Substitution possibilities for energy in the Italian economy: a general to specific econometric analysis

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Abstract

The paper considers a neoclassical model set in the cost function approach to estimate primary energy factor demands for the Italian economy, using a translog cost function specification. Cointegration theory is employed to estimate the long-run factor share model, and the general to specific methodology to derive an error correction formulation for the short-run adjustment process. Both quarterly and yearly series, for the period 1978q1-1994q4 and 1960-1994, respectively, have been considered in the analysis. The different energy sources substitution pattern obtained by the quarterly and annual series and the super exogeneity property of the annual model suggest the importance of using low frequency data rather than quarterly data in estimating long-run relationships.

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1 Introduction

A neoclassical model has been used to focus on the long-run role of prices in explaining changes in primary energy demand in Italy. This model is set in the cost function approach, employed to retrieve the energy input derived demand functions. The aggregate production function, specified in terms of aggregate capital, labour, energy and raw materials, has been studied in its dual representation, under the assumption of weak separability of the energy inputs from the other arguments of the production function. A translog cost function has been employed to derive, via Shephard's lemma, the specification of the energy share equations. Four primary energy sources have been considered, namely, oil, electricity, natural gas and coal.

Recent advances in time series econometric theory have devised tools to model jointly the long-run equilibrium relationships and their associated short-run dynamics. The Engle and Granger (1987) and the Engle and Yoo (1989) cointegration approach has been utilised to estimate the long-run relationships while the general to specific methodology has been followed to derive error correction formulations, in which short-run dynamics and equilibrium relationships could be modelled together. Both yearly (1960-1994) and quarterly (1978q1-1994q4) series have been considered in this study. High frequency data are generally available for a shorter time horizon than low frequency data. Time series econometric studies usually tend to favour the number of observations over the time span of the data set. This is mainly due to the fact that most of the results provided by time series econometrics are justified only asymptotically. Both quarterly and yearly models have, then, been fitted to asses whether similar results were yielded by data sampled at different frequencies and covering only partially overlapping time spans.

2 The underlying economic model

Suppose the aggregate production function is explicited with respect to capital (K), labour (L), energy (E), raw materials (M) and the level of technology (T):

$$Y = f\left(K, L, E, M, T\right) \qquad \textbf{(1)}$$

The production function is assumed to be weakly separable in the four main categories of primary energy sources, namely oil (E_o), electricity (E_e)[1], gas (E_g) and coal (E_c), so that the technical rates of substitution between any two of the fuel inputs is independent of the quantities of the other non-energy inputs. The level of technology can be thought of as composed by an index A of neutral technical progress and a number of indexes A_i representing factor saving technical progress.

Under the assumption of homothetic weak separability of the production function in the energy inputs and of exogeneity of the factor prices and the output level (Shephard, 1953) the corresponding cost function in efficiency units is

$$C = C\left(\frac{P_E}{A_E}\left(\frac{P_o}{A_o}, \frac{P_e}{A_e}, \frac{P_g}{A_g}, \frac{P_c}{A_c}\right), \frac{P_K}{A_K}, \frac{P_L}{A_L}, \frac{P_M}{A_M}, A, Y\right) \qquad \textbf{(2)}$$

and the optimisation problem may be solved in two stages (Denny and Fuss, 1975). In the first stage the economic agents optimise with respect to the fuel mix, while in the second stage the optimisation is concerned with capital, labour, materials and energy. Since the econometric model studied is concerned with the first stage alone, the hypothesis of homotheticity of the energy subsector has not been assumed a priori but has been explicitly tested. That is, total energy production enters as an explanatory variable in the energy model and its statistical significance is tested.

When the aggregate energy price index $\frac{P_E}{A_E}$ is approximated by a translog cost function it may be written as

$$\ln C_t = a_0 + \sum_i a_i \ln \frac{p_{it}}{A_{it}} + \frac{1}{2} \sum_i \sum_j a_{ij} \ln \frac{p_{it}}{A_{it}} \ln \frac{p_{jt}}{A_{jt}} + a_E \ln E_t +$$

$$+ \frac{1}{2} a_{EE} (\ln E_t)^2 + \sum_i a_{Ei} \ln \frac{p_{it}}{A_{it}} \ln E_t \tag{3}$$

where *i,j=o,e,g,c*.

Partial differentiation with respect to factor prices (Shephard's Lemma) gives a set of cost share equations and the generic cost share may be written as

$$S_{it} = a_i + \sum_j a_{ij} \ln \left(\frac{p_{jt}}{A_{jt}} \right) + a_E \ln E_t \tag{4}$$

where the index *i* refers to the four energy shares, the index *j* to the energy prices, and the index *t* is the temporal index.

3 The econometric model

When an exponential augmentation form of the type $A_{jt} = A_0 e^{bt}$ is selected, the cost share equation system may be written as

$$S_{ot} = a_o + b_o t + \sum_j a_{oj} \ln p_{jt} + a_{oE} \ln E_t + u_{ot}$$

$$S_{et} = a_e + b_e t + \sum_j a_{ej} \ln p_{jt} + a_{eE} \ln E_t + u_{et}$$

$$S_{gt} = a_g + b_g t + \sum_j a_{gj} \ln p_{jt} + a_{gE} \ln E_t + u_{gt}$$

$$S_{ct} = a_c + b_c t + \sum_j a_{cj} \ln p_{jt} + a_{cE} \ln E_t + u_{ct}$$

(5)

[1]: The electricity primary input is an aggregate input composed largely of hydroelectric, geothermal and nuclear energy. Nuclear energy is no longer produced internally, but imported from France in the form of electricity.

where $\alpha_i = -\sum_j \alpha_{ij} A_0$, $\beta_i = -\sum_j \alpha_{ij} \beta$, $i, j = o, e, g, c$ and u_{it} is an additive disturbance term. The restrictions of price symmetry and homogeneity require $\alpha_{ij} = \alpha_{ji}\ \forall i, j$ $i \neq j$ and $\sum_j \alpha_{ij} = 0\ \forall i,\ i, j = o, e, g, c$, respectively.

Since coal demand is likely to have been more affected by regulatory measures than the other energy sources over the time span considered, the share of coal has been omitted and the remaining three cost shares, namely oil, electricity and natural gas have been estimated by FIML. This allows to overcome the problem of the singular disturbance variance-covariance matrix associated to the full share equations system (adding-up restriction). Moreover, the use of ML guarantees that parameter estimates, estimated standard errors, and log-likelihood values are invariant to the choice of which equation is deleted (Berndt, 1991).

3.1 The statistical framework

Economic theory supports the reduction of the eight equation VAR to a system of three share equations. This is on the basis of the assumed exogeneity of prices and total energy for the long-run parameters. In addition, all of the series seemed to behave as I(1) processes.[2] Therefore, the Engle and Granger (1987) and Engle and Yoo (1989) methods have been followed to obtain the estimates of the long and short-run parameters for each share equation, assuming the weak exogeneity of prices and total energy. This assumption has been successively tested following the procedure indicated by Johansen (1992a) and Urbain (1992).[3]

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[2]: The statistical properties of the series have been assessed by ADF (Dickey and Fuller, 1981) and HEGY (Hylleberg, Engle, Granger and Yoo, 1990) tests. The results suggested the plausibility of modelling the cost share equations system as non-stationary, treating all of the series as I(1) process. As far as seasonality is concerned, the HEGY tests suggested a deterministic specification.

[3]: When the parameters of interest are the cointegrating vectors and the error correction coefficients, a necessary and sufficient condition for a conditioning variable to be weakly exogenous is to be not-error correcting. Then, weak exogeneity has been tested via *t*-ratio and *F* tests conducted on the marginal models, to check that the coefficients of the added error correction terms are zero. The null of weak exogeneity has been found to be non rejected in all of the cases.

The Engle and Granger (1987) method consists of two successive steps. In the first stage a static regression (explaining the long-run) is run and the residuals are tested for stationarity. In the second stage, the stationary residuals are plugged in to the error correction system (explaining the short-run variation) and an estimate of the speed of adjustment of the system is obtained. The Engle and Granger method gives consistent estimates of the long-run parameters. The Engle and Yoo (1989) third-step makes the consistent estimates of the cointegrating vector asymptotically efficient, and makes the distribution of the estimator of the cointegrating vector standard.

3.2 Testing for cointegration

The specification of the static unrestricted systems is the following

$$\mathbf{S}_t = \boldsymbol{\beta}\mathbf{x}_t + \boldsymbol{\Psi}\mathbf{D}_t + \mathbf{u}_t \qquad t = 1,\ldots,T \qquad \textbf{(7)}$$

where $\boldsymbol{\beta}$ and $\boldsymbol{\Psi}$ are matrices of coefficients of dimension (3×5) and $(3\times d)$ with $d=2$ for the yearly model and $d=5$ for the quarterly model. $\mathbf{x}'_t = \left[\ln p_{ot} \quad \ln p_{et} \quad \ln p_{gt} \quad \ln p_{ct} \quad \ln E_t\right]$, $\mathbf{S}'_t = \left[S_{ot} \quad S_{et} \quad S_{gt}\right]$, $\mathbf{u}'_t = \left[u_{ot} \quad u_{et} \quad u_{gt}\right]$, and $\mathbf{D}_t$ includes an intercept, a deterministic linear trend and three centred seasonals in the quarterly model $\left(\mathbf{D}'_t = \left[1 \quad t \quad cs_1 \quad cs_2 \quad cs_3\right]\right)$, and includes the intercept and the trend alone in the yearly model $\left(\mathbf{D}'_t = \left[1 \quad t\right]\right)$. The error terms are assumed to be independently normally distributed $\mathbf{u} \approx \mathbf{IN}(\mathbf{0},\boldsymbol{\Omega})$. Model (7) has been estimated by FIML.

The interpretation of the static regressions as cointegrating regressions is supported by the Dickey-Fuller (ADF) test for cointegration (Engle and Granger, 1987) and by the cointegrating regression Durbin-Watson test (CRDW; Sargan and Bhargava, 1983).

Table 1: Cointegration tests.[4]

	Quarterly data		Yearly data	
shares	**ADF**	**CRDW**	**ADF**	**CRDW**
So	-5.34**	1.19*	-5.00*	2.06*
Se	-5.06**	1.15*	-4.72*	1.98*
Sg	-8.27***	1.65*	-3.55	1.89*

As is shown in table 1, the null of non-cointegration is always rejected for both the quarterly and yearly models. For the yearly share of gas series the null of non-cointegration can be rejected on the basis of the CRDW test. However, as Kremers *et al.* (1992) have shown, in testing for cointegration with the Dickey-Fuller test a possibly invalid common factor restriction is imposed a priori on the data. Therefore, the presence of the error correction term in the conditional process should be considered as the strongest evidence in favour of the cointegration hypothesis.

The static cointegrating regression may be justified asymptotically because the lagged differenced terms are of a lower order of magnitude than the variables in levels, that is they are $\mathrm{I}(0)$ while the non-stationary regressors are $\mathrm{I}(1)$. However, neglecting the dynamic adjustment process has been found to matter even for large but finite samples (Hendry and Neale, 1987), so that the Engle and Yoo (1989) third step correction has been carried out successively to make the estimates of the long-run parameters fully efficient and to validate statistical inference.

3.3 The short-run adjustment mechanism

A general to specific methodology has been followed to achieve parsimonious short-run dynamic specifications for the econometric models. As Anderson and Blundell (1982) have shown, given the long-run share relationships in system form

$$\mathbf{s}_t = \boldsymbol{\beta}\mathbf{x}_t + \mathbf{u}_t \qquad t = 1, \ldots, T \qquad \textbf{(8)}$$

[4]: For the ADF tests, "*" indicates rejection of the null of no-cointegration at the 10% significance level, "**" at the 5% significance level, and "***" at the 1% significance level. The critical values are from Engle and Yoo (1989), Table II. For the CRDW tests , "*" indicates rejection of the null of no-cointegration at the 5% significance level. The critical values are from Banerjee *et al.* (1993), Table 7.1.

where $\boldsymbol{\beta}$ is a matrix of coefficients of dimension (3×7) for both the yearly and quarterly models[5], $\mathbf{S}'_t = [S_{ot} \quad S_{et} \quad S_{gt} \; S_{ct}]$, $\mathbf{u}'_t = [u_{ot} \quad u_{et} \quad u_{gt} \; u_{ct}]$ with $\mathbf{u} \approx \mathbf{N}(\mathbf{0}, \boldsymbol{\Sigma})$ and $\mathbf{x}'_t = [1 \quad t \quad \ln p_{ot} \quad \ln p_{et} \quad \ln p_{gt} \quad \ln p_{ct} \quad \ln E_t]$, a general dynamic version of the model can be written as

$$\mathbf{B}^*(L)\mathbf{S}_t = \boldsymbol{\Gamma}^*(L)\mathbf{x}_t + \mathbf{u}_t \qquad \textbf{(9)}$$

where $\mathbf{B}^*(L)$ and $\boldsymbol{\Gamma}^*(L)$ are polynomial matrices in the lag operator L,

$$\mathbf{B}^*(L) = \mathbf{I} + \mathbf{B}^*_1 L + \mathbf{B}^*_2 L^2 + \ldots + \mathbf{B}^*_p L^p$$

$$\boldsymbol{\Gamma}^*(L) = \boldsymbol{\Gamma}^*_0 + \boldsymbol{\Gamma}^*_1 L + \boldsymbol{\Gamma}^*_2 L^2 + \ldots + \boldsymbol{\Gamma}^*_q L^q \, .$$

and reparameterized in the observationally equivalent generalised error correction form

$$\Delta\mathbf{S}_t = -\mathbf{B}(L)\Delta\mathbf{S}_t + \Gamma(L)\Delta\tilde{\mathbf{x}}_t - \mathbf{K}\left[\mathbf{S}_{t-p} - \boldsymbol{\beta}\mathbf{x}_{t-q}\right] + \mathbf{u}_t \, , \qquad \textbf{(10)}$$

where $\mathbf{B}(L) = \sum_{i=1}^{p-1}\left(\sum_{j=0}^{i}\mathbf{B}^*_j\right)L^i \quad p > 1$, $\boldsymbol{\Gamma}(L) = \sum_{i=0}^{q-1}\left(\sum_{j=0}^{i}\tilde{\boldsymbol{\Gamma}}^*_j\right)L^i \quad q \geq 1$,

$\mathbf{K} = \sum_{j=0}^{p}\mathbf{B}^*_j = I + B^*_1 + B^*_2 + \ldots + B^*_p$, $\tilde{\boldsymbol{\Gamma}}^*_j$ is $\boldsymbol{\Gamma}^*_j$ with the first column deleted and $\tilde{\mathbf{x}}$ is $\mathbf{x}_t$ with the first element (the intercept) deleted.

Because of the singularity problem, the error correction model as it is cannot be estimated. As in the static case an estimable model may be obtained by deleting one of

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[5]: The centred seasonals have been neglected for simplicity.

the equation (the share of coal dynamic equation) from the model. The dynamic model becomes

$$\Delta\mathbf{S}_t = -\mathbf{B}^n(L)\Delta\mathbf{S}_{nt} + \Gamma(L)\Delta\tilde{\mathbf{x}}_t - \mathbf{K}^n\left[\mathbf{S}_{nt-p} - \boldsymbol{\beta}_n\mathbf{x}_{t-q}\right] + \mathbf{u}_t\,, \qquad \textbf{(11)}$$

where $\mathbf{B}^n(L)$ and $\mathbf{K}^n$ are $\mathbf{B}(L)$ and $\mathbf{K}$ with the *n*th column deleted, respectively, and $\mathbf{S}_{nt}$ and $\boldsymbol{\beta}_n$ are $\mathbf{S}$ and $\boldsymbol{\beta}$ with the *n*th row deleted. The adding-up restriction then implies that each column of the coefficient matrices $\mathbf{B}^{jn}$ $j = 1, ..., p$ and $\mathbf{K}^n$ adds up to zero, that is

$$\mathbf{i}'\mathbf{B}_j^n = \mathbf{0} \quad j = 1, ..., p \qquad \textbf{(12)}$$

$$\mathbf{i}'\mathbf{K}^n = \mathbf{0} \qquad \textbf{(13)}$$

In particular, the restriction $\mathbf{i}'\mathbf{K}^n = \mathbf{0}$ implies that in the case of a diagonal system, each share has to adjust at the same speed.

Since the adding-up restrictions strongly constrain the selection of dynamics, I have preferred to let the data determine the adjustment process, by constraining the feedback matrix only. The adding-up restrictions as the other economic restrictions (price homogeneity and symmetry), however, have been imposed on the long-run structure of the model. This would guarantee that the estimates of the long-run parameters are invariant with respect to which share equation is deleted.[6]

The error correction models have been derived by simplification, assisted by testing for whiteness, normality, and structural stability. The final specification has been obtained in two successive steps. In the first step the starting model, after testing, has been reduced to the most parsimonious form. A first interesting finding is that the data did not reject the hypothesis of a diagonal feedback matrix for both the quarterly and yearly models, so that each equilibrium relationship enters only into the corresponding equation of the short-run model. The diagonalisation of the adjustment process implies that the disequilibria in the energy market are not interrelated, so that the adjustment in each share equation depends only on the gap between the actual and long-run values of each corresponding share. The likelihood ratio tests for the diagonalisation of the feedback matrix are $\chi^2(6) = 8.9040\ [0.1790]$ and $\chi^2(6) = 10.4289\ [0.1077]$ for the quarterly and yearly models, respectively.

In the second step, the economic restrictions have been tested and imposed on the long-run components. To do this, the quarterly and yearly error correction terms have been entered in unrestricted form in the corresponding dynamic econometric models and the restriction of price symmetry has been tested via a LR test, after having imposed the price homogeneity constraint. The restrictions were not rejected by the data for both the quarterly and yearly data ($\chi^2(3) = 3.88\ [0.2794]$ for the quarterly data and $\chi^2(3) = 1.7820\ [0.6189]$ for the yearly series).

Finally, the restricted error correction terms have also been constrained to assume the same speed of adjustment ($\alpha_{11} = \alpha_{22} = \alpha_{33}$), restriction which the data have not rejected ($\chi^2(3) = 3.9009\ [0.1422]$ and $\chi^2(3) = 5.0599\ [0.0797]$ for the quarterly and yearly models, respectively).

The estimated residuals have been tested for autocorrelation (Godfrey, 1978; Breusch, 1978), heteroscedasticity (White, 1980), normality (Doornik and Hansen, 1994) and for autoregressive conditional heteroscedasticity (Engle, 1982). The econometric models have been also tested for parameter stability.[7]

6: The dynamic model was estimated also with the adding-up restrictions fully imposed. The estimated speed of adjustment parameters did not show any significant difference with respect to the estimates reported.

7: Parameter constancy has been analysed by rolling system Chow (1960) tests conducted over 29 periods in the quarterly model (1986q1-1994q4) and over 12 periods in the yearly model (1983-1994). Following Doornik and Hendry (1994), three different kinds of system Chow tests have been calculated, namely 1-step

Table 4: Yearly econometric model (FIML estimates).

REM	###So		###Se		###Sg	
variable	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
###S*g*_2	-0.4230	.1438			0.6307	.0986
###ln*E*	0.1769	.0317	-0.1480	.0306		
###ln*Po*	0.2239	.0118	-0.1639	.0136	-0.0354	.0050
### ln*Po*_2	0.0382	.0100	-0.0627	.0134	0.0385	.0059
###ln*Pe*	-0.2287	.0340	0.2517	.0351	-0.0219	.0088
### ln*Pe*_1	-0.1200	.0328	0.1084	.0312		
### ln*Pe*_2					0.0472	.0079
###ln*Pg*			-0.0624	.0102	0.0849	.0063
### ln*Pg*_2			0.0612	.0144	-0.0809	.0104
###ln*Pc*	0.0466	.0219	-0.0880	.0210		
*ECT*_1	-0.5511	.0809	-0.5511	.0809	-0.5511	.0809

Table 5: Quarterly econometric model (FIML estimates).

REM	###So		###Se		###Sg	
variable	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
###S*o*_1			-0.2205	.0791		
###S*e*_1			-0.2601	.0882		
###S*e*_2	0.2804	.0618			-0.2786	.0542
###S*g*_1			-0.3450	.0842		
###S*g*_2	0.7872	.1001	-0.3188	.0827	-0.5228	.0709
###ln*E*	0.1333	.0277	-0.3130	.0255	0.1795	.0188
###ln*Po*	0.2774	.0281	-0.1113	.0264	-0.1247	.0181
### ln*Po*_2	0.0931	.0247			-0.0913	.0218
### ln*Pe*_1	-0.1445	.0342	0.1736	.0337		
###ln*Pg*	-0.0698	.0259	-0.0891	.0243	0.1696	.0168
### ln*Pg*_2	-0.0932	.0234			0.0733	.0206
###ln*Pc*	-0.0334	.0131			-0.0263	.0115
*ECT*_1	-0.4121	.0709	-0.4121	.0709	-0.4121	.0709

Chow tests, break point Chow tests and forecasts Chow tests. Both models are stable at the 5% level. See figures A1 and A2 in the appendix.

Table 6: Yearly econometric model, properties of the residuals.[8]

REM	###So	###Se	###Sg	Vector ###S
s	0.0104	0.0114	0.0037	
AR 1-4	2.9151 [0.06]	4.1536 [0.02]	1.4813 [0.26]	1.0608 [0.44]
ARCH 1	0.5041 [0.49]	0.5984 [0.45]	2.6826 [0.12]	
Normality	1.76 [0.42]	0.8219 [0.66]	0.4210 [0.81]	6.46 [0.37]

Table 7: Quarterly econometric model, properties of the residuals.[9]

REM	###So	###Se	###Sg	Vector ###S
s	0.0184	0.0177	0.0116	
AR 1-8	1.3275 [0.26]	1.3909 [0.23]	1.4589 [0.20]	1.1213 [0.30]
ARCH 4	2.2923 [0.08]	2.1239 [0.09]	0.6298 [0.64]	
Normality	0.2633 [0.88]	0.4849 [0.78]	2.0976 [0.35]	5.7402 [0.45]
Heterosc.	0.3521 [0.99]	0.5590 [0.93]	0.1900 [1.00]	0.5573 [0.99]

An interesting feature of the short-run adjustment process for the cost share of oil is that no oil share lagged values have been found statistically significant at the 5% level of significance. Only the lagged electricity and natural gas shares (###S*e*_2, ### S*g*_2) enter the quarterly specification and just the lagged share of natural gas the yearly one. The adjustment process seems to be mainly explained by contemporaneous and lagged values of the exogenous regressors. The main difference between the quarterly and yearly specifications is given by the absence of the natural gas price variable in the yearly model and by the weaker role played by the electricity price variables in the quarterly model.

As far as the shares of electricity are concerned, the results indicate that the adjustment process for the quarterly and yearly shares are quite different. In fact, in the quarterly model both lagged shares and contemporaneous exogenous variables have a role in explaining the share of electricity dynamics while in the yearly model lagged share values do not have any role and the adjustment process is entirely explained by contemporaneous and lagged values of the exogenous regressors.

8: AR 1-4 = F(4,15); Normality = c^2(2); ARCH 1 = F(1,17); Vector AR 1-4: F(36,30); Vector Normality = c^2(6). The heteroscedasticity test has not been carried out on the yearly statistical model due to lack of degrees of freedom.

As for the share of oil, the adjustment process for the quarterly and yearly shares of gas is explained by both lagged shares (###S*e*_2 and ###S*g*_2) and contemporaneous and lagged exogenous variables. In the yearly model, however, only the lagged share of natural gas has been found statistically significant at the 5% significance level.

To summarise the results, the selected models meet the diagnostic criteria satisfactorily and the residuals approximate Gaussian white noise processes. The main difference would seem to be the lower speed of adjustment shown by the quarterly econometric model. In all of the cases, however, the error correction term has been found negative and strongly significant, a result which gives support to the cointegration analysis, since as Boswijk and Franses (1992) have indicated, in the context of valid conditional error correction modelling, the statistical significance of the error correction term dominates, as a cointegration test, both the Johansen maximal eigenvalue test and the ADF test. Moreover, more parsimonious specifications have been selected for the yearly models than the quarterly ones. Finally, the yearly models, on average, show lower standard error of the regression than the quarterly ones.

3.4 The long-run

The Engle and Yoo (1989) third step correction has been applied to the estimates of the cointegrating vector provided in the first step. The three share equations have been estimated jointly with the price homogeneity and simmetry restrictions imposed.

9: AR 1-8 = F(8,42); Normality = $c^2(2)$; ARCH 4 = F(4,42); Heteroscedasticity = F(30,19); Vector AR 1-8 =

Table 8: Quarterly corrected restricted economic model.

cREM	**So**		**Se**		**Sg**		**Sc**	
variable	**Coeff.**	**S.E.**	**Coeff.**	**S.E.**	**Coeff.**	**S.E.**	**Coeff.**	**S.E.**
ln*E*	0.2999	.1041	-0.3682	.1053	0.0583	.0729	0.0100	.0509
ln*Po/Pc*	0.2517	.0480	-0.0982	.0352	-0.0935	.0389	-0.0600	.0443
ln*Pe/Pc*	-0.0982	.0352	0.1896	.0442	-0.0495	.0292	-0.0419	.0436
ln*Pg/Pc*	-0.0935	.0389	-0.0495	.0292	0.129	.0389	0.0140	.0282
Trend	-0.0027	.0004	0.0012	.0004	0.0016	.0003	-0.0001	.0002
CSeas	-0.0444	.0160	0.0029	.0156	0.0346	.0107		
CSeas_1	0.0253	.0251	-0.0087	.0247	-0.0265	.0164		
CSeas_2	0.0764	.0273	-0.0477	.0268	-0.0399	.0183		
Const	-0.3245	.3530	1.2121	.3461	-0.0078	.2381	0.1202	.0551

Table 9: Yearly corrected restricted economic model.

cREM	**So**		**Se**		**Sg**		**Sc**	
variable	**Coeff.**	**S.E.**	**Coeff.**	**S.E.**	**Coeff.**	**S.E.**	**Coeff.**	**S.E.**
ln*E*	0.3973	.0218	-0.287	.0254	-0.0502	.0088	-0.0601	.0088
ln*Po/Pc*	0.2141	.0128	-0.1314	.0116	-0.04	.0080	-0.0427	.0212
ln*Pe/Pc*	-0.1314	.0116	0.2047	.0143	-0.033	.0059	-0.0403	.0331
ln*Pg/Pc*	-0.04	.0080	-0.033	.0059	0.0702	.0075	0.0028	.0325
Trend	-0.0110	.0006	0.0044	.0006	0.0055	.0002	0.0011	.0003
Const	-0.9433	.1082	1.2166	.1279	0.3158	.0461	0.4109	.0541

The estimates for the parameters and standard errors of the omitted share of coal have been retrieved using the homogeneity constraints. By comparing the estimates of tables 8 and 9 it can be noticed that the quarterly and yearly corrected parameter estimates, apart from the estimated parameter for total energy in the share of natural gas equation, show the same signs. The major difference in magnitude between the quarterly and yearly parameter estimates is given by the estimated parameters of the total energy variable and by the biases of technical progress. In particular the biases of technical progress are much larger in the yearly model than in the quarterly one. This may be imputed to the longer time span covered by the yearly data, which includes periods of intense adjustments in the Italian energy market following the two major oil price shocks of 1973 and 1979-1980. According to the estimates, over the period

F(72,90); Vector Normality = $c^2(6)$; Vector Heteroscedasticity = F(180,126).

1960-1994, technical progress would have biased the system towards saving oil and using electricity, natural gas and coal. Moreover, by comparing the quarterly and yearly implicit estimates of the share of coal parameters it can be noticed that the large standard errors make some of the shares of coal parameters statistically non significant.

4 Price and substitution elasticities

The Allen partial elasticities of substitution and the partial price elasticities have been calculated for the corrected yearly and quarterly models at the average share point of approximation. Substitution (σ) and price elasticities (ε) have been calculated by the formulas

$$\sigma_{ij} = \frac{\left(\beta_{ij} + S_i S_j\right)}{S_i S_j} \quad i \neq j \qquad\qquad \sigma_{ii} = \frac{\left(\beta_{ii} + S_i^2 - S_i\right)}{S_i^2} \tag{14}$$

$$\varepsilon_{ij} = S_j \sigma_{ij} = \frac{\left(\beta_{ij} + S_i S_j\right)}{S_i} \quad i \neq j \qquad\qquad \varepsilon_{ii} = S_i \sigma_{ii} = \frac{\left(\beta_{ii} + S_i^2 - S_i\right)}{S_i} \tag{15}$$

where $i, j = o, e, g, c$.

Table 9: Allen elasticities of substitution and partial price elasticities, yearly estimates.

average share	Elasticities of substitution				Partial price elasticities			
	o	**e**	**g**	**c**	**o**	**e**	**g**	**c**
o	-0.14	0.21	0.25	-0.39	-0.07	0.02	0.06	-0.01
	(.05)	(.09)	(.14)	(0.69)	(0.02)	(0.03)	(0.01)	(0.04)
e		-0.13	0.02	-1.09	0.11	-0.04	0.00	-0.07
		(0.14)	(0.22)	(1.73)	(0.04)	(0.04)	(0.02)	(0.10)
g			-2.14	1.45	0.13	0.01	-0.22	0.09
			(0.69)	(5.16)	(0.07)	(0.06)	(0.07)	(0.31)
c				6.71	-0.20	-0.35	0.15	0.40
				(2.53)	(0.35)	(0.55)	(0.54)	(0.15)

Table 10: Allen elasticities of substitution and partial price elasticities, quarterly estimates.

average share	Elasticities of substitution				Partial price elasticities			
	o	e	g	c	o	e	g	c
o	0.11 (0.16)	0.25 (0.34)	-0.10 (0.46)	-0.85 (1.37)	0.01 (0.09)	0.06 (0.08)	-0.02 (0.07)	-0.05 (0.08)
e		0.16 (0.79)	-0.36 (1.07)	-2.01 (3.13)	0.14 (0.15)	0.04 (0.19)	-0.05 (0.15)	-0.12 (0.18)
g			-0.06 (1.64)	2.55 (3.11)	-0.06 (0.26)	-0.08 (0.19)	-0.01 (0.25)	0.15 (0.18)
c				9.42 (5.32)	-0.47 (0.75)	-0.48 (0.74)	0.39 (0.48)	0.55 (0.31)

A different pattern of substitution among the energy inputs is indicated by the quarterly and yearly estimates. According to the quarterly elasticities, oil and electricity are substitutes while oil and electricity are complement for gas and coal. In addition, gas and coal are substitutes. The complementarity of oil and electricity with coal is particularly strong, as it is the substitutability of gas and coal. On the other hand, the yearly estimates indicate that oil, electricity and natural gas would be substitutes while coal would be a complement for oil and electricity, and a substitute for natural gas. In addition, the quarterly estimated own price elasticities reveal a pervasive violation of the maintained hypothesis of concavity and monotonicity at the point of approximation selected. In fact, the own price elasticities for oil, electricity and coal are positive, although they are not statistically significant at the usual levels. As far as the yearly estimates are concerned, the own price elasticities are negative for oil, electricity, and natural gas.

5 Testing for super exogeneity

By applying the third-step correction on the cointegrating vectors estimated in the first-step, inference on the estimated parameters is allowed for. More economically meaningful magnitudes than the cointegrating vectors are, however, the price and substitution elasticities which have been derived from them. Since the most relevant use of these parameters is for policy analysis, to establish whether the estimates provided are subject to the Lucas (1976) critique is of some importance. Super exogeneity is the concept introduced by Engle, Hendry and Richard (1983) to neutralise the Lucas' 1976 critique. If the conditioning variables are held as superexogenous, then changes in the parameters of the marginal model do not affect the conditional model and valid policy analysis is allowed for.

Following Engle and Hendry (1993), a test for the super exogeneity null hypothesis may be set up by specifying a comprehensive alternative hypothesis allowing for failures of both weak exogeneity and invariance. Considering the bivariate process (y_t, x_t) with joint normal distribution conditioned on the information set $I_t = (Y_{t-1}, X_{t-1}, W_t)$ composed of lagged values of y and x and of current and lagged values of a set of valid conditioning variables w_t, that is

$$\begin{pmatrix} y_t \\ x_t \end{pmatrix} \Big| I_t \approx \mathbf{N}\left[\begin{pmatrix} \mu_t^y \\ \mu_t^x \end{pmatrix}, \begin{pmatrix} \sigma^{yy} & \sigma^{yx} \\ \sigma^{yx} & \sigma^{xx} \end{pmatrix}\right] = \mathbf{N}(\boldsymbol{\mu}_t, \boldsymbol{\Sigma}_t), \quad \textbf{(16)}$$

the conditional model for y_t may be written as

$$y_t \big| x_t, I_t \approx N\left[\phi_t\left(x_t - \mu_t^x\right) + \mu_t^y, \omega_t\right] \quad \textbf{(17)}$$

where $\phi = \dfrac{\sigma_t^{yx}}{\sigma_t^{xx}}$ and $\omega_t = \sigma_t^{yy} - \left(\sigma_t^{yx}\right)^2 / \sigma_t^{xx}$. Suppose the parameters of interest are β and ψ in the theoretical model

$$\mu_t^y = \beta_t(\delta_t)\mu_t^x + \mathbf{w}_t'\boldsymbol{\varphi} \quad \textbf{(18)}$$

in which by the term $b_t(d_t)$ the parameter b is allowed to be influenced by changes in the parameters of the marginal process of x_t (d_t) and the relationship between b and d_t itself is allowed to be time-varying. By substituting (18) into (17) yields

$$y_t \big| x_t, I_t \approx N\left[b_t(d_t)x_t + \mathbf{w}'\boldsymbol{\varphi} + \{f_t - b_t(d_t)\}(x_t - m_t^x), w_t\right] \qquad \textbf{(19)}$$

As Engle and Hendry (1993) have shown, three conditions are necessary for valid inference on the parameters of interest $(b, \boldsymbol{\varphi})$ in the conditional regression model

$$y_t = bx_t + \mathbf{w}_t'\boldsymbol{\varphi} + e_t \qquad \textbf{(20)}$$

namely,

i) x_t is weakly exogenous for the parameters of interest $(b, \boldsymbol{\varphi})$. In this case m_t^x and s_t^x do not enter the conditional model. A necessary condition is then $f_t = b_t(d_t)$;

ii) the regression coefficient of x_t, b_t, is constant. This implies $f_t = f \quad \forall t$ and $w_t = w$ if $s_t^{yy} = w + fs_t^{xx}$;

iii) b_t is invariant to changes in d_t, that is $b_t(d_t) = b \quad \forall t$.

Conditions *i*), *ii*) and *iii*) imply, in fact, that $f = b$, so that the conditional model (19) may be written as

$$y_t \big| x_t, I_t \approx N\left[bx_t + \mathbf{w}'\boldsymbol{\varphi}, w\right] \qquad \textbf{(21)}$$

An alternative hypothesis against which to test the null hypothesis of super exogeneity may be constructed by expanding $\beta(\delta_t)$ in terms of μ_t^x and σ_t^x. Following Psaradakis and Sola (1996), this can be written as

$$\beta\left(\mu_t^x, \sigma_t^{xx}\right) = \beta_0 + \beta_1 \mu_t^x + \beta_2 \sigma_t^{xx} + \beta_3 \sigma_t^{xx}\left(\mu_t^x\right)^{-1} \qquad \textbf{(22)}$$

with $\mu_t^x \neq 0 \ \forall t$.

By substituting (22) into (19) and allowing ϕ_t to be time-varying according to the form $\phi_t = \phi_0 + \phi_1 \sigma_t^{xx}$, the following regression model is obtained

$$y_t = \beta_0 x_t + \mathbf{w}_t' \boldsymbol{\varphi} + \left(\phi_0 - \beta_0\right)\left(x_t - \mu_t^x\right) + \phi_1 \sigma_t^{xx}\left(x_t - \mu_t^x\right) + \beta_1\left(\mu_t^x\right)^2 + \beta_2 \sigma_t^{xx} + \beta_3 \mu_t^x \sigma_t^{xx} + \varepsilon_t . \qquad \textbf{(23)}$$

A direct test for super exogeneity may be carried out from (23) by testing sequentially for weak exogeneity and invariance. Weak exogeneity of x_t for the parameters of interest $(\beta, \boldsymbol{\varphi})$ requires $(\phi_0 - \beta_0) = 0$ while invariance requires $\beta_1 = \beta_2 = \beta_3 = 0$. As noticed by Engle and Hendry (1993), although weak exogeneity and invariance are theoretically necessary and sufficient conditions for super exogeneity, in practice also parameter constancy is required $(\phi_1 = 0)$. Since the terms μ_t^x and σ_t^x are unknown, proxies for the first two conditional moments of each of the conditioning variables have been obtained by the fitted values of the marginal processes. The fitted values $\Delta \ln \hat{Z}_{i,t}$ ($i = P_o, P_e, P_g, P_c, E;\ t = 1, \ldots, n$) have been employed to proxy the conditional means $\hat{\mu}_t^{Z_i}$, while the conditional variances $\hat{\sigma}_t^{Z_i Z_i}$ have been proxied by the terms $\hat{\sigma}_t^{Z_i Z_i} = (\Delta \ln Z_{i,t} - \Delta \ln \hat{Z}_{i,t})^2$. Given the number of additional terms to be tested (25) and the possible problems of multicollinearity, *F*-tests for omitted variables have been carried out on each conditional share equation model, considering the variables in blocks. That is the null of super exogeneity has been tested sequentially in three stages. In the first stage the conditioning regressors have been tested for weak exogeneity by augmenting the each share equation conditional model with the terms $\hat{\gamma}_t^{z_i} = \left(\hat{\sigma}_t^{z_i z_i}\right)^{1/2}$ and testing their significance. Analogously, in the second stage a test for invariance has been carried out by augmenting each conditional share equation model by the terms $\left(\hat{\mu}_t^{z_i}\right)^2$, $\hat{\mu}_t^{z_i}\,\hat{\sigma}_t^{z_i z_i}$, $\hat{\sigma}_t^{z_i z_i}$ and their statistical significance has been jointly tested. Finally, in the third stage a test for parameter constancy has been carried out by checking the joint statistical significance of the set of terms $\hat{\sigma}_t^{z_i z_i}\,\hat{\gamma}_t^{z_i}$.

Table 11: Quarterly and yearly conditional models, *F*-tests.[10]

null hypot.	Quarterly data			Yearly data		
	###So	**###Se**	**###Sg**	**###So**	**###Se**	**###Sg**
weak exog.	2.18 [.07]	2.16 [.07]	0.66 [.66]	2.56 [.06]	0.79 [.56]	0.79 [.57]
invariance	1.16 [.34]	1.70 [.09]	0.32 [.99]	2.70 [.08]	1.99 [.16]	0.97 [.54]
stability	1.71 [.15]	1.10 [.37]	0.26 [.93]	2.28 [.09]	1.00 [.45]	0.96 [.47]

According to the sequential testing procedure followed, the hypothesis of super exogeneity would not be rejected for both the quarterly and yearly models at the 5% level of significance. The calculated statistics are, however, quite close to the rejection zone for the quarterly share of electricity equation and for the yearly share of oil model. Moreover, as suggested by Psaradakis and Sola (1996), tests of super exogeneity may suffer from very low power, so that testing this hypothesis in different ways is advised. Thus, an additional test for invariance, consisting of testing the significance of the interventions variables considered in the marginal equations in the conditional model has been carried out. That is, system (11) has been augmented by the set of dummies $\mathbf{D}'_t = \left[i81q4_t \quad i86q1_t \quad s86q1_t \quad i88q1_t \quad i92q1_t \quad i92q2_t \quad i92q4_t\right]$ in the

[10]: Quarterly data: ###So: weak exo. = F(5,48); invariance = F(15,38); parameter stability = F(5,48); ###Se: weak exo. = F(5,49); invariance = F(15,39); parameter stability = F(5,49); ###Sg: weak exo. = F(5,49); invariance = F(15,39); parameter stability = F(5,49). Yearly data: ###So: weak exo. = F(5,18); invariance = F(15,8); parameter stability = F(5,18); ###Se: weak exo. = F(5,18); invariance = F(15,8); parameter stability = F(5,18); ###Sg: weak exo. = F(5,19); invariance = F(15,9); parameter stability = F(5,19).

quarterly specification and by $\mathbf{D}_t' = [i74_t \quad s74_t \quad i86_t \quad s86_t]$ in the yearly one.[11] With the prefixes *i* and *s* have been indicated the dummies of the *i*mpulse and *s*tep type, respectively.

Invariance requires that the instability detected in the marginal model should not affect the conditional one. As is shown in table 12, only in the quarterly system have some of the dummy variables been found significant. In particular, in the shares of electricity and natural gas have been found significant the impulse dummies for the years 1981 and 1992, which were included in the marginal models to account for the 1979-1980 oil crisis and for the 1992 *oil in the ground* OPEC policy.

Table 12: Quarterly conditional model, *t*-tests.

0bs.	###So			###Se			###Sg		
	Coeff.	HCSE	t-value	Coeff.	HCSE	t-value	Coeff.	HCSE	t-value
81-q4				-0.019	0.0032	-6.0	0.0146	0.0030	4.87
92-q1				-0.119	0.0025	-4.8	0.0109	0.0022	4.96
92-q4							0.0120	0.0025	4.8

On the basis of the lack of invariance, therefore, the super exogeneity hypothesis could be rejected only for the quarterly conditional model. Thus, the yearly model should be preferred for policy analysis.

6 Concluding remarks

Over the 35 years considered in the analysis the primary energy sources cost shares have shown considerable variation, with the shares of oil and natural gas increasing by 25% and 13%, respectively, and the shares of electricity and coal falling by 35% and 3%, respectively. The analysis conducted has shown that both relative price and biased technical progress effects would have played a role in determining these dynamics, although the very low estimated price elasticities would indicate that primary energy inputs are not very responsive to price changes, even in the long-run.

[11]: These dummies were included in the specification of the marginal models.

In addition, a slow but steady substitution process of oil and coal for natural gas and electricity would appear to be at work. The direction of this substitution process would seem to be appropriate to face the present climate change concern, since methane gas has a lower carbon content than oil and coal. These findings suggest that, in addition to relative prices, other key factors may have played a relevant role in determining the dynamics of the energy market. As a consequence, policies aiming at reducing carbon emissions by only taxing energy sources could require very large tax rates. Modelling data sampled at different frequencies side by side has provided interesting results. In particular, yearly data have shown to provide more reliable information about the long-run aspects of the energy market than the quarterly ones. The same conclusion would seem to hold for policy analysis as well. This result is, to a certain extent, not surprising and give strength to the view that to determine the deep feature of the economy a longer time span is more relevant than a larger sample size obtained by using higher frequency data but sampled on a shorter time span.

Appendix

A1 Data construction

Energy Prices

Source: OECD (IEA): Energy Prices and Taxes

Frequency: Quarterly (seasonally unadjusted): 1978q1-1994q4. Yearly: 1960-1994.

Definition: Industry end-user prices (MTOE) in national currency.

Po: oil industry price; the series is a weighted average of the HFOIP (heavy oil fuel industry price)and LFOIP (light oil fuel industry price) series, with weights calculated from the Quarterly Oil Statistics and Energy Balances (OECD).

Pg: natural gas industry end-user price

Pc: steam coal industry end-user price

Pe: electricity industry end-user price

Energy Quantities:

Source: OECD (IEA): Quarterly Oil Statistics and Energy Balances and Energy Balances of OECD Countries.

Frequency: Quarterly (seasonally unadjusted): 1978q1 1994q4. Yearly: 1960-1994.

Definition: Quantities are expressed in million tons of oil equivalent (MTOE), and refers to primary energy supply.

$\mathbf{Q_c}$: coal

$\mathbf{Q_o}$: oil

$\mathbf{Q_g}$: natural gas

$\mathbf{Q_e}$: electricity; the series includes the generation of electricity from hydro/geothermal, nuclear, and the provision from electricity (net) trade.

A2 Data analysis

The statistical properties of the series have been assessed by ADF (Dickey and Fuller, 1981) and HEGY (Hylleberg, Engle, Granger and Yoo, 1990) tests.

Unambiguous evidences of unit roots at the zero frequency were found in all of the quarterly series, with the exception of the steam coal price series (lnP_c) and the share of natural gas (S_g). For the price of coal, non-stationarity seemed to be more of the deterministic type. For the share of natural gas the null of unit root at the zero frequency was rejected at the 5% confidence level, but not at the 1% level. None of the series showed stochastic seasonality at the 5% confidence level, and six series out of twelve showed signs of deterministic seasonality (lnP_e, S_o, S_e, S_c, S_g, lnE). The findings have been found to be quite robust to the introduction of seasonal break dummies in correspondence of the year 1986 to control for a structural change in the seasonal pattern. When the seasonal break dummies are considered, the rejection of the unit root at the zero frequency for the share of natural gas got stronger, reinforcing the importance of the deterministic trend component. For the energy price series, apart from the price of natural gas for which the null was rejected, the null of unit root at the zero frequency was not rejected at the 1% level. The remaining series did not seem to have been affected by the dummies used. Overall, the inclusion of the break dummies had the effect of pushing the statistics towards the rejection of the unit-root hypothesis, as Perron (1989) suggested.

For all of the yearly series, the unit root hypothesis could not be rejected at the 5% level. When the ADF auxiliary regression was augmented to control for possible segmented trends in correspondence of the years of the major oil shocks, only the share of coal showed a clear rejection of the null. For the price series of oil and electricity, the null was not rejected at the 1% level. For all of the other series, the introduction of the break dummies did not seem to affect the results significantly.[12]

Unit root tests, therefore, would suggest the plausibility of modelling the cost share equations system as non-stationary, treating all of the series as I(1) process. The works of Shiller and Perron (1985) and Pierse and Snell (1995), support this conclusion, showing that, in general, the power of unit-root tests depends more on the sample size than on the sample frequency. Then, a conclusion about the order of integration should preferably be drawn by low frequency data covering a longer time span. Finally, as far as seasonality is concerned, the HEGY tests would suggest a deterministic specification.

A3 Marginal models

Table A1: Yearly marginal models, estimation sample: 1963-1994.

variable	**###ln*Po***	**###ln*Pe***	**###ln*Pg***	**###ln*Pc***	**###ln*E***
	Coeff. (S.E)	**Coeff. (S.E)**	**Coeff. (S.E)**	**Coeff. (S.E)**	**Coeff. (S.E)**
Constant		-0.06 (.033)			0.04 (.013)
### ln*Po*_1		0.13 (.042)			
### ln*Po*_2		0.18 (.052)			0.16 (.038)
### ln*Pe*_1				0.62 (.166)	-0.19 (.068)
### ln*Pe*_2	-1.11 (.352)		-1.19 (.239)		
### ln*Pg*_2	0.63 (.170)		0.52 (.122)		-0.15 (.049)
### ln*Pc*_1					
### ln*E*_1	1.40 (.619)	0.73 (.328)			0.44 (.145)
### ln*E*_2	-1.49 (.598)			0.76 (.218)	
i1974	0.65 (.137)				
s1974	0.22 (.043)	0.28 (.041)	0.34 (.043)	0.07 (.039)	
i1986	-0.98 (.130)	-0.19 (.055)	-0.85 (.098)		
s1986		-0.16 (.032)	-0.20 (.052)	-0.12 (.043)	-0.03 (.013)

12: Since the relative price series do not show clear break points, only standard ADF tests have been carried out.

Tale A2: Quarterly marginal models, estimation sample: 1979q2-1994q4.

variable	**###ln*Po***	**###ln*Pe***	**###ln*Pg***	**###ln*Pc***	**###ln*E***
	Coeff. (S.E)	**Coeff. (S.E)**	**Coeff. (S.E)**	**Coeff. (S.E)**	**Coeff. (S.E)**
Constant	0.04 (.010)	0.01 (.005)			
###		0.16 (.037)			
ln*Po*_1					
###	-0.27 (.122)	0.13 (.038)			
ln*Po*_2					
###		0.18 (.052)			
ln*Po*_3					
###			-0.26 (.092)		
ln*Po*_4					
###				0.64 (.187)	
ln*Pe*_1					
###		-0.21 (.085)	0.49 (.119)		
ln*Pe*_2					
###		0.34 (.078)	0.58 (.107)	0.40 (.181)	
ln*Pe*_4					
###					
ln*Pg*_1					
###	0.35 (.106)				
ln*Pg*_2					
###	0.19 (.075)	-0.11 (.046)			
ln*Pg*_3					
###			0.23 (.080)		
ln*Pg*_4					
###		0.20 (.037)		-0.38 (.087)	
ln*Pc*_1					
###	0.16 (.062)		0.22 (.047)	-0.21 (.081)	-0.36 (.109)
ln*E*_1					
###	0.58 (.185)				-0.51 (.109)
ln*E*_2					
###					-0.35 (.108)
ln*E*_3					
###	0.45 (.185)	0.16 (.036)			0.48 (.109)
ln*E*_4					
i81q4		-0.15 (.036)		-0.26 (.071)	
i86q1	-0.70 (.086)		-0.81 (.057)	0.31 (.082)	
s86q1			0.03 (.010)		
i88q1	-0.48 (.076)	-0.08 (.034)	-0.27 (.058)	-0.21 (.082)	
i92q1			-0.21 (.062)		
i92q2	-0.17 (.074)			-0.27 (.081)	
i92q4		-0.09 (.036)		0.22 (.084)	-0.11 (.038)

Table A3: Quarterly marginal models, properties of the residuals.[13]

variable	###ln*Po*	###ln*Pe*	###ln*Pg*	###ln*Pc*	###ln*E*
s	0.0709	0.0319	0.0551	0.0798	0.0372
AR 1-8	0.89 [.5361]	0.55 [.8119]	1.33 [.2520]	0.90 [.5282]	2.50 [.0228]*
ARCH 4	1.20 [.3238]	1.58 [.1967]	1.04 [.3953]	0.23 [.9183]	0.36 [.8370]
Norm.	2.55 [.2789]	1.14 [.5643]	4.40 [.1109]	2.60 [.2730]	3.84 [.1468]
Het.	0.57 [.8807]	1.45 [.1760]	2.37 [.0169]	1.36 [.2175]	0.98 [.4666]

Table A4: Yearly marginal models, properties of the residuals.[14]

variable	###ln*Po*	###ln*Pe*	###ln*Pg*	###ln*Pc*	###ln*E*
s	0.1252	0.0508	0.0912	0.0786	0.0281
AR 1-2	0.88 [.4299]	1.73 [.1997]	0.48 [.6271]	0.13 [.8753]	2.67 [.0900]
ARCH 1	0.93 [.3456]	1.69 [.2077]	0.17 [.6807]	0.07 [.7951]	0.08 [.7819]
Norm.	0.16 [.9245]	1.93 [.3811]	2.97 [.0488]	4.64 [.0982]	0.64 [.7240]
Het.	0.53 [.8536]	0.91 [.5561]	0.34 [.9234]	0.61 [.7173]	0.83 [.5990]

A4 Weak exogeneity analysis

Weak exogeneity tests were carried out by testing the significance of the error correction terms one at a time and jointly in each marginal equation using *F*-tests for omitted variables.

Table A5 reports the outcome of the single error correction term analysis, while table A6 gives the joint error correction terms *F*-tests.

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[13]: ###ln*Po*: AR 1-8 = F(8,45); ARCH 4 = F(4,45); Normality = c^2(2); Heteroscedasticity = F(15,37). ### ln*Pe*: AR 1-8 = F(8,41); ARCH 4 = F(4,41); Normality = c^2(2); Heteroscedasticity = F(23,25). ###ln*Pg*: AR 1-8 = F(8,43); ARCH 4 = F(4,43); Normality = c^2(2); Heteroscedasticity = F(19,31). ###ln*Pc*: AR 1-8 = F(8,46); ARCH 4 = F(4,46); Normality = c^2(2); Heteroscedasticity = F(14,39); ###ln*E*: AR 1-8 = F(8,50); ARCH 4 = F(4,50); Normality = c^2(2); Heteroscedasticity = F(9,48).

[14]: ###ln*Po*: AR 1-2 = F(2,23); ARCH 1 = F(1,23); Normality = c^2(2); Heteroscedasticity = F(11,13). ### ln*Pe*: AR 1-2 = F(2,22); ARCH 1 = F(1,22); Normality = c^2(2); Heteroscedasticity = F(11,12). ###ln*Pg*: AR 1-2 = F(2,25); ARCH 1 = F(1,25); Normality = c^2(2); Heteroscedasticity = F(7,19). ###ln*Pc*: AR 1-2 = F(2,26); ARCH 1 = F(1,26); Normality = c^2(2); Heteroscedasticity = F(6,21). ###ln*E*: AR 1-2 = F(2,14); ARCH 1 = F(1,24); Normality = c^2(2); Heteroscedasticity = F(9,16).

Table A5: Weak exogeneity tests (*F*-tests).[15]

***F*-test**	**Quarterly data**			**Yearly data**		
var.	**ECTSo_1**	**ECTSe_1**	**ECTSg_1**	**ECTSo_1**	**ECTSe_1**	**ECTSg_1**
### ln***Po***	2.3 [.14]	0.6 [.46]	0.2 [.63]	0.0 [.98]	0.0 [.89]	0.2 [.68]
### ln***Pe***	0.0 [.99]	0.0 [.86]	0.1 [.81]	0.0 [.98]	0.0 [.89]	0.1 [.78]
### ln***Pg***	2.2 [.15]	0.9 [.35]	0.1 [.75]	1.3 [.26]	1.2 [.28]	1.5 [.23]
### ln***Pc***	2.2 [.14]	5.3 [.03]	0.1 [.92]	0.6 [.45]	0.0 [.92]	1.6 [.21]
### ln***E***	0.2 [.66]	0.0 [.89]	0.9 [.35]	0.1 [.93]	0.2 [.67]	0.5 [.83]

In the table above the rows correspond to each marginal process and the terms ECTS*i*_1 $i = o, e, g$ refer to the residuals from the unrestricted static cointegrating regressions estimated previously. In brackets are reported the *p*-values of the corresponding *F*-tests. As is shown above, in none of the quarterly and yearly marginal processes have the error correction terms been found significant (1% level). In particular, the error correction term relative to the share of electricity (ECTSe_1) enters the price of coal equation with an implausible magnitude since the estimated speed of adjustment is above unity. The estimated parameter for this variable is, in fact, -1.31 with standard error equal to 0.57.

Table A6 shows the outcome of the weak exogeneity analysis carried out considering jointly the three error correction terms.

Table A6: Weak exogeneity tests, joint analysis.[16]

***F*-test**	**weak exogeneity**	
variable	**Quarterly data**	**Yearly data**
###ln*Po*	1.15 [0.34]	0.06 [0.98]
###ln*Pe*	0.03 [0.99]	0.18 [0.91]
###ln*Pg*	0.85 [0.47]	1.12 [0.36]
###ln*Pc*	2.10 [0.08]	1.11 [0.36]
###ln*E*	0.31 [0.82]	0.64 [0.60]

As is shown in the table above, the joint analysis supports the conclusions reported in table A5 that energy prices and total energy can be taken as weakly exogenous for the long-run

[15]: Quarterly data: ###ln*Po*: *F*(1,52); ###ln*Pe*: *F*(1,50); ###ln*Pg*: *F*(1,53); ###ln*Pc*: *F*(1,53); ###ln*E*: *F*(1,57). Yearly data: ###ln*Po*: *F*(1,24); ###ln*Pe*: *F*(1,23); ###ln*Pg*: *F*(1,26); ###ln*Pc*: *F*(1,27); ###ln*E*: *F*(1,25).

[16]: Quarterly data: ###ln*Po*: F(3,50); ###ln*Pe*: F(3,48); ###ln*Pg*: F(3,51); ###ln*Pc*: F(3,51); ###ln*E*: F(3,55). Yearly data: ###ln*Po*: F(3,22); ###ln*Pe*: F(3,21); ###ln*Pg*: F(3,24); ###ln*Pc*: F(3,25); ###ln*E*: F(3,23).

parameters. This result would validate the construction of an error correction model for each share equation, conditioned on relative prices and total energy or jointly modelling the three share error correction equations as a single system. This latter way has been followed in the analysis and in the second step the three share equations have been estimated jointly in the error correction form.

A5 Parameter stability analysis

Figures A1 and A2 report the outcome of the parameter constancy tests for the quarterly and yearly statistical systems, respectively. The calculated Chow tests statistics are scaled by the corresponding critical values, so that values above the straight line at unity indicate rejection of the null hypothesis of parameter constancy at the chosen level of significance. In the figures, the 1-step residuals have been reported as well.

Figure A1: Quarterly econometric model, stability analysis.

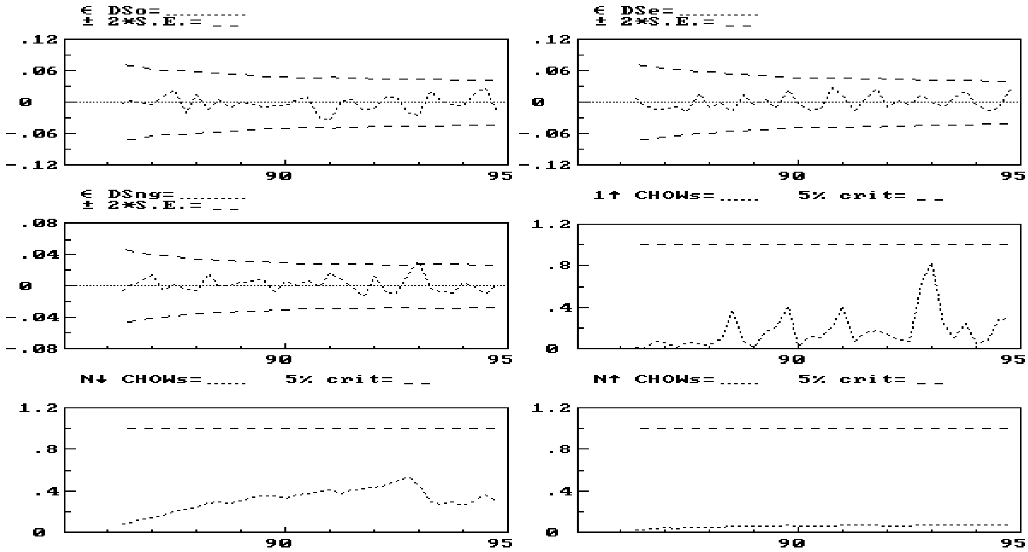
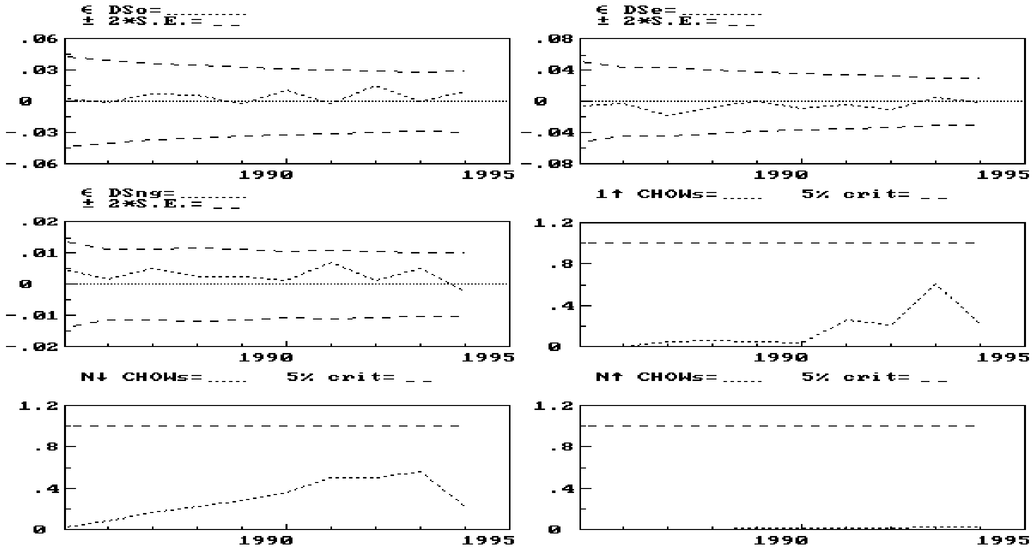


Figure A2: Yearly econometric model, stability analysis.



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