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Abstract

In this paper we estimate residential electricity demand elasticities and conduct an analysis of the causal relationship between electricity demand, disposable income and electricity price for a group of several OECD members. We apply panel cointegration and Granger causality testing to a data set consisting of eighteen countries in the cross-sectional dimension and the years 1981–2008 in the time domain. Our results for the whole panel indicate a near unity income elasticity and an inelastic price elasticity of approximately –0.4 in the long run. These results are robust with regard to the estimation methods employed (group-means panel FMOLS and DOLS). In the short run, our estimates from an ECM indicate an income elasticity of 0.2 and a price elasticity of approximately –0.1. Moreover, our tests on Granger causality provide an indication for a bidirectional causal relationship between electricity consumption and economic growth. Hence, our findings are in favor of the feedback hypothesis.

JEL classification: Q41; Q43

Keywords: Residential electricity demand; Elasticities; Panel cointegration; Granger-causality; OECD countries

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1. Introduction

Insights concerning the responsiveness of energy demand to other economic factors have a crucial relevance for policy advice with regard to objectives such as the security of energy supply and GHG abatement. Especially in the light of the current debate and concern regarding anthropological causes of global warming and its effects on the environment, these topics have risen again on the agenda of policy-makers and researchers alike.

A possible policy instrument for the purpose of energy conservation is a Pigouvian energy tax. Therefore, knowledge about the response of energy consumption to changes in energy prices is of paramount importance for optimal policy design.

Furthermore, the assessment of the causal relationship between energy use and economic activity is valuable for an appraisal of potentially conflicting policy objectives, such as the trade-off between energy conservation and economic growth. There are four opposing economic hypotheses regarding the causal mechanisms underlying the energy consumption – economic growth nexus, which are currently heavily under debate in the energy economics literature (for a useful survey see Payne, 2010): the *conservation hypothesis*, the *growth hypothesis*, the *feedback hypothesis* and the *neutrality hypothesis*. While the conservation hypothesis implies causality running from economic growth to energy consumption, the growth hypothesis implies the opposite causal relationship. The feedback hypothesis combines the latter hypotheses, by claiming an interdependent causal relationship between both variables. Finally, the neutrality hypothesis states that both variables are only of little importance in determining each other.

Most previous studies on estimating residential electricity demand elasticities are based on time series data. One major problem of economic time series is the likely existence of stochastic trends in the variables. This requires applying econometric approaches, i.e. unit root and cointegration methods, which take the nonstationarity of the data-generating process (DGP) underlying the considered variables explicitly into account. However, traditional unit root and cointegration tests in a pure time series context are known to suffer from the problem of very low power and size. Hence, increasing the number of observations by including a cross-sectional dimension helps to reduce this problem. The added cross-sections can be interpreted as repeated draws from the same distribution, that increase power and hence permit more reliable statistical inference.

The aim of this paper is twofold. First, we estimate residential electricity demand elasticities with regard to income and own price for a group of OECD countries, thereby differentiating between the short and the long run in an error correction framework. The long-run relationship is estimated by applying the fully modified ordinary least squares (FMOLS) and dynamic OLS (DOLS) group-means panel estimators (Pedroni, 2000 and 2001) to a panel of eighteen countries in the cross-sectional dimension and the period 1981–2008 in the time domain. The countries considered are: Austria, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Mexico, the Netherlands, Norway, Portugal, South Korea, Spain, Switzerland, the UK and the US.

Second, we investigate the causal relationship between residential electricity demand, disposable income and electricity price by utilizing the concept of Granger causality (Granger, 1969). For this purpose we employ the pooled mean group estimator (PMGE) proposed by Pesaran *et al.* (1999) for estimating a trivariate panel vector error correction model (PVECM). A set of tests on short-run, long-run and joint causality are conducted, in order to assess the direction of causality between the variables at hand.

Following the arguments in Bernstein and Madlener (2010), we choose a disaggregate approach in the sense that, firstly, we restrict our analysis to one energy carrier (electricity). Secondly, we only examine one economic sector (households), aiming at capturing the behavior of a relatively homogeneous group of economic agents.

The econometric estimation of energy demand elasticities has a long tradition, going back as far as the early 1950's. Among the first studies are Houthakker (1951), Fisher and Kaysen (1962), Halvorsen (1975) and Pindyck (1979). Similarly, the related literature strand aimed at determining the causal relationship in the energy-growth nexus was sparked by the early work of Kraft and Kraft (1978), and has received considerable interest in the last ten to fifteen years. However, it was only with the introduction of cointegration analysis, which was triggered by the seminal paper of Engle and Granger (1987), that the problem of spurious regressions was starting to be adequately dealt with in econometric applications.

Table 1 gives an overview of selected recent studies on residential electricity demand elasticities, most of which employ pure time series methods for single countries:[1] Athukorala

[1] We only consider studies here that employ time-series or panel estimation techniques, that account for non-stationarity in the DGP, and that were published after the year 2000.

and Wilson (2010) for Sri Lanka; Dergiadis and Tsoulfidis (2008) for the US; Halicioglu (2007) for Turkey; Holtedahl and Joutz (2004) for Taiwan; Hondroyiannis (2004) for Greece; Nakajima (2010) for Japan; Nakajima and Hamori (2010) for the US; Narayan and Smyth (2005) for Australia; Narayan *et al.* (2007) for the G7 countries; Zachariadis and Pashourtidou (2007) for Cyprus and Ziramba (2008) for South Africa. The long-run demand elasticities from these studies range between 0.25 and 1.57 with regard to income and between –0.14 and –1.56 with regard to own price. For the short run, the elasticity estimates range between 0.10 and 0.44 with regard to income and between –0.11 and –0.46 with regard to own price.

Unlike the majority of the studies reported, Nakajima and Hamori (2010), Nakajima (2010) and Narayan *et al.* (2007), use cointegration techniques that are based on panel data in their analyses. In order to attain a cross-sectional dimension, the former two studies make use of available data on geographical regions within Japan and the US, respectively. The latter study is closest to ours with regard to the countries[2] analyzed and the econometric approach used, but interestingly comes to different conclusions. Specifically, Narayan *et al.* (2007) analyze a panel consisting of the G7 countries as the cross-sectional dimension for the time period 1978–2003. They employ both the OLS and the DOLS method (Stock and Watson, 1993; Kao and Chiang, 2000) to estimate the cointegration relationship between electricity demand, income, electricity price, and natural gas price. For the whole panel they estimate an inelastic income elasticity of 0.25 (0.31) and an elastic price elasticity of –1.56 (–1.45) using DOLS (OLS). Consequently, they come to the conclusion that pricing policies aimed at reducing residential electricity consumption and hence GHG emissions are bound to be successful.

Table 2 displays the three studies from the set of studies mentioned above that conduct Granger causality analyses for residential electricity consumption, income and electricity price. Dergiades and Tsoulfidis (2008) and Halicioglu (2007) both find unidirectional long-run causality from income and price to electricity consumption. For the short run, the former find causality from electricity price to income, while the latter finds causality from income and price to electricity consumption.[3] Zachariadis and Pashourtidou (2007) find evidence for a

[2] Our set of countries also includes the G7 countries, except for Canada, for which recent data on the residential electricity price were lacking.

[3] Dergiades and Tsoulfidis (2008) also consider other variables such as oil price and cooling and heating degree days, for which we do not report the results.

Table 1
Recent residential electricity demand studies and elasticity estimates

Study	Country	Method	Data	Elasticity estimates	
				Income	Price
Athukorala & Wilson (2010)	Sri Lanka	Johansen / VECM	Time series, 1960–2007 (annual)	*L*: 0.78 *S*: 0.32	*L*: –0.62 *S*: –0.16
Dergiades & Tsoulfidis (2008)	US	Bounds testing / ARDL	Time series, 1965–2006 (annual)	*L*: 0.27 *S*: 0.10	*L*: –1.07 *S*: –0.39
Halicioglu (2007)	Turkey	Bounds testing / ARDL	Time series, 1968–2005 (annual)	*L*: 0.49 to 0.70 *S*: 0.37 to 0.44	*L*: –0.52 to –0.63 *S*: –0.33 to –0.46
Holtedahl & Joutz (2004)	Taiwan	Johansen / VECM	Time series, 1955–1995 (annual)	*L*: 1.04 to 1.57 *S*: 0.22	*L*: –0.15 *S*: –0.15
Hondroyiannis (2004)	Greece	Johansen / VECM	Time series, 1986–1999 (monthly)	*L*: 1.56 *S*: 0.20	*L*: –0.41
Nakajima (2010)	Japan	Panel cointegration, DOLS	Panel data, 1975–2005 (annual), $T \times N$: 31 × 46 = 1426	*L*: 0.60 to 0.65	*L*: –1.13 to –1.20
Nakajima & Hamori (2010)	US	Panel cointegration, DOLS	Panel data, 1993–2008 (quarterly), $T \times N$: 32 × 49 = 1568	*L*: 0.38 to 0.85	*L*: –0.14 to –0.33
Narayan & Smyth (2005)	Australia	Bounds testing / ARDL	Time series, 1969–2000 (annual)	*L*: 0.32 to 0.41 *S*: *0.01* to *0.04*	*L*: –0.47 to –0.54 *S*: –0.26 to –0.27
Narayan *et al.* (2007)	G7	Panel Cointegration, OLS & DOLS	Panel data, 1978–2003 (annual), $T \times N$: 26 × 7 = 182	*L*: 0.25 to 0.31 *S*: *–0.19*	*L*: –1.45 to –1.56 *S*: –0.11
Zachariadis & Pashourtidou (2007)	Cyprus	Johansen / VECM	Time series, 1960–2004 (annual)	*L*: 1.18	*L*: –0.43
Ziramba (2008)	South Africa	Bounds testing / ARDL	Time series, 1978–2005 (annual)	*L*: 0.31 to 0.87 *S*: 0.30	*L*: *–0.01* to *–0.04* *S*: *–0.02*

Notes: S and *L* denote estimates for the short and the long run, respectively. Elasticity estimates which are not statistically significantly different from zero on conventional levels are printed in *italics*. *T*: Number of time series observations; *N*: Number of cross-sections. DOLS: Dynamic OLS; ARDL: Autoregressive Distributed Lag.

long-run causal relationship, running from income and price to electricity consumption, and from electricity consumption and price to income. For the short run, they only find causality running from electricity consumption to income. Hence, so far there is empirical evidence for the conservation as well as the feedback hypothesis concerning the causal relationship between residential electricity demand and real income.

Table 2
Results from causality analyses on residential electricity demand

Study	Country	Data	Direction of causality	
			Long-run	Short-run
Dergiades & Tsoulfidis (2008)	US	Time series, 1965–2006 (annual)	$Y, P \rightarrow E$	$P \rightarrow Y$
Halicioglu (2007)	Turkey	Time series, 1968–2005 (annual)	$Y, P \rightarrow E$	$Y, P \rightarrow E$
Zachariadis & Pashourtidou (2007)	Cyprus	Time series, 1960–2004 (annual)	$Y, P \rightarrow E$ $E, P \rightarrow Y$	$E \rightarrow Y$

Notes: *Y*: Income; *P*: Electricity price; *E*: Electricity consumption; → denotes the direction of causality.

Our paper proceeds as follows. In Section 2, we provide the analytical framework for the econometric analysis undertaken. Section 3 gives a methodological overview of the applied estimation and testing procedures applied, while Section 4 discusses the data, the application of the model and the results from the analysis. Section 5 concludes.

2. Analytical framework

The long-run relationship between residential electricity demand and its determinants can be characterized by the general function

$$E_{i,t} = f\left(Y_{i,t}, P_{i,t}, X_{i,t}\right), \tag{1}$$

with the subscripts i (i = 1,…, N) and t (t = 1,…, T), denoting the cross-sectional and time dimension, respectively. Eq. (1) states that residential electricity consumption per capita ($E_{i,t}$) is a function of real (disposable) income per capita ($Y_{i,t}$) and real residential electricity price ($P_{i,t}$).[4] Previous studies on residential electricity demand have included further control variables (X_t), as for example the real price of an electricity substitute (e.g., Narayan *et al.*, 2007), heating and cooling degree days (e.g., Zachariadis and Pashourtidou, 2007),

[4] In the following, real disposable income per capita and real residential electricity price will be referred to as "income" and "electricity price" for brevity.

urbanization (e.g., Holtedahl and Joutz, 2004), and capital variables (e.g., Silk and Joutz, 1997).

Following the principle of Occam's razor, we choose a parsimonious specification, which only includes real disposable income per capita and real residential electricity price as determinants of electricity demand. Furthermore, as the following analysis comprises a panel of various countries, problems with the availability of data on additional variables would have brought further restrictions to the data set with regard to the cross-sections and/or the time period studied.[5] Finally, having less parameters to estimate has the advantage of attaining more degrees of freedom in the estimation of the core explanatory variables' coefficients. More specifically, the demand model on which our econometric analysis is based takes the following standard constant elasticity functional form:

$$E_{i,t} = C_i \, Y_{i,t}^{\beta_{y,i}} \, P_{i,t}^{\beta_{p,i}} \, e^{\varepsilon_{i,t}}, \tag{2}$$

where the subscripts i ($i = 1, \ldots, N$) and t ($t = 1, \ldots, T$) represent the cross-sectional dimension (the eighteen countries considered) and the time dimension (the years 1981–2008 considered), respectively. C_i are country-specific drift terms, e is Euler's number, $\varepsilon_{i,t}$ are random error terms, and $\beta_{y,i}$ and $\beta_{p,i}$ are the long-run elasticities to be estimated with regard to income and electricity price, respectively. A higher income is expected to increase electricity demand on account of higher economic activity, whereas a higher electricity price is naturally expected to decrease electricity demand. Moreover, the price elasticity is expected to be inelastic, as in general electricity is characterized by a lack of substitutability.

3. Methodology

3.1. Panel unit root tests

As a preliminary analysis, we check for the order of integration of the single series by employing a number of panel unit root tests. In detail, these are: the LLC (Levin *et al.*, 2002),

[5] Also, not all variables, such as cooling degree days or the price of natural gas, are likely to have the same relevance or, more specifically, similar explanatory power for all the considered countries.

the UB (Breitung, 2000), the IPS (Im *et al.*, 2003), the ADF-Fisher (Maddala and Wu, 1999), and the PP-Fisher (Maddala and Wu, 1999).

All the mentioned tests are based on the following AR(1) panel regression model:

$$x_{i,t} = \delta_i \, x_{i,t-1} + \Gamma_i \, X_{i,t} + u_{i,t}, \qquad i = 1, 2, \ldots, N; \quad t = 1, 2, \ldots, T, \tag{3}$$

where δ_i are the autoregressive parameters, $X_{i,t}$ represents exogenous variables and/or fixed effects and cross-section-specific time trends and $u_{i,t}$ are stationary error terms. In the case that $|\delta_i| < 1$, x_i is referred to be weakly trend-stationary. In contrast, if $|\delta_i| = 1$, x_i is considered to be a unit root process.

The five aforementioned tests can be divided into two different groups with respect to the assumptions about the δ_i. The LLC test and the UB test assume that all cross-sections have a common unit root, i.e. $\delta_i = \delta$ for all i. On the other hand, the IPS test, the ADF-Fisher test and the PP-Fisher test all assume that the δ_i can be heterogeneous across the cross-sections. In order to save space we refer the reader to the original articles for further details on these tests.

In the case where the test results for a variable in levels indicate a rejection of the null hypothesis, whereas the test results for the same variable in first differences does not reject the null at conventional significance levels, this variable is assumed to be integrated of order one, denoted $I(1)$.

3.2. Panel cointegration tests

Given that the variables electricity consumption, income and electricity price are all integrated of order one, i.e. $e_{i,t}$, $y_{i,t}$, $p_{i,t} \sim I(1)$, we can include them in the cointegration analysis, in order to be able to estimate the long-run relationship described by Eq. (2). Taking natural logarithms of Eq. (2) yields the econometric specification of our long-run residential electricity demand function:

$$e_{i,t} = c_i + \beta_{y,i} y_{i,t} + \beta_{p,i} p_{i,t} + \varepsilon_{i,t}, \qquad i = 1, 2, \ldots, N; \quad t = 1, 2, \ldots, T, \tag{4}$$

where $e_{i,t} = \ln(E_{i,t})$, $y_{i,t} = \ln(Y_{i,t})$, $p_{i,t} = \ln(P_{i,t})$, c_i are country-specific fixed effects and $\varepsilon_{i,t}$ are the error terms, which are interpreted as deviations from long-run equilibria. The country-specific slope coefficients $\beta_{y,i}$ and $\beta_{p,i}$ are the long-run elasticities to be estimated with regard to income and electricity price, respectively. Hence, this specification allows for the cointegrating vectors to vary across the single countries of our panel.

Pedroni (1999, 2004) extends the cointegration testing approach of Engle and Granger (1987), which is based on examining the stationarity properties of the residuals from a regression using $I(1)$ variables, to a panel data setting. Following this approach, Eq. (4) is estimated by OLS, and the residuals obtained, $\hat{\varepsilon}_{i,t}$, are used for the following auxiliary autoregression for every i:

$$\hat{\varepsilon}_{i,t} = \rho_i \hat{\varepsilon}_{i,t-1} + \sum_{j=1}^{n_i} \varphi_{i,t} \Delta \hat{\varepsilon}_{i,t-j} + w_{i,t}, \tag{5}$$

Where the ρ_i are autoregressive parameters, n_i are the lag lengths in the augmented case, and w_i are stationary error terms.

Under the null hypothesis of no cointegration, the $\hat{\varepsilon}_{i,t}$ should be found to be $I(1)$. This is the case if

$$H_0: \rho_i = 1, \quad \forall\, i = 1, ..., N$$

is not rejected. For each of the seven statistics provided by Pedroni (1999, 2004) this is the null hypothesis. Concerning the alternative hypothesis, the tests can be divided into two classes. For the so-called (within-dimension) panel statistics tests (i.e. the Panel-v, the Panel-PP-ρ, the Panel-PP-t and the Panel-ADF-t test) the alternative hypothesis is

$$H_1: \rho_i < 1, \quad \forall\, i = 1, ..., N,$$

whereas for the so-called (between-dimension) group statistics tests (i.e. the Group-PP-ρ, the Group-PP-t and the Group-ADF-t test) the alternative hypothesis is

$$H_1: (\rho_i = \rho) < 1, \quad \forall\, i = 1, ..., N.$$

Hence, the group statistics tests are less restrictive in the sense that they allow for heterogeneity across countries.

3.3. Estimation of long- and short-run elasticities

Given that the panel cointegration tests indicate a significant cointegration relationship, we apply the fully modified OLS (FMOLS) and the dynamic OLS (DOLS) group-means panel estimators proposed by Pedroni (2000, 2001) for estimating the long-run demand relationship characterized by Eq. (4). Both estimators allow for standard normal inference through incorporating corrections for endogeneity bias and serial correlation.[6] While the FMOLS estimator employs a semi-parametric correction using $\Delta y_{i,t}$, $\Delta p_{i,t}$ and $\hat{\varepsilon}_{i,t}$, the DOLS estimator employs a parametric approach by augmenting Eq. (4) with lead and lag dynamics of $\Delta y_{i,t}$ and $\Delta p_{i,t}$ as follows:

$$e_{i,t} = \mu_i + \beta_{y,i} y_{i,t} + \beta_{p,i} p_{i,t} + \sum_{l=-l_i}^{l_i} \lambda_{y,i,l} \Delta y_{i,t-l} + \sum_{l=-l_i}^{l_i} \lambda_{p,i,l} \Delta p_{i,t-l} + \upsilon_{i,t} \tag{6}$$

where l_i is the lead and lag length, μ_i is the country-specific fixed effect and $\upsilon_{i,t}$ is the error term.

The group-means FMOLS and DOLS panel estimates for the slope coefficients, $\hat{\beta}_G^{FMOLS}$ and $\hat{\beta}_G^{DOLS}$, and their corresponding t-statistics, $t\left(\hat{\beta}_G^{FMOLS}\right)$ and $t\left(\hat{\beta}_G^{DOLS}\right)$, are calculated as follows:

$$\hat{\beta}_{G,y}^{m} = \frac{1}{N}\sum_{i=1}^{N}\hat{\beta}_{y,i}^{m} \tag{7}$$

$$t\left(\hat{\beta}_{G,y}^{m}\right) = \frac{1}{\sqrt{N}}\sum_{i=1}^{N} t\left(\hat{\beta}_{y,i}^{m}\right) \tag{8}$$

$$\hat{\beta}_{G,p}^{m} = \frac{1}{N}\sum_{i=1}^{N}\hat{\beta}_{p,i}^{m} \tag{9}$$

$$t\left(\hat{\beta}_{G,p}^{m}\right) = \frac{1}{\sqrt{N}}\sum_{i=1}^{N} t\left(\hat{\beta}_{p,i}^{m}\right) \tag{10}$$

where the superscript m is a place holder denoting either the FMOLS or the DOLS estimation method; $\hat{\beta}_{y,i}$ and $\hat{\beta}_{p,i}$ are the country-specific estimates of income and price elasticity, respectively.

[6] Harris and Sollis (2003) provide an excellent exposition on this topic.

Considering both the FMOLS and the DOLS panel approach has the advantage of being able to provide some evidence on the robustness of our results with regard to the estimation method.

In order to estimate the short-run elasticities and the speed of adjustment to long-run equilibrium, the residuals from the cointegrating regressions, which resemble the deviation from long-run equilibrium in any given period *t*, are used as error correction terms (*ECT*) in country-specific and panel error correction models (ECM). The latter takes on the following form

$$\Delta e_{i,t} = \gamma_{0,i}^{m} + \alpha_{i}^{m} ECT_{i,t-1}^{m} + \gamma_{y,i}^{m} \Delta y_{i,t} + \gamma_{p,i}^{m} \Delta p_{i,t} + v_{i,t}^{m}, \tag{11}$$

where m denotes the estimation method (FMOLS or DOLS), $\gamma_{0,i}$ is a country-specific constant, α_i is the speed of adjustment coefficient, ECT_i is the aforementioned error correction term lagged by one period, $\gamma_{y,i}$ and $\gamma_{p,i}$ are the short-run income and price elasticities, respectively, and $v_{i,t}$ are the error terms. To ensure an error correction mechanism via adjustments of electricity consumption, α_i has to be negative.

3.4. Panel Granger causality

A cointegration relationship between a set of variables necessarily implies Granger causality in at least one direction. For the residential electricity demand relationship in Eq. (2), the direction of Granger causality between the single variables is tested by the two-step Engle-Granger procedure (Engle and Granger, 1987) on the basis of the following PVECM:

$$\Delta e_{i,t} = \gamma_{1,i}^{m} + \phi_{1,i}^{m} ECT_{i,t-1}^{m} + \sum_{j=1}^{k} \theta_{11,i,j}^{m} \Delta e_{i,t-j} + \sum_{j=1}^{k} \theta_{12,i,j}^{m} \Delta y_{i,t-j} + \sum_{j=1}^{k} \theta_{13,i,j}^{m} \Delta p_{i,t-j} + \varphi_{1,i,t}^{m} \tag{12a}$$

$$\Delta y_{i,t} = \gamma_{2,i}^{m} + \phi_{2,i}^{m} ECT_{i,t-1}^{m} + \sum_{j=1}^{k} \theta_{21,i,j}^{m} \Delta e_{i,t-j} + \sum_{j=1}^{k} \theta_{22,i,j}^{m} \Delta y_{i,t-j} + \sum_{j=1}^{k} \theta_{23,i,j}^{m} \Delta p_{i,t-j} + \varphi_{2,i,t}^{m} \tag{12b}$$

$$\Delta p_{i,t} = \gamma_{3,i}^{m} + \phi_{3,i}^{m} ECT_{i,t-1}^{m} + \sum_{j=1}^{k} \theta_{31,i,j}^{m} \Delta e_{i,t-j} + \sum_{j=1}^{k} \theta_{32,i,j}^{m} \Delta y_{i,t-j} + \sum_{j=1}^{k} \theta_{33,i,j}^{m} \Delta p_{i,t-j} + \varphi_{3,i,t}^{m} \tag{12c}$$

where m again denotes the estimation method, ECT_i is the error correction term (obtained from estimation of Eq. (4)), k is the lag length, the ϕ_i are the speed of adjustment coefficients, the γ_i are country-specific constants, the θ_i are short-run coefficients and the $\varphi_{i,t}$ are serially uncorrelated error terms. The equation system (12a-c) is estimated by the PMGE method proposed by Pesaran *et al.* (1999).

For the long run, the direction of Granger causality can be checked for by testing whether the adjustment coefficient of the respective equation is significantly different from zero. Besides long-run Granger causality, short-run Granger causality can occur through the lagged first differences of the independent variables in each equation of the system. Hence we check for short-run Granger causality from income and electricity price to electricity consumption by testing $H_0 : \theta_{12,i,j} = 0,\ \forall\, i,j$ and $H_0 : \theta_{13,i,j} = 0,\ \forall\, i,j$ in Eq. (12a). For short-run Granger causality from electricity consumption and electricity price to income in Eq. (12b), we test $H_0 : \theta_{21,i,j} = 0,\ \forall\, i,j$ and $H_0 : \theta_{23,i,j} = 0,\ \forall\, i,j$, respectively. In Eq. (12c) we test $H_0 : \theta_{31,i,j} = 0,\ \forall\, i,j$ and $H_0 : \theta_{32,i,j} = 0,\ \forall\, i,j$ for short-run Granger causality running from electricity consumption and income to electricity price, respectively. Finally, we test on the joint significance between the lagged *ECT* and the lagged differences of the independent variables in the respective equations in order to check for strong causality in each equation.

4. Empirical analysis

4.1. Data

We make use of annual data on residential electricity consumption, net disposable income and residential electricity price, which are all available for a reasonable length of time (1978–2008) for eighteen OECD countries already mentioned in the introduction.

The time series are transformed as follows: In order to obtain real values, disposable income and residential electricity price are deflated to the 2000 levels using the consumer price index (CPI) of the respective country. After that, the 2000 exchange rate conversion factors are used to standardize both the series for every non-Euro country to €. Electricity consumption and disposable income are divided by total population in order to get per capita values for each country. Finally, all the series are log-transformed.[7]

[7] Note, that the German unification in October 1990 is treated by imposing trends for West-Germany to the series of unified Germany in 1991.

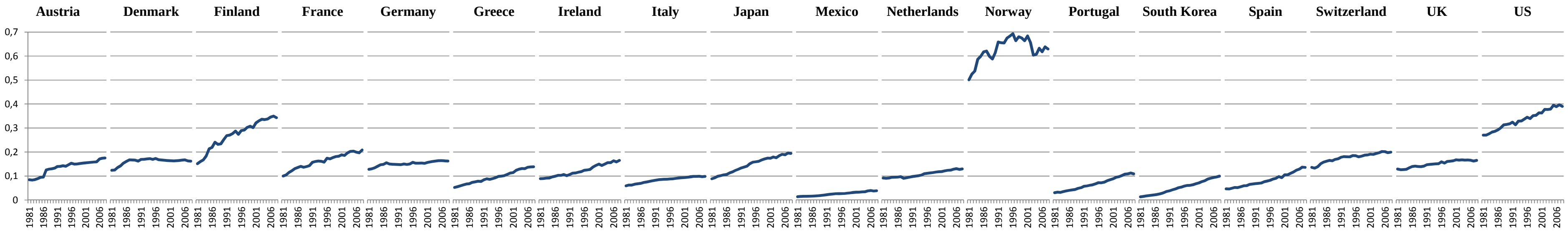
Data on residential electricity consumption and residential electricity price are taken from the International Energy Agency (IEA) database (http://www.iea.org/stats) 'Energy Balances of OECD Countries' and 'Energy Prices & Taxes', respectively, while the currency exchange rates are from the European Central Bank (ECB, http://www.ecb.int/stats). All other variables, i.e. national net disposable income, total population and CPIs, are adopted from the OECD database (http://stats.oecd.org).

The second oil price shock in 1979/1980 will most likely have introduced exogenous structural breaks to the DGP of the time series at hand. In order to circumvent this problem, we truncate the initially available period (1978–2008) accordingly, which leaves us with a sample ranging from 1981 to 2008 for every country. This yields a panel with the dimensions $N = 18$ and $T = 28$. Hence, our analysis is based on 504 observations in total.

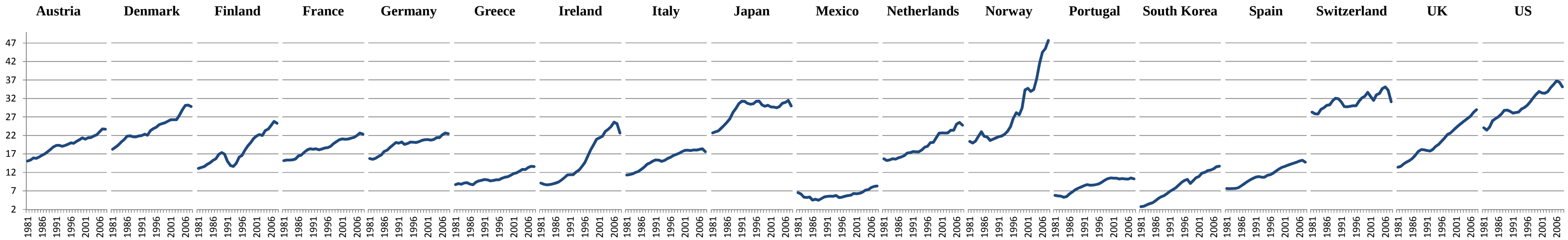
Fig. 1, plots (a) – (c) displays the individual time series of electricity consumption (measured in tons of oil equivalent, toe) per capita, real disposable income (measured in constant 1000 €, base year 2000 = 100) per capita, and real electricity price (measured in constant 1000 €/toe, base year 2000 = 100), respectively.

Both, electricity consumption and real income show an overall upward trend in all countries considered. The decreasing effect of the latest financial crisis in 2008 on the level of these variables can be seen in nearly all of the income series. One anomaly worth mentioning is the Norwegian electricity consumption, which features very high per capita values and reveals a very high volatility compared to other countries. Both characteristics presumably are due to the fact that Norwegian households predominantly (approximately 70%) use electricity for heating purposes, which of course is highly subject to the weather conditions in a given year.

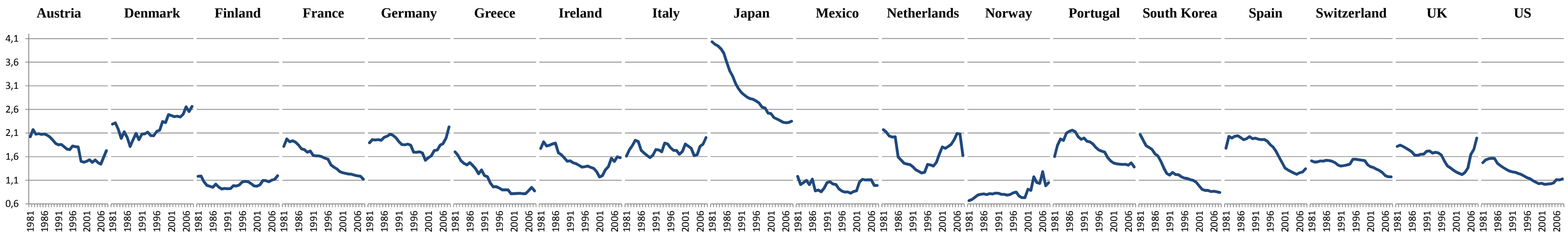
For the price of electricity it is difficult to identify an overall trend for all countries. One price series that stands out is the one for Japan, which compared internationally, starts on a very high level in the beginning of the 1980's and then shows a sharp decline throughout the time period under consideration. One of the reasons for this high initial level is a very high proportion of oil in the power generation mix at that time and the all-time peak in oil prices following the oil crises. The subsequent decline of electricity prices can, amongst other possible reasons, be attributed to the following decrease of oil prices. Moreover, since the early 1990's, the Japanese government has undertaken several efforts with regard to power



(a) Residential electricity consumption



(b) Real net disposable income



(c) Real residential electricity prices

Fig. 1: Visual inspection of data by country, 1981–2008; Notes: **(a)** Residential electricity consumption (toe) per capita. **(b)** Real net disposable income (1000 €) per capita, 2000 = 100. **(c)** Real residential electricity prices (1000 €/toe), 2000 = 100. Data sources: IEA, ECB, OECD, own calculations and illustration.

market liberalization in order to lower electricity price to an internationally more competitive level (see IEA, 2008).

4.2. Panel unit root tests

To check for the properties of the time series, we apply the panel unit root tests outlined in Section 3.1.[8] The test statistics and the corresponding *p*-values are reported in Table 3. While the test regressions for the variables in levels (e, y, p) contain an intercept and a time trend, the test regressions for the variables in first differences (Δe, Δy, Δp) only contain an intercept. For the first four tests, the lag lengths are selected according to the usual information criteria.

Table 3
Panel unit root tests

Method	e	Δe	y	Δy	p	Δp
LLC	–0.67	–15.09***	0.29	–3.45***	0.12	–9.77***
	[0.25]	[0.00]	[0.61]	[0.00]	[0.55]	[0.00]
UB	1.07	–6.09***	5.37	1.15	3.65	–1.53*
	[0.86]	[0.00]	[1.00]	[0.87]	[1.00]	[0.06]
IPS	0.23	–15.02***	–0.33	–7.25***	–0.26	–10.76***
	[0.59]	[0.00]	[0.37]	[0.00]	[0.40]	[0.00]
ADF-Fisher	31.47	179.25***	46.62	123.87***	39.29	95.69***
	[0.68]	[0.00]	[0.11]	[0.00]	[0.32]	[0.00]
PP-Fisher	43.88	282.98***	19.63	132.09***	44.41	218.85***
	[0.17]	[0.00]	[0.99]	[0.00]	[0.16]	[0.00]

Notes: *** and * denote significance at the 1% and 10% level, respectively. *p*-values are reported in squared brackets. LLC: Levin *et al*. (2002); UB: Breitung (2000); IPS: Im *et al*. (2003); ADF-Fisher and PP-Fisher: Maddala & Wu (1999).

The results of the first five tests on the variables in levels do not allow for a rejection of the panel unit root hypothesis at the conventional significance levels and hence indicate an order of integration of at least one for all three variables. However, the tests on the first differences of the variables reject the null hypothesis at least at the 10% level. An exception is the UB test statistic, which indicates a second panel unit root for real income. As this order of integration is very unlikely for variables such as income, and the other tests clearly reject the panel unit

[8] We also considered the test of Hadri (2000), but the results, which we do not report here, indicate the existence of a second unit root for all three variables. In a simulation study, Hlouskova and Wagner (2006) find that the Hadri test tends to overreject the null hypothesis of stationarity, which is consistent with the findings of most empirical applications.

root hypothesis for the first differences of the income variable, we conclude for all three variables that they are indeed integrated of order one, $I(1)$.

4.3. Panel cointegration tests

Having established a non-stationary behavior for all the series, we proceed to testing for a long-run cointegrating relationship by applying Pedroni's panel cointegration tests described in Section 3.2. The results for both the within- and the between-dimension tests are summarized in Table 4. Except for the Panel PP-ρ and the Group PP-ρ tests, the null hypothesis of no cointegration is rejected at the 1% level. Using Monte Carlo simulations, Pedroni (2004) investigates the small-sample properties of the Panel v, the Panel PP-ρ, the Panel PP-t, the Group PP-ρ and the Group PP-t statistics for different dimensions of the panel. In the case where $N = 20$, which is fairly close to our case, the Panel PP-t and Group PP-t perform best regarding size and power when $T < 130$ and $T < 40$, respectively. Hence, we have some evidence for a stationary behavior of the residuals from Eq. (4) and conclude that there exists a panel-cointegrating relationship between residential electricity consumption, real disposable income and real residential electricity price.

Table 4
Pedroni's panel cointegration test

Method	**Unweighted**		**Weighted**		**Inference**
	Statistic	[Prob.]	Statistic	[Prob.]	
Alternative hypothesis: common AR coefficients (within-dimension)					
Panel v	8.78***	[0.00]	4.52***	[0.00]	$\hat{\varepsilon}_{i,t} \sim I(0)$
Panel PP-ρ	–0.51	[0.30]	–0.67	[0.25]	$\hat{\varepsilon}_{i,t} \sim I(1)$
Panel PP-t	–3.10***	[0.00]	–3.59***	[0.00]	$\hat{\varepsilon}_{i,t} \sim I(0)$
Panel ADF-t	–4.39***	[0.00]	–4.46***	[0.00]	$\hat{\varepsilon}_{i,t} \sim I(0)$
Alternative hypothesis: individual AR coefficients (between-dimension)					
Group PP-ρ	0.44	[0.67]			$\hat{\varepsilon}_{i,t} \sim I(1)$
Group PP-t	–3.11***	[0.00]			$\hat{\varepsilon}_{i,t} \sim I(0)$
Group ADF-t	–4.14***	[0.00]			$\hat{\varepsilon}_{i,t} \sim I(0)$

Notes: Normalization: $e_{i,t} = c_i + \beta_{y,i}\, y_{i,t} + \beta_{p,i}\, p_{i,t} + \varepsilon_{i,t}$. Null hypothesis: no cointegration. Tests assume individual intercepts. Lag lengths selected by Schwarz Information Criterion (SIC). *** denotes significance at the 1% level. p-values are reported in squared brackets.

4.4. Long-run and short-run elasticities

Long-run elasticities

As a next step we employ Pedronis' group-means FMOLS and DOLS panel estimators described in Section 3.3 for estimating the long-run demand relationship characterized by Eq. (4). The coefficient estimates and the corresponding *t*-statistics from both the individual tests and the panel tests are summarized in Table 5. Comparing the FMOLS and DOLS estimates of the slope coefficients with each other, we find that most of them are in agreement concerning sign and magnitude, rendering our results robust with regard to the estimation method. Furthermore, the plausibility of the cointegrating vector estimates in terms of sign and magnitude, as well as the statistical significance, indicate that we have indeed attained a reasonable approximation of the true equilibrium relationship. In the following, we summarize the results in more detail.

For the whole panel the estimate of the income elasticity is 0.96 (0.91) from the FMOLS (DOLS) estimator. This implies that a 1% rise in income is associated with a 0.9%–1.0% increase in electricity demand. The estimate of the price elasticity is –0.39 (–0.38) from the FMOLS (DOLS) estimator, implying a change in the magnitude of –0.39% (–0.38%) in electricity demand in response to a 1% increase in electricity price.

In the country-specific regressions, the coefficients on real disposable income from both FMOLS and DOLS have a positive sign and are significant at the 1% level in sixteen out of the eighteen countries considered. The magnitudes of the significant coefficients vary between 0.38 (0.35) in the UK and a high 2.04 (1.92) in Austria from the FMOLS (DOLS) estimator. The price elasticities have a negative sign and are significant at the 10% level in ten out of eighteen countries for both the FMOLS and DOLS estimations. The magnitudes of the significant coefficients vary between –0.12 (–0.14) in the UK and –1.36 (–1.37) in Japan from the FMOLS (DOLS) estimator.

The only country which exhibits neither a significant income nor a significant price elasticity is Norway. A possible explanation can be found in the observations from the graphical inspection of Fig. 1, plot (a), in Section 4.1. Accordingly, in contrast to other countries, weather conditions presumably have a much higher explanatory power for the variation in electricity consumption in Norway, as compared to, say, income or the price of electricity.

Table 5
Long-run elasticities by (group-means panel) FMOLS and DOLS estimation

Countries	Fully modified OLS (FMOLS)		Dynamic OLS (DOLS)	
	y	*p*	*y*	*p*
Group	**0.96*** **	**–0.39*** **	**0.91*** **	**–0.38*** **
	(38.33)	(–9.76)	(27.96)	(–13.19)
Austria	**2.04*** **	0.37	**1.92*** **	0.27
	(6.53)	(1.24)	(5.81)	(0.81)
Denmark	**0.82*** **	**–0.87*** **	**0.79*** **	**–0.80*** **
	(9.15)	(–7.63)	(7.25)	(–5.37)
Finland	**0.87*** **	**–1.03** **	**0.99*** **	–0.85
	(5.70)	(–2.32)	(4.95)	(–1.51)
France	**1.59*** **	0.08	**1.64*** **	0.03
	(3.10)	(0.23)	(3.23)	(0.10)
Germany	**0.48*** **	–0.04	**0.42*** **	**–0.16** **
	(8.80)	(–0.65)	(6.01)	(–2.25)
Greece	**1.14*** **	**–0.58*** **	**0.98*** **	**–0.56*** **
	(7.89)	(–6.60)	(4.82)	(–6.04)
Ireland	**0.53*** **	0.03	**0.53*** **	0.06
	(15.46)	(0.31)	(10.93)	(0.42)
Italy	**0.99*** **	0.04	**0.90*** **	0.10
	(34.75)	(0.57)	(25.09)	(0.97)
Japan	–0.09	**–1.36*** **	–0.05	**–1.37*** **
	(–0.43)	(–11.32)	(–0.55)	(–27.51)
Mexico	**1.61*** **	**–1.28*** **	**1.57*** **	**–1.16** **
	(5.92)	(–2.88)	(5.00)	(–2.21)
Netherlands	**0.74*** **	0.00	**0.56*** **	–0.01
	(20.55)	(0.09)	(6.13)	(–0.13)
Norway	0.10	–0.13	0.14	–0.25
	(0.95)	(–0.63)	(0.87)	(–0.97)
Portugal	**1.42*** **	**–0.80*** **	**1.60*** **	**–0.67* **
	(7.68)	(–2.88)	(4.91)	(–1.88)
South Korea	**0.97*** **	**–0.58** **	**0.86*** **	**–0.66** **
	(6.78)	(–2.23)	(6.63)	(–2.83)
Spain	**1.21*** **	**–0.34** **	**1.24*** **	**–0.30** **
	(9.79)	(–2.35)	(9.61)	(–2.30)
Switzerland	**1.72*** **	–0.09	**1.34*** **	–0.23
	(4.38)	(–0.33)	(3.11)	(–0.70)
UK	**0.38*** **	**–0.12* **	**0.35*** **	**–0.14** **
	(10.12)	(–1.93)	(9.17)	(–2.13)
US	**0.72*** **	**–0.23** **	**0.66*** **	**–0.21** **
	(5.50)	(–2.08)	(5.64)	(–2.41)

Notes: Dependent variable: Electricity consumption (*e*). *t*-statistics appear in parentheses below the respective coefficients. ***, ** and * denote significance at the 1%, 5% and 10% level, respectively. Coefficient estimates significant at the 10% level are highlighted in boldface.

Table 6 compares country-specific estimates of long-run elasticities from our analysis to estimates from previous studies:

- For France, our estimate of the income elasticity is fairly high (1.64) and of similar magnitude as the one obtained in Narayan *et al.* (2007) of 1.49. However, our estimate of the price elasticity is not significant, while their estimate is, and amounts to –0.50.
- For Germany, our income elasticity is 0.42 compared to 0.54 (not significant) from Narayan *et al.* (2007). Concerning the price elasticity we find a considerable difference: While we estimate an elasticity of –0.16, their estimate takes on a value of –4.20, which seems implausibly high in magnitude, as electricity price elasticities in general are expected to be fairly inelastic due to limited substitution possibilities for electricity.
- For Greece, we have very similar results with regard to the price elasticity as Hondroyiannis (2004), i.e. –0.56 versus –0.41, respectively. However, while we estimate a near unity income elasticity, his estimate is somewhat higher (1.56).
- For Italy, our estimate of the income elasticity amounts to 0.90, while the estimate of the price elasticity is statistically not significant. In Narayan *et al.* (2007) both estimates are not significant.
- For Japan, our estimate for the price elasticity, which appears to be fairly high in magnitude (–1.37), is supported by both Nakajima (2010) and Narayan *et al.* (2007), whose estimates are –1.20 and –1.49, respectively.
- For the UK, our estimate of the income elasticity amounts to 0.35, while the estimate of the price elasticity amounts to –0.14. The estimates from Narayan *et al.* (2007) are both not significant.
- For the US, our results are much closer to Nakajima and Hamori (2010) than to Dergiades and Tsoulfidis (2008). While our estimate for the income elasticity (price elasticity) is 0.66 (–0.21), which is fairly close to the 0.85 (–0.33) from Nakajima and Hamori, Dergiades and Tsoulfidis find an income elasticity (price elasticity) of 0.27 (–1.07). In Narayan *et al.* (2007) both estimates are not significant.

Comparing our group estimates of electricity demand elasticities to the ones from Narayan *et al.* (2007) we find a striking difference in magnitudes, despite the similarity with regard to the time span analyzed and the estimation method used. While we find a near unity income elasticity (0.9 to 1) and an inelastic price elasticity (–0.4) for the whole panel, their estimates reveal an inelastic income elasticity (0.2 to 0.3) and an elastic price elasticity (–1.6 to –1.5).

We suspect that their group estimate for price elasticity is most likely biased to a considerable extent by the unreasonably high estimate for Germany (–4.20), especially bearing in mind that the sample only consists of seven countries in the cross-sectional dimension.

Table 6
Comparison of long-run country-specific demand elasticities

Countries	BM		NSP		H		DS		N		NH	
	Y	*P*	*Y*	*P*	*Y*	*P*	*Y*	*P*	*Y*	*P*	*Y*	*P*
France	**1.64**	0.03	**1.49**	**–0.50**	—	—	—	—	—	—	—	—
Germany	**0.42**	**–0.16**	0.54	**–4.20**	—	—	—	—	—	—	—	—
Greece	**0.98**	**–0.56**	__	__	**1.56**	**–0.41**	—	—	—	—	—	—
Italy	**0.90**	0.10	–0.49	0.08	—	—	—	—	—	—	—	—
Japan	–0.05	**–1.37**	–0.89	**–1.49**	—	—	—	—	**0.65**	**–1.20**	—	—
UK	**0.35**	**–0.14**	0.66	0.60	—	—	—	—	—	—	—	—
US	**0.66**	**–0.21**	0.40	0.33	—	—	**0.27**	**–1.07**	—	—	**0.85**	**–0.33**

Notes: *Y* and *P* denote income and price elasticities, respectively. Elasticity estimates statistically significantly different from zero at conventional levels are highlighted in boldface. BM: our estimations; NSP: Narayan *et al.* (2007); H: Hondroyiannis (2004); DS: Dergiades & Tsoulfidis (2008); N: Nakajima (2010); NH: Nakajima & Hamori (2010). For simplicity of illustration we only report the DOLS estimates from BM and only the most recent estimates from NH.

Hence, quite contrary to the findings of Narayan *et al.* (2007), the implications from our results indicate that in general the potential effectiveness of pricing policies aimed at reducing residential electricity consumption are very limited.

Short-run elasticities

Using the residuals from the FMOLS and DOLS cointegrating regressions, we estimate two ECMs for each country and the whole panel according to Eq. (11). The coefficients and the corresponding *t*-statistics are reported in Table 7. As expected, electricity consumption adjusts negatively to deviations from the long-run equilibrium both in the panel and in the country-specific models.

For the panel, the adjustment coefficients have the expected negative sign to ensure error correction to the long-run equilibrium. The short-run income elasticity estimate is 0.23 and 0.17, while the short-run price elasticity is estimated at –0.06 and –0.05, depending on whether the *ECT*s are estimated by FMOLS or DOLS. Hence, the elasticities are lower in magnitude than their long-run counterparts, which complies with intuition and the results from previous studies.

Table 7
Short-run elasticities

Countries	*ECT*s from FMOLS estimation				*ECT*s from DOLS estimation			
	Constant	Δy	Δp	ECT_{t-1}	Constant	Δy	Δp	ECT_{t-1}
Group	0.02***	**0.23*****	**–0.06*****	**–0.17*****	0.02***	**0.17*****	**–0.05****	**–0.27*****
	(8.04)	(4.99)	(–2.68)	(–6.40)	(8.38)	(4.28)	(–2.40)	(–6.81)
Austria	0.02	0.79	–0.04	**–0.36****	0.02	0.73	–0.10	**–0.35****
	(1.57)	(1.22)	(–0.18)	(–2.21)	(1.10)	(1.14)	(–0.49)	(–2.22)
Denmark	0.00	**0.45***	**–0.21****	**–0.45*****	0.00	0.40	**–0.19***	**–0.43*****
	(0.62)	(1.84)	(–2.14)	(–3.74)	(0.62)	(1.66)	(–1.99)	(–3.61)
Finland	0.03***	0.03	0.15	**–0.23*****	0.03***	–0.03	0.10	**–0.19*****
	(3.54)	(0.19)	(0.91)	(–3.57)	(3.68)	(–0.17)	(0.58)	(–2.99)
France	0.03***	–0.40	–0.22	**–0.36*****	0.03***	–0.45	–0.24	**–0.29****
	(2.88)	(–0.98)	(–0.80)	(–2.92)	(3.05)	(–1.06)	(–1.05)	(–2.36)
Germany	0.01	0.24	0.04	–0.33	0.01*	–0.05	–0.05	**–0.49***
	(1.19)	(0.85)	(0.36)	(–1.63)	(1.70)	(–0.23)	(–0.50)	(–1.75)
Greece	0.03***	0.19	–0.17	**–0.24***	0.03***	0.13	–0.07	**–0.52*****
	(4.42)	(0.86)	(–1.52)	(–1.90)	(4.72)	(0.67)	(–0.74)	(–3.06)
Ireland	0.02**	0.10	**–0.19***	**–0.49*****	0.02***	–0.01	–0.14	**–0.56****
	(2.43)	(0.86)	(–1.93)	(–3.09)	(3.29)	(–0.08)	(–1.30)	(–2.41)
Italy	0.01***	**0.34****	–0.02	**–0.29****	0.01***	**0.28****	–0.04	**–0.55*****
	(3.91)	(2.43)	(–0.45)	(–2.25)	(4.02)	(2.20)	(–0.74)	(–2.78)
Japan	0.02***	0.16	–0.30	**–0.26****	0.02**	0.16	–0.30	**–0.26****
	(3.69)	(0.78)	(–1.02)	(–2.15)	(3.69)	(0.78)	(–1.02)	(–2.15)
Mexico	0.04***	**0.23***	0.08	–0.05	0.04***	0.22	0.09	**–0.16***
	(5.13)	(1.74)	(0.90)	(–1.08)	(4.88)	(1.65)	(1.03)	(–1.89)
Netherlands	0.01*	0.18	0.02	–0.27	0.01*	0.03	0.02	–0.60
	(1.88)	(0.77)	(0.33)	(–1.61)	(1.85)	(0.11)	(0.37)	(–1.63)
Norway	0.01	0.07	**–0.12****	**–0.29*****	0.01	0.06	**–0.11***	**–0.40*****
	(1.14)	(0.49)	(–2.10)	(–3.00)	(0.75)	(0.42)	(–1.85)	(–2.96)
Portugal	0.04***	0.27	0.02	–0.12	0.05***	0.05	0.16	**–0.33*****
	(4.73)	(1.37)	(0.09)	(–1.37)	(6.67)	(0.32)	(0.83)	(–3.01)
South Korea	0.03***	**0.53*****	–0.27	**–0.23****	0.04***	**0.45*****	**–0.29***	**–0.46*****
	(3.76)	(4.53)	(–1.57)	(–2.10)	(4.51)	(4.43)	(–1.91)	(–3.09)
Spain	0.03***	0.30	0.01	**–0.27***	0.03***	0.14	–0.14	**–0.52*****
	(3.02)	(0.89)	(0.06)	(–1.72)	(3.53)	(0.48)	(–0.64)	(–2.81)
Switzerland	0.02***	**0.31****	**0.61*****	**–0.27*****	0.02***	**0.28****	**0.45****	**–0.21*****
	(4.54)	(2.76)	(3.84)	(–4.53)	(4.07)	(2.16)	(2.64)	(–3.29)
UK	0.02**	–0.21	**–0.16***	**–0.31***	0.01*	–0.16	**–0.15***	**–0.33***
	(2.03)	(–0.86)	(–2.00)	(–1.76)	(1.78)	(–0.66)	(–1.94)	(–1.95)
US	0.01	**0.29***	–0.21	**–0.47****	0.01	**0.29***	**–0.23***	**–0.85*****
	(1.13)	(1.80)	(–1.52)	(–2.72)	(1.58)	(2.01)	(–1.85)	(–4.03)

Notes: Dependent variable: First difference of electricity consumption (Δe). *t*-statistics appear in parentheses below the respective coefficients. ***, ** and * denote significance at the 1%, 5% and 10% level, respectively. Slope coefficient estimates which are significant at the 10% level are highlighted in boldface.

In the country-specific ECMs the adjustment coefficients are all negative and mostly significant in both the FMOLS-based ECMs and the DOLS-based ECMs. In both models for the Netherlands, significance is only found at a level of approximately 12%. For the most part, the country-specific short-run elasticities are not significantly different from zero. Where they are, the signs and magnitudes are reasonable, and, in magnitude well below the corresponding long-run elasticities. The only exception is the short-run price elasticity of Switzerland, which takes on a significant positive value.

4.5. Panel Granger causality testing

As the last step in our analysis of residential electricity demand in OECD countries, we conduct tests on Granger causality between electricity consumption, disposable income and electricity price. As described in Section 3.4, we estimate two PVECMs according to Eq. (12a-c), both of which are based on the residuals of the FMOLS and DOLS regressions, respectively, using the PMGE method proposed by Pesaran *et al.* (1999). The lag length is chosen such that serially uncorrelated residuals are ensured. The Schwarz Information Criterion points at an optimal lag length of one, $k = 1$. The results are reported in Table 8.

First of all, in the short run we find evidence for causality running from disposable income to electricity consumption in both, the FMOLS- and DOLS-based models. Vice versa, the null hypothesis of no short-run causality from electricity consumption to disposable income is only rejected in the FMOLS-based PVECM. In both models, the *ECT*s in the electricity consumption and disposable income equations are significant and have the right sign in order to ensure error correction to long-run equilibrium after a shock occurs. This bidirectional long-run Granger causality between electricity consumption and disposable income is also supported by the results of the *F*-tests on joint significance. The magnitudes of the adjustment coefficients (not reported in Table 8) are –0.13 (–0.25) and 0.17 (0.12) for the electricity consumption and disposable income equation in the FMOLS-based (DOLS-based) model. Thus, near-complete adjustments (of at least 95%) to long-run equilibrium induced by changes in electricity consumption and disposable income take approximately seven to nine years.

For the electricity price equation there is only evidence for an adjustment to equilibrium, and hence long-run causality, in the FMOLS-based model and only at the 10% level of significance. However, the F-tests on joint significance are unable to reject the null hypotheses. This indicates that electricity price in our models is to be viewed as strongly exogenous.

Table 8
Panel Granger causality tests

Dependent variable	Independent variables (sources of causation)						
	Short-run			Long-run	Joint		
	Δe	Δy	Δp	*ECT*	*ECT* & Δe	*ECT* & Δy	*ECT* & Δp
Panel A: ECTs from FMOLS							
(12a) Δe	–	2.87*** [0.00]	0.56 [0.58]	–6.52*** [0.00]	–	26.00*** [0.00]	21.65*** [0.00]
(12b) Δy	1.40 [0.16]	–	–0.68 [0.49]	8.48*** [0.00]	38.24*** [0.00]	–	36.89*** [0.00]
(12c) Δp	0.22 [0.83]	0.83 [0.40]	–	–1.82* [0.07]	1.67 [0.19]	2.07 [0.13]	–
Panel B: ECTs from DOLS							
(12a) Δe	–	2.15** [0.03]	0.03 [0.98]	–9.58*** [0.00]	–	50.53*** [0.00]	46.31*** [0.00]
(12b) Δy	1.97** [0.05]	–	0.26 [0.79]	4.13*** [0.00]	11.34*** [0.00]	–	8.73*** [0.00]
(12c) Δp	–0.06 [0.95]	1.26 [0.21]	–	–0.76 [0.45]	0.30 [0.74]	1.18 [0.31]	–

Notes: t-statistics are displayed for short-run and long-run causation. Partial F-statistics are displayed for joint causation. p-values are in brackets. ***, ** and * denote significance at the 1%, 5% and 10% level, respectively. Tests are based on a PVECM(1).

Altogether, our results suggest that the feedback hypothesis holds at least in the long run, while in the short run the results are mixed. This implies that policy measures aiming at electricity conservation at the same time will involve a trade-off in the form of a decreasing effect on economic growth at least in the long run.

Comparing our results to the three studies listed in Table 2, we find support for the results of Zachariadis and Pashourtidou (2007), who also find evidence for the feedback hypothesis. Halicioglu (2007) and Dergiades and Tsoulfidis (2008), on the other hand, both find evidence for the conservation hypothesis in the long run.

5. Conclusions

In this paper we have conducted an analysis on the responsiveness of residential electricity demand for a set of OECD countries. We estimate electricity demand elasticities with regard to disposable income and own price, thereby differentiating between the short and the long run in an error correction framework. Furthermore, we conduct tests on Granger causality between the variables. In order to circumvent the problem of low power and size of more traditional unit root and cointegration tests based on time series data, we make use of available panel data for eighteen countries in the cross-sectional dimension and the years 1981–2008 in the time domain. The methods employed for estimating the cointegrating vectors are Pedroni's group-means fully modified OLS (FMOLS) and dynamic OLS (DOLS) panel estimators. The Granger causality tests are based on a panel vector error correction model estimated using the pooled mean group estimator (Pesaran *et al.*, 1999).

Our findings for the whole panel indicate a near unity income elasticity and an inelastic price elasticity of approximately –0.4 in the long run. These results are robust to the two estimation methods employed. In the short run, our estimates from error correction models indicate an income elasticity of 0.2 and a price elasticity of approximately –0.1. When compared with the results of previous studies on residential electricity demand, our elasticity estimates bear some resemblance to other estimates on the individual country level.

Overall, our results imply that the steering effect of tax-induced price increases on residential electricity demand has a very limited potential for energy conservation, and hence a reduction of GHG emissions. These policy implications are in stark contrast, e.g., to the ones of Narayan *et al.* (2007).

Furthermore, our tests on Granger causality indicate that both electricity consumption and income adjust toward the long-run equilibrium after a shock hits the system. Thus, a bidirectional causal relationship between electricity consumption and economic growth exists in the long run. Our findings are, therefore, in favor of the feedback hypothesis. Hence, a reduction of electricity consumption will be associated with a trade-off with regard to per capita income.

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