \ln(GDP/c)_{it} + \beta_3(\ln(GDP/c))^2_{it} + \\ &+ \beta_4(\ln(GDP/c))^3_{it} + \lambda \ln(P/c)_{it-1} + \theta_{it} \end{aligned} \tag{5}$$

where $\theta_{it} = \varepsilon_{it} - \lambda\, \varepsilon_{it-1}$.

Coefficient β_1 that will be obtained from the estimation of equation (5) can be interpreted as the short-run elasticity of government spending on pollution, while the marginal effect of the income variable may be interpreted as the short-run income elasticity. Long run elasticity of government spending is given by β_1 / (1-λ) while the long-run income elasticity can be obtained, respectively, by dividing the short-run elasticity of income by the term (1-λ).

4. Results

Before turning to the estimation of per capita pollution equations we should examine the time series properties of the main variables used. Testing for unit roots in panel data requires both the asymptotic behavior of the time-series dimension T, and the cross-section dimension N, to be taken into consideration. Since the panel data set we examine consists of both $N \rightarrow \infty$ and $T \rightarrow \infty$ dimensions, the tests of stationarity performed are based on the Fisher-type Phillips-Peron unit root test. The test allows heterogeneity of the autoregressive parameter and although in its general form does not control for cross-sectional dependence, is more powerful than Levin et al. (2002) in that case[3]. Table 1a presents the results of the Phillips-Perron unit root tests on the variables of interest. There is evidence against stationarity in levels, since in all cases our variables are I(1).

Table 1a: Panel data unit root tests

Variable	no trend c-s means (1)	no trend minus c-s means (2)	with trend c-s means (3)	with trend minus c-s means (4)
Log SO_2/c	0.673	0.707	0.316	0.604
Δ(Log SO_2/c)	0.000	0.000	0.000	0.000
Log CO_2/c	0.049	0.361	0.273	0.880
Δ(Log CO_2/c)	0.000	0.000	0.000	0.000
Log Government share lagged	0.224	0.034	0.479	0.043
Δ(Log Government share lagged)	0.000	0.000	0.000	0.000
Log GDP/c	1.000	0.925	1.000	1.000
Δ(Log GDP/c)	0.000	0.000	0.000	0.000
Log GDP/c^2	1.000	0.975	1.000	1.000
Δ(Log GDP/c)^2	0.000	0.000	0.000	0.000
Log GDP/c^3	1.000	0.998	1.000	1.000
Δ(Log GDP/c)^3	0.000	0.000	0.000	0.000

Note: Fisher-type Phillips-Perron unit root tests performed on each panel including zero or one Newey-West lag. All values reported are probabilities. C-s means stands for cross-sectional means.

Additionally, application of the DFE method requires that the variables in the model are cointegrated meaning that there is a long-run relationship among them.

[3] We also compute the mean of the series across panels and subtract this mean from the series (columns 2 and 4 in Table 1a) to mitigate the impact of cross-sectional dependence.

Table 1b presents the Pedroni and the Kao (Engle based) cointegration tests for the two pollutants equations. We reject the null hypothesis of no-cointegration at the conventional statistical significance level of 0.05 in six of the eight cases for the SO_2 equation and in five cases for CO_2. However, in terms of raw power of the statistics for relatively small values of T the rho and panel-v statistics are the most conservative and show a tendency to not reject (Pedroni, 2004), suggesting that evidence of cointegration is even stronger than that depicted in Table 1b.

Table 1b: Pedroni residual cointegration test for the two pollution equations

	SO_2/c		CO_2/c	
	Statistic	Probability	Statistic	Probability
Panel v-statistic	4.331	0.000	7.118	0.000
Panel rho-statistic	7.799	1.000	0.181	0.572
Panel PP-statistic	-8.798	0.000	-2.623	0.004
Panel ADF-statistic	-22.60	0.000	-9.173	0.000
Group rho-statistic	12.02	1.000	3.886	0.999
Group PP-statistic	-8.238	0.000	-0.757	0.225
Group ADG-statistic	-25.18	0.000	-9.576	0.000
Kao test (Engle based)	-33.88	0.000	-34.29	0.000

Table 2 provides the estimates of per capita pollution emissions. In our model, as mentioned, according to the Hausman test, the FE estimation is preferred to RE. Hence, for each pollutant we report FE and DFE estimates. Dynamics are taken into account in the estimates reported in columns 2 and 4 of the Table. Comparing the Mean Group (MG) and Pooled Mean Group (PMG) estimators, with the use of a Hausman test, we found that the PMG estimator, the efficient estimator under the null hypothesis, is preferred indicating that the assumption of equal long-run coefficients across panels is more appropriate in our panel. Additionally, another application of the Hausman test suggests that the simultaneous equation bias between the error term and the lagged dependent variable is minimal in our panel and we may conclude that the DFE model is the most appropriate. In addition, the error correction term in the DFE estimator for both pollutants is statistically significant at the 1% level for both pollutants, suggesting that taking into account dynamics is necessary.

Table 2: Estimates of per capita pollution emissions for the world sample

	SO_2/c		CO_2/c	
	FE	DFE	FE	DFE
Log government share lagged	-0.379**	-0.663**	-0.052	-0.070
	(0.155)	(0.287)	(0.086)	(0.121)
Log GDPc	-27.31**	-22.52	-17.12***	-22.74***
	(11.44)	(13.76)	(4.889)	(7.805)
(Log GDPc)^2	3.849**	3.284*	2.269***	2.943***
	(1.444)	(1.716)	(0.586)	(0.914)
(Log GDPc)^3	-0.174***	-0.153**	-0.094***	-0.121***
	(0.060)	(0.071)	(0.023)	(0.035)
Constant	57.74*		39.36***	
	(29.67)		(13.29)	
Error correction term		-0.137***		-0.118***
		(0.055)		(0.016)
Turning Points	380/6,673	298/5,502	419/23,242	573/19,228
R^2	0.305		0.392	
F test	0.000		0.000	
Hausman FE v. RE	0.000		0.000	
Hausman MG v. PMG		0.510		0.527
Hausman PMG v. DFE		0.010		0.997
Nobs/Countries	2,190/71	2,119/71	2,698/71	2,627/71

Note: Robust standard errors are in parentheses. All tests' values reported are probabilities.
*Significant at 10%. **Significant at 5% ***Significant at 1%.

Both pollutants have a significant inverted N-shaped cubic relationship with per capita income in all estimates (for similar findings see for example Cole, 2007). Interestingly, taking into account dynamics in the DFE estimates produces slightly lower turning points for both pollutants. However, the initial turning point is particularly low in all estimates that essentially for the in sample income observations the derived EKCs have the conventional quadratic form.

On the other hand, a negative direct effect of government share of income on pollution is estimated by all models. Concentrating on DFE estimates the government share of income possesses a negative relationship with SO_2/c which is significant at 5%, however the effect on CO_2/c is insignificant. In particular, an increase of government expenditure by 1%, ceteris paribus, may result in a 0.663% reduction of SO_2/c emissions but has no effect on CO_2/c emissions.

In Table 3 we present the estimates of the pollution equations employing the Koyck transformation. Results are presented for the whole sample, as well as for two sub-samples, namely the OECD group of countries and one with the rest countries of our sample that do not belong in OECD. It is interesting to note that the estimated coefficients of the lagged pollutant variables are significant in all cases at the 1% level. In addition, for SO_2 the coefficient of the lagged pollution variable greatly differs between the two subgroups, suggesting different adjustment rates and return to equilibrium pollution levels after a change in the explanatory variables.

Table 3: Estimates of per capita pollution emissions using Koyck transformation

	SO_2/c			CO_2/c		
	World	OECD	Non-OECD	World	OECD	Non-OECD
Log government share lagged	-0.090^{**}	-0.082^{**}	-0.112^{**}	-0.016	-0.023	-0.015
	(0.039)	(0.030)	(0.553)	(0.014)	(0.017)	(0.017)
Log GDPc	-3.569^{**}	-0.084^{***}	1.301^{*}	-2.707^{**}	0.848^{**}	-2.632^{**}
	(1.769)	(0.022)	(0.660)	(1.076)	(7.805)	(1.174)
(Log GDPc)^2	0.514^{**}		-0.076^{*}	0.358^{***}	-0.044^{**}	0.345^{**}
	(0.226)		(0.039)	(0.126)	(0.020)	(0.139)
(Log GDPc)^3	-0.024^{**}			-0.015^{***}		-0.014^{**}
	(0.014)			(0.005)		(0.005)
Log SO_2/c lagged	0.857^{***}	0.985^{***}	0.771^{***}	0.870^{***}	0.896^{***}	0.864^{***}
	(0.054)	(0.010)	(0.089)	(0.014)	(0.021)	(0.016)
Constant	7.410^{*}	0.913^{***}	-6.386^{*}	6.332^{**}	-3.953^{**}	6.160^{*}
	(4.428)	(0.249)	(3.095)	(2.996)	(1.840)	(3.255)
Turning Points	387/4,103	-	5,215	485/16,751	15,312	412/33,089
Long-run gov. expend. elasticity	-0.629	-5.466	-0.489	-	-	-
Short-run income elasticity	-0.019	-0.084	0.105	0.131	0.006	0.197
Long-run income elasticity	-0.133	-5.600	0.459	1.008	0.058	1.447
R^2	0.821	0.963	0.687	0.878	0.902	0.875
F test	0.000	0.000	0.000	0.000	0.000	0.000
Hausman FE v. RE	0.000	0.032	0.000	0.000	0.000	0.000
Nobs/Countries	2,190/71	828/26	1,362/45	2,698/71	988 /26	1,710/45

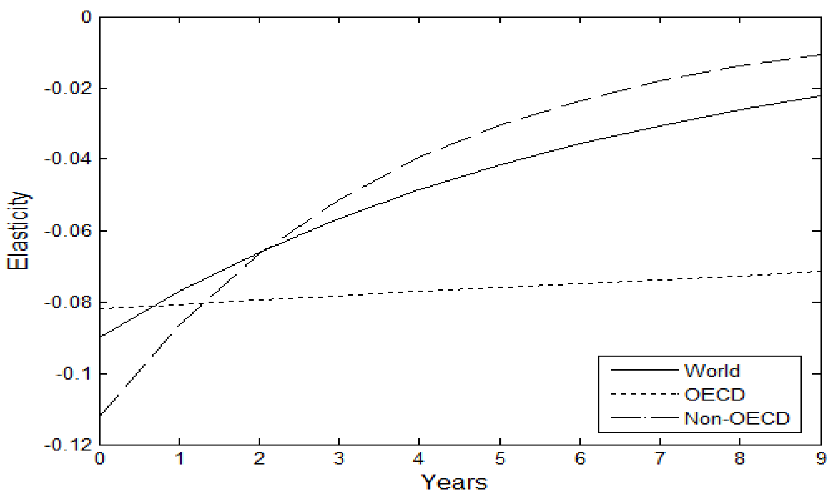
Note: Robust standard errors are in parentheses. All tests' values reported are probabilities. Short-run income elasticity, as well as long-run income and government expenditure elasticities are calculated at the sample median level of per capita income of each sub-sample which are $4,565, $14,319 and $2,605 for the World, OECD and Non-OECD groups, respectively.

*Significant at 10%. **Significant at 5% ***Significant at 1%.

Consistent with the previous results, the estimated effect of government expenditure is negative in all cases but remains significant only for SO_2. However, the specification of the pollution equation depends on the sample of countries used. In particular, for SO_2, there is evidence for an inverted N-shaped EKC in the full sample and for a quadratic form for the Non-OECD countries; however results suggest a monotonic relationship for the OECD countries. On the other hand, for CO_2 the derived EKC is inverted N-shaped for the World and Non-OECD countries and inverted U-shaped for the OECD sample.

The estimated long-run elasticities of government expenditure on SO_2 are greater than their short-run counterparts in all cases. Focusing on SO_2, the estimated short-run elasticities of government share are of similar magnitude among the different groups, however the same does not hold for the long-run elasticities. The latter are much greater, in absolute value, for OECD countries suggesting that a sustained increase of 1% in government share, ceteris paribus, leads to a long-run reduction of 5.466% in SO_2 emissions, a result which is more than 10 times greater than for the Non-OECD countries. This relationship is depicted in Figure 1 where the partial effect of an 1% increase in government expenditure on SO_2 is shown for the following 10 years, for each of the three country groups.

Figure 1: The partial effects of government share on SO2/c



The elasticities of income on SO_2 are negative for the world and OECD countries group but positive in the Non-OECD countries. In particular, the estimate of the long run elasticity of income on SO_2 for the median income OECD country implies that following a 1% sustained increase in income, ceteris paribus, there will be a 5.6% reduction in SO_2 emissions. On the contrary, a 1% sustained increase in income, ceteris parbus, is estimated to cause a 0.459% increase in SO_2 emissions in a Non-OECD country. For the CO_2 emissions the income elasticities are positive in all samples. However, both in the short- and long-run the effect is much larger in the Non-OECD countries group. Figures 2 and 3 depict these relationships.

Figure 2: The partial effects of income on SO2/c

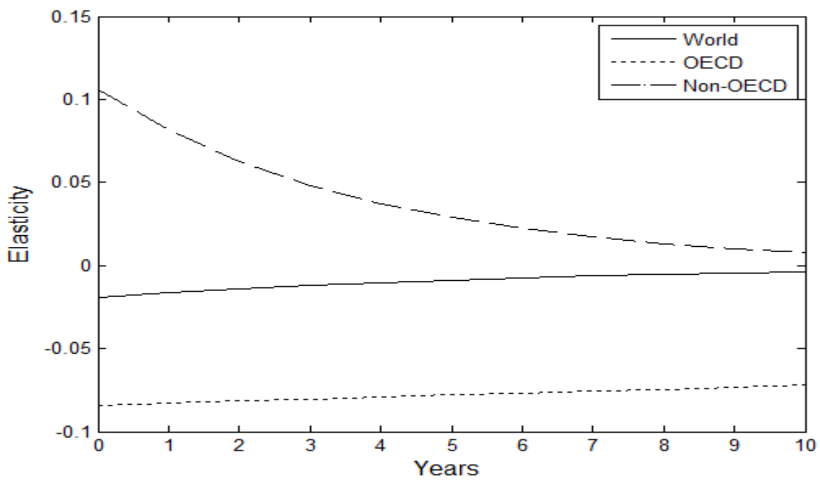
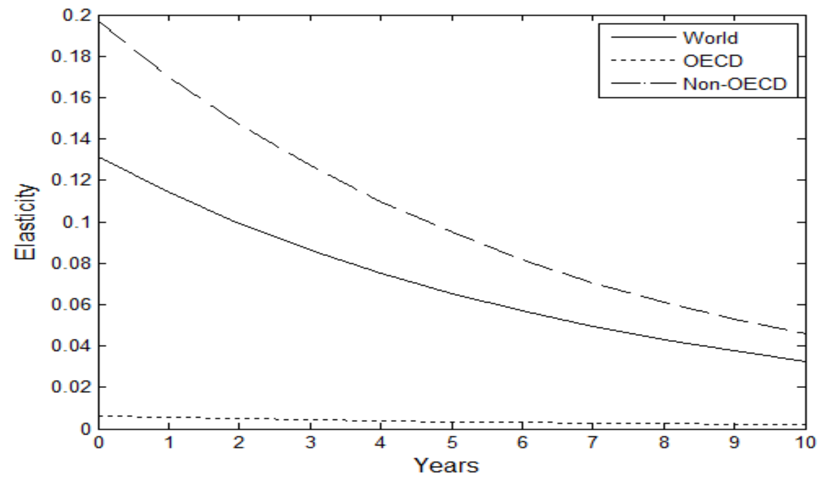


Figure 3: The partial effects of income on CO2/c



4.1 Sensitivity analysis

We test the existence of potential biases from omitted time-variant variables. Table 4 reports the results from estimating the effect of government expenditure under a series of relative correlation restrictions, using the method proposed by Krauth (2011). To account for country fixed-effects, each variable is expressed in terms of deviation from the corresponding country-level average. The results suggest that the estimated effect for SO_2/c is robust, while the same does not hold for CO_2/c, as expected. We find that for the effect on SO_2/c to cease being strictly negative the correlation between government expenditure and unobservables would need to be 6.25 times larger than the correlation with the observables, which seems highly unlikely. However, for CO_2/c a relative correlation of only 40% or greater, implies that the point estimate of the effect includes zero and thus is not strictly negative.

Table 4: Robustness checks for omitted variables bias

Relative correlation restriction (Λ)	Bounds on Government share effect by pollutant $[\theta_L(\Lambda), \theta_H(\Lambda)]$	
	SO_2/c	CO_2/c
{0.00}	-0.363**	-0.025
	(-0.645, -0.081)	(-0.200, 0.151)
[0.00, 0.50]	[-0.457, -0.363]	[-0.025, 0.006]
	(-0.753, -0.110)	(-0.189, 0.193)
[0.00, 1.00]	[-0.554, -0.363]	[-0.025, 0.038]
	(-0.921, -0.119)	(-0.184, 0.271)
[0.00, 5.00]	[-1.634, -0.363]	[-0.025, 0.353]
	(-3.873, -0.120)	(-0.181, 1.430)
[0.00, 10.00]	$(-\infty, \infty)$	$(-\infty, \infty)$
	$(-\infty, \infty)$	$(-\infty, \infty)$
λ	6.25	0.40

Note: Bounds on the effect of government share of GDP on per capita pollution emissions, given relative correlation restrictions. Intervals in square brackets are the bounds themselves, while the intervals in the round brackets are the Imbens-Manski 95% cluster-robust asymptotic confidence intervals.
**Significant at 5%

5. Conclusions

This paper, using a sample of 71 countries for the period 1970-2008 examines the effect of government size and income on pollution taking into account the dynamic nature of the relationships. Our results stress the importance of the long-term effects of a change in income or government expenditure on pollution.

The estimated direct effect of government expenditure is negative and significant for SO_2, but insignificant for CO_2. Estimation of a non-positive direct effect of government size on SO_2 is in line with recent findings by Lopez et al. (2011) and Lopez and Palacios (2010). Specifically, the results suggest that the direct effect of government spending on pollution is insignificant and considerably smaller, in absolute values, for CO_2. This finding may be explained by considering both pollutants' impact on human health, as well as the technological capabilities of reducing their levels in the atmosphere. In particular, SO_2 emissions externalities are local and immediate while CO_2 emissions externalities are global and occur mostly in the future. Local environmental degradation, as in the case of SO_2, increases demand for technological improvements to diminish that impact. The difference in magnitude and significance between the estimated direct effects of government expenditure on SO_2 and CO_2 could also be attributed to the different responsiveness of the pollutants to certain policies. In particular, the regulation of production generated pollutants, like SO_2, is expected to be more straightforward and this is reflected in the estimated effects.

Policy implications, occurring from the analysis, differ according to the level of income in a country. Many studies have shown that government size reduces prosperity (Folster and Henrekson, 2001; Bergh and Karlsson, 2010). However, cutting government expenditure should be undertaken with particular care in some levels of GDP. Combining our results with those of Halkos and Paizanos (2013), who also take into account the indirect effect of government expenditure on the environment through its impact on income, reducing government size in developing countries leads to deterioration of environmental quality. Therefore, cutting government expenditure in these countries should be accompanied by appropriate environmental regulation along with the establishment of international environmental treaties. On the other hand, in countries with higher income levels, cutting government expenditures leads to improvements in both income and environmental quality, while our results suggest that these effects are even greater in the long-run. In particular, countries with income level at the decreasing area of the EKC, i.e. developed countries, are more likely to have already established appropriate environmental legislation and to have undertaken public expenditures for the improvement of environmental quality, thus they are susceptible to diminishing returns from a further increase in government size. In that context, cutting out public spending items that increase market failure would be the most beneficial (Lopez et al., 2011).

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