Short and long-run elasticities of nitrogen demand:

an application of cointegration to Italian data

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Elasticités à court et à long terme de la demande d'azote: la cointégration appliquée aux données italiennes

Résumé – Nous nous efforçons d'évaluer l'élasticité prix de la demande d'azote minéral en Emilie-Romagne, l'une des régions les plus touchées par l'eutrophisation des eaux. Il est important de savoir, pour des besoins de politique économique, si la demande d'engrais est élastique par rapport à son prix, lorsque l'on souhaite réduire la consommation d'intrants polluants. Or, les auteurs ne sont pas d'accord sur le niveau de l'élasticité prix de la demande d'azote. Nous utilisons un modèle dynamique à une équation, sous contrainte de maximisation du profit, où la demande optimale d'azote dépend de son prix, des prix escomptés à la production et des surfaces occupées. Les modèles dynamiques à une équation sont utilisés depuis longtemps pour mesurer la demande d'engrais, la relation d'équilibre étant associée à une hypothèse d'ajustement partiel. Afin d'éviter les problèmes afférents à cette hypothèse, nous estimons les paramètres d'équilibre à l'aide d'un modèle à correction d'erreurs (ECM). Pour cela, il est indispensable que les variables utilisées soient cointégrées. Nous vérifions la cointégration de deux façons: en utilisant les régressions statiques d'Engle-Granger ou la méthode du maximum de vraisemblance mise au point par Johansen.

Avec ce modèle à correction d'erreur:

1) l'ajustement à une situation d'équilibre de court terme est d'environ 29% par an;

2) l'élasticité prix de la demande d'azote est de – 0,87 à court terme, et de – 1,82 à l'équilibre;

3) une hausse de 2% du prix réel de l'engrais conduirait à faire baisser sa demande de 1,6% à court terme et de 3,8% à long terme;

4) une réduction prolongée des surfaces utilisées aboutit à réduire la consommation d'azote à long terme, même si l'élasticité par rapport aux facteurs fixes est nettement inférieure à 1.

En simulant la mise en vigueur d'une taxe verte, nous montrons que la consommation d'azote a des chances d'être considérablement réduite quand on augmente son prix de 10 à 20%. Nous constatons aussi que le recours à une taxe peu élevée, mais sur une longue période, a plus d'incidence sur la consommation d'azote qu'une taxe plus élevée. Cela signifierait donc qu'une évolution de la technologie peut être induite par une hausse prolongée des prix.

Mots-clés:
politique de l'environnement, demande d'engrais, Italie, cointégration, modèles à correction d'erreur, taxe verte, élasticité prix

Short and long-run elasticities of nitrogen demand: an application of cointegration to Italian data

Summary – *This is an attempt to measure the price elasticity of agricultural demand for chemical nitrogen in Emilia Romagna, one of the Italian regions worst hit by the problem of water eutrophication. We build up a single equation model and use the theory of cointegration and an error correction model to estimate the price elasticity of demand for nitrogen. Nitrogen turns out to be elastic, in the long run, both to its own price and to the expected output price. A simulation exercise has been carried out for the implementation of a "green tax".*

Key-words: *nitrogen demand, cointegration, error correction models, green taxes*

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THIS paper deals with the problem of water pollution by fertilisers in the Adriatic sea facing Emilia Romagna in Italy[1]. According to recent studies, 40 % of the eutrophication of the Adriatic sea is caused by agricultural chemicals (Agassi, 1992) most of which has been attributed to phosphates. Two aspects seem to be particularly relevant in this context:

a) the Adriatic Sea between Chioggia e Fiorenzuola shows the highest level of eutrophication in Italy (Regione Emilia Romagna, 1984).

b) The consumption of chemicals, in Emilia Romagna, is much higher than the average national level and there has been a substitution of nitrogen fertilisers for phosphates (Regione Emilia Romagna, 1984).

We believe that nitrogen fertiliser has a certain impact on the Adriatic problem for the following reasons:

1) Despite strict regulation of organic fertilisers (manure) the problem of eutrophication of the Adriatic sea has not improved over the last decade.

2) There has been a sharp increase in the consumption of manure against chemicals (Agra Europe, 1991).

3) There has been, over the last eighteen years, a substitution of nitrogen fertilisers for phosphates (OECD, 1989). It is interesting to note that the quantity of nitrogen fertiliser in the soil has risen from 192.058 tons in 1950 to 1,121.709 tons in 1989 with an increase of 584 %. The quantity of nitrogen fertiliser in the soil rose from 15 kg/ha in 1950 to 90.8 kg/ha in 1989.

4) According to OECD indicators, the quality of Italian water, in particular of the Po river, expressed by the ratio between oxygen content and nitrate content of river waters, in mgN/l, is deteriorating due to the higher content of nitrate compounds.

The aim of this paper is to estimate the long-run price elasticity of demand for nitrogen. This is a crucial parameter when dealing with issues of environmental policy, for instance in order to use economic instruments, such as taxes, to reduce the consumption of polluting inputs.

We will use a single equation dynamic model consistent with profit maximization in which the optimal nitrogen demand depends on its price, the expected output price and the quantity of land. The use of single equation dynamic models to estimate fertilizer demand is a long established tradition which goes back to the seminal work of Griliches

[1] I am in debt to Prof. Rizzi for his continuous help and advice. He has constantly supervised and corrected my work. I also wish to thank two anonymous referees of this journal. Any mistake is, however, my own responsibility.

(1958), in which the equilibrium relation is combined with a partial adjustment hypothesis. The problem with this model is that, when there is a more general dynamic structure (Hendry *et al.*, 1984), the estimation of the long-run elasticities is going to be biased. This is perhaps what explains the great variability of results obtained with the partial adjustment hypothesis (Burrell, 1989).

In order to avoid these difficulties we will obtain the equilibrium parameters using an error correction model (ECM). Before specifying the ECM we should make sure the variables are cointegrated. We do that in two ways: with Engle and Granger (1987) static regressions and with the maximum likelihood procedure developed by Johansen (1988) and Johansen and Juselius (1990).

From the ECM representation, the direct price elasticity of demand for nitrogen in Emilia Romagna agriculture over the period 1954-1990 is – 0.87 in the short run (for one year) and – 1.82 in equilibrium. In the final part of our work we will value the change in nitrogen consumption when a green tax is temporarily applied.

OPTIMAL DEMAND FOR NITROGEN AND COINTEGRATION

Starting from a variable profit function as a dual representation of a production function, we obtain the optimal demand of each input by differentiating the negative of the profit function with respect to the price of each input (Chambers, 1988). In this way the optimal demand for nitrogen (Q_N) depends on its price (P_N), on the vector of other variable inputs prices ($\boldsymbol{P}_X$), on the expected output price (P_Y) and on the quantity of the fixed input land (SA):

$$Q_N = f(P_N, \boldsymbol{P}_X, P_Y, SA)$$

Moreover if we assume strong separability of the inputs we can isolate the demand for nitrogen from the rest of the system which now becomes: $Q_N = f(P_N, P_Y, SA)$. As an empirical version of $f(.)$ we use a linear approximation in the logs of the variables:

$$qn = \beta_n pn + \beta_y py + \beta_a sa \tag{1}$$

where lower case letters denote $\log_e$ of corresponding capital letters whereas β_n and β_y are the coefficients of the long-run price elasticity of demand. In this *ad hoc* demand equation only the essential variables are present, as suggested by economic theory, but we do not impose homogeneity in the prices: $\beta_n + \beta_y = 0$

The fact that economic theory suggests a static relationship such as (1), which is valid in the long run, makes it difficult to find out the value of the β elasticities, because the data describe an adjustment towards equilibrium. In recent years however, Engle and Granger (1991) and many others have developed the statistical notion of time series cointegration, which can be considered as corresponding to the concept of long-run economic equilibrium [2].

LONG- AND SHORT-RUN PRICE ELASTICITIES OF NITROGEN DEMAND

Two variables (but this is also true in general), y_{1t} and y_{2t} are said to be cointegrated of order (p,q) if: i) they are integrated of the same order p, $I(p)$; ii) there exists a value such that the linear combination $y_{1t} - \alpha y_{2t} = u_t$ is $I(p\text{-}q)$. Usually we assume $p = q = 1$. In this case, y_{1t} and y_{2t} will each have an infinite variance if they are $I(1)$, but their linear combination, u_t, will be stationary $I(0)$, if they are cointegrated. Given that economic theory postulates stationary relationships, the existence of cointegration among the variables relevant for the theory under examination, is the necessary condition in order to be able to estimate equilibrium relationships.

In order to implement our analysis, we use annual data 1951-1990. The regional consumption of nitrogen (in tons) for six types of fertilisers and the total, the aggregate price of output (index number 1985 = 1.0) and the land cultivated (ha) are taken from Istituto Nazionale di Statistica (ISTAT); nitrogen prices, for six fertilisers, are taken from Scartoni (1990). The price of nitrogen P_N has been obtained as a translog aggregate (1985 = 1.0) of the six elementary prices, as in Rizzi (1986). Under the hypothesis of static expectations, the output price is given by the implicit deflator of gross regional output at the beginning of the period: $Py_t = Py_{t-1}$. From what we have already postulated, the first thing to do, in order to estimate the equilibrium parameters in the demand for nitrogen, is to verify that the variables (qn_t, pn_t, py_t, sa_t) are cointegrated. This means that we first have to verify whether they are integrated of the same order and then if there are cointegrating vectors.

ORDER OF INTEGRATION OF THE VARIABLES

There are many tests available to check the order of integration of a time series, parametrical and non parametrical tests, which assume as the null hypothesis the presence of a unit root or stationarity, but none

[2] The statistical theory of integration and cointegration of time series has produced in recent years a large number of contributions, often of a very technical nature. Besides the essays contained in the Engle and Granger volume (1991), other references are: Banerjee *et al.* (1993) and Davidson and Mackinnon (1993).

of them is particularly powerful in small samples (Davidson and MacKinnon, 1993). In this paper we follow a sequential procedure starting from the null hypothesis of 2 unit roots (Dolado *et al.*, 1990).

Test 1: **2 versus 1 unit roots**

$$\Delta^2 x_t = \alpha_0 + \beta_2 \Delta x_{t-1} + \varepsilon_t$$

compare the t-value of β_2 with the critical value of -2.93 at the 5% significance level; if $t(\beta_2) < -2.93$, we reject the null hypothesis of two unit roots and go to test 2.

Test 2: **1 versus 0 unit roots**

$$\Delta^2 x_t = \alpha_0 + \beta_1 x_{t-1} + \beta_2 \Delta x_{t-1} + \varepsilon_t$$

compare the t-value of β_1 with the critical value of -2.93 at the 5% level of significance; if $t(\beta_1) < -2.93$, we reject the hypothesis of 1 unit root in favour of stationarity; if we accept the null hypothesis of a $I(1)$ variable we go to test 3.

Test 3: **Random walk versus random walk with drift**

$$\Delta x_t = \gamma_0 + \gamma(L) \Delta x_t + \varepsilon_t$$

compare the t-value on intercept against the critical value given in traditional t-distribution tables at the 5% level of significance 1.68. If $t(\gamma_0) > 1.68$ we reject the null hypothesis of a zero intercept and go to test 4.

Test 4: **Random walk with drift versus deterministic trend**

$$\Delta x_t = \gamma_0 + \delta_1 x_{t-1} + \delta_{2t} + \varepsilon_t$$

compare the F-value associated with setting the coefficients on lagged level and trend equal to zero against critical values given in Dickey-Fuller (1981, table VI). If the F test turns out to be significant, the series is trend stationary. We present below the table of sequential tests performed on the variables.

Table 1. Sequential unit root tests

	qn	*pn*	*py*	*sa*
Test 1	– 4.70*	– 3.50*	– 3.36*	– 3.96*
Test 2	– 1.67	– 0.14	0.44	– 0.33
Test 3	1.74*	1.37	2.12*	– 1.27
Test 4	3.01		3.09	

* indicates statistically significant values

For all variables we cannot reject the hypothesis of one unit root and there is evidence of a drift term in the quantity of nitrogen and in the output price.

We have also performed the augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root test (Davidson, MacKinnon, 1993). Results of the ADF and PP tests based on the t-statistics are summarized below. The critical value for trended variables is – 3.53 according to the MacKinnon tables (1991). As we can see from table 2 we accept the null hypothesis of one unit root in the variables. Therefore both the ADF and the PP tests lead to the conclusion that the variables are $I(1)$.

Table 2.
Augmented Dickey-Fuller and Phillips-Perron unit root tests

	qn (1)	*pn* (2)	*py* (2)	*sa* (1)
Augmented Dickey-Fuller test	– 3.308	– 2.185	– 2.479	– 2.704
Phillips-Perron test	– 2.514	– 1.856	– 1.786	– 2.112

(.) is the number of included lags, critical value at the 5% of significance – 3.53 (MacKinnon, 1991).

COINTEGRATION ANALYSIS

Having found that the variables in the vector $y_t = (qn_t, pn_t, py_t, sa_t)$ are $I(1)$, we have tried to verify whether they are cointegrated, that is whether it is possible to find at least a linear combination which is $I(0)$; in order to do that we have used two methods, among those available: the Engle and Granger procedure (1987) and the VAR method by Johansen (1988).

The Engle and Granger procedure implies the estimation of four OLS (ordinary least squares) static regressions, using each time one of the four components of y_t as the dependent variable and the other three as independent. In each regression we verify the residuals' stationarity. Under the null hypothesis of non-cointegration the residuals will be $I(1)$, whereas under the alternative hypothesis they will be $I(0)$.

In table 3 we present the coefficients of the four static regressions normalized over *qn*, which represent a super-consistent estimation of the long-run elasticities, if the residuals are stationary, and three cointegration tests based on the residuals: the Durbin-Watson test (CRDW: cointegrating regression D. W.), and in table 4 we present two non parametric tests by Phillips and Ouliaris (1990). From these two estimations, the long-run nitrogen demand seems to be elastic with respect to its price, with a value going from – 1.37 to – 1.96. On the other hand, according to the CRDW test, the hypothesis of non cointegration could be rejected, but the other two tests do not allow to reject the null hypothesis of non cointegration even at the 10% level of confidence[(3)].

[(3)] However if one tries to add power to the tests by imposing a priori constant returns to scale in the long run, $\beta_{ij} = 1$, the null hypothesis of non cointegration of $y_t\ (qn_t - sa_t,\ pn_t,\ py_t)$ could be refused even using the non parametric tests. In this case, β_{ij} approximately equals – 3.0.

Table 3. **Cointegrating regressions** (coefficients normalized over qn)

Regressand/ regressors	qn	pn	py	sa
qn	1.0	1.0	1.0	1.0
pn	– 1.380	– 1.961	– 1.753	– 1.366
py	1.736	2.204	2.054	1.728
sa	0.792	0.784	0.789	0.793
SER%	15.41	9.37	8.16	19.44
CRDW	0.804	0.821	0.863	0.798

Table 4. **Phillips-Ouliaris tests on residuals** ($\Delta u_t = (\rho - 1)u_{t-1} + const$)

	qn	pn	py	sa
$\rho - 1$	– 0.417	– 0.405	– 0.427	– 0.414
t_{ADF}	– 3.52	– 3.12	– 3.26	– 3.51
$Z\alpha$	– 18.89	– 16.64	– 18.14	– 19.19
SER%	12.34	7.63	6.79	15.52

Critical values 10% $t_{ADF} = -4.03$ (MacKinnon, 1991)
$Z\alpha\,(T = 50) = -27.28$ (Haug, 1992, table 2).

According to the Johansen procedure we need to estimate a VAR model with p lags of vector y_t reparametrised as:

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + \dots + \Gamma_{p-1} \Delta y_{t-p+1} - \Pi y_{t-p} + v_t \qquad (2)$$

subject to the hypothesis that matrix Π has rank $r < 4$, where r indicates the number of cointegrating vectors. In order to verify the hypothesis on rank r we use trace and maximum eigenvalue statistics. In the trace test, the null hypothesis is that the number of cointegrating vectors be $\leq r$, ($r = 0, 1, 2, 3$); however the null hypothesis is tested against the general alternative. The maximum eigenvalue statistics is similar to trace statistics, but the alternative hypothesis is explicit; the null $r = 0$ is tested against $r = 1$, $r \leq 1$ is tested against $r = 2$ and so on. From the estimation of the VAR model with two lags we obtain the values presented in table 5. Starting from the top we reject $r = 0$ (no cointegration) against finding at least one cointegrating vector, *i.e.*, the trace and max statistics exceed their 5% critical values. In the second row we find that the calculated statistics do not exceed their 5% critical values. Thus the procedure defines $r = 1$, one cointegrating vector, that we consider as the long-run nitrogen demand.

The maximum likelihood estimations of β_n and β_y are – 1.337 and 0.842[(4)].

(4) If we set the constraint $\beta_a = 0.8$, we obtain $\beta_n = -1.919$ and $\beta_y = 2.161$, but the constraint is rejected with $\chi^2\,(1) = 12.76$. If we set homogeneity in the prices together with $\beta_a = 0.8$, we obtain $\beta_n = -\beta_y = -3.373$ with $\chi^2\,(2) = 14.87$.

Table 5.
Johansen tests for the number of co-integrating vectors
(1953-1991)

Eigenvalues: 0.54,	0.34,		0.26, 0.00008				
	Trace statistic			Max. eigenvalues statistics			
H_0	H_1	score	(0.95)	H_0	H_1	score	(0.95)
$r = 0$	$r \geq 1$	57.62*	47.21	$r = 0$	$r = 1$	29.92*	27.07
$r \leq 1$	$r \geq 2$	27.70	29.68	$r \geq 1$	$r = 2$	15.87	20.97
$r \leq 2$	$r \geq 3$	11.84	15.41	$r \geq 2$	$r = 3$	11.8	14.07
$r \leq 3$	$r = 4$	0.003	3.76	$r \geq 3$	$r = 4$	0.003	3.76

Max. lag in VAR $p = 2$; trended case with trend in GDP
* indicates statistical significance at the 5 % level

MODELLING NITROGEN DEMAND AS AN ECM

Although the two procedures give similar estimations of the long-run price elasticities only for β_n, the Johansen test has proved the existence of a single cointegrating vector. Thus the Engle and Granger results (1987) guarantee that the change in qn_t has an ECM representation:

$$qn_t - \beta_n pn_t - \beta_y py_t - \beta_a sa_t = u_t$$

From a theoretical point of view, Salmon (1982) and Nickell (1985) have shown that the error correction model can be derived from the dynamic optimising behaviour of economic agents. From an empirical standpoint, we specify the nitrogen demand as an ECM because this representation allows more flexibility in short-run dynamics while the model is constrained to return to long-run equilibrium. We obtain the ECM as a reparametrization of an autoregressive distributed lags (ADL) model (Banerjee *et al.*, 1990, Davidson and MacKinnon, 1993).

An ADL (2, 2, 3, 2) for nitrogen consumption can be written as:

$$\begin{aligned} qn_t = {} & A_1 qn_{t-1} + A_2 qn_{t-2} + B_{10}\, pn_t + B_{11}\, pn_{t-1} + B_{12}\, pn_{t-2} \\ & + B_{21}\, py_{t-1} + B_{22}\, py_{t-2} + B_{23}\, py_{t-3} + B_{30}\, sa_t \\ & + B_{31} sa_{t-1} B_{32} sa_{t-2} + const + e_t \end{aligned} \tag{3}$$

with $0 < (1 - A_1 - A_2) < 1$.

Without the equilibrium terms the variables should converge to their stationary long-run value which is given by:

$$\begin{aligned} qn[1 - A_1 - A_2] = {} & [B_{10} + B_{11} + B_{12}]\, pn \\ & + [B_{21} + B_{22} + B_{23}]\, py \\ & + [B_{30} + B_{31} + B_{32}]\, sa \end{aligned} \tag{4}$$

solving for qn as a function of pn, py, sa we obtain:

$$qn = \beta_n\, pn + \beta_y\, py + \beta_a\, sa \tag{5}$$

where the long-run elasticities are:

$$\begin{aligned} \beta_n &= [B_{10} + B_{11} + B_{12}] / [1 - A_1 - A_2] \\ \beta_y &= [B_{21} + B_{22} + B_{23}] / [1 - A_1 - A_2] \\ \beta_a &= [B_{30} + B_{31} + B_{32}] / [1 - A_1 - A_2] \end{aligned} \tag{6}$$

On this basis, we can reformulate the ADL model in the ECM form as follows:

$$\begin{aligned} \Delta qn_t = {} & \gamma\,[qn_{t-2} - \beta_n\, pn_{t-2} - \beta_y\, py_{t-3} - \beta_a sa_{t-2}] + \phi_1 \Delta pn_t \\ & + \phi_2 \Delta pn_{t-1} + \phi_3\, \Delta py_{t-1} + \phi_4\, \Delta py_{t-2} + \phi_5\, \Delta sa_t + \phi_6\, \Delta sa_{t-1} \\ & + \phi_7\, \Delta qn_{t-1} + const + e_t \end{aligned} \tag{7}$$

with:

$\gamma = -[1 - A_1 - A_2]$; $\phi_1 = B_{10}$; $\phi_2 = [B_{10} + B_{11}]$; $\phi_3 = B_{21}$

$\phi_4 = [B_{21} + B_{22}]$; $\phi_5 = B_{30}$; $\phi_6 = [B_{30} + B_{31}]$; $\phi_7 = -[1 - A_1]$

In equation 7, the term in brackets is the ECM; γ measures the adjustment towards equilibrium while the coefficients ϕ_j represent the short-run dynamics. In particular we interpret ϕ_1 and ϕ_3 as the short-run elasticities of nitrogen demand with respect to *pn* and *py*[5].

After some attempts on the non-constrained ADL model, we have set $A_1 = 1$, $B_{12} = B_{30} = B_{31} = 0$, which can be easily accepted with $\chi^2(4) = 3.14$; therefore the final version of the model, estimated with non linear least squares, is the following:

$$\begin{aligned} \Delta qn_t = {} & -\underset{(0.065)}{0.2094}\, [qn_{t-2} + \underset{(0.371)}{1.815} pn_{t-1} - \underset{(0.294)}{2.019} py_{t-3} - \underset{(0.007)}{0.781} sa_{t-2}] \\ & - \underset{(0.155)}{0.868} \Delta pn_t + \underset{(0.211)}{0.772} \Delta py_{t-1} + \underset{(0.231)}{1.467} \Delta py_{t-2} \end{aligned} \tag{8}$$

$T = 37$ $\quad R^2 = 0.59$ $\quad$ SER% $= 7.70$ $\quad$ DW $= 1.82$

The fit is acceptable with $R^2 = 0.59$, the Durbin-Watson statistics fall in the inconclusive region, but a LM (maximum likelihood) test for first order and second order autocorrelation refuses the hypothesis of serial correlation in the residuals with $\chi^2(1) = 2.17$ and $\chi^2(2) = 1.21$. The White test for heteroskedasticity and the ARCH (autoregressive conditional heteroskedasticity) test for conditional autoregressive heteroskedasticity produce respectively $\chi^2 = 0.101$ and $\chi^2 = 0.43$ thus rejecting the hypothesis of heteroskedasticity. A Ramsey RESET test for the func-

[5] We could collect the ADL terms in a different way and obtain an equivalent representation in which the ECM has the form $\gamma\,[qn_{t-1} - \beta_n\, pn_t - \beta_y\, py_{t-1} - \beta_a sa_t]$ and the short-run coefficients have a different structure (Banerjee *et al.*, 1990, Davidson and MacKinnon, 1993). In this case the coefficients would for instance take the form: $\phi_1 = -[B_{11} + B_{12}]$; $\phi_2 = -B_{12}, \ldots ; \phi_7 = -A_2$.

tional form gives a $\chi^2(1) = 2.179$ which allows us to assume that the functional form is correctly specified. All tests have been carried out at the 5 % level of significance[(6)].

From the ECM estimation we obtain the following results:

i) the adjustment to the short-run disequilibrium is about 29 % per year;

ii) the long-run demand for nitrogen is elastic both to its own price (– 1.82) and to the expected output price (2.02);

iii) in the short run, the elasticity of nitrogen demand with respect to its price is – 0.87 and 0.77 with respect to *py*;

iv) an increase in the real price of fertiliser by 2 % (due to an increase by 1 % in *pn* and to a decrease by 1 % in *py*) would cause a reduction of its demand by 1.6 % in the short run and by 3.8 % in the long run;

v) the continuous reduction in the quantity of agricultural land used determines a further reduction in the consumption of nitrogen in the long run, even though the coefficient for the elasticity with respect to the fixed factor is significantly smaller than 1 (0.78).

When we estimate the ECM imposing homogeneity in the prices ($\beta_n = -\beta_y$) the adjustment coefficient shrinks to $\gamma = -0.194$ (0.034), the long-run elasticities become $\beta_n = -\beta_y = -2.445$ (0.300), $\beta_a = 0.766$ (0.006), and the short-run elasticities become $\phi_1 = -0.919$ (0.134), $\phi_3 = 0.671$ (0.230). The likelihood ratio $\chi^2(1) = 2.65$ has a p-value of 0.104. Since the ECM is a more general dynamic representation than the partial adjustment hypothesis, we have estimated again the previous equation setting the coefficients of differentiated variables at zero. In this case the adjustment coefficient is $\gamma = -0.251$ (0.062) and the long-run elasticities become $\beta_n = -2.313$ (0.562), $\beta_y = 2.492$ (0.487), $\beta_a = 0.798$ (0.008). The likelihood ratio test is $\chi^2(3) = 19.93$, which clearly rejects the null hypothesis, in favour of the alternative hypothesis in which the change in the exogenous variables also helps to explain nitrogen consumption in the short-run. This having been demonstrated, we can conclude that the ECM without imposed homogeneity in the prices is the model most easily accepted by the data. This equation is also the basis of our simulation exercise.

A GREEN TAX ON NITROGEN CONSUMPTION

In this last section we describe the effect of a green tax on nitrogen consumption in the Emilia Romagna region. The basic scenario is obtained through a dynamic simulation (1954-2005) in which we assume

[(6)] Following the observations of Kremers *et al.* (1992) on the power of cointegration tests, it is interesting to notice that, if we use in the ECM the steady state solution calculated from the constrained ADL, the t statistics for γ ($\tau_{ecm} = -5.08$) exceed the critical MacKinnon (1991) value at the 5% level of confidence = – 4.41, thus confirming the existence of cointegration.

that, during the period 1991-2005, the prices P_n and P_y increase at an annual average rate of 0.010, whereas the land used decreases at a rate equal to – 0.0005. This simulation corresponds to column A in table 6. In the B column we have implemented, for 5 years (1994-1998) starting from 1994, a tax of 200 lire/kg, about 20% of the nitrogen price in 1985. Finally, in the C simulation, the tax is implemented for 10 years (1994-2003), but only for an amount of 100 lire/kg.

As we can see from table 6 (col. B), during the five years of implementation of the heavier tax nitrogen consumption decreases up to a maximum of 32%, but comes back to the pre-tax level within four years, after the tax has been phased out. On average, nitrogen consumption decreases by 27% in the period 1995-1999, but only by 15% in the period 1995-2003. The effect of a lower, but prolonged tax (17-18%), on the other hand, is more regular (tab. 6 col. C). With the lower tax, nitrogen consumption declines on average by 17% in the first 5 years (1995-1999) and by 16% in the period 1995-2003. In figure 1 QNT1 represents nitrogen consumption with the first tax scenario, whereas QNT2 represents consumption in the second tax scenario.

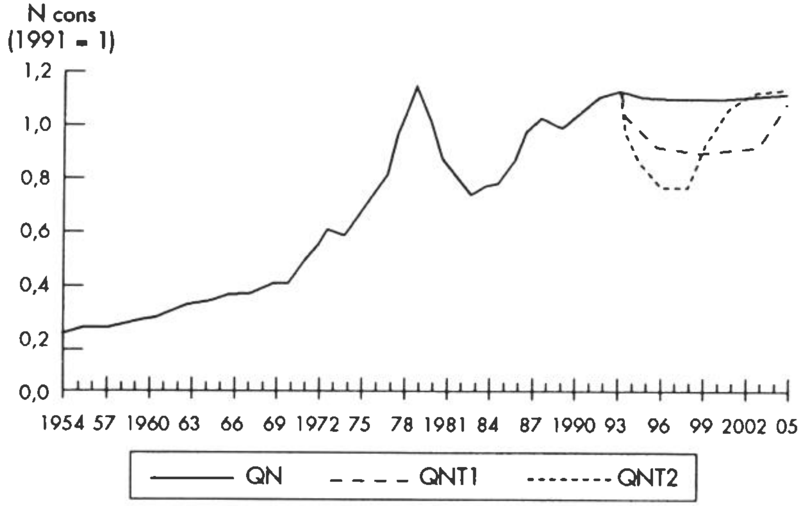
Table 6. **Simulated values for nitrogen consumption** in tons (1993-2005)

	A	B	(B/A)A%	C	(C/A)/A%
1993	131968	131968	0.00	131968	0.00
1994	131179	109070	– 16.85	119031	– 9.26
1995	130215	96799	– 25.66	111396	– 14.45
1996	129549	91002	– 29.75	107577	– 16.96
1997	129229	88743	– 31.67	105728	– 18.15
1998	129168	88259	– 31.67	105728	– 18.15
1999	129264	105646	– 18.27	105963	– 18.03
2000	129442	118342	– 8.57	106402	– 17.80
2001	129654	125791	– 2.98	106882	– 17.56
2002	129877	129377	– 0.39	107342	– 17.35
2003	130102	130760	0.51	107770	– 17.16
2004	130324	131133	0.62	118169	– 9.33
2005	130545	131159	0.47	125116	– 4.16
Mean values:					
1995/1999	129485	94090	– 27.34	107342	– 17.10
1995/2004	129682	110585	– 14.75	108328	– 16.47

A = basic simulation
B = tax of 200 lire/kg (20% of 1985 price) for 5 years (1994-1998)
C = tax of 100 lire/kg (10% of 1985 price) for 10 years (1994-2003)

Thus, a green tax should be an efficient means of reducing nitrogen consumption in Emilia Romagna, as the long-run demand is elastic to its own price. A lower tax implemented over a longer period seems to be preferred to the higher tax, thus suggesting an adaptation of the technology in agriculture to the new structure of the prices.

Fig. 1
Effects of taxation



CONCLUSIONS

There is no agreement in the literature as far as the price elasticity of demand for nitrogen is concerned. Using cointegration theory, we estimate a long-run relationship between the consumption of nitrogen and its price. In order to do so we use a fairly long series of data which give some degree of certainty to the results. As far as the methodology is concerned it should be noticed that cointegration amongst the variables allows us to use an ECM as a consistent representation of the dynamics of nitrogen demand, and thus to calculate the long-run parameters. We have shown that the demand for nitrogen is elastic to its price. By simulating the implementation of a green tax we have shown that the consumption of nitrogen can be significantly reduced by increasing its price by 10 or 20%. Our simulation results suggest that the effect on nitrogen consumption is greater when we apply a lower but prolonged tax rather than a higher tax. This suggests that changes in the technology might be induced when the increase in the price lasts longer.

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