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Regulatory Implications of Heterogeneity
-
the Case of
Nitrogen Fertilizer Demand from Danish Crop Farms

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Abstract

In this paper we estimate nitrogen fertilizer demand elasticities for Danish crop farms using the dual profit function approach on micro panel data. The model includes several farm specific parameters allowing us to estimate the mean demand elasticity and test for homogeneity of elasticities across panel farms. We find a mean own price elasticity for nitrogen of -0.45 and a significant standard deviation from this mean for individual farms of 0.24.

Heterogeneity of demand elasticities implies that regulating fertilizer application through mandated uniform percent reductions, as is currently used in Danish nitrogen regulation, increases abatement costs when compared to tax regulation. Somewhat surprisingly, this only causes the abatement costs of quota regulation to be 8% larger than with tax regulation. Simulation results indicate that the primary threat to the efficiency of uniform reduction schemes comes from inaccurate estimation of baselines rather than from heterogeneity of elasticities.

1. Introduction

Nitrate leaching is considered to be a serious environmental problem in Denmark, and agricultural application of nitrogen fertilizer has been regulated for a number of years. As in a number of other countries, Danish regulation of fertilizer application is essentially based on standards and norms.

Under the present regulations, crop farms are required to reduce nitrogen fertilizer application to 90% of the profit maximizing level, with each farm's baseline being calculated so as to take account of land quality, land allocation to each crop and crop rotation etc. on the individual farm. Noncompliant farmers are fined with a fee calculated as a progressive function of the magnitude of noncompliance. Depending on political priorities, if the reduction in farm profits and land values caused by such a system are substantially smaller than for tax regulation, (as is confirmed in this study), quotas may be viewed as having a distributional advantage over taxes. On the other hand, proportional reduction quotas are less efficient than e.g. a fertilizer tax, if fertilizer price elasticities vary across farms and/or baselines are inaccurately estimated. Furthermore, the quota system requires monitoring of detailed farm level data for calculation of baselines and for deterring illicit inter-farm fertilizer trading.

In this paper we quantify the abatement costs and distributional effects of reducing nitrogen fertilizer application on Danish crop farms through proportional reduction quotas, grand fathered tradable quotas and a fertilizer tax. This is done by estimating nitrogen fertilizer demand for Danish crop farms using the dual profit function approach on micro panel data. The model includes several farm specific parameters so that heterogeneity of elasticities among panel farms is allowed and can be tested against a hypothesis of homogeneous elasticities. We reject the homogeneous elasticity hypothesis and find a sizable variation in farm elasticities (with a standard deviation for individual farm elasticities of over 50% of the panel mean). Somewhat surprisingly, this only causes the abatement costs of quota regulation to be 8% larger than with tax regulation. Simulation results indicate that the primary threat to the efficiency of uniform reduction schemes comes from inaccurate estimation of baselines rather than from heterogeneity of elasticities.

Since regulation of agricultural nitrogen fertilizer application is also undertaken in a number of other countries, these findings may be of more general interest.

Our analysis may also be of methodological interest since we exploit the possibilities of our panel data by specifying a variant of the trans-log profit function that allows heterogeneity of elasticities and makes testing against a hypothesis of homogeneous elasticities possible. Although a number of recent papers have estimated the fertilizer demand of industrialised farmers, (e.g. Burrell (1989) - including a good review of older studies, Denbaly and Vroomen (1993), Rayner and Cooper (1994), Garcia and Randall (1994), Mergos and Stoforos (1997)), heterogeneity of elasticities and its regulatory implications has received little attention.

In section 2 we present the economic model and functional specification to be estimated. The data are described in section 3, and model estimation is presented in section 4. Section 5 summarises the results, and conclusions are drawn in section 6.

2. The Model

In the following we assume that crop farm production is described by a well-behaved

production function $f(x,n,z)$ where x is an aggregate crop output (positively signed), n is nitrogen fertilizer input (negatively signed) and z is input of cultivated land which is considered fixed for the scope and time horizon of the analysis. The production of the intermediate fixed cultivated land input is specified as follows:

$$z = z(\text{capital, labour, materials, uncultivated land})$$

so that the specified primary inputs can be dropped from the model since the intermediate input is observed in the data.

We assume that cultivated land is produced using land, tractor and harvesting equipment, and fuel and labour as inputs (i.e. the cost of cultivating a hectare of land is assumed to be independent of the amount of fertilizer applied and the crop yield harvested). The remaining non-nutrient cropping input (drying costs, sowing seed and pesticide costs) is assumed to be proportional to crop production (i.e. a limitation production relationship). Proportionality to crop output is also assumed for phosphorous and potash fertilizer since, under Danish farming conditions they are usually applied to insure an ample stock of these nutrients for growing crops (i.e. the nutrient restricting crop production is normally nitrogen). Thus, only nitrogen fertilizer input and land enter into a complex production relationship with crop output.

While retaining flexible estimation of behaviour with respect to key nitrogen flows, this simple specification makes it possible to estimate farm specific fertilizer demand elasticities and test for homogeneity.

We denote the dual profit function $B(p^x, p^n, z, \mathbf{2})$ where p^x is a price index of aggregate crop output, p^n is the price of nitrogen fertilizer and $\mathbf{2}$ is a vector of parameters. For farmer i , the complete system of profit and derived demand and supply functions becomes:

$$\begin{aligned} B_i' & B(p_i^x, p_i^n, z_i, \mathbf{2}_i) \\ x_i' & \frac{*B}{*p_x}(p_i^x, p_i^n, z_i, \mathbf{2}_i) \\ n_i' & \frac{*B}{*p_n}(p_i^x, p_i^n, z_i, \mathbf{2}_i) \end{aligned} \quad (1)$$

The shadow values or rent of the fixed cultivated land inputs r , can be derived from the profit function as:

$$r_i' = \frac{*B}{*z_i}(p_i^x, p_i^n, z_i, \mathbf{2}_i) \quad (2)$$

Regulatory efficiency

After estimating $\mathbf{2}_i$, the effect on farm i of introducing a nitrogen tax t can be simulated by inserting $p_i^n + t$ into (1). We denote initial farm profit and nitrogen fertilizer application $B_i^I, B(p_i^x, p_i^n, z_i, \mathbf{2}_i)$

and $n_i^I, \frac{*B}{*p_n}(p_i^x, p_i^n, z_i, 2_i)$ respectively and after tax profit and fertilizer application are denoted $B_i^T, B(p_i^x, p_i^n, z_i, 2_i)$ and $n_i^T, \frac{*B}{*p_n}(p_i^x, p_i^n, z_i, 2_i)$ respectively. Thus the change in abatement costs (i.e. profit reduction less tax payment) c_i^T and the change in nitrogen application n_i^T on farm i that is induced by a fertilizer tax become:

$$c_i^T, B_i^I \& B_i^T \& t(n_i^T) \quad (3)$$

$$n_i^T, n_i^I \& n_i^T \quad (4)$$

Exploiting the profit/production function duality virtual price result of Neary and Roberts (1980), the corresponding effect of a nitrogen quota Q_i , can be simulated by inserting a farm specific fertilizer tax q_i , which exactly ensures that nitrogen demand equals the quota (i.e. that

$Q_i, \frac{*B}{*p_n}(p_i^x, p_i^n, q_i, z_i, 2_i)$) while refunding tax revenue. Remembering that fertilizer input and quotas are negatively signed, farm profit under the quota becomes:

$$B_i^Q, B(p_i^x, p_i^n, q_i, z_i, 2_i) \& q_i(Q_i) \quad (5)$$

and the corresponding abatement cost and nitrogen application change on farm i become:

$$c_i^Q, B_i^I \& B_i^Q \quad (6)$$

$$n_i^Q, n_i^I \& Q_i \quad (7)$$

For a regulator wishing to implement an aggregate reduction goal N , it is well known that a fertilizer tax will ensure the distribution of reductions among farmers that minimizes aggregate abatement costs:

$$c_j^T \# c_j^Q \quad (8)$$

when $j \in \{1, \dots, N\}$, n_j^T, n_j^Q, N

The current Danish fertilizer quota system aims to reduce fertilizer application to 90% of

¹ We use fertilizer application as an indicator of nitrogen loss/leaching in the following. This is clearly not a reasonable effect indicator for livestock farms where manure application is important, but for crop farms without manure application it may be acceptable. Furthermore, the aggregation level of the model estimated here implies that the two measures, by definition, are proportional, making distinction redundant for the following analysis.

the profit maximizing level on crop farms (for farms with animal husbandry, the ambition is to increase manure utilization as well). This is done by calculating a profit maximizing baseline for each farm using self reported crop rotation and land allocation and mandating a fertilizer quota of 90% of this baseline. Quotas are controlled by random checks of farm accounts etc. If the regulator is able to generate exact estimates of each farm's baseline fertilizer application, and if fertilizer demand elasticities are homogeneous across farms, then such a quota system will induce abatement costs equal to those induced by a fertilizer tax (i.e. $t = q_1 = \dots = q_i = \dots = q_m$). If, on the other hand, these assumptions do not apply, proportional reduction quotas will be less efficient and may imply substantially higher abatement costs.

Distributional effects

Distributional effects may be an important dimension of instrument choice in general, and perhaps especially so when farmers are regulated. This is because regulation not only affects current farm profits, but we would also expect future profit losses to be capitalised in land values so that in addition current owners may suffer substantial capital losses.

Consider first the distributional effects of a fertilizer tax. The core effect is the reduction of annual farm profits ΔB_i^T i.e.:

$$\Delta B_i^T = B_i^I - B_i^T \quad (9)$$

Since land is an immobile factor of production, a reduction in farm profit will probably affect the value of farmland. If farms are typically sold as one unit (a going concern), a simplistic indicator of the effect on farm values is the capitalized value of the annual profit reduction " L_i^T " i.e.:

$$L_i^T = \Delta B_i^T / i \quad (10)$$

where i is the interest rate.

At present, however, the Danish agricultural sector is undergoing substantial concentration and farmers expanding land holdings have for some time been dominant on the demand side of the land market. If existing farmers engaging in marginal adjustments of land holdings dominate the market, a more relevant indicator of the effect on land values will be the effect on marginal land rent, since this indicates the effect of regulation on what existing farmers are willing to pay for additional land. We denote the initial rent of the marginal hectare of cultivated land $r_i^I = \frac{\partial B_i}{\partial z_i}(p_i^x, p_i^n, z_i, 2_i)$ and

the after tax rent $r_i^T = \frac{\partial B_i^T}{\partial z_i}(p_i^x, p_i^n, z_i, 2_i)$. Since capital and labour input to cultivated land are

mobile factors of production, it seems reasonable to assume that the entire rent change accrues to the immobile factor land. A simplistic indicator of the effect on land values in this case is the capitalized value of the change in marginal land rent ΔL_i^T i.e.:

$$\Delta L_i^T = (r_i^I - r_i^T) z_i / i \quad (11)$$

Though both indicators are crude, the land value effects they span may indicate the magnitude of the land value effects of regulation.

Now we consider the distributional effects of a quota. The core profit effect and capitalized value hereof are:

$$) B_i^Q, B_i^I \text{ \& } B_i^Q \quad (12)$$

and

$$" L_i^Q,) B_i^Q/i \quad (13)$$

Marginal land rent under a quota is the marginal effect on profit (defined in (5)) of increasing cultivated land input i.e.:

$$r_i^Q, \frac{dB_i^Q}{dz_i}, \frac{*B_i^Q}{*z_i} \% \frac{*B_i^Q * q_i}{*p_n * z_i} \text{ \& } \frac{*q_i}{*z_i} Q_i \text{ \& } q_i \frac{*Q_i}{*z_i} \quad (14)$$

By inserting the definition $Q_i = \frac{*B}{*p_n} (p_i^x, p_i^n, q_i, z_i, 2_i)$ the equation is reduced to:

$$r_i^Q, \frac{*B_i^Q}{*z_i} \text{ \& } q_i \frac{*Q_i}{*z_i} \quad (15)$$

Thus the capitalized value of this effect becomes :

$$) L_i^Q, (r_i^I \text{ \& } r_i^Q) z_i/i \quad (16)$$

3. The Data

Estimations are based on a panel data set provided by Landbrugets Rådgivningscenter (The Danish Agricultural Advisory Centre). The panel contains annual data, and is unbalanced covering ten growing seasons (1982 to 1991) with, on average, 1350 farms represented each year with each farm participating 3.9 years on average. Data are sampled from detailed gross margin accounts through a voluntary programme where only a small part of the more business oriented farmers participate. On the one hand, voluntary participation is an advantage in that participating farmers a priori are motivated and have an incentive to provide data of high quality. On the other hand, the sample of farmers in the data set is not representative of the population of Danish farmers.

For this analysis a group of specialised crop farms were selected. The criteria for selecting was that the farmer had no animal husbandry (cattle, pigs, chickens etc.) Specialised crop farms comprise about 15% of all farms in the data set. In the estimation we further restrict the sample to farms that have participated in the panel for more than three seasons.

For each farm the data include detailed annual accounts of variable costs along with corresponding accounts of quantitative flows of most nitrogen relevant inputs and outputs (i.e. fertilizer and crop yield). This allows an analysis of production and calculation of output and input prices at the farm level. Coefficients indicating average nitrogen content of crop outputs have been added enabling us to calculate annual farm level mass balances for nitrogen and residual nitrogen loss (see Hansen and Jensen (1998) for details). Mean values of the price index for each are reported in table 1.

Table 1 Means of farm specific price indexes for specialized crop farms

<i>Year</i>	<i>Crops</i>	<i>N-fertilizer</i>
1982	1.10	0.88
1983	1.22	1.12
1984	1.37	1.14
1985	1.17	1.26
1986	1.17	1.12
1987	1.10	0.83
1988	1.07	0.80
1989	1.10	0.89
1990	0.97	0.93
1991	0.93	1.07

Prices of nitrogen input were calculated directly for each farmer as cost divided by volume. Using a common base observation (containing all crop types), Fisher price indexes for crops were constructed based on the individual farmer's price. These were calculated as income (net of proportional input costs i.e. costs of drying, sowing seed, pesticides and phosphorous and potash fertilizer) divided by volume. Thus prices for nitrogen input and crop output vary across farms as well as over time. Both indexes exhibit substantial variation over the data period though no trend is apparent.

Mean values for key production variables and environmental indicators are shown in table 2. As noted above applied nitrogen fertilizer volume is registered in farm accounts while nitrogen loss is calculated using registered volumes and standard (average) nitrogen coefficients (again see Hansen and Jensen 1998 for details). Profit shares are defined as shares of gross profits before deduction of fixed costs and costs of cultivation. Thus this profit concept equals income from crop sales net of proportional input costs minus the cost of nitrogen fertilizer, so that the sum of the two profit shares (the input share being negatively signed) by definition equals one for each observation.

Table 2. Means of production and emission variables for specialized crop farms

<i>Profit share - Crops</i>	1.034
<i>Profit share - N-fertilizer</i>	-0.034
<i>Cultivated area</i>	188.9 hectares
<i>Applied nitrogen fertilizer</i>	136.2 kg/hectare
<i>Nitrogen loss</i>	62.6 kg/hectare
<i>Number of farms</i>	194
<i>Number of observations</i>	967

The participating crop farms are substantially larger than is typical for Danish crop farms, while fertilizer application and nitrogen loss per hectare correspond to that of typical Danish crop farms (see e.g. Brouwer et al. (1995) for nitrogen balances and a useful cross country comparison).

4. Estimation

The profit function in (1) is assumed to have the trans-log functional form with the following estimable specification:

$$\ln(B_{i,t}) = a_i + b_i^x \ln(p_{i,t}^x) + b_i^n \ln(p_{i,t}^n) + b_i^z \ln(z_{i,t}) + \frac{1}{2} \begin{bmatrix} \ln(p_{i,t}^x) & \ln(p_{i,t}^n) & \ln(z_{i,t}) \end{bmatrix} \begin{bmatrix} c_i^{x,x} & c_i^{x,n} & c_i^{x,z} \\ c_i^{n,x} & c_i^{n,n} & c_i^{n,z} \\ c_i^{z,x} & c_i^{z,n} & c_i^{z,z} \end{bmatrix} \begin{bmatrix} \ln(p_{i,t}^x) \\ \ln(p_{i,t}^n) \\ \ln(z_{i,t}) \end{bmatrix} \quad (17)$$

yielding derived profit share equations of the following form:

$$s_{i,t}^x = b_i^x + c_i^{x,x} \ln(p_{i,t}^x) + c_i^{x,n} \ln(p_{i,t}^n) + c_i^{x,z} \ln(z_{i,t}) \quad (18)$$

$$s_{i,t}^n = b_i^n + c_i^{n,x} \ln(p_{i,t}^x) + c_i^{n,n} \ln(p_{i,t}^n) + c_i^{n,z} \ln(z_{i,t}) \quad (19)$$

where i is a farm index, t indicates the time period and $s_{i,t}^x = \frac{p_{i,t}^x x_{i,t}}{B_{i,t}}$ and $s_{i,t}^n = \frac{p_{i,t}^n n_{i,t}}{B_{i,t}}$ are the

profit shares of crop output and nitrogen fertilizer input. Note that a_i , b_i^x , and b_i^n are farm specific parameters allowing fixed effects in each budget share equation as well as in the profit equation.

Furthermore $c_i^{x,x}$, $c_i^{x,n}$, $c_i^{n,n}$ and $c_i^{n,x}$ are permitted to vary across farms in a structured way (specified below) that allows for homogeneity as well as varying degrees of heterogeneous price elasticities across farms.

The complete system is estimated in two steps. First, the system of derived profit share equations (without the profit function) is estimated using maximum likelihood estimation. Restricting the parameters to ensure symmetry and homogeneity in prices (but not in the fixed land input (i.e. $b_i^x \% b_i^n$, 1 , $c_i^{x,z}$, $c_i^{n,z}$ and $c_i^{n,n}$, $c_i^{n,x}$, $c_i^{x,n}$, $c_i^{x,x}$ for all i) we eliminate the crop equation ($s_{i,t}^x$) to avoid singularity (the maximum likelihood procedure ensures that estimates are invariant as to which equation is eliminated). Technically, the c parameters in (19) are estimated using within farm transformed variables, eliminating time invariant farm specific constant b_i^n and the homogeneity/symmetry restrictions, i.e.:

$$\tilde{s}_{i,t}^n = c^{n,n}(\tilde{l}\tilde{n}(p_{i,t}^n) \& \tilde{l}\tilde{n}(p_{i,t}^x)) \% c^{n,z}\tilde{l}\tilde{n}(z_{i,t}) \quad (20)$$

where $\tilde{\cdot}$ indicates within transformed variables. The fixed effects b_i^n are then estimated as $\hat{b}_i^n = \bar{s}_{i,t} \& \hat{c}_i^{n,n}(\bar{l}\tilde{n}(p_{i,t}^n) \& \bar{l}\tilde{n}(p_{i,t}^x)) \& \hat{c}_i^{n,z}\bar{l}\tilde{n}(z_{i,t})$ for each i with $\bar{\cdot}$ indicating the mean value of the variable taken over the time periods that farm i participates in the panel.

The degrees of freedom allowed by the data makes estimation of unrestricted farm specific $c_i^{n,n}$ coefficients infeasible. Instead we generalize the standard trans-log model by letting $c_i^{n,n}$ depend on farm specific mean input prices using the following quadratic specification:

$$c_i^{n,n} = \alpha_1 \% \alpha_2 \bar{s}_i^n \% \alpha_3 (\bar{s}_i^n)^2 \quad (21)$$

The quadratic functional form is a flexible generalisation of the usual uniform coefficient assumption, which is attractive for our purposes since, when inserted into the trans-log own price elasticity

formula $e_i^n = \bar{s}_i^n \% \frac{c_i^{n,n}}{\bar{s}_i^n} \& 1$ the mean price elasticity becomes:

$$e_i^n = \bar{s}_i^n \% \frac{(\alpha_1 \% \alpha_2 \bar{s}_i^n \% \alpha_3 (\bar{s}_i^n)^2)}{\bar{s}_i^n} \& 1$$

$$\bar{s}_i^n \% \frac{1}{\bar{s}_i^n} \% \alpha_2 \% \alpha_3 \bar{s}_i^n \& 1 \quad (22)$$

This is nice because, if $\alpha_1 = 0$ and $\alpha_3 = -1$ then e_i^n becomes constant (homogeneous) across farms

which makes it possible to test the hypotheses of elasticity homogeneity. We can also test for applicability of the usual simple trans-log specification (i.e. restrictions $\alpha_2 = 0$ and $\alpha_3 = 0$).

After inserting (21) into (20) and regrouping we have the following equation to be estimated:

$$\tilde{s}_{i,t}^n = \alpha_1 ([\ln(p_{i,t}^n) \ \& \ \ln(p_{i,t}^x)] \ % \ \alpha_2 ([\bar{s}_i^n (\ln(p_{i,t}^n) \ \& \ \ln(p_{i,t}^x))] \ \% \ \alpha_3 [(\bar{s}_i^n)^2 (\ln(p_{i,t}^n) \ \& \ \ln(p_{i,t}^x))] \ \% \ c^{n,z} [\ln(z_{i,t})] \ \% \ u_{i,t} \quad (23)$$

where square parentheses indicate the independent variables used in estimation and $u_{i,t}$ is an error term. The error term is assumed to be normally distributed and equation (23) was estimated with SAS PROC MODEL using full information maximum likelihood. Results of this estimation are reported in table 3 (parameters for the crop equation were calculated residually).

Table 3 Common parameters estimated in the first step

Parameter	Estimated general model			Test of restrictions (likelihood ratio)	
	Estimate	Approx Std Err	Approx Prob	Homogeneous Elasticities	Standard Trans-Log
α_1	-0.0051	0.0085	0.5299	0	
α_2	0.5714	0.3062	0.0624		0
α_3	5.0789	2.2783	0.0260	-1	0
$c^{n,z}$	0.0072	0.0052	0.1705		
	DF	Model sig.	R ²	P ² (2)=17.81 (0.0001)	P ² (2)= 4.96 (0.0837)
	769	P ² (4)=34.30 (0.0000)	0.0348		

Table 4 contains the mean, median and standard deviation of the distribution of residually calculated farm specific fixed effects.

Table 4 Farm specific fixed effects estimated in the first step

Parameter	mean of estimates	median of estimates	std. dev. of estimates
b_i^n	-0.0733	-0.0709	0.0195
b_i^{x*}	1.0733	1.0709	0.0195

In the second step the remaining parameters of the profit function were estimated treating fixed effects and parameters estimated in the derived system as known, i.e. :

$$\ln(B_{i,t}) \& b_i^x \ln(p_{i,t}^x) \& b_i^n \ln(p_{i,t}^n) \& \ln(p_{i,t}^x) c_i^{z,x} \ln(z_{i,t}) \& \ln(p_{i,t}^n) c_i^{z,n} \ln(z_{i,t})$$

$$\frac{1}{2} \begin{bmatrix} \ln(p_{i,t}^x) & \ln(p_{i,t}^n) \end{bmatrix} \begin{bmatrix} c_i^{x,x} & c_i^{x,n} \\ c_i^{n,x} & c_i^{n,n} \end{bmatrix} \begin{bmatrix} \ln(p_{i,t}^x) \\ \ln(p_{i,t}^n) \end{bmatrix} \quad (24)$$

$$, \quad a_i \ \% \ b^z \ln(z_{i,t}) \ \% \ \ln(z_{i,t}) c^{z,z} \ln(z_{i,t})$$

If the parameters estimated in the first step are unbiased, this will also be the case for parameters estimated in the second step. However, the procedure may generate heteroscedasticity and invalidate the usual inference statistics. Results of the second step (also estimated with SAS PROC MODEL using full information maximum likelihood) are presented in table 5.

Table 5 Parameters estimated in the second step

<i>Parameter</i>	<i>Estimate</i>	<i>Approx Std Err*</i>	<i>Approx Prob*</i>
b^z	1.1028	0.0756	0.0001
$c^{z,z}$	1.0471	0.3119	0.0008
	<i>mean of Estimates</i>	<i>median of Estimates</i>	<i>std. dev. of Estimates</i>
a_i	-4.9924	-4.6416	4.2655
	DF	Model sig.	R ²
	771	P ² (2)=137.52 (0.0000)	0.1326

Note: * inference statistics are conditional on known first step parameters, i.e. a lower bound for true standard errors.

The first step estimation is highly significant though most individual parameters are not. Estimated share equations are consistent with monotonicity and convexity. Inspection of error correlation matrices did not reveal serious serial correlation. Statistical tests showed that the error distribution is significantly different from the normal distribution though residual plots indicated that deviation is not substantial. Residual plots showed clear signs of heteroscedastic error terms. Though non-normality and heteroscedasticity may invalidate inference tests, parameter estimates are still unbiased. The second step estimation is also highly significant and error correlation matrices and residual plots did

not indicate serious serial correlation or heteroscedasticity, but significant non-normality was detected.

Estimated price elasticities (see appendix) have the expected signs in over 95% of all single observations. All single observations of rents of cultivated land have the expected sign. In conclusion the estimated model performs well statistically and generates plausible behavioural inferences.

Accepting the general model the restriction tests reported in table 3 can be interpreted. Though qualified by non-normality and heteroscedasticity the clear rejection of the homogeneity hypothesis indicates that nitrogen demand elasticities are heterogeneous across panel farms. This is confirmed by the bootstrap confidence interval around the calculated standard deviation of farm elasticities from the panel mean that is reported below. The restriction resulting in the simple standard trans-log specification can, on the other hand, only be rejected at the 10% level. In conclusion elasticities are heterogeneous across the panel, but the general model specification only results in a slightly better description of this heterogeneity than the standard trans-log specification. In the following section we base results on the general model specification.

5. Results

In table 6 we report mean elasticities and standard deviations of individual farm elasticities from these means (as a measure of the heterogeneity of elasticities in the sample) using the parameters reported in the previous section. Mean and standard deviations of farm elasticities are calculated using one observation for each farm evaluated at mean farm prices.²

Table 6 Own price elasticity of nitrogen fertilizer

	<i>Point estimate</i>	<i>90% confidence inter-</i>
<i>val*</i>		
<i>Mean of panel farm elasticities</i>	-0.45	[-0.58 : -0.31]
<i>Standard deviation of farm elasticities from panel mean</i>	0.24	[0.01 : 0.41]

Note: One elasticity observation per farm evaluated at mean farm prices.

* Calculated by bootstrapping based on 1250 data re-samplings.

The mean own price elasticity for nitrogen fertilizer of -0.45 reported here has a short run Marshallian interpretation. Given this, it is in line with or somewhat larger numerically than the elasticities found in several recent studies (all based on aggregate time series covering all or most of the

² Note that the reported measures are almost identical to the corresponding measures after price deflation to the mean panel price level (see appendix). Deflation gives consistent estimates of aggregate behaviour, however, deflation may widen confidence intervals around the calculated individual farm elasticities. In our case the discussion turns out to be academic since the resulting elasticities are almost identical.

agricultural sector), e.g. Burrell (1989) finds a short run elasticity of between -0.4 and -0.6, and Rayner and Cooper (1994) find an elasticity of between -0.1 (short run) and -0.25 (long run), for the UK, while Denbaly and Vroomen (1993) report between -0.2 (short run) and -0.4 (long run) for the USA, all with Marshallian interpretations as ours.

Three studies of fertilizer demand by Danish agriculture are of special interest. In an older study using aggregate time series covering the entire Danish agricultural sector, Dubgaard (1987) estimates a Marshallian nitrogen fertilizer own price elasticity of -0.19. Jensen (1996), using an estimated aggregate model of Danish agriculture, finds a Marshallian long run elasticity for composite NPK-fertilizer of -1.8 and Kristensen and Jensen (1999), using cost functions estimated on panel data, find a short run Hicksian elasticity for composite NPK-fertilizer of -0.52 for Danish crop farms. Our finding is somewhat larger numerically than Dubgaard's result but smaller than the comparable elasticities implied by the two more recent studies.³

The main contribution of this paper is the quantification of the heterogeneity of elasticities across farms. As a measure of this, the standard deviation of individual farm elasticities from the mean of panel elasticities is found to be 0.24 as indicated in table 6. The bootstrap confidence interval indicates that this is significantly different from zero at the 5% level, supporting the clear rejection of the homogeneous elasticity hypotheses in the previous section. Thus, the standard deviation is significant and sizable: with 50% of the panel farms having elasticities outside the -0.33 to -0.57 range around the mean elasticity.

If all farms had had the same fertilizer demand elasticity quotas, requiring uniform percentage reductions in fertilizer application would result in an optimal allocation of cutbacks across farms. When elasticities are heterogeneous, however, uniform percentage reductions become inefficient, inducing aggregate abatement costs that are larger than the minimum level that would be ensured by tax regulation. This is the classical efficiency argument for preferring economic incentive regulation to standards and norms.

In practical applications, a regulator using uniform percentage reduction quotas also faces the problem of estimating fertilizer application baselines for each farm. This is not a trivial task since baseline applications depend on a number of farm specific factors which are not observed perfectly by the regulator (farm land quality, local weather conditions, capacity and quality of fixed capital inputs, entrepreneurial talent etc.). Thus baselines will generally be estimated with some degree of error, implying that the applied individual farm quotas will deviate in some random way from the uniform percentage reduction goal. In addition, if the farm specific information used for calculating baselines depends on farm production decisions, incentives that distort these decisions may be generated.

Using our model of panel farms we are able to quantify the abatement cost increase of reduction quotas caused by elasticity heterogeneity and the additional abatement cost increase that would result from various magnitudes of random errors in baseline measurement. Because of the

³ When consistently estimated results for both composite fertilizer and nitrogen are reported (see Burrell, 1989 and some of the older studies summarised in Burrell's paper) the nitrogen elasticity tends to be the same size or numerically larger. Thus the Hicksian short run composite elasticity reported by Kristensen and Jensen (1999) indicates a short run Marshallian nitrogen elasticity numerically larger than -0.52 while the composite Marshallian long run elasticity reported by Jensen (1996) probably implies a comparable elasticity substantially larger than the one found here.

aggregation level of the estimated model, the effect of distorting incentives generated by the baseline calculation system cannot be evaluated.

Results are presented in the first row of table 7 where abatement costs (DKK/kilo reduction in nitrogen application) are indicated for a 10% reduction in applications. The fertilizer tax rate t is found by iteration and abatement costs calculated using equation (3) parameterised with the estimated model. For the quota system the baseline for each farm is calculated as the farm's optimal unregulated fertilizer application plus a random error term drawn from a normal distribution with zero mean and a standard deviation of the indicated percentage of the farm's optimal application level. Thus the abatement cost difference between the tax and the quota with 0% measurement error can be interpreted as the isolated cost effect of the estimated panel elasticity heterogeneity, whereas differences to the remaining columns include the added cost effect of the indicated baseline measurement error.⁴ In the final column the abatement cost of a corresponding tradable quota system is presented. These are calculated under the assumption that the quota market is perfect (i.e. resulting in a post trading allocation of quotas equal to the applications induced by a fertilizer tax), so that abatement costs are independent of the initial quota distribution.

By design, the fertilizer tax generates incentives that minimize abatement costs. We see that the abatement cost increase under the non-tradable quota scheme induced by the sizable elasticity heterogeneity estimated here is relatively small, amounting to a cost increase of less than 8%. This is about the magnitude of the abatement cost increase caused by a 2.5% baseline measurement error, whereas larger measurement errors cause substantially larger cost increases. Note that the table does not include distorting effects or administrative costs of the non-tradable quota systems baseline estimation procedure. If these are important, the efficiency advantage of using taxes will be greater than indicated in the table. Likewise the table does not include added administrative costs or incentive distortions of the tradable quota systems grand fathering procedure, which may cause such a system to be less efficient than tax regulation.

⁴ Specifically farm quotas Q_i are set to a uniform percentage of each farm's estimated baseline application (including measurement error). Then farm specific tax rates q_i which implement these are found by iteration for each farm after which abatement costs are calculated using equation (6). The uniform percentage of baselines used to calculate farm specific quotas that ensure an aggregate 10% application reduction are found by iteration in an outer loop. Without measurement error, a uniform percentage of 90 ensures a 10% reduction. With a normally distributed measurement error, some 90% quotas may exceed actual optimal farm application so that the uniform percentage must be increased in order to avoid exceeding the 10% reduction goal.

Table 7 Abatement costs and agricultural income effects when aggregate N-fertilizer application is reduced by 10% (DKK per kilo N-fertilizer reduction)

	Tax	Non-tradable Quotas* proportional percent reduction in application with baseline measurement error of					Grand-fathered Tradable quotas*
		0 %	2½ %	5 %	7½ %	10 %	
Average abatement costs	0.376	0.406	0.426	0.483	0.566	0.668	0.376
Average farm profit	7.377	0.406	0.426	0.483	0.566	0.668	0.376
Average farm capital loss	73.8 - 490.5	4.1 - 25.5	4.3 - 26.8	4.8 - 30.5	5.7 - 35.7	6.7 - 42.2	3.8 - 24.2

Note: Measurement errors are normally distributed with standard deviation of the indicated per cent of baseline. Calculations are at 1989 fertilizer and crop prices. The fertilizer tax rate is 0.776 DKK/kilo N-fertilizer. The applied quota percentages of baseline application are 90.0% for tradable quotas and non-tradable quotas with a 0% error, 90.2% for a 2½% error, 90.5% for a 5% error, 90.9% for a 7½% error, and 91.5% for a 10% error. The interest rate used in the calculation of capital loss is 10%.

* Excluding the possibly distorting effect on farm production decisions of the baseline estimation procedure/initial quota distribution procedure.

In rows two and three the policy effect on annual farm profit and land values are presented (per kilo fertilizer reduction for cross row comparability). The effect on annual profit is calculated using equations (9) and (12) respectively. The lower bound on the land value effects is calculated using equations (10) and (13), and equations (11) and (16) are used to find the upper bound (the trans-log specification of (10),(11) and (13) is straight forward, see appendix for specification of (16)).

Closely matching the revenue raised by a fertilizer tax, the reduction it causes in farm income is between 11 and 18 times the income reduction caused by a quota system. The corresponding land value reductions are between 11 and 19 times as large. The substantial differences in distributional effects are not surprising given the inelasticity of nitrogen demand and the relatively small abatement cost advantage of the fertilizer tax. The distributional effects of the tradable quota system are smaller than for the non-tradable system, reflecting the efficiency advantage.⁵ Again table 7 assumes that the baseline estimation procedure/quota distribution procedure has negligible distortionary effects.

6. Conclusions

Using farm level panel data and a generalized trans-log profit function specification with a number of farm specific parameters we find a short run Marshallian price elasticity for nitrogen fertilizer demand by Danish crop farms of -0.45. This is within the span of elasticities found internationally and in other

⁵ Since baseline measurement error only affects the initial distribution of quotas in the tradable quota system, this does not affect the efficiency or aggregate farm income presented here. However measurement error may cause substantial income redistribution among individual farmers.

Danish studies.

The main contribution of this paper is the quantification of the heterogeneity of elasticities across farms. As a measure of this, the standard deviation of individual farm elasticities from the mean of panel farm elasticities is found to be 0.24. This indicates sizable heterogeneity with 50% of the panel farms having elasticities outside the -0.33 to -0.57 range around the mean elasticity.

The classical economic argument for incentive regulation is that heterogeneity of elasticities causes abatement costs of uniform percentage reduction quotas to increase above the minimum level ensured by tax regulation. However, the abatement cost increase induced by the (sizable) estimated heterogeneity amounts to less than 8% of the abatement cost minimum that is ensured if reductions are induced by a fertilizer tax. This result is somewhat surprising, suggesting that the classical argument for incentive regulation may be relatively weak even in regulatory situations with substantial polluter heterogeneity. Our simulation results indicate that inaccurate estimation of baselines may be a more important threat to the efficiency of uniform reduction schemes than heterogeneity of elasticities.

The substantial agricultural income and land value reductions associated with a fertilizer tax can be avoided without efficiency loss through a system of grandfathered tradable quotas (assuming that a non-distorting quota distribution procedure is utilized). If policymakers, for ethical or other reasons, forbid trading, our results indicate that the efficiency loss due to heterogeneity is limited. There may still, however, be a substantial efficiency loss due to baseline measurement error and to distorting incentives generated by the baseline estimation procedure. At any rate, our results indicate that careful designing of baseline estimation procedures should be given a high priority whenever uniform percentage reduction schemes are used.

It is important to stress that the specific implications regarding the Danish nitrogen quota system drawn here only cover its application to crop farms. Interpretation as an evaluation of the system in general would be misleading since relative performance of different regulatory instruments on farms with livestock may be very different. First of all, because behavioural relations for pig and dairy farms may differ substantially and second, because the Danish quota system has special rules for livestock farms reducing quotas by the calculated fertilizer value of manure.

Appendix

Price elasticities derived from the general model

Own and cross price elasticities and land rent are calculated for expected input/output profit shares for each data observation and for each farm at mean farm prices both for non-deflated prices and after deflating to the mean panel price level. The mean, median and standard deviation of the distribution of individual observation and farm elasticities/rents and the proportion of positive elasticities/rents are reported for all input/output price combinations.

<i>Input/ output</i>	<i>Individual observations</i> <i>non-deflated prices</i> <i>(967 observations)</i>	<i>Farm observations</i> <i>non-deflated mean farm prices</i> <i>(194 observations)</i>
	Prices	Prices
<i>x</i>	<i>p^x</i> <i>pⁿ</i>	<i>p^x</i> <i>pⁿ</i>
<i>Mean</i>	0.019 -0.019	0.019 -0.019
<i>median</i>	0.014 -0.014	0.014 -0.014
<i>Std Dev</i>	0.023 0.023	0.025 0.025
<i>%>0</i>	95% 5%	97% 3%
 <i>n</i>		
<i>Mean</i>	0.442 -0.442	0.447 -0.447
<i>median</i>	0.469 -0.469	0.457 -0.457
<i>Std Dev.</i>	0.285 0.285	0.238 0.238
<i>%>0</i>	95% 5%	97% 3%
	Land rent	Land rent
	DKK/hectare	DKK/hectare
<i>Mean</i>	41285.21	38412.67
<i>Median</i>	41040.62	38218.60
<i>Std Dev.</i>	13154.98	11068.25
<i>%>0</i>	100%	100%

<i>Input/ output</i>	<i>Individual observations deflated prices</i> (967 observations)	<i>Farm observations deflated mean farm prices</i> (194 observations)
	Prices	
<i>x</i>	p ^x p ⁿ	p ^x p ⁿ
<i>Mean</i>	0.019 -0.019	0.020 -0.020
<i>median</i>	0.014 -0.014	0.014 -0.014
<i>Std Dev</i>	0.023 0.023	0.025 0.025
<i>%>0</i>	96% 4%	97% 3%
 <i>n</i>		
<i>Mean</i>	0.443 -0.443	0.459 -0.459
<i>median</i>	0.467 -0.467	0.464 -0.464
<i>Std Dev.</i>	0.286 0.286	0.232 0.232
<i>%>0</i>	96% 4%	97% 3%
	Land rent	
	DKK/hectare	
<i>Mean</i>	39183.77	39324.94
<i>Median</i>	39027.06	39290.13
<i>Std Dev.</i>	11624.48	11109.20
<i>%>0</i>	100%	100%

Derivation of the quota effect on marginal land rent

The general specification of marginal land rent under quotas from equation (15) is:

$$r_i^Q, \frac{*B_i^Q}{*z_i} \& q_i \frac{*Q_i}{*z_i} \quad (\text{a.1})$$

where the trans-log specification of $\frac{*B_i^Q}{*z_i}$, $\frac{*B_i^Q}{*z_i} \left(b^z \% c^{z,x} \ln(p_{i,t}^x) \% c^{z,n} \ln(p_{i,t}^n) \% c^{z,z} \ln(z_{i,t}) \right)$ is straightforward. The trans-log specification of Q_i (i.e. the fertilizer demand induced by the tax adjusted price $(p_i^n \% a_i)$) is $Q_i' \frac{*B_i^Q s_i^n}{(p_i^n \% a_i)}$ so that:

$$\begin{aligned} \frac{*Q_i}{*z_i}, & \left(\frac{*B_i^Q}{*z_i} s_i^n \% B_i^Q \frac{*s_i^n}{*z_i} \right) / (p_i^n \% a_i) \\ & , \\ & \left(\frac{*B_i^Q}{*z_i} s_i^n \% B_i^Q \frac{*s_i^n}{*\ln(z_i)} \frac{*\ln(z_i)}{*z_i} \right) / (p_i^n \% a_i) \\ & , \\ & \left(\frac{*B_i^Q}{*z_i} s_i^n \% \frac{B_i^Q c^{n,z}}{z_i} \right) / (p_i^n \% a_i) \end{aligned}$$

which, after inserting into (a.1), gives the following trans-log specification of marginal land rent:

$$\begin{aligned} r_i^Q, & \frac{*B_i^Q}{z_i} \left(b^z \% c^{z,x} \ln(p_{i,t}^x) \% c^{z,n} \ln(p_{i,t}^n) \% c^{z,z} \ln(z_{i,t}) \right) \& \\ & \frac{q_i}{p_i^n \% a_i} \left(\frac{*B_i^Q}{*z_i} s_i^n \% \frac{B_i^Q c^{n,z}}{z_i} \right) \end{aligned} \quad (\text{a.2})$$

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