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CERE Working Paper, 2013:5

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February 23, 2014

Abstract

We provide consistent, cross-country estimates of price and income elasticity for the residential sector of 11 OECD countries. Using survey data from 2011 on annual consumption of electricity and sample-derived average electricity price, we provide country-specific price elasticity estimates and average income elasticity estimates. For most countries in our sample, we find strong price responsiveness, with elasticities varying between -0.27 for South Korea and -1.4 for Australia, with most countries elasticity being above (in absolute value) -0.5. Exploiting the presence of many attitudinal indicators, we find evidence for non-price related factors, in particular households' self-reported energy saving behavior, to reduce energy demand between 2 and 4%. In contrast, we find very weak income response, with income elasticities varying from 0.07 to 0.14 and no evidence for heterogeneity across the countries in our sample. Our results regarding price elasticity are in contrast with many existing studies which find low-to-moderate price responsiveness, and adds to a few recent studies indicating more policy space for demand reduction than previously thought.

Keywords: Residential Electricity Demand, Price Elasticity

JEL Classification: Q4, Q41, C5, D12

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*The data used in this work come from an survey on Environmental Policy and Individual Behaviour Change (EPIC) periodically conducted by the Environment Directorate. The views expressed here do not necessarily reflect those of the Organisation for Economic Cooperation and Development (OECD) or its member countries.

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1 Introduction

The issue of climate change mitigation has led to a resurgence of studies regarding household electricity demand. Such studies have focused on three distinct issues; understanding price responsiveness, appliance choice and the impact of policy on energy demand, including the issue of "rebound" effects. Most studies on the first and third issue work with two types of data, household level or aggregate (local, regional or state/province/country-level). Aggregate data are able to distinguish larger-scale patterns, such as regional variation in key parameters, which household data typically do not allow one to infer (primarily due to the lack of panel household data sets). On the other hand, parameters estimated from such aggregate studies are difficult to interpret or apply to policy, in addition to suffering from well known aggregation and other biases (see e.g. Fell et al. (2011)).

Household data sets, however, are being increasingly used in recent studies in the US (Fell et al. (2011); Alberini et al. (2011)) to provide detailed estimates of household-level parameters at the regional level. In general, the substantial heterogeneity which exists in consumption behaviour across space is acknowledged but rarely addressed in much of the existing literature.

The current study addresses, for 11 OECD countries, the issue of estimation of price- and income-elasticity of electricity demand, using a cross-section of survey data. This is, to our knowledge, the first systematic attempt at assessing important demand parameters from *household data* in a *cross-country* setting. In particular, this study addresses the following issue: estimation of income and price elasticities, exploring both the commonly used double-log and a more reasonable alternative, the trans-log functional form. In addition, the issue of possible parameter heterogeneity is addressed by estimating country-specific coefficients on income and price, using the double-log functional form. Finally, the study also provides some indications of the potential empirical significance regarding non-price-related factors ("attitudinal" factors), preference heterogeneity and the issue of split incentives.

It is important to mention at the outset that the price elasticities are estimated here based on a moderate-sized sample, with *average prices* derived from the survey data; as a result, these elasticity estimates need further validation. Unlike in recent papers which have used either utility-level aggregate data (Fell et al. (2011), Alberini et al. (2011)) or instruments (Alberini & Filippini (2011)), our data set does not provide information regarding either utility-level or region-level price (or other utility data), and we can use neither of the above strategies. We therefore derive the household-specific average price from survey-reported annual quantity estimates. We note that consumers in our survey are in general billed on a per-unit basis, and indicate that they are aware of the marginal price (about 90% of the sample, see Kriström (2013)); nonetheless, there is little empirical evidence regarding the basis (marginal or average price) upon which consumption decisions are made.

In using average price data (at the household level) we assume, explicitly as in some of the recent literature (e.g. Borenstein (2009), Fell et al. (2011) and Ito (2012)) and implicitly in others (e.g. Metcalf & Hassett (1999)), that consumers make consumption decisions based on *average*, not *marginal* prices. In our case, in common with Alberini et al. (2011), we resort to using average price due to missing data on marginal price. We are unable to obtain data regarding marginal price (and fixed components of price) both due to the magnitude of the task (of collecting utility-region level prices for different countries, a point emphasized in (Alberini et al. 2011) for the US), as well as due to lack of data regarding the utility serving the surveyed households.

An increasingly important topic of current research has been in understanding which prices are actually used by consumers in their decision making (Ito (2012) and Borenstein (2009)). The most common approach to estimating electricity demand, beginning at least with Taylor (1975), is to model a consumer who is perfectly informed about the marginal price (or price schedules) and can adjust consumption to the optimal costlessly, an assumption which "...seems to be at odds with the way nearly everyone thinks about residential (water and) electricity consumption" ((Borenstein 2009, pp 3)). In particular, given the restrictions on consumers' ability (or willingness) to respond to every price change, many studies report average prices to be more consistent with observed consumer demand (Shin (1985), Bushnell & Mansur (2005), Borenstein (2009), Ito (2012)).

We stress that we do not take any position regarding which measure of price, marginal or average, is actually used by individuals in our sample to make their decision. Nonetheless, we believe that there are a few reasons for consumers to be unable to optimize as commonly used approaches assume (see also (Fell et al. 2011, pp 2-3) and Borenstein (2009)). First, the billing period is variable across consumers (typically a month, but as much as 3 months or more in certain cases); second, the marginal price, while constant, is typically unknown when consumption decisions are made (except for countries with highly regulated prices e.g. France) and finally, given the aggregation to an annual level, consumers are more likely to have either a certain share of income or a fixed amount of nominal value of consumption in mind.

As indicated in Borenstein (2009), there are at least three measures of prices available: ex-post marginal price[1], expected marginal price and an average price based on the ex-post quantity consumed[2]. Observe that whenever the marginal price schedule is non-linear (as for the US, and a few countries in our sample) or there is a fixed component to the price (e.g. network charges)–as in our case–the average price differs from the marginal. Given constant marginal price and moderately large consumption (as in our case when annual consumption is used), the difference between the marginal and average price is likely much less than the fixed cost as a fraction of the bill. As a result, numerically, the effect of using average price when marginal price is the correct one, is likely small.

We briefly summarize the main results from our base specification, a double-log de-

[1] Marginal price paid by the consumer at the end of the consumption period, as opposed to one known before making consumption decisions, as is commonly implicitly assumed.

[2] Here on in, our references to "marginal price" and "average price", unless otherwise indicated, are to the ex-post marginal price and the average price based on observed ex-post consumption.

mand function with country-specific fixed effects. We find a very high, in certain cases unitary, price elasticity for most countries in the sample (ranging from -0.25 for Korea to -1.4 for Australia), with most countries having a high price elasticity. Netherlands and South Korea are the only countries with an elasticity below -0.5, at about -0.25 and -0.26 respectively. We also find low, between 0.05 and 0.14, and rarely significant, income elasticity. Further, in our main specification, we find no evidence for variation of income elasticity across countries. In addition, we find only modest differences in elasticity between home owners and renters in terms of electricity demand, after controlling for other observables including total number of appliances. We find that an increase in the energy behaviour index reduces energy consumption by 2 to 4%, while membership in an environmental organization has no significant impact on demand.

We also find the usual relationships to hold with the control variables: households with greater size, larger homes, more appliances and electric space heating (or cooling) use more electricity. While the income elasticity estimated here is well within the range of estimates for other settings in the literature, we find the price elasticities to be very large in comparison to most of the recent literature. Only Fell et al. (2011) and Alberini et al. (2011), among recent studies, report elasticities as high as we do.

Our results are robust to a range of potential issues, including: possible endogeneity of price for a few countries in the sample which follow the increasing-block-pricing structure, regional differences in policies (accounted for by using country-region-specific fixed effects) and outliers. Since the double log functional form, in addition to its well documented drawbacks (see for instance Plourde & Ryan (1985)), cannot be consistent with any utility function when price (or income) elasticities are close to unity[3], we also provide

[3]To see this, note that the indirect utility function for an individual with the double-log demand function can be written (Hausman (1981, equation (21))) as: $v(p,y) = -e^{\Gamma X}\frac{p+\gamma}{1+\gamma} + \frac{y^{1-\beta}}{1-\beta}$, using notation from eq. (2). It is evident that, for $\gamma \to -1$, this utility function is infinite, $\forall(p,y)$. We note that Fell et al. (2011) ignore this issue in their estimation using the double-log form, an issue particularly relevant for them due to their finding of a unitary price elasticity.

elasticity estimates from an alternative and encompassing functional form, the translog, which does not suffer from such drawbacks. Thus, our results are also robust to usual criticisms pertaining to choice of functional form. Finally, we emphasize that our identification of the income elasticity does not depend upon which price is used by consumers to make these decisions in our main specification, the double-log functional form (see section 2).

We briefly comment on the issue of endogeneity of average prices, when computed based on actual consumption. Note that in a specification with quantity of electricity as the dependent and price as an independent variable, endogeneity of price arises whenever average price (computed as ratio of expenditure to actual usage) is used as the measure of price[4]. This type of endogeneity is essentially mechanical, and is independent of the non-linearity of the marginal pricing scheme. However, using the case of the double-log demand function, it can be shown for our case (see section 2) that the degree of bias depends non-linearly upon (marginal) prices, and is likely not large enough to explain away our results.

To summarize, if consumers make consumption decisions based upon marginal prices, instead of average as we assume, then the magnitude of bias on our estimated price (and income) elasticity is unlikely to be sufficient to explain the magnitude of the price elasticity we find. Finally, given that the magnitude of bias in price- and income-elasticity is likely dependent upon functional form and data configuration, the fact that our main findings–of high price elasticity and low income elasticity–are identical across functional forms (translog) and a variety of data configurations (exclusion of a group of countries, pooling all countries, accounting for regional variation and outliers; see section 4.2) provides some confidence that our results are not severely biased [5].

[4]There is another, more common, reason for endogeneity of the marginal price, whenever marginal price is used on the r.h.s of a demand equation with observed quantity on the l.h.s. In the face of increasing-block-pricing, when the full schedule of marginal prices are unavailable, when any single measure of marginal price (such as average marginal price) is used as a regressor, the price variable is "endogenous" due to being simultaneously determined with quantity. See Reiss & White (2005) for an approach to estimation of price elasticity with this type of endogeneity.

[5]Whenever household-level (marginal or average) prices are unavailable, some aggregation of price is used instead of the actual price (a very common procedure is to use state- or utility-level aggregate prices; e.g. Branch (1993), Alberini et al. (2011) and Fell et al. (2011) or most of the studies cited in section 4.3), leading

The plan of the paper is as follows. We lay out the econometric approach in section 2, discuss the data and summary statistics in section 3 and present our results in the context of existing literature in section 4. Section 5 provides a discussion and concluding remarks.

2 Econometric Approach

Denote by $E_{i,j}, Q_{i,j}$ & $P_{i,j}$ the levels of (annual) expenditure (euro), quantity consumed (KwH) and (marginal) price of electricity (euro) faced by consumer $j = 1,\ldots,J_i$ in country $i = 1,\ldots,I$. Data on $E_{i,j}$ is available, while that on one of $Q_{i,j}, P_{i,j}$ is not. Posit the usual relationship between $Q_{i,j}$ and $P_{i,j}$

$$\ln(Q_{i,j}) = \beta \ln I_{i,j} + \gamma \ln P_{i,j} + \Gamma X_{i,j} + \varepsilon_{i,j} \tag{1}$$

where X is a vector of control variables, including appliance holdings, individual characteristics and other relevant non-economic factors. Writing out explicitly the equation for expenditure, $E_{ij} = Q_{ij}P_{ij}$, in log and substituting eq. (1), we have

$$\ln\left(E_{i,j}\right) = \beta \ln I_{i,j} + \underbrace{(1+\gamma)}_{\tilde{\gamma}} \ln(P_{i,j}) + \Gamma X_{i,j} + \varepsilon_{i,j} \tag{2}$$

Thus, eq. (2) is easily estimable, with data on P and E. In a large number (seven) of the countries sampled here, endogeneity of marginal price is not an issue (see footnote 4), given the constant marginal price (with substantial variation at the region and home-type level). Presuming data on quantity consumed (KwH) are available (as in our case), it is possible to compute $\bar{P}_{i,j} := \dfrac{Q_{i,j}}{E_{i,j}}$. Provided a sample of data on $E_{i,j}$ and $\bar{P}_{i,j}$ is available,

to measurement errors. Only recently (Alberini & Filippini (2011); see also Alberini et al. (2011)) has this issue been investigated in some detail (in panel data settings); the tentative conclusion of these US-based studies–which assume classical measurement error– is that the resulting bias is modest, at about 10-15%. Given the lack of additional data at the country-level, obtaining valid instruments in our cross-section setting is infeasible. Nonetheless, the findings in Alberini et al. (2011) provide some comfort that the large price elasticity (and very small income elasticity) found in our setting is unlikely to be an artifact of measurement errors in price.

it is straightforward to estimate the parameters of interest, γ and β, the price and income elasticity[6,7].

We note here that we have not indexed any of the coefficients with i, but that is purely for notational convenience; in the setting above, equations (1)-(2) may well be thought of as being country-specific, with country-specific coefficients. This speaks to our issue of addressing (observed) heterogeneity; estimated elasticities vary substantially over countries, even within a similar methodology. In addition, for the larger countries, there is likely to be substantial heterogeneity within a country, and finally, heterogeneity in preferences across individuals.

Our strategy for addressing this issue is the following: we do not wish to impose parameter homogeneity in our estimation of eq. (2) (eq. (1)) above. Instead, we take the approach of testing for parameter homogeneity. The strategy is to test for similarity of coefficients between the unrestricted model, with many coefficients being country-specific, and a restricted model with only country-fixed-effects, a traditional approach to addressing heterogeneity.

In more detail, we consider testing between the following two models:

$$\ln\left(E_{i,j}\right) = \alpha_i + \beta_i \ln I_{i,j} + \tilde{\gamma}_i \ln P_{i,j} + \Gamma X_{i,j} + \varepsilon_{i,j} \qquad \text{(Unrestricted)}$$

and

$$\ln\left(E_{i,j}\right) = \alpha_i + \beta \ln I_{i,j} + \tilde{\gamma} \ln P_{i,j} + \Gamma X_{i,j} + \varepsilon_{i,j} \qquad \text{(Restricted)}$$

It is then straightforward to estimate the two models and to test the following hypothesis:

$$H_0 : \tilde{\gamma}_i = \tilde{\gamma} \;\&\; \beta_i = \beta \;\forall i \tag{3}$$

[6]Note that the coefficient on $\ln P_{i,j}$ in the equation for expenditure, denoted $\tilde{\gamma}$, is not the elasticity of interest, γ; evidently, $\gamma = \tilde{\gamma} - 1$, and its standard error can be computed from that of $\tilde{\gamma}$.

[7]Using data on E and Q to estimate P, as we do, it is evident that both eq. (1) and eq. (2) yield identical estimates. We write out both equations for purposes of comparability to other frameworks, including that in Fell et al. (2011). In subsequent developments, we will always estimate the quantity equation, eq. (1).

This approach, while intuitive, also illustrates the limitation of this framework. There is no reason to restrict differences in elasticities to be only dependent on the country; any other observed characteristic (such as possession of an energy efficient appliance, of being a home owner, or belonging to an environmental organization etc) can, in addition, be envisaged as giving rise to differences. At the extreme, such an approach would lead to an estimation of a very large number of parameters, one for each country-category combination. In order to obtain interpretable and usable results, we restrict such testing to only country-specific coefficients, although interactions between other key variables and countries were not found to be significant[8].

It is to be noted that the coefficients estimated in eq. (Unrestricted) are not equivalent to country-by-country estimation, for two reasons: there are many cross-equation restrictions (most coefficients are restricted to be equal across countries) and the error variance is identical across countries. An important advantage of estimating demand parameters for many different countries in the same framework and using comparable data is the consistency in estimation and resulting comparability of parameters across countries. In addition, in the context of our data, the small and highly variable sample sizes at the country-level (varying from 85 for Australia to 218 for Sweden, see table 1), the estimation of country-by-country regressions lead to difficulties in both estimation and interpretation.

Note that while price elasticity is not estimated using country-by-country regressions, country-specific price elasticities are identified based on within-country variation in price (when country-specific fixed effects are used). Country-specific sample sizes (varying between 85 and 218, see section 3.2) are moderate on average, by modern demand estimation standards. Nonetheless, identification challenges are considerably mitigated in our case by two factors: the substantial within-country variation in price (documented in section 3.2) and the rich source of preference-related data, allowing for more controls than are typi-

[8]Two other indicators were postulated as being of some interest, in this context: possession of an energy efficient appliance and of belonging to an environmental organization, one at a time.

cal in demand estimation. In addition, price elasticity estimated from the regression with country-specific effects is very similar to that obtained from a variety of data configurations (see section 4.2), including from the pooled model with 1410 observations (a much larger sample size), mitigating any such concerns.

Yet another issue worth highlighting is the dependence of interpretation on functional form: for a general functional form (even, for instance, a linear form), elasticities are a function of the level of a variable, as well as possibly of all other variables in the model. As a result, we present estimates of elasticity using an alternative and encompassing functional form, the translog, both as a check of robustness of results using the double-log form and for behavioral validity .

We provide results using the trans-log[9] functional form:

$$\begin{aligned}\ln(Q_{i,j}) = \beta_0 \ln I_{i,j} + \gamma_o \ln P_{i,j} + \Gamma X_{i,j} + \beta_1 \left(\ln I_{i,j}\right)^2 + \\ \gamma_1 \left(\ln P_{i,j}\right)^2 + \delta \left(\ln P_{i,j} \ln I_{i,j}\right) + \varepsilon_{i,j}\end{aligned} \tag{4}$$

Even in the absence of country-specific coefficients, the elasticities of price and income each depend on all other parameters and variables and, in particular, on each other. The price and income elasticities, denoted η_P, η_I respectively, are: $\eta^P_{ij} = \gamma_0 + 2\gamma_1 \ln P_{i,j} + \delta \ln I_{i,j}$ and $\eta^I_{ij} = \beta_0 + 2\beta_1 \ln I_{i,j} + \delta \ln P_{i,j}$. Thus, each observation has its own elasticity.

In this functional form, on the other hand, there is no consistent and parsimonious way of allowing for country-specific *coefficients*[10]. Therefore, for this specification, we do not allow for country-specific coefficients on any variable; nonetheless, since the price (income) elasticities are, in addition to their own levels, a function of income (price), there is

[9]Note that the term “trans-log”, as used here, refers to the quadratic-in-prices-and-income aspect of the demand function for electricity. Unlike in the case of demand system estimation, there are no restrictions on the parameters for consistency with theory. Essentially, this form is simply viewed, in much of the literature on electricity and other fuel demand (see Wadud et al. (2010), Liu (2011) and references therein), as a quadratic approximation to an arbitrary functional form in I , P.

[10]If country-specific coefficients are postulated on price and income, then all transformations of price and income, including the squared and interaction terms, must also have country-specific coefficients, for sake of consistency. Given the significant increase in the number of parameters which results (an additional 30 parameters, compared to equation (2)), we do not pursue this route here.

substantial variation in country-specific price (income) elasticities. This variation is essentially a reflection of the country-specific variation in price (income).

We report, for eq. (4), country-specific (and overall) mean elasticities, computed at the country-specific mean of price and grand mean of income (with standard errors computed using the delta-method). More precisely, $\overline{\eta_i^P} = \widehat{\gamma}_0 + 2\widehat{\gamma}_1 \ln \overline{P_i} + \widehat{\delta} \ln \overline{I}$ and $\overline{\eta^I} = \widehat{\beta}_0 + 2\widehat{\beta}_1 \ln \overline{I} + \widehat{\delta} \ln \overline{P}$ are reported. Finally, since for four countries (South Korea, Australia, Japan and Israel) in the sample, certain regions (or all of South Korea, which is served by a single utility), use increasing-block-pricing, we exclude these countries and re-estimate eq. (1) and eq. (4)[11]. We emphasize that this is in the nature of a robustness check.

To summarize, for the double-log form of the demand function, separate price coefficient (price-elasticity) is estimated across countries while for the trans-log form in eq. (4), coefficients on all variables are restricted to be identical across countries. The non-linearity of the functional form (and country-specific variation in income and prices), however, yields country-specific price elasticity estimates.

We briefly turn to the task of speculating, for our main specification in eq. (1), about the possible magnitude of bias from using average price, which is endogenous. Define $\overline{P}$, computed as $\overline{P}_{i,j} = \dfrac{P_{i,j} \times Q_{i,j} + \kappa}{Q_{i,j}}$, as the average price, with κ a fixed component of the electric bill, potentially as high as 30% of the total bill. We rewrite eq. (1) below in terms of average (instead of marginal) price, as:

$$\ln(Q_{i,j}) = \gamma \ln \overline{P}_{i,j} + \Gamma X_{i,j} + \varepsilon_{i,j} \tag{5}$$

[11]These countries have relatively complicated residential consumer pricing schemes. The block-pricing structure in South Korea is sophisticated, with non-linear block-pricing for low and high-voltage consumers while for Japan, block-pricing schemes depend upon Ampere contracted, time-of-day, and a so-called "fuel cost adjustment" that varies by month. Finally, in Australia, pricing of electricity is complex, eg. rebates are available for certain types of consumers.

Using the definition of $\overline{P}$, the following version of eq. (1) is obtained[12]:

$$\ln(Q_{i,j}) = \gamma \ln P_{i,j} + \Gamma X_{i,j} + \underbrace{\left(\varepsilon_{i,j} + \frac{\gamma\kappa}{P_{i,j} \times Q_{i,j}}\right)}_{v_{i,j}} \tag{6}$$

A few points about eq. (6) are worth noting. First, unlike in the cases considered in (Fell et al. 2011, eq. (2)) (where both Q and P are estimated), we have information on two components, E and Q, mitigating the degree to which we use estimated information. Second, assuming the true value of γ is far from 0 (else the bias is likely insignificant), while the composite error term, $v_{i,j}$, is correlated with both $\ln P_{i,j}$ and (via Q) $X_{i,j}$, the correlation is likely low for a variety of reasons, including the highly non-linear nature of the relationship[13]. Finally, as shown in Halvorsen (1975), if both supply and demand equations–based on average prices– are specified in the double-log form for a consumer who makes decisions based on marginal price, the coefficient on average price in a regression such as that in eq. (5), is unbiased.

To summarize, while we are unable to either use instrumental variables approaches or estimate supply and demand jointly to address endogeneity (due to data limitations), the impact of endogeneity on our estimated price- and income-elasticity is likely small. We also emphasize again that an important reason for this conclusion is the constant marginal price faced by most consumers in our sample[14], as opposed to the increasing-marginal-

[12] Substituting $\overline{P}_{i,j} = \frac{P_{i,j} \times Q_{i,j} + \kappa}{Q_{i,j}}$ in eq. (5), using the decomposition of $\ln(PQ + \kappa)$ into $\ln(PQ) + \ln\left(1 + \frac{\kappa}{PQ}\right)$ and then, noting that $\kappa < P \times Q$, using the first order Taylor series expansion of $\ln\left(1 + \frac{\kappa}{PQ}\right)$ as $\frac{\kappa}{PQ}$, results in eq. (6).

[13] The non-linear relationship obtained here is in contrast to the results in Fell et al. (2011, eq. (2), pp 5). Due to the fact that both Q and P are estimated there, the final equation has a residual which is a linear function of the included average price, leading to potentially higher bias as a result.

[14] The nature of the marginal price significantly affects the interpretation and magnitude of endogeneity. To see this, note that the discussion in Taylor (1975) regarding endogeneity assumes a positive correlation between average and marginal price. While this is true when price schedules are of increasing-block type, for constant marginal price with a fixed cost, the correlation is actually negative, since average prices fall (limiting to the marginal) with consumption.

price scenario typically considered for the case of the US (e.g. Taylor (1975), Halvorsen (1975) among the earlier studies, and Fell et al. (2011) and Alberini et al. (2011) among the more recent ones).

3 Data Description and Summary Statistics

3.1 The Survey

Data for the analysis was drawn from the OECD's project on *Greening Household Behaviour*, as part of which a periodic survey on Environmental Policy and Individual Behaviour Change (EPIC), covering a number of countries and areas, is carried out. The second survey was conducted in 2011 and included 11 countries: Australia, Canada, Chile, France, Israel, South Korea, Japan, the Netherlands, Spain, Sweden and Switzerland. We provide a very brief description of the survey and refer to OECD (2013, Annex B) for details.

About 1100 individuals in each country were surveyed using an internet-based questionnaire, for a total sample size of 12,200 households. The questionnaire collected information regarding household behaviours in five distinct areas (apart from household characteristics and environmental attitudes): residential energy use, waste generation and recycling, food consumption, personal transport, and water consumption. The present analysis uses data from the energy section.

Sample selection followed a strategy of stratification based on income, age-group, region and gender. In order to account for non-random sampling, ex-post probability weights were provided, which are used to render estimation results using this sample comparable to those using random samples from country-level population distributions.

We note that individuals were requested to provide data on their electricity bill (annual) and quantity consumed in Kwh (annual). Very few individuals provided billing data, and of those, a few provided quantity data, allowing computation of the average price. As a result, the final sample size, about 1100, is a fraction of the usable responses of approximately $11,000$ households[15].

3.2 Variable Descriptions and Summary Statistics

We provide a brief discussion of selected summary statistics for the regression sample, as well as discuss prior expectations (based on existing literature) regarding the signs of many of the control variables appearing in eq. (1).

We begin with summarizing the reported electricity usage and billing data. The mean expenditure varies widely across countries, from a low of 175€ in South Korea to a high of 1950€ in Sweden, which is consistent with the variance in usage, South Korea being the lowest and Sweden, the highest. The derived average price mirrors the variation above; noteworthy however is the relatively low Swedish electricity prices, at 16 (euro) cents, compared to both the Netherlands and Spain, at 27 and 28 (of the EU members). Overall, while colder countries (Canada, Sweden and Switzerland) appear to consume more electricity, cross-country variation in prices and possible existence of substitute sources of energy render such comparisons tricky. All of these figures are consistent with country-reported data (see Kriström (2013) for a more detailed analysis of the EPIC 2011 data on energy).

Finally, given that obtaining meaningful price variation is typically a challenge in studies of electricity demand, we check to see that there is indeed sufficient variation in within-

[15] We note that concerns regarding sample selection in our context are unlikely to be important. This follows from the fact that only if households do not report expenditure (quantity consumed) based on *unobserved* factors unrelated to the explanatory variables (i.e. selection on *unobservables*) are the estimates biased. In the case of electricity demand, expenditure (quantity) is a function of exogenous price, the individual's income and other household characteristics, as well as unobserved habits/preferences. In our case, both because there is no reason to believe that (unobserved) preferences are related to the unobserved factors in the demand equation (eq. (1)) and because we have many preference-related variables at the household level, selection on *observables* is a more likely scenario, in which case OLS estimation of eq. (1) has the usual properties (see e.g. (Wooldridge 2010, §19.4.1)). It is likely, in our sample, that individuals who did not present expenditure data differ from those who do on *observable* characteristics, a fact confirmed by an inspection of the summary statistics by response status (tables available upon request). Individuals who did not respond differ from those who do in terms of having lower income and home owning rates, smaller homes and larger families, lower rates of environmental membership. Thus, inclusion of these controls in the main regression is very likely sufficient to account for the effects of non-response.

	Australia	Canada	Chile	France	Israel	Japan	South Korea	Netherlands	Spain	Sweden	Switzerland	Overall
Income (euro)	50990.99	47066.9	14654.11	38811.58	25922.61	41989.08	30724.24	41048	32550.74	46587.77	69201.16	39089.53
	(2828.436)	(3158.455)	(1030.951)	(1424.585)	(1626.451)	(2193.368)	(1768)	(1302.828)	(1811.545)	(1379.915)	(3119.411)	(1077.377)
Electricity price (euro/kwh)	0.222	0.133	0.239	0.150	0.147	0.213	0.057	0.274	0.284	0.164	0.166	0.200
	(0.028)	(0.012)	(0.018)	(0.011)	(0.009)	(0.013)	(0.01)	(0.016)	(0.028)	(0.009)	(0.008)	(0.007)
Electricity expenditure (euro)	785.409	873.349	503.397	826.638	891.746	901.320	175.504	918.451	703.780	1948.849	760.370	840.299
	(63.612)	(61.578)	(41.405)	(44.595)	(68.844)	(60.222)	(33.995)	(51.662)	(38.237)	(98.488)	(84.058)	(30.077)
Electricity Consumption (khw)	4978.372	10097.7	3561.876	7167.865	7529.249	4971.784	3961.334	3816.945	3840.819	13795.36	5402.354	5625.268
	(457.588)	(1225.294)	(575.835)	(473.498)	(656.548)	(322.187)	(485.625)	(205.651)	(325.312)	(808.825)	(594.964)	(196.699)
Members in household	2.712	2.373	3.991	2.536	3.331	2.534	3.453	2.722	2.948	2.388	2.780	2.730
	(0.129)	(0.143)	(0.188)	(0.098)	(0.18)	(0.118)	(0.148)	(0.095)	(0.095)	(0.075)	(0.141)	(0.063)
Size of primary residence (sqr. mtr.)	179.760	133.524	96.329	108.705	105.208	93.875	89.893	133.211	113.129	119.448	124.269	105.781
	(13.423)	(7.4)	(4.713)	(3.906)	(5.522)	(4.672)	(4.569)	(4.964)	(5.223)	(2.991)	(5.984)	(2.542)
Age of respondent	47.504	52.144	49.240	51.732	43.833	52.344	44.658	47.096	49.458	51.950	46.877	50.791
	(1.851)	(1.321)	(1.49)	(1.026)	(1.698)	(1.277)	(3.506)	(1.004)	(1.33)	(0.93)	(1.265)	(0.673)
Years of post-secondary education	3.184	3.068	4.006	2.449	3.239	3.814	2.759	4.325	3.392	2.421	3.134	3.419
	(0.302)	(0.246)	(0.262)	(0.182)	(0.245)	(0.28)	(0.458)	(0.186)	(0.25)	(0.157)	(0.27)	(0.137)
Energy behaviour index	7.939	7.407	8.562	7.944	7.613	7.166	8.078	7.097	8.109	5.571	7.203	7.489
	(0.19)	(0.183)	(0.118)	(0.131)	(0.165)	(0.18)	(0.277)	(0.141)	(0.132)	(0.114)	(0.168)	(0.089)
Number of appliances	15.199	16.265	14.032	13.339	15.816	16.588	13.722	14.492	14.536	15.006	12.782	15.343
	(0.897)	(0.764)	(0.565)	(0.491)	(0.758)	(0.716)	(1.149)	(0.407)	(0.686)	(0.421)	(0.537)	(0.347)
Observations	85	88	153	147	104	175	53	163	118	218	106	1410

Notes: Mean (standard error) reported above using probability weights to account for survey sampling. The use of probability weights implies that the means (standard errors) reported above differ from the raw sample means (standard errors).

Table 1: Summary statistics of regression sample for continuous variables.

	Australia	Canada	Chile	France	Israel	Japan	South Korea	Netherlands	Spain	Sweden	Switzerland	Overall
Home Owner (1=Owner)	0.777	0.879	0.697	0.767	0.734	0.614	0.802	0.816	0.866	0.859	0.575	0.714
	(0.05)	(0.035)	(0.052)	(0.037)	(0.046)	(0.048)	(0.06)	(0.032)	(0.033)	(0.024)	(0.055)	(0.024)
Home Type (1=Multi-dwelling)	0.050	0.134	0.170	0.253	0.720	0.386	0.531	0.186	0.744	0.235	0.549	0.367
	(0.023)	(0.038)	(0.04)	(0.038)	(0.046)	(0.048)	(0.118)	(0.033)	(0.044)	(0.03)	(0.055)	(0.025)
Urban area (1=Urban)	0.180	0.382	0.745	0.200	0.675	0.350	0.776	0.216	0.518	0.176	0.149	0.380
	(0.042)	(0.055)	(0.053)	(0.035)	(0.049)	(0.048)	(0.068)	(0.034)	(0.05)	(0.026)	(0.039)	(0.025)
Gender (1=Male)	0.514	0.526	0.591	0.541	0.519	0.484	0.498	0.596	0.599	0.614	0.614	0.526
	(0.059)	(0.057)	(0.057)	(0.043)	(0.054)	(0.048)	(0.113)	(0.041)	(0.049)	(0.035)	(0.054)	(0.025)
Member of Envt. Org. (1=Yes)	0.194	0.140	0.273	0.123	0.163	0.072	0.372	0.161	0.120	0.161	0.288	0.131
	(0.054)	(0.039)	(0.052)	(0.029)	(0.037)	(0.034)	(0.133)	(0.03)	(0.031)	(0.026)	(0.051)	(0.019)
Energy efficient appliances (1=Yes)	0.734	0.673	0.446	0.786	0.698	0.560	0.622	0.729	0.821	0.743	0.741	0.644
	(0.051)	(0.054)	(0.055)	(0.035)	(0.048)	(0.047)	(0.114)	(0.037)	(0.036)	(0.031)	(0.05)	(0.024)
Space heat/cool (1=Yes)	0.748	0.544	0.566	0.486	1.000	0.872	0.288	0.378	0.639	0.421	0.148	0.678
	(0.055)	(0.056)	(0.054)	(0.043)	(0)	(0.028)	(0.079)	(0.04)	(0.049)	(0.035)	(0.04)	(0.021)

Notes: (i) Mean (standard error) reported above, using probability weights to account for survey sampling. The use of probability weights implies that the means (standard errors) reported above differ from the raw sample means (standard errors).
(ii) Mean value for categorical variables (i.e. all variables above) are interpreted as the proportion of “1s”.
(iii) “Energy efficient appliances” is an indicator variable for having access to a top-rated energy efficient appliance; “space heat” is an indicator for having electricity as the main space heating (cooling) source.

Table 2: Summary statistics of regression sample for categorical variables.

country price. From fig. 1, it is evident that there is sufficient within-country price variation (except for South Korea) to assuage such concerns. This is an important point for our estimation, since identification of country-specific elasticities is based on within-country price variation (see section 2, page 10).

Other key control variables in our regression include household characteristics such as household size, home size, age of respondent, years of post-secondary education, urban location, behavioral attributes such as energy behaviour index and membership in an environmental organization, appliance characteristics such as possession of at least one top-rated energy efficient device (labelled "energy efficient device"), appliance stock and an indicator for electric space cooling/heating. We summarize them (in table 1 and table 2), by country in order to highlight the cross-country variation and exhibit country-specific characteristics which motivate us, in our regressions, to allow for (observed) heterogeneity in demand parameters across countries.

Average income, at 39100€, is consistent with measures derived from national income statistics but certain countries, South Korea and Switzerland in particular, appear to have far higher income than respective country-averages. In addition to issues with exchange rate dynamics (a concern particularly for Switzerland over this period), this highlights a strong urban and high income bias for South Korea, with 77% of the respondents residing in urban areas. Chile has the lowest income in the sample countries (average of 14700€), but also suffers from a strong urban bias, with 75% of the sample located in urban areas, in addition to having higher-than-average years of education, at 4.1.

There are smaller variations in household size and years of education across countries, with Chile, Netherlands and Japan having households with more years of education of the respondent (above 3.5) and Chile, Israel and South Korea having larger family sizes (above 3.3). Israel and South Korea have the youngest responders, (average age below 45 years) while Sweden, Japan and Canada have the oldest (average age above 51). We note that the sample average age of 50 is likely marginally lower than in other surveys due to the exclusion of respondents older than 70 (65 for Korea) from the sample, as well as possibly due to an internet bias.

Home size (*size of primary residence*) varies widely across countries, with (as anticipated) South Korea, Japan (and Chile, to a lesser extent) having the smallest (less than 90 sqr. mtr.) and Australia by far the largest, at 179 sqr. mtr. Home ownership does not vary substantially, with only Japan and Switzerland having significantly lower rates than the average, of 71%. Turning next to urbanisation, we find that, apart from Chile, South Korea and Israel, there is not a substantial urban bias, with a mean urban residence of 38%; Australia is the least urban, at 18%. We categorize homes into two types focused on in the literature, flats or "multi-dwellings" and (semi-)detached houses, and find that there is substantial variation across countries along this dimension. Australia has the smallest proportion of households in flats (5%) and Spain the largest (74%).

Finally, membership in environmental organizations, postulated to be a key preference attribute impacting energy demand (see e.g. Kotchen & Moore (2008)), varies across countries, from very little participation in Japan (7%) to more than a quarter of the sample in Switzerland (28%). Overall, we find that many of the key variables such as price, consumption, income, membership and size of home, vary widely across countries and mostly as anticipated[16].

We turn, next, to summarizing our prior expectations, based on existing literature, regarding the effect of control variables on energy demand. Income, members in household and size of residence (in sq. mtrs) are typically thought of as positively related to electricity consumption, with price negatively related. Recent literature postulates electric space heating (cooling) and age to be positively related, the latter operating possibly through the channel of energy efficiency and habits, while urban residence, residence in apartments ("multi-dwelling") and home ownership are posited to negatively influence electricity de-

[16]This is not always true; e.g. the very low values of energy behaviour index–lowest of all countries–for Sweden, at 5.5, are unanticipated, see page 19.

mand.

These factors essentially reflect the spatial and incentive structures, since typically, urban residences have higher density and lower area while home ownership alters incentives for investment in energy efficient appliances, the so-called "split incentives" issue (see Davis (2012) for details and references). It is important to emphasize that some of the these indicators act through similar channels, and it is sometimes difficult to interpret these variables in isolation[17].

There has been a recent emphasis on the importance of non-price-related "attitudinal" characteristics determining energy demand (e.g. Martinsson et al. (2011), Wilson & Dowlatabadi (2007)); all of these essentially involve some form of inertia ("habits") or "lock-in" and posit different behavior depending on attitude towards the environment (see also Maréchal (2010)). To account for these effects, the survey provides at least two indicators[18]: membership in an environmental organization, and an energy behavior index. The former is an indicator for unobserved differences in behavior related to the environment and, presumably, to energy, while the latter is a more direct, self-reported measure of certain measures taken to save energy. This index is defined to lie between 1 and 10, with higher values indicating greater tendency to save energy.

Energy behaviour index has comparatively lower variability than many other variables, with only Spain, Chile and South Korea having high scores, above 8; somewhat surprisingly, Sweden has the lowest score by far, at 5.7, the only country below 6. Both variables are anticipated to lead to reductions in total electricity usage, conditional on other factors.

[17] For instance, higher income households tend to have larger homes and to be more urban; thus, many of these variables can sometimes represent very similar effects. However, dropping certain variables which had identical channels of effect did not appreciably alter results, and therefore, we do not refer any further to this issue.

[18] There are many other indicators which can be used to indicate different attitudes to environmental issue, both general and specific, such as "importance of environmental issues", "importance of climate change" etc. None of these indicators, however, are directly related to the issue of energy; moreover, revealed measures, such as membership or energy behaviour index, are less ambiguous indicators of individual preferences than those based on (more vague) "importance" measures.

4 Results

We begin with discussing the results from our preferred specification, the double-log with country-specific coefficients on price and all other coefficients restricted to be identical across countries, and subsequently turn to assessing the robustness of the results obtained to a variety of factors and data configurations.

4.1 Preferred Specification

The regression results from our preferred specification are reported in table 3 and table 4. For models with probability weights, the covariance matrix estimated is the standard Huber-White sandwich matrix while for the model without weights, we estimate the more robust bootstrap standard errors, with independent resampling within countries. We note that both approaches account for heteroscedasticity across countries.

In all tables below, the "Model Significance Test" reported is a (joint) Wald test of all coefficients except the country-specific fixed-effects; the test statistic is either χ^2- or $F-$distributed for unweighted and weighted models respectively and, in all cases we report, is significant. The "Model Comparison Test statistic" reported is a (joint) Wald test of equality of all coefficients, i.e. the test indicated in eq. (3); as for the Model Significance Test, the test statistic is either χ^2- or $F-$distributed and, in all cases here, is significant.

The test results for individual coefficients indicated that, except for price and home ownership, all other coefficients, in particular income, were identical across countries. For brevity, we report results from a model in which only the price coefficient was allowed to vary over countries, noting that allowing home ownership to vary across countries does not substantively alter the results on price and other coefficients. The column reading "Exogenous Only", in all tables below, provides results from regressions restricted to only countries in which increasing-block-pricing is not applied for any region/utility (i.e. excludes Australia, Israel, Japan and South Korea). Finally, all regressions, except those columns labelled “unweighted” (in tables table 3 and table 6) use probability weights to account for survey sampling. The unweighted regressions essentially are used to highlight differences across coefficients for key variables when sampling does make a difference.

We turn next to discussing the results and begin with by noting that in double- (or semi-) logarithmic equations, the coefficient on dummy variables can be (approximately) interpreted as a proportional effect[19]. We find, in most specifications (from table 3, table 6 and table A.1) that the coefficients on these control variables accord with both intuition and prior results, whenever significant. Household size and home size (not shown) are positively associated with energy demand, and so are the number of appliances. Urban residence is negatively associated with demand, with a coefficient which varies in magnitude depending upon the specification. In unweighted specifications, using the full sample, the effect is quite strong, about $9-11\%$, and significant; restricted to exogenous-only countries, the effect is a more modest 7% and not significant. Overall, it is clear that the impact of urban residence varies across countries and that larger sample sizes are possibly necessary before any definitive assessments can be made.

Age is positively associated, and significant, while education is never significant. Home ownership is never significant, for all samples and specifications, and we speculate that it is due to the inclusion of the channel through which it operates (energy efficient device(s)). It is curious that the dummy for presence of at least one top-rated energy efficient appliance, while of the correct sign in most specifications, is never significant. This result is similar to that in Alberini et al. (2011, table 9).

A surprisingly strong result is that regarding the difference between apartments and

[19]To see this, let D be a dummy variable for urbanisation (for instance), and the effect of urbanisation on electricity demand as a proportion be $u = \dfrac{Q_1 - Q_o}{Q_1}$ where $_1$ and $_0$ indicate respectively urban and non-urban data points, and c be the estimated coefficient on D in eq. (1). Then, as indicated in Halvorsen & Palmquist (1980), $c \neq u$, in general; to be more precise, $u = \exp(c) - 1$. In our case, the difference between c and u (computed from the regression results), for all dummy variables included in the regression, were found to be very small, and we therefore discuss and interpret c (the estimated coefficient) as if it were u (the proportional effect of interest).

	Full Sample		Exogenous only Sample	
	weighted	**unweighted**	**weighted**	**unweighted**
Income elasticity	0.0669	0.0941**	0.151**	0.131**
	(0.0528)	(0.0392)	(0.0614)	(0.0531)
house hold size (number of members)	0.0949***	0.107***	0.0683***	0.0930***
	(0.0283)	(0.0180)	(0.0241)	(0.0203)
Home Ownership (1=Yes)	-0.0103	0.0579	0.121	0.0326
	(0.0959)	(0.0588)	(0.0849)	(0.0691)
Home Type(1=Multi-dwelling)	-0.284***	-0.352***	-0.220**	-0.386***
	(0.104)	(0.0595)	(0.0869)	(0.0769)
Urban Area (1=Urban)	-0.233***	-0.0921*	-0.126*	-0.0754
	(0.0846)	(0.0511)	(0.0717)	(0.0596)
Age of respondent	0.00853***	0.00923***	0.00926***	0.00897***
	(0.00264)	(0.00181)	(0.00241)	(0.00190)
Gender(1=Male)	0.0595	-0.0163	0.0535	0.00581
	(0.0652)	(0.0438)	(0.0569)	(0.0459)
Years of Post-secondary Education	0.0142	-0.000962	-0.00518	-0.0120
	(0.0163)	(0.00955)	(0.0123)	(0.0111)
Member of Envt. Organization (1=Yes)	-0.123	0.0317	0.0587	0.0498
	(0.170)	(0.0559)	(0.0577)	(0.0579)
Energy Behaviour Index	-0.0285	-0.0390***	-0.0247	-0.0345***
	(0.0249)	(0.0117)	(0.0168)	(0.0131)
Own an Energy Efficient Appliance (1=Yes)	-0.00325	-0.0361	0.0324	-0.00390
	(0.0640)	(0.0425)	(0.0682)	(0.0510)
Number of Appliances	0.0178**	0.0150***	0.00994*	0.0153***
	(0.00740)	(0.00418)	(0.00569)	(0.00472)
Space Heat/Cool Elect.(1=Yes)	0.221***	0.326***	0.394***	0.392***
	(0.0619)	(0.0446)	(0.0549)	(0.0469)
Observations	1410	1410	993	993
R-squared	0.474	0.483	0.395	0.488
Model Significance Test	16.94	623.3	14.97	509.2
Model Comparison Test Statistic	3.62	169		

***** p<0.01, ** p<0.05, * p<0.1**

Notes: (i) Standard errors for (a) weighted regressions are the Huber-White sandwich estimators (b) the unweighted regressions are based on the bootstrap (with countries independently resampled).
(ii) Columns labelled "Exogenous-only" refer to the sample obtained when countries with increasing-block-pricing schemes(Australia, Israel, Japan and South Korea) are excluded.

Table 3: Regression results for double log specification in eq. (1)

	Full data				Exogenous only			
	Trans-log		Double Log		Trans-log		Double Log	
Country	**Elasticity**	**S.E**	**Elasticity**	**S.E**	**Elasticity**	**S.E**	**Elasticity**	**S.E**
Overall	**-0.7456**	**0.0466**			**-0.7537**	**0.05657**		
Australia	-0.7592	0.0569	-1,4338	0,3308				
Canada	-0.6441	0.0518	-1.1089	0.1882	-0.6817	0.0728	-1.1079	0.1873
Chile	-0.8609	0.0708	-1.0966	0.0923	-0.8807	0.0730	-1.0921	0.0968
France	-0.6942	0.0496	-0.8518	0.1419	-0.7133	0.0639	-0.8131	0.1473
Israel	-0.7245	0.0463	-0.8876	0.1178				
Japan	-0.7857	0.0543	-0.5002	0.1171				
South Korea	-0.3936	0.0745	-0.2623	0.084				
Netherlands	-0.8232	0.0582	-0.2455	0.0859	-0.7784	0.0632	-0.2419	0.0863
Spain	-0.8372	0.0534	-1.0441	0.1326	-0.8045	0.0558	-1.0304	0.1291
Sweden	-0.7146	0.0555	-0.6893	0.1737	-0.7120	0.0678	-0.6788	0.1701
Switzerland	-0.6977	0.0710	-0.5363	0.1917	-0.6774	0.0860	-0.4801	0.2061
R-squared	0.45		0.483		0.4548		0.488	
N	1410		1410		993		993	

Price elasticities for the double-log (in eq. (1)) and trans-log (in eq. (4)) specifications; all regressions use probability weights and country fixed-effects. For the translog model, the country-specific price elasticities are obtained by estimating the elasticity at the country-specific mean , while the **Overall** coefficient is obtained by estimating at the grand mean. Standard error computations for the double-log specification are as in table 3. For the trans-log specification, standard error computations are based on the delta method. For a definition of the "Exogenous-only" sample, refer to table 3.

Table 4: Country-specific price elasticities

(semi-) detached housing; across all specifications and all data configurations, residence in apartments has a very strong negative effect on electricity demand, varying between 22% and 38%, and always significant. There are several factors that may explain this finding. First of all, the area needed to be cooled/heated/lit is, in general, smaller (in this sample, at least).[20] Second, there will typically be less number of windows and walls in an apartment compared to a detached house.

Finally, in common with Alberini et al. (2011), we find that having electric heating (and/or cooling) makes a substantial impact on electricity demand; across specifications, it varies between 22 and 41%, and is always significant.

We turn now to understanding the impact of attitudinal determinants of energy demand, membership in an environmental organization and energy behaviour index. From the results in table 3, table 6, table A.1 and table A.2, we see that energy behaviour index is indeed negatively associated with energy demand across all specifications. Significance however, is restricted either to only the "exogenous only" countries in the main specification, with unweighted (i.e. raw) samples or with the trans-log specification. Essentially, significance depends upon assumptions regarding the population, and is sensitive to data configurations. Overall, however, there appears a significant negative effect, varying from 2 – 4%, on electricity demand. On the other hand, membership in an environmental organization is not, in any specification, significantly associated with electricity demand and is often of the incorrect sign, a somewhat surprising result, and one which (for our sample, at least) contradicts the results in Kotchen & Moore (2008), who report a strong negative effect on demand of environmental membership.

We note that income elasticity is rather low and varies widely across specifications, between 0.08 and 0.14. Typically, when weighting is resorted to, income turns insignificant, but is significant in all specifications with unweighted data. The magnitude and variability

[20] We estimate the average living area per person to about 48 m^2 across countries. For Sweden this number is about 54m^2 , which can be compared to the official statistics of 56 m^2 per person `http://www.scb.se/statistik/HE/HE0103/2006A03/HE0103_2006A03_SM_BO23SM0801.pdf`.

(depending upon specification) of income elasticity are consistent with those in, for instance, Fell et al. (2011) and Reiss & White (2005), as well as with much of the existing literature. A possible explanation for the lack of significance of income elasticity when sampling weights are used is the oversampling of relatively higher income individuals in many countries leading to higher than anticipated average country income, a fact already remarked upon in section 3.2.

Our key result (column 4 in table 4) is a unitary price elasticity for four countries, Australia, Canada, Chile and Spain and very high, but less-than-unitary elasticity for France and Israel. Only in the case of Korea and the Netherlands is the elasticity below 0.5. Price elasticity estimates as high[21] as this are somewhat beyond the typical long-run elasticities reported in the literature (which are much lower), about 0.5 or below. Only Fell et al. (2011) and Alberini et al. (2011), using average prices, document very high price elasticity, unitary in the former and between 0.6 and 0.87 in the latter. Borenstein (2009) and Ito (2012), using average prices, find much smaller elasticities, less than 0.3, using a sample restricted to only (parts of) California. However, some of the older literature, beginning with Halvorsen (1975) and summarized in Bohi & Zimmerman (1984, table 1, pp 116-118), find a substantially higher price elasticity, closer to the ones obtained in this study.

4.2 Robustness Checks

We provide evidence to the effect that our results are robust to a variety of factors which can either confound the estimation or lead to biased estimates. In particular, we consider four types of issues (i) endogeneity of marginal price (ii) functional form chosen (iii) method of accommodating for heterogeneity (in estimating country-specific coefficients) and (iv) of outliers and extreme values in the sample.

[21] In the discussion of results here, "high" and "low" price elasticity refers to the absolute value of elasticity, consistent with both intuition and usage in the literature. In the same spirit, we also do not append a $-$ sign to price elasticities, to simplify understanding of 'high' and 'low'.

	Regional FE		Trimmed		Winsorized	
Country	**Elasticity**	**S.E**	**Elasticity**	**S.E**	**Elasticity**	**S.E**
Australia	-1.5266	0.4121	-0.7471	0.504	-1.2557	0.2706
Canada	-1.2727	0.2473	-1.1373	0.2718	-1.0257	0.135
Chile	-1.1397	0.1476	-1.3091	0.2584	-1.0244	0.0817
France	-1.0045	0.1394	-1.2428	0.2551	-0.9293	0.1544
Israel	-0.9589	0.1392	-1.0385	0.1549	-0.9701	0.1257
Japan	-0.5051	0.107	-0.8896	0.2087	-0.4748	0.0815
South Korea	-0.181	0.0651	-0.9275	0.6036	-0.803	0.0486
Netherlands	-0.2884	0.1117	0.0386	0.0734	-0.1766	0.1008
Spain	-1.0852	0.1295	-0.8821	0.1975	-1.0816	0.1723
Sweden	-0.7447	0.2618	-0.6734	0.2159	-0.6981	0.1442
Switzerland	-0.6703	0.2701	-0.7161	0.3017	-0.6069	0.1729

Notes: Country-specific price elasticity, computed for various data configurations, are shown. The regression specification, and standard error computation, is identical to that in table 3. The only differences are in fixed effects used (the model labelled "Regional FE" replaces country-specific fixed effects with country-region specific fixed-effects), sample size (model labelled "trimmed" reports results from a specification in which the upper and lower 5% of observations on price are dropped) or sample composition (the model labelled "winsorized" reports results from a specification in which price and quantity (expenditure) observations beyond the upper and lower 5% are replaced with the values at the 95^{th} and 5^{th} percentile, respectively).

Table 5: Robustness checks for price elasticity

	Full Sample: Pooled		Exogenous only Sample: Pooled	
	weighted	unweighted	weighted	unweighted
Price elasticity	-0.9902*	-1.1017***	-1.2468**	-1.1796**
	(0.0542)	(0.0368)	(0.0516)	(0.0517)
Income elasticity	0.0715	0.104**	0.145**	0.144***
	(0.0734)	(0.0422)	(0.0602)	(0.0540)
house hold size (number of members)	0.102***	0.106***	0.0690***	0.0979***
	(0.0302)	(0.0196)	(0.0234)	(0.0225)
Home Ownership (1=Yes)	-0.00738	0.0514	0.0693	0.00114
	(0.110)	(0.0657)	(0.0766)	(0.0691)
Home Type(1=Multi-dwelling)	-0.227**	-0.324***	-0.226***	-0.378***
	(0.114)	(0.0663)	(0.0866)	(0.0749)
Urban Area (1=Urban)	-0.279***	-0.113**	-0.118	-0.0691
	(0.0922)	(0.0566)	(0.0719)	(0.0618)
Age of respondent	0.00851***	0.00969***	0.00959***	0.00918***
	(0.00279)	(0.00199)	(0.00248)	(0.00214)
Gender(1=Male)	0.0721	-0.0133	0.0505	0.00267
	(0.0705)	(0.0445)	(0.0577)	(0.0467)
Years of Post-secondary Education	0.0243	0.00650	-0.000560	-0.00775
	(0.0191)	(0.00992)	(0.0116)	(0.0104)
Member of Envt. Organization (1=Yes)	-0.0840	0.00446	-0.000561	0.00653
	(0.176)	(0.0576)	(0.0625)	(0.0600)
Energy Behaviour Index	-0.0343	-0.0392***	-0.0295*	-0.0369***
	(0.0259)	(0.0126)	(0.0170)	(0.0135)
Own an Energy Efficient Appliance (1=Yes)	-0.114	-0.0557	0.0353	-0.00938
	(0.0832)	(0.0445)	(0.0635)	(0.0564)
Number of Appliances	0.0229**	0.0172***	0.0131**	0.0180***
	(0.00907)	(0.00417)	(0.00607)	(0.00485)
Space Heat/Cool Elect.(1=Yes)	0.292***	0.363***	0.410***	0.402***
	(0.0676)	(0.0488)	(0.0562)	(0.0503)
Observations	1410	1410	993	993
R-squared	0.419	0.437	0.361	0.445
Model Significance Test	17.45	411.0	17.02	354.1

Robust standard errors in parentheses
***** p<0.01, ** p<0.05, * p<0.1**

Notes: Standard error computation is as for the double-log specification in table 3. For a definition of the "Exogenous-only" sample, refer to table 3.

Table 6: Regression results for the “pooled” model.

Turning first to endogeneity of price (see footnote 4), we note that when four countries with increasing-block-pricing are excluded from the regression sample (column 8 in table 4), the three remaining countries (Canada, Chile and Spain) nonetheless have almost identical (unitary) elasticities as before[22]. In addition, price elasticity of the other four countries do not appreciable change (except for a minor reduction for Switzerland).

An alternative way to see this is using the pooled model, where all coefficients (except the intercept) are assumed identical across countries, and country-specific fixed effects are used to allow intercepts to vary. In this specification (table 6, columns 2 and 4), we note two important points: first, that the pooled model yields unitary income elasticity (-0.99), with the full sample, and second, that excluding the countries with suspected endogeneity from the regressions leads to slightly *increased* price elasticity, to -1.18. In other words, the very high estimates of elasticity are not an artifact either of increasing-block-pricing induced endogeneity in a part of the sample or a result of allowing for country-specific elasticities.

Second, given the substantial variation in elasticity estimates across functional forms and in general, the somewhat unrealistic nature of the double-log form (as noted in Plourde & Ryan (1985) and Ryan & Plourde (2008), among others), we estimate a more general and encompassing trans-log functional form (which has been found to provide more reasonable price elasticity estimates than the double-log for other fuels, for instance, gasoline, in Wadud et al. (2010), Liu (2011)).

Results (column 2, table 4) clearly indicate a very high price elasticity. Two points are noteworthy (i) apart from South Korea, no country has an elasticity lower than 0.35 (ii) no country has a unitary price elasticity, although countries with unitary elasticity in the double-log form have rather high price elasticity. In fact, price elasticity for *five* countries

[22] Recall that our estimation is not equivalent to country-by-country estimation (see discussion on page 9). Nonetheless, estimates of coefficients which are unrestricted typically are not altered significantly by moderate changes to the sample, such as we undertake here. It is not however, always the case, as may be observed by a comparison of the elasticity for the trans-log specification, for the Netherlands.

(Japan, Korea, the Netherlands, Sweden and Switzerland) are higher with the trans-log form. In general, the trans-log functional form yields high but plausible elasticity estimates, very similar to those from the double-log form.

These results are unaltered when countries with suspected price endogeneity (as a result of increasing-block-pricing; see footnote 4) are excluded (column 6). We note that, for all model specifications for the translog, we are able to reject the following null, via a Wald test: $\beta_1 = \gamma_1 = \delta = 0$, which is essentially a test of the double-log versus the trans-log functional form[23]. We do not emphasize the model selection aspect here, since both models provide results which are qualitatively consistent and quantitatively comparable.

Third, we note that there is likely substantial regional variation, in both electricity pricing (e.g. in Australia and Canada), and important policies (e.g. climate policies at the regional level in Canada), within countries. The more robust approach to dealing with these would be to use country-region-specific fixed effects in models with country-specific coefficients[24]. In our base model (table 3 and table 4), we do not take such an approach due to the substantial increase in the number of parameters estimated, relative to the sample size (152 regional fixed effects, as opposed to 11 country-fixed effects).

Nonetheless, to check if accounting for the substantial regional differences alter the magnitude of the estimated price elasticity, we estimate the same model as the base (eq. (Unrestricted)) with country-region-specific fixed-effects instead of country-fixed effects. We find virtually no qualitative difference in our results (table A.2, column labelled "Regional-FE"); in fact, the estimated price elasticities (table 5, column labelled "Regional FE") are now higher for most countries (except for South Korea, for which elasticity is moderately

[23]It is important to point out that the double-log model which is estimated (eq. (Unrestricted)) is not derived from the trans-log in eq. (4) by imposing the restrictions indicated above: $\beta_1 = \gamma_1 = \delta = 0$. This is because imposing these restrictions gives rise to a double-log model with identical coefficients on price–the "pooled" model whose estimates are reported in table 6. As a result, the R^2 for the double-log and trans-log models in table 4 are not comparable.

[24]This approach, of using country-specific region-fixed-effects, controls for factors at the sub-national level which influence consumption but are not accounted for by country-specific factors proxied by country-specific fixed effects.

reduced). The only other change is that standard errors on most coefficients are moderately higher as a result of the substantial number of additional parameters estimated (although significance levels are unaltered).

Fourth, there are many outlier or implausible values for expenditure and/or price (provide examples). In particular, our primary concerns centered around (i) wildly overstated prices (ii) substantial changes for households which report low spending and (iii) very different behaviour of high income households from those with very low income. In order to investigate if outliers substantially alter the magnitude of the estimates we obtain, we carry out two sets of regressions. In the first, we winsorized data at the 5% threshold, for price and quantity (expenditure). On repeating the analyses with these data for main specification (table A.2, column IV, labelled "winsorized"), we found no qualitative difference in our main results and only very minor quantitative differences. The two important differences were the wider variation of income elasticity (not significant) and slightly reduced price elasticity for most countries, with only Korea having a significant change, a four-fold increase to 0.8 (table 5, column labelled "winsorized").

In the second set, we dropped observations which had very high and very low prices (the upper and lower 5%) and re-estimated the main specification (table A.2, column III, labelled "trimmed"). Again, the qualitative direction of our main results are clearly unchanged (there are only minor quantitative differences for other coefficients); however, income is now significant, and with a larger elasticity, close to 0.1. Price elasticities (table 5, column labelled "trimmed") are qualitatively very similar, with mostly minor changes in magnitude; for many countries, elasticity increases (Chile, France, Japan, South Korea and, to a smaller extent, Switzerland and Israel) while Australia's is almost halved, to 0.7, and Netherlands' turns insignificant (and positive). Nonetheless, four countries still have unitary price elasticity, and our results for the main specification are more or less reproduced here.

To summarize, we find that our main results with the base specification(from section 4.1), very high (unitary for four countries) price elasticity and low and insignificant income elasticity, are robust to a variety of plausible issues–involving different data configurations, functional forms and regional heterogeneity–likely to adversely affect or weaken them.

4.3 Comparison to prior literature

We turn next to providing a brief comparison of our results to those in the existing literature. We note that for the Netherlands, France and Chile, we have been unable to obtain references to any country-specific studies and so cannot provide explicit comparisons. An explicit comparison is also complicated by the very different methodologies (time-series versus panel versus cross-section) used across studies. Nonetheless, we provide comparisons with long-run estimates of elasticity wherever possible[25]. It is also pertinent to note that most studies cited here (except for Bernard et al. (2011)) use data at some degree of aggregation, instead of household, and (except for Filippini (2011)) some measure of average price. A common problem with many studies is insufficient price variation or lack of marginal price data at the household level.

The only study for Australia using individual-level data (at the annual frequency), Fan & Hyndman (2011), found a relatively low price elasticity, between -0.36 and -0.43, using average electricity price. However, it is rather difficult to compare our estimates to those in this study due to the unrepresentative sample(drawn from only one town in Southern Australia). For Canada, our estimate, of -1.1, is far larger than the short-run coefficients in the current literature, which vary between -0.3 and -0.67 but are quite close to the long-run elasticity of about -1.3 reported in (Bernard et al. 2011, table 4) (using marginal prices invariant over the cross-section in a repeated cross-section setting for households in Quebec).

For Israel, our estimate of -0.85 is higher than the range (-0.21 to -0.59) reported

[25] Since at least Baltagi & Griffin (1984), cross-section elasticity estimates are interpreted as long-run estimates, and we follow that convention here, at least for comparing our elasticities with those in the literature.

in Beenstock et al. (1999) (using quarterly household aggregate data, and possibly average price), while for Japan, our estimates of -0.5 to -0.8 are lower than that in Nakajima (2010) (using prefecture-level data, with both consumption and price derived from aggregates), of -1.3, but comparable to those in the prior literature (Nakajima (2010, table 1)), which lie between -0.37 & -0.47.

For South Korea, estimated price elasticity varies widely between the trans-log and the double-log, at -1.04 and -0.27 respectively, possibly as a result of data collection or sample selection issues (already referred to above). Neither of our estimates accord with the estimated (long-run) elasticity of -0.5 , in the most recent study, Saad (2009) (estimated using national-level aggregate data on consumption and price), but the double-log estimate is almost identical to the one obtained for Seoul (-0.24) in Yoo et al. (2007). For Spain, our estimates, for all specifications, are quite close, between -0.9 and -1; however, existing elasticity estimates for Spain using household level data are widely varying, and range between -0.87 (Labandeira et al. (2006)) and -0.25 (Labandeira et al. (2012)). Our estimates are closer to the former which, like ours, are estimated for survey data (annual data on expenditure, with average aggregated prices derived from a variety of sources) .

For Sweden, our estimates, which vary between -0.68 and -1, are very similar to those reported in the literature using household data. Damsgaard (2003) provides varied estimates, based on type of heating, between -0.37 and -1.35, while Andersson (1997), considering only households with electric heating, finds a rather high elasticity of -1.37. Finally, for Switzerland, our estimates vary between -0.49 and -0.63, which are remarkably similar to the short-run elasticities estimated using house-hold data for a large sample, in Filippini (2011), which vary between -0.65 and -0.85 (estimated at the city level, for 22 cities, using city-specific price).

5 Conclusions

Consistent estimates of price and income elasticity are an important part of understanding the policy space for climate change mitigation in many OECD countries. There are few existing studies estimating these important parameters across countries using a consistent household dataset and methodology. This study is a first attempt at addressing the aforementioned gap.

The primary objective of our study is an estimation of price- and income-elasticity across 11 OECD countries using survey data, while accommodating observed heterogeneity through country-specific coefficients. In addition, this study is among a few studies on electricity demand to have gone beyond the double-log functional form, in order to overcome its well known deficiencies, using, in addition, the encompassing trans-log demand function.

Unlike in some of the existing literature, we find very strong price responsiveness, ranging from a low of -0.25 (for the Netherlands) to a high of -1.4 (for Australia). In general, except for the Netherlands and South Korea, we find price responsiveness to be substantial. In contrast, our estimates of income elasticity, varying between 0.07 and 0.14 (and always significant in the unweighted specifications), are quite low but consistent with those in the existing literature.

In addition, due to the unique nature of the dataset, we find that non-price related attitudinal measures are also of some importance in explaining energy demand, a theme increasingly being explored in the literature. In particular, self-reported energy saving measures undertaken by households reduce energy demand between 2 and 4%, depending upon the specification.

Our findings suggest that households appear to exhibit substantial response to average price changes. However, before these can be taken directly to policy (as in Fell et al. (2011)), at least three important issues have to be addressed. First and foremost, what prices do consumers actually respond to? Evidence that consumers respond to average prices is rather sparse and rely on very low samples in specific regions, while evidence that consumers respond to marginal price appears to contradict the rather large literature that stresses household's inability to respond in a rational manner to economic incentives. Second, obtaining estimates similar to those we do in a panel context (with its attendant benefits), would likely provide a stronger evidence base upon which policies may be designed.

An important drawback of the current study is the lack of sufficient instruments to explicitly address the issue of endogeneity of price. Presumably, future rounds of the OECD-survey will provide researchers with additional data (such as region-specific marginal prices, current and past), allowing construction of instruments to explicitly deal with the identification challenges raised by the endogeneity of average price. A feasible and possibly immediate extension of our work would be to couple two rounds of the OECD surveys (the next round begins in 2014) to accommodate (the well-documented) variability over time of price- and income-responsiveness.

Elasticities to the type estimated here play an important role in policy evaluation for country groups, where consistency and comparability of elasticities used are important. For instance, these elasticities (when re-estimated across groups of interest) can be used to investigate welfare implications of relative price changes (e.g. of increases in/imposition of carbon taxes for a cross-country setting) or in computable-general-equilibrium-based evaluation of regional policies on climate change or other energy sector policies.

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A Additional Tables

	Full Data	Exogenous Only
ln(Income)	1.563**	1.383**
	(0.704)	(0.657)
ln(Price)	-1.795***	-2.026***
	(0.566)	(0.749)
ln(Income)*ln(Price)	0.0581	0.100
	(0.0561)	(0.0717)
$ln(Income)^2$	-0.0667*	-0.0520
	(0.0343)	(0.0333)
$ln(Price)^2$	-0.120***	-0.0639
	(0.0210)	(0.0406)
household size	0.116***	0.104***
	(0.0201)	(0.0224)
Home Size (sq mtr)	0.000461	0.00143***
	(0.000570)	(0.000468)
Home Ownership (1=Yes)	0.0741	0.0463
	(0.0620)	(0.0711)
Home Type(1=Multi-dwelling)	-0.337***	-0.378***
	(0.0660)	(0.0798)
Urban Area (1=Urban)	-0.0906*	-0.0657
	(0.0516)	(0.0598)
Age of HH head	0.00891***	0.00839***
	(0.00183)	(0.00213)
Gender(1=Male)	-0.0140	0.00677
	(0.0449)	(0.0509)
Years of Post-secondary Education	0.00112	-0.0109
	(0.00955)	(0.0109)
Member of Envt. Organization (1=Yes)	0.0165	0.0383
	(0.0539)	(0.0560)
Energy Behaviour Index	-0.0394***	-0.0361***
	(0.0123)	(0.0135)
Own an Energy Efficient Appliance (1=Yes)	-0.0661	-0.0321
	(0.0435)	(0.0575)
Number of Appliances	0.0139***	0.0141***
	(0.00404)	(0.00469)
Space Heating Elect.(1=Yes)	0.344***	0.400***
	(0.0491)	(0.0502)
Observations	1 410	993
R-squared	0,533	0.622

Notes: Standard error computations are based on the huber-white sandwich estimator.
For a definition of the "Exogenous Only" sample, refer to table 3.

Table A.1: Complete regression results for the trans-log specification in eq. (4).

	Regional FE	Trimmed	Winsorized
Income elasticity	0.0648	0.0907*	0.0722
	(0.0503)	(0.0509)	(0.0454)
house hold size (number of members)	0.0900***	0.0881***	0.0876***
	(0.0288)	(0.0294)	(0.0203)
Home Ownership (1=Yes)	-0.0272	-0.0428	0.0142
	(0.104)	(0.107)	(0.0659)
Home Type(1=Multi-dwelling)	-0.267***	-0.303***	-0.204***
	(0.103)	(0.111)	(0.0648)
Urban Area (1=Urban)	-0.253***	-0.186**	-0.187***
	(0.0935)	(0.0888)	(0.0557)
Age of respondent	0.00645**	0.00824***	0.00604***
	(0.00269)	(0.00286)	(0.00188)
Gender(1=Male)	0.0433	0.0430	0.0648
	(0.0619)	(0.0672)	(0.0471)
Years of Post-secondary Education	0.00856	0.0121	0.00552
	(0.0158)	(0.0180)	(0.0112)
Member of Envt. Organization (1=Yes)	-0.190	-0.132	-0.0920
	(0.192)	(0.188)	(0.0941)
Energy Behaviour Index	-0.0366	-0.0339	-0.0319*
	(0.0266)	(0.0262)	(0.0163)
Own an Energy Efficient Appliance (1=Yes)	-0.000768	0.0218	0.0496
	(0.0636)	(0.0664)	(0.0509)
Number of Appliances	0.0199**	0.0214***	0.0153***
	(0.00803)	(0.00824)	(0.00482)
Space Heat/Cool Elect.(1=Yes)	0.239***	0.202***	0.235***
	(0.0671)	(0.0659)	(0.0490)
Observations	1,410	1,274	1,410
R-squared	0.543	0.343	0.494
Model Significance Test	21.10	22.95	18.30
***** p<0.01, ** p<0.05, * p<0.1**			

Notes: The regression specification, and standard error computation, is identical to that in table 3. The only differences are in fixed effects (the model labelled "Regional FE" replaces country-specific fixed effects with country-region specific fixed-effects), sample size (model labelled "trimmed" reports results from a specification in which the upper and lower 5% of observations on price are dropped) or sample composition (the model labelled "winsorized" reports results from a specification in which price and quantity (expenditure) observations beyond the upper and lower 5% are replaced with the values at the 95^{th} and 5^{th} percentile, respectively).

Table A.2: Complete regression results for robustness checks.

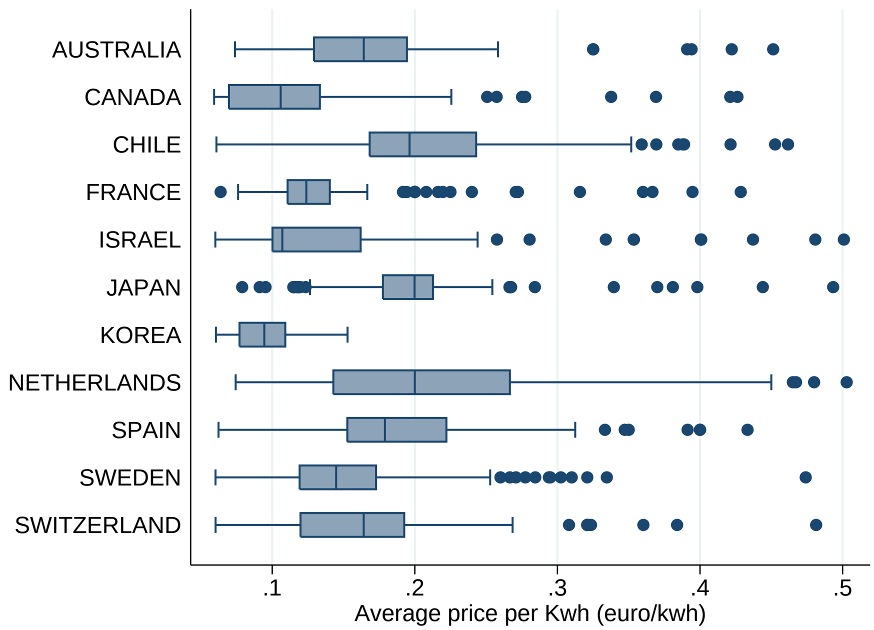

Figure 1: Within-country variation in average price per kwh

Notes: Price variable is trimmed at the top and bottom 5^{th} percentile, for ease of visualization.