Environ Sci Pollut Res (2017) 24:15044–15054
DOI 10.1007/s11356-017-9109-z

RESEARCH ARTICLE

Are renewable energy policies upsetting carbon dioxide emissions? The case of Latin America countries

José Alberto Fuinhas[1] · António Cardoso Marques[1] · Matheus Koengkan[2]

Received: 21 December 2016 / Accepted: 21 April 2017 / Published online: 10 May 2017

Abstract The impact of renewable energy policies in carbon dioxide emissions was analysed for a panel of ten Latin American countries, for the period from 1991 to 2012. Panel autoregressive distributed lag methodology was used to decompose the total effect of renewable energy policies on carbon dioxide emissions in its short- and long-run components. There is evidence for the presence of cross-sectional dependence, confirming that Latin American countries share spatial patterns. Heteroskedasticity, contemporaneous correlation, and first-order autocorrelation cross-sectional dependence are also present. To cope with these phenomena, the robust dynamic Driscoll-Kraay estimator, with fixed effects, was used. It was confirmed that the primary energy consumption *per capita*, in both the short- and long-run, contributes to an increase in carbon dioxide emissions, and also that renewable energy policies in the long-run, and renewable electricity generation *per capita* both in the short- and long-run, help to mitigate *per capita* carbon dioxide emissions.

Keywords Latin America · CO_2 emissions · Renewable energy policies · Panel autoregressive distributed lag

Research supported by: NECE, R&D unit funded by the FCT—Portuguese Foundation for the Development of Science and Technology, Ministry of Education and Science.

Responsible editor: Philippe Garrigues

✉ José Alberto Fuinhas
fuinhas@ubi.pt

[1] Management and Economics Department, University of Beira Interior, Rua Marquês d´Ávila e Bolama, 6201-001 Covilhã, Portugal

[2] NECE-UBI and University of Beira Interior, R. Marquês de Ávila e Bolama, 6201-001 Covilhã, Portugal

Introduction

The increasing level of carbon dioxide emissions (CO_2) has set off an alarm signal worldwide, causing major concern in the political context and in society in general (Arce et al. 2016). The Latin American (LAM) countries have seen major increases in CO_2 emissions, which have more than doubled during the last three decades (Al-Mulali et al. 2015). For example, in 2010, the region accounted for about 11% of Global Greenhouse Gas (GHG) (Vergara et al. 2013). Despite this continuous increase, the LAM region is a small contributor to the world's GHG (Schipper et al. 2011), but must still be an active player in combating climate change. Policymakers face the dilemma of how to pursue the development of their economies without substantially damaging the environment. Therefore, it is essential that policymakers develop measures to attain economic growth, while mitigating climate change (Sakamoto and Managi 2016). Consequently, several countries have attempted to implement a policy mix of decreasing fossil fuel consumption, while increasing the deployment of renewable energy, with the goal of reducing CO_2 emissions (Sakamoto and Managi 2016). Europe, as well as other regions such as LAM, has adopted policies promoting renewable energy sources (RES). RES policies began in the LAM in the mid-1970s with establishment of the *ProÁlcool* biofuels programme in Brazil in 1975, and geothermal laws in Costa Rica in 1976 and Nicaragua in 1977 (IRENA 2015).

The LAM region is one of the regions with the largest shares of RES, due to hydropower and, more recently, the contribution of biofuels and biomass to the energy supply. Some countries in the LAM region are also becoming among the most dynamic markets for wind, solar and geothermal. Recently, the region has experienced a rapid growth in RES, and there is a visible interest in developing these sources. The LAM region has been confronted with fast economic changes, high energy prices in most of its countries, rapid energy demand, energy security concerns, and enormous biodiversity, and in order to fully explore the potential of RES (e.g. hydropower, wind, solar, geothermal) in most LAM countries, there could be necessity of exports and RES policies. All these facts are a fertile ground for the deployment of renewable energy technologies that have become more attractive and competitive due to the recent decreases in certain technology costs (IRENA 2015). This situation has been incorporated into several policies and laws to support renewable energy sources.

In the literature, the impact of RES policies on CO_2 emissions has scarcely been researched. One example is Arce et al. (2016) who investigated whether RES policies, namely carbon taxes, FITs, premium payments, and quota obligations are efficient in reducing CO_2 emissions. The authors found that carbon taxes are the most cost-effective policy for reducing these emissions. Arce and Sauma (2016) analysed the efficiency of carbon taxes, FITs, premium payments, and quota systems on CO_2 emissions. They found evidence that FITs and premium payments are more cost effective in reducing CO_2 emissions than carbon taxes and quota systems. Redondo and Collado (2014) investigated the impact of premium payments on RES consumption in Spain and found that the use of premium payments implies positive externalities valued at 493 million euros for avoided CO_2 emissions.

The aim of this study is to answer the following question: are renewable energy policies upsetting carbon dioxide emissions? To answer this question, the impact of RES policies on CO_2 emissions will be analysed for ten LAM countries, for the period from 1991 to 2012, using a panel autoregressive distributed lag (ARDL) approach. Indeed, the use of panel data is one of our limited options when we only have short time spans to investigate renewable energy policies in the LAM region. The article addresses the impact of RES polices on CO_2 emissions to identify if these policies are efficient and also makes a contribution to expand the scarce literature on these impacts in LAM countries. Additionally, the choice of LAM countries has the attraction of being a region that (i) has experienced rapid growth in renewable energy investment and is very interested in developing those resources, (ii) has been a pioneer in designing and implementing specific RES promotion mechanisms, and (iii) has been an important player in the innovation and development of RES policies. Based on the results identified in our literature review, our central hypothesis is that RES policies can mitigate CO_2 emissions.

The paper is organised as follows: the next section presents the literature review. The following sections present the model specifications and databases used, the results and the discussion, and finally, the conclusions.

Literature review

The impact of RES policies on CO_2 emissions has barely been researched in the literature. The studies on RES policies have been centred in seven policies (e.g. Arce et al. 2016; Verma and Kumar 2013) specifically: (i) carbon taxes, (ii) feed-in tariffs (FITs), (iii) premium payments, (iv) quota systems, (v) auctions, (vi) cap systems, and (vii) trade systems. There is evidence in the literature that these policies have paved the way for RES and helped to restrain CO_2 emissions. Table 1 presents a summary of the literature review, namely authors, periods, countries, policies, and main conclusions.

The literature provides evidence that premium payments, quota systems, cap systems and trade systems, i.e. all RES policies have paved the way for renewable energy and, therefore, have contributed to the mitigation of greenhouse gas emissions. The following section will highlight the most common RES policies in LAM countries, as well as the main findings of the literature.

Renewable energy policies in LAM countries

The fast growth of RES policies seen in LAM countries could be attributed to the interrelated energy challenges they faced. The region will need a substantial amount of new electricity generation to meet growth in demand and replace ageing infrastructure (Jacobs et al. 2013). Currently, many countries in the LAM region have energy mixes that expose them to fossil fuel price instability. This could significantly affect their national budgets through pass-through provisions in electricity supply contracts and/or climate variability (including droughts), especially those with heavy hydropower structures (Jacobs et al. 2013). These energy challenges have led to an increased interest in the developing of RES in LAM countries. RES policies began in LAM countries in the mid-1970s (IRENA 2015) with establishment of (i) the *ProÁlcool* biofuels programme in Brazil in 1975, (ii) geothermal laws in Costa Rica in 1976, and (iii) an assessment of geothermal resources in Nicaragua in 1977, with the "Master Plan for Electrical Development 1977–2000". From this initial period, a range of different mechanisms emerged that drove growth in the renewable energy market. The most common mechanisms in the region (IRENA 2015) are (i) national renewable energy targets, (ii) auctions, (iii) FITs, (iv) certificate systems, (v) net metering and self-supply, (vi) biofuel blending mandates, (vii) solar mandates, and (viii) local content requirements.

A few authors have focused on the analysis of the impact of RES policies on CO_2 emissions in LAM countries. For instance, Pereira et al. (2011) analysed the best strategies for maintaining the high share of RES in Brazil's electric power generation system. The authors found that the introduction of the energy compensation mechanism had the advantage of being a mechanism that compensated producers, who invested in plants emitting less CO_2. Jacobs et al. (2013) studied FITs in 12 Latin America and Caribbean (LAC) countries. The results indicated that some LAC countries, namely Argentina, Dominican Republic, Ecuador, Honduras, and Nicaragua, have used FITs to promote renewables and

Table 1 Summary of literature review

Author(s)	Period	Country(ies)	Policy(ies)	Conclusion(s)
Arce et al. (2016)	n. a.	n. a.	Carbon taxes; FITs; premium payments; quota obligations.	Carbon tax is the most cost-effective policy for reducing CO_2 emissions.
Thapar et al. (2016)	n. a.	India	Grant/subsidies; accelerated depreciation; tax concessions/exemptions; preferential tariffs; renewable purchase obligations.	Results indicate a high financial impact of these instruments (support of US$ 3–5/MW over applicable tariff), which becomes neutralised when tax inflow is considered. Lower carbon abatement cost (US$ 3–6/t CO2 eq) indicates higher environmental efficacy.
Arce and Sauma (2016)	n. a.	n. a.	Carbon taxes; FITs; premium payments and quota systems.	The FITs and premium payments are more cost effective in reducing CO_2 emissions than carbon taxes and quota systems.
Redondo and Collado (2014)	2011	Spain	Premium payments	The use of premium payments implies positive externalities valued at 493 million euros in terms of avoided CO_2 emissions.
Ortega et al. (2013)	2002–2011	Spain	FITs	FITs encourage the use of RES and the reduction of CO_2 emissions.
Verma and Kumar (2013)	n. a.	n. a.	Carbon quotas; Cap-and-trade and bilateral IPPs.	All policies contribute to the reduction of CO_2 emissions.
Stokes (2013)	1997–2012	Canada	FITs	FITs can reduce the cost of renewable energy, and speed deployment, supporting much-needed decarbonisation.
Hinrichs-Rahlwes (2013)	1998–2009	Germany	FITs	Mitigate climate change in the best possible way.
Green et al. (2007)	n. a.	n. a.	Carbon taxes	Carbon tax policies could help to reduce CO_2 emissions associated with conventional energy.
Wüstenhagen and Bilharz (2006)	1973–2003	Germany	FITs	This policy contributes to the reduction of greenhouse gases emissions.
Palmer and Burtraw (2005)	n. a.	n. a.	REPC and RPS	RPS policies appear to be more cost-effective than REPC policies in both promoting renewables and reducing carbon.

n. a. not available, *FITs* feed-in tariffs, *REPC* renewable energy production credit, *RPS* renewables portfolio standard, *IPPs* independent power producers, *CO_2* carbon dioxide emissions, *RES* renewable energy sources, *MWs* megawatts.

reduce CO_2 emissions, and that FITs are becoming increasingly popular. If well designed, they can mitigate investor risk in RES. Zwaan et al. (2016) investigated opportunities for energy technology deployment as part of climate change mitigation efforts in LAM countries. The authors project several RES policy scenarios up to 2050, which could reduce CO_2 emissions.

Data and model specification

Data

The availability of data was the main criterion for selecting both the countries and time span of the analysis. Indeed, all the available data was used, but even so, the need to use aggregated data was unavoidable, as in the case of public policies supporting renewables. Most of these policies became effective only just prior to the 1990s, and even so, most of the individual policies were not used by several countries. As a consequence, the individual series are plagued by the well-known 'excess of zeros problem'. This fact could be overcome by using the policies all together, which is not new in the literature (e.g. Polzin et al. 2015; Marques and Fuinhas 2012; Johnstone et al. 2010). This indicator of public policies on renewable energy has the shortcoming of not capturing the strength of policies, as it only registers their deployment. A precise measurement of the intensity of policies is nearly impossible because of both the unavailability of data and the diverse particularities of countries (e.g. Zhao et al. 2013). Overall, this was not a severe constraint given that the objective was to assess the effectiveness of the public intervention. It is worthwhile to note that the LAM countries extensively use hydropower to generate electricity, which could be seen as a natural barrier to the diversification of renewable sources, making the reduction the CO_2 emissions more difficult. As such, the inclusion of a total renewable energy policy variable in the model is an effort to determine whether the intensity of public policy interventions were able to stimulate the deployment of the renewables.

Ten LAM countries, namely Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Mexico, Nicaragua, Peru, and Uruguay, meet the criteria of having data available for the entire period for RES consumption, CO_2 emissions, primary energy consumption, and RES policies. More specifically, the variables are (i) carbon dioxide emissions, from consumption of energy (million metric tons); (ii) renewable energy consumption (kilowatt-hours, from hydroelectric, geothermal, wind, solar, tide, wave, and biomass); (iii) primary energy consumption (in quadrillion Btu) from fossil fuels and other sources, transformed into per capita values; (iv) Gross Domestic Product (GDP), in constant local currency unity (LCU) and transformed into per capita values; and (v) renewable energy policies. This latter variable includes all policies defined by the International Energy Agency (IEA), namely: (a) Economic Instruments; (b) Information and Education; (c) Policy Support; (d) Regulatory Instruments; (e) Research, Development and Deployment (RD&D); and (f) Voluntary Approaches. All the variables, except the renewable energy policies, were transformed into per capita values. The uses of per capita values let us control for disparities in population growth among the Latin America countries. Table 2 shows the name, the definition, and the source of raw data.

Given that renewable energy policies are likely to require time to produce their full effect on CO_2, a panel ARDL model approach was used. The properties of this estimation method allow the decomposition of the total effect into its short- and long-run dimensions. Accordingly, to achieve the goal of decomposing the global effects in the short- and long-run, we balanced the longest available time span with the maximum possible number of LAM countries that have renewable energy policies available. The EViews 9.5 and Stata 14.2 software were used.

Considering that we are working upon a macro panel, the best econometric practices strongly recommend testing for the presence of heterogeneity, which could arise when a long time span is used. Indeed, long time spans exacerbate the potential occurrence of a panel with parameter slope heterogeneity and the presence of cross-section dependence (CSD). Indeed, in LAM countries, it is expected that the existence of CSD due the countries would share common characteristics. When the presence of CSD is not controlled, it can produce both biased estimates and a severe identification problem (e.g. Eberhardt and Presbitero 2013), which require appropriate estimators to handle them. The descriptive statistics, the CSD, and the order of integration of the variables are analysed to capture the features of both series and crosses. Table 3 reveals both the descriptive statistics (for panel descriptive statistics, see Table 9) and the CSD of variables. Hereafter the prefixes (L) and (D) denote natural logarithm, and first differences of the variables, respectively.

The CSD-test points to the presence of cross-section dependence in the variables both in levels and in first differences, except for the RES generation in differences (DLRE). A possible answer for this result is that the generation of RES is largely country specific and conditional on the intermittence that characterises its generation (for example, solar and wind sources). The presence of CSD shows evidence of interdependence between the cross-sections, i.e. that the countries share common shocks.

Model specification

To analyse the impact of RES policies on CO_2 emissions an unrestricted error correction model (UECM) form of the ARDL model was used. This model decomposes the total effect of a variable into its short- and long-run components (e.g., Srinivasan et al. 2012). Moreover, it generates consistent and efficient parameter estimates, as well as the inference of parameters based on standard test. The dynamic general UECM form of the ARDL model, which decomposes the total effect into short- and long-term effects, is used in this empirical analysis follow the specification of Eq. (1):

$$\begin{aligned} DLCO2_{it} = {} & \alpha_i + \delta_i\, TREND_t + \sum_{j-1}^{k} \beta 1_{ij}\, DLRE_{it-j} + \sum_{j-1}^{k} \beta 2_{ij}\, DLPOL_{it-j} + \sum_{j-1}^{k} \beta 3_{ij}\, DLE_{it-j} + \\ & \sum_{j=0}^{k} \beta 4_{ij}\, DLY_{it-j} + \gamma 1_i\, LCO2_{it-1} + \gamma 2_i\, LRE_{it-1} + \gamma 3_i\, LPOL_{it-1} + \\ & \gamma 4_i\, LE_{it-1} + \gamma 5_i\, LY_{it-1} + \varepsilon i_t \end{aligned} \tag{1}$$

where αi denotes the intercept, δi , βkij , $k = 1, \ldots, m$, and γim are the estimated parameters, and εi the error term.

Given that the potentially strong relationships between variables such as CO_2 emissions, renewable energy consumption, and GDP, it is prudent to assess if multicollinearity is a concern in estimation of models. The variance inflation factor (VIF) provides an indication of the impact of multicollinearity on the accuracy of estimated regression coefficients (e.g. O'Brien 2007). As such, both the VIF statistics and correlation coefficients between variables were computed (see Table 4).

As can be seen in Table 4, only the correlation between the consumption of primary energy (LE) and CO2 emissions

Table 2 Variable description

Variable	Acronym	Source
Carbon dioxide emissions	CO2	Energy Information Administration (EIA)
Renewable electricity consumption	RE	EIA
Primary electricity consumption	E	EIA
Gross domestic product	Y	World Bank—World DataBank
Renewable energy policies	POL	International Energy Agency (IEA)

(LCO2) reveals a high correlation coefficient. Given that this correlation is between an independent variable and the dependent one, this does not give rise to any econometric problem. The low VIF statistics support the argument that multicollinearity is of no great concern in the model.

To assess the order of integration of the variables, first- and second-generation unit root tests were used. The first-generation unit root tests of LLC (Levin et al. 2002), ADF-Fisher (Maddala and Wu 1999), and ADF-Choi (Choi 2001), were used. The second-generation unit root test CIPS of Pesaran (2007) was used. The CIPS test is robust to heterogeneity and CSD, and tests for the null of non-stationary, under a nonstandard distribution. Table 5 shows the results of unit root tests.

The test LLC, ADF-Fisher, ADF-Choi, and CIPS are consensual in supporting that all the variables in levels, except LRE, are integrated of order one, I(1), i.e. they have one unit root. The LRE and all the variables in first differences are stationary.

The Hausman test of the RE against the FE specification was applied, to identify the presence of RE or FE in the model. This test has the null hypothesis that the best model is RE. The statistically highly significant Hausman test ($X^2_{10} = 70.03$) allows us to select the FE model. Moreover, the FE model is appropriate for analysing the influences of variables over time, as well as removes all time invariant features from the independent variables. To check the cointegration of results, the second-generation cointegration test of Westerlund (2007) was used. This test has a null hypothesis the existence of no-cointegration between the variables. Indeed, the Westerlund cointegration test is based in an error correction model, where all variables are stationary (Fuinhas et al. 2015).

In the macro panels, the presence of long time spans and many cross-sections makes testing for the slope heterogeneity of parameters highly advisable. This testing could be of two types: (i) heterogeneity of parameters in the short- and long-run and (ii) heterogeneity of parameters only in the short-run. To deal with heterogeneity, the mean group (MG) or pooled mean group (PMG) estimators could be applied. The MG is a flexible technique, which creates regressions for each individual and then computes for all individuals an average coefficient (Pesaran et al. 1999). Indeed, this estimator is consistent in long-run average, while in presence of slope homogeneity, the model is not efficient (Pesaran et al. 1999). The PMG is an estimator that in long-run parameters makes restrictions among cross-sections but not in short-run and in adjustment speed term. Moreover, the PMG estimator is more efficient and consistent in the existence of homogeneity in the long-run if compared with MG estimator (Fuinhas et al. 2015).

Finally, a battery of diagnostic tests was performed: (i) modified Wald test for groupwise heteroskedasticity. This test has the null hypothesis of homoscedasticity; (ii) Pesaran (2004) test of cross-section independence, to identify the presence of contemporaneous correlation

Table 3 Descriptive statistics and cross-section dependence test

	Descriptive statistics					Cross-section dependence (CSD)			
	Obs	Mean	Std. Dev.	Min.	Max.	CD test		Corr.	Abs(Corr)
LCO2	220	−13.2156	5.5640	−14.6042	−12.2706	19.13	*	0.608	0.627
LRE	220	−14.2404	8.4719	−16.4111	−12.7685	8.30	**	0.264	0.366
LPOL	220	1.18910	1.0711	0.0000	3.66356	25.16	*	0.800	0.800
LE	220	−17.1651	6.0144	−18.6029	−16.2434	24.01	*	0.763	0.763
LY	220	10.8001	2.6872	7.7480	16.1225	28.63	*	0.910	0.910
DLCO2	210	0.2362	0.6815	−2.7756	2.2509	2.75	***	0.089	0.187
DLRE	210	0.1466	1.6448	−6.1359	8.0862	−0.52		−0.017	0.214
DLPOL	210	1.1703	2.6883	−1.5415	1.9459	1.43	***	0.047	0.175
DLE	210	0.2164	0.6641	−2.0858	2.6329	7.18	**	0.233	0.278
DLY	210	0.2292	0.3326	−1.2644	0.9999	11.05	*	0.360	0.360

Pesaran (2004) CD test has N (0,1) distribution, under the H_0: cross-section independence. ***, **, * denote statistically significant at 1, 5, and 10% level, respectively.

Table 4 Matrices of correlations and VIF statistics

	LCO2		LRE		LPOL		LE		LY
LCO2	1.0000								
LRE	0.4717	***	1.0000						
LPOL	0.2872	***	0.0591	**	1.0000				
LE	0.9594	***	0.6769	***	0.2794	***	1.0000		
LY	0.3073	***	0.3792	***	0.0148		0.3383	***	1.0000
VIF			1.99		1.13		2.10		1.19
Mean VIF					1.60				
	DLCO2		DLRE		DLPOL		DLE		DLY
DLCO2	1.0000								
DLRE	-0.2746	***	1.0000						
DLPOL	-0.0060		0.0444		1.0000				
DLE	0.4504	***	0.3155	***	0.0153		1.0000		
DLY	0.2753	***	0.0916		−0.0410		0.3346	***	1.0000
VIF			1.11		1.00		1.24		1.13
Mean VIF					1.12				

*** denotes statically significant at 1% level.

among cross-sections. The null hypothesis of this test specifies that the residuals are not correlated and it follows a normal distribution; (iii) Breusch and Pagan (1980) Langrarian Multiplier test of cross-sectional independence that follows chi-square distribution was performed to measure whether the variances across individuals are correlated; (iv) Frees (1995, 2004) test of cross-sectional independence; (v) Friedman (1937) test of cross-sectional independence; and (vi) Wooldridge (2002) test, to check for the existence of serial correlation.

Results and discussion

As stated earlier, the aim of this research is to examine the effect of RES policies on CO_2 emissions in LAM countries. It is worthwhile noting that the results are based on per capita data. There is evidence of the presence of CSD in the variables (see Table 3). The test of unit roots (see Table 5) points to the possibility of stationary of LRE. Table 6 shows the results of the Westerlund cointegration tests.

The Westerlund cointegration tests, considering both the panel as a whole, and each country individually reject cointegration. Moreover, the non-detection of cointegration among variables points to the use of econometric techniques that are less stringent about integration variables, i.e. ARDL models.

The MG and PMG estimators were tested against the dynamic fixed effects (DFE) estimator. The Driscoll and Kraay (1998) estimator was applied (e.g. by Fuinhas et al. 2015; Hoechle 2007) to cope with the presence of heteroskedasticity, contemporaneous correlation, first-order autocorrelation, and cross-sectional dependence (spatial dependence or spatial regimes). Moreover, this estimator is a matrix estimator that

Table 5 Unit roots tests

	1st generation unit root tests						2nd generation unit root tests			
	LLC		ADF-Fisher		ADF-Choi		CIPS (Zt-bar)			
	Individual intercept and trend						Without trend		With trend	
LCO2	−1.0714		24.3974		−0.6699		−0.776		0.969	
LRE	−4.2044	***	39.3896	***	−2.6446	***	−1.337	***	−1.300	***
LPOL	−0.8576		17.7105		0.1587		−0.404		1.056	
LE	−0.5597		27.8676		−1.0734		−0.678		1.259	
LY	0.6792		18.9771		1.0888		−1.199		−0.750	
DLCO2	−6.7437	***	83.8301	***	−6.6476	***	−4.976	***	−4.710	***
DLRE	−13.0036	***	139.080	***	−9.6431	***	−6.254	***	−5.157	***
DLPOL	−6.0603	***	65.5947	***	−5.1363	***	−4.038	***	−3.413	***
DLE	−7.3999	***	113.166	***	−8.0571	***	−3.290	***	−1.868	***
DLY	−6.5306	***	68.1892	***	−5.3035	***	−3.826	***	−2.377	***

*** denotes statistically significant at 1% level. The LLC test has H_0: unit root (common unit root process), the test controls for individuals' effects, individual linear trends, has a lag length 1, and Newey-West automatic bandwidth selection and Bartlett kernel were used; the ADF-FISHER and ADF-Choi test has H_0: unit root (individual unit root process), the test controls for individual effects, individual linear trends, has a lag length 1. The CIPS test has H_0: series are I(1).

Table 6 Westerlund cointegration tests

	Westerlund cointegration test								
	None			Constant			Constant and trend		
Statistics	Value	*Z* value	*P* value robust	Value	*Z* value	*P* value robust	Value	*Z* value	*P* value robust
Gt	−1.777	0.622	0.228	−2.547	−0.337	0.098	−2.484	1.323	0.343
Ga	−5.517	1.940	0.088	−6.916	2.493	0.063	−4.622	4.661	0.466
Pt	−5.445	−0.263	0.116	−5.235	1.439	0.384	−5.957	2.354	0.406
Pt	−5.758	0.154	0.026	−5.856	1.439	0.111	−4.063	3.609	0.429

Bootstrapping regression with 800 reps. H_0: No cointegration; H_1 Gt and Ga test the cointegration for each country individually, and Pt and Pa test the cointegration of the panel as whole.

generates robust standard errors for several phenomena found in the sample errors. The DFE estimator, DFE robust standard errors, and DFE Driscoll and Kraay (DFE D.-K.) were computed. Finally, a set of specification tests as (i) modified Wald test, (ii) Pesaran test, (iii) Breusch and Pagan Langrarian multiplier test, (iv) Frees test, (v) Friedman test, and (vi) Wooldridge test was applied.

Table 7 shows the results of the estimations for the MG, PMG, DFE models, and the outcome of the Hausman test and also exhibits the short-run semi-elasticities and long-run elasticities for the DFE, DFE Robust, and DFE D.-K models and, finally, demonstrates the results of the specification tests. The semi-elasticities were computed by adding the coefficients of variables in the first differences. The elasticities are computed by dividing the coefficient of the variables by the coefficient of LCO2, both lagged once and multiplying the ratio by −1.

The Hausman test indicates that the DFE is the appropriate estimator, i.e. there is evidence that the panel is 'homogeneous'. The estimations resulting from the DFE estimator, DFE robust standard errors, and DFE Driscoll and Kraay (DFE D.-K.) point to the presence of long memory in the variables and the ECM term is statistically significant at 1%

Table 7 Estimation results

Models (dependent variable DLCO2)								
	Heterogeneous estimator				Fixed effects			
	MG (I)		PMG (II)		Coefficient	DFE (III)	DFE Robust (IV)	DFE D.-K. (V)
Constant	−2.7910		−4.5068	***	−5.2545	***	***	***
Trend	−0.0013		−0.0027	**	0.0006			
					Short-run (semi-elasticities)			
DLRE	−0.2279	***	−0.1676	***	−0.1854	***	***	***
DLPOL	0.0173		0.0502		0.0061			
DLE	0.8176	***	0.7630	***	0.5822	***	***	***
DLY	0.5204	***	0.5203	***	0.4276	***	***	***
					Long-run (elasticities)			
LRE(−1)	−1.0157		−0.1163	***	−0.1965	***	***	***
LPOL(−1)	−0.0414		0.0078		−0.0358	***	***	***
LE(−1)	2.7717		0.6951	***	0.7082	***	***	***
LY(−1)	−1.2185		0.3588	***	0.4776	***	***	***
					Speed of adjustment			
ECM(−1)	−0.9598	***	0.6763	***	−0.5850	***	***	***
	Hausman test				Specification tests			
	MG vs PMG		PMG vs DFE		Modified Wald	Pesaran		Wooldridge
	$\chi^2_{11} = 20.30$		$\chi^2_{11} = 0.00$***		$\chi^2_{11} = 172.31$***	$N(0,1) = 4.71$***		$F(1,9) = 111.47$***
					Breusch-Pagan LM	Frees		Friedman
					$\chi^2_{45} = 65.597$**	−0.038		37.460***

Hausman results for H_0: difference in coefficients not systematic; ECM denotes error correction mechanism; the long-run parameters are computed elasticities; for H_0 of Modified Wald test: sigma(i)^2 = sigma^2 for all I; results for H_0 of Breusch-Pagan LM test, Pesaran's test, Frees' test, and Friedman's: cross-sectional independence in the residuals; results for H_0 of Wooldridge test: no first-order autocorrelation.

*** and ** denote statistically significant at 1 and 5% level, respectively level and has a negative sign. This result also confirms the presence of Granger causality from statistically significant variables to CO2 emissions. Indeed, the unrestricted error correction model (UECM) form of an autoregressive distributed lag (ARDL) model allows us to discriminate between short- and long-run Granger causality. Indeed, the UECM-ARDL model is widely known as the Cointegration and Error Correction version of Granger causality (e.g. Jouini 2015; Mehrara 2007). Furthermore, the ARDL methodology is robust to the presence of endogeneity of variables. Given that the ECM parameter is statistically significant and is negative, we can thus consider that when a parameter is statistically significant it will be identical for testing Granger causality. Given the Cointegration and Error Correction version of Granger causality, we can ensure that both the causality and the magnitude of the effects are revealed by the elasticities themselves. Finally, the battery of specification test, like the modified Wald test, points to the statistically highly significant presence of heteroskedasticity. The Breusch-Pagan LM test, the Pesaran's test, and the Friedman's test identified the presence of cross-sectional independence in the residuals. The Wooldridge test that checks for the existence of serial correlation proved to be statistically highly significant, pointing to the presence of first-order autocorrelation.

The results show that the long-run elasticities of the RES policies variable exert a negative outcome on CO_2 emissions. RES consumption reduces CO_2 emissions both in the short- and long-run. Primary energy consumption (LE) and economic growth (LY) increase CO_2 emissions in both the short- and long-run. The negative sign of LPOL could be due to the implementation of RES policies that increase the introduction of RES into the energy mix. The finding that primary energy consumption and economic growth increase CO_2 emissions in both the short- and long-run may result from the evidence that the LAM economies are still highly dependent on fossil fuels. This dependence may be due to the fact that many of these countries are major fossil fuel producers, such as Argentina, Brazil, Colombia, Ecuador, Mexico, Peru, and Venezuela, or because they are dependent on imports, such as the Central American countries and Chile.

The capacity of RES policies to reduce CO_2 emissions is probably also related to the efficiency gains associated with these policies. For example, in LAM countries, the most efficient policies are the *national renewable energy targets*, which provide a clear indication about the intended level of renewable energy development and the timeline envisioned by governments. For this reason, several countries in LAM have also established their own formal renewable energy targets by legislation or decree. Another extremely effective and very popular policy in LAM countries is that of *auctions*. Auctions refer to competitive bidding procurement processes for electricity from renewable energy sources or where renewable energy technologies are eligible. Moreover, RES auctions in LAM countries usually offer a long-term power purchase agreement (PPA), with durations ranging from 10 to 30 years to successful bidders. In others words, state participation, through laws, decrees, and auctions to incentivize investments in RES and spread incorporation of RES in the energy matrix of the region, is very large. However, the vast state participation is because the region faces a series of interrelated energy challenges. On the one hand, the LAM region will need a substantial amount of new electricity generation to meet growth in demand and to replace ageing infrastructure. On the other hand, many LAM countries have undiversified energy portfolios and are very exposed to fossil fuel price instability that could seriously affect their national budgets. Moreover, the investments in renewable energy sources could result from the availability of enormous biodiversity and the abundance of renewable sources (e.g. hydropower, wind, solar, and geothermal) in most LAM countries. This abundance encourages the deployment of RE technologies, brings new investments and, consequently, creates jobs, fosters economic growth, and reduces CO_2 emissions (e.g. Alvarado and Toledo 2016). In addition, they have a strong incentive to implement low-carbon generation into their energy systems to reduce CO_2 emissions and take advantage of the financial resources available throughout the international climate negotiations. It is worthwhile to note that RES policies, when implemented, even if inactive, continues producing stimuli over time.

Robustness check

As is well known, the LAM countries suffered several economic and political shocks that have affected carbon dioxide emissions in various ways. These shocks were caused by a series of both domestic and external crises in LAM countries that began in the 1990s and had impacts on the real economy. In 2001, Bolivia (BOL) experienced social tensions that led to the blocking of roads and violent clashes between army troops and peasants who opposed the eradication of coca crops and the *Aguas de Ley* (Water Laws), preventing the operation of networks (Bandeira 2002). Additionally, these social tensions generated many economic and political impacts. During the severe crisis of peso convertibility that affected Argentina between 2001 and 2002, many of the country's customers withdrew their dollar deposits held in Uruguayan banks. This caused a crisis in the financial system in 2001 (Brun and Licandro 2005). Chile (CHL), in 2007, saw a reverse trend, mainly driven by the international crisis that affected industry, forestry, and steel and further supplemented by reduced diesel generation due to the rearrangement of the matrix of power generation to move away from using this fuel (BCG 2013). Finally, Uruguay (URY) in 2009 was also impacted by the international crisis in 2008–2009; however, the impact has been very moderate and the country only showed a decline in GDP in the first quarter of 2009, having continued to grow thereafter (IMF, International Monetary Fund 2010). The shocks that affect CO_2 emissions were confirmed in the residuals of the model for the years 2001, 2007, and 2009. By using dummy variables (BOL2001, URY2001, CHL2007, and URY2009), the statistical significance of these shocks was properly tested. Table 8 shows the results of the estimation of semi-elasticities and the elasticities for the DFE, DFE Robust, and DFE D.-K. models including the dummy variables.

To select the better model (with or without dummy variables) the likelihood-ratio test was applied, which performs a likelihood-ratio test of the null hypothesis that the parameter vector of a statistical model satisfies some mild constraint. The results of the likelihood-ratio test ($\chi^2_4 = 56.25$), suggest that the unrestricted model is better, as it is statically significant at 1% level. The shocks proved to be statistically significant at 1% level. Furthermore, as can be seen by comparing Tables 7 and 8, the results of both models are the same, proving the robustness of the approach pursued, even in the presence of shocks.

Table 8 Estimation results with shocks

Models (dependent variable DLCO2)				
	Fixed effects			
	Coefficient	FE (VI)	FE Robust (VII)	FE D.-K. (VIII)
Constant	−4.7448	***	***	***
Trend	0.0006			
	Dummy variables			
BOL2001	−0.1081	***	***	***
URY2001	−0.2054	***	***	***
CHL2007	−0.1573	***	***	***
URY2009	0.1424	***	***	***
	Short-run (semi-elasticities)			
DLRE	−0.1634	***	***	***
DLPOL	−0.0055			
DLE	0.5822	***	***	***
DLY	0.3733	***	***	***
	Long-run (elasticities)			
LRE(−1)	−0.1433	***	***	***
LPOL(−1)	−0.0415	***	***	***
LE(−1)	0.6945	***	***	***
LY(−1)	0.4953	***	***	***
	Speed of adjustment			
ECM(−1)	−0.5494	***	***	***

*** denotes statistically significant at 1% level.

Conclusion

The impact of renewable energy policies on CO_2 emissions was analysed for ten Latin America countries, for the period from 1991 to 2012. Using a panel Auto-Regressive Distributed Lag approach, the presence of cross-sectional dependence was proven, thus confirming that these countries share spatial patterns, heteroskedasticity, contemporaneous correlation, and first-order auto-correlation. The results show that renewable energy policies reduced carbon dioxide emissions in the long-run. Indeed, the ability of renewable energy policies to reduce CO_2 emissions, in the long-run, was probably also related to the efficiency gains associated with these policies. Renewable energy consumption decreases CO_2 emissions both in short- and long-run. This result could be a consequence of renewable energy policies, which substitute the use of fossil fuels with the production and use of RES in Latin American countries. This finding supports the relevance of designing public policies to diversify the renewable mix, in order to reduce CO2 emissions. Primary energy consumption increases CO_2 emissions in both the short- and long-run. One possible explanation for this occurrence could be the presence of fossil fuels in the energy matrix in some Latin America countries, such as Mexico, Chile, Colombia, and Bolivia. Finally, economic growth increases CO_2 emissions both in the short- and long-run. This result shows that Latin American economies are still highly dependent on fossil fuels for growth. This dependence may be due to the fact that many of these countries are major fossil fuel producers, such as Argentina, Brazil, Colombia, Ecuador, Mexico, Peru, and Venezuela, or because they are dependent on imports, such as the Central American countries and Chile.

The robustness of the model was proven by identifying and including in the model the main shocks that occurred in the Latin American countries. Moreover, this evidence points to the necessity of creating new renewable energy policies to promote production and consumption. Indeed, the impact of renewable energy policies on CO_2 emissions is small. These findings indicate the need for policymakers to change the current energy mix to a more sustainable one, as well as for the need to develop new renewable policies, designed to promote economic growth and environmental sustainability. Indeed, renewable energy policies have the capacity to bring new investments in RES and foster the economy of countries or regions, given that renewable energy policies have a propensity to generate income and contribute to reducing CO_2 emissions.

Acknowledgments The financial support of the NECE-UBI, Research Unit in Business Science and Economics, sponsored by the Portuguese Foundation for the Development of Science and Technology, project UID/GES/04630/2013, is gratefully acknowledged.

Appendix

Table 9 Panel descriptive statistics

Variables		Mean	Std.Dev	Min.	Max.	Observations	
DLCO2	Overall	0.0236	0.0682	−0.2776	0.2251	N =	210
	Between		0.0126	0.0022	0.0416	n =	10
	Within		0.0671	−0.2778	0.2135	T =	21
DLRE	Overall	0.0147	0.1645	−0.6136	0.8086	N =	210
	Between		0.0082	−0.0010	0.0246	n =	10
	Within		0.1643	−0.5979	0.8243	T =	21
DLPOL	Overall	0.1170	0.2688	−0.1542	1.9459	N =	210
	Between		0.0314	0.0523	0.1402	n =	10
	Within		0.2672	−0.1747	1.9309	T =	21
DLE	Overall	0.0216	0.0664	−0.2086	0.2633	N =	210
	Between		0.0109	0.0052	0.0374	n =	10
	Within		0.0656	−0.2070	0.2649	T =	21
DLY	Overall	0.0229	0.0333	−0.1264	0.0999	N =	210
	Between		0.0083	0.0110	0.0365	n =	10
	Within		0.0323	−0.1299	0.0956	T =	21
LCO2	Overall	−13.2157	0.5564	−14.6042	−12.2706	N =	220
	Between		0.5615	−14.1508	−12.4872	n =	10
	Within		0.1566	−13.6987	−12.7419	T =	22
LRE	Overall	−14.2405	0.8472	−16.4111	−12.7685	N =	220
	Between		0.8713	−15.7312	−13.0923	n =	10
	Within		0.1771	−14.9204	−13.5535	T =	22
LPOL	Overall	1.1891	1.0712	0.0000	3.6636	N =	220
	Between		0.5256	0.5249	1.8890	n =	10
	Within		0.9475	−0.6978	3.4368	T =	22
LE	Overall	−17.1651	0.6014	−18.6029	−16.2434	N =	220
	Between		0.6113	−18.3019	−16.4363	n =	10
	Within		0.1546	−17.6587	−16.6639	T =	22
LY	Overall	10.8001	2.6873	7.7480	16.1225	N =	220
	Between		2.8219	7.9272	15.8541	n =	10
	Within		0.1469	10.3558	11.2190	T =	22

The Stata command *xtsum* was used to achieve the results for the panel's between and within statistics

References

Al-Mulali U, Tang CF, Ozturk I (2015) Estimating the environment Kuznets curve hypothesis: evidence from Latin America and the Caribbean countries. Renew Sust Energ Rev 50:918–924

Alvarado R, Toledo E (2016) Environmental degradation and economic growth: evidence for a developing country. Environ Dev Sustain:1–14

Arce MP, Sauma E (2016) Comparison of incentive policies for renewable energy in an oligopolistic market with price-responsive demand. Energy J 37(3):159–179

Arce MP, Sauma E, Contreras J (2016) Renewable energy policy performance in reducing CO_2 emissions. Energy Econ 54:272–280

Bandeira LAM (2002) As políticas neoliberais e a crise na América do Sul. Rev Bras Polít Int 45(2):135–146

BCG (The Bonton Consulting Group) (2013) Inventario de emisiones de GEI 1990–2010, proyección de emisiones a 2040 y matrices de abatimiento de CO2–Chile:1–134

Breusch TS, Pagan AR (1980) The Lagrange multiplier test and its applications to model specification tests in econometrics. Rev Econ Stud 47:239–253

Brun J, Licandro G (2005) To hell and back, crisis management in a dollarized economy: the case of Uruguay. ISSN 1688-7565:1–40

Choi I (2001) Unit root test for panel data. J Int Money Financ 20(1):249–272

Driscoll J, Kraay AC (1998) Consistent covariance matrix estimation with spatially dependent data. Rev Econ Stat 80(4):549–560

Eberhardt M, Presbitero AF (2013) This time they are different: heterogeneity and nonlinearity in the relationship between debt and growth. IMF, Working Paper 13/248

Frees EW (1995) Assessing cross-sectional correlations in panel data. J Econ 69(2):393–414

Frees EW (2004) Longitudinal and panel data: analysis and applications in the social sciences. Cambridge University Press

Friedman M (1937) The use of ranks to avoid the assumption of normality implicit in the analysis of variance. J Am Stat Assoc 32(200): 675–701

Fuinhas JA, Marques AC, Couto AP (2015) Oil rents and economic growth in oil producing countries: evidence from a macro panel. Econ Chang Restruct 48(3):257–279

Green K, Hayward S, Hassett K (2007) Climate change: caps vs. taxes. American Enterprise Institute. Environmental and Energy Economics

Hinrichs-Rahlwes R (2013) Renewable energy: paving the way towards sustainable energy, security lessons learnt from Germany. Renew Energy 49:10–14

Hoechle D (2007) Robust standard errors for panel regressions with cross-sectional dependence. Stata J 7(3):281–312

IMF (International Monetary Fund) (2010) Uruguay: 2009 article IV consultation-staff report; staff statement and supplement; public Information notice on the executive board discussion and statement by the executive director for Uruguay. IMF Country Rep N 10(48): 1–40

IRENA (2015) Renewable energy in Latin America 2015: an overview of policies, The International Renewable Energy Agency

Jacobs D, Marzolf N, Paredes JR, Rickerson W, Flynn H (2013) Analysis of renewable energy incentives in the Latin America and Caribbean region: the feed-in tariff case. Energ Policy 60:601–610

Johnstone N, Haščič I, Popp D (2010) Renewable energy policies and technological innovation: evidence based on patent counts. Environ Resour Econ 45(1):133–155

Jouini J (2015) Economic growth and remittances in Tunisia: bi-directional causal links. J Policy Model 37(2):355–373

Levin A, Lin C-F, Chu C-SJ (2002) Unit root test in panel dada: asymptotic and finite-sample properties. J Econ 108(1):1–24

Maddala GS, Wu S (1999) A comparative study of unit root test with panel data a new simple test. Oxf Bull Econ Stat 61(1):631–652

Marques AC, Fuinhas JA (2012) Are public policies towards renewables successful? Evidence from European countries. Renew Energy 44: 109–118

Mehrara M (2007) Energy consumption and economic growth: the case of oil exporting countries. Energ Policy 35(5):2939–2945

O'Brien RM (2007) A caution regarding rules of thumb for variance inflation factors. Qual Quant 41(5):673–690

Ortega M, Del Rio P, Montero EA (2013) Assessing the benefits and costs of renewable electricity. The Spanish case. Renew Sust Energ Rev 27:294–304

Palmer K, Burtraw D (2005) Cost-effectiveness of renewable electricity policies. Energy Econ 27(6):873–894

Pereira AO Jr, Pereira AS, La Rovere EL, Barata MML, Villar SC, Pires SH (2011) Strategies to promote renewable energy in Brazil. Renew Sust Energ Rev 15(1):681–688

Pesaran MH (2004) General diagnostic tests for cross section dependence in panels, Cambridge Working Papers in Economics No. 0435, Faculty of Economics, University of Cambridge

Pesaran MH (2007) A simple panel unit root test in the presence of cross-section dependence. J Appl Econ 22(2):256–312

Pesaran MH, Shin Y, Smith RP (1999) Pooled mean group estimation of dynamic heterogeneous panels. J Am Stat Assoc 94(446):621–634

Polzin F, Migendt M, Täube FA, von Flotow P (2015) Public policy influence on renewable energy investments—a panel data study across OECD countries. Energ Policy 80:98–111

Redondo AJG, Collado RR (2014) An economic valuation of renewable electricity promoted by feed-in system in Spain. Renew Energy 68: 51–57

Sakamoto T, Managi S (2016) Optimal economic growth and energy policy: analysis of non-renewable and renewable energy. Environ Econ Policy Stud Springer J 18(1):1–19

Schipper L, Deakin E, McAndrew C (2011) Carbon dioxide emissions from urban road transport in Latin America: CO_2 reduction as a co-benefit of transport strategies. Transport moving to climate intelligence, part of the series transportation research. Econ Policy 111–127

Srinivasan P, Kumar PK, Santhosh, Ganesh L (2012) Tourism and economic growth in Sri Lanka: an ARDL bounds testing approach. Rom Econ J 15(45):211–226

Stokes L (2013) The politics of renewable energy policies: the case of feed-in tariffs in Ontario, Canada. Energ Policy 56:490–500

Thapar S, Sharma S, Verma A (2016) Economic and environmental effectiveness of renewable energy policy instruments: best practices from India. Renew Sust Energ Rev 66:487–498

Vergara W, Rios RR, Galindo LM (2013) The climate and development challenge for Latin America and the Caribbean: options for climate-resilient, low-carbon development no. 978–1–59782-165-0 (IADB)

Verma YP, Kumar A (2013) Potential impacts of emission concerned policies on power system operation with renewable energy sources. Int J Electr Power Energy Syst 44(1):520–529

Westerlund J (2007) Testing for error correction in panel data. Oxf Econ Stat 31(2):217–224

Wooldridge JM (2002) Econometric analysis of cross section and panel data. The MIT Press, Cambridge, Massachusetts

Wüstenhagen R, Bilharz M (2006) Green energy market development in Germany: effective public policy and emerging customer demand. Energ Policy 34(13):1681–1696

Zhao Y, Tang KK, Wang L (2013) Do renewable electricity policies promote renewable electricity generation? Evidence from panel data. Energ Policy 62:887–897

Zwaan BVD, Kober T, Calderon S, Clarke L, Daenzer K, Kitous A, Labriet M, Lucena AFP, Octaviano C, Sbroiavacca ND (2016) Energy technology roll-out for climate change mitigation: a multi-model study for Latin America. Energy Econ 56:526–542