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A PANEL DATA ANALYSIS OF ELECTRIC CONSUMPTIONS IN THE RESIDENTIAL SECTOR

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A Panel Data Analysis of Electric Consumptions

in the Residential Sector

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Abstract

The present work analyses a dataset on electricity consumption in an Italian region for a period of 11 years, using aggregated data on municipalities. Our aim is to estimate the value of the price elasticity, of the income elasticity and cross-price elasticity of other fuels and to analyse the influence of socio-economic, demographic, structural and climate variables on consumption levels and trends. The FEVD procedure, recently proposed by Plumper and Troeger (2007), is adopted to estimate panel data characterised by the presence of time invariant variables, or variables which vary rarely in time.

KEYWORDS: electricity consumption, demand analysis, panel data
JEL Code: Q41, Q48, C23

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1. Introduction

The reduction of electricity consumption in the household sector is a fundamental objective of the EU green paper on energy, yet the trends are on the increase in spite of an analogous tendency observed for the electricity price. Better knowledge of the determinants of electricity consumption and its trends would help insight on which policy instruments could be used to induce households to a more responsible consumption behaviour.

The literature of electricity demand for domestic uses is quite extensive. A large part is devoted to the analysis of the microeconomics determinants of electricity consumption. The aim of these works is to give some insight on what characteristics and habits determine a higher level of consumption.

The vast majority studies utilise a Marshallian demand and insert the average price and a linear or linearised budget constraint. Taylor (1975) and Nordin (1976) propose to insert the marginal price and to add another variable that have to take into account the existence of block of consumption in tariff systems. This variable, called "difference variable", is the difference between the total amount currently paid by a household minus what would have been paid if all units had been charged at the higher marginal price. Shin (1985) indicates that the "difference variable" has an explanatory power only if its size is large enough to become significant and this clearly links it to the characteristics (size and number of the blocks) of tariff system.

Filippini (1995) analyses the consumption of 40 Swiss cities for a period of 4 years. He inserts as covariates the income, the household size, a dummy variable for the availability of the natural gas and a weather variable. Filippini finds that income is not significant, and the same result is obtained by Garcia Cerrutti (2000) and Reiss and White (2005). Reiss and White (2005) argue that this is due to the fact that electricity demand is a conditional demand, so income influences the stock of durable energy goods purchased more than the final electricity consumption. Garcia Cerrutti (2000), Halvorsen and Larsen (2001b), Benavente et al. (2004) insert prices of other fuels, and they verify that gas (Garcia Cerrutti, 2000), wood and heating oil (Halvorsen and Larsen, 2001b) are complements of electricity, but LGP (Benavente et al., 2004) is a substitute. Baker and Blundell (1991) in their analysis of UK households consumption introduce the price of the other fuels and some dummy variable that indicate the presence of a central heating system and the fuel used. Schuler et al. (2000) introduce a dummy variable on the characteristics of the heating system (central heating system or single heating system, single ovens, independent boilers), the quantity of fuels purchased (electricity, coke, oil etc.), a dummy variable for the main fuel used for house heating, the system used to heat water (without tap water boilers, central boiler, independent boilers, central and independent boilers).

The household size variable is positive and significant in almost all the works analysed (see, for example, Filippini, 1995a; Halvorsen e Larsen, 2001b). Recently, the number of variables inserted as covariates has been enlarged with variables on the age of the population and the housing characteristics. In order to understand if age determines different levels of consumption because of varying lifestyles and purchasing powers, many studies introduce variables on the age of inhabitants (Leth-Petersen, 2002). Most of these studies show that older people consume less electricity (Casey e Yamada (2002) e Matsukawa (2007)). Some studies insert variables on the presence of housewives (Filippini, 1995b) retired (Liao e Chang (2002), Higgs et al. (2006)) and unemployed (Rehdanz, 2007; Rehdanz e Stowhase, 2007) family members. These studies aim at seeing if people who spend more time at home consume more electricity, but the empirical analysis gives ambiguous results.

The size of the house and the number of rooms influence positively electricity consumption, even if Baker and Blundell (1991) find a non monotonic relationship. Leth-Petersen e Togeby (2001) e Levison e Niemann (2004) analyse the differences in consumption between owners and renters. They make the assumption that renters consume more electricity because rent includes the electricity bill. The empirical analysis confirms this assumption.

Several studies analyse the influence of durable electric goods on the level and the trend of consumption. Reiss and White (2004) in their study on Californian households find that the price elasticity is influenced by the type of electric appliance used. They find that the price elasticity is lower for more common appliances which are primary goods. Halvorsen and Larsen (2001b) find that the increasing trend of consumption in Norway is due an increasing use of new appliances, like dishwashers and tumble-dryers. A lot of studies are concentrated on the consequences of the spread of use of air conditioning systems, which are considered the hungriest electricity appliance (Matsukawa e Ito, 1998;

Matsukawa, 2004; Ruth e Lin, 2006) and strictly linked to global warming (Sailor and Pavlova, 2003).

Most studies use a climate variable in order to verify how the climate influences the consumption. Filippini (1995), Halvorsen and Larsen (2001b) insert a variable for the number of days in which the temperature was less than 18 degrees (HDD) so that heating was not required. Other studies use a variable for days requiring air conditioning because of temperatures higher than 28°C (CDD) (Garcia Cerruti, 2000; Bernstein et al., 2006).

Recently many studies have been devoted to the relationship between climate and electric consumption (see more on this in: Bigano, Bosello e Marano (2006), De Cian, Lanzi e Roson (2007) and Henley and Pearson (1998)). Summerfield et al. (2007), Levison e Niemann (2004) concentrated on the perceived climate and on the temperature inside the house because this is the important variable when households choose their level of consumption. Filippini (1995b), Baker e Blundell (1991), Rehdanz (2007) introduce some dummies for the city or the region or the district where households live. Reiss and White (2005) insert a dummy variable to distinguish between urban or rural households.

Most studies devoted to the analysis of the microeconomic determinants of electric consumption and to the estimation of the value of the price elasticity, are cross section or time series analysis and use data from sample surveys. The few analyses based on aggregated data adopt a panel data approach. We can cite Shin (1985), Filippini (1995), Garcia Berruti (2000), Benavente et al. (2004), Bernstein e Griffin (2006). With the exception of Filippini (1995) who uses household size, these studies do not consider demographic variables that capture the characteristics of the population and of the household size, but just the income, the price of other fuels (not inserted by Filippini, 1995), the size of the total population of the municipality and climate variables. Dynamic estimators or Fixed Effect estimators are most commonly used; only Shin (1985) and Filippini (1995) use a static approach: Shin (1985) compares the results of a LSDV with a 2SLS, and Filippini (1995) uses a Random Effect estimator.

The present work analyses a dataset on electricity consumption in an Italian region for a period of 11 years (1995-2005), using aggregated data on municipalities obtained through official statistics.

Our aim is to estimate the value of the price elasticity, of the income elasticity and cross-price elasticity of other fuels, in order to analyse the effect of price instruments on the electricity consumption choices; but in addition we aim at analysing the influence of socio-economic, demographic characteristics (household size, age, education etc.), housing (size, number of rooms, renovation, property, etc.), heating systems (central vs. single, fuel used), geography and climate on consumption levels and trends. A new climate variable, the aridity index, which summarizes all the climate variable generally used in literature, will be proposed. Unfortunately, we are not able to evaluate the influence of electric durable goods, since these data are not collected by the National Census and the other surveys on households consumption are only at regional level.

Another interesting feature of our work is the application of a new procedure to estimate panel data characterised by the presence of time invariant variables, or variables which vary rarely in time: the FEVD (Fixed Effect Vector Decomposition) procedure, recently proposed by Plumper and Troeger (2007). It will be seen that adoption of this procedure will allow to obtain reliable estimates even in a situation, as it turns out for our application, where other models would give unsatisfactorily results.

2. The area of study

Sardinia is the second largest island of the Mediterranean sea. It is characterised by mild winters and hot-humid summers. Most of its 1,600,000 inhabitants live in coastal areas where there is a high population density, while 80% of the municipalities have less than 5,000 inhabitants.

Natural gas is not available in the region. The most common heating system is the single heating appliance which heats only a part of the house (70% in municipalities, especially in villages with less than 1000 inhabitants (82%), less in those with more than 30000 inhabitants (25%)). The heating fuel most used is wood (76%, again especially in small municipalities), then electricity (15%), diesel oil (12%), LGP and other gases (8%); while electricity is the most used energy for water heating (76%). Alternative energy sources (solar panels, etc.) are not yet widely diffused, even though recent changes in policy, with relevant subsidies, are expected to induce higher adoption levels. Data on national survey on household consumption (ISTAT, 2005) show that air conditioning systems are increasingly diffused in Sardinia (Table 1).

The dataset includes 374 municipalities (over 377) for a period of 11 years. In this period, the service was managed by a public monopolist (ENEL). The tariffs were based on an increasing tariff block system. Marginal prices per block were established by CIP (Comitato Interministeriale Prezzi) until 1996 and then by AEEG (Autorità Nazionale per l'Energia Elettrica ed il Gas).

3. Methodology

The data has a balanced panel static structure with large N (374 towns) and short T (11 years). The equation to be estimated is assumed to be linear in logarithms.

The static model can be written as:

$$DEPVAR_{it} = \beta\ X_{it} + \gamma\ Z_i + \alpha_i + \varepsilon_{it} \qquad [1]$$

where *i* indexes individuals and *t* indexes time periods, $ELCON_{it}$ is the dependent variable, X_{it} is a 1xK vector of time varying regressors (Cross Sectional Time Series variables, CS-TS forward) and Z_i is a 1xG vector of time invariant regressors (Cross Sectional variables, CS forward). α_i is an individual specific and time invariant error component, assumed *iid* N $(0, \sigma^2_\alpha)$ and ε_{it} is a classical mean zero disturbance, *iid* N $(0, \sigma^2_\varepsilon)$. β and γ are vectors of parameters associated with regressors. α_i is the component of variation not explained by the equation. That is, any factor that is specific to each town and that has not been included among the independent variables will be included in α_i and may be correlated with parts of X and Z. ε_{it} is assumed to be uncorrelated with both the explanatory variables and the effect α_i.

Different estimators can be used depending on specific assumptions on individual heterogeneity.

The simplest model is the Pooled OLS model. It consists of an OLS estimation of model [1]. If we assume that α_i is identical for every town (so individual heterogeneity is all explained in regressors and in the usual error term), the OLS estimates are unbiased and consistent. If heteroskedasticity is present, it is still possible to obtain a correct variance – covariance matrix, using White-Huber-"Sandwich" correction in order to obtain a robust estimation.

If individual heterogeneity is present, we can insert a dummy variable for each individual observation (Least Squared Dummy Variables) but such a model can be difficult to manage if N is quite large.

In that case, it is possible to use a panel specific estimator in order to take into account the individual heterogeneity. If we assume that individual effects are fixed for each town, we can apply the within transformation. It consists of an estimation of that model taking all variables as deviation from individual (town) means. The Fixed Effect (FE) model can be written as

$$\Delta DEPVAR_{it} = \alpha_i + \beta \ \Delta X_{it} + \gamma \ \Delta Z_i + \Delta\varepsilon_{it} \qquad [2]$$

where $\varepsilon_{it} = iid \, \mathrm{N} \, (0, \ \sigma^2_{\varepsilon})$

and $\Delta DEPVAR_{it} = \left(DEPVAR_{it} - \overline{DEPVAR_i}\right)$, $\Delta X_{it} = \left(X_{it} - \overline{X_i}\right)$, $\Delta Z_{it} = \left(Z_i - \overline{Z_i}\right)$ e $\Delta\varepsilon_{it} = \left(\varepsilon_{it} - \overline{\varepsilon_i}\right)$.

The individual effects are modelled by the intercept which varies across all observations. This estimation is consistent and unbiased even if the independent variables are correlated with the individual error, but the within transformation drops out all time invariant regressors from the model.

However, time invariant variables can be important to explain an economic behaviour. Socio economic variables are generally time invariant or are not available as time series, and the Fixed Effect specification reduces the explanatory power of the model. In order to maintain time invariant variables in the model we can adopt a Random Effect – GLS procedure (RE). The model can be written as

$$DEPVAR_{it} = \alpha + \beta X_{it} + \gamma \ Z_i + \left(\alpha_i + \varepsilon_{it}\right) \qquad [3]$$

In this specification individual effects are random variables and individual heterogeneity is explained by a second error term, α_i, *iid* N $(0, \sigma^2_{\alpha})$, and εit is the idiosyncratic error, *iid* N $(0, \sigma^2_{\varepsilon})$. The Random Effect model is consistent and more efficient than Fixed Effect if there is no correlation between α_ι and regressors.

A Breusch – Pagan test allows us to compare the Pooled OLS and the Random Effect Estimator. The null hypothesis is that variance of α_i is zero. If the null is not rejected, there is no individual heterogeneity so Pooled OLS is unbiased and consistent; if the null is rejected the Random Effect Estimator should be preferred.

A Hausman test (1978) can be used to compare Fixed and Random Effects. The Hausman test verifies exogeneity of individual effects Rejection of the null hypothesis on no systematic differences between FE and RE coefficients means that there is correlation between regressors and unobserved individual heterogeneity. If this is the case, RE, as well as OLS, are not consistent and should be rejected.

Unfortunately, as discussed above, the Fixed Effects model can be a very poor choice when the focus of the study is on the effect of time invariant or rarely variant variables on the dependent variable. A possibility to keep time invariant variable in the model even in presence of correlation of regressors and individual heterogeneity is to apply an instrumental variable estimator. Three approaches are most popular in literature: the Hausman – Taylor (HT, 1978), the Amemiya-MaCurdy (AM, 1986) and the Breusch, Mizon and Schmidt approach (BMS, 1989). The last two estimators require stronger exogeneity assumptions than HT. HT requires only that the means of the X_i variables be uncorrelated with the unobserved effect, α_ι while AM and BMS require that variables are not correlated at each point of time. The HT estimator is at least as precise as the within estimator and may avoid the inconsistency of the GLS estimator. Use of the HT and AM estimators requires identifying both correlated and uncorrelated time invariant and time variant variables. The model to be estimated is

$$DEPVAR_{it} = \beta_1 X_{1it} + \beta_2 X_{2it} + \gamma_1 Z_{1i} + \gamma_2 Z_{2i} + \alpha_i + \varepsilon_{it} \qquad [4]$$

where X_{1it} e Z_{1i} are time variant and time invariant uncorrelated variables respectively and X_{2it} e Z_{1i} are time variant and time invariant correlated variables respectively. Unfortunately, the procedure can work well only if instruments are uncorrelated with errors and unit effects and highly correlated with the endogenous regressors, and if instruments are not weak (Plumper and Troeger, 2007). Moreover, individuation of endogenous regressors might be problematic.

Recently, Plumper and Troeger (2007) proposed a new model, alternative to the instrumental variable approaches. They argue that the Fixed Effect estimator may be inappropriate in some circumstances, not only because it drops out all time invariant variables, but also because the estimates of rarely changing variables are inefficient. They suggest a new estimator, called FEVD (Fixed Effect Vector Decomposition). The FEVD is a three-stage estimator. In first stage, the fixed effects vector decomposition procedure estimates a standard fixed effects model to obtain estimates of the unit effects. The dependent variable is regressed on time invariant and rarely changing variables. In the second stage the unit effects from the first stage is regressed on the observed time invariant and rarely changing variables to obtain the unexplained part of fixed effects. In the third stage, the full model (time variant, time invariant, rarely changing variables) is run, including the unexplained part of the decomposed unit fixed effect vector obtained in stage 2. This stage is estimated by pooled OLS.

The authors show, through a Monte Carlo analysis, that FEVD has a better performance than pooled OLS, Random Effect Estimator, and the Hausman – Taylor Estimator if time variant and time invariant variables are correlated with fixed effects. Moreover, FEVD is more efficient than the Fixed Effect estimator to estimate coefficients of rarely changing variables.

3. Econometric analysis

The dependent variable is the log of consumption of electricity (average consumption per user in 374 municipalities, expressed in KWh (ELCON)). Variables inserted as covariates are:

- price of electricity (the mean of the marginal price per the higher block of consumption (MP),
- “Difference” variable of Taylor and Nordin (DIFF),
- average taxpayer income (INCOME),
- price of other fuels (the annual average price of LGP and of wood (PLGP, PWOOD)),
- demographic variables (household size (HHSIZE), proportion of inhabitants older than 65 (OL65), proportion of not working inhabitants (NWI), proportion of inhabitants with secondary schooling or university degree (EDUCATION)),
- housing characteristics (average size of houses (SURFACE), average number of rooms (NROOMS), proportion of property homes (OWNERS), proportion of houses not renovated in the period 1991-2001 (NORENOV)
- heating systems (proportion of homes with electric boilers (WATELECT), proportion of homes with electric heating systems (HELECT), homes with wood heating systems (WELECT), homes with LGP heating system (HLGP)),
- tourist economy variable (indicating the degree of tourist specialisation of municipalities (TOUR))
- climate variable (summer aridity index (AIsummer) and winter aridity index (AIwinter)).

All the variables inserted do not required any explanation, except for the "difference" variable and the aridity index.

The variable "difference" used here, is the inverse of the one proposed by Taylor and Nordin. It is the difference between what would have been paid if all units were charged at the higher marginal price and the actual bill. As suggested by Renwich and Green (2000), the value of this difference variable will grow as consumption increases, thus indicating an implicit subsidy due to the fact the inter-marginal units are paid at the lower block price. Its expected sign is positive.

The summer aridity index and the winter aridity index are calculated as the ratio between precipitation and evapotranspiration. The evapotranspiration represents the environmental demand for water. This demand is determined by the interaction between climate and land use and is strictly linked to solar radiation, temperature, humidity and wind speed. The aridity index gives information on the perceived climate; it varies between 0 and 1, where 1 means a less arid climate, a lower value of evapotranspiration and a higher value of precipitation; otherwise, when it raise a value close to zero, it means that we will have an arid climate with higher value of evapotranspiration and lower value of precipitation. When the summer aridity index is close to 1, we will have a more tolerable perceived climate, so we will expect to use less air conditioning; conversely, when the winter aridity index is close to one, it means a colder perceived climate, so we will expect a higher consumption due to greater use of the heating system. Ceteris paribus, areas with higher evapotranspiration and lower precipitation in the summer months are characterised by higher perceived temperatures. This situation leads to a more intense use of air conditioning. Areas with a lower evapotranspiration level in winter months and higher precipitation are associated with a lower perceived temperature and, we suppose, a higher use of heating systems.

The cross sectional – time series variables used in our study are: the dependent variable, marginal price of the electricity and "difference" variable, income, proportion of inhabitants older than 65 and climate variable. Prices of other fuels vary only across time but not across municipalities. The tourist variable is a categorical variable. All the other variables are only cross sectional variables. A statistical description of variables can be found in Table 2 in the Appendix.

The final specification, selected through standard procedures (Wald and F-test), is the following:

$$
\begin{aligned}
&Log(ELCOM_{it}) = \text{intercept} + \beta_1 \log(MP_{it}) + \beta_2 \log(DIFF_{it}) + \beta_3 \log(INCOME_{it}) + \\
&\beta_4 \log(PGPL10_{it}) + \beta_5 \log(PWOOD) + \beta_6 \log(HHSIZE_i) + \beta_7 \log(OL65_{it}) + \beta_8 \log(NLF_i) + \\
&\beta_9 \log(EDUC_i) + \beta_{10} \log(NROOMS_1) + \beta_{11} \log(SURFACE_i) + \beta_{12} \log(OWNERS_i) + \\
&\beta_{13} \log(NORENOV_i) + \beta_{14} \log(WATELECT_i) + \beta_{15} \log(HELECT_i) + \beta_{16} \log(HWOOD_i) + \\
&\beta_{17} \log(HGPL_i) + \beta_{18}(TOUR_i) + \beta_{19} \log(AISUMMER_{it}) + \beta_{20} \log(AIWINTER_{it}) + u_{it}
\end{aligned}
$$

All estimations have been obtained by STATA 9.2. The results are shown in Table 2.

We start the analysis with the simplest model, the Pooled OLS then we compare it with the the Random Effect Estimator, the Fixed Effect Estimator, the Hausman-Taylor and the FEVD model.

We used some tests to verify the goodness of the models and to select the best fitting.

The Breusch – Pagan test (test value: 6988.99 – p.value: 0.000) for the presence of random effect reject the null hypothesis, which implies that the Pooled OLS is inefficient and the Random Effect estimator should be preferred.

In order to compare the Random Effect and the Fixed Effect Estimators we use a Hausman test. The Hausman test (test value: 30.90 – p.value: 0.000) rejects the null hypothesis of no systematic differences in coefficients, but the matrix of the squares of the differences of the variances of the coefficients is not positive definite. In this case, the Hausman test is not reliable. This problem may occur when the variance of Fixed Effect coefficients is lower than the variance of the Random Effect coefficients. This means that the Random Effect model is not efficient because of the existence of a correlation between regressors and the unobserved individual heterogeneity. A Wooldridge test confirms rejection of the Random Effect model.

Further tests for heteroskedasticity and serial correlation in data (Breusch – Pagan test for heteroskedasticity: test value: 43.53; p-value: 0.000; Wooldridge test for serial correlation of order 1: test value: 132.35; p-value: 0.000) show that both problems are present. All models have been estimated using the Robust Variance – Covariance Matrix of White – Huber – "Sandwich" in order to correct for heteroskedasticity and to obtain robust standard errors.

The supposed presence of a correlation between the unobserved individual heterogeneity and the regressors, lead us to consider alternative estimators to the Random Effects model, because the estimator is inefficient and not consistent. On the other hand, the Fixed

Effects model is inappropriate for our study, since the within transformation drops out all the time – invariant variables.

The solution widely proposed in literature is to use the Hausman-Taylor or the Amemiya-MaCurdy instrumental variable estimators. These models require the researcher to select the variables that are supposed to be correlated with unobserved individual heterogeneity. Evaluating the correlation between the varying part of fixed effect and the regressors in the Fixed Effect model we find that the correlation is most probably due to the marginal price, the difference variable and the price of LGP.

The estimation obtained through these models is quite satisfactory in terms of signs and coefficients, but some uncertainty may be left because of the arbitrariness in the choice of the "endogenous" regressors. The recent proposal of Plumper and Troeger (2007) allows to estimate an efficient and more reliable model. As suggested by the authors, the model is estimated using a correction for the heteroskedasticity, for the contemporaneous correlation between time variant regressor and the error term and the serial correlation of order 1.

The heteroskedasticity of the data is clearly present even in previous estimators. The contemporaneous correlation can be due to the presence of omitted variables (Wooldridge, p. 308): in our case is probably due to the absence of variables related to the stock of electric appliances and especially the stock of air conditioning. The Durbin Watson test shows the presence of serial correlation (DW: 0.83): the statistics calculated after the Cochrane-Orcutt correction indicate a quite satisfactory correction of the problem (DW: 1.60).

The analysis of the results show that the marginal price is significant and with the correct negative sign. The value of the elasticity is – 0.065. Shin (1985), who uses a marginal price, obtains a value of – 0.143/-0.120 in the two estimated models. Filippini (1995a) obtains a marginal price of –0.60. Garcia Cerrutti (2000) obtains a value of –0.04 but Benavente et al. (2006) a value of –0.0548. These studies use as explanatory variable the average price.

The variable "Difference" proposed by Taylor and Nordin is significant and with a positive sign as expected. As discussed above, this variable is used to take into account the presence of increasing blocks even if we used a linearised budget constraint. The adopted formulation represents the implicit subsidy on consumptions for the presence of blocks. This subsidy is higher when the consumption increases. The subsidy indicates the saving due to the presence of blocks because of the consumers purchase inter-marginal units at a price lower than the higher marginal price.

The income variable is not significant. This result is found also in previous studies (Durbin and McFadden, 1984; Filippini, 1995a; Henley and Pearson, 1998; Garcia Cerrutti, 2000; Reiss and White, 2005). This result can be explained considering that the demand for electricity is a demand derived from the use of electric appliances purchased and not a direct request of electricity (Reiss and White, 2005).

The price of LGP is quite significant and with a positive sign, which implies that this fuel is a substitute for electricity (Benavente et al., 2004), but the price of wood is not significant. This result can be explained considering that in small towns people do not purchase wood but collect it in public lands (see Halvorsen and Nesbakken, 2004 for a similar situation).

The household size is significant and with a positive sign, as in previous studies (Filippini, 1995a; Halvorsen and Larsen, 2001b, etc.). The value of the coefficient is 0.74, not far from the 0.81 found by Filippini (1995).

The proportion of inhabitants older than 65 is significant and with a negative sign, as in other studies (Casey and Yamada (2002) e Matsukawa (2007)). The negative sign can be explained considering that older people have a lower purchase power and a reduction in the social life habits that influence consumption (Liao and Chang, 2002).

The education variable is rarely seen in previous studies on the demand of energy. However, it is quite often used in literature on the estimation of energy requirement and efficiency, because higher education may imply a better understanding of energy savings practices (Mansouri et al., 1996; Shen e Sajio, 2008). In our study, this variable is significant and with a negative sign.

The proportion of not working inhabitants is significant and with a negative sign. In literature similar variables (housewives, retired and unemployed people) are inserted to understand if people that spend more time at home consume more. In our case the result is probably due to the lower purchasing power of these people.

Considering housing characteristics, we find that the proportion of property homes is significant with a negative sign. In literature it is widely recognised that renters consume more either because rent is inclusive of the electric bill, or because property homes are built with more efficient energy standard criteria, and the stock of electric appliance is more efficient (see more on: Leth-Petersen and Togeby, 2001; Levison and Niemann, 2004).

The size of the house in squared meters is a positive and significant variable, while the number of rooms is negative and significant. The first result – common in literature – suggests that large houses need more electricity both for heating and for conditioning. The number of rooms in literature shows ambiguous results (Baker and Blundell, 1991; Kalulumia e Green, 2004). In our study the negative sign implies that an increasing number of rooms is associated to less consumption: probably only some rooms in the house are normally used, so only some rooms need heating or conditioning. In recent years the building trend is to have fewer and larger rooms. This evidently would produce an increase in electricity consumption because larger rooms require more light spots, and more energy for heating and conditioning.

The variable on the proportion of houses that have not renovated the electric plant or the windows is significant but negative. We hypothesized that renovation should reduce consumption but the empirical results show the opposite. The data on electric heating systems show that there has been an increase in the diffusion and in the stock of electric heating in the period between 1991 to 2001. Moreover, data collected by ISTAT (2005) show that air conditioning are quite common in Sardinian houses. It can be thought that the renovation is not made to substitute electricity heating with other fuel heating systems but to substitute other fuel heating and to introduce electricity for heating and for conditioning (similar results can be found in: Matsukawa and Ito (1998), Sailor and Pavlova (2003).

The heating system variables show quite intuitive results: houses with electric heating for both air and water consume more electricity and houses with wood heating system consume less. The coefficient of LGP is not significant (while its price is quite significant): probably the low diffusion of this fuel in Sardinian towns can explain this result.

The tourist variable indicates that towns with a higher tourist specialisation consume more electricity. This result is probably due to the presence of a high number of tourist in secondary houses rented in summer months that are officially considered as residents houses.

The climate variables are significant and with the expected sign: the summer aridity index is significant with a negative sign. This variable shows that when the aridity index is close to 1, the perceived climate is more tolerable so there are less electricity consumption because of we do not need to use air conditioning. The winter aridity index is significant with a positive sign. This variable shows that when the aridity index is close to 1, the perceived climate is colder, and more energy is necessary to heat water and house.

5. Conclusions and policy indications

Some interesting policy indications can be found analysing our results. The value of the price elasticity is quite low, so increasing more the price level cannot be useful and can have regressive effects. Considering our results, the recent introduction of a social tariff for low income households should not induce higher consumption level but only have a distributional effect.

Other fuel prices coefficients indicate that LGP is significant with a positive sign, so it can be considered a substitute good; while the price of wood is not significant, probably due to the fact that in small towns people do not purchase wood but collect it in public lands. The variable that indicates the proportion of houses with a LGP heating system is not significant maybe because of LGP is not yet popular in Sardinia: it can be appropriate to provide incentives to adopt LGP and other alternative fuels. On the contrary, the proportion of houses with a wood heating system is associated to a low level of electricity consumption. Taking into account that wood heating systems produce less CO_2 emissions, policies should be implemented to further promoting this type of heating.

Looking at the influence of socioeconomic variables, we first see that the household size is associated to higher levels of consumption, but in a less than proportional way: this means that there are some economies of scale. People with a higher level of education consume less electricity: this can be due to the fact that they are more aware of energy efficiency and energy saving practices, and they take into account these concepts when buying or using an electric appliance. This fact suggests that more extensive and efficient information campaigns should be promoted, in order to induce less educated people to consume less. The proportion of not working people is associated with low consumption: people who probably spend more time at home are not associated with higher levels of consumption, as found in other studies.

Analysing housing characteristics, we found that the number of rooms is negatively associated but the size of the house is positively associated with electricity consumption. This would imply that the current building trend – with less but larger rooms– may induce an increase in consumption. In addition, the proportion of houses that was not be renovated show a negative association with consumption: this result, which goes against ours expectations, may be due to the fact that renovations may often place in an air conditioning system.

We can notice that towns with a colder climate during winters are associated to higher levels of consumptions (probably due to a more frequent use of the electric heating system) and towns with a less warm climate during summers are associated to an inferior level of consumption (probably due to a less frequent use of air conditioning).

Overall, these results suggest to keep promoting good building practices, and to promote the adoption of alternative heating systems and renewable energy to compensate the increase of consumption due to the growing diffusion of air conditioning systems.

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Table 1. Diffusion of air conditioning in Italian Regions

	Air conditioning				
Geographical areas	**2001**	**2002**	**2003**	**2004**	**2005**
North– West	8.8	10.3	10.3	16.2	16.3
North – East	19.8	22.1	26.9	32.6	31.2
Middle Region	7.5	9	13.1	15.2	15.2
South	5.7	8	12.6	16	17.2
Isole (Sardinia and sicily)	**17.4**	**22.4**	**27.9**	**31.9**	**32.2**
Italy	10.9	13.1	16.4	20.8	20.9

SINGLE REGIONS (Year 2005)	**Air conditioning**
Sicily	30.7
Sardinia	37.0

Tabella 2. Table of statistics *(follow on the next page)*

	Media	Std. Dev.	Data	Definition	Source
ELCON	2646. 68	445.04	CS-TS	average consumption per user in 374 municipalities (in KWh) per year 1995-2005	ENEL
MP	0.12	0.059	CS-TS	the mean of the marginal price per the higher block of consumption per municipalities per year 1995-2005	ENEL
DIFF	142.792	110.631	CS-TS	Taylor – Nordin "Difference" Variable per municipalities per year (1995-2005)	Our computation
INCOME	10073.3	2583.44	CS-TS	Average taxpayer income per municipalities per year (1995-2005)	Ministry of Treasure
PLGP	10.729	2.704	TS	Annual average price of LGP per year (1995-2005)	Camera di Commercio di Cagliari
PWOOD	10.150	3.473	TS	Annual average price of wood per year (1995-2005)	Camera di Commercio di Cagliari
HHSIZE	2.723	0.238	CS	Average household size	Istat, 2001
OV65	0.191	0.544	CS-TS	Proportion of inhabitants older than 65 per municipalities per year (1995-2005)	Istat, 1995-2005
EDUCATION	0.208	0.057	CS	Proportion of inhabitants with secondary schooling or university degree	Istat, 2001
NWI	0.571	0.053	CS	Proportion of not working inhabitants	Istat, 2001
OWNERS	0.845	0.068	CS	Proportion of property homes	Istat, 2001
NROOMS	4.551	0.449	CS	Average number of rooms	Istat, 2001
SURFACE	103.431	11.922	CS	Average size of houses	Istat, 2001
NORENOV	0.741	0.061	CS	Proportion of houses not renovated in the period 1991-2001	Istat, 2001
WATELECT	0.715	0.115	CS	Proportion of homes with electric boilers	Istat, 2001

HELECT	0.151	0.117	CS	Proportion of homes with electric heating systems	Istat, 2001
HWOOD	0.759	0.187	CS	Homes with wood heating systems	Istat, 2001
HLGP	0.081	0.058	CS	Homes with LGP heating system	Istat, 2001
TOUR	0.393	0.991	Categorical dummy	Homes with LGP heating system (homes with LGP heating system)	Our computation
AISUMMER	0.157	0.099	CS-TS	Summer aridity index per municipalities per year (1995 – 2005)	SAR
AIWINTER	1.207	0.747	CS-TS	Winter aridity index per municipalities per year (1995 – 2005)	SAR

Table 3. Estimation results *(follow in the next page)*

	OLS Robust S.E.	FIXED EFFECT Robust S.E	RANDOM EFFECT – GLS Robust S.E.	HT[1] bootstrap standard errors	FEVD[2] pcse and ar1
INTERCEPT	4.577*** (33.32)	6.91*** (55.77)	5.398*** (26.45)	5.50*** (21.83)	5.67*** (120.15)
MP	-0.187*** (-16.55)	-0.080*** (-8.30)	-0.093*** (-9.63)	-0.081*** (-5.53)	-0.065*** (-3.44)
DIFF	0.159*** (26.76)	0.052*** (7.22)	0.064*** (9.23)	0.052*** (4.39)	0.047*** (4.65)
INCOME	0.065*** (7.16)	-0.002 (0.26)	0.008 (1.18)	0.009 (0.09)	-0.011 (-0.79)
PGPL10	0.088*** (4.02)	0.146*** (10.30)	0.143*** (10.03)	0.148*** (9.99)	0.132*** (3.52)
Pwood	-0.006 (-0.46)	-0.012 (-1.37)	-0.013 (-1.36)	-0.013 (-1.28)	-0.021 (-0.77)
HHSIZE	0.476*** (17.84)	- -	0.610*** (10.60)	0.634*** (5.91)	0.759*** (18.13)
OV65	-0.136*** (-13.08)	-0.132*** (-6.44)	-0.140*** (-8.31)	-0.136*** (-3.87)	-0.069*** (-3.40)
EDUCATION	-0.013*** (-5.04)	- -	-0.010* (-1.72)	-0.010 (-1.33)	-0.009*** (-4.60)
NLF	-0.068*** (-3.57)	- -	-0.089** (-2.06)	-0.095 (-1.57)	-0.135*** (-11.53)
OWNERS	-0.135*** (-5.84)	- -	-0.171*** (-3.41)	-0.178** (-2.37)	-0.203*** (-15.75)
ROOMS	-0.241*** (-7.90)	- -	-0.277*** (-4.35)	-0.287*** (-3.06)	-0.353*** (-30.08)
SURFACE	0.194*** (7.89)	- -	0.228*** (4.55)	0.235*** (3.43)	0.259*** (21.53)
NORENEW	-0.082*** (-4.43)	- -	-0.099** (-2.12)	-0.101 (-1.53)	-0.094*** (-5.64)
WATELECT	0.109*** (11.03)	- -	0.125*** (5.18)	0.126*** (4.58)	0.116*** (12.66)
HELECT	0.028*** (11.66)	- -	0.035*** (6.39)	0.036*** (4.28)	0.038*** (37.15)
HWOOD	-0.077*** (-13.29)	- -	-0.092*** (-8.16)	-0.094*** (-4.69)	-0.096*** (-30.42)
HGPL	-0.002 (-0.98)	- -	-0.002 (-0.32)	-0.002 (-0.27)	-0.000 (-0.24)
TOUR	0.012*** (6.11)	- -	0.015*** (3.39)	0.016*** (2.68)	0.017*** (8.52)
AIsummer	-0.006**	-0.003**	-0.003**	-0.003*	-0.009***

	(-2.24)	(-2.12)	(-2.17)	(-1.70)	(-3.15)
AIwinter	0.003 (1.04)	0.011*** (5.51)	0.010*** (4.89)	0.010*** (4.77)	0.011** (2.57)
N	4114	4114	4114	4114	3740
R-squared	0.74	0.57	0.71	0.70	0.92
Rho	-	0.87	0.54	0.71	-

In brackets: t statistics for OLS, FE, RE-GLS and FEVD, z values for Instrumental Variables Estimators; * 10% significance level, ** 5% significance level and *** 1% significance level.
1 = marginal price, difference variable and the price of GPL are considered correlated with unobserved individual heterogeneity; 2 = FEVD with pcse (for heteroskedasticity and contemporaneous correlation) and ar1 (for serial correlation) correction option

Table 4. Test results

Test	Valore della statistica del test	p- value
Breusch – Pagan test for no heteroskedasticity in pooled OLS	43.53	0.0000
Wooldridge test for no autocorrelation in panel data	132.355	0.000
Durbin Watson test for serial correlation in FEVD (model without correction for heterosckedasticity and serial correlation)	0.83	Lower DW bound: about 1.89
Durbin Watson test for serial correlation in FEVD (model with Cochrane – Orcutt correction for serial correlation)	1.60	Lower DW bound: about 1.89
Breusch Pagan LM Test for no random effect in panel data	6988.99	0.000
Hausman test for no endogeneity between FE and RE	30.90*	0.0001
Hausman test for no endogeneity between FE and HT	Hausman test cannot be run, because of the fitted model fails to meet the asymptotic assumption of Hausman test	
Wooldridge test for no differences between FE and RE	384.60	0.000
Hausman test for no differences between FE and HT	363.62	0.000

* matrix of coefficients is not positive definite

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