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Energy Consumption and Economic Growth: New Insights into the Cointegration Relationship[1]

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Abstract. This paper examines the long-run relationship between energy consumption and real GDP, including energy prices, for 25 OECD countries from 1981 to 2007. The distinction between common factors and idiosyncratic components using principal component analysis allows to distinguish between developments on an international and a national level as drivers of the long-run relationship. Indeed, cointegration between the common components of the underlying variables indicates that international developments dominate the long-run relationship between energy consumption and real GDP. Furthermore, the results suggest that energy consumption is price-inelastic. Causality tests indicate the presence of a bi-directional causal relationship between energy consumption and economic growth.

JEL-Classification: C33, O13, Q43

Keywords: Energy consumption, panel unit roots, panel cointegration, vector error-correction models, Granger causality

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1 Introduction

The question of whether or not energy conservation policies affect economic activity is of great interest in the international debate on global warming and the reduction of greenhouse gas emissions. Although the causal relationship between energy consumption and economic growth has been widely studied, no consensus regarding this so-called energy consumption-growth nexus has yet been reached. The direction of causality is highly relevant for policy makers. For instance, if causality runs from energy consumption to economic growth, energy conservation policies that have the aim of reducing energy consumption may have a negative impact on an economy's growth. The literature proposes four different hypotheses regarding the possible outcomes of causality Apergis and Payne (2009a,b).[1] The growth hypothesis suggests that energy consumption is a crucial component in growth, directly or indirectly as a complement to capital and labour as input factors of production. Hence, a decrease in energy consumption causes a decrease in real GDP. In this case, the economy is called 'energy dependent' and energy conservation policies may be implemented with adverse effects on real GDP. By contrast, the conservation hypothesis claims that policies directed towards lower energy consumption may have little or no adverse impact on real GDP. This hypothesis is based on a uni-directional causal relationship running from real GDP to energy consumption. Bi-directional causality corresponds with the feedback hypothesis, which argues that energy consumption and real GDP affect each other simultaneously. In this case, policy makers should take into account the feedback effect of real GDP on energy consumption by implementing regulations to reduce energy use. Additionally, economic growth should be decoupled from energy consumption to avoid a negative impact on economic development resulting from a reduction of energy use. A shift from less efficient energy sources to more efficient and less polluting options may establish a stimulus rather than an obstacle to economic growth (Costantini and Martini, 2010). Finally, the neutrality hypothesis indicates that reducing energy consumption does not affect economic growth or vice versa. Hence, energy conservation policies would not have any impact on real GDP.

Our analysis of the relationship between energy consumption and economic activity is based on a sample of 25 OECD countries from 1981 to 2007 and uses recently developed panel-econometric methods. We explore an additional channel of causality by introducing energy prices. As energy prices have been neglected in many previous studies, the long-run parameters and the evidence of causality may be biased, see Masih and Masih (1997) and Asafu-Adjaye (2000). But in contrast to these two studies, we examine the original energy price index rather than the consumer price index (CPI) as a proxy. Income and price elasticities provide policy makers a suggestion of the extent to which prices need to increase, in the form of energy taxes, in order to reduce energy consumption and the potential for the market to conserve energy (Lee and Lee, 2010). Additionally, energy companies need this information to design their demand management policies. But only a few papers have estimated income and price elasticities for energy consumption in a panel framework. Furthermore, the long-run equilibrium relationship is studied in both directions, i.e. with either energy consumption or real GDP as a dependent variable (Costantini and Martini, 2010).

The innovative contribution of our paper is to determine the long-run relationship between energy consumption, real GDP and energy prices in more detail. In contrast to other studies concerning the energy consumption-GDP growth nexus, we distinguish between national and international trends as potential drivers of the long-run equilibrium between energy consumption, real GDP and energy prices. To analyse these issues, each variable is decomposed into common and idiosyncratic compo-

[1] The different directions of causality between energy consumption and economic growth have been described in many previous studies. Also, the phrase 'neutrality hypothesis' has often been used. The denotations of the other causal relations were proposed by Apergis and Payne (2009a,b).

nents. The idiosyncratic component is the part of a variable that is driven by national developments, whereas the common component represents international trends in the evolution of the variables. These might, however, have a different relevance for individual countries. Taking this decomposition as a starting point, cointegration between the common components means that the common components of energy consumption, real GDP and energy prices move together in the long run and do not deviate permanently from one another. Hence, cointegration between the common components suggests that the relationship between these variables depends to a great extent on international developments. Instead, cointegration between idiosyncratic components refers to developments relevant exclusively on the national level (Dreger and Reimers, 2009). Depending on the results of the cointegration tests, this distinction has important implications for policy makers. If the common components cointegrate, national energy policies may not have a large impact on economic growth. Indeed, this paper delivers empirical evidence that energy consumption, real GDP and energy prices are cointegrated in their common factors, but not in their idiosyncratic components.

The remainder of this paper is organised as follows. Section 2 briefly reviews the literature related to the causal relationship between energy consumption and economic growth. Section 3 presents the data, discusses the econometric methods and presents the empirical results. Section 4 provides conclusions and policy implications.

2 Literature review

The empirical literature provides mixed and conflicting evidence with respect to the energy consumption-growth nexus. This discrepancy in results is due largely to the use of different econometric methods and time periods, besides country-specific heterogeneity in climate conditions, economic development and energy consumption patterns, among other things. From a methodological perspective, four generations of contributions can be identified. First generation studies applied a traditional vector autoregression (VAR) model in the tradition of Sims (1972). For example, the seminal work of Kraft and Kraft (1978), using a VAR model, found evidence in favour of causality running from income to energy consumption in the United States for the period 1947-1974. Further, studies of the first generation examined the direction of causality assuming stationarity of the underlying variables (see, e.g. Erol and Yu, 1987; Yu and Choi, 1985; Abosedra and Baghestani, 1989). Second generation studies accounted for non-stationarity in the data and performed cointegration analysis to investigate the long-run relationship between energy consumption and growth. This second generation literature, based on the Engle and Granger (1987) two-step procedure, studied pairs of variables to check for cointegration relationships and used estimated error-correction models to test for Granger causality (see, e.g. Nachane et al., 1988; Cheng and Lai, 1997; Glasure and Lee, 1998). Third generation studies used multivariate estimators in the style of Johansen (1991). Johansen's multivariate approach also allows for more than two variables in the cointegration relationship (see, e.g. Masih and Masih, 1997; Stern, 2000; Asafu-Adjaye, 2000; Soytas and Sari, 2003; Oh and Lee, 2004a,b). Finally, fourth generation studies employ recently developed panel-econometric methods to test for unit roots and cointegration relations. This literature estimates panel-based error-correction models to perform Granger causality tests (see, e.g. Lee, 2005; Al-Iriani, 2006; Mahadevan and Asafu-Adjaye, 2007; Lee and Chang, 2007, 2008; Apergis and Payne, 2009a,b; Lee and Lee, 2010; Costantini and Martini, 2010). Some selected studies and their empirical setups are summarized in Table 1.

Most of the studies dealing with the energy consumption-growth nexus focus on production side models, which often include capital stock and labour in addition to energy consumption and GDP. If one concentrates on energy demand, trivariate models with energy prices as an additional variable

Table 1: Overview of selected studies

Study	Method	Countries	Result
Kraft and Kraft (1978)	Bivar. Sims Causality	USA	Growth → Energy
Yu and Choi (1985)	Bivar. Granger test	South Korea	Growth → Energy
		Philippines	Engery → Growth
Erol and Yu (1987)	Bivar. Granger test	USA	Energy ∼ Growth
Yu and Jin (1992)	Bivar. Granger test	USA	Energy ∼ Growth
Masih and Masih (1996)	Trivar. VECM	Malaysia, Singapore & Philippines	Energy ∼ Growth
		India	Energy → Growth
		Indonesia	Growth → Energy
		Pakistan	Energy ↔ Growth
Glasure and Lee (1998)	Bivar. VECM	South Korea & Singapore	Energy ↔ Growth
Masih and Masih (1998)	Trivar. VECM	Sri Lanka & Thailand	Energy → Growth
Asafu-Adjaye (2000)	Trivar. VECM	India & Indonesia	Energy → Growth
		Thailand&Philippines	Energy ↔ Growth
Hondroyiannis et al. (2002)	Trivar. VECM	Greece	Energy ↔ Growth
Soytas and Sari (2003)	Bivar. VECM	Argentina	Energy ↔ Growth
		South Korea	Growth → Energy
		Turkey	Energy → Growth
		Indonesia & Poland	Energy ↔ Growth
		Canada, USA & UK	Energy ↔ Growth
Fatai et al. (2004)	Bivar. Toda and Yamamoto (1995)	Indonesia & India	Energy → Growth
		Thailand&Philippines	Energy ↔ Growth
Oh and Lee (2004b)	Trivar. VECM	South Korea	Energy ↔ Growth
Wolde-Rufael (2004)	Bivar. Toda and Yamamoto (1995)	Shanghai	Energy → Growth
Lee (2005)	Trivar. Panel VECM	18 developing nations	Energy → Growth
Al-Iriani (2006)	Bivar. Panel VECM	Gulf Cooperation C.	Growth → Energy
Lee and Chang (2008)	Mulitv. Panel VECM	16 Asian countries	Energy → Growth
Lee et al. (2008)	Trivar. Panel VECM	22 OECD countries	Energy ↔ Growth
Narayan and Smyth (2008)	Multiv. Panel VECM	G7 countries	Energy → Growth
Apergis and Payne (2009a)	Multiv. Panel VECM	11 countries of the Commonwealth of Independent States	Energy ↔ Growth
Apergis and Payne (2009b)	Multiv. Panel VECM	6 Central American countries	Energy → Growth
Lee and Lee (2010)	Multiv. Panel VECM	25 OECD countries	Energy ↔ Growth

Notes: X → Y means variable X Granger-causes variable Y.

should be used (see Oh and Lee, 2004b). The studies by Masih and Masih (1998), Asafu-Adjaye (2000), Fatai et al. (2004) as well as Mahadevan and Asafu-Adjaye (2007) take the consumer price index (CPI) as a proxy of the energy price. However, as the CPI is known not to capture the energy price very well, we employ the real energy price index, such as Lee and Lee (2010) and Costantini and Martini (2010). Masih and Masih (1998) and Asafu-Adjaye (2000) previously used the vector error-correction model (VECM); Fatai et al. (2004) applied the autoregressive distributed lag (ARDL) approach; and Mahadevan and Asafu-Adjaye (2007), Lee and Lee (2010) as well as Costantini and Martini (2010) used a panel vector error-correction specification for the trivariate model.

In this paper, we study the cointegration property in more precise terms within a panel-econometric framework. Firstly, in order to distinguish between national and international trends that might drive the overall cointegration relationship, each variable is separated into common and idiosyncratic components by a principal component analysis. As a second step, we test common and idiosyncratic components separately for unit roots and their cointegration properties. Lastly, we apply Granger causality tests within a panel error-correction model.

3 Data, methodology and empirical results

In our study we use annual data from 1981 to 2007 for 25 OECD countries. These are Australia, Austria, Belgium, Canada, the Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Japan, Slovakia, South Korea, Luxembourg, Mexico, the Netherlands, Portugal, Poland, Spain, Sweden, the United Kingdom, and the United States. The sample period has been chosen such that the second oil crisis of 1979/80 is not included.[2] Data on real GDP per capita in constant 2000 U.S. dollars using purchasing power parities (PPPs) are used as a proxy of economic growth (Y).[3] Furthermore, time series data for the final energy consumption in kilotonnes of oil equivalent (ktoe) (E) and for the energy price index (P) in U.S. dollars (PPP) have been collected. All variables are in natural logarithms and have been obtained from the International Energy Agency's (IEA) online database.[4]

It is widely known that standard unit root and cointegration tests based on individual time series have low statistical power, especially when the time series is short (Campbell and Perron, 1991). Panel-based tests represent an improvement in this respect by exploiting additional information that results from the inclusion of the cross-sectional dimension. However, first generation panel unit root and cointegration tests often assume that the cross-section members are independent. This condition is often likely to be violated, for example, because of common oil price shocks. Therefore, our study controls for cross-section dependencies by taking into account a common factor structure. Suppose that the underlying model can be expressed as

$$Y_{i,t} = \alpha_i + \beta_i X_{i,t} + \varepsilon_{i,t} \quad (1)$$

$$\varepsilon_{i,t} = \xi_i' F_t + E_{i,t}, \quad (2)$$

[2] As expected, the time pattern of the energy prices exhibits considerable peaks at the period of the oil crisis. We originally collected data from 1978 to 2007.

[3] We use per capita data because they are less sensitive to territorial changes and provide the variables in the same units for countries of different sizes Lanne and Liski (2004).

[4] More precisely, data for real GDP per capita and final energy consumption are taken from IEA publications on energy balances of OECD countries (annual), while data for energy prices are drawn from IEA statistics on energy prices and taxes (quarterly).

where $i = 1, ..., N$ represents the cross-section member and $t = 1, ..., T$ refers to the time period, F denotes the common factors and E the idiosyncratic components. Hence, the error ε can follow a common factor structure. The common and idiosyncratic components can be either integrated of order one, I(1), or stationary, I(0), and we therefore have to test both components separately for unit roots and cointegration relationships. Cointegration implies that both the common and idiosyncratic parts of the error term are stationary.

3.1 Variable decomposition

The first and novel step of this paper regarding the energy consumption-growth nexus is to decompose each variable into two uncorrelated components, i.e. a common and an idiosyncratic component, by principal component analysis. The idiosyncratic component is a residual, which captures the impact of shocks affecting the respective variable of one specific country. These country-specific shocks, such as natural disasters, may have large but geographically concentrated effects. The common component of a variable is 'common' in the sense that it depends on a small number of common shocks, which affect the respective variable of all the countries. The decomposition is based on differenced data because of potential non-stationarity of the levels of the variables, as suggested by Bai and Ng (2004). After estimating the common factors they are re-cumulated to match the integration properties of the original variables. We obtain the idiosyncratic components from a regression of the original series on their common factors. For all three variables we use two common components, which is enough to capture more than 50 percent of the overall variance. Any further component would add only a small proportion and the evidence shows that results do not qualitatively change.
As a second step, the common factors and idiosyncratic components are tested seperately for unit roots and cointegration relationships, assuming the following form of the underlying variables

$$Y_{i,t} = \xi_{1i} F_{1t} + E_{1i,t} \tag{3}$$
$$X_{i,t} = \xi_{2i} F_{2t} + E_{2i,t}, \tag{4}$$

where F denotes the common factors and E the idiosyncratic components of the respective variables. A cointegration relationship between Y and X

$$Y_{i,t} - \beta_i X_{i,t} = \xi_{1i} \left(F_{1t} - \beta_i \frac{\xi_{2i}}{\xi_{1i}} F_{2t} \right) + E_{1i,t} - \beta_i E_{2i,t} \tag{5}$$

requires that the null hypothesis of no cointegration can be rejected for both the common and the idiosyncratic components. If the common factors are I(1), while the idiosyncratic components are I(0), the non-stationarity in the panel will be driven entirely by a reduced number of international stochastic trends. In that case, cointegration between the series requires that the common factors of the variables cointegrate (Dreger and Reimers, 2009).

3.2 Unit root tests

In the analysis of the common components of energy consumption, real GDP per capita and energy prices, standard time series unit root tests can be applied. To ensure robustness we use several unit root tests, including the augmented Dickey and Fuller (1979) (ADF) test, the Phillips and Perron (1988) (PP) test, as well as the Kwiatkowski, Phillips, Schmidt, and Shin (1992) (KPPS) test. The latter tests the null of stationarity whereas the former two investigate the null of a unit root. We do not further discuss the details of these well-known time series unit root tests but instead call attention to Maddala and Kim (1998) for their excellent treatment of ADF, PP and KPSS.[5] According to our results, displayed in Table 2, the common components of energy consumption, real GDP per capita and energy prices all turn out to be integrated of order one, I(1).

Table 2: Time series unit root tests for common components of energy consumption, GDP and energy prices

Variable	ADF	PP	KPSS
E	-2.67(0)	-3.05[2]	0.17[3]**
ΔE	-3.08(2)***	-3.09[4]***	0.15[2]
Y	-2.61(1)	-1.72[2]	0.14[3]*
ΔY	-2.47(0)**	-2.47[0]**	0.23[2]
P	-0.56(0)	0.79[1]	0.19[3]**
ΔP	-3.73(0)***	-3.71[3]***	0.25[1]

Notes: Δ denotes first differences. Numbers in parentheses are lag levels based on the Schwarz Information Criterion. Numbers in brackets represents the automatic Newey-West bandwidth selection using the Bartlett kernel. We include a constant and a linear time trend to test the variables in level form for unit roots and we include neither a constant nor a linear time trend to test the variables in differenced form with the exception of KPPS, which always includes a constant. ***, ** and * indicate significance at the 1%, 5% and 10% levels.

Since the defactored series are independent by construction, stochastic trends in the idiosyncratic components are efficiently explored by first generation panel unit root tests to exploit the additional information provided by the cross-sectional data. We apply the tests suggested by Levin, Lin, and Chu (2002) (LLC) and Im, Pesaran, and Shin (2003) (IPS). The underlying autoregressive model can be expressed as follows:

$$y_{i,t} = \rho_i y_{i,t-1} + \delta_i X_{i,t} + \varepsilon_{i,t}, \tag{6}$$

where $i = 1, ..., N$ and $t = 1, ..., T$ represent panel members and time periods, respectively. $X_{i,t}$ refers to the predetermined variables including any fixed effects and individual time trends, ρ_i are the autoregressive coefficients and $\varepsilon_{i,t}$ represents a white noise error process. If $|\rho_i| < 1$, $y_{i,t}$ behaves as a weakly (trend-)stationary process. In contrast, if $|\rho_i| = 1$, $y_{i,t}$ contains a unit root.

The LLC test examines the null hypothesis that each individual time series contains a unit root against the alternative that each time series is (trend-)stationary. In a first step, the LLC test performs separate ADF regressions for each cross-section

$$\Delta y_{i,t} = \rho_i y_{i,t-1} + \sum_{k=1}^{p_i} \varphi_{i,k} \Delta y_{i,t-k} + \delta_i X_{i,t} + \varepsilon_{i,t}, \tag{7}$$

[5]Further details of the ADF tests can also be found in Harris and Sollis (2003).

where Δ is the first-difference operator. The lag order, p_i, is allowed to vary across the i cross-sections. In a second step, the LLC test computes pooled t-statistics to check for the null hypothesis of non-stationarity. The test t-values are asymptotically distributed under the standard normal distribution. A limitation of the LLC test is its assumption that all cross-sections have the same first order autoregressive parameter, i.e. $\rho_i = \rho$.

The IPS test relaxes this assumption by allowing heterogeneity in this coefficient for all cross-section units. The null hypothesis is that the series under study is non-stationary for all panel members. The alternative is not that all processes are stationary, such as in the case of the LLC test, but that at least one individual process is stationary. As a first step, separate ADF regressions are obtained from estimating (7), allowing for different lag orders across cross-sections. As a second step, the average of the individual t-statistics from the ADF regressions

$$\bar{t} = \frac{1}{N}\sum_{i=1}^{N} t_{\rho_i}, \tag{8}$$

is adjusted to arrive at the desired test statistic. IPS standardize their test statistic based on simulations of the mean and variance of the t_ρ series

$$z\left(\bar{t}\right) = \frac{\sqrt{N}\left(\bar{t} - E(t_\rho)\right)}{\sqrt{Var(t_\rho)}} \tag{9}$$

and show that the z-statistic has an asymptotic standard normal distribution. The simulated values of the mean and the variance are tabulated in Im et al. (2003).

Although we prefer the IPS tests, our study also reports the results of the LLC test to provide an additional check for robustness. In contrast to the time series unit root evidence for the common components, the LLC and IPS panel unit root tests propose that the idiosyncratic components are widely stationary (see Table 3).

Table 3: Panel unit root tests for the idiosyncratic components of energy consumption, GDP and energy prices

Variable	LLC	IPS
E	-1.94**	-3.33***
Y	-2.45***	-2.19**
P	-1.36*	-3.15***

Notes: Probabilities were computed on the assumption of asymptotic normality. The choice of lag levels is based on the Schwarz Information Criterion. No time trend, only a constant, is included. The LLC tests were computed using the Bartlett kernel with automatic bandwidth selection. ***, ** and * indicate significance at the 1%, 5% and 10% levels.

Hence, the results suggest that random walks in the data are driven mainly by international developments. As a consequence, cointegration i.e. a long-run relationship may exist between the common rather than the idiosyncratic components.

3.3 Cointegration analysis

As integration of order one is established for the common components of the variables under investigation, the next step is to determine whether a long-run relationship exists.[6] Cointegration between the common components can be investigated using standard time series tests such as the Johansen reduced rank approach (Johansen, 1995). As aforementioned, small sample size can induce biased Johansen test statistics. Hence, we apply the small sample modification proposed by Reinsel and Ahn (1992) and Reimers (1992), who suggest the multiplication of the Johansen statistics with the scale factor $(T - pk)/T$, where T is the number of observations, p the number of variables and k the lag order of the VAR. This approach corrects for small sample bias such that a proper inference can be made. The empirical realisations of the modified Johansen trace statistic as well as those of the modified Johansen maximum eigenvalue statistic suggest evidence in favour of a long-run relationship between the common factors of energy consumption, real GDP per capita and energy prices (see Table 4).

Table 4: Results of Johansen's tests for cointegration among common components

H_0	Trace Statistic	Critical Value	Max. Eigenvalue Statistic	Critical Value
None	47.02*	42.92	28.19*	25.82
At most 1	18.83	25.87	13.35	19.39
At most 2	5.48	12.52	5.48	12.52

Notes: Potential small sample bias is corrected by multiplying the Johansen statistics with the scale factor $(T - pk)/T$, where T is the number of observations, p the number of variables and k the lag order of the underlying VAR model in levels, see Reinsel and Ahn (1992) and Reimers (1992). Critical values are taken from MacKinnon et al. (1999), and are also valid for the small sample correction. A * indicates the rejection of the null hypothesis of no cointegration at least at the 5% level of significance.

As a next step, we estimate the long-run relationships using the dynamic ordinary least squares (DOLS) estimator proposed by Mark and Sul (2003). The DOLS estimator corrects standard OLS for bias induced by endogeneity and serial correlation. First, the endogenous variable in each equation is regressed on the leads and lags of the first-differenced regressors from all equations to control for potential endogeneities. Then the OLS method is applied using the residuals from the first step regression. The DOLS estimator is preferred to the non-parametric FMOLS estimator because of its better performance. According to Wagner and Hlouskova (2010), the DOLS estimator outperforms all other studied estimators, both single equation estimators and system estimators, even for large samples. Furthermore, Harris and Sollis (2003) suggest that non-parametric approaches such as FMOLS are less robust if the data have significant outliers and also have problems in cases where the residuals have large negative moving average components, which is a fairly common occurrence in macro time series data. The estimated models are:

$$
\begin{aligned}
E_{i,t} &= \alpha_i + \delta_i t + \beta_i Y_{i,t} + \gamma_i P_{i,t} + \upsilon_{i,t} \\
Y_{i,t} &= \alpha_i + \delta_i t + \beta_i E_{i,t} + \gamma_i P_{i,t} + \varepsilon_{i,t} \\
P_{i,t} &= \alpha_i + \delta_i t + \beta_i E_{i,t} + \gamma_i Y_{i,t} + \eta_{i,t},
\end{aligned}
\tag{10}
$$

[6] Since the panel unit root tests of the idiosyncratic components suggest stationarity, we do not test for cointegration between the idiosyncratic components of energy consumption, GDP and energy prices.

where $i = 1, ..., N$ refers to each country in the panel and $t = 1, ..., T$ denotes the time period, α_i and δ_i are country-specific fixed effects and time trends, respectively. Since all variables are in natural logarithms, the estimated long-run coefficients can be interpreted as elasticities. The long-run income elasticity of energy consumption is 0.55, positive and statistically significant at the 1% level. This implies that a 1% increase in real GDP per capita increases total energy consumption by 0.6%. Energy consumption is relatively price-inelastic in view of a price elasticity of -0.14, which is statistically significant at the 1% level. The impact of energy prices on energy consumption is estimated to be negative, as expected from theory. Taking real GDP per capita as the dependent variable, income also turns out to increase by 0.6% if energy consumption grows by 1%. This elasticity is also significant at the 1% level. The price elasticity of income reveals a positive sign, but is insignificant as energy prices have no impact on real GDP per capita. The positive impact of income and energy consumption on each other implies that they are important determinants of each other.

A comparison with other studies reporting estimated long-run elasticities reveals that our empirical results are actually within the range of previous analyses. For instance, the very recent study by Lee and Lee (2010) also reports estimates of the income and price elasticities of energy demand for OECD countries. They come up with an estimated income elasticity of 0.52 and an estimated price elasticity of -0.19.

3.4 Dynamic panel causality

Having established a cointegration relationship, we estimate a panel-based error-correction model to test for Granger causality among energy consumption, real GDP per capita and energy prices. A two-step procedure is applied. First, the long-run equations specified in (10) are used to obtain the deviations from the long-run equilibrium (υ, ε and η). Then the error-correction model is estimated with the one-period lagged residuals from the first step as dynamic error-correction terms, as proposed in Holtz-Eakin et al. (1988):

$$\Delta E_{i,t} = \alpha_{1i} + \sum_{k=1}^{h} \theta_{11i,k} \Delta E_{i,t-k} + \sum_{k=1}^{h} \theta_{12i,k} \Delta Y_{i,t-k} + \sum_{k=1}^{h} \theta_{13i,k} \Delta P_{i,t-k} + \lambda_{1i} \upsilon_{i,t-1} + u_{1i,t} \tag{11}$$

$$\Delta Y_{i,t} = \alpha_{2i} + \sum_{k=1}^{h} \theta_{21i,k} \Delta E_{i,t-k} + \sum_{k=1}^{h} \theta_{22i,k} \Delta Y_{i,t-k} + \sum_{k=1}^{h} \theta_{23i,k} \Delta P_{i,t-k} + \lambda_{2i} \varepsilon_{i,t-1} + u_{2i,t} \tag{12}$$

$$\Delta P_{i,t} = \alpha_{3i} + \sum_{k=1}^{h} \theta_{31i,k} \Delta E_{i,t-k} + \sum_{k=1}^{h} \theta_{32ik} \Delta Y_{i,t-k} + \sum_{k=1}^{h} \theta_{33i,k} \Delta P_{i,t-k} + \lambda_{3i} \eta_{i,t-1} + u_{3i,t}, \tag{13}$$

where Δ is the first-difference operator, k is the lag length, λ_i is the speed of adjustment and $u_{i,t}$ is the serially uncorrelated error term with mean zero. The differenced form takes care of the OLS estimation problem, which is due to correlation between country-specific effects and explanatory variables. But differencing introduces the problem of simultaneity because the lagged dependent variables are correlated with the differenced error term. Additionally, heteroscedasticity in the errors across the cross-section members is expected to occur. We have to apply an instrumental variable estimator to cope with these problems. A widely used estimator for the system in (11) - (13) is the panel generalized method of moments (GMM) estimator proposed by Arellano and Bond (1991). Predetermined lags of the system variables are used as instruments to obtain consistent results. According to our empirical investigations, a lag length of $k = 2$ proves to be necessary to remove serial correlation in the error term. Hence, we employ variables lagged three and four periods as instruments for the lagged dependent variables.

The direction of causality can be determined by testing for the significance of the coefficients of each dependent variable in equations (11) to (13). First, to check for short-run causality we test $H_0 : \theta_{12ik} = 0, \forall ik$, and $H_0 : \theta_{13ik} = 0, \forall ik$, i.e. to detect whether causality runs from real GDP per capita and/or energy prices to energy consumption in (11). The underlying null hypotheses for testing whether short-run causality runs from energy consumption and/or energy prices to real GDP per capita in equation (12) are $H_0 : \theta_{21ik} = 0, \forall ik$, and $H_0 : \theta_{23ik} = 0, \forall ik$. Further, for short-run causality running from energy consumption and/or real GDP per capita to energy prices in (13) we test $H_0 : \theta_{31ik} = 0, \forall ik$, and $H_0 : \theta_{32ik} = 0, \forall ik$. Second, we check for long-run causality by testing the significance of the speed of adjustment, i.e. we test whether the coefficient of the respective error-correction term represented by λ is equal to zero. Finally, we test for strong causality by applying joint tests including the coefficients of the respective explanatory variables and the respective error-correction term of each equation (Y and P each with υ; E and P each with ε; E and Y each with η). This specific notion of causality denotes which variables bear the burden of a short-run adjustment to re-establish a long-run equilibrium, following a shock to the system (Asafu-Adjaye, 2000; Oh and Lee, 2004a,b). In the case of no causality in either direction the neutrality hypothesis holds. Since all variables are represented in stationary form we can use standard Wald F-tests when testing the various null hypotheses. Table 5 shows the results of our corresponding Granger causality tests.

Table 5: Panel causality test results for energy consumption, GDP and energy prices

Dependent variable	Sources of causation (independent variables)						
	Short-run			Long-run	Strong causality		
	ΔE	ΔY	ΔP	ECT	ΔE, ECT	ΔY, ECT	ΔP, ECT
ΔE	-	1.17	7.92**	4.33*	-	8.66**	5.33**
ΔY	3.57*	-	7.07**	8.03**	18.18**	-	6.91**
ΔP	4.59*	4.34*	-	15.63**	8.81**	5.57**	-

Notes: We report empirical realisations of the Wald F-test statistics. Potential heteroscedasticity of the error terms is corrected by using White robust standard errors. ECT represents the coefficient of the error-correction terms υ, ε and η, respectively. ** and * indicate that the null hypothesis of no causation is rejected at the 1% and 5% levels.

Our empirical exercise reveals that there are mutual causal relationships between ΔE, ΔY and ΔP in all three cases, i.e. short-run, long-run and strong causality. Energy consumption Granger-causes real GDP per capita and vice versa in the long run, which implies that an increase in energy consumption leads to an increase in economic growth and vice versa. In contrast, a rise in energy prices has a negative effect on energy consumption. Economic growth and energy consumption also have an impact on energy prices. Further, the significance of all error correction terms (ECT) indicates that all three variables readjust towards a common international equilibrium relationship after a shock occurs.

Since in the OECD countries a large portion of energy prices is related to energy taxes, energy regulations in terms of taxes will possibly have a negative impact on energy consumption and economic growth. Simultaneously, significant changes in economic growth and energy use patterns can influence the development of energy prices.

With respect to the widely studied energy consumption-growth nexus, a bi-directional causal relationship between energy consumption and economic growth is also reported by Masih and Masih (1997), Glasure and Lee (1998), Asafu-Adjaye (2000), Ghali and El-Sakka (2004), Lee et al. (2008), Apergis and Payne (2009a) as well as by Lee and Lee (2010). However, compared with other previous studies our findings contradict, on the one hand, those of Kraft and Kraft (1978) and Al-Iriani (2006), who found a uni-directional causal relationship running from economic growth to energy consumption, and, on the other, those of Lee (2005), Lee and Chang (2008) and Apergis and Payne (2009b), who inferred that causality runs from energy consumption to economic growth. Further, our empirical results also refute the neutrality hypothesis as proposed by Erol and Yu (1987), Yu and Jin (1992) and Masih and Masih (1996).

4 Conclusions and policy implications

In our contribution, we study the causal relationship between energy consumption and economic growth for 25 OECD countries from 1981 to 2007, explicitly taking into account the role of energy prices. We provide new empirical insights into the long-run relationship among these variables by applying factor decomposition to distinguish between common factors and idiosyncratic components as potential drivers of this relationship. The distinction between common factors and idiosyncratic components has important policy implications. Cointegration between the common components suggests that international developments dominate the long-run relationship whereas cointegration between idiosyncratic components relates to developments relevant on the national level. Hence, national energy policies may not have a large impact if international developments dominate the relationship between energy consumption, economic growth and energy prices.

Indeed, our main empirical finding is that only the *common* components of energy consumption, economic growth and energy prices are cointegrated. This result highlights the relevance of international developments to explain energy demand. Hence, policy makers should take into account the international impact on energy demand when designing efficient energy policies. Additionally, energy companies need accurate information concerning energy demand in order to be able to predict the future requirements and to take account of the necessary capacity to satisfy future energy consumption.

Further analysis of the cointegration relationships suggests that energy consumption is relatively price-inelastic. This underlines the theoretical expectation that energy use is mostly a necessity. The established long-run causality in the energy demand equation means that energy consumption readjusts after a shock towards an international rather than a national equilibrium relationship. In this light, national energy policies may have only a limited impact on energy consumption. The same holds for economic growth, such that national energy conservation policies may not have a large impact on economic growth either. What is more, bi-directional causality between energy consumption and economic growth in the long run suggests that no variable leads the other. An increase in energy consumption leads to an increase in economic growth and vice versa. Hence, it seems that OECD countries exhibit a kind of energy-dependence in the sense that an adequately large supply of energy seems to ensure higher economic growth (Lee and Lee, 2010). The bi-directional causal relationship indicates that the feedback hypothesis holds. This suggests that energy consumption and economic growth are interrelated. If this is true, the design of efficient energy conservation policies should imply the consideration of the direct impact of energy consumption on economic growth and the feedback effect of economic growth on energy consumption. In order to ease the trade-off between energy consumption and economic growth, energy policies devoted to a reduction in greenhouse gas

emissions should emphasise the use of alternative energy sources rather than exclusively try to reduce overall energy consumption. The shift from less efficient and more polluting energy sources to more efficient energy options may establish a stimulus rather than an obstacle to economic development (Costantini and Martini, 2010).
One main task of energy policy is the conservation of energy which means a more efficient use of energy and a reduction in greenhouse gas emissions using alternative energy options. In order to achieve these ambitious objectives, it should be noted that efficient energy conservation policies cannot be designed without considering other economic and environmental factors than the underlying variables in our study. Furthermore, according to the results of our study, not only national factors such as energy supply infrastructure, energy efficiency considerations or institutional constraints, but also international developments should be taken into account in the future.

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