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Modeling industrial energy demand in Greece using cointegration techniques

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Abstract

This paper attempts to shed light on the determinants of industrial energy demand in Greece. For this purpose we used cointegration analysis in order to capture short-run and long-run elasticities for oil and electricity industrial demand, respectively. The sample spans the period 1970–2004. From the empirical analysis and the Johansen's maximum likelihood procedure we found cointegration for oil and electricity demand. The results suggest that industrial energy demand is inelastic both in the short and the long run, while electricity and oil are substitutes. Also, oil and electricity prices are weakly exogenous both in the short and the long run.

Keywords: Elasticities; Long-run relationship; Weakly exogenous

1. Introduction and related literature

Industrial energy consumption accounts roughly for a third of total energy consumption in the world. According to the International Energy Agency (IEA) the share of global industrial consumption in total final energy consumption (transportation, agriculture, residential sector, etc.) is estimated nearly to 32%. During the last two decades, it has shown a significant decrease. The main reasons for this are connected to the decrease of energy intensity of the economy and conservation. The former was followed by a rapid change in economic policy mix in most of the developed countries targeted at the provision of investment incentives in the tertiary sector and the increase in services sector expenses as a share of GDP (de-industrialization process). In other words, the decrease of the energy intensity of the economy as a whole is a result of the growth of the GDP share of services. While on the other hand, the rapid decline of the industrial energy intensity might be the result of shifts towards less energy intensive sectors (e.g. from petrochemicals to pharmaceuticals). Finally, energy conservation took the form of employing energy-saving technologies in the industrial production process that are more enhanced and efficient.

Within the last decade, industrial energy demand has shown modest growth rates compared to other uses of energy in the Greek economy (transport or residential sector). It is estimated that within the period 1990–2003 industrial energy demand has increased by 9% compared to 15.2% for the period 1985–2003. Industry sector in Greece is one of the major polluting ones. However, greenhouse gas emissions (GHG) show a decreasing trend until 1994 due to the reduction of industrial activity in Greece, which is followed, despite some fluctuations by an increasing trend (MINENV, 2004). More specifically, in 2002, GHG emissions, derived from manufacturing has shown an increase of 3.7% compared to 1990, with an average annual rate of 0.3% for the period 1990–2002. Industrial energy demand is a key category of carbon dioxide emissions (CO_2) a major cause of the greenhouse effect. CO_2 emissions in 2002 increased by approximately 3.6% compared to 1990 emissions, while CH_4 (methane) emissions almost tripled from 1990 (MINENV, 2004).

The reasons for modeling (industrial) energy consumption are as varied as the ways in which energy is used in production process, making it one of the hardest end-uses

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to analyze, model and forecast. Several authors have attempted to examine industrial energy demand (Andrikopoulos and Vlachou, 1995; Beenstock et al., 1999; Bose and Shukla, 1999; Christopoulos, 2000; Christopoulos and Tsionas, 2002; Donatos and Mergos, 1989; Floros and Vlachou, 2005; Kamerschen and Porter, 2004; Urga, 1999; Vlachou et al., 1996). The economic approach adapted in most of these studies is related to the estimation of a system of factor demand equations derived from a typically generalized translog cost function, originally developed by Christensen et al. (1973) and extensively applied in studies investigating the energy demand of industry. For the estimation of this function, the iterative Zellner method or the seemingly unrelated regression estimation (SURE) is used. This method, which is equivalent to maximum likelihood estimation, gives consistent and asymptotically efficient estimates (Floros and Vlachou, 2005).

The empirical findings from the decomposition of industrial energy demand in Greece indicate that the demand for each main energy inputs (i.e. electricity, liquid fuels, solid fuels) is inelastic (Andrikopoulos and Vlachou, 1995; Christopoulos, 2000; Floros and Vlachou, 2005; Vlachou et al., 1996). In the recent study of Floros and Vlachou (2005), own-price elasticities for all the industrial sectors range from −0.076 (basic metal industry) to −0.830 (wearing apparel) while the demand for diesel exhibits the highest price responsiveness among the three energy types (electricity, diesel and heavy fuel oil).[1] The cross-price elasticities suggest that electricity and diesel and electricity and heavy fuel oil are substitutes, while diesel and heavy fuel oil are complement in most manufacturing sectors. The same results hold for similar studies. Christopoulos (2000) found an elastic demand for diesel and substitutability between diesel and electricity by using data for the period 1970–1990. However, the own-price elasticity for electricity was found to be inelastic but not statistically different from zero.

An alternative econometric approach is followed in the studies of Beenstock et al. (1999); Bose and Shukla (1999) and Kamerschen and Porter (2004), which is consisted of applying cointegration techniques (Engle-Granger and Johansen methodology) in order to estimate own and cross-price elasticities. In the study of Kamerschen and Porter (2004), two alternative econometric models were used (partial adjustment model and simultaneous equations model) in order to assess residential and industrial electricity demand in the USA for the period 1973–1998. The results of this study suggest that simultaneous equation models, which are estimated by using three stages least squares method, are more appropriate since they provide negative price elasticity estimates for the residential, industrial and total electricity samples. The industrial own-price estimates for electricity, range from −0.34 to −0.55, while cross-price elasticities do not exceed 0.13 revealing a substitutability relationship between electricity and natural gas in the industrial sector.

In order to estimate the relevant elasticities of industrial electricity demand in Israel, Beenstock et al. (1999), employ three different econometric models (dynamic regression, error-correction and vector autoregression model) by using quarterly data for the period (1973–1994). The parameter estimates of the three econometric models do not vary systematically suggesting that in the long-run, electricity demand is inelastic to fluctuations in electricity price with the magnitude of these estimates ranging from −0.002 (ECM) to −0.435 (VAR).

A different approach is followed in the study of Bose and Shukla (1999). This paper examines the econometric relationship between industrial electricity consumption (small and medium industries, and large industries) and its main determinants (income, price of electricity and diesel) for India.[2] The results show that electricity consumption in large industrial sectors are income elastic (>1), while in small and medium industries are income inelastic (<1). The short-run price elasticities vary from −0.45 in large industry and −0.26 in small and medium industry.

The purpose of this paper is to examine the main factors that determine the industrial energy demand in Greece and on the basis of our findings to attempt to draw some conclusions on the policy choices that the country has to make. The recent oil price shock, the compliance with the Kyoto target (+25% for the period 2008–2012), the restructuring of energy taxation are some of the main parameters that bring again to the fore the need for a more determined energy approach. In other words the government should implement measures facing the abovementioned issues. This paper differs from earlier works in the subject in two ways. First, our approach is quite different from that of some previous studies (see Andrikopoulos and Vlachou, 1995; Caloghirou et al., 1997; Floros and Vlachou, 2005; Vlachou et al., 1996), by taking into account the "spurious" regression issue.[3] Second, we try to examine if there are structural breaks in the industrial energy demand in Greece, to which rather scant attention has been paid by the earlier studies.

The paper consists of five sections. In Section 2 a brief overview of the industrial energy sector in Greece is provided, together with a short presentation of the level of taxation by type of fuel. Section 3 offers a detailed description of the model, while Section 4 reports the main results of the cointegration techniques. Finally, in Section 5 we summarize our findings and attempt to draw some policy implications.

[1] However, energy demand in few industrial sectors seems to be quite elastic (i.e. metal products, furniture, other transport equipment, etc.).

[2] The study uses a linear demand equation by pooling data across 19 states spread over nine years (1985/1986) and (1993/1994).

[3] In the studies mentioned before, the estimated relationships were not tested if they were meaningful in the long-run (cointegrated).

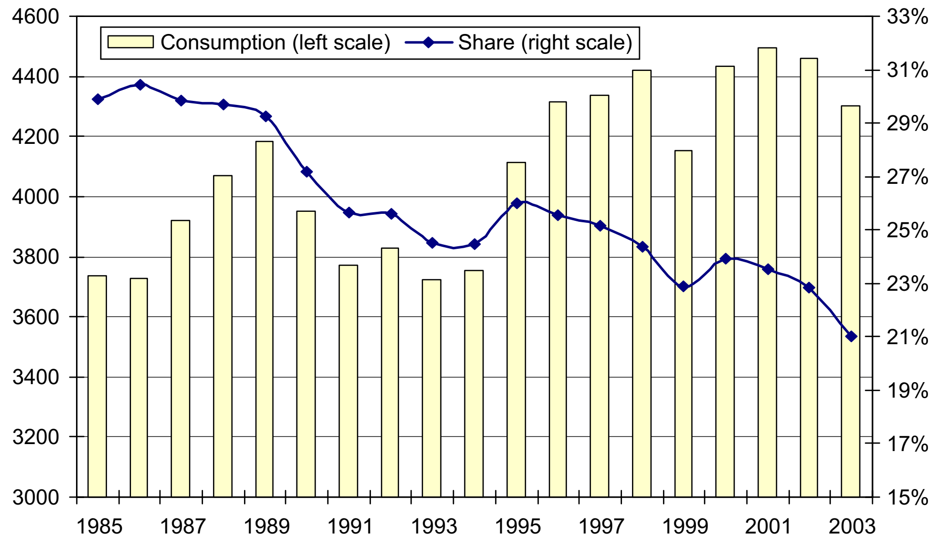

Fig. 1. Total final industrial consumption (000 toe) and share of industrial consumption in total energy consumption (1985–2003). *Source*: Eurostat—New Cronos Theme 8 (Environment and Energy)—www.europa.eu.int/newcronos.

2. The characteristics of industrial energy demand in Greece

Industry sector constitutes one of the most energy-intensive sectors in Greece with a total final consumption of 4.3 million tonnes of oil equivalent (TOE) in 2003 (Fig. 1). Since the mid 1970s, a rise of energy consumption in industrial sector has been observed, faster than that of total energy consumption, mostly reflecting a rapid industrialization of the country and a shift from the agriculture sector to manufacture. Total industrial consumption in Greece has increased by 0.8% annually within the period 1985–2003, while total final energy consumption has shown an increase by about 2.8% within the same period.

The share of industrial consumption in total energy demand has shown a gradually decrease during the last 20 years reaching the level of 21% in 2003. Energy intensity has shown a decline in recent years as a result of an in-depth restructuring of the sector. This evolution consisted of significant increase in the level of fixed capital investment—boosted by the Community Support Framework III—that induced a technological change and created a decline trend in the level of average industrial energy consumption (RAE, 2003).[4]

The excise tax on heavy fuel oil (low sulphur fuel oil) has increased within the period 1978–2003, by about 16.6% annually, while the excise tax levied on light fuel oil has shown an increase for the same period by about 15.3% per annum reaching 120 euros/1000l in 2003. It is worth mentioning that electricity for industry is not subject to any form of taxation (excise tax or VAT). Energy taxation has varied significantly during the previous mentioned periods, since it was used by the Greek Government not only as an important source of state revenues, but in certain circumstances also as an anti-inflationary tool, which was quite high over the previous decades (Rapanos and Polemis, 2005).

Greek manufacturing mainly consumes oil products (diesel oil and heavy fuel oil) and electricity. This proportion may change significantly within the next few years due to the liberalization of electricity and natural gas markets and the need to comply with environmental regulations (Kyoto protocol). The biggest share of industrial consumption in 2004 corresponds to crude oil and petroleum products (45%) and follows electricity (30%) which is mainly used by the basic metal industries. Solid fuels hold the 13% of the total final consumption while the share of natural gas in total demand of the sector does not exceed 12%. However, a historical substitution from solid fuels to oil and electricity occurred during the first years of the reference period. Although it may be minimal at the end of the estimation period, it had been very significant in the early part of the sample (especially through the 1980s) and was a major factor in explaining industrial energy demand in Greece.

According to Table 1, almost all industrial sectors were oil-intensive in 2003, while many of these sectors are also dependent on electricity (i.e. iron and steel industry, engineering and other metal industry and textile, leather and clothing industry). This seems to be in part explained by the small use of natural gas in energy balance of Greece due to the fact that it has been introduced to the energy mix of Greece since 1997.[5] It might also indicate substitutability between electricity and liquid fuels over the period.

[4]More efficient technologies as well as heat pumps played a catalytic role in industrial energy conservation.

[5]According to the provisions of the Gas Directive (2003/55), Greece is an emerging gas market and has been granted derogation from the implementation of this Directive until November 2006. Therefore the Greek gas market is currently operating in a monopolistic way, under the provisions of law 2364/1995 (RAE, 2005).

Table 1
Energy structure of Greek industrial sectors (2003)

Sectors	Sector's share in industrial consumption (%)	Shares (%)[a]			
		Electricity	Oil products	Natural gas	Solid fuels
Iron and steel industry	5.8	53.8	22.3	23.9	0.0
Non-ferrous metal industries	19.3	38.9	32.7	6.3	22.1
Chemical industry	4.9	25.6	64.0	10.4	0.0
Non-metallic mineral products industry	25.8	16.1	41.5	5.4	37.0
Ore extraction industry	2.0	9.3	90.7	0.0	0.0
Food drink and tobacco industry	16.1	31.2	55.4	13.4	0.0
Textile, leather and clothing industry	4.4	45.3	38.9	15.8	0.0
Paper and printing industry	4.0	28.6	55.4	16.0	0.0
Engineering and other metal industry	2.9	72.6	24.2	0.8	2.4
Other non-classified industries	14.8	21.7	76.9	1.5	0.0
Total	100.0	—	—	—	—

Source: Eurostat—New Cronos Theme 8 (Environment and Energy)—www.europa.eu.int/newcronos.
[a]Fuel share in total energy consumption of the sector.

However, since a high percentage (nearly 65%) of electricity is generated by lignite-fired stations, high electricity consumption by industrial sectors tends to be the cause of the increase of CO_2 emissions (Floros and Vlachou, 2005).

3. Data description and empirical methodology

The data used in the empirical estimation are annual time series that cover the period 1970–2004. Electricity consumption (ELEC) is measured in kilowatt hours (KWH) and oil consumption (OIL) is expressed in metric tonnes (MT). Natural gas appears to be used in manufacturing since 1997 so that we had to exclude it from our empirical analysis. These data are available from the Ministry of Development and the IEA.[6] Industrial production index is available from the National Statistical Service of Greece (NSSG).

Energy prices for electricity (RPELEC) and oil (RPOIL) are taken from "Energy Prices and Taxes" (IEA) and have been deflated by the wholesale price index. Moreover, the variable of heating degree days (HDD) captures the temperature effect on oil and electricity consumption and is available from the IEA database. The HDD variable may not be very relevant for industrial electricity demand as it comes out as insignificant in most estimations of the paper (see Section 4). However, due to the fact that data for cooling degree-days are missing, we chose to keep the relevant variable as a proxy for variations in temperature. Finally, the number of medium voltage electricity consumers (CONS) is taken from the Public Power Corporation (PPC).

[6]The data for the period 1978–2004 on industrial energy demand come from "Energy Balances in OECD Countries". The data prior to the above-mentioned period (1970–1978) are taken from the Greek Ministry of Development.

Industrial energy demand decomposed in two distinct sub-sectors (oil and electricity) can be modeled with two autoregressive distributed lag specifications:

$$a(L)\mathrm{OIL}_t = \mu + \beta(L)\mathrm{DBP}_t + \gamma(L)\mathrm{RPOIL}_t + \delta(L)\mathrm{RPELEC}_t + u_t, \quad (1)$$

$$c(L)\mathrm{ELEC}_t = \varepsilon + \zeta(L)\mathrm{DBP}_t + \eta(L)\mathrm{RPELEC}_t + \theta(L)\mathrm{RPOIL}_t + \lambda(L)\mathrm{CONS}_t + e_t, \quad (2)$$

where L is the lag operator,

$$a(L) = 1 - a_1 L - \cdots - a_1 L^{\mathrm{O}},$$

$$c(L) = 1 - c_1 L - \cdots - c_1 L^{\mathrm{P}},$$

$$\beta(L) = 1 + \beta_1 L + \cdots + \beta_{\mathrm{Q}} L^{\mathrm{Q}},$$

$$\gamma(L) = 1 + \gamma_1 L + \cdots + \gamma_{\mathrm{R}} L^{\mathrm{R}},$$

$$\delta(L) = 1 + \delta_1 L + \cdots + \delta_{\mathrm{S}} L^{\mathrm{S}},$$

$$\zeta(L) = 1 + \zeta_1 L + \cdots + \zeta_{\varepsilon} L^{\mathrm{T}},$$

$$\eta(L) = 1 + \eta_1 L + \cdots + \eta_{\varepsilon} L^{\mathrm{X}},$$

$$\theta(L) = 1 + \theta_1 L + \cdots + \theta_{\varepsilon} L^{\mathrm{Y}},$$

$$\lambda(L) = 1 + \lambda_1 L + \cdots + \lambda_{\varepsilon} L^{\mathrm{Z}}.$$

Capital letters O, P, Q, R, S, T, X, Y and Z represent the optimal number of lags of the polynomials $\alpha(L)$, $c(L)$, $\beta(L)$, $\gamma(L)$, $\delta(L)$, $\zeta(L)$, $\eta(L)$, $\theta(L)$ and $\lambda(L)$, respectively. With OIL_t and ELEC_t we indicate oil and electricity consumption, respectively, DBP_t is the industrial production index. RPOIL_t and RPELEC_t represent real price of oil and electricity, respectively, while CONS is the number of medium voltage electricity consumers; u_t and e_t are the disturbance terms, respectively. All variables are log-transformed.

Having specified the two ARDL models, we test for the existence of a long-run relationship among the variables of the two models (Eqs. (1) and (2)) applying cointegration techniques (Johansen procedure).[7]

In order to capture the short-run dynamics (elasticities) we use estimates of the two error-correction models (oil and electricity demand) that help us to understand the exogeneity of the variables entered the models.

The basic statistical assumption underlying this approach is that the variables are stationary with the first two moments of the underlying data generation process not depending on time. In other words, we have to check for the presence of unit roots because if variables are non-stationary $I(1)$ processes, then there may exist a linear combination which may well be stationary $I(0)$ process. If this is the case then the variables are cointegrated and long-run effects (elasticities) can be estimated. For this reason, we employed different tests for the unit root hypothesis (ADF, Phillips–Perron and Kwiatkowski–Phillips–Schmidt–Shin or KPSS tests).[8]

However, since the existence of structural breaks will be investigated, we must mention that unit root tests that do not account for structural breaks may falsely fail to reject the unit root null hypothesis against the trend stationary alternative when the data generating process is trend stationary with a one-time break. This means that when a break is found, it is necessary to conduct the appropriate unit root tests instead of convention Augmented Dickey Fuller (ADF) and Phillips–Perron tests.[9] As it will be shown we did not find structural breaks in industrial energy demand in Greece. As a result the usual unit root tests (ADF and Phillips–Perron) are quite valid and appropriate.

Looking at the data over the examined period, we observe that our variables are probably non-stationary $I(1)$. In order to examine the order of integration we have applied a series of diagnostic tests both in levels and first differences of these variables.

The results of the above tests are presented in Table 2. Applying the relevant tests, we observe that the null-hypothesis of a unit root cannot be rejected at 5% critical value for almost all of the variables. In other words all the series—except for HDD—are non-stationary. For HDD, however, the unit root hypothesis is strongly rejected so that the variable is stationary or integrated $I(0)$.

From the results of Dickey–Fuller test, we observe that the levels of the variable CONS appear to be stationary at $a = 0.05$ level of significance. However, the combined results of Philips–Perron and KPSS tests show little support for the stationary hypothesis. By taking first differences of the non-stationary variables the hypothesis of stationarity cannot be rejected at $a = 0.05$ level of significance. Therefore, all the variables except for HDD are non-stationary $I(1)$.[10] This implies that there can be no long-run effect from the number of HDD on industrial energy demand, but this does not preclude the existence of short-run effect (Bentzen and Engsted, 1993; Rapanos and Polemis, 2005).[11]

Although the relevant variables are not stationary $I(1)$, the above-mentioned series (oil and electricity demand) may be cointegrated. Testing for cointegration is equivalent to test whether the residuals of the static (long-run) regressions are $I(0)$. In order to assess industrial consumption decomposed by type of energy (oil and electricity) we estimated two VAR's for oil and electricity demand, respectively. The reason for using cointegration techniques is that non-stationary time series result to spurious regressions and hence do not allow statistical interpretation of the estimations (Charemza and Deadman, 1997). In order to overcome this problem, we apply the Johansen, technique (see Johansen, 1988, 1992). Johansen's method has enjoyed widespread adoption within the last years since it is main advantage is that it allows estimation of multiple cointegrating vectors where they exists (Vita De et al., 2006). In other words it allows us to examine whether there is a long run co-movement of the variables.

Our next step is to choose the lag value of the multivariate system of equations. By setting the lag length equal to one we ensure that the residuals of the two VAR models are white noise (Gausian errors).[12] Log likelihood ratio tests are used for testing the deletion of two dummy variables from the two VAR models. The first dummy variable (D1986) reflects the rapid decrease of the international price of oil and the second dummy (D1997) accounts for the introduction of natural gas in the industry sector.[13] The one-time dummy variable for 1986 helps in the cointegration analysis in order to filter out a large fall in oil prices which might conceal the long-run price effects on industrial energy use, while the other dummy variable

[7]The static (long-run) regressions for oil and electricity industrial demand can be represented by the following equations: $\mathrm{OIL}_t = a_0 + a_1\mathrm{DBP}_t + a_2\mathrm{RPOIL}_t + a_3\mathrm{RPELEC}_t + e_t$ and $\mathrm{ELEC}_t = b_0 + b_1\mathrm{DBP}_t + b_2\mathrm{RPELEC}_t + b_3\mathrm{RPOIL}_t + b_4\mathrm{CONS}_t + u_t$.

[8]For the empirical estimation of the models the econometric package Econometric Views (version 5) was used.

[9]Perron, proposes three different formulations for unit root testing depending on the nature of the break: (a) if it is a one-time change in level of the test variable, (b) if there is only a change in the slope or (c) both a change in level and slope. Of course there are also other unit root tests that accommodate structural breaks and even multiple breaks as suggested by Lumsdaine and Papell (1997).

[10]According to Φ_2 statistic all the examined variables except RPOIL include deterministic terms (trend and intercept).

[11]However, we must bear in mind that the size and power of unit root tests is low in small samples and similar concerns are valid for cointegration analysis as well (Harris and Sollis, 2003). Since the data on industrial energy demand in Greece are lacking the analysis covers a short period (1970–2004).

[12]To determine the lag length of the VAR's, an extensive diagnosing testing of the OLS residuals is employed for various lag lengths. Each equation of the VAR systems passes a series of diagnostic tests including serial correlation based on the autocorrelation functions of the residuals as well as the reported Langrange multiplier (LM test).

[13]The dummies D1986 and D1997 take the value of 1 in the corresponding year and 0 in all other years, respectively.

Table 2
Tests for unit roots

Variable	(Augmented) Dickey–Fuller						Phillips–Perron		KPSS				Order of integration
	Lags	τ_t	Φ_3	τ_μ	Φ_2	τ	τ_t	τ_μ	$l=1$		$l=4$		
									n_μ	η_τ	n_μ	η_τ	
Levels													
DBP	0	−2.40	3.05	−1.58	6.24*		−2.53	−1.88	0.83**	0.13*	0.47*	0.09	$I(1)$
ELEC	0	−3.52	7.21	−0.68	6.72*		−3.47*	0.24	1.26**	0.08	0.62*	0.08	$I(1)$
OIL	1	−3.08	5.95	−1.49	5.97*		−1.62	−1.81	0.81**	0.24**	0.44*	0.14*	$I(1)$
RPELEC	0	−3.53	6.01	1.32	5.75*		−3.47	−1.28	0.71*	0.25**	0.34	0.14*	$I(1)$
RPOIL	0	−2.72	4.13	−2.87	4.75	1.02	−2.75	−2.87	0.43*	0.20*	0.46*	0.08	$I(1)$
CONS	0	−3.42*	13.20**				−3.22	−2.92	1.33**	0.27**	0.63*	0.16*	$I(1)$
HDD	0	−4.52**	—	—	—	—	−4.71**	−3.87**	0.43	0.06	0.37	0.10	$I(0)$
First differences													
Δ(DBP)	0	−5.03**	—	—	—	—	−5.20**	−5.21**	0.07	0.07	0.09	0.09	$I(0)$
Δ(ELEC)	0	−6.36**	—	—	—	—	−13.1**	−12.19**	0.05	0.04	0.14	0.11	$I(0)$
Δ(OIL)	1	−3.63*	—	—	—	—	−3.35*	−3.16*	0.16	0.05	0.17	0.07	$I(0)$
Δ(RPELEC)	0	−3.75*	—	—	—	—	−3.62*	−3.03*	0.56*	0.22**	0.33	0.15	$I(0)$
Δ(RPOIL)	0	−3.73*	—	—	—	—	−3.63*	−3.76**	0.13	0.10	0.16	0.12	$I(0)$
Δ(CONS)	0	−3.62*	—	—	—	—	−3.62*	−3.01*	0.33	0.16*	0.49*	0.11	$I(0)$
Δ(HDD)	—	—	—	—	—	—	—	—	—	—	—	—	—

Notes: The relevant tests are derived from the OLS estimation of the following autoregression for the variable involved:

$$\Delta Y_t = \delta + \beta Y_{t-1} + \gamma t + \sum \alpha_i^* \Delta Y_{t-i} + u_t. \tag{1a}$$

τ_μ is the t-statistic for testing the significance of β when a time trend is not included in Eq. (1) and τ_τ is the t-statistic for testing the significance of β when a time trend is included in Eq. (1). The calculated statistics are those reported in Dickey and Fuller (1981). The critical values at 5% and 1% for $N = 50$ are given in Dickey and Fuller (1981). The lag length structure of α_i of the dependent variable y_t is determined using a recursive procedure in the light of a Lagrange multiplier LM autocorrelation test for orders up to 4 which is asymptotically distributed as chi-squared distribution and the value of t-statistic of the coefficient associated with the last lag in the estimated autoregression. The critical values for the Phillips Perron unit root tests are obtained from Dickey and Fuller (1981). n_μ and n_τ are the KPSS statistics for testing the null hypothesis that the series are $I(0)$ when the residuals are computed from a regression equation with only an intercept and intercept and time trend, respectively. The critical values are given in Kwiatkowski et al. (1992).

*Indicates significance at the 1% level.

**Indicates significance at the 5% level.

(D1997) encapsulates the rapid introduction of natural gas from industries. The reason for not using the price of natural gas as an exogenous variable is that there is not available an extended data set. All the relevant tests cannot reject the null hypothesis of the deletion of the two dummy variables from the corresponding VAR models. Finally, according to Schwarz Criterion (SC) the estimation procedure for the VAR models assumes the existence of an intercept.

4. Econometric results

4.1. Maximum likelihood procedure

Table 3 presents the maximum likelihood eigenvalue statistics. It becomes clear from the table that the null hypothesis (no cointegration) is rejected at 1% level. The estimated likelihood ratio tests and eigenvalues indicate that there is one cointegrating vector for each model (oil and electricity).

Having specified the number of cointegrating relations for each VAR, we investigate whether all variables enter statistically significant in the cointegrating vectors. Table 4, reports the likelihood ratio tests as analytically described in Johansen and Juselius (1992). The results suggest that all the variables except for the number of medium voltage electricity consumers (CONS), enter statistically significant into the cointegrating vectors.

The resulting normalized parameter estimates of our oil model from the cointegration analysis are as follows:

$$\underset{}{\mathrm{OIL}} = \underset{(0.22)}{1.25\mathrm{DBP}} \quad \underset{(0.08)}{-0.18\mathrm{RPOIL}} \quad \underset{(0.03)}{+0.11\mathrm{RPELEC}} \quad \underset{(0.10)}{+8.12} + U, \qquad (3)$$

where the numbers in parentheses denote standard errors. From Eq. (4), it is evident that all the coefficients of the variables have the anticipated signs. More specifically, industrial oil demand (light and heavy fuel oil), seems to be highly elastic to the responses of industrial production index (DBP) with the relevant elasticity equals to 1.25. This means that a 1% increase of industrial production will lead to an increase of oil demand by 1.25%. Furthermore, an increase in real price level of oil leads to a decrease in the level of oil consumption, with the relevant (own price) elasticity equals to −0.18. This means that when the price of oil increases (decreases) by 1%, industrial oil consumption will marginally decrease (increase) by 0.18%. In addition, low price sensitivity means that, taxing oil products (e.g. diesel, heavy fuel oil, etc.) can be a good source of revenues in the long-run. The small magnitude of own price elasticity combined with the small value of cross-price elasticity of demand (0.11) indicates the absence of alternative energy sources that can be used as inputs in the

Table 4
LR test of restrictions that each variable does not enter in the cointegrating vector

Variables	Oil	Electricity
OIL	4.50[a]	—
ELEC	—	4.80[a]
DBP	7.58[b]	8.25[b]
RPOIL	6.98[b]	6.67[b]
RPELEC	9.85[b]	22.18[b]
CONS	—	0.11

LR tests are distributed as χ^2 distribution with degrees of freedom the number of cointegrating vectors.
[a]Rejection of the null hypothesis at $\alpha = 0.01$.
[b]Rejection of the null hypothesis at $\alpha = 0.05$.

Table 3
Johansen's maximum likelihood method test for cointegration relationship

Null hypothesis H_0	Alternative hypothesis H_1	Eigenvalue	Critical values	
			95%	99%
Oil demand: *VAR* = 2, *variables*: *OIL*, *DBP*, *RPOIL*, *RPELEC*				
Maximum eigenvalues				
$r = 0$	$r = 1$	37.21[a]	31.46	36.65
$r \leqslant 1$	$r = 2$	23.83	25.54	30.34
Trace statistic				
$r = 0$	$r \geqslant 1$	74.80[b]	62.99	70.05
$r \leqslant 1$	$r \geqslant 2$	38.74	42.44	48.45
$r \leqslant 2$	$r \geqslant 3$	21.75	25.32	30.45
Electricity demand: *VAR* = 1, *variables*: *ELEC*, *DBP*, *RPOIL*, *RPELEC*, *CONS*				
Maximum eigenvalues				
$r = 0$	$r = 1$	75.84[a]	37.52	42.36
$r \leqslant 1$	$r = 2$	18.60	31.46	36.65
Trace statistic				
$r = 0$	$r \geqslant 1$	117.39[a]	87.31	96.58
$r \leqslant 1$	$r \geqslant 2$	41.54	62.99	70.05

r: denotes the number of cointegrating equations.
[a]Significance at $\alpha = 0.01$.
[b]Significance at $\alpha = 0.05$.

industrial sectors (i.e. natural gas, renewable energy sources, etc.).

The estimation period for this study covers the somewhat volatile time of the international price of crude oil in 1986 and the introduction of natural gas for industry in 1997. Hence it is crucial to check the cointegration relationship (long-run) for the existence of structural breaks. For this purpose we apply econometric techniques to capture possible changes in the oil demand, such as recursive OLS (i.e. CUSUM tests, recursive residuals, one step forecast test and recursive coefficients). The results indicate the absence of structural break suggesting that the long-run industrial oil demand in Greece remained quite unchanged during the estimation period.[14]

To sum up, the estimated cointegrating vector of oil demand is affected by changes in industrial production and real price level of oil and electricity, respectively, while the magnitude of long-run oil demand elasticities does not deviate substantially from other empirical studies (Christopoulos, 2000; Floros and Vlachou, 2005; Vlachou et al., 1996). Moreover, no sign of structural shift was found in the long-run oil demand equation reflecting the low level of efficiency in the industrial energy sector in Greece.

Having specified the oil demand model we apply the VAR methodology to electricity model. The resulting normalized parameter estimates of our model are as follows:

$$\begin{array}{lllll} \mathrm{ELEC} & = 0.85\mathrm{DBP} & +0.18\mathrm{RPOIL} & -0.85\mathrm{RPELEC} & +3.65 + \mathrm{U}, \\ & (0.15) & (0.03) & (0.38) & (0.58) \end{array} \tag{4}$$

where the numbers in parentheses stand for the resulting standard errors. From the above normalized equation, several main points can be drawn. First, industrial electricity demand, seems to be inelastic to the responses of the industrial production (DBP) and the relevant elasticity is estimated to 0.85. Second, own price elasticity is negative (−0.85). This value is close to the one reported (−0.84) in the study of Christopoulos and Tsionas (2002), who estimated total industrial energy demand in Greece for the period 1970–1990. Finally, cross-price elasticity comes with a positive sign and less than unity (0.18). This finding supports the argument that there is a substitutability relationship between oil and electricity demand in the industrial process. The absence of close substitutes such as natural gas, LPG, bio-fuels, hydrogen that could have significant impact on the environment (i.e. combating CO_2 emissions) may support the argument that the scope of energy switching in industrial sector in Greek is still limited. However, the increasing use of natural gas in many industrial sectors within the next years as a result of the liberalization of the market might trigger a substitution effect between the industrial energy sources.

[14]Due to space limitations, the results are available from the author upon request.

Furthermore, the number of industrial consumers (medium voltage), does not affect industrial electricity demand since the relevant coefficient is not statistical significant and therefore omitted from the long-run equation. A possible explanation for this result might be that in Greece the average level of industrial electricity prices are distorted (OECD, 2001). This happens due to the fact that Public Power Corporation (PPC), which is a vertical integrated company into all aspects of the electricity sector in Greece (generation, transmission, distribution and supply), supplies large quantities of electricity at about half price to the selected industrial sectors such as basic metal industries, where commercial and small industrial customers connected with the medium voltage system pay prices well above their cost of supply (OECD, 2001).[15]

These results can be compared to previous studies (Christopoulos, 2000; Floros and Vlachou, 2005; Vlachou and Samouilidis, 1986) who estimated own and cross-price elasticities for electricity, liquid and solid fuels in Greek manufacturing. These studies found that demand for each energy type (electricity, diesel, heavy fuel oil, etc.) is inelastic. Moreover, they found that electricity and liquid fuels (diesel, heavy fuel oil) are substitutes in the major manufacturing sectors. This outcome comes in agreement with this study, denoting that in the last decades the share of electricity in total industrial energy demand has been raised significantly.[16]

The existence of structural stability in the long-run electricity demand equation can be found by using relevant tests that employ recursive OLS methodology. From their results, we concluded that in the long-run there is no sign of structural break in the electricity demand equation.[17]

Having specified long-run elasticities from the Johansen procedure, we opt to estimate vector error-correction models in order to obtain short-run responses (elasticities). Table 5 shows the results from the estimation of the two VECM's for industrial energy demand. Each coefficient of the variables denotes the short-run elasticity. The specific table does not include the results from the other equations (DBP, RPOIL, RPELEC), because only energy demand is of interest.

All the coefficients of the variables of the oil demand (column 1) are in alignment with the theory and statistical significant except the number of heating degree-days D(HDD). Industrial oil demand appears to be inelastic to fluctuations in the level of production output with the elasticity equals to 0.72.

The magnitude of this elasticity denotes that in the short-run industrial oil demand is changing more slowly

[15]This is so-called cross-subsidy in prices and shrinks the level of competition in electricity sector. However, these subsidies are due to be phased out soon.

[16]This result suggests that manufacturing could move away from expensive imported liquid fuels by using electricity, produced by indigenous sources like lignite or renewable energy sources (Floros and Vlachou, 2005).

[17]The results are available from the author upon request.

Table 5
Results of vector error-correction models (VECM)

Variables	D(OIL) (1)	D(ELEC) (2)
D(DBP)	0.72[a]	0.61[b]
	(1.74)	(4.19)
D(RPOIL)	−0.13[c]	0.18[b]
	(−1.87)	(3.65)
D(RPELEC)	0.07[c]	−0.35[b]
	(−2.12)	(−2.77)
D(OIL(-1))	0.94[b]	—
	(2.85)	
ECT_{t-1}	−0.77[a]	−0.24[b]
	(−1.85)	(−4.25)
Diagnostics		
R^2	0.53	0.64
Adjusted R^2	0.26	0.51
F-statistic	5.04	7.01
	[0.00]	[0.00]
Durbin–Watson	1.94	2.32
Breusch–Godfrey LM Test	0.05	2.23
	[0.81]	[0.15]
LM (5)	0.57	1.41
	[0.71]	[0.29]
White test	0.74	0.87
	[0.69]	[0.96]
J. Bera	2.71	3.96
	[0.45]	[0.13]
ARCH test	0.47	1.61
	[0.49]	[0.27]
LMARCH (5)	0.17	0.37
	[0.96]	[0.85]

Note: The numbers in parentheses and in the square brackets are *t* and *p* values, respectively. LM (5) and LMARCH (5) are Langrance multiplier tests for fifth order autocorrelation and fifth order autoregressive conditional heteroskedasticity (ARCH), respectively.

[a]Significance at 0.10.

[b]Significance at 0.01.

[c]Significance at 0.05.

compared to fluctuations in the level of production. In other words, oil demand and production output are positively linked to each other. The short-run elasticity with respect to own price is estimated to be less than unity (−0.13), while cross-price elasticity is 0.07. This means that oil and electricity are substitutes in the industrial production process. The above result is confirmed by other relevant studies (Christopoulos, 2000; Floros and Vlachou, 2005). However, the low level response of oil demand to its own price fluctuations reveals the relevant difficulty to substitute oil (light and heavy fuel oil) with other energy products (natural gas, renewable energy sources, etc.). In addition to, growth in oil consumption a year before the current consumption (DOIL(-1)) has a statistical significant positive effect (0.94) reflecting that industrial oil demand is affected by changes in the past level of oil consumption. Finally, the error correction term $(ECT)_{t-1}$ is strongly significant with an adjustment coefficient of −0.77, implying that in the case we are off the long-run demand curve, oil demand adjusts towards its long-run level with about 77% of this adjustment taking place within the first year.

The dynamic oil demand function appears to be well behaved to the diagnostic tests including the serial correlation (LM test), the autoregressive conditional heteroskedasticity test (ARCH test) and the White test for heteroskedasticity. Also the estimated regression is tested whether it is stable throughout the sample, using Chow breakpoint and forecast tests.

The estimation of electricity demand in the short-run gives similar results (column 2). All the coefficients of the variables are in alignment with the theory, smaller than their long-run counterparts and statistical significant except the number of industrial consumers D(CONS) and heating degree-days D(HDD). Analytically, electricity demand seems to be rather inelastic with respect to price with the relevant elasticity (own-price) estimated below unity (−0.35). This evolution, reflects that when electricity prices are raised (i.e. tax imposition, price shock) by a significant proportion (e.g. 10%) provided that all the other determinants affecting electricity demand stay unchanged, electricity demand will fall by about 3.5%, respectively. Electricity demand seems to have a substitutability relationship with oil since cross-price elasticity is equal to 0.18.

Furthermore, electricity demand does not seem to be affected by the alterations in the level of temperature (HDD). This evolution is consistent with the study of Kamerschen and Porter (2004), in which the industrial HDD parameter estimates although negative they were not statistically significant. Relevant results are obtained in the study of Beenstock et al. (1999). Finally, the ECT (or long-run adjustment coefficient) is estimated to −0.24. This means that nearly 24% of long-run disequilibrium is adjusted to the current period.

The above results pass a series of diagnostic tests for serial correlation, heteroskedasticity, ARCH and normality of residuals. Overall, none of these tests, given the power for which they are designed over the sample size could find any significant evidence of departures from standard assumptions.

The next step is the examination of the exogeneity or the endogeneity of the variables. In other words, we can check for the direction of Granger causality between the selected variables of the energy models. In this way, we can check the validity of the hypothesis that the estimated equations are demand relationships. For the equations (oil and electricity demand) to be demand curves, the prices (at minimum) would have to be weakly exogenous. Table 6 reports the findings of the endogeneity of the variables for all two categories of energy consumption (oil and electricity consumption), based on vector error corrections models. More specifically, we provide joint Wald *F*-statistics tests of the lagged explanatory variables in the ECM. These tests give an indication of the significance of short-run causal effects (Asafu-Adjaye, 2000). We also employ *t*-statistics for the coefficients of the ECTs which

Table 6
Summary of tests for weak and strong exogeneity of variables in the VECM's

Equations	ECT (weak exogeneity)	Short-run dynamics (non-causality)							Strong exogeneity
		D(OIL)	D(ELEC)	D(DBP)	D(RPOIL)	D(RPELEC)	D(HDD)	D(CONS)	
D(OIL)	−2.81[a]	—	—	0.68	0.27	0.0010	0.002	—	1.78
D(DBP)	0.03	0.19	—	—	0.008	0.52	0.15	—	0.95
D(RPOIL)	−0.17	0.04	—	0.10	—	1.66	1.32	—	4.13
D(RPELEC)	0.05	1.56	—	0.18	0.23	—	9.42[b]	—	3.56
D(HDD)	−0.09	9.18[b]	—	1.67	11.7[b]	2.70	—	—	1.88
D(ELEC)	−2.34[a]	—	—	0.51	0.79	0.36	0.03	0.04	3.29
D(DBP)	0.13	—	0.39	—	0.11	0.12	0.14	0.0002	1.09
D(RPOIL)	0.08	—	0.01	0.62	—	0.14	0.22	0.001	5.72
D(RPELEC)	−0.20	—	0.02	0.89	0.78	—	4.75[b]	8.70[a]	2.78
D(CONS)	0.05	—	0.84	0.006	0.94	0.07	0.76	—	5.16
D(HDD)	0.09	—	0.01	0.15	2.04	3.04	—	1.69	10.08

Note: Numbers beneath the column "ECT" refer to *t*-statistics testing the H_0 that the error-correction term is equivalent to zero (weak exogeneity). In the "Short-run" column the figures refer to significance levels associated with *F*-statistics testing the H_0 that the lags of the variable indicated are jointly equivalent to zero.
[a]Significance at 0.05.
[b]Significance at 0.01.

give an indication of long-run causal effects, while we do test for the strong exogeneity of the variables entering the VECMs.

From the estimates of the parameters, we see that the ECT is significant both in the equation of industrial oil and electricity consumption. This evolution means that industrial oil and electricity consumption have the tendency to restore equilibrium and take the burden of the shock to the system. In other words given a deviation of industrial oil demand from the long-run equilibrium relationship as defined by ECT, DBP, RPOIL, RPELEC and HDD, all four variables interact in a dynamic fashion to restore long-run equilibrium. The same result holds for the industrial electricity demand. However, the *t*-tests for the ECTs, indicate that all the other variables of the two distinct energy models including the price level of oil and electricity industrial consumption are weakly exogenous. The non-significance of the *F*-statistics for price implies that energy consumption (oil and electricity) bears the burden of the short-term adjustment to long-term equilibrium.

In the short-run dynamics, the Wald tests indicate that a relationship holds in the electricity price level equation. More specifically, the results of the non-causality tests suggest that in the short-run changes in the electricity price level are affected by changes in the temperature level and electricity consumption. In other words, there is a causal relationship between the above-mentioned variables (short-run endogeneity). Finally, the empirical results do not reject the null hypothesis of strong exogeneity of all variables.

From the above empirical analysis we can conclude that in the long-run industrial oil consumption in Greece should not be considered as exogenous to its main determinants.[18] The same does hold for industrial electricity consumption. However, in the short-run, industrial oil and electricity demand are weakly exogenous variables not affected by changes in prices and the level of industrial production. Finally, industrial prices of oil and electricity are weakly exogenous denoting that energy consumption should be treated as a necessity with a perfectly inelastic demand curve.

[18]Similar results are obtained by Hondroyiannis et al. (2002).

5. Conclusions

In this paper we have attempted to investigate the main determinants of the industrial energy demand in Greece. For this purpose we estimated the elasticities of oil and electricity demand in Greece during the period 1970–2004. We then decompose industrial energy demand into its main fuel inputs (oil and electricity) and examine the issue of price sensitivity of both in the short and the long-run. In order to capture short-run dynamics, we utilized the (vector) error correction methodology and estimate the relevant elasticities. The main results of the estimated econometric models can be summarized as follows.

First, maximum likelihood procedure (Johansen methodology) revealed cointegration for both oil and electricity industrial demand, respectively. Second, the empirical findings from maximum likelihood procedure indicate that the demand for each energy type (oil and electricity) is inelastic both in the short and the long-run. Furthermore, the demand for oil exhibits the lowest price responsiveness compared to industrial electricity demand both in the short and the long-run. This is not unusual since almost all of the industry sectors in Greece are oil-intensive. The cross-price elasticities reveal that oil and electricity are substitutes in the industrial process as suggested in most of the empirical studies for Greece. Industrial electricity demand does not seem to be influenced by variations in temperature and changes in the number of consumers.

The results of this paper do not substantially differ from earlier works. Industrial energy demand in Greece seems to be rather inelastic while there is strong evidence that a substitutability relationship between fuels (oil and electricity) holds. Also, industrial energy demand appears to be inelastic on other countries as well. These results could have serious policy implications in terms of tax structure and environmental concerns. For instance, should the Greek government need to increase the level of its tax revenues, more tax could be levied on oil than electricity since we found the price elasticity for oil to be much lower than for electricity. Furthermore, the differences in own price elasticities across energy source provide useful evidence of informing environmentally oriented energy taxation if the government wishes to implement this policy objective.

Moreover using the vector-error correction mechanism, we detect (both in the long and the short-run) the existence of Granger causality and the endogeneity or the exogeneity of each variable. From the empirical results, we saw that in the long-run oil and electricity consumption variations should be considered as endogenous variables to price and industrial production. In the short-run, both oil and electricity consumption are weakly exogenous variables not affected by changes in prices and the level of industrial production. In addition the use of strong exogeneity tests uncover the Granger exogeneity of industrial oil and electricity prices, respectively, denoting that industrial energy consumption has an inelastic demand curve as suggested by the economic theory.

The study's finding of weakly exogeneity in industrial energy prices has a number of potential implications for energy policy analysis and forecasting. The most significant one is that industrial energy consumption in Greece should be treated as a necessity and thus price variations tend to have a minimum effect in the level of energy demand both in the short and the long-run. This finding, also calls for caution in the use of single equation regressions of industrial energy demand for conducting econometric forecasts. Our results suggest that both the industrial energy prices and the level of industrial production in Greece could be treated as exogenous variables and therefore single equation forecasts of one or the other could not be misleading. In particular, any analysis which does not employ the error-corection terms might likely give unreliable results.

Lastly we must mention that none of the two dummy variables that capture the rapid decline of the international price of oil and the introduction of natural gas in industry had a significant impact on industrial energy demand. This evolution, combined with the results of the structural stability tests imply that there was not found signs of structural change in industrial demand. That is because Greek manufacturing is rather sluggish and is strongly characterized by low efficiency as a result of the low degree of substitutability between the sources of energy. In other words, Greece is a country with a low degree of industrialization and low energy consumption that could not curtail the increasing needs for energy of its economy.

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