The relationship between renewable energy and retail electricity prices: Panel evidence from OECD countries

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Abstract

The centrality of electricity to everyday life is indisputable, and the price thereof can have significant implications. Previous literature is inconclusive over the effect of the renewable energy share in the electricity energy mix on retail electricity price as country-specific regulatory policy has a significant impact on retail electricity prices. The purpose of this paper is to determine the effect of the increasing renewable electricity share on retail electricity prices for 34-OECD countries, considering the change in market structure for 23 EU countries, due to data availability. The results show that the influence of the renewable energy share in the energy mix towards retail electricity prices is positive and statistically significant. This should not deter policymakers from promoting renewable energy as the effect is marginal and is expected to decline in coming years. The benefits of employing renewables far outweigh the environmental cost.

Keywords: Retail electricity prices; Renewable energy, Electricity market

1. Introduction

The centrality of electricity to everyday life is indisputable, and the price thereof can have significant implications. The European Commission (2016) states that while low electricity prices "raise purchasing power," and increases both living standards and industry competition, high electricity prices act as a signal to move to cleaner energy and improve energy efficiency. Studying the effect of increasing renewables on electricity prices is crucial in understanding market signals. The purpose of this study is to examine the effect of the continuously increasing share of Renewable Energy Sources (RES) in the energy mix on electricity prices in 34 OECD countries from 1997 to 2015, considering the change in electricity market structure. Our study extends on the research done by Moreno, López and García-Álvarez (2012) broadening the country group from European Union (EU) countries to include OECD countries as only EU countries fall under the emission trading scheme (EU- ETS). Extending the time period from 1997 to 2007 to include the most recent time period that was characterized by the financial crisis and its aftermath, allows us to view the subsequent constraints on investment as well as the decrease in the cost of renewable energy technologies.

Many IEA member countries embarked on the process of electricity market liberalization in the 1990s motivated by sectoral inefficiencies, the increasing trend of privatizing state-owned companies and the declining transitional cost towards a different system. Electricity market liberalization contributed to significant economic benefits, as competition increased efficiency within the sector, producing long-term consumer benefits. However, the system requires government involvement in upholding checks and balances. (IEA, 2005)

The European Union fully liberalized majority of their electricity markets in 2014, all member countries except Bulgaria and Malta are compliant. Market liberalization for numerous countries has been in effect for 5 to 10 years, excluding Cyprus who was the last to join the trend (European Commission, 2014). Wholesale electricity prices (referring to industry) decreased in Australia, Denmark, Finland, Norway, Sweden, the United Kingdom, and the United States after market liberalization. However, retail electricity prices (referring to households) have seen an increase, largely due to increasing fuel cost and cost associated with CO_2 emissions within Europe (IEA, 2005). Trujillo-Baute, del Río and Mir-Artigues (2018) attribute the sharp increase in retail elasticity prices over the years to the increase in renewable energy sources, while the European Commission (2016) states that wholesale electricity prices have decreased significantly although retail electricity prices tend to increase due to the “network price component, taxes, and levies". Transmission and distribution networks along with fuel cost are essential components of electricity prices, the former resulting from developments in economic regulation and not from increased competition in the electricity market. Fuel efficiency and ultimately energy efficiency has been a significant risk indicator for investors and determines the economic efficiency of a power plant (IEA, 2005). The International Renewable Energy Agency (IRENA, 2018) recently projected that within the next two years all renewable energy sources would be price competitive with fossils fuels. This new development is likely to increase the renewable energy share even further. The cost of renewable energy technologies decreased over the years and is expected to decrease further, which could potentially lower electricity prices in the future.

Full market liberalization reduces electricity prices by increasing competition; moreover, the European Union has committed to reducing emissions under the Kyoto protocol and therefore employed renewable energy support schemes to encourage the implementation of renewable technologies (IEA, 2001) (UNFCCC, Kyoto, 11 December 1997). This paper attempts to analyze the degree and magnitude of retail price change for 34 OECD countries from 1997 to 2015 considering the proportion of the share of electricity generated from renewable sources (RES-E) controlling for other variables including market structure.

Section 2 analyzes the current OECD electricity market and renewable policies, providing background to the issue at hand, followed by a detailed literature review in section 3. Section 4 discusses the panel data techniques employed to estimate the model. To do so, we will test for the presence of unit roots, if the results of the test conclude on a stationarity series, a pooled estimation will follow. Alternatively, Pedroni's (1999, 2004) panel cointegration test, which allows for "heterogenous short-run dynamics" will be pursued. We discuss the estimation results, followed by concluding remarks and a discussion in section 5.

2. OECD Electricity market and Renewable Policy

The Paris Agreement reached at the United Nations Framework Convention on Climate Change, Conference of the Parties (COP 21) in 2015 strengthens policies on climate change and incentives the transition towards a low-carbon energy system. The average CO_2 intensity of electricity needs to decrease by 96% from 2015 to 2050 to prevent global temperatures from increasing with 2°C (UNFCCC, 1997).

Pressures on CO_2 emissions motivated carbon-tax in several OECD member countries, including the European Union implemented in 2003, certain states of the United States, Canada in 2008, South Korea in 2015 and most recently China in 2016 (IEA, 2016). The IEA (2016) states that high carbon prices and increasing shares of RES can generate sufficient revenue to recover the fixed cost of low-carbon power sources, potentially increasing renewable investments. Wholesale electricity prices are decreasing due to the increased supply of low marginal-cost electricity generation, referred to as the merit order effect. Once the fixed cost of renewable technologies are covered, the marginal cost associated is low and consequently renewables places first in the merit order. The aftermath results in decreased revenue for all operators as well as overcapacity. For effective investment to happen, these issues have to be addressed for the transitional period towards low-carbon technologies. Support policies include fixed-price instruments, shared risk instruments, and subsidies.

Consumer decisions are made based on retail electricity prices, which have not seen the same decreases as wholesale electricity prices. Figure 1 depicts the correlation between average retail electricity prices and average renewable electricity share. Numerous renewable policies are financed by levies, increasing the consumer price. The electricity price to end-users comprises of energy – which is 43% of total price in Europe 2012-, distribution (30%), energy tax (13%) and value added tax (14%) (IEA, 2016). As a result of increased retail prices and decreasing cost of rooftop solar PV, household self-consumer or behind-the-meter electricity generation is increasing, ultimately resulting in an implicit feed-in-tariff. However, this does not eliminate the need for meter generation. Inefficiencies within the retail price, such as "tariff structure, taxation and the lack of time-varying options" give incorrect price signals to the consumer. Producers are encouraged to disclose the cost structure more openly depicting the actual value of the product to consumers. (IEA, 2016)

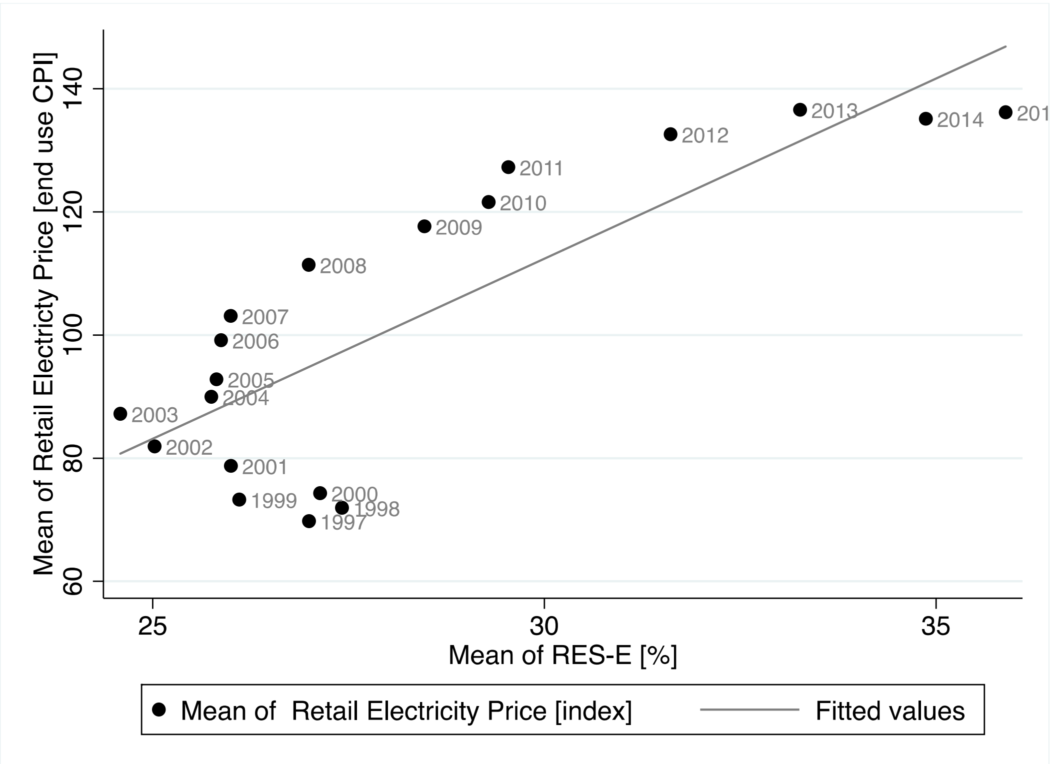

Figure 1: Correlation between average retail electricity prices and average renewable electricity share for all OECD countries 1997-2015 [Source: own calculations from IEA and World Bank (2018)]

3. Literature review

Previous literature is inconclusive over the effect of the renewable energy share in the energy mix on retail electricity price, as country-specific regulatory policy and market structure has a significant impact on retail electricity prices (Trujillo-Baute, del Río, and Mir-Artigues, 2018). Ballester and Furió (2015) found that in the Spanish case, an increasing share of renewable energy results in lower retail electricity prices for the period 2010 to 2013 if they consider "peak and off-peak prices" separately. The opposite is true for the period 2002 to 2009, due to less intense employment of renewable technologies and higher price associated with these technologies during this period. The weight of RES increased from 29% to 59% from 2008 to 2013, consisting of up to 80% of the daily supply on occasion since 2011. The authors employed a model adapted by Cartea and Figueroa (2005) which is a "stochastic process with mean reversion that includes a discrete jump process" which allows for price volatility to be captured. They arrived at the conclusion that the relationship between the RES-E share and electricity prices is only significant for peak prices.

Würtzburg, Labandeiria and Linares (2013) amongst others (Sensfuß, Ragwitz, and Genoese, 2008; Gelabert, Labandeira and Linares, 2011; Sáenz de Miera, del Río and Vizcaíno, 2008) studied the merit-order effect of the renewable energy share on wholesale electricity prices. Würtzburg et al. (2013) thoroughly reviewed previous literature, which allowed them to isolate trends and patterns. They found that the merit-order effect is much larger for smaller markets opposed to larger markets. Conducting an empirical analysis, they found that for each GWh of average hourly predicted renewable energy generation the day-ahead electricity price was reduced by 2% for the German-Austrian market. This price effect is not directedly transferred to the consumer retail price. The authors found that Germany's withdrawal from nuclear energy did not affect the merit order effect. They found weak evidence supporting string merit-order effects for high electricity demand, possibly due to the use of fossil fuels during these periods. Sańenz de Miera et al. (2008) states that this decrease in marginal cost to the producer may offset the initial setup cost and act as an incentive to invest in renewable technologies.

Iimura and Cross (2018) analyze the effects of renewable energy on household electricity prices in liberalized electricity markets in 7 OECD countries. The results indicate a "strong path dependency" for household electricity prices, while market reforms resulted in more significant price decreases than policy anticipated. There is no significant relationship between higher prices and increased renewable deployment. The authors suggest that renewables are more likely to be traded with neighboring countries than deployed by the host country due to the merit-order effect.

This study's approach aligned with the research conducted by Moreno, López and García-Álvarez (2012) who developed an econometric panel data model to estimate this relationship for the European Union from 1998 to 2009. They found that a 1% increase in renewable energy results in a 0.018% increase in retail electricity prices at a 1% level of significance. Noting that while the effect is small, it mostly influenced by “RES-E support schemes financed by the electricity market.” Public RES-E support schemes may effectively mitigate the retail price increase, which has been limited to what is deemed truly necessary (European Commission, 2015).

A 1% increase in Electricity Generation Concentration (EGC) results in a 0.005% decrease in retail electricity prices, this result is contradictory to the authors initial expectation. Moreno et al. (2012) note that market liberalization did not result in a less concentrated market structure, while observing that countries with less concentrated market structures do not necessarily have lower electricity prices.

A 1% increase in Energy Industries Emissions (EIE) results in a 0.025% increase in retail electricity prices while a 1% increase in GDP per capita results in a 1.345% increase. A 1% increase in Energy Dependency (ED) results in a 0.004% increase in retail electricity prices, Moreno et al. (2012) explain that the increase is due to the EU’s large energy dependency which is linked to international energy commodity prices. Noting that increased renewable deployment could result in lower levels of energy dependency. All the coefficients are statistically significant.

Our research builds on that of Moreno et al. (2012) to see if the degree and magnitude of the effect of renewable energy on household electricity prices are similar for 1997 to 2015 than it was for 1998 to 2009. Improvements in technology, increased competition and “a large base of experienced, internationally active project developers” have been singled out as key drivers in the cost reduction of renewable technologies (IRENA, 2018). Could the effect of the reduction of public RES-E support schemes on household electricity prices be offset by the decreased price of renewable technologies?

4. Methodology and Data

4.1 Theoretical Framework

This paper’s methodology is based on one of the models employed by Moreno, López and García-Álvarez (2012).

Our model is defined as:

$$\ln(y_{it}) = \alpha_i + \beta_1 \ln(RESE_{it}) + \beta_2 \ln(GDPPC_{it}) + \beta_3 \ln(EIE_{it}) + \beta_4 ED_{it} + \beta_5 \ln(EGC_{it}) + u_{it} \quad (1)$$

Where y is *household electricity prices* as the dependent variable along with RES-E, GDPPC, EIE, ED and EGC as the explanatory variables, defined below:

- *“Electricity generated from renewable sources* as a percentage of total gross electricity production (RES-E,")”, this variable will provide information on the share of renewable energy sources employed. Theory suggests that a positive relationship exist as public support schemes fund projects. It is important to note the impact of RES-E on electricity prices given of the reduction of public support schemes;

- *Gross Domestic Product per capita*, measured in constant 2010 US dollars (GDPPC), this will measure the relative economic activity of each country;
- *"Greenhouse gas emissions by the energy sector* as a percentage of total greenhouse gas emissions (EIE)", since the countries engage in an Emissions trading scheme, fluctuation in this variable has a direct effect on the marginal cost of energy production;
- *Energy Dependency* (ED), the degree of which the countries are dependent on natural resource importation connects the price of electricity to the price of these resources;
- *Market share of the largest generator in the electricity market* (EGC, Electricity Generation Concentration)", increased competition should have a significant role in the reduction of electricity prices.

4.2 Econometric methodology

The Im, Pesaran, and Shin (2003) panel unit root test is used to confirm stationary variables, the test allows for heterogeneous autoregressive coefficients. Apergis and Payne (2010), and Inglesi-Lotz (2016) explain that Im et al. (2003) uses "the average of the Augmented Dickey-Fuller unit root test while allowing for different orders of serial correlation". Which yields the following equation:

$$y_{it} = \rho_i Y_{i_{t-1}} + \sum_{j=1}^{p=1} \varphi_{ij} \varepsilon_{it_{-j}} + \delta X_{it} + u_{it} \qquad (2)$$

Where i=1, … , N for each country in the data span; t=1, … , T is the year; X_{it} represents the combined exogenous variables including two way fixed effects; ρ_i indicates the number of lags and ε_{it} the disturbance term.

To ensure robustness of the unit root test results, we also conducted the Levin, Lin and Chu (2002) t-statistics, the Breitung (2001) t-statistic, the Im, Pesaran and Shin (2003) W-statistic, the ADF Fisher Chi-square, and the Phillips-Perron (1988) Fisher Chi-square test (Maddala and Wu, 1999). If we reject the null hypothesis, therefore concluding on a stationary series, we will proceed with a pooled estimation. If not, we have to proceed with Pedroni's (1999, 2004) panel cointegration test which allows for cross-country interconnections.

$$y_{it} = \alpha_i + \delta_i t + \gamma_{1i}\ln(RESE_{it}) + \gamma_{2i}\ln(GDPPC_{it}) + \gamma_{3i}\ln(EIE_{it})$$

$$+\gamma_{4i}ED_{it} + \gamma_{5i}\ln(EGC_{it}) + \varepsilon_{it} \quad (3)$$

Where i = 1, … , N for each country in the data span ; t=1, … , T is the year; α_{it} and δ_i represents country-specific time invariant effects, and ε_{it} represents the estimated residuals deviations from the long-run value.

Pedroni (1999, 2004) make use of four statistics: panel ν, panel ρ, panel PP and panel ADF-statistic. Large positive values signify rejection in the panel ν statistic, where larger negative values indicate rejection in panel ρ, panel PP and panel ADF-statistic. "These statistics allow for heterogeneous fixed effects and deterministic trends and also for heterogenous short-run dynamics." Pedroni (1999, 2004) found that panels with T=20 the ADF group and ADF panel statistic, followed by the panel ρ statistic generally fair the best concerning power, size, and reliability, while the panel and group PP are somewhere between the ADF panel and panel ρ statistic.

If we reject of the null hypothesis of no cointegration with the Pedroni test, the Hausman (1978) test will be conducted to conclude whether a pooled or fixed effects estimation will follow. If the Hausman test concludes on a fixed effects model, we will proceed with a two-way fixed effects estimation, since we assume that each country and year has specific and unique time-invariant characteristics that influence the significant differences in household electricity prices amongst the 34-OECD countries. As such we control for the assumed correlation between the error term and our explanatory variables, this is denoted by α_i and is treated as a regression parameter (Wooldridge, 2015). We will conduct two separate models for the OECD and the EU for comparative purposes. As data on EGC is only available for EU countries, we want to see how our results differ when including only EU countries with RES-E, GDPPC, EIE and ED as explanatory variables.

4.3 Dataset

The data utilized in this evaluation were obtained from the International Energy Agency (EIA), the World Bank, OECD and Eurostat databases (the source of each indicator is described in Table 1, last column). Since data availability for wholesale electricity prices is restricted, only retail electricity prices will be examined form 1997-2015. Data availability for electricity price is not reported for the entire data span, and EGC is only available for EU countries (1997-2015), leading to an unbalanced panel dataset. The dependent variable is retail electricity prices in index form along with renewable energy share, electricity generation concentration, GDP per capita, energy industry emissions and energy dependence as explanatory variables. The variables expressed in natural logarithms can be expressed in terms of elasticities.

Table 1: Descriptive statistics of the variables

	Mean	SD	Min	Max	OBS	Sources
Retail electricity price [Real end use consumer price index]	105.9	75.3	3	638.7	623	IEA (2018)
RES-E [%]	28.2	27.5	0.04	100	665	World Bank (2018)
GDP per capita [CUS$]	36497.52	21479.79	5857.01	111968.3	665	World Bank (2018)
EIE [%]	27.8	13.8	0.06	73	636	OECD (2018)
ED[%]	18.6	130.5	-843.5	98.6	645	World Bank (2018)
EGC [%]	54.1	25.2	15.3	100	352	Eurostat (2018)

4 Empirical results

As discussed in the methodology section the unit root test proposed by Im, Pesaran, and Shin (2003) was done to determine the stationarity of the variables. In each case (except for ln(EGC) and ln(GDPPC) regarding trend and intercept) the null hypothesis was accepted, as such, each series in the panel dataset contains a unit root and is non-stationary . Given the results, we proceeded by testing for the existence of cointegration; results are represented in Table 2.

Each model has ln(Retail Price) as the dependent variable followed by the following explanatory variables, where model 1-4 looks at OECD countries and model (5a) and (5b) looks at only EU countries:

Model (1) - ln(RES-E)

Model (2) - ln(RES-E), ln(GDPPC)

Model (3) - ln(RES-E), ln(GDPPC), and ln(EIE)

Model (4) - ln(RES-E), ln(GDPPC), ln(EIE) and ED

Model (5a) - ln(RES-E), ln(GDPPC), ln(EIE) and ED

Model (5b) - ln(RES-E), ln(GDPPC), ln(EIE), ED, and ln(EGC)

Table 2: Panel cointegration test results

		Model 1		*Model 2*		*Model 3*	
		Panel	**Group**	**Panel**	**Group**	**Panel**	**Group**
Intercept and Trend	V-Statistic	8.553***		11.704***		12.709***	
	Rho-Statistic	1.005	3.216	2.357	4.845	2.320	5.273
	PP-Statistic	-2.040*	-0.639	-0.889	0.656	-4.374***	-3.199***
	ADF-Statistic	-6.410	-0.665	0.831	0.328	0.028	-0.957
Intercept	V-Statistic	-1.592		1.824**		-0.031	
	Rho-Statistic	0.489	1.101	-0.201	1.663	1.903	2.797
	Pp-Statistic	-1.719*	-3.901***	-3.431***	-2.802***	-1.457*	-4.034***
	ADF-Statistic	0.159	-1.155	-2.962***	-1.218	-0.461	-0.297
None	V-Statistic	-4.187		0.642		-0.654	
	Rho-Statistic	0.994	2.714	-0.014	2.174	-0.104	2.499
	PP-Statistic	-1.204	-1.835**	-2.244**	-1.989**	-4.114***	-3.759***
	ADF-Statistic	-0.472	-0.7255	-0.579	-0.548	-1.927**	-0.809
		Model 4		*Model 5a*		*Model 5b*	
		Panel	**Group**	**Panel**	**Group**	**Panel**	**Group**
Intercept and Trend	V-Statistic	9.797***		9.324***		-0.208	
	Rho-Statistic	3.149	6.294	2.831	5.755	4.329	6.412
	PP-Statistic	-6.741***	-4.110***	-6.508***	-3.94***	-3.587***	-1.168
	ADF-Statistic	-1.616*	0.734	-1.661*	0.679	0.011	1.124
Intercept	V-Statistic	1.262		1.156		-1.105	
	Rho-Statistic	2.207	3.794	2.061	3.447	2.570	4.339
	PP-Statistic	-2.856***	-3.958***	-2.528***	-3.248***	-3.213***	-5.114***
	ADF-Statistic	-2.609***	-0.406	-2.405***	-0.255	0.523	-0.389
None	V-Statistic	-0.784		-0.594		-1.391	
	Rho-Statistic	1.432	3.582	1.413	3.301	1.931	3.977
	PP-Statistic	-2.557***	-2.852***	-2.221***	-2.715***	-2.929***	-3.269***
	ADF-Statistic	1.100	-0.283	0.906	-0.381	-0.369	-0.475

Notes: *(**)[***] denotes 1% (5%) and [10%] levels of statistical significance → The null hypothesis of No cointegration is rejected

Inglesi-Lotz (2016) explains that these statistics "are based on the average values of the individual autoregressive coefficients linked with the unit root test for each country in the panel." Within all deterministic structures, the PP-statistics indicates the rejection of the null hypothesis of no cointegration at a 1% level of significance. As such, the results confirm a long-run relationship between retail electricity prices and electricity generated from renewable sources, including the control variables such as market share for the 22 EU countries. Given the results of the Hausman test, a two-way fixed effects estimation followed (Table 3), to account for heterogeneity between cross- and year-sections while seeing the effects of each control variable as our number of cross sections change significantly when controlling for market structure.

Model (1), (2) and (3) all have positive and statistically significant coefficients for RES-E, indicating that electricity generated from renewable sources does have a significant effect for the period examined on retail electricity prices when controlling for GDP per capita, and energy industry emissions. These results are in line with that of Moreno et al. (2012).

Table 3:Two-way fixed effects estimation

	(1) OECD	(2) OECD	(3) OECD	(4) OECD	(5a) EU	(5b) EU
ln(RES-E)	0.0535** (0.005)	0.0399** (0.033)	0.0376* (0.055)	0.0272 (0.163)	0.0443** (0.008)	0.0462** (0.005)
ln(GDPPC)		0.650*** (0.000)	0.726*** (0.000)	0.895*** (0.000)	0.383*** (0.000)	0.429*** (0.000)
ln(EIE)			-0.0751* (0.100)	-0.0334 (0.461)	0.114 (0.209)	0.157* (0.093)
ED				0.000261 (0.532)	-0.000712 (0.425)	-0.000220 (0.811)
ln(EGC)						-0.0912** (0.049)
constant	3.914*** (0.000)	-2.633** (0.014)	-3.174** (0.004)	-5.029*** (0.000)	-0.194 (0.858)	-0.466 (0.669)
N	623	623	601	581	349	349
adj. R*2*	0.578	0.604	0.606	0.631	0.760	0.762
BIC	-127.2	-160.5	-156.1	-165.1	-364.6	-363.0
F	47.61*** (0.000)	50.00*** (0.000)	46.49*** (0.000)	47.39*** (0.000)	154.5*** (0.000)	55.19*** (0.000)
Hausman test	4.643*	15.563***	14.952***	17.998***	30.67***	32.358***

Notes: *(**)[***] denotes 1% (5%) and [10%] levels of statistical significance

Once we include energy dependence in model (4), the coefficient for RES-E remains positive but is not statistically significant in the 34 OECD countries. Most OECD countries are net energy importers, with the exception of Australia, Canada, Denmark, Mexico and New Zealand. On average OECD members are less energy dependent that EU members, with an energy dependence rate of 21% and 53.9% in 2015 respectively, this difference could explain why the coefficient for RES-E becomes insignificant with the inclusion of ED for OECD countries. Further investigation into the sources of energy imports are required to make proper inference on this relationship. Figure 2 indicates energy dependency for the year 1997 and 2015 respectively, where positive values indicate net energy importers and negative values, net energy exporters.

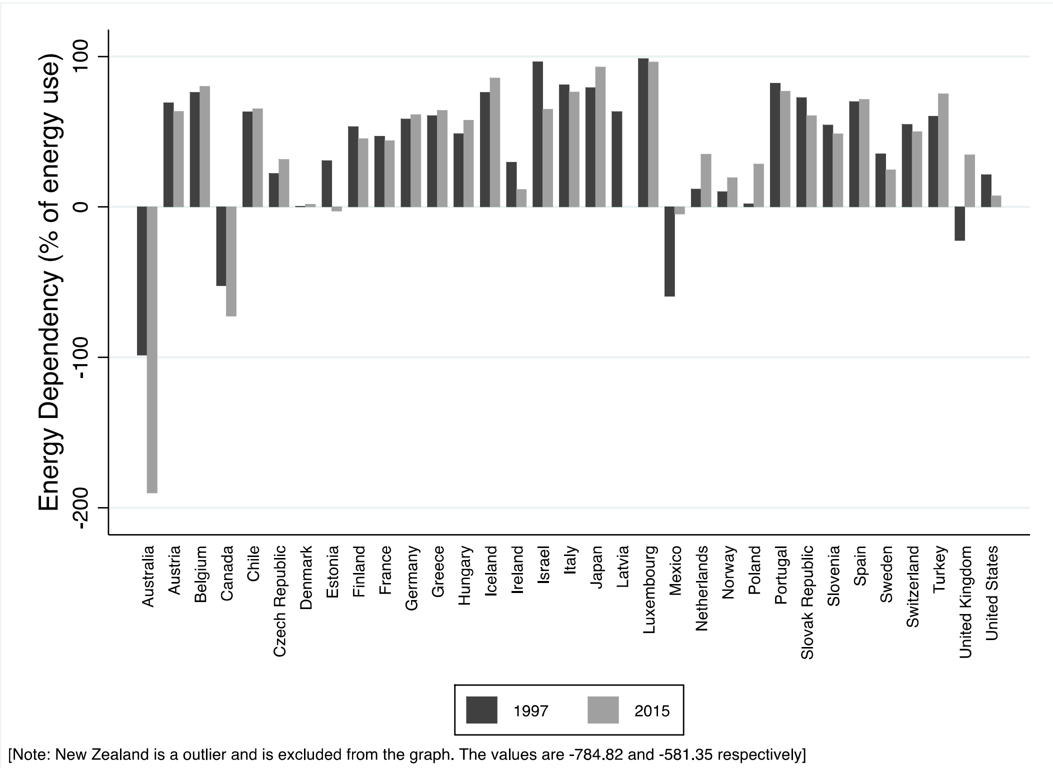

Figure 2: Energy Dependency for all OECD countries for the year 1997 and 2015 respectively [Source: World Bank 2018]

Staying true to Moreno et al. (2012), the results indicate the need to control for electricity generation market share. As EGC is only available for EU members reducing our number of cross sections from 34 to 23, we need to consider a replication of model (4) for only EU countries, represented by model (5a). A 1% increase in the share of RES-E results in a 0.044% increase in the share of retail electricity prices. While GDP per capita remains positive and statistically significant, EIE and ED remain statistically insignificant. Model (5b) controls for market structure, where we see that all variables are statistically significant, except for ED. A 1% increase in the share of RES-E results in a 0.046% increase in the share of retail electricity prices, the coefficient is slightly large than Moreno et al. (2012). A 1% increase in GDP per capita leads to a 0.429% increase in retail electricity prices, while a 1% increase in EIE leads to 0.157% increase. The effect of EIE is much larger than in Moreno et al. (2012), indicating that the effect of emission trading schemes increased from 2007 to 2015. EIE does not have a significant effect in Model (4) which contains the OECD countries of which not everyone has an emissions trading scheme in contrast to Model (5a) which contains only EU countries all employing a emissions trading scheme, where EIE is still statistically insignificant. This result indicates that controlling for market share (EGC) is a necessary condition to see the effects of the emissions trading scheme.

ED has a negative sign, indicating that a 1% increase in energy dependency could result in a decrease in retail electricity prices, this could be a result of the low-cost fossil fuel energy imports, but the coefficient is not statistically significant. A 1% increase in EGC leads to a 0.091% decrease in retail electricity prices, which is contradictory to theory indicating that increased competition leads to decreases in prices, but in line with the findings of Moreno et al. (2012) as they explained that countries with higher market concentration have more government subsidies decreasing electricity prices. All the results except for ED are in line with that of Moreno et al. (2012).

5. Conclusion and Discussion

The purpose of this paper is to determine the effect of the increasing renewable electricity share on retail electricity prices for 34-OECD countries, considering the change in market structure for 23 EU countries in a panel data framework from 1997 to 2015. The study confirmed the existence of a long-run relationship between retail electricity price, the share of renewable electricity formation, GDP per capita, greenhouse gas emissions by the energy sector, energy dependence and electricity market concentration.

The two-way fixed effects estimation results are as follows; a 1% increase in the share of renewable energy leads to a 0.046% increase in retail electricity prices; while a 1% increase in the share of GDP per capita leads to a 0.429% increase; a 1 % increase in share of Emissions by the Energy Industry leads to a 0.157% increase in retail electricity prices; a 1% increase in energy dependency leads to a decrease in retail electricity prices, but is not significant; 1% increase in electricity market concentration leads to a 0.091% decrease in retail electricity prices. The current increase of RES-E on electricity prices is marginal and is largely due "RES-E support schemes financed by the electricity market" (Moreno et al., 2012).

The results hold important implications for future policies encouraging renewable energy sources and understanding price signals as a consumer. Strengthening renewable electricity generation could shield against threats to electricity security, as well as providing efficient and affordable access to electricity. As such renewables requires a cost-effective match between demand and supply, encouraging investment while phasing out support schemes in the long-run (Tuerk & Fazekas, 2016). The electricity market and regulatory framework need to be adjusted to accommodate a "higher degree of flexibility from the network to compensate for renewable intermittency" (IAE, 2016). Increases in electricity prices towards the consumer should act as signal to move towards cleaner energies, as the marginal cost associated is low while the environmental benefit is high.

The current increase of RES-E on electricity prices is marginal and is largely due "RES-E support schemes financed by the electricity market" (Moreno et al., 2012). IRENA (2018) projected that renewable energy sources would be price competitive with fossil fuels within the next two years, we suspect that with future data the relationship will eventually be negative. Encouraging private RES-E support schemes could effectively mitigating the increases in retail electricity prices bringing about this relationship sooner. Emissions trading schemes by the energy industries only hold a significant effect for EU countries. Most countries' energy dependency changed over the period, declining in both energy exports and imports (Dedeoğlu and Kaya, 2013) and holds no significant effects for retail electricity prices in this analysis. Our findings regarding RES-E, GDP per capita, Energy Industry Emissions and Electricity Generation Concentration aligned with the findings of Moreno et al. (2012). However, our coefficient for Energy Dependency is negative and statistically insignificant where that of Moreno et al. (2012) was positive and significant. As mentioned before this could be the result of low cost fossil fuel energy imports.

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