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*Analysing Residential Energy Demand: An Error Correction
Demand System Approach for Ireland*

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Abstract: This paper analyses the Irish residential energy demand system by using variants of Deaton and Muellbauer's Almost Ideal Demand System model. Annual data from 1970 to 2013 is employed to estimate a demand system for solid fuels, oil, gas and electricity with the models incorporating quadratic and demographic terms to estimate long-run price and expenditure elasticities. This is the first attempt in an Irish context to estimate an energy demand system for the residential sector. Error correction models were also estimated to recover short-run elasticities. Against the backdrop of onerous climate and energy efficiency policy targets, and given the residential sector's substantial energy use it is important to update energy demand elasticity estimates to better inform policy instrument design.

Keyword(s): Energy Demand, AIDS, QUAIDS, residential energy expenditure

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1 Introduction

The curtailment and more efficient use of energy is an objective of several environmental policies driven by the fact that fossil fuel combustion is one of the main sources of the growth in global greenhouse gas emissions since 1970 (IPCC, 2014). For example, in January 2014 the EU Commission published a 2030 Climate and Energy Policy Framework that included a target reduction in greenhouse gas emissions of 40% below the 1990 level, an EU-wide binding target for renewable energy of at least 27%, and renewed ambitions for energy efficiency policies (European Commission, 2014). This framework was subsequently endorsed by the European Council in October 2014 (European Council, 2014). In recent years carbon taxes have been advocated as a mechanism to control emissions associated with fossil fuel combustion, whereas historically energy taxes were an important source of government revenues or earmarked for specific purposes like road infrastructure (OECD, 2013). The efficacy of energy taxes and especially carbon taxes depends on the magnitude of the underlying price elasticity of demand for which there is an ongoing policy need for such information.

There is a wide ranging debate in the literature on energy elasticities. Energy price elasticities vary across countries due to cultural or climatic differences though there was some consensus in the 1970s and early 1980s that own-price elasticity of energy varied in the range of -0.03 to -0.5 (Kouris, 1983; Taylor, 1977). But since then price elasticities for energy have been gradually decreasing over time and there is also evidence that energy demand is more price responsive in poorer countries (Seale Jr and Solano, 2012). Price elasticities are fuel and sector specific (Smyth, 1996; Di Cosmo and Hyland, 2013) and differ in the short and long run (Blázquez et al., 2013; IMF Staff, 2011). In addition estimated elasticity values can vary depending on estimation methods (Davis and Kilian, 2011; Menegaki, 2014). Therefore, to inform policy decisions there is an ongoing need to update estimates of energy elasticities by country, sector and fuel. This is particularly relevant in the residential sector given its high energy demand.

The residential sector represents a large share of final energy consumption, in Ireland it is approximately 25% (Howley and Holland, 2013). Nearly 40% of European final energy consumption occurs in buildings, and specifically within residential buildings, two-thirds of energy use is for space heating (European Commission, 2011). Against the backdrop of onerous climate and energy efficiency policy targets, and given the residential sector's substantial energy use it is important to gauge the sector's response to policy instru-

ments to better inform policy decisions. This paper seeks to update and further disaggregate energy elasticity estimates for the Irish residential sector. It builds on an extensive literature estimating energy elasticities for the Irish residential sector (Leser, 1962, 1964; Pratschke, 1969; O’Riordan, 1975; Murphy, 1976; McCarthy, 1977; Reilly, 1986; Conniffe and Scott, 1990; Conniffe, 2000; Scott et al., 2008; Hennessy and FitzGerald, 2011; Di Cosmo and Hyland, 2013). This paper also complements an international literature that shows wide variability in elasticity estimates across countries (e.g. Asche et al. (2008); Alberini et al. (2011)).

The current paper is the first in an Irish energy context to estimate a sector-specific multi-fuel demand system based on the almost ideal demand system (AIDS) of Deaton and Muellbauer (1980). The analysis focuses on the residential sector and estimates a series of AIDS models incorporating both quadratic expenditure and demographics terms. We also estimate error-correction models to recover short-run as well as long-run equilibrium elasticity estimates. The estimated results complement and extend elasticity estimates for the residential sector, in particular for oil and solid fuels.

The remainder of the paper is organised as follows. Section 2 reviews energy demand estimation literature. Section 3 describes the equilibrium and error corrected AIDS/QUAIDS models we estimate. Section 4 reports the empirical results, the policy implications of which are discussed in Section 5. Section 6 concludes.

2 Literature

The energy demand literature is quite heterogeneous in methodological approaches. For instance, in a review of methods used for modelling energy use in the residential sector, Swan and Ugursal (2009) reference a range of engineering and statistical/econometric techniques applied to top-down and bottom-up approaches. Econometric models include single demand equations, e.g. Haas and Schipper (1998); vector autoregressive (VAR) models, e.g. Azgun (2011); and autoregressive distributed lag (ARDL) models, e.g. Dergiades and Tsoulfidis (2008). In this paper we estimate an energy demand system with individual fuels for one sector of the economy, of which there are few examples in the literature. Labandeira et al. (2006) is a notable application, where they estimate a demand system for residential energy demand in Spain. An advantage of the demand system approach is that it can more easily estimate cross-price effects between different energy products. Labandeira et al. utilise the AIDS/QUAIDS demand models, similar to the

approach here, but their data is from household panel surveys across three years compared to aggregate household time series data used in this application, meaning that the results will not be precisely comparable. While data and research questions influence methodological approach, which in turn can affect estimated elasticity values (Menegaki, 2014), all the approaches provide insight on energy demand preferences.

Residential energy demand studies are more numerous than for other sectors. This may be due to the sector's large share of final energy demand but may also reflect the complexity of residential energy demand or the greater attention of policy makers on the sector. Some studies focus on aggregate energy demand (Haas and Schipper, 1998; Agostini et al., 1992) finding that residential energy demand across OECD countries is price inelastic, at less than -0.33. These estimates echo results for the United Kingdom, Scotland and Northern Ireland (Smyth, 1996) and the Republic of Ireland (Conniffe and Scott, 1990), though Lyons et al. (2009) more recently estimate a slightly higher elasticity at -0.53.

Empirical evidence suggests that elasticity values are fuel specific (Di Cosmo and Hyland, 2013) and accordingly many analyses of residential energy demand focus on individual fuels. Jamil and Ahmad (2011) review over twelve studies of electricity demand, roughly half of which relate to the residential sector. The empirical consensus is that electricity demand is inelastic, for the most part less than -0.5 but with estimates as low as -0.04. Asche et al. (2008) also find electricity to be highly inelastic across many European countries. However, Alberini et al. (2011) for the United States and Krishnamurthy and Kriström (2015) for 11 other OECD countries, estimate substantially higher electricity price elasticities ranging between -0.67 and -1.5. But these latter studies use survey data, whereas the lower value elasticity estimates are generally from time series analyses. In the case of gas, Asche et al. also estimate price elasticities that vary considerably across country. Regardless of whether the elasticity estimates are economy-wide or sector-specific, the estimates vary substantially by country, which highlights the need for sector, country, as well as fuel specific elasticity estimates.

Asche et al. (2008) also highlight an important 'non-result' in their analysis. Similar to Maddala et al. (1997) they find that some of their elasticity estimates have implausible signs and values. For example, finding a gas price elasticity of +0.765 for Spain and an electricity price elasticity of +0.106 for the Netherlands. These types of implausible results are less likely to be reported in the published literature but they are a feature of demand model

estimation, one which we encounter in our estimates. Asche et al. partly attribute the implausible estimates to model estimators but are unable to explain why such estimates persist in their preferred models.

Several of the previous studies on Irish energy demand are based on data from cross-section household expenditure surveys and from an elasticity estimation perspective are restricted to income (or expenditure) elasticities of fuel demand (e.g. Leser (1962, 1964); Pratschke (1969); Murphy (1976); Conniffe (2000); Scott et al. (2008)). The studies that estimate price elasticities follow a variety of methodological approaches. Some estimate residential sector demand systems covering all consumer expenditure and in those instances energy is treated as an aggregate fuel product (O’Riordan, 1975; McCarthy, 1977). Given that elasticity estimates vary by fuel type, this approach is less useful for policy purposes. Where demand systems have been estimated by fuel type (e.g. coal, gas, etc.) the demand system was for the entire economy rather than by economic sector (Reilly, 1986; Conniffe and Scott, 1990; Lyons et al., 2009). The associated elasticity estimates provide useful information for policy makers or energy companies though obviously the analysis is unable to reveal sector specific behaviours. The papers by Hennessy and FitzGerald (2011) and Di Cosmo and Hyland (2013) take a different approach, modelling the entire energy system by sector and fuel, estimating demand equations for electricity and non-electricity fuels. Their estimates do not allow for cross price effects but are the only papers to date, along with Lyons et al. (2009), that have estimated both long- and short-run energy price elasticities for Ireland.

AIDS models are probably the most popular demand system specifications estimated, reflecting the benefit of flexible functional forms and also ease of estimation. To simplify estimation Deaton and Muellbauer (1980) suggested using a linear approximation with the Stone Price Index, commonly referred to as the linear approximate almost ideal demand system (LA/AIDS). A consequence of this approximation was that several approaches to computing elasticities were proposed, some of which lead to significant errors (Green and Alston, 1990). The most common method in the literature for calculating elasticities in LA/AIDS models is the special case where expenditure shares are assumed constant.¹ Alston et al. (1994) using Monte Carlo methods show that this approach provides quite accurate estimates compared to the correct formula for elasticities in LA/AIDS models. We continue the approach in the literature assuming constant expenditure shares when calculating elasticities

¹This approach has been attributed, among others, to Chalfant (1987)

LA/AIDS models.

3 Methods

3.1 AIDS models

The AIDS model developed by Deaton and Muellbauer (1980) is a commonly used demand system specification. The popularity of the AIDS model is due in part to its flexible functional form, as it does not impose *a priori* restrictions, and theoretical restrictions can be easily imposed or tested. The model is also easy to estimate. To simplify estimation Deaton and Muellbauer (1980, p.316) initially suggested approximating the specified transcendental logarithm price index function with the linear Stone Geometric Price Index. Since then advances in computing and software mean that fitting non-linear compared to linear systems is not much more difficult and a number of extensions of the model have been developed. Banks et al. (1997) developed a generalisation that includes a quadratic expenditure term, calling their model QUAIDS. The QUAIDS model itself has also been extended to incorporate demographic variables (Ray, 1983; Blacklow et al., 2010). The AIDS model is nested within these more general models.

The AIDS model is derived from an expenditure function that belongs to the price independent generalized logarithmic (PIGLOG) class of preferences and satisfies the necessary conditions for consistent aggregation across consumers. With appropriate choice for functional forms the associated market demand equations are consistent with the behaviour of a rational representative economic agent (Deaton and Muellbauer, 1980). For utility maximising consumers the demand functions in budget share form for the AIDS model can be written as follows

$$S_{it} = \alpha_i + \sum_{j=1}^k \gamma_{ij} \log p_{jt} + \beta_i \log \left\{ \frac{m_t}{a(p)} \right\} \quad (1)$$

where S_{it} is the i^{th} budget share in year t , p_{jt} is the price of j^{th} product and m_t represents total expenditure. The function $a(p)$ is the price index function which we define later. Lower-case Greek letters represent parameters to be estimated. The AIDS budget share equation incorporating demographics is given by equation (2)

$$S_{it} = \alpha_i + \sum_{j=1}^k \gamma_{ij} \log p_{jt} + (\beta_i + \eta'_i z) \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \quad (2)$$

where z represents a vector of r demographic variables, and following Ray (1983) $m_0(z) = 1 + \rho'z$ with ρ being a vector of parameters to be estimated. Following Poi (2012) the QUAIDS budget share equation excluding and including demographic variables are given by equations (3) and (4).

$$S_{it} = \alpha_i + \sum_{j=1}^k \gamma_{ij} \log p_{jt} + \beta_i \log \left\{ \frac{m_t}{a(p)} \right\} + \frac{\lambda_i}{b(p)} \left[\log \left\{ \frac{m_t}{a(p)} \right\} \right]^2 \quad (3)$$

$$S_{it} = \alpha_i + \sum_{j=1}^k \gamma_{ij} \log p_{jt} + (\beta_i + \eta'_i z) \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} + \frac{\lambda_i}{b(p)c(p, z)} \left[\log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right]^2 \quad (4)$$

where functions $b(p)$ and $c(p, z)$ are the Cobb-Douglas price aggregators

$$b(p) = \prod_{i=1}^k p_i^{\beta_i} \text{ and } c(p, z) = \prod_{j=1}^k p_j^{\eta'_j z} \quad (5)$$

In model estimation we use each of the budget share equations (1)-(4) but specify two separate functional forms for the logarithm of the price function $a(p)$. Deaton and Muellbauer (1980) originally specified the transcendental logarithm function for the AIDS model.

$$\log(a(p)) = \alpha_0 + \sum_{i=1}^k \alpha_i \log p_i + \frac{1}{2} \sum_{i=1}^k \sum_{j=1}^k \gamma_{ij} \log p_i \log p_j \quad (6)$$

While in theory the parameter α_0 can be estimated, in practice its estimation proves difficult. Standard practice is to preassign a value, usually slightly less than the lowest value of $\log m_t$ observed in the data (Poi, 2012).

Deaton and Muellbauer also suggested approximating the translog with the Stone Price Index to overcome estimation difficulties with a non-linear price index. Moschini (1995) shows that, as the Stone Price Index is not invariant to the units of measurement of prices, it may affect the approximation

properties of the model. He suggests the loglinear analogue of the Paasche index as a possible alternative that retains the properties of the Stone index, which Moschini ascribes as the “corrected” Stone Price Index.

$$\log(a(p)) = \sum_{i=1}^k S_{it} \log \left(\frac{p_{it}}{p_i^0} \right) \quad (7)$$

where p_i^0 is the price in the base period. With modern computing capabilities the requirement to use the Stone approximation (7) is less. For example the ‘quads’ command in StataTM, which facilitates estimation of (1)-(4), only uses the translog price index (6) without allowing other price index alternatives. However, because of limited degrees of freedom in the models we estimate, especially in the error-correction variants, we specify both (6) and (7) in the models we estimate. Estimation of (1)-(4) using (7) is implemented via Stata’s ‘nlsur’ command for estimating non-linear systems of equations.

Economic theory requires aggregation, homogeneity and symmetry conditions to hold. Within the AIDS/QUAIDS models these restrictions can be either imposed or tested and imply the following:

$$\begin{aligned} \sum_{i=1}^k \alpha_i = 1, \quad \sum_{i=1}^k \beta_i = 0, \quad \sum_{i=1}^k \gamma_{ij} = 0, \quad \sum_{i=1}^k \lambda_i = 0, \quad \sum_{j=1}^k \eta_{rj} = 0, \\ \sum_{j=1}^k \gamma_{ij} = 0, \quad \gamma_{ij} = \gamma_{ji} \end{aligned} \quad (8)$$

Due to limited degrees of freedom and thus to improve the efficiency of our estimates these demand restrictions are imposed for the main model estimates. The adding-up conditions (e.g. Engle and Cournot aggregation) are imposed automatically by not estimating one of the equations in the system, whereas the remaining axioms are imposed by parameter restrictions. Negativity cannot be parametrically imposed.

3.2 Elasticities

The estimated model parameters are difficult to interpret and therefore not of direct interest. Instead the associated Marshallian price and expenditure elasticities are of policy relevance. The calculation of these elasticities differ depending on the combination of budget share equation (1)-(4) and price

index (6)-(7) selected for estimation. The formulae are reported in Tables 1 and 2 where ϵ_{ij}^M refers to the uncompensated price elasticity of fuel i with respect to changes in the price of fuel j ; θ_i is the expenditure elasticity for good i ; and δ_{ij} is the Kronecker delta. Compensated price elasticities and Allen elasticities of substitution can be recovered via the Slutsky equation as $\epsilon_{ij}^C = \epsilon_{ij}^M + \theta_i S_j$ and $\sigma_{ij} = \epsilon_{ij}^M / S_j + \eta_i$

3.3 Short- and long-run dynamics

It is widely acknowledged that many demand time series present non-stationary dynamics. The models specified above are presumed to be long-run or equilibrium energy demand relationships and to assess whether these relationships are economically meaningful or merely spurious it is necessary to investigate the time-series properties of the data used in estimation. Augmented Dickey-Fuller unit root tests are used to detect the presence of non-stationarity (Dickey and Fuller, 1981). However, it is possible to have a cointegrated relationship even though the variables of interest have different time series properties and thus, a different order of integration. Therefore, we ultimately test for long-run equilibrium cointegration relationships by using an augmented Dickey-Fuller test on the estimated model residuals. We follow the two step Engel-Granger procedure for cointegration modeling. The first stage of this method is to model the long-run relationship. The residuals from this first-stage regression are tested for stationarity. If they are stationary, a cointegrating relationship exists and we can proceed. The second stage is to estimate dynamic short-run relationships. The short-run regression includes the lagged residuals from the first step as the error correction term. The results of these short-run equations tell us the speed at which each variable adjusts to its long-run equilibrium value.

Using an error correction representation of the AIDS model is one approach for incorporating short-run dynamics (e.g. Nzuma and Sarker (2010); Eakins and Gallagher (2003); Karagiannis et al. (2000)). An error correction model (ECM) (Engle and Granger, 1987) is a restricted form of a vector autoregression (VAR), which is commonly used to examine time series dynamics. The ECM specification allows for short-run disequilibrium by assuming that the estimation error associated with the long-run demand relationship is a disturbance from the equilibrium. The ECM for the QUAIDS model with demographics (4) is specified as

$$\begin{aligned} \Delta S_{it} = & \Delta \varphi S_{it-1} + \sum_{j=1}^k \gamma_{ij} \Delta \log p_{jt} + (\beta_i + \eta'_i z) \Delta \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \\ & + \frac{\lambda_i}{b(p)c(p, z)} \left[\Delta \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right]^2 + \psi_i \mu_{it-1} + \nu_i \end{aligned} \quad (9)$$

where Δ is the difference operator, μ_{it-1} are the lagged residuals from the estimated cointegration equation (in this case equation (4)). The coefficient ψ_i measures the speed of adjustment to the long-run equilibrium following a disturbance from the equilibrium budget allocation related to fuel i in period $t - 1$. The ECM for the QUAIDS model (3) is nested within (9). The short-run elasticities can be recovered by substituting the estimated parameters from (9) into the appropriate formulae in Table (2).

3.4 Data

We estimate a demand system for the residential sector with four fuels and an aggregate non-fuel consumer good. Fuel quantity data is taken from sector energy balance sheets published by Sustainable Energy Authority of Ireland (SEAI).² Fuel prices are sourced from the ESRI database³ and supplemented with data from SEAI's Domestic Fuel Cost Archive.⁴ Solid fuel is an aggregate fuel comprising sod peat, peat briquettes and coal and its price calculated as a quantity-weighted average. Oil comprises kerosene, diesel, LPG and petroleum coke with its price also calculated as a quantity-weighted average. Data for the non-fuel aggregate consumer good is sourced from National Income and Expenditure Accounts of the Central Statistics Office (CSO) and its price is the consumer price index excluding energy products. Data on the number of households was linearly extrapolated from Census data, which occur at five year intervals. The dataset covers the years 1970 to 2013.

Figures 1 and 2 display historical residential fuel demand and prices. The composition of demand has evolved over the period with oil, gas and electricity demand increasing over the period, whereas solid fuel demand has declined since the late 1980s. Nominal fuel prices have trended upwards, though price increases were most pronounced for electricity. Electricity prices

²http://www.seai.ie/Publications/Statistics_Publications/Energy_Balance/

³http://www.esri.ie/irish_economy/databank/

⁴http://www.seai.ie/Publications/Statistics_Publications/Fuel_Cost_Comparison/

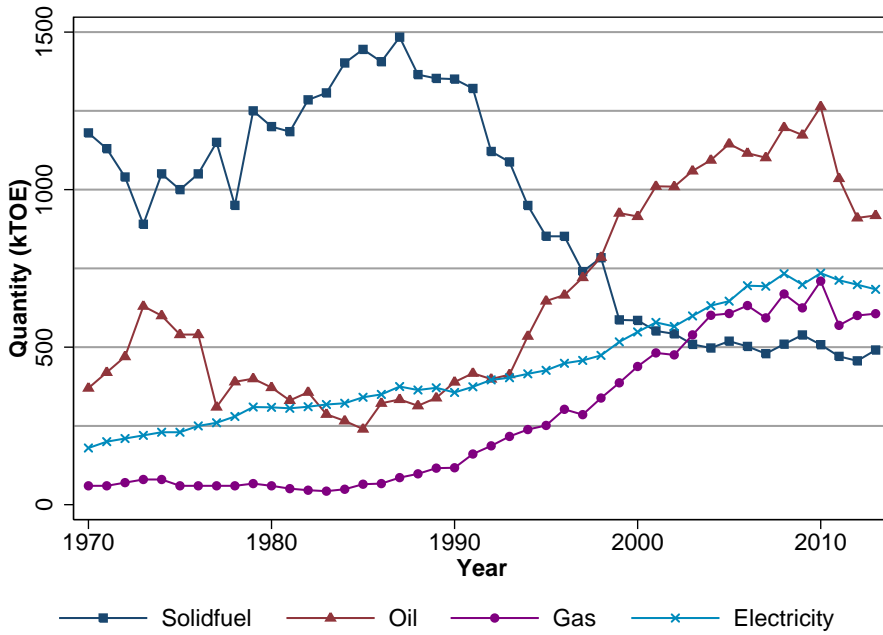


Figure 1: Quantity of Residential Fuel Consumption

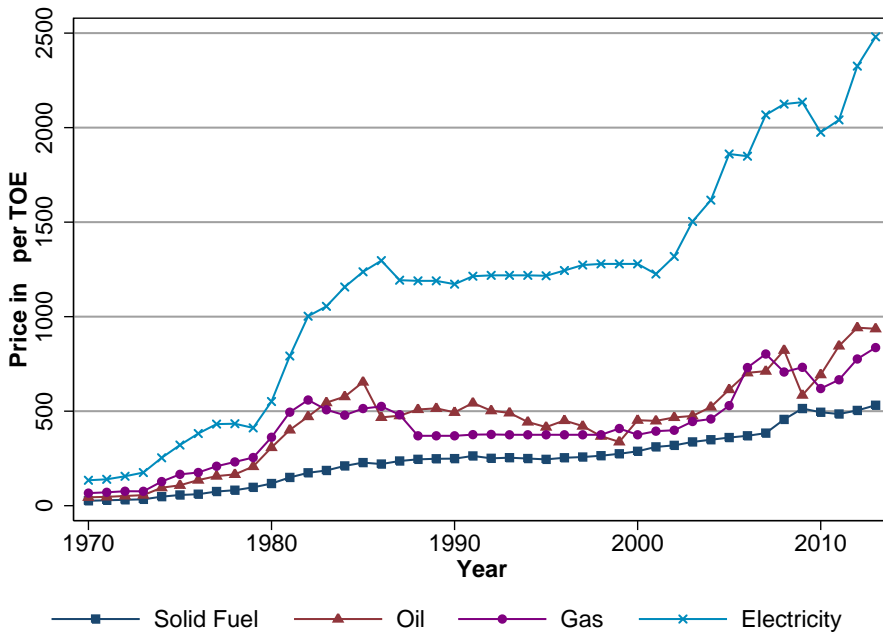


Figure 2: Fuel Price per TOE

increased quite dramatically in the early 1980s and again after 2000. By contrast, solid fuel prices increases were relatively moderate.

Table 1: AIDS model elasticities

Model	Translog Price Index	Stone Price Index
Excluding Demo-graphics	ϵ_{ij}^M $-\delta_{ij} + \frac{1}{S_i} \left[\gamma_{ij} - \beta_i \left(\alpha_j + \sum_k \gamma_{jk} \log p_k \right) \right]$	$-\delta_{ij} + \frac{1}{S_j} [\gamma_{ij} - S_j \beta_i]$
Excluding Demo-graphics	θ_i $1 + \frac{\beta_i}{S_i}$	$1 + \frac{\beta_i}{S_i}$
Including Demo-graphics	ϵ_{ij}^M $-\delta_{ij} + \frac{1}{S_i} \left[\gamma_{ij} - [\beta_i + \eta'_i z] \times \left(\alpha_j + \sum_k \gamma_{jk} \log p_k \right) \right]$	$-\delta_{ij} + \frac{1}{S_j} [\gamma_{ij} - S_j (\beta_i + \eta'_i z)]$
Including Demo-graphics	θ_i $1 + \frac{(\beta_i + \eta'_i z)}{S_i}$	$1 + \frac{(\beta_i + \eta'_i z)}{S_i}$

Table 2: QUAIDS model elasticities

Model	Translog Price Index	Stone Price Index
Excluding Demo-graphics	ϵ_{ij}^M $-\delta_{ij} + \frac{1}{S_i} \left[\gamma_{ij} - \left[\beta_i + \frac{2\lambda_i}{b(p)} \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right] \right]$ $\times \left(\alpha_j + \sum_k \gamma_{jk} \log p_k \right)$ $-\frac{\beta_i \lambda_i}{b(p)} \left[\log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right]^2 \right]$	$-\delta_{ij} + \frac{1}{S_i} \left[\gamma_{ij} - \left[\beta_i S_j + \frac{2\lambda_i S_j}{b(p)} \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right] \right]$ $-\frac{\beta_j \lambda_i}{b(p)} \left[\log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right]^2 \right]$
Excluding Demo-graphics	θ_i $1 + \frac{1}{S_i} \left(\beta_i + \frac{2\lambda_i}{b(p)} \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right)$	$1 + \frac{1}{S_i} \left(\beta_i + \frac{2\lambda_i}{b(p)} \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right)$
Including Demo-graphics	ϵ_{ij}^M $-\delta_{ij} + \frac{1}{S_i} \left[\gamma_{ij} - \left[\beta_i + \eta'_i z + \frac{2\lambda_i}{b(p)c(p,z)} \right] \right]$ $\log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \times \left(\alpha_j + \sum_k \gamma_{jk} \log p_k \right)$ $-\frac{(\beta_i + \eta'_i z) \lambda_i}{b(p)c(p,z)} \left[\log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right]^2 \right]$	$-\delta_{ij} + \frac{1}{S_i} \left[\gamma_{ij} - S_j (\beta_i + \eta'_i z) \right]$ $-S_j \frac{2\lambda_i}{b(p)c(p,z)} \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\}$ $-\frac{(\beta_j + \eta'_j z) \lambda_i}{b(p)c(p,z)} \left[\log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right]^2 \right]$
Including Demo-graphics	θ_i $1 + \frac{1}{S_i} \left(\beta_i + \eta'_i z + \frac{2\lambda_i}{b(p)c(p,z)} \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right)$	$1 + \frac{1}{S_i} \left(\beta_i + \eta'_i z + \frac{2\lambda_i}{b(p)c(p,z)} \log \left\{ \frac{m_t}{m_0(z)a(p)} \right\} \right)$

4 Results

Our estimation procedure is to first estimate variations of long-run equilibrium AIDS/QUAIDS models using the two alternative price indices and compare the outcome of the various models. We report parameter estimates in Table 4. The restrictions implied by demand theory were imposed during estimation. However, when unrestricted models are estimated we find that the symmetry axiom generally holds but not homogeneity.

4.1 Long-run models

The stationarity properties of our demand system are tested using the augmented Dickey-Fuller unit root test, with the test statistics reported in Table 3. While we generally fail to reject the existence of a unit root in our data in levels, non-stationarity is rejected for the first differences. Given the rising trend in fuel quantities and prices, as shown in Figures 1 and 2, our tests also allow for a trend. Non-stationary time series may however be cointegrated as short-run deviations between dependent and explanatory variables may converge to an underlying long-run cointegration relationship. We test for a cointegrating relationship using the residuals from the long-run equations reported in Table 4. We reject the null hypothesis of no cointegration in our demand system.

Residuals from the estimated models were examined to detect the presence of structural breaks. We inspected if the residuals were mean-reverting. In this respect the QUAIDS models incorporating demographic variables were the best fit for all four fuels. In the remaining models the residuals in the oil and gas equations exhibited a tighter mean-reverting behaviour than in the electricity and solid fuels equations. Across all the models estimated, the residuals were generally within two standard deviations, with the exceptions only occurring for at most a few periods (i.e. 1-3 years). Across the models there was not strong evidence of a structural break, and particularly so for the more flexible QUAIDS models incorporating demographic variables.

We use likelihood ratio tests to test model assumptions both on price index and model specifications. Regardless of price index specification, the hypothesis that the more flexible model specifications (e.g. QUAIDS or QUAIDS with demographics versus AIDS, etc.) are a better fit cannot be rejected ($p < 0.01$). These test results reflect the fact that when demographic variables are incorporated in the models, the estimates of the associated pa-

Table 3: Augmented Dickey Fuller stationarity and cointegration tests

Series	Level series		First differences		Cointegration
	No trend	Trend	No trend	Trend	No trend
<i>Budget shares (S_i)</i>					
Solid Fuel	-0.91	-1.56	-8.11	-8.01	-3.64
Oil	-2.32	-2.92	-7.34	-7.25	-4.48
Gas	-0.01	-1.81	-7.64	-7.91	-4.99
Electricity	-1.37	-1.64	-4.06	-4.01	-4.08
log price solid fuel	-4.48	-2.29	-3.74	-4.54	
log price oil	-2.91	-2.11	-5.27	-5.68	
log price gas	-2.69	-2.18	-4.30	-4.45	
log price electricity	-3.52	-2.20	-3.43	-3.86	
log total expenditure	-6.52	-0.36	-1.82	-3.19	
log households	2.50	0.13	-3.58	-3.91	
5% Critical values	-2.95	-3.52	-2.95	-3.52	-1.68

parameters, η_i , which govern how demographics scale the expenditure function, are each generally significant. However estimates of ρ , which determines how demographics deflate household expenditure, are insignificant in the AIDS models but significant in the QUAIDS models. The test results also reflect that the quadratic parameters λ_i are mostly significant. The implication from the tests is that the more flexible models are better suited to estimating the energy demand system. We proceed by first discussing the estimates of all the long-run models. We subsequently report on the dynamic short-run model estimates, which are confined to estimates of the more flexible QUAIDS model.

Looking at the parameter estimates, there are noticeable differences between models depending on whether the translog or Stone price indices are used. The differences occur both in parameter magnitude and sign. As the elasticity formulae differ depending on whether a translog or Stone price specification is used, we would expect the underlying parameter estimates to differ and consequently specific parameters are difficult to interpret. Comparison of the associated elasticity estimates allows us to judge whether the translog or Stone price specification makes a practical difference in terms of policy inference.

4.2 Long-run elasticities

Own-price elasticities are presented in Table 5. As expected, nearly all of the own-price elasticity estimates for fuels are negative but electricity is a notable exception. The positive and significant own-price elasticity in some of the models estimated is implausible. In a couple of the estimated QUAIDS models the elasticity is negative ranging between -0.05 and -0.23 but the estimate is not statistically significant.⁵ Previous studies have estimated a negative price elasticity ranging between -0.07 and -0.31 (Hennessy and FitzGerald, 2011; Asche et al., 2008; Di Cosmo and Hyland, 2013).

Labandeira et al. (2006) estimate a model for Spain that most closely resembles the model here and as such is potentially useful as an international benchmark, though their data is a panel across just three time periods. Their elasticity estimate is -0.79.⁶ By contrast both Blázquez et al. (2013) and Asche et al. (2008) using alternative models for Spain estimate long-run price elasticities of -0.19 and -0.31 and both estimate short-run elasticities of -0.07. So while Labandeira et al.'s results are substantially different from our Irish estimate, it is not clear if that is a real difference or due to methodological reasons.

The price elasticity estimates for solid fuels is roughly -0.27 across the models estimated but none of the estimates are statistically significant. Di Cosmo and Hyland (2013) did not find a significant price effect for solid fuels either but much earlier work by Conniffe and Scott (1990) estimated a coal price elasticity of -1.39. So solid fuels do not appear to be very price responsive. However, there are significant cross-price elasticity estimates for solid fuels with respect to a change in the price of gas or electricity of roughly 0.2 and 0.3 respectfully. Solid fuels are a substitute for gas and electricity but the cross-price elasticities are not symmetric. Cross-price elasticities for gas and electricity with respect to solid fuels are approximately 0.7 and 0.15 respectfully.

The estimates of the own-price elasticity of oil roughly averages -0.8, ranging from -0.51 to -0.95, which is broadly consistent with an estimate of -0.52 by Conniffe and Scott (1990) and a -0.73 estimate by Hennessy and FitzGerald (2011) for non-electric energy in the residential sector. This

⁵Bootstrap methods were used to calculate confidence intervals for the elasticity estimates of the models that specified a Stone price index (6). For the translog specification the 'quaid' command in Stata returned standard errors.

⁶Labandeira et al. (2006) do not report standard errors on their elasticity estimates.

contrasts with an insignificant price elasticity estimate for light fuel oil in Ireland by Asche et al. (2008).

Price elasticity estimates for gas are in the vicinity of -0.9, ranging from -0.64 to -1.11 (though in one case there is the implausible estimate of -13.04). Again these are consistent with earlier estimates by Conniffe and Scott (1990) and also similar in magnitude to estimates for Switzerland but substantially higher than those for Austria, Belgium, Denmark, Germany and the UK (Asche et al., 2008). Gas and electricity were found to be complementary fuels, which is consistent with previous Irish (Conniffe and Scott, 1990) and international literature (Akmal et al., 2001). The cross-price elasticity of gas for a change in the price of electricity is -0.5 and of electricity with respect to gas prices is -0.1. The remaining cross-price elasticity estimates were statistically insignificant.

Expenditure elasticities are reported in Table 6. The estimates for solid fuels are negative and range from -0.35 to -1.74. These estimates are consistent with Conniffe (2000) who estimate expenditure elasticities for coal ranging from -0.43 to -0.56 for the residential sector. While at first appearing unusual, negative expenditure elasticities for solid fuels can be explained by considering coal and peat as traditional fuels that consumers substitute away from as their consumption set expands. The estimated expenditure elasticity for oil range from 0.64 to 1.95 and for gas range from 0.92 to 1.98. Estimates for both fuels are broadly consistent with previous estimates by Conniffe (2000) and Labandeira et al. (2006) for Spain. Expenditure elasticity estimates for electricity vary between 0.29 to 0.84 across the models estimated.

4.3 Error correction models

Table 7 presents the results for the coefficient estimates of the QUAIDS error correction models. Due to limited degrees of freedom the error correction models are estimated using only the Stone price index (7). The error correction coefficient, ψ_i , captures the speed at which fuel demand adjusts to its long-run equilibrium. In one model the estimates of ψ_i imply that roughly up to one third of the adjustment to long-run equilibrium occurs in one year across the four fuels. In the QUAIDS with demographics ECM model adjustment in oil and gas demand was estimated to occur much faster, with between 60-80% adjustment within one year. This latter estimate compares favourably to a 72% demand adjustment rate for a aggregate fuel product in

Table 4: Parameter Estimates for long-run AIDS & QUAIDS models

Model:	AIDS	AIDS	AIDS with demo	AIDS with demo	QUAIDS	QUAIDS	QUAIDS with demo	QUAIDS with demo
Price Index:	Translog	Stone	Translog	Stone	Translog	Stone	Translog	Stone
<i>Coefficient</i>								
α_1	-0.030*** (0.005)	0.252*** (0.019)	-0.025*** (0.005)	0.133* (0.069)	-0.033*** (0.005)	0.507* (0.269)	-0.017*** (0.004)	0.336* (0.197)
α_2	-0.002 (0.003)	0.065*** (0.018)	0.003 (0.003)	-0.053 (0.077)	0.008** (0.004)	0.133 (0.208)	0.001 (0.002)	0.112 (0.253)
α_3	0.007*** (0.002)	-0.034*** (0.006)	0.002 (0.001)	0.072*** (0.019)	0.007*** (0.002)	0.481*** (0.078)	0.002 (0.001)	0.193*** (0.050)
α_4	-0.051*** (0.004)	0.042*** (0.011)	-0.046*** (0.003)	0.069 (0.047)	-0.057*** (0.004)	-0.186 (0.202)	-0.037*** (0.003)	0.158 (0.212)
β_1	-0.012*** (0.001)	-0.012*** (0.001)	0.064* (0.037)	0.002 (0.007)	-0.018*** (0.004)	-0.033 (0.022)	0.764*** (0.130)	-0.021 (0.017)
β_2	-0.003*** (0.001)	-0.003*** (0.001)	-0.180*** (0.035)	0.009 (0.008)	0.013*** (0.004)	-0.008 (0.017)	-0.495*** (0.157)	0.007 (0.021)
β_3	0.002*** (0.000)	0.002*** (0.000)	-0.049*** (0.014)	-0.010*** (0.002)	0.008*** (0.001)	-0.041*** (0.007)	-0.080 (0.052)	-0.016*** (0.005)
β_4	-0.004*** (0.001)	-0.004*** (0.001)	0.150*** (0.018)	-0.006 (0.005)	-0.017*** (0.002)	0.015 (0.017)	0.471*** (0.065)	-0.024 (0.017)
γ_{11}	0.008** (0.004)	0.008** (0.004)	0.007** (0.003)	0.014*** (0.004)	0.008** (0.003)	-0.006 (0.014)	0.008*** (0.003)	0.011 (0.007)
γ_{12}	0.002 (0.002)	0.002 (0.002)	0.002 (0.002)	0.005*** (0.002)	0.002 (0.002)	0.004 (0.004)	-0.000 (0.001)	0.002 (0.002)
γ_{13}	0.003*** (0.001)	0.003*** (0.001)	0.002** (0.001)	-0.001 (0.001)	0.002** (0.001)	-0.012 (0.010)	0.002** (0.001)	-0.004 (0.002)
γ_{14}	0.003 (0.002)	0.003 (0.002)	0.002 (0.002)	0.003 (0.002)	0.003* (0.002)	0.009 (0.007)	0.002* (0.001)	0.004 (0.003)
γ_{22}	0.002 (0.002)	0.002 (0.002)	0.004*** (0.001)	0.002 (0.002)	0.003* (0.002)	0.001 (0.003)	0.005*** (0.001)	-0.005 (0.010)
γ_{23}	-0.000 (0.001)	-0.000 (0.001)	-0.000 (0.001)	-0.001*** (0.000)	-0.000 (0.001)	-0.005 (0.008)	-0.000 (0.001)	-0.005 (0.004)
γ_{24}	0.001 (0.001)	0.001 (0.001)	-0.001 (0.001)	0.002** (0.001)	-0.000 (0.001)	0.004 (0.005)	0.000 (0.001)	0.003 (0.004)
γ_{33}	0.000 (0.000)	0.000 (0.000)	-0.000 (0.001)	0.002*** (0.000)	-0.000 (0.000)	-0.019*** (0.007)	-0.000 (0.000)	-0.001 (0.001)
γ_{34}	-0.002** (0.001)	-0.002** (0.001)	0.001 (0.001)	-0.002*** (0.001)	-0.000 (0.001)	0.007 (0.009)	0.000 (0.001)	-0.000 (0.003)
γ_{44}	0.022*** (0.002)	0.022*** (0.002)	0.020*** (0.002)	0.021*** (0.002)	0.022*** (0.002)	0.018*** (0.007)	0.018*** (0.001)	0.019*** (0.002)
λ_1					-0.002 (0.002)	0.000 (0.000)	0.029*** (0.005)	0.000 (0.000)
λ_2					0.007*** (0.002)	0.000 (0.000)	-0.013** (0.006)	0.000 (0.000)
λ_3					0.002*** (0.001)	0.001*** (0.000)	-0.001 (0.002)	0.000*** (0.000)
λ_4					-0.005*** (0.001)	-0.000 (0.000)	0.014*** (0.003)	-0.000 (0.000)
η_1			-0.006** (0.003)	0.011			-0.055*** (0.009)	0.000 (0.000)
η_2			0.013*** (0.003)	-0.001** (0.000)			0.035*** (0.011)	-0.002*** (0.000)
η_3			0.004*** (0.001)	-0.001* (0.000)			0.006 (0.004)	0.000 (0.000)
η_4			-0.011*** (0.001)	0.000*** (0.000)			-0.034*** (0.005)	0.001*** (0.000)
ρ			-0.018 (0.012)	0.000 (0.000)			-0.035*** (0.002)	-0.069*** (0.000)

Standard errors in parentheses; *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$, demo = demographics

Table 5: Long-run own-price elasticities

Model:	AIDS	AIDS	AIDS	AIDS	QUAIDS	QUAIDS	QUAIDS	QUAIDS
			with	with			with	with
			demo	demo			demo	demo
Price Index:	Translog	Stone	Translog	Stone	Translog	Stone	Translog	Stone
ϵ_{11}^M : Solid Fuel	-0.28	-0.27	-0.31	0.12	-0.29	-2.26	-0.22	0.33
ϵ_{22}^M : Oil	-0.84***	-0.84 ^b	-0.60***	-0.81	-0.74***	-0.95 ^b	-0.51***	-0.97
ϵ_{33}^M : Gas	-0.87***	-0.87 ^b	-1.11***	-0.44	-1.11***	-13.04 ^b	-1.08***	-0.64 ^b
ϵ_{44}^M : Electricity	0.19*	0.19 ^b	0.05	0.27	0.18**	-0.23	-0.05	0.03

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$, ^b bootstrapped $p < 0.1$

Table 6: Long-run expenditure elasticities

Model:	AIDS	AIDS	AIDS	AIDS	QUAIDS	QUAIDS	QUAIDS	QUAIDS
			with	with			with	with
			demo	demo			demo	demo
Price Index:	Translog	Stone	Translog	Stone	Translog	Stone	Translog	Stone
θ_1 : Solid fuel	-0.14	-0.14	-0.35**	15.25	-0.12	-0.18	-1.74***	0.23
θ_2 : Oil	0.71***	0.72 ^b	1.28***	1.08	0.65***	0.64 ^b	1.79***	1.95 ^b
θ_3 : Gas	1.55***	1.55 ^b	1.98***	-4.28	1.41***	1.45 ^b	1.98***	0.92 ^b
θ_4 : Electricity	0.79***	0.79 ^b	0.57***	-6.52	0.84***	0.78 ^b	0.29***	0.51 ^b

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$, ^b bootstrapped $p < 0.1$

a nine category AIDS model for Ireland estimated by Lyons et al. (2009).

4.4 Short-run elasticities

The calculation of short-run own-price elasticities using the parameter estimates from Table 7 and elasticity formulae in Table 2 yielded implausible estimates and are not reported. Instead, Table 8 reports elasticities calculated using only ECM parameter estimates in Table 7 that are statistically significantly different from zero. The short-run own-price elasticity estimates are smaller than their long-run equivalent, which is consistent with the hypothesis that households' full response to price changes is not instantaneous but instead take time to adjust. The short-run price elasticity estimate for solid fuel is -0.12, with the estimate from the QUAIDS with demographics model not credible. The short-run own-price elasticity for oil is roughly -0.3, gas roughly -0.25 and electricity -0.2. By contrast Lyons et al. (2009) estimate a short-run own price elasticity for an aggregate fuel product of -0.53 within a nine category AIDS model estimated for Ireland covering the period 1976–2003.

Table 7: Parameter Estimates for error corrected QUAIDS modes

Model:	QUAIDS	QUAIDS with demo.	QUAIDS	QUAIDS with demo.	
Price Index:	Stone	Stone	Stone	Stone	
<i>Coefficient</i>			<i>Coefficient</i>		
α_1	-0.268*** (0.100)	-0.233*** (0.079)	λ_1	-0.005 (0.079)	-0.215*** (0.057)
α_2	-0.030 (0.126)	0.116 (0.129)	λ_2	0.072 (0.092)	0.038 (0.074)
α_3	0.015 (0.101)	0.174* (0.103)	λ_3	-0.003 (0.020)	0.021 (0.017)
α_4	0.080 (0.058)	0.109* (0.059)	λ_4	0.053 (0.037)	0.003 (0.031)
β_1	-0.016*** (0.006)	-0.076*** (0.018)	η_1		2.867*** (0.845)
β_2	-0.008 (0.007)	0.005 (0.024)	η_2		-0.273 (1.119)
β_3	0.001 (0.002)	0.008 (0.005)	η_3		-0.373 (0.236)
β_4	-0.009*** (0.003)	-0.020** (0.010)	η_4		0.493 (0.441)
γ_{11}	0.010*** (0.004)	0.016*** (0.003)	ρ		1.367*** (0.229)
γ_{12}	0.003** (0.001)	0.001 (0.001)	ψ_1	-0.223** (0.104)	-0.281** (0.111)
γ_{13}	0.002** (0.001)	-0.000 (0.001)	ψ_2	-0.310*** (0.098)	-0.614*** (0.133)
γ_{14}	-0.000 (0.001)	0.001 (0.001)	ψ_3	-0.369*** (0.096)	-0.795*** (0.129)
γ_{22}	0.008*** (0.002)	0.006*** (0.002)	ψ_4	-0.322***	-0.367***
γ_{23}	-0.000 (0.000)	-0.000 (0.000)			
γ_{24}	0.001 (0.001)	0.001* (0.001)			
γ_{33}	0.002*** (0.000)	0.003*** (0.000)			
γ_{34}	-0.001 (0.001)	-0.001 (0.001)			
γ_{44}	0.016*** (0.001)	0.015*** (0.001)			

Standard errors in parentheses; *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$, demo = demographics

Table 8: Short-run own-price elasticities

Model:	QUAIDS ECM	QUAIDS ECM with demo
Price Index:	Stone	Stone
ϵ_{11}^M : Solid fuel	-0.12	a
ϵ_{22}^M : Oil	-0.24	-0.36
ϵ_{33}^M : Gas	-0.30	-0.21
ϵ_{44}^M : Electricity	-0.18	-0.20

a: Calculated elasticity was not plausible

5 Discussion

It was clear from the existing literature that energy price elasticities are likely to be country, sector and fuel specific. Additionally, in a meta-analysis study Menegaki (2014) conclude that elasticity estimates are not independent of econometric analysis. On that basis it is likely that there will not be consensus on the magnitude of an elasticity estimate. Depending on the policy application for which the elasticity is required, a particular methodological approach or dataset may be most useful, rather than the most recent estimate.

5.1 Solid fuels

For the equilibrium models there are a number of general conclusions that can be drawn. Solid fuel is an ‘inferior’ fuel, with less demanded as total expenditure increases. While most of the models indicated a negative own price elasticity none of the estimates were statistically significant. Our solid fuel category comprises products such as coal, sod peat, and peat briquettes and potentially our implicit assumption that these goods could be treated as an homogeneous product was not reasonable. However, we did find evidence that solid fuels can be considered as substitutes for gas and electricity.

5.2 Oil

Estimates of the own-price elasticity of oil ranged between -0.51 and -0.95, so demand is moderately inelastic. The short-run response is more muted with the elasticity estimate roughly -0.3. There was a wide spread in the expenditure elasticity estimates from 0.64 to 1.95 with several of the estimates above one. The least restrictive QUAIDS models, which were the best

fit models on the basis of the likelihood ratio tests discussed earlier, yield expenditure elasticity estimates of 1.79 and 1.95. On that basis if we view oil predominantly as a heating fuel (it is also used for cooking), as levels of residential expenditure increase households will spend proportionately more on additional home comfort.

5.3 Gas

The findings for gas are surprisingly similar to those of oil. Surprising because mains gas is an urban fuel whereas oil is the predominant fuel in rural areas and previous research indicated differences in fuel preferences by urban/rural location (Conniffe, 2000). Own-price elasticity estimates ranged from -0.64 to -1.11. The least restrictive QUAIDS models yielded estimates at the two extremes so there is no guidance on which might be closer to the true value. The variation in expenditure elasticity estimates was somewhat similar, ranging from 0.92 to 1.98, though with the exception of one model all estimates exceeded one. The short-run price elasticity was roughly -0.25, quite similar to that for oil. Short-run responses to price shocks are likely to be quite small but over time have the potential to be significantly greater. Similar to oil, as expenditure grows households are likely to spend proportionality more on gas.

5.4 Electricity

The results for electricity are more difficult to understand. There was no clear conclusion across the models estimated on the long-run own price elasticity. In the models where there is a statistically significant estimate, its sign is positive. Where estimates have a negative sign they are statistically insignificant. Regardless of sign it may be reasonable to conclude that its value is close to zero. A recent estimate by Di Cosmo and Hyland (2013) is -0.07. The short-run elasticity estimates are approximately -0.2, so short-run responses to price shocks may be more dramatic. Such a finding may be reconciled with the fact that electricity is such a critical part of everyday life that households find it difficult to maintain short-run changes in behaviours associated with price changes.

5.5 Implications for climate and energy policy

Climate and energy policy ambitions are for lower but more efficient fossil energy use with the objective of reducing greenhouse gas emissions. Carbon or energy taxes are potential policy instruments to achieve such a goal but

on the basis of the estimates in this paper it is likely that achieving substantial reductions in energy (or ultimately emissions) would require quite large additional taxes on energy or carbon. Furthermore, the full effect of such a policy mechanism would take several years to be realised. The implication is that any new policy measures focusing on the residential sector should not rely solely on a price effect.

Solid fuels, whether peat or coal, are the most emissions intensive fuels. What also distinguishes solid from the other fuels in the analysis in this paper are negative expenditure elasticities. Demand for solid fuels will decline with rising incomes. However, that is not useful information for designing policy to curtail emissions in the short term. But it does point to the fact that there may be an income effect preventing households switching from solid fuels to cleaner fuels, such as gas. For instance, initial capital costs of gas using equipment may be prohibitive. If that is the case a grant scheme supporting low income, solid fuel using households to switch to alternative fuels may accelerate the long term trend away from solid fuels as expenditure (or income) increases.

6 Conclusions

This is the first attempt in an Irish context to estimate an energy demand system for individual fuels within residential sector. The methodology follows Deaton and Muellbauer's AIDS model incorporating both quadratic expenditure and demographics terms. We also estimate error-correction models to recover short-run as well as long-run equilibrium elasticity estimates. The estimated results complement and extend demand elasticity estimates for the Irish residential sector.

A brief review of the energy demand literature finds that there is little consensus on the magnitude of demand elasticities, even within countries or sectors. Different methodological approaches appear to yield widely different estimates for what is nominally the same parameter. Nonetheless, there is an ongoing need to update and inform policy decisions. The paper's contribution is that it is the first paper to estimate demand system for individual fuels in Ireland for the residential sector, complementing earlier single equation estimates.

With respect to price elasticities, we find that residential energy products are rather price inelastic. Oil and gas demand are most responsive to price

changes, whereas solid fuels and electricity are not very responsive at all. In the case of electricity that may reflect how reliant modern life is on electricity. Our treatment of solid fuels as a homogeneous product may have concealed information on price responsiveness. However, there is evidence that solid fuels are substitute fuels for gas and electricity, presumably as a space heating fuel.

The policy implications are relatively straightforward. With demand for energy products by the residential sector being quite price inelastic it is likely that any policy measures intended to reduce demand (and thereby reduce associated greenhouse gas emissions) by increasing price are likely to be relatively ineffective for marginal changes in prices. If emissions reduction is the objective, a policy ambition might be to switch residential fuel demand away from carbon intensive solid fuels. Based on negative expenditure elasticities it may be reasonable to assume that there is an income effect preventing households switching away from solid fuels. Policy schemes that address barriers to fuel-switching, such as the capital costs of conversion, may be quite successful but requires further research.

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