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Modelling OECD Industrial Energy Demand: Asymmetric Price Responses and Energy – Saving Technical Change

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ABSTRACT

The industrial sector embodies a multifaceted production process consequently modelling the 'derived demand' for energy is a complex issue; made all the more difficult by the need to capture the effect of technical progress of the capital stock. This paper is an exercise in econometric modelling of industrial energy demand using panel data for 15 OECD countries over the period 1962 – 2003 exploring the issue of energy-saving technical change and asymmetric price responses. Although difficult to determine precisely, it is tentatively concluded that the preferred specification for OECD industrial energy demand incorporates asymmetric price responses but not exogenous energy-saving technical change.

JEL Classification Numbers: *C33, Q41.*

Keywords: *OECD Industrial energy demand; Asymmetry; Energy-saving technical change; Modelling.*

Modelling OECD Industrial Energy Demand: Asymmetric Price Responses and Energy – Saving Technical Change

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1. Introduction

Given the importance of the global environmental agenda, never before has it been so important to understand the determinants of industrial energy demand in the developed world in order to assist international policy makers in their deliberations. They require sound and dependable models to support their projections of future industrial energy demand to underpin policy; for example, the allocation of emission trading permits. However, the industrial sector embodies a multifaceted production process so that modelling the 'derived demand' for energy is a complex issue; made all the more difficult because of the need to capture the effect of technical progress of the capital stock and its subsequent effect on improved energy efficiency and hence energy consumption. Consequently, an understanding of this issue is vitally important – whatever modelling approach is adopted. This paper is an exercise in econometric modelling of OECD industrial energy demand in a panel context in order to explore the relationship between energy-saving technical change and asymmetric

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price responses since, as far as is known, modelling of the industrial sector has not been undertaken in this way before. In particular, an attempt is made to determine whether an industrial energy demand model that incorporates *either* asymmetry in the price response *or* exogenous energy-saving technical change *or* both is accepted by the data in order to understand better the determination of OECD industrial aggregate energy consumption.

The sharp increases in crude oil prices during the early 1970s stimulated a significant interest in energy demand research. This interest was maintained with the further increases in crude oil prices in the late 1970s and early 1980s followed by the collapse in the mid-1980s. The effect of these changes on the real OECD industrial energy demand price is illustrated in Figure 1 along with the index of production and energy consumption.[1] It can be seen that OECD industrial energy demand was rising consistently until the early 1970s and the first crude oil price hike, but since then has fluctuated with total consumption in 2003 for the countries in the sample being below that in 1974. At first sight this would appear to suggest an asymmetric price response with the large increases in the real energy price causing a significant reduction in consumption that was not reversed as prices subsequently eased.

{Figure 1 about here}

Most of the earlier studies of industrial energy demand followed the seminal work of Berndt and Wood (1975) and concentrated on factor substitution and subsequently inter-fuel substitution models. However, these models were based on a 'strict'

[1] This refers to the 15 OECD countries used in this study. The definition and calculation of these data are given in section 3 below.

neoclassical production and cost structure (normally represented by the translog function) that were often at odds with the data and, as Waverman (1992) states, the results from such models were "based mainly on intuition and thus incorrect" (p. 23). More recently, as Table 1 illustrates, a number of studies of industrial energy demand published since 1990 have continued to employ factor substitution models but in addition a number of studies have used a single equation approach often with a constant elasticity of demand (linear in logs) function. This procedure has become standard in energy demand estimation given its simplicity, straightforward interpretation, and limited data requirements and, as noted by Pesaran et al. (1998), it generally outperforms more complex specifications across a large variety of settings.

Table 1 also illustrates that all cited studies assume that the estimated elasticities are symmetric; however, they do differ in terms of the country or countries, data frequency and period, the dynamic specification, the econometric technique used and the allowance for technical progress (or the underlying energy demand trend). For example, Hunt and Lynk (1992) estimated a cointegrating error correction model (ECM) for the UK manufacturing sector with a deterministic trend using annual data from 1952 to 1988. Hunt et al (2003a) and Dimitropoulos et al (2005) used the structural time series model (STSM) to capture a non-linear underlying trend with an autoregressive dynamic lag (ARDL) model with UK quarterly and annual data respectively. Whereas Jones (1995 and 1996), estimated a dynamic linear–logit factor substitution model using annual data but for the USA and the G-7 countries. Chang and Martinez-Chombo (2003), on the other hand, used cointegrating ECM with time varying parameters to estimate electricity demand in Mexico using annual data but made no allowance for exogenous energy-saving technical change – instead implicitly assuming all technical progress is induced through the price effects. Medlock III and Soligo (2001) utilised a log quadratic specification to estimate the relationship between energy and income for a panel of mixed OECD and non-OECD countries by allowing for non-constant income elasticity with no deterministic trend or time dummies in their preferred model. Kamerschen and Porter (2004) estimated an electricity demand relationship using an adjustment factor on USA energy prices to reflect consumer expectation of future prices and included a deterministic trend as a measure of technical progress. Casler (1997) and Dahl and Erdogan (2000) estimated factor substitution models for the USA and Turkey respectively, but neither included a time trend in their model. Not surprisingly therefore, as Table 1 illustrates, there is a fairly wide range of estimates for the long run OECD aggregate industrial energy demand price elasticity whereas, when estimated, the OECD aggregate industrial energy demand long run income elasticity is about 0.7[2].

As noted, all the studies cited in Table 1 used a symmetric price elasticity approach, but Table 1 also illustrates that there is some variation in the way technical progress, or the more general underling energy demand trend, is captured. This reflects the long-running debate about whether a deterministic trend is an appropriate specification in such circumstances. A number of studies, (including Beenstock and Willcocks, 1983; Jones, 1994) have attempted to determine whether a simple deterministic trend should or should not be included in the estimation to capture the effect of technical progress but more recently Hunt et al (2003a, 2003b) have argued that it is unrealistic to expect a simple deterministic time trend to capture technical progress *and* other important exogenous factors (such as government policies,

[2] Although it should be noted that these are predominantly UK estimates.

important changes in economic structure etc.) so that they allow the underlying energy demand trend (UEDT) to be stochastic.

In summary, there is no consensus on how to estimate industrial energy demand, in particular how the effect of technical change (and possible other important exogenous factors) is captured. Therefore the principle suggested by Hunt et al (2003a and 2003b) is followed in that any estimated 'general' model should be as flexible as possible and any restricted version is only accepted if supported by the data. Consequently, for the framework adopted here, the general model allows for both asymmetric price responses and energy saving technical progress. Nevertheless, it could be argued that asymmetry is less likely for industrial sector energy demand than for whole economy energy or oil demand; for example, the introduction of more fuel-efficient cars is unlikely to be reversed (fully) by an oil price decline. However, if an energy price rise *does* stimulate the installation of more efficient capital in the industrial sector it is arguably still unlikely to be reversed if the price rises again. Therefore, by utilising the 'general to specific' approach alluded to above it allows the data to determine whether the asymmetry is statistically important (relative to the energy saving technical change dummies) and if so whether it the responses are 'weaker' than for other sectors of the economy. Furthermore, to date as far as is known, there has been no attempt to model asymmetric price responses for the industrial sector as an alternative way to capture induced technical change.[3] Therefore in order to undertake this exercise for the OECD industrial sector the approach adopted here follows and extends the approach applied to the whole

[3] Although Chang and Martinez-Chombo, 2003 do allow the elasticities to change over time using a time varying parameter (TVP) model.

economy in a series of papers by Gately and Huntington (2002), Griffin and Schulman (2005), and Huntington (2006).

Gately and Huntington (2002) – hereafter GH – estimated aggregate energy (and oil) demand functions by allowing for asymmetric price elasticities[4] in a similar way to that originally used in the agricultural supply literature[5] and applied in the energy field, for example by Gately and Dargay (1995a, 1995b, 1997), and Dargay (1990) – but not for industrial energy demand in the OECD. However, Griffin and Schulman (2005) – hereafter GS – suggest that that price asymmetry methodology made popular by GH is merely acting as a proxy for energy-saving technical change. To explore this, GS included time dummies as a proxy for induced technical progress arguing that this better represents the underlying trend demand than the asymmetries. Huntington (2006) in response showed that if the restrictions are actually tested on the GS results then it is possible to conclude that both asymmetric price responses and the exogenous time dummies have a role to play.

The use of 'top down' econometrics in this way arguably gives valuable insights into the aggregate effects of 'macro' trends in response to changing macro economic variables. However, unlike more 'bottom-up' engineering approaches it does not capture the intricacies of various new and developing technologies for the many industrial sub-sectors. In forecasting the future, it is usually preferable therefore to combine both 'top-down and 'bottom-up' techniques; however, for the remainder of this paper the focus is on the 'top-down' aspect. Therefore the question that this paper attempts to answer is whether OECD industrial sector energy demand is best

[4] GH actually tested for asymmetric income elasticities but overall they were rejected by the data in favour of symmetric income elasticities.
[5] For example, Wolfram (1971), Houck (1977) and Traill et al (1978).

modelled by the use of 'time dummies', 'asymmetric price responses' or 'both' when modelled using panel econometric techniques.

The plan of this paper is therefore as follows: Section 2 details the methodology used in the study; Section 3 presents the data and estimation results; with Section 5 providing a summary of the findings and some general conclusions.

2. Methodology

As highlighted above the theoretical foundations for the exercise undertaken here draw principally from the work of GH, GS, and Huntington (2006).

2.1 Gately and Huntington (2002)

Following the GS notation, and applying the GH methodology based on the Koyck model, a general symmetric model is specified where industrial energy demand in natural logs (e_t) is dependent upon real industrial output in natural logs (y_t) and a distributed lag on past real industrial energy prices in natural logs ($\gamma(L)p_t$) as follows:[6]

$$e_t = f[y_t, \gamma(L)p_t] \qquad \textbf{(1)}^{7}$$

where L is the lag operator. Assuming a linear specification and that the lag distribution on prices follows a geometric lag distribution, equation (1) may be written as:

$$e_t = \alpha + \beta y_t + \frac{\lambda p_t}{1-\lambda L} + \mu_t \qquad \textbf{(2)}$$

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[6] Note that when modelling whole economy energy (and oil) demand GH and GS used per capita energy consumption and per capita income. However this is not appropriate for the industrial sector given the different shares of the total economy the industrial sector will cover in the different countries.

[7] To follow the GH model completely equation (1) should actually be $e_t = f[\beta(L)y_t, \gamma(L)p_t]$ since they initially allowed for a differential speed of adjustment for income as well as price. However, given that subsequent modelling rejected this assumption it is not included here, similar to GS.

where μ_t is the random error term assumed to be $N(0, \sigma_\mu^2)$.

Equation (2) may be transformed by the lag operator, *L,* to obtain:

$$e_t = \alpha^* + \beta(y_t - \lambda y_{t-1}) + \gamma p_t + \lambda e_{t-1} + \varepsilon_t \tag{3}$$

where $\alpha^* = \alpha(1-\lambda)$ and $\varepsilon_t = \mu_t - \lambda\mu_{t-1}$. But given that a panel of OECD countries is utilised, like GH, equation (2) is re-written in a panel context and augmented with country dummies in order to allow for a different constant for each country, *i,* (the fixed effects approach) as given by:

$$e_{it} = \alpha^* + \beta(y_{it} - \lambda y_{it-1}) + \gamma p_{it} + \lambda e_{it-1} + \delta_i D_i + \varepsilon_{it} \tag{4}$$

where $\varepsilon_{it} = \mu_{it} - \lambda\mu_{it-1}$ and δ_i represent the differential constants for the individual countries relative to the base constant α* with all other parameters assumed to be constant across countries.

Furthermore, in their initial specification, GH identified two asymmetry phenomena related to prices and income. In particular, they argue that the energy demand response to a price increase is not necessarily reversed completely by an equivalent price decrease, nor is the demand response to an increase in the maximum historical price necessarily the same as the response to a price recovery (sub-maximum increase). A similar approach was also adopted for income however, in subsequent estimation the income response for the OECD countries is found to be symmetric so that, similar to GS, only asymmetry effects of prices are explored here. Therefore, following GH the real industrial energy price is decomposed as follows:

$$p_t = p_1 + p_{\max,t} + p_{cut,t} + p_{rec,t} \quad \textbf{(5)}$$

Where:

p_1 = log of price in starting year t=1

$p_{\max,t}$ = cumulative increases in log of maximum historical prices; monotonically non-decreasing: $p_{\max,t} \geq 0$

$p_{cut,t}$ = cumulative decreases in log of prices; monotonically non-increasing: $p_{cut,t} \leq 0$

$p_{rec,t}$ = cumulative sub-maximum increases in log of prices; monotonically non-decreasing: $p_{rec,t} \geq 0$

Substituting the decomposed price variable given in equation (5) into equation (4) and combining the constants into a single constant gives:

$$e_{it} = \alpha^{**} + \beta(y_{it} - \lambda y_{it-1}) + \gamma_m p_{\max,it} + \gamma_c p_{cut,it} + \gamma_r p_{rec,it} + \lambda e_{it-1} + \delta_i D_i + \varepsilon_{it} \quad \textbf{(6)}$$

where $\alpha^{**} = \alpha(1-\lambda) + \gamma p_1$ and $\varepsilon_{it} = \mu_{it} - \lambda \mu_{it}$; which is similar to the main GS specification and may be estimated by non-linear least squares. However GH assumed that ε_{it} is *not* autocorrelated but independently and normally distributed – thus ignoring the first order Moving Average MA(1) structure that comes about from the Koyck derivation. This issue is discussed further below.

2.2 Griffin and Schulman (2005) methodology

In their paper, GS hypothesise that the asymmetry effects are only acting as a proxy for energy-saving technical progress. They posit that the theoretical problem with using price asymmetry to proxy for energy-saving technical change is that it is driven by the volatility in price.[8] Thus the model would probably explain the energy/price

[8] Arguably the GS approach misses an important point since price asymmetry can be imposed as a temporary rather than a permanent response (see for example Vande Kamp and Kaiser, 1999). If so it might be better to model this directly rather than use time dummies which are arguably difficult to

relationship in the 1970s and 1980s when the annual price was very volatile however, not thereafter when the annual price fluctuated less widely. Therefore, according to GS the GH price asymmetry produces estimates that are observationally equivalent to intercept changes that capture energy-saving technical change but cannot explain this for long periods.[9] They therefore suggest that equation (1) is augmented as follows:

$$e_t = f[y_t, \gamma(L)p_t, Z_t] \quad \textbf{(7)}$$

where, Z_t represents a technical index of energy efficiency. However, given that indices of energy-saving technical change are unobservable GS suggested the use of fixed time (or year) dummy variables to capture the energy efficiency effects in the panel estimation.[10] Therefore equation (4) is augmented with time dummies as follows:

$$e_{it} = \alpha^* + \beta(y_{it} - \lambda y_{it-1}) + \gamma p_{it} + \lambda e_{it-1} + \delta i D_i + \theta_t D_t + \varepsilon_{it} \quad \textbf{(8)}$$

where again $\varepsilon_{it} = \mu_{it} - \lambda\mu_{it-1}$ and θ_i represent the differential time dummy coefficients for each year of the sample relative to the base; which is similar to the main equation proposed by GS. However, they also estimated a combined model with asymmetry and allowing for exogenous energy-saving technical change; so augmenting the decomposed equation (6) with time dummies gives the following:

$$e_{it} = \alpha^{**} + \beta(y_{it} - \lambda y_{it-1}) + \gamma_m p_{\max,it} + \gamma_c p_{cut,it} + \gamma_r p_{rec,it} + \lambda e_{it-1} + \delta_i D_i + \theta_t D_t + \varepsilon_{it} \quad \textbf{(9)}$$

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interpret. Moreover, the fixed time effects impose the same pattern across all countries when the pattern is probably being generated by a set of socioeconomic and structural conditions that may vary across countries. However, this is beyond the scope of the present paper but is an issue that is returned to in the summary and conclusion when discussing future work.

[9] That is, the different approaches can be considered 'substitutes' for each other.

[10] The effect is a 'non-linear' underlying energy demand trend since the estimated coefficients for the time dummies are unlikely to be consistently downwards and will have an unpredictable path picking up not only the energy efficiency improvements but also the effect of other exogenous socio-economic and structural effects. Arguably, this can be thought of as a similar approach to that proposed by Hunt et al (2003a and 2003b) in a time series context where they estimate non-linear underlying energy demand trends using the structural time series model.

where again $\varepsilon_{it} = \mu_{it} - \lambda\mu_{it}$. Equation (8) and equation (9) are therefore similar to the two equations proposed by GS and may also be estimated by non-linear least squares. GS, similar to GH, presented results where they assumed that ε_{it} was *not* autocorrelated but did report in a footnote (p. 10) that they tested for this.

2.3 Huntington (2006)

In his reply to GS, Huntington (2006) argued that the choice between the three competing models, represented here by equation (6), equation (8), and equation (9), should be guided by sound statistical tests. He therefore used the standard F-test for parameter restrictions to identify the model with better explanatory power.[11] In particular he specified two testable null hypotheses:

$$H_0 : \gamma_m = \gamma_r = \gamma_c \qquad \textbf{(10)}$$

and

$$H_0 : \theta_t = 0 \qquad \textbf{(11)}$$

The restriction given by equation (10) is that imposed when moving from equation (9) (asymmetry *with* fixed time effects) to equation (8) (symmetry *with* fixed time effects); whereas, the restriction given by equation (11) is that imposed when moving from equation (9) (asymmetry *with* fixed time effects) to equation (6) (asymmetry *without* fixed time effects).

Huntington (2006) therefore allows for the two specific equations after imposing the restrictions to be formally tested and in general finds for aggregate OECD energy and

[11] The F-test is given by: $F_{(k,DF_{UR})} = [(SSR_R - SSR_{UR})/k]/[SSR_{UR}/DF_{UR}]$
Where *SSR* = the sum of squared residuals, *DF* = degrees of freedom, *UR* = unrestricted model, *R* = restricted model, and *k* = $DF_R - DF_{UR}$

oil demand that the restrictions are rejected, hence suggesting that *both* asymmetry effects *and* time dummies for exogenous energy-saving technical progress have a role to play – that is they 'complement' each other.

2.4 Estimation

The approach adopted in this study therefore is to follow the 'general to specific' testing procedure suggested by Huntington (2006) by estimating equation (6), equation (8) and equation (9) for OECD industrial energy demand using non-linear least squares over the period 1963-2003 (to allow for the lag) and testing the two restrictions accordingly.[12] However, this is further extended by utilising a non-nested test between equation (6) and equation (8).

A non-nested test is utilised to attempt to test between the asymmetry model without fixed time effects, equation (6), and the symmetry model with fixed time effects, equation (8). Since the two equations are non-nested, the traditional F-test becomes inappropriate therefore; a J-test (Davidson and McKinnon, 1993) for comparison of non-nested models is applied. The J-test in this case, is a comparison of equation (6) with equation (8), by including the estimated fitted values from equation (8) (symmetric model with fixed time effects), $\hat{e}_{it}^{II}$, in equation (6) as follows:

$$e_{it} = \alpha^{**} + \beta(y_{it} - \lambda y_{it-1}) + \gamma_m p_{\max,it} + \gamma_c p_{cut,it} + \gamma_r p_{rec,it} + \lambda e_{it-1} + \delta_i D_i + \psi \hat{e}_{it}^{II} + \varepsilon_{it} \qquad \textbf{(12)}$$

and test the null hypothesis $\psi = 0$ using the conventional t-test for ψ. Similarly,

[12] All estimation is done using EViews version 5.

a comparison of equation (8) is made with equation (6), by including the estimated fitted values from equation (6) (asymmetric model without fixed time effects), $\hat{e}_{it}^{I}$, in equation (8) as follows

$$e_{it} = \alpha^* + \beta(y_{it} - \lambda y_{it-1}) + \gamma p_{it} + \lambda e_{it-1} + \delta i D_i + \theta_t D_t + \varphi \hat{e}_{it}^{I} + \varepsilon_{it} \qquad \textbf{(13)}$$

and test the null hypothesis $\varphi = 0$ using the conventional t-test for φ.

Thus rejection of the null hypothesis in either case suggests that there is additional information from the alternative model not captured in the base model whereas acceptance suggests the opposite. Thus the J-test allows for a further assessment of whether, when modelling OECD industrial energy demand, asymmetry price elasticities just measure energy-saving technical change (better captured by time dummies) as suggested by GS or in fact they both have a role to play.

In addition, as noted above, the Koyck derivations of all three main equations result in the inclusion of a lagged dependent variable and an error term with a MA(1) process.[13] Given this, as noted by GS, the use of non-linear least squares may be subject to specification error from simultaneous equations bias and/or autocorrelated errors that are correlated to the lagged dependent variable. Therefore following GS we also conduct Hausman tests for endogeneity and tests for autocorrelated errors.[14]

[13] Ideally equations (6), (8), and (9) should be estimated with an allowance for the MA(1) error process to avoid potential specification errors; but, as far is known, is not possible with current available econometric software.
[14] The test for autocorrelated errors is a Lagrange Multiplier based test undertaken by estimating an auxiliary regression by non-linear least squares. The residuals from the original estimated equation are regressed on the set of original explanatory variables in their non-linear form (other than the first time/year dummy) plus the residuals lagged one period. The test statistics in Table 1 below being the t-statistics for the coefficient on the lagged residuals. Note this actually tests the null hypothesis of no autocorrelated errors for an AR(1) process rather than a MA(1) process given by the Koyck derivation but still gives an indication of any problems.

3. Empirical Results

3.1 Data

The panel data used in the analysis consists of 15 OECD countries (Austria, Belgium, Canada, France, Greece, Italy, Japan, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, the UK, and the USA) and covers the period from 1962 to 2003. The primary source of these data is the International Energy Agency (IEA) database *Energy Statistics of OECD Countries* available at www.ieaa.org.[15] This includes each country's aggregate industrial energy consumption in thousand tonnes of oil equivalent (ktoe) and the index of industrial output (2000=100) over the whole period 1962 – 2003. Whereas the index of industrial real energy price (2000=100) is only available from the IEA for the period 1978 – 2003. Consequently this is spliced with an index for each country derived from data in Baade (1981) calculated from different fuel price indices: [16] the real industrial gas price, the real industrial coal price, and the real industrial electricity price weighted by their fuel consumption shares. This produces industrial real aggregate energy price indices for each country in 1972 prices (1972 = 100) over the period 1962 to 1980. The two series (1962 – 1980; 1972=100) and (1978 – 2003; 2000=100) are subsequently spliced using the ratio from the overlap year 1978 to obtain the series for the whole period 1962 to 2003 at 2000 prices (2000=100). The aggregate/weighted average data for the 15 OECD countries in the sample are illustrated in Figure 1 as previously discussed and the GH (2002) real price decompositions, equation (5) are presented in Figure 2. This illustrates that prior to the first oil price crises the total price is driven by price reductions whereas from the early 1970s to the mid 1980s it is driven by price rises

[15] The 2005 version.

[16] This source was used in a similar way by Prosser (1985) to calculate real price indices for the whole economy.

above the previous maximum. However, since the early 1980s the maximum price has been constant despite variation in both the price rise and price cut components.

{Figure 2 about here}

3.2 Estimation results

The results from estimating equations (6), (8), and (9) are presented in Table 2 represented by Models I, II and III respectively. Considering the diagnostic tests first it can be seen that the Hausman test for endogeneity suggests that the least squares estimates are consistent, however, the tests for autocorrelated errors suggests that there is a problem with Model II and Model III at the 4% and 6% levels of significance respectively. The models were therefore re-estimated with an allowance for an AR(1) process in the errors but the results showed no discernable difference[17]; therefore discussion of the results focuses on the results without the adjustment.

{Table 2 about here}

Turning to the nested tests used by Huntington (2006), it can be seen that Model III is preferred to Model I since the null hypothesis of no fixed time effects is clearly rejected *and* Model II since the null hypothesis of symmetry is clearly rejected. Therefore these tests suggest that the most general Model III (asymmetry with time effects) is preferred over model I (asymmetry with no time effects) and Model II

[17] The most common way of dealing with autocorrelated errors is to use techniques such as Cochrane-Orcutt, Prais-Winsten, etc. However, EViews5 is used for all estimation which uses a non-linear regression technique for dealing with autocorrelated errors given the drawbacks of the more traditional models when there is a lagged dependent variable.

(symmetry with time effects). This suggests that asymmetry and the time dummies 'complement' each other and both should be retained.

The non-nested tests, however, clearly cannot reject the null hypothesis for both Model I and Model II suggesting that for both there is no extra statistical information available by adding the missing components from the other models; that is, there is nothing to be gained by adding the fixed time effects to an asymmetric model or decomposing the price variable in a symmetric model with fixed time effects. In other words the tests do not determine which model is 'statistically' superior, which arguably suggests that asymmetry and the time dummies are 'substitutes' for each other.

Turning to the estimated parameters it can be seen that again there is an inconsistency in the results.[18] In all three models the income variable is very significant suggesting a long run industrial income elasticity of about 0.6 for Models II and III with the time dummies but about 0.8 for Model I without the time dummies; the relativities being similar to that found by GS for the whole economy. However, the price effects are not well determined for all models; in Model II the total price is not significant at conventional levels and in Model III only the price-cut variable is significant with price-max not only insignificant but also positive. For Model I, on the other hand, the price-max and price-rec variables are negative and very significant whereas the price-cut variable is not significant at conventional levels. Moreover, the long run elasticity for price-max is less (in absolute terms) than price-rec – which is contrary to a-priori expectations. However, again this is similar to a number of results for the whole

[18] Given the range of country sizes and energy price shocks there is potential for the errors to be heteroscedastic, therefore White's heteroscedasticity-consistent standard errors are used for determining the significance levels for the estimated parameters.

economy presented in GS. Finally, a striking feature of the results is the very high (and very significant) estimated coefficient for λ, all being in excess of 0.9 suggesting that energy demand may have a unit root.[19] This is therefore tested using an array of panel unit root tests (Levin, Lin and Chu, 2002; Breitung, 2000; Im, Pesaran and Shin, 2003; Fisher-type tests using ADF and PP tests by Maddala and Wu, 1999 and Choi, 2001; and Hadri, 2000) given in Table 3 which shows that, although not totally consistent, the majority of the tests reject the presence of a unit root, in particularly the most commonly used tests by Levin, Lin and Chu (2002) and Im, Pesaran and Shin (2003). It is therefore assumed that this is not a problem despite the relatively high estimated value of λ.

{Table 3 about here}

To further consider the models it is interesting to consider the plots of the implied long run technical efficiency improvements given by the estimated coefficients of time dummies for Model II and Model III presented in Figure 3, similar to GS.[20] This shows that the estimated long run coefficients do not show a consistent downward pattern with considerable variation from year to year with both upward and downward movement which is consistent with the UEDT approach adopted by Hunt et al (2003a and 2003b) in a time series context. However, overall both models do trend downward as shown. Furthermore, the trend for Model III (asymmetry price responses) is steeper than that for Model II (symmetric price responses). This suggests that the there is *less* exogenous energy-saving technical progress when symmetry is imposed. This is somewhat surprising since intuitively it is expected

[19] We are grateful to an anonymous referee for pointing this out.
[20] Calculated by $\theta_t/(1-\lambda)$.

that the general model allowing for asymmetry price effects would capture a larger part of the induced energy-saving technical progress with the trend capturing the remainder, being less than that in a symmetrical model where more of the technical progress (and other exogenous effects) is forced to be captured by the underlying trend. This possibly suggests some model miss-specification as suggested by the problems with the determination of the price effects in the models.

{Figure 3 about here}

In summary, the above results give a mixed message. Model III is generally accepted on pure statistical grounds and gives a sensible shaped underlying trend. However, the individual price parameters are not significant and do not conform to prior expectations. Moreover, the non-nested tests are inconclusive in trying to choose between Model I and Model II and the relative slopes of the underlying trends between Model II and Model III are contrary to expectations.

4. Summary and Conclusion

This paper is an exercise in estimating a panel data model of OECD industrial energy demand for 15 countries with data covering the period 1962-2003 based on the models and procedures developed by GH, GS, and Huntington (2006). The results, discussed in the previous section, show that the results give mixed messages. In particular, unlike Huntington (2006) for the whole economy energy (and oil) demand, it is not possible to conclusively conclude that both asymmetric price responses and time dummies have a role to play; that is they are 'complements' rather than 'substitutes' as GS imply. Although the tests suggested by Huntington (2006) do support a similar conclusion for OECD industrial energy demand, the estimated price coefficients are not well determined. The coefficients are not in line with economic theory; with the coefficients on the price-max and price-rec variables being either positive and/or insignificantly different from zero leaving the price-cut variable the only significant price variable. Therefore Model III is rejected on economic grounds, hence the idea that asymmetry and time dummies are 'complements' is rejected leaving the choice between the 'substitutes': Model I and Model II.

However, the choice between Model I and Model II is not an easy one; although all price variables have the right sign the total price variable in Model II and the price-cut variable in Model I are not significant and the relative sizes of the coefficients on price-max and price-cut are not as expected.[21] For pragmatic reasons Model I is therefore chosen over Model II on the grounds that, despite the relative signs of price-max and price-rec not being as expected they are statistically significant[22] – whereas the total price term in model II is not. Although it should be stressed that this is nowhere clear cut and further research is needed to try and disentangle this complicated relationship.

Taking Model I as the preferred model suggests that the estimated long-run income elasticity of OECD industrial energy demand is 0.8. Furthermore the estimated long run elasticity of OECD industrial energy demand with respect to a price rise above the previous maximum and with respect to a price rise below the previous maximum

[21] Note that this problem of relative sizes of the coefficients on the decomposed price variable applies to some of the whole economy energy demand estimates in GS.

[22] And the fact that price-cut is not significantly different from zero is not a problem on the grounds that the response to a fall in price is expected to be less than a price rise.

are -0.5 and -0.6 respectively, whereas the estimated long run elasticity of OECD industrial energy demand with respect to a price cut is -0.3 (although this is based on the short run coefficient which is not statistically significant from zero). The estimated income elasticity is therefore very close to the previous estimates presented in Table 1 for the UK. A comparison of the estimated price elasticities is more difficult given that all the cited previous studies of industrial energy demand assumed symmetric elasticities, however, the estimated price elasticities for a price-max and price rise are generally towards the more central range of the previous estimates for the total price elasticity.

This exercise shows that when estimating energy demand models and considering the important issue of energy-saving technical progress (and other exogenous trends) a general flexible approach should initially be adopted. The chosen model should be the one that is accepted by the data while at the same time conforming to economic theory – but this should be estimated and tested rather than imposed at the outset. However, this exercise also illustrates that even then a clear favoured statistical model may not be found without the recourse to economic intuition and theory.

In conclusion, it has been shown that econometric modelling of OECD industrial energy demand is not an easy task and further research is needed before 'definitive' estimates are obtained. Nevertheless this exercise has illustrated the importance, when modelling industrial energy demand in a panel context, of using a general flexible framework allowing for asymmetric price responses and time dummies to capture the underlying energy demand trends driven by technical progress and other exogenous factors. Although the results are not conclusive they do show that assuming a specific model or imposing, rather than testing, particular assumptions can be equally misleading and wherever possible the data should be allowed to determine the model - but guided by economic intuition and theory.

However, the exercise has also exposed a number of weaknesses that need further research. A number of restrictions imposed by the GH and GS panel framework are adopted here, in particular the assumption that the slope and time coefficients are constant across a wide range of countries.[23] Given the diverse nature of the countries used in this study it could be argued that these restrictions are unrealistic.[24] Each country's share of industrial output in GDP is likely to be different and also involve different industrial structures, institutions and socioeconomic patterns. Therefore the imposition of the same pattern of underlying energy demand trend across each country via the fixed time effects appears particularly restrictive (at least without formal testing). Furthermore, although this is a panel of developed OECD countries, they are still likely to be at different sages of development. Given these factors it is not surprising that it is difficult to obtain statistically sound and economically consistent estimates with sensible elasticities that apply across *all* countries with an identical underlying energy demand trend. Future research will therefore aim to investigate these matters further, by testing pooling restrictions across countries[25] and, if as expected the pooling restrictions are generally rejected, estimate industrial demand models for each country separately, again starting with a general specification that allows for asymmetric responses but with a non-linear underlying

[23] Also the assumption of a Koyck lag structure could be questioned, despite this being quite popular in the asymmetric price response literature.
[24] Arguably this is also true of the whole economy work undertaken by GH and GS.
[25] Arguably, pooling restriction tests are also required for the whole economy models used by GH and GS.

trend to capture energy-saving technical change and other exogenous effects and test the models accordingly.

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Figure 1a: OECD Real Industrial Energy Price (2000=100)

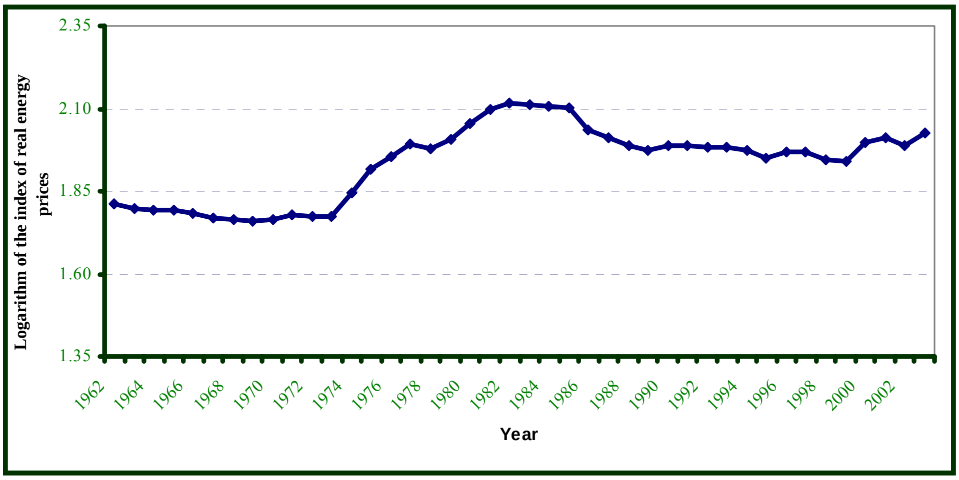


Figure 1b: OECD Industrial Energy Consumption (ktoe)

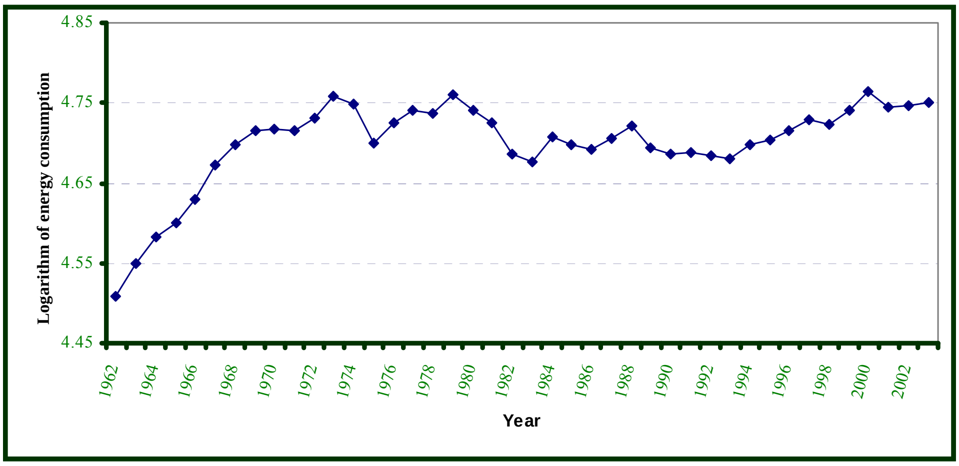
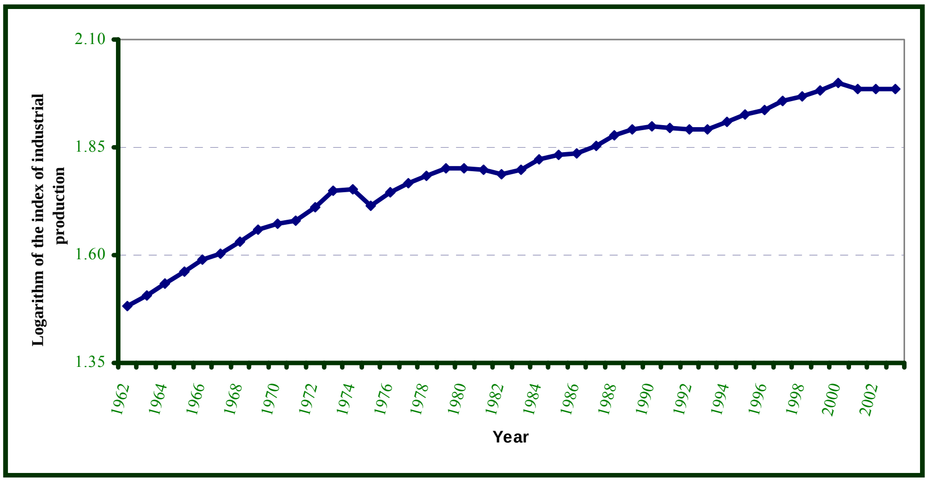


Figure 1c: OECD Index of Industrial Production (2000=100)



Notes

(i) The price series is a weighted average aggregate energy index price for the 15 OECD countries in the sample (weights being energy consumption)..

(ii) The activity series is a weighted average index of industrial production for the 15 OECD countries in the sample (weights being real GDP).

Figure 2: Decomposed (weighted) OECD Industrial Energy Price (2000=100)

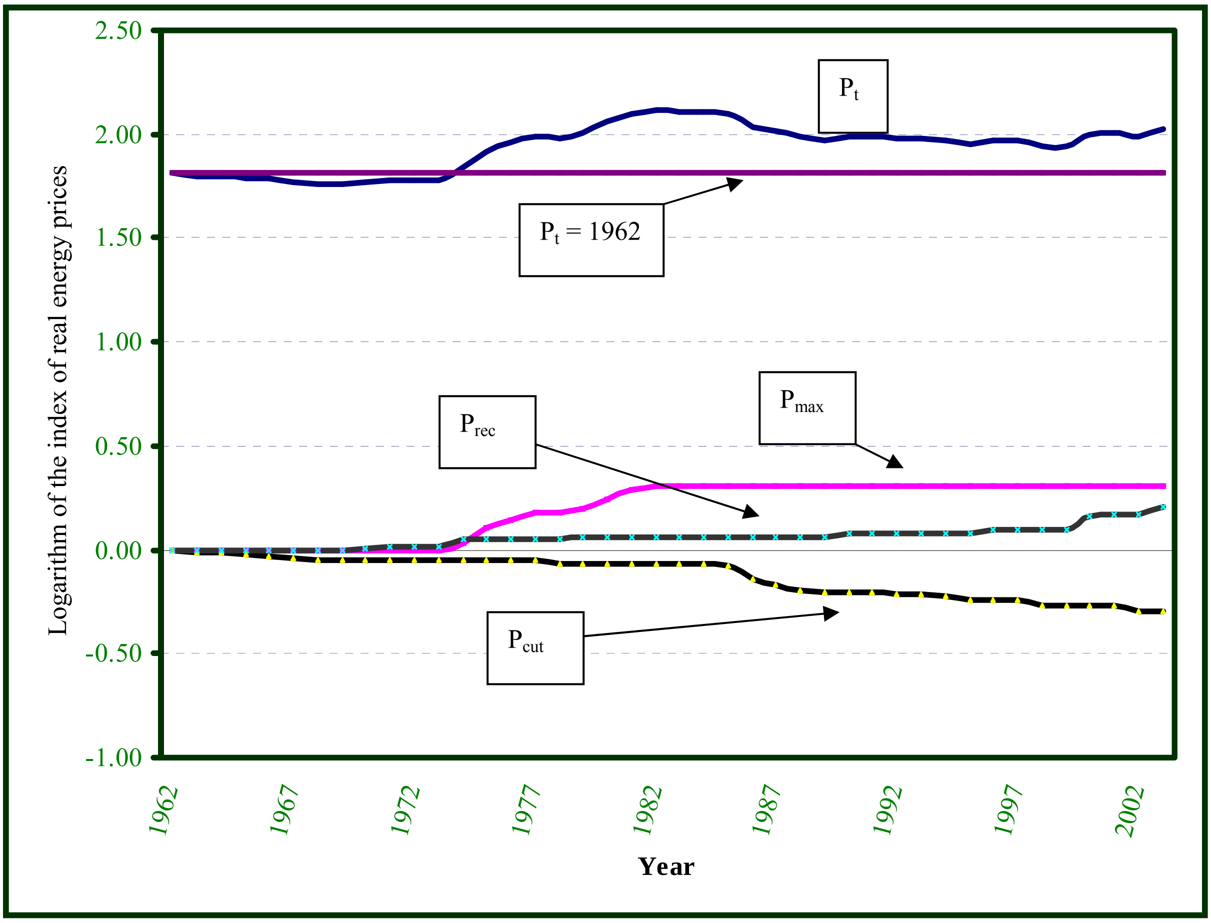


Figure 3: Estimates of energy-saving technical change for Model II and III in Table 2

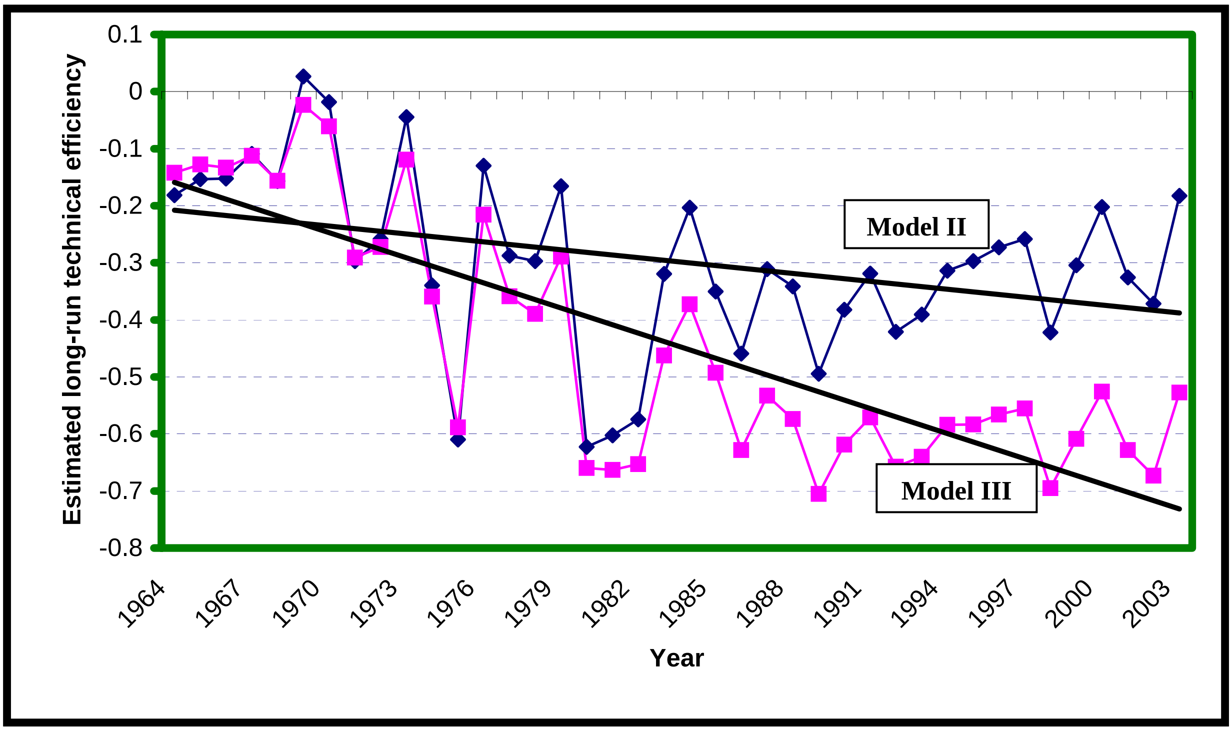


Table 1: Selection of previous studies on industrial energy demand published since 1990

Study	Types of energy	Country	Technique/ Model used	Treatment of trend	Data used	Estimated LR Elasticities
Hunt & Lynk (1992)	Aggregate energy	UK	Cointegration; ECM	Deterministic trend	Annual data 1952-1988	Price = -0.3 Income = 0.7
Jones (1995)	Various fuels	USA	Dynamic linear-logit factor substitution model	Deterministic trend	Annual data 1960-1992	Price = -0.1 to -1.3 Income = NE
Jones (1996)	Aggregate energy	G-7 countries	Dynamic linear-logit factor substitution model	No trend included	Annual data 1960-1991	Price = -0.23 to -2.5 Income = NE
Casler (1997)	Aggregate energy	USA	Various factor substitution model: SURE	No trend included	Annual data 1947-1971	Price = -0.5 to -0.8 Income = NE
Dahl & Erdogan (2000)	Aggregate energy	Turkey	Translog factor substitution model: SURE; 3SLS	No trend included	Annual data 1963-1992	Price = -0.2 to -0.3 Income = NE
Medlock III & Soligo (2001)	Aggregate energy	28 OECD/ non-OECD countries.	2SLS	No trend or time dummies in preferred model	Annual panel data 1978-1995	Price = -0.3 Income = 3.9
Hunt et al (2003a)	Aggregate energy	UK	STSM; ARDL	Stochastic trend	Quarterly data 1971Q1-1995Q4	Price = -0.2 Income = 0.7
Chang & Martinez-Chombo (2003)	Electricity	Mexico	Cointegration & ECM with TVP	No trend included	Monthly data 1985M01-2000M05	Price = 0.1 Income = 0.2 to 0.3
Kamerschen & Porter (2004)	Electricity	USA	PAM; 3SLS; simultaneous equation	Deterministic trend	Annual data 1960-1992	Price = -0.4 to -0.6 Income = 0.01 to 0.13
Dimitropoulos et al (2005)	Aggregate energy	UK	STSM; ARDL	Stochastic trend	Annual data 1967-1999	Price = -0.2 Income = 0.7

Notes

(i) NE = not estimated; ECM = error correction mechanism; SURE = seemingly unrelated regression equations; 2SLS = two stage least squares; . 3SLS = three stage least squares; STSM = structural time series model; ARDL = auto regressive distributed lag; TVP = time varying parameters; PAM = partial adjustment mechanism; GMM = generalized method of moments.

(ii) Some of the studies also include estimates for other sectors, but are estimated separately so not included here.

(iii) Some studies actually refer to the manufacturing sector rather than the industrial sector.

(iv) Jones (1996) did include estimates for individual countries but the aggregate G7 results across the four fuels coal, oil electricity and gas are given above.

Table 2: OECD Industrial Energy Demand

Estimated Parameters	Model I	Model II	Model III
β (income)	0.777***	0.562***	0.551***
	[0.00]	*[0.00]*	*[0.00]*
γ (price)		-0.014	
		[0.21]	
γ_m (price-max)	-0.036***		0.019
	[0.00]		*[0.19]*
γ_r (price-rec)	-0.047***		-0.020
	[0.00]		*[0.31]*
γ_c (price- cut)	-0.021		-0.073***
	[0.16])		*[0.00]*
λ (lagged energy)	0.931***	0.938***	0.921***
	[0.00]	*[0.00]*	*[0.00]*
Time Dummies Included	*No*	*Yes*	*Yes*
Long-run elasticities			
Income	0.78	0.56	0.55
Total Price		-0.22	
Price – max	-0.52		0.24
Price – rec	-0.68		-0.25
Price – cut	-0.30		-0.92
No. of observations	615	615	615
No. of estimated parameters	20	58	60
SSR	0.226660	0.188842	0.186338
Diagnostics			
Autocorrelated errors	$t_{(579)}$= -0.66*[0.51]*	$T_{(542)}$= -2.04*[0.04]***	$t_{(540)}$= -1.89*[0.06]**
Hausman	$t_{(594)}$= -0.40*[0.69]*	$t_{(556)}$= -0.92*[0.36]*	$t_{(554)}$= -0.74*[0.46]*
Nested Restriction Tests			
No fixed time effects: $\theta_t = 0$			$F_{(40,555)}$=3.00*[0.00]****
Symmetric price response: $\gamma_{max} = \gamma_{rec} = \gamma_{cut}$	$F_{(2,595)}$=9.24*[0.00]****		$F_{(2,555)}$=3.73*[0.00]****
Non-Nested J Tests			
$\psi = 0$	$t_{(594)}$= -0.05*[0.96]*		
$\varphi = 0$		$t_{(556)}$= 0.90*[0.37]*	

Notes:

(i) All estimation undertaken in Eviews5.

(ii) Models I, II, and III correspond to equations (6), (8), and (9).

(iii) The adjusted R^2 for all the specifications were very high at over 0.99.

(iv) *[...]* indicate probability values, with ***, **, and * representing significance at the 1% level, 5% level, and 10% level respectively (for the estimated parameters the probabilities are calculated using White's heteroscedastic consistent variances).

(v) The long-run price elasticity is calculated as $\gamma/(1-\lambda)$ for the total price and $\gamma_{max}/(1-\lambda)$, $\gamma_{rec}/(1-\lambda)$, and $\gamma_{cut}/(1-\lambda)$ for the long run Price-max, Price-rec and Price-cut elasticities respectively.

(vi) The test for autocorrelated errors tests for the presence of an AR(1) process.

Table 3: Summary of Panel Unit Root Tests for OECD Industrial Energy Demand (e_{it})

Method	**Statistic *[probability]***
Null: Unit root (assumes common unit root process)	
Levin, Lin & Chu t*	-4.88 *[0.00]****
Breitung t-stat	0.42 *[0.66]*
Null: Unit root (assumes individual unit root process)	
Im, Pesaran and Shin W-stat	-2.21 *[0.01]****
ADF - Fisher Chi-square	47.65 *[0.02]***
PP - Fisher Chi-square	67.73 *[0.00]****
Null: No unit root (assumes common unit root process)	
Hadri Z-stat	8.60 *[0.00]****

Notes:

(vii) All estimation undertaken in Eviews5.

(viii) Includes linear trends and lags of one year.

(ix) *[...]* indicate probability values, with ***, **, and * representing significance at the 1% level, 5% level, and 10% level respectively.

(x) Probabilities for Fisher tests are computed using an asymptotic χ^2 distribution. All other tests assume asymptotic normality.

Note:

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