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SHORT RUN AND LONG RUN DYNAMICS OF RESIDENTIAL ELECTRICITY CONSUMPTION: HOMOGENEOUS AND HETEROGENEOUS PANEL ESTIMATIONS FOR OECD

Abstract: *The purpose of this paper is to reveal the short run and long run* dynamics of *residential electricity consumption for 11 OECD countries within annual period 1979-2006. To this end, this paper first explores the findings from related literature evidence and, later, follows panel cointegration equations (CEs) and panel error correction models (ECMs). CEs give long run relations of the variables in residential electricity demand function. ECMs include both long run and short run parameter estimates of the per capita residential electricity demand in terms of residential electricity price, residential light fuel oil price, residential natural gas price and per capita income. For both ECs and ECMs, the techniques of panel OLS, panel adjusted OLS and panel dynamic OLS are utilized. Finally, this paper yields short term and long term elasticities of residential electricity consumption together with error correction terms through homogeneous and heterogeneous variance structures.*

Keywords: *electricity consumption, elasticities, homogeneous and heterogeneous variance structures, panel error correction model, panel dynamic ordinary least squares.*

JEL Classification: C33, C51, D12, Q43

1. Introduction

Throughout energy literature, the response of energy consumption to the explanatory variables of its own price, prices of other energy sources, income and other determinants has been great interest to researchers and policy makers. In this paper, specifically, it is aimed at observing the behavior of residential electricity consumption. Accordingly, after analyzing the related articles, paper yields its empirical findings for the determinants of residential electricity consumption in short and long terms.

Inglesi (2010) employs cointegration (CE) and error correction model (ECM) with annual South African data for the period 1980-2005 to reveal the long run and short run parameters of electricity demand in South Africa and finds income elasticity of 0.415 and price elasticity of - 0.564 in the long run and income elasticity of 0.820 and population elasticity of 3.467 in the short run. Mitchell (2006) follows CE, ECM and vector error correction model (VECM) by employing 1960-2005 annual data for Barbados and concludes that long run petroleum energy demand's price and income elasticities of ECM are -0.166 and 0.575, respectively, while those of VECM are -0.289 and 0.430, respectively. As for the short term, he finds income elasticity of 0.357 by ECM and insignificant relation between price and demand by ECM and VECM analyses. Nakajima and Hamori (2010), using 1993-2008 quarterly panel data for the US states, apply panel CE to explore long run elasticities of residential electricity demand in the USA. They reach income elasticities ranging from 0.33 to 1.00, price elasticities ranging from -0.12 to -0.34, a unit effect of heating degree days on electricity consumption varying from 0.14 to 0.22 and that of cooling degree days varying between 0.54 and 0.82. Al-Iriani (2005) observes monthly data spanning from 1981 to 2000 for United Arab Emirates (UAE) and analyses demand for electricity demand with respect to climate conditions in UAE. He follows Autoregressive Integrated Moving Average (ARIMA) model to estimate his electricity consumption model employing cooling degrees and time trend as independent variables and obtains climate elasticity of 0.027. Filippini and Pachauri (2004), using survey data for 30.000 households obtained in 1993-1994 in India, investigate electricity consumption in winter, monsoon and summer seasons. Following their cross sectional model of electricity consumption in India, they conclude that electricity is price and income inelastic and that parameters such as household size, the region where the households live and the age of households have significant effects on electricity consumption. Narayan et al. (2007) conduct panel cointegration analyses to explore long run and short run price and income elasticities of per capita residential electricity consumption of G7 countries for the period 1987-2003 and, by considering the results of panel DOLS results, they reveal that income elasticity, own price elasticity and cross price (natural gas) elasticity are 0.245, -1.563 and 2.965, respectively in the long run. In the short run, they find insignificant parameters of income and natural gas price on electricity consumption while significant price elasticity value of -0.107. Sa'ad (2009) investigates electricity consumption in South Korea with the structural time series analyses over the period from 1973 to 2007. He takes per capita electricity consumption as a function of price, per capita income, structural and life style factors and stocks of appliances used by households. His model states that, in the long run, per capita electricity consumption has income and price elasticities of 2.35 and -0.50, respectively. Dergiades and Tsoulfidis (2008) carry out time series ARDL cointegration technique to estimate per capita residential electricity demand function in USA by observing annual data spanning from 1965 to 2006. Their model estimates long run elasticities of per capita residential electricity consumption regarding per capita income, residential electricity price, cooling and heating degrees-days index, average price of oil for heating purposes and per capita owned electrical appliances as 0.272, -1.065, 0.726, 0.199 and 1.543, respectively. The short term estimations, on the other hand, indicate that short run elasticities of per capita residential electricity consumption in terms of per capita income, residential electricity price, cooling and heating degrees-days index, oil price, lagged oil price and per capita stock of appliances are 0.101, -0.386, 0.263, 0.014, -0.049 and 0.560, respectively. Hondroyiannis (2004) employs monthly data from 1986 to 1999 for Greece to obtain long run and short run elasticities of residential demand for electricity in Greece, and, through his VECM results, concludes that long run income, price and temperature elasticities of residential electricity consumption are 1.56, -0.41 and -0.19, respectively, while, in the short run, only income elasticity appears to be significant with the value of 0.20. Athukorala and Wilson (2009) carry out CE and ECM analyses for the annual period over 1960-2007 to examine the per capita electricity consumption in Sri-Lanka. They figure out that long run per capita electricity consumption has the elasticities of per capita income (0.785), own price (-0.616), price of kerosene oil (0.142) and that short run per capita electricity consumption elasticities for the same variables are 0.32, -0.16 and 0.10, respectively.

This paper specifically focuses on two possible variance structures of the estimations to overcome the possible less appropriate statistical properties that might arise when the variance structure of panel data is explicitly assumed solely either homogeneous or heterogeneous. Instead, when the cross sections' population variances are not known, a comparison of these two types of variance structures in the analyses would give researcher more confident statistical output of the panel estimations through the models of ordinary least squares and dynamic ordinary least squares. To this end, this paper aims at revealing statistically significant parameter estimations of residential demand function for electricity consumption including its own price, price of a substitute energy source, price of a complementary energy source and income by employing the related data of OECD and IEA as described in details in Section 3. After evaluating the literature findings and aim of this paper, Section 2 gives the details of panel data analyses carried out and Section 3 remarks on parameter estimations of residential electricity consumption for OECD countries.

2. Econometric Methodology

Let time series $\{y_{i,t}\}$ be integrated of order one $I(1)$, as $i = 1,2,..,N$ and $t = 1,2,..,T$. In this case $y_{i,t}$ is random walk but $u_{i,t}$ from Eq. (1) is stationary.

$$(1-L)y_{i,t} = (y_{i,t} - y_{i,t-1}) = \delta_i y_{i,t-1} + v_{i,t}'\psi_i + u_{i,t} \quad (1)$$

where L is lag operator, $\delta_i = (\rho_i - 1)$ and $v_{i,t}$ denote the optional variables of individual constants and/or trends. If $|p_i| \geq 1$, y_{it} has the variance increasing through time and thus, it is not stationary. The power of L or integration order I(d) determines the number of unit roots, hence y_{it}, as is defined above, has one unit root. The homogeneous unit root null hypothesis (or common unit root null hypothesis or common AR null hypothesis) is H_0: $p_i = p = 1$ for all *i,* (or equivalently H_0: $\delta_i = \delta = 0$, for all i), as alternative hypothesis is H_A: $p_i = p < 1$ for all *i,* (or equivalently H_A: $\delta_i = \delta < 0$, for all i). The heterogeneous unit root null hypothesis (or individual unit root null hypothesis or individual AR null hypothesis) is H_0: $p_i = 1$ for all *i,* (or equivalently H_0: $\delta_i = 0$, for all i), whereas alternative hypothesis is H_A: $p_i < 1$ for all *i,* (or equivalently H_A: $\delta_i < 0$, for all i). In homogeneous tests, therefore, all δ_i are identical across members and thus individual effects are not considered. And, in heterogeneous tests, all δ_i are not identical across members and, hence, individual effects are specifically observed. When a variable is found $I(1)$, it implies that, although time series follows random walk in level, it is difference stationary. In this study both homogeneous tests of Breitung (2000) and heterogeneous tests of Im, Pesaran and Shin, IPS, (2003) are run to test the unit root null hypotheses against the alternative hypotheses. In case of obtaining different order of integration, for instance, under homogeneous variance structure (when some variables are found $I(1)$ while others are found $I(0)$), the homogeneous unit root test of Hadri (2000) will be performed, as well. It should be noted here that Hadri's null hypothesis is H_0: $p_i = p < 1$ against alternative one that H_A: $p_i = p = 1$. By the same token, if IPS (2003) tests yield mixed results, another heterogeneous unit root test will be launched.

In applied economics, the regression model may employ $I(1)$ variables provided that they are cointegrated. If they are not cointegrated, the parameter estimates and corresponding test statistics would be biased and inconsistent. This paper follows cointegration equation to be tested by considering Eq. (2).

$$ec_{i,t} = \alpha_i + \emptyset_i ep_{i,t} + \partial_i fp_{i,t} + \theta_i\, ngp_{i,t} + \vartheta_i y_{i,t} + \varepsilon_{i,t} \quad (2)$$

where $ec_{i,t}$, α_i, $ep_{i,t}$, $fp_{i,t}$, $ngp_{i,t}$, $y_{i,t}$, $\varepsilon_{i,t}$ denote electricity consumption, constant, electricity price, light fuel oil price, natural gas price, income and residuals from panel regression for individual *i* at time *t*, respectively. The cointegration test is implemented by Eq. (3).

$$\varepsilon_{it} = \gamma_i \varepsilon_{it-1} + e_{it} \quad (3)$$

The null hypothesis of no cointegration is H_0: $\gamma_i = 1$ for all *i*. The alternative hypothesis is H_A: $\gamma_i = \gamma < 1$ for all *i* under homogeneous variance structure (within dimension) and alternative hypothesis is H_A: $\gamma_i < 1$ for all *i* under heterogeneous variance structure (between dimension). This paper carries out both within dimension tests (panel tests) by Kao (1999) and Pedroni (1995) and between dimension tests (group tests) by Pedroni (1999, 2004).

The rejection the null of no cointegration states the evidence of long run relation between variables. The parameters $\emptyset_i, \partial_i, \theta_i$ and ϑ_i from Eq. (2), then, become long run parameters. The short run estimations are obtained through error correction model (ECM) as depicted by Eq. (4).

$$\Delta ec_{i,t} = \sigma_i + \pi_i \left[ec_{i,t-1} - \alpha_i - \emptyset_i ep_{i,t-1} - \partial_i fp_{i,t-1} - \theta_i\, ngp_{i,t-1} - \vartheta_i y_{i,t-1}\right]$$
$$+ a_{1i}\Delta ep_{i,t} + a_{2i}\Delta fp_{i,t} + a_{3i}\Delta ngp_{i,t} + a_{4i}\Delta y_{i,t} + \omega_{i,t} \quad (4)$$

where $\Delta ec_{i,t}$, $\Delta ep_{i,t}$, $\Delta fp_{i,t}$, $\Delta ngp_{i,t}$ and $\Delta y_{i,t}$ represent the first differences of $ec_{i,t}$, $ep_{i,t}$, $fp_{i,t}$, $ngp_{i,t}$ and $y_{i,t}$, respectively. The term in brackets is equal to error correction term at time *t*-1 (ECT_{t-1}) which corresponds to $\varepsilon_{i,t-1}$. Thus, ECT_{t-1} is deviation from long run equilibrium at time *t*-1. The constant, adjustment parameter, short run parameters and residuals in Eq. (4) are denoted by σ_i, π_i, a_{ji} and $\omega_{i,t}$, respectively for all *i*, as *j* = 1, 2, 3, 4. and *t* = 1, 2, 3,…,T. The discrepancy from long run equilibrium of Eq. (2) at time *t*-1 disappears by the estimation value of adjustment parameter, $\hat{\pi}_i$, each time *t* period for all *i*. Therefore, the ECM given by Eq. (4) employs both long run and short run dynamics of demand function for electricity; $ec = f(ep, fp, ngp, y)$.

Finally, the next econometrical issue is to choose the estimation model. Throughout empirical panel studies, Ordinary Least Squares (OLS), Dynamic Ordinary Least Squares (DOLS) and Fully Modified Ordinary Least Squares (FMOLS) are used to obtain the panel regression coefficients. On the other hand, according to Monte Carlo simulations of Kao and Chiang (2000), FMOLS estimations for the panel consisting of less than 20 cross sections and less than 60 time series observations provide serious biasedness and, in this case, DOLS estimations are superior to FMOLS. To this end, in this paper, DOLS model is followed to reach the conclusion of the panel estimations. For the purpose of comparison, panel OLS and panel adjusted OLS are also analyzed together with panel DOLS. Reconsidering the Eq. (2), the panel fixed effect model in matrix form can be written as below.

$$ec_{i,t} = \alpha_i + x'_{i,t}B + \varepsilon_{i,t},\ i = 1,2,..,N \text{ and } t = 1,2,..,T. \quad (5)$$

where $ec_{i,t}$ is n×1 vector, α_i is n×1 vector of constants, $x_{i,t}$ is k×n matrix of explanatory variables, $x'_{i,t}$ is transpose of matrix x, B is k×1 vector of slope parameters and $\varepsilon_{i,t}$ is n×1 vector of residuals. Eq. (5) can be modified to obtain DOLS regression by adding differenced leads and lags to correct the serial correlation and endogeneity of $\varepsilon_{i,t}$ from OLS as is shown below.

$$ec_{i,t} = \alpha_i + x'_{i,t}B + \sum_{j=-k}^{k} \mu_{i,k}\Delta x_{i,t+k} + n_{i,t} \quad (6)$$

In the existence of cointegration relationship, both $\{ec_{i,t}\}$ and $\{x_{i,t}\}$ are $I(1)$ and long run parameters of OLS and DOLS are obtained by following Kao and Chiang (2000) and Pedroni (2001).

$$\hat{B}_{OLS} = \left(\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{i,t} - \bar{x}_i)(x_{i,t} - \bar{x}_i)'\right)^{-1}\left(\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{i,t} - \bar{x}_i)(ec_{i,t} - \overline{ec}_i)\right) \quad (7)$$

$$\hat{B}_{DOLS} = N^{-1}\sum_{i=1}^{N}\left(\sum_{t=1}^{T} z_{it} z'_{it}\right)^{-1}\left(\sum_{t=1}^{T} z_{i,t}\left(ec_{i,t} - \overline{ec}_i\right)\right) \quad (8)$$

where $z_{i,t}$ represents $2(K+1) \times 1$ vector of explanatory variables including $(x_{i,t} - \bar{x}_i,\ \Delta x_{i,t-K}, \ldots, \Delta x_{i,t+K})$.

3. Data and Estimation Results

In this work, the 11 OECD countries; Austria, Finland, France, Ireland, Japan, Luxembourg, the Netherlands, Spain, Switzerland, UK and the USA are observed as cross sections in panel data spanning annually from 1979 to 2006. The natural log of per capita residential electricity consumption in thousand KWh units (*ec*) is taken to be a function of natural log of residential electricity price (*ep*), natural log of residential light fuel oil price (*fp*), natural log of residential natural gas price (*ngp*) and natural log of per capita income (*y*). The data for *ep*, *fp* and *ngp* come from IEA, Energy End-Use Prices (USD/unit) guided by 2Q2008 Documentation of IEA, Energy prices and Taxes: Beyond 2020 Edition. The data for *ec* is obtained from OECD International Energy Agency (IEA) and OECD Population and Employment. The variable *y* (USD/unit) is provided by Heston et al., (2009). The empirical results of panel unit root tests, panel cointegration tests and panel long run and short run analyses are given by Table 1, Table 2 and Table 3, respectively.

In Table 1 and Table 2, homogeneous (common) AR and heterogeneous (individual) AR structures correspond to within dimension estimation (panel) and between dimension estimation (group), respectively. In Table 1, all unit root tests include constant and trend terms and the probabilities are shown in parentheses. The differenced terms *Δec, Δep, Δfp, Δngp* and *Δy*, represent the first differences of *ec, ep, fp, ngp* and *y*.

Table 1: Panel Stationarity Tests for Panel Data over the Period 1979-2006

Homogeneous AR	***ec***	***ep***	***fp***	***ngp***	***y***
Breitung t (2000)	0.997	-1.429	2.477	0.88000	-2.231
	(0.840)	(0.076)	(0.993)	(0.810)	(0.012)
Hadri Z (2000)	6.774	5.065	7.427	5.728	6.996
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Heteregeneous AR	***ec***	***ep***	***fp***	***ngp***	***y***
IPS W (2003)	-2.006	0.205	3.672	1.666	-0.311
	(0.022)	(0.581)	(0.999)	(0.952)	(0.377)
Homogeneous AR	***Δec***	***Δep***	***Δfp***	***Δngp***	***Δy***
Breitung t (2000)	-8.015	-4.481	-2.935	-3.289	-2.037
	(0.000)	(0.000)	(0.000)	(0.000)	(0.020)
Heteregeneous AR	***Δec***	***Δep***	***Δfp***	***Δngp***	***Δy***
IPS W (2003)	-14.283	-8.296	-12.190	-10.300	-4.330
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

Breitung (2000) *t* test statistics reveal that *ec*, *fp* and *ngp* are nonstationarity, whereas *ep* and *y* are stationary at 10% and 5% significances, respectively. Hadri (2000) Z test statistics, on the other hand, find all variables non stationary at 1% significance level. Hence, the null hypothesis of common unit root is accepted for all panel data. IPS (2003) Wald test statistics also fail to reject the null hypothesis of individual unit root for all variables except *ec*. Considering the differences of the series, Breitung (2000) and IPS (2003) tests indicate that all variables are *I*(1) at 1% level of significance, with the exemption that Breitung (2000) test statistic for Δy is significant at 5% level.

Table 2 homogeneous (common AR) panel cointegration test results give the evidence of cointegration relation between variables. All Kao (1999) and Pedroni (1995) test statistics confirm this finding at 1% level of significance, while null hypothesis of no cointegration is rejected by DF *t* rho (Kao, 1999) at 5% level and not accepted by DF *t* rho star and ADF (Kao, 1999) at 10% level of significance.

Table 2: Panel Cointegration Tests for Panel Data over the Period 1979-2006[a]

Homogeneous AR	**Statistic**	**Probability**
DF rho (Kao, 1999)	-4.457	0.000
DF *t* rho (Kao, 1999)	-1.918	0.027
DF rho star (Kao, 1999)	-9.216	0.000
DF *t* rho star (Kao, 1999)	-1.474	0.070
ADF (Kao, 1999)	-1.598	0.055
t rho NT (Pedroni, 1995)	-153.602	0.000
TN1 rho (Pedroni, 1995)	-12.877	0.000
TN2 rho (Pedroni, 1995)	-12.857	0.000
Heteregeneous AR	**Statistic**	**Probability**
Group rho (Pedroni, 1999, 2004)	1.570	0.941
Group PP (Pedroni, 1999, 2004)	-3.042	0.001
Group ADF (Pedroni, 1999, 2004)	-4.221	0.000

[a] Gauss code of Kao and Chiang (2000), NPT 1.3 Program written by Chiang and Kao (2002) and Pedroni's RATS code (1999) are carried out to estimate the cointegration test statistics.

The overall result is in favor of cointegration with common AR. The heterogeneous (individual AR) cointegration tests by Pedroni (1999, 2004) disclose the result of cointegration except Group rho test statistics of Pedroni (1999, 2004). Hence, both common unit root and individual unit root null hypotheses result in existence of long run relation between *ec*, *ep*, *fp*, *ngp* and *y*.

Table 3 shows long run and short run parameter estimates of per capita residential electricity consumption function. ECT_{-1} denotes error correction term at time t_{-1}. The long run and short run estimations by OLS and adjusted OLS are similar to those of homogeneous DOLS. The parameter estimations of homogeneous DOLS, and heterogeneous DOLS, on the hand, seem to be close each other but differ in terms of significance of natural gas in the long run.

Table 3: Long Run and Short Run Dynamics of Demand for Residential Electricity for Panel Data over the Period 1979-2006[a]

	OLS[b]	OLS[c]	DOLS[d]	DOLS[e]
Long run				
ep	- 0.648 (0.000)	-0.665 (0.000)	-0.740 (0.000)	-0.664 (0.000)
fp	0.129 (0.035)	0.130 (0.001)	0.156 (0.000)	0.116 (0.000)
ngp	ns[f]	ns	ns	-0.033 (0.000)
y	0.611 (0.000)	0.617 (0.000)	0.615 (0.000)	0.576 (0.000)
Short run				
Δep	ns	ns	ns	ns
Δfp	ns	ns	ns	ns
Δngp	-0.040 (0.064)	-0.040 (0.064)	-0.040 (0.064)	-0.040 (0.065)
Δy	ns	ns	ns	ns
ECT_{-1}	-0.011 (0.086)	-0.011 (0.088)	-0.011 (0.076)	-0.010 (0.097)

[a] Gauss code of Kao and Chiang (2000) and NPT 1.3 Program written by Chiang and Kao (2002) with some modifications are run to estimate the parameters.
[b] Conventional OLS estimators under homogeneous covariance structure and probabilities of *t* test statistics in parentheses.
[c] Bias corrected OLS estimators under homogeneous covariance structure and probabilities of adjusted *t* test statistics in parentheses.
[d] DOLS estimators with one lead and two lags under homogeneous covariance structure and probabilities of *t* test statistics in parentheses.
[e] DOLS estimators with one lead and two lags under heterogeneous covariance structure and probabilities of *t* test statistics in parentheses.
[f] (ns) indicates not significant parameter.

As DOLS estimation yields 5% significance for *fp*, all other long run estimations are found significant at 1% level. Homogeneous long run estimates of DOLS have correct signs and bear less than unit elasticities. As the coefficient of *ngp* is found insignificant, the panel variables of *ep*, *fp* and *y* receive the elasticities of -0.740, 0.156 and 0.615, respectively. Heterogeneous DOLS estimates find all long run coefficients are significant and slightly differ from homogeneous ones. The per capita residential electricity consumption falls by 0.66% in return for 1% increase in residential electricity price, rises by 0.12% as residential fuel oil price increases by 1%, decreases by 0.03% when natural gas price goes up by 1%, and finally, increases by 0.58% if there is 1% increase in per capita income, as other things given. In terms of income elasticity of 0.58, the residential electricity consumption is considered a necessity good. A necessity good brings in lower price elasticity of consumption. Although cross elasticities are very low, one may imply that the residential fuel oil is a substitute of residential electricity and residential natural gas is a complementary energy source of residential electricity consumption.

The short run estimations indicate that the per capita residential electricity consumption is not sensitive at all to electricity price, fuel oil price and income but slightly sensitive to natural gas price in the short run. The per capita residential electricity consumption decreases only 0.04% as natural gas price increases 1% in the short term.

ECT_{-1} is the deviation from log run equilibrium in previous period. The coefficient value of ECT_{-1} shows the speed of adjustment to reach the cointegration equilibrium again at current period. The coefficient values of ECT_{-1} from homogeneous DOLS and heterogeneous DOLS are -0.011 and -0.010, respectively. Both adjustment values have the statistical significances at 10% levels. The mean, maximum and minimum values of ECT_{-1} from homogeneous DOLS are -0.0003, 0.6907 and -0.8517, respectively. The mean, maximum and minimum values of ECT_{-1} from heterogeneous DOLS are -0.00158, 0.6861 and -0.8181, respectively.

The mean, maximum and minimum values of ECT_{-1} from homogeneous DOLS are -0.0003, 0.6907 and -0.8517, respectively. The mean, maximum and minimum values of ECT_{-1} from heterogeneous DOLS are -0.00158, 0.6861 and -0.8181, respectively. Let one observe, by coincidence, the maximum and minimum values of ECT_{-1} obtained from heterogeneous DOLS among panel ECT_{-1} of 297 observations. The minimum value, -0.8181, is ECT_{-1} of Spain in 1980 and the maximum value, 0.6861, belongs to ECT_{-1} of Finland in 1997. This means that per capita residential electricity consumption is above its long run equilibrium by 0.6861 units in 1997 and the next year, in 1998, this electricity consumption will decrease by 0.006861 units, [(0.6861) × (-0.010) = -0.006861] to restore the long run equilibrium. The per capita residential electricity consumption, on the other hand, is below its long run equilibrium by -0.8181 units in 1980 and this residential consumption will increase by 0,008181 units in 1981 to correct the error from cointegration equilibrium. Eventually, one may note that, according to homogeneous and heterogeneous DOLS estimations, per capita residential electricity consumption in OECD, together with its deviations, keeps long run equilibrium through 1979-2006 with 297 observations.

4. Summary and Conclusion

This paper explores the long run and short run dynamics of per capita residential electricity consumption by observing the panel data for 11 OECD countries from 1979 to 2006. This work, in which residential electricity price, residential light fuel oil price, residential natural gas price and per capita income are employed as explanatory variables, eventually aims at finding the long run and short run estimates of the electricity consumption. For this purpose, homogeneous and heterogeneous analyses of panel unit root tests, panel cointegration tests and panel error correction models are conducted within given period and cross sections.

The homogeneous findings report that the long run elasticities of electricity consumption's own price, fuel oil price and income are -0.740, 0.156 and 0.615, respectively. The heterogeneous DOLS estimations result that own price elasticity, cross price elasticity with respect to fuel oil price, cross price elasticity with respect to natural gas price, and income are -0.664, 0.116, -0.033 and 0.576, respectively. On the other hand, in the short run, electricity consumption does not respond to changes in electricity price, fuel oil price and income as it reacts slightly to the changes in natural gas prices. All elasticities are found less than unity (inelastic) with expected signs. This has several outcomes. First, the inelastic demand for residential electricity implies that increase in residential electricity price results in higher expenditures on residential electricity. Second, residential fuel oil is a weak substitute for residential electricity and residential natural gas a weak complementary energy source of residential electricity. Third, residential electricity is a necessity good. Finally, in the study, the error correction terms are found significant. This significance reveals that cointegration equilibrium of per capita residential electricity consumption with regard to residential electricity price, residential light fuel oil price, residential natural gas price and per capita income is restored after some deviations from log run.

This paper may suggest that, throughout the literature of energy demand functions, one may consider also the impact of stock prices on oil prices, as in Lai et al. (2011), and on other energy prices. Or, alternatively, one may follow Aldea and Ciobanu (2011) to run nonparametric techniques through bootstrapping analyses to understand the electricity market. One may also carry out alternative electricity demand work in which different substitute(s) and complementary energy source(s) are taken into consideration. Or, a researcher may observe demographical, geographical and seasonal dummies as additional exogenous variables to understand the behavior of electricity consumption. Finally, another possible electricity consumption model might be proposed to seek for specifically the behaviors of commercial electricity and industrial electricity consumptions.

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