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Technology and efficiency in a panel of Italian dairy farms: an SGM restricted cost function approach

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Abstract

This article employs a short-term specification of the symmetric generalised McFadden (SGM) cost function capable of accommodating quasi-fixed factors and variable returns. Temporary equilibrium and scale economies are investigated while maintaining the consistency of the estimated model with microeconomic theory and approximation properties. It also makes use of a two-step procedure to estimate first the technology parameters and then time-varying efficiency at farm level. No distributional assumptions are required on efficiency as we consider a fixed effect model. A balanced panel of Italian dairy farms during the years from 1980 to 1992 serves as the case study. The results suggest a rigid productive structure during the pre- and post-quota period. Moreover, Italian milk producers are found to exhibit considerable excess capacity and rather low input technical efficiency.

JEL classification: Q12; C33

Keywords: Restricted cost function; SGM; Capacity utilisation; Input technical efficiency; Panel data

1. Introduction and background

In applied economics, aggregate output is commonly related to a list of inputs through a production function and, when dealing with time series data, to a proxy for technology which is often represented by a linear trend. This framework does not recognise the short-run fixity of some factors that in agriculture, as virtually in any sector, represents a structural constraint. In spite of its general acceptance, this notion has not found an adequate representation at the empirical level because of both data and model inadequacy.

Our study is primarily concerned with contributing towards this strand of literature and investigating the productive behaviour of the dairy sector in Italy. We assume the existence of a short-run aggregate technology and depict it from the dual by means of the symmetric generalised McFadden (SGM) cost function. This flexible form has been proposed by Diewert and Wales (1987) and subsequently adapted in one way or another (Kumbhakar, 1990, 1994; Peeters and Surry, 2000). To our knowledge, only few authors investigate short-run behaviour within the SGM framework. Rask (1995), e.g. estimates a modified SGM function for Brazilian sugarcane production. His model, though, is not fully quadratic and places unnecessary restrictions on the underlying cost structure. This paper builds upon Kumbhakar et al.'s (1989) work in specifying a restricted cost model that accommodates quasi-fixed

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inputs and maintains the consistency of the estimated function with microeconomic theory and approximation properties. As case study we use a balanced panel of dairy farms in the plain of the Po river. This area accounts for more than 60% of Italian milk supply. The analysis covers the years from 1980 to 1992. The productive technology consists of one aggregate output, three variable inputs (purchased feed, other intermediate inputs and hired labour), two quasi-fixed factors (family labour and capital).

A second major purpose of this study is to examine whether the introduction of milk quotas in 1984 had any impact on farmer decisions. Few subjects have drawn more attention in European agricultural policy than the organisation of the dairy sector in the EU (Burrell, 1989; Boussard, 1985; Petit et al., 1987, among others). However, no studies have examined how Italian producers initially reacted to this policy. The history of the Italian milk quotas is one of long delays in applying Community legislation, huge fines, deceitful activities and protests. Italy was granted a quota of 9.9 million tonnes (mt) on the basis of national statistics. In 1984 the Italian Minister of Agriculture argued that there was a discrepancy of 1.5 mt between this figure and actual production. Disagreements over the actual production level and number of producers were to continue for many years. As Italy imports 40% of its milk requirement, it was decided to treat the whole country as a single entity for 2 years and not to allocate individual quotas in order to exploit fully the national allotment. In 1988, again in order to ensure full use of it, it was decided to allocate the quota to producer associations each of which would act as a single producer, and Unalat was created for this purpose. However, Unalat decided to apply the legislation on a voluntary basis. Failure to apply the system meant that Italy ran up a fine in the order of 300 billion lira each year. By 1992 the fine had reached 4000 billion lira. Italy maintained that the quota was inadequate and requested a backdated increase. In 1993 a first attempt was made to collect data to establish quotas. Individual quotas were published, but their sum exceeded the national ceiling. It was then decided to rely on a system whereby farmers provide their own data on production. This system was an invitation to irregular practices (Borroni et al., 2001; Pieri and Rama, 1996; Senior, 2002).

How did dairy enterprises respond to the contradictory announcements of Italian institutions? We argue that the adjustment process that took place is not trivial and attempt to answer this question by estimating the productive efficiency of dairy farms individually, testing for the presence of structural breaks and analysing the trends of capacity utilisation (CU).

The paper proceeds by presenting the analytical model and the two-step procedure used to estimate first the parameters of the cost function and then farm level efficiency. Section 3 provides a brief discussion of our balanced panel and variable construction. In Section 4, production elasticities and input technical efficiency are discussed. Given the short-term standpoint, cost flexibility is decomposed into scale economies and capacity utilisation. The effect of technological bias on input use is also investigated. In the second step, no distributional assumptions are required on technical efficiency as we consider a fixed effect model. Since our analysis covers 13 years, individual effects are allowed to vary according to a flexible second-degree polynomial. Section 5 concludes.

2. Methodological framework

2.1. The SGM restricted cost function

In this study we maintain that the objective of the farmers in the sample is to minimise the cost of producing a given level of output, conditional on input prices, stocks of quasi-fixed inputs and technology. Under some regularity conditions, duality principles ensure consistency between variable cost and production functions, so that either one will describe the farming activity equally well (Chambers, 1988). The restricted cost function is given by

$$G = G^*(Y, W, Z, t) \tag{1}$$

where G is variable cost, Y the output, $W \equiv (W_1, W_2, \dots, W_N)'$ the vector of input prices, $Z \equiv (Z_1, Z_2, \dots, Z_M)'$ the vector of fixed inputs and t is the time trend used as proxy for technology.[1]

[1] The cost function is linearly homogeneous, non-decreasing and concave in W, non-decreasing in Y, non-increasing and convex in Z, non-negative, continuos and twice continuously differentiable in all its arguments.

Empirically, we depict G^* by means of the SGM form because it is flexible, in the sense of providing a second-order approximation to an unknown function at any given point (Diewert, 1976); it has a Hessian of constants, thus the curvature properties hold globally and can be tested and possibly imposed without destroying flexibility; finally, it is invariant to normalisation. In this study, we depart from Diewert and Wales (1987) by adding quasi-fixed inputs (Kumbhakar, 1989). The model estimated is

$$\begin{aligned} G = {} & g(W)Y + \sum_i^N b_{ii} W_i Y + \sum_i^N b_i W_i \\ & + \sum_i^N b_{it} W_i t Y + b_t(\alpha' W)t + b_{YY}(\beta' W)Y^2 \\ & + b_{tt}(\gamma' W)t^2 Y + \sum_i^N \sum_k^M d_{ik} W_i Z_k \\ & + \sum_k^M c_{kY}(\delta' W) Z_k Y + \sum_k^M c_{kt}(\lambda' W) Z_k t \\ & + 0.5 \sum_j^M \sum_k^M c_{jk}(\eta' W) \frac{Z_j Z_k}{Y} \end{aligned} \tag{2}$$

where $g(W)$ is defined by

$$g(W) = \frac{W'SW}{2\theta' W} = \frac{\sum_i^N \sum_h^N s_{ih} W_i W_h}{2\sum_i^N \theta_i W_i} \tag{3}$$

and S is a $N \times N$ symmetric negative semidefinite (nsd) matrix such that $S'W^* = 0$ with $W^* \gg 0$, and i, h denote variable inputs and j, k fixed inputs. Since W^* is chosen to be the vector of ones, $\sum_h s_{ih} = 0$ for all i, and the rank of S is $(N - 1)$. $\theta = (\theta_1, \ldots, \theta_N)'$ is a vector of non-negative constants not all zero and $c_{jk} = c_{kj}$.

It can be shown that G is a flexible (linearly homogeneous in W) restricted cost function at any point (Y^*, W^*, Z^*, t^*) provided that $W^* \gg 0$, $\theta' W^* > 0$, $\alpha' W^* \neq 0$, $\beta' W^* \neq 0$, $\gamma' W^* \neq 0$, $\delta' W^* \neq 0$, $\lambda' W^* \neq 0$, $\eta' W^* \neq 0$. Moreover, G is globally concave in W if $S = \{s_{ih}\}$ is negative semidefinite and globally convex in Z if $C = \{c_{jk}\}$ is positive semidefinite and $\eta' W^* > 0$.[2] For the SGM cost function to be parsimonious, i.e. provide the second-order approximation using a minimal number of parameters, the vector θ, along with α, β, γ, δ, λ and η, need to be exogenously given.[3] Thus, there are $(N + M)(N + M + 1)/2 + 2(N + M) + 3$ free parameters to be estimated—just enough for the SGM variable cost function to be flexible at the point (Y^*, W^*, Z^*, t^*).

If the estimated S matrix does not conform to concavity criteria, negative semidefiniteness can be imposed[4] by reparameterising it as $S = -TT'$, where T is a lower triangular matrix. Global convexity in quasi-fixed inputs Z can be stated (imposed) analogously upon the positive semidefiniteness of the estimated matrix C.

For econometric implementation, a set of cost-minimising variable input demands can be derived using Shephard's lemma. Here, optimal input–output coefficients ($X_i/Y = (\partial G/\partial W_i)/Y$) are considered to reduce possible heteroskedasticity:

$$\begin{aligned} \frac{X_i}{Y} = {} & \left\{ \frac{S^{(i)} W}{\theta' W} - \frac{\theta_i}{2} \frac{W'SW}{(\theta' W)^2} \right\} + b_{ii} + \frac{b_i}{Y} \\ & + b_{it} t + \frac{\alpha_i b_t t}{Y} + \beta_i b_{YY} Y + \gamma_i b_{tt} t^2 + \sum_k^M d_{ik} \frac{Z_k}{Y} \\ & + \delta_i \sum_k^M c_{kY} Z_k + \lambda_i \sum_k^M c_{kt} \frac{Z_k t}{Y} \\ & + \frac{\eta_i}{2} \sum_j^M \sum_k^M c_{jk} \frac{Z_j Z_k}{Y^2} \end{aligned} \tag{4}$$

Note that the system (4) contains all the relevant parameters; hence, we need not provide an enlarged set of equations. However, greater efficiency in estimation can be gained by forcing more structure on the data,

[2] The statements on flexibility and concavity follow theorems 10 and 11 in Diewert and Wales (1987).

[3] The inner products can be seen as fixed-weight price indexes. We assume that they have the Laspeyres form with weights given by the mean quantities (Diewert and Wales, 1987; Kohli, 1993) as well as $\theta = \alpha = \beta = \gamma = \delta = \lambda = \eta$. In this case, $\theta' W^* > 0$ and $\theta > 0$, and similarly for the remaining indexes. For the flexibility proof, see Appendix A in Kumbhakar (1989).

[4] The imposition of required curvature at each data point does not destroy flexibility. However, by reducing the rank of the reparameterised Hessian we hamper the range of second-order effects and move to a semiflexible version (Diewert and Wales, 1987; Moschini, 1998; Ryan and Wales, 1998). Empirically, the rank reduction is equal to the number of the Hessian eigenvalues with wrong signs.

e.g. including additional information in the form of shadow value equations. Such equations represent the potential reduction in variable cost from an additional unit of quasi-fixed input ($-\partial G/\partial Z_k = F_k$).[5] Variable returns to scale prevent us from equating the residual measure of returns to multiple quasi-fixed inputs, $PY - G$, where P is output price, with the shadow fixed cost, $\sum F_k Z_k$ (Morrison, 1988). So, for estimation purposes one either assumes that shadow prices are proportional to ex ante user costs or merely omits them. On theoretical as well as empirical grounds, we opt for the second alternative.

2.2. Elasticities and capacity utilisation

The proposed model ascribes a central role to relative prices: in the short run, they determine the demand for variable inputs and, via shadow prices, contribute towards the explanation of capacity utilisation; in the long run, they determine the optimal levels of quasi-fixed factors. These effects can be measured using the conventional elasticity coefficients.

Short-run price elasticities are calculated as $\varepsilon_{ih} = \partial \ln X_i/\partial \ln W_i$, with $\sum_h \varepsilon_{ih} = 0$, $\forall i$. They are proportional to Allen-Uzawa measures of substitution, defined as $\sigma_{ih} = \varepsilon_{ih}/\omega_h$, where $\omega_h = X_h W_h/G$ is the cost share. Concerning scale and capacity induced impacts we have $\varepsilon_{iY} = \partial \ln X_i/\partial \ln Y$, and $\varepsilon_{ik} = \partial \ln X_i/\partial \ln Z_k$, respectively. Shadow price responses are defined analogously, $\varphi_{kh} = \partial \ln F_k/\partial \ln W_h$, with $\sum_h \varphi_{kh} = 1$, $\forall k$. These parameters are interpretable as indirect measures of utilisation: $\varphi_{kh} > 0$, e.g. means that an increase in W_h brings about a positive change in F_k. Thinking of the shadow price as the marginal reward of desired stock, its increase materialises in a higher degree of utilisation of the relevant asset. On the other hand, flexibilities, $\varphi_{kj} = \partial \ln F_k/\partial \ln Z_j$, convey information on the long-run behaviour of quasi-fixed inputs, k and j being substitutes (complements) when $\varphi_{kj} < 0$ ($\varphi_{kj} > 0$).

All economic measures of capacity utilisation derive from the comparison of temporary and long-run equilibria (Berndt and Fuss, 1986). In particular, a dual indicator of the deviation of quasi-fixed inputs from their long-run levels is given by $\mathrm{CU} = C^*/C$, where C is total cost and C^* is total shadow cost, i.e. total cost with quasi-fixed inputs evaluated at their shadow prices. Under constant returns to scale (CRTS), short-run cost flexibility and CU coincide (Morrison, 1985):

$$\mathrm{CU} = 1 - \sum_k \varepsilon_{Ck} = \varepsilon_{CY} \tag{5}$$

where $\varepsilon_{CY} = \partial \ln C/\partial \ln Y$ and $\varepsilon_{Ck} = \partial \ln C/\partial \ln Z_k = (W_k - F_k)Z_k/C$. Using the notion of shadow price, one can determine whether the stock Z_k is in excess ($W_k > F_k$) or falls short ($W_k < F_k$) of its equilibrium level. In turn, over ($\mathrm{CU} > 1$) or under ($\mathrm{CU} < 1$) utilisation will prevail depending upon the algebraic contribution of each ε_{Ck}. If shadow and rental prices coincide ($W_k = F_k$), $\varepsilon_{Ck} = 0$, $\forall k$, and capacity is fully utilised ($\mathrm{CU} = 1$).

When variable returns to scale and sub optimal utilisation coexist, short-run cost flexibility necessarily captures both effects. However, under homotheticity, the two components are

$$\varepsilon_{CY} = \varepsilon_{\mathrm{CY}}^{L}\left(1 - \sum_k \varepsilon_{Ck}\right) = \varepsilon_{CY}^{L}\,\mathrm{CU} \tag{6}$$

where $\varepsilon_{CY}^{L} = d \ln C/d \ln Y = d \ln Z_k/d \ln Y$ ($\forall k$), i.e. all output elasticities of quasi-fixed inputs are the same and equal to the long-run (inverse of) returns to scale, ε_{CY}^{L}.

Finally, we define the rate of technological progress (regress) as the percentage reduction (increase) in variable cost over time, $\varepsilon_{Gt} = \partial \ln G(\cdot)/\partial t$. Generally, the advancement of knowledge manifests itself in a non-neutral manner; this bias can be expressed by the rate of change in factor proportions, $B_i = \partial \ln \omega_i/\partial t$, $\forall i$. Recalling that the SGM demand functions are in terms of input level, it can easily be seen that: $B_i = \varepsilon_{it} - \varepsilon_{Gt}$, where $\varepsilon_{it} = \partial \ln X_i/\partial t$. The semi-elasticities ε_{it}'s are not independent of one another, in that $\varepsilon_{Gt} = \sum_i \omega_i \varepsilon_{it}$ and, consequently, $\sum_i \omega_i B_i = 0$. Technological change is defined to be input i-using ($B_i > 0$), saving ($B_i < 0$), or neutral ($B_i = 0$), depending on whether relative change in input i is larger, smaller or the same as the rate of cost reduction, respectively. When all inputs are affected equiproportionally, i.e. $B_i = 0$, $\forall i$, overall neutrality is implied.

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[5] The shadow price equations F_k are used only to derive the parametric expressions of elasticities.

2.3. Input technical efficiency

In principle, one can distinguish between two notions of technical efficiency: output oriented, which reflects the capability of producing maximal output from a given set of inputs, and input oriented, which corresponds to producing a given output using a minimum amount of inputs. The two coincide if and only if constant returns to scale prevail (Färe and Lovell, 1978). In the present case, this correspondence vanishes, with important empirical implications. For example, the input oriented measure of technical efficiency does not enter the derived demands, but rather appears in the restricted cost function alone. The individual frontier can then be written as $(1/b_f)G_f$, where $(1/b_f)$, $0 < b_f \leq 1$, reflects the cost of radial over-utilisation of inputs (Bauer, 1990; Atkinson and Cornwell, 1994) on the fth farm.

Recent developments in parametric frontier modelling can be found in Fried et al. (1993) and Kumbhakar and Lovell (2000), among others.[6] As no single approach seems to prevail in terms of theoretical properties and/or empirical advantages, we opt for the fixed effect model (Schmidt and Sickles, 1984). This panel estimator is distribution-free, allows for correlation between efficiency and regressors, and becomes consistent as the temporal dimension approaches infinity (Nickell, 1981). Individual effects are accounted for by specific intercepts, which may be interpreted as reflecting unobserved structural heterogeneity such as input quality and/or managerial skill. Time-varying fixed effects seems a realistic assumption, which can be represented either according to parameterised functions of time or discretely by means of temporal dummies. Examples include Cornwell et al. (1990), Kumbhakar and Hjalmarsson (1993), Ahmad and Bravo-Ureta (1995, 1996) and Cuesta (2000), to name a few. Here, time-varying efficiency is approximated by a flexible second-degree polynomial.

A two-step estimator is considered (Ahmad and Bravo-Ureta, 1995; Cornwell et al., 1990; Kumbhakar and Heshmati, 1995; Kumbhakar and Hjalmarsson, 1993). First, the minimised variable cost is obtained using the parameter estimates of the demand system (4). Observed and fitted variable costs are related as follows:

$$\ln G_{ft} = \ln \hat{G}_{ft} + e_{ft}, \tag{7}$$

where $\hat{G}_{ft} = \sum_i W_{ift}\hat{X}_{ift}$. The first-step estimated residual, e_{ft}, is composed of two terms:

$$e_{ft} = \mu_{ft} + v_{ft} \tag{8}$$

where $\mu_{ft} = \ln(1/b_{ft})$, which is restricted to be non-negative, includes both the farm-specific effect and technical efficiency, and v_{ft} is the statistical noise, which is heteroskedastic by construction. In the second-step, individual effects and time-varying efficiencies can be estimated by the least squares procedure, as

$$e_{ft} = \sum_f (\mu_f + \mu_{1f}t + \mu_{2f}t^2)D_f + v_{ft} \tag{9}$$

where μ_f, μ_{1f}, and μ_{2f} are unknown parameters, D_f is a dummy whose value is 1 for the fth farm and 0 otherwise, and v_{ft} is assumed iid normal with mean zero and finite covariance matrix.[7] The predicted value $m_{ft} = (m_f + m_{1f}t + m_{2f}t^2)$ is the basis for calculating efficiency scores at the farm level:

$$\mathrm{TE}_{ft} = \frac{[\min_f \exp(m_{ft})]}{\exp(m_{ft})} \tag{10}$$

The numerator of (10) is the least predicted value in each cross-section of the panel, i.e. the best practice or the reference against which all others are compared in that year.

3. Data and variable construction

The farm production data is drawn from the Farm Accountancy Data Network (FADN) and consists of annual observations on 41 specialised dairy farms (defined here as those farms where 75% or more of total revenue is derived from dairy enterprise) in Lombardia. This northern region provides more than one-third of the milk supply in Italy. The investigation period covers the years from 1980 to 1992 and the panel is balanced. The analysis was restricted to a single region

[6] A review is also given in Bauer (1990), Battese (1992), Bravo-Ureta and Pinheiro (1993), and Coelli (1995).

[7] The HETERO option of LSQ command causes TSP to compute standard errors which are consistent even in the presence of unknown heteroskedasticity (White, 1980).

Table 1
Summary statistics of selected variables in the data set

Variable	Unit	Panel mean	Standard deviation	Minimum	Maximum
Output	10^6 (1990, Italian lira)	490.27	265.81	100.27	2004.08
Land	Hectare	69.25	43.02	14.70	190.00
Livestock	Cow equivalent	155.63	89.67	30.40	824.30
Purchased feeds	10^6 (1990, Italian lira)	117.83	75.17	5.15	483.04
Other inputs	10^6 (1990, Italian lira)	99.02	66.23	13.64	310.40
Hired labour	Worker equivalent	2.66	1.71	1.00	9.00
Family labour	Worker equivalent	3.32	1.55	1.00	9.00
Capital	10^6 (1990, Italian lira)	2432.54	1455.54	381.71	6848.15
Price index of other inputs	1990 = 1	0.86	0.14	0.51	1.03
Price index of feeds	1990 = 1	0.93	0.14	0.59	1.14
Wage of hired labour	10^6 (1990, Italian lira)	21.20	7.95	4.60	43.96
Age	Years	55.15	11.47	27	80

in order to ensure as much homogeneity as possible in input quality as well as technological and structural conditions. Accordingly, only farms with hired labour and located in the plain of the Po river were considered. The observed holdings are medium- to large-size compared to national standards.

The FADN does not provide farm gate prices of variable inputs or outputs, with the exception of hired labour and milk; hence, the relevant information is provided by Divisia indexes obtained by aggregating regional prices of the elementary components weighted by farm-specific cost (revenue) shares.[8] The resulting series are farm-specific due to differences in input and output compositions. Quantities are obtained by dividing the values of output and variable inputs by the farm-specific price index.

The vast bulk of output consists of milk. Some beef, mostly as a joint product to milk, deficiency payments and other production subsidies are also included. Aggregate output does not include categories such as intermediate inputs (feed grains, roughage, milk and so on produced on the farm).

Variable costs consist of three input categories: (1) purchased feeds; (2) other intermediate inputs; and (3) hired labour. Feed costs include aggregate outlays on concentrates, forages, feed grains and so on. The second group consists of the remaining intermediate inputs (mainly fertiliser, pesticides, seed, fuel, energy, veterinary costs, as well as overheads, i.e. the costs of repair and maintenance of capital equipment, insurance and rent). The wage rate per hour (with social cost included) is taken from the FADN.

The quasi-fixed inputs consist of the service flows from family labour and capital. The latter aggregates four assets: land, breeding livestock, machinery and buildings. Quantities of the fixed assets are calculated by dividing the invested capital by a price index of the corresponding services. User cost is defined as the sum of interest and depreciation cost at the farm level replacement value. Family labour is expressed in equivalent fully employed workers (2200 h per year). Technological change is represented by a time trend and is not farm-specific. The base year is 1990. Table 1 gives an overview of selected variables in the data set.

If subscript f refers to the farm ($f = 1, \dots, 41$) and t to the year ($t = 1980, \dots, 1992$), data are organised as a sequence of time series so that the slowest varying index is f. Each stacked vector contains 533 observations. An additive error term is appended to the behavioural Eq. (4). Parameters are estimated using iterative Zellner techniques[9] under the typical assumption that the error term v_{ift} for the ith equation is iid across units f over years t.

[8] We had no choice in this respect as prices are not available at the sub-regional level. However, since the analysis focuses on a sole region and the productive technology of the observed holdings is homogeneous, one can reasonably assume that the input market (defined here as the region) is also unique.

[9] The command used is LSQ of TSP 4.4.

Table 2
Test statistics for alternative model specifications

Test	H_0	Degrees of freedom (d.f.)	log-likelihood	Test statistic $\chi^2_{(\text{d.f.})}$	Critical value (at 1% level)
No differential intercepts	$b_{ii}D_M = b_{ii}D_L = 0$ $(i = 1, 2, 3)$	6	−7446.5	204.1[a]	16.8
CRTS	$b_i = b_t = b_{yy} = c_{kY} = 0$ $(i = 1, 2, 3;\ k = 1, 2)$	7	−7454.3	219.8[a]	18.5
Time independence	$b_{it} = b_t = b_{tt} = c_{kt} = 0$ $(i = 1, 2, 3;\ k = 1, 2)$	7	−7393.5	98.2[a]	18.5
Parameter stability	1980–1992 vs. 1980–1984	984	−2887.2	184.4[b]	1164.3

[a] Based on the likelihood ratio (LR) test: LR $= 2[\mathrm{l}(\mathrm{H}_1) - \mathrm{l}(\mathrm{H}_0)]\tilde{\chi}^2_{(\text{d.f.})}$, where l is the log-likelihood and d.f. the number of independent restrictions.

[b] Based on the following test $2[(N/N_1)\mathrm{l}(\mathrm{H}_1) + 0.5N\log(N/N_1) - \mathrm{l}(\mathrm{H}_0)]\tilde{\chi}^2_{(\text{d.f.})}$, where N and N_1 are the observations under H_0 and H_1, respectively, and d.f. $= m(N - N_1)$, with m number of estimated equations.

4. Empirical implementation and discussion

4.1. Specification tests

A number of formal statistical tests have been performed in search of an appropriate specification and the presence of structural change. Parameter restrictions under the null hypotheses, resulting statistics and critical values are shown in Table 2, while parameter estimates and approximated standard errors are reported in Appendix A.

First, we distinguish between holdings according to hectares of land[10] and check whether the intercept b_{ii} varies across classifications. Small, medium and large farms are defined as having less than 50 ha, between 50–100 and more than 100 ha, respectively. There are 17 small farms, 15 medium farms and 9 large farms. A restricted system, which does not account for heterogeneity, is compared to an unrestricted version (small farms being the reference), using the likelihood ratio (LR) test. The resulting LR-statistic is 204.1, suggesting that the null hypothesis is strongly rejected at the 1% significance level. The size-related dummy variables in Appendix A indicate that larger farms tend to have higher i/o coefficients of hired labour and lower i/o coefficients of purchased feed, ceteris paribus. This size effect fits in well with the established practice of producing forage on dairy farms in the Po plain.

The second question concerns scale economies. The null of CRTS in the long-run amounts to the following parameter restrictions: H_0: $b_i = b_t = b_{YY} = c_{kY} = 0$ $(i = 1, 2, 3;\ k = 1, 2)$. Since the sample statistic is 219.8, which is well in excess of the critical value $\chi^2_{(7)} = 18.5$ at the 1% level of significance, the CRTS hypothesis is decisively rejected.

Third, we explore whether the production function exhibits any exogenous technical change. Since ε_{Gt} represents the rate of cost diminution, the null hypothesis can be expressed as: H_0: $b_{it} = b_t = b_{tt} = c_{kt} = 0$ $(i = 1, 2, 3;\ k = 1, 2)$. The resulting LR-statistic is 98.2, meaning that this hypothesis is also rejected at 1% level of significance.

Based on the above test sequence, we believe that the regional dairy technology can be confidently identified with the estimated SGM, which is monotonic in W and Y (non-decreasing) at all sample points, and in Z (non-increasing) at the approximation point (globally), concave in prices and (globally) convex in quasi-fixed inputs. Given the monotonicity results and the negative (positive) semidefiniteness of the estimated $S(C)$ matrix, curvature criteria are satisfied globally by the proposed specification.

Finally, since our analysis refers to years in which policy changes are deemed to have affected the economic behaviour of milk producers, we look for structural change using the test proposed by Anderson and Blundell (1984). The model is estimated twice, over the entire period (1980–1992) and over the pre-quota years (1980–1984), with theoretical and

[10] We are aware that there can be more convenient proxies of farm size. Notwithstanding, Italian dairy farms with hired labour traditionally produce forage; hence, land is positively and highly correlated with other possible indicators (e.g. number of cows, specialisation index, gross revenue and/or economic size unit).

Table 3
Variable input elasticities (at the sample mean, approximated standard errors in parenthesis)

All farms, 1980/1992	Feeds	Other inputs	Hired labour	Output	Family labour	Capital
Feeds	−0.312 (0.144)	0.162 (0.134)	0.150 (0.039)	1.483 (0.041)	−0.116 (0.028)	−0.368 (0.038)
Other inputs	0.237 (0.196)	−0.354 (0.192)	0.117 (0.049)	0.430 (0.046)	0.129 (0.032)	0.288 (0.041)
Hired labour	0.321 (0.084)	0.171 (0.071)	−0.492 (0.066)	0.699 (0.058)	−0.354 (0.040)	0.161 (0.055)

approximation properties embedded. The resulting statistic is 184.4, which suggests that the null hypothesis of no parameter differences cannot be rejected, i.e. the model structure appears to be fairly stable between the pre- and post-quota periods.

4.2. Input demand elasticities

Since the results show little variation over time and farms, we discuss only panel mean estimates in order to conserve space. Table 3 reports variable input elasticities.[11] On the whole, input use is much more responsive to the scale of production than to prices. Hence, short-run changes in factor proportions mainly depend on the output level. Own-price and cross-price elasticities indicate that coefficients are accurately estimated and all responses are much smaller than unity, which suggests a rather rigid structure.

Direct responses of feeds and other inputs (which contains fertiliser) are comparatively low, indicating that feeding strategies and hence production of forage for the dairy herd are, to some extent, fixed within each production year. The own-price elasticity of hired labour (−0.49) shows a relatively higher degree of responsiveness.[12] It is slightly lower than Tiffin's (1991) estimate for the dairy sector in UK and Wales before supply control. The lower value found here may be due to the fact that in Tiffin's study the response (derived from a profit function) is uncompensated and labour (defined as family and hired labour) is variable, suggesting that the number of hours worked by family labour is more flexible. On the other hand, Stefanou et al. (1992) find that numerous responses of Germany dairy farmers did change with the introduction of the production quota. The change between periods appears to be more dramatic for the variable input demands; in particular, the substitution elasticity of hired labour declines in the post-quota period. Cross-effects show an overall substitutability, regardless of period and size group. In general, these responses are in the range of the estimates derived by Maietta (2000) under similar modelling assumptions for a panel of Italian dairy producers.

Table 3 also reports elasticities with respect to output and quasi-fixed inputs. Because of variable returns, scale elasticities do not resemble each other. A unit increase in output has a stronger effect on purchased feeds (1.5), whereas the responses of hired labour (0.7) and other inputs (0.43) are less than proportional. Feeds and other inputs adjust consistently to both fixed inputs, albeit in opposite directions, while the sign of hired labour adjustment depends upon which stock is changing. In particular, both family labour (−0.12) and capital (−0.37) substitute for purchased feeds. Hence, e.g. an increase in capacity due to land acquisition or renting for forage production aims, to some extent, at substituting the costly concentrates, which make up more than 46% of variable costs (Table 5). The two stocks and other inputs are complement; finally, family labour substitutes for and capital behaves as a complement of hired labour. Most of these adjustments are modest and within the range of price effects.

For the symmetry relationships attributable to the twice continuous differentiability of the cost function, we have that quasi-fixed input demand elasticities and shadow price elasticities (Table 4) share similar information.[13] For example, since self-employed farmers substitute for both purchased feeds and hired

[11] Formulas for the SGM elasticities are given in Kumbhakar (1990). We modify them to include quasi-fixed factors. Analytical derivatives and approximated standard errors are obtained through the TSP commands DIFFER and ANALYZ, respectively.

[12] Farm group elasticities, and other results not presented here, can be obtained from the authors.

[13] Namely, $\partial X_i/\partial Z_k = -\partial F_k/\partial W_i$. This proposition can be re-phrased in terms of elasticities as: $\varepsilon_{ik} = -(\omega_i^*/\omega_k^*)\varphi_{ki}$, where ω_i^* and ω_k^* are the input shares on shadow cost C^*, and φ_{ki} gives the impact of W_i on the quasi-rent of stock k.

Table 4
Shadow price elasticities (at the sample mean, approximated standard errors in parenthesis)

All farms, 1980/1992	Feeds	Other inputs	Hired labour	Output	Family labour	Capital
Family labour	0.601 (0.106)	−0.460 (0.197)	0.859 (0.162)	−0.298 (0.451)	−0.642 (0.348)	0.217 (0.051)
Capital	3.853 (1.919)	−2.066 (1.331)	−0.787 (0.632)	0.889 (0.498)	0.437 (0.245)	−1.974 (1.106)

labour, an increase of their market prices makes the marginal value product of family labour increase in the short run. The opposite holds for a change in the price of other inputs. Responses are normally higher for capital. In particular, its quasi-rent, and thus utilisation, increases much more than proportionally with feed prices (3.9), whereas wages (−0.8) and especially other input prices (−2.1) have negative impacts. As is evident from the standard deviations, output flexibilities are not statistical significant. Finally, cross-flexibilities seem to indicate that the two quasi-fixed inputs are weak complements. Complementarity of family labour and capital is a result that has been also observed by Pierani and Rizzi (1994) in a study that is not specific to dairy but to Italian agriculture.

From Table 5, it appears that variable cost has declined by 3.5% per year. The semi-elasticities indicate that the advancement of knowledge has had statistically significant impacts on factor intensities, independent of both relative prices and scale adjustments. The bias turns out to be towards the use of other inputs (0.22) and economising in both hired labour (−0.16) and purchased feeds (−0.08). We normally expect the ε_{it}'s to be negative since their weighted average equals ε_{Gt}. However, some of them may be positive. This is the case with other inputs (0.18); while feeds (−0.11) and especially hired labour (−0.19) register negative rates of change. This result is not surprising. The high rate of technical change and the complementarity of other inputs with the given capacity in the short run are coherent with practical observations. In the study area, almost the 4/5 of the cow herd is made up of highly selected breeding livestock, and purchased feeds have been progressively replaced by forage produced on farm land (which requires fertiliser among others) over the sample period (Osservatorio sul Mercato dei Prodotti Lattiero-Caseari, 2001). This technology leads to variable cost savings.

Table 5
Cost shares, technological biases and rates of change of inputs (at the sample mean, approximated standard errors in parenthesis)

All farms, 1980/1992	ω_i	B_i	ε_i
Feeds	0.465 (005)	−0.076 (0.023)	−0.111 (0.029)
Other inputs	0.318 (0.004)	0.218 (0.029)	0.184 (0.033)
Hired labour	0.217 (0.003)	−0.157 (0.037)	−0.191 (0.043)
Weighted sum	1	0	−0.035 (0.018)

4.3. Scale economies and capacity utilisation

The estimates of cost flexibility, ε_{CY}, scale economies, ε_{CY}^{L}, and capacity utilisation, CU, by farm group, are presented in Table 6. The prevalence of scale economies is evident. At the panel mean, ε_{CY} is about 0.61; the proportionate cost saving is not as strong in the long run and it is estimated around 14% (from 10 to 20% according to farm group). The discrepancy can be explained by the fact that more than one-fourth (29%) of the overall capacity is in excess. Both quasi-fixed factors contribute to the disequilibrium, though at notably different levels: the utilisation elasticity is about 0.09 and 0.2 for family labour and capital, respectively. This finding is evidence of their sub-optimal use. The estimates also suggest a positive relationship between farm size and excess capacity. In their study on dairy technology on Vermont farms, for example, Quiroga and Bravo-Ureta (1992) find that family labour and herd size (used as proxy for the scale of the operation) are significantly lower than their optimal levels. These findings were interpreted as consistent with a continuing shift toward fewer and larger farms and partly attributed to the existence of a price support system in the US agriculture.

The situation has changed notably over the years. Utilisation elasticities slope monotonically downward, with capital always pacing faster (Fig. 1) so that its value halves (from 0.33 in 1980 to 0.16 in 1992) and comes closer to that of family labour (0.08 in 1992).

Table 6
Dual measures of scale effects, capacity utilisation and elasticities of utilisation by group (at the sample mean, approximated standard errors in parenthesis)

Classification, 1980/1992	ε_{CY}	$\varepsilon_{CY}^{L}=\eta$	CU	ε_{CL}	ε_{CK}
<50 ha	0.664 (0.20)	0.904 (0.020)	0.735 (0.021)	0.101 (0.018)	0.165 (0.012)
50–100 ha	0.597 (0.019)	0.841 (0.025)	0.710 (0.019)	0.094 (0.014)	0.196 (0.016)
>100 ha	0.559 (0.024)	0.802 (0.035)	0.697 (0.024)	0.093 (0.017)	0.209 (0.019)
All farms	0.612 (0.018)	0.862 (0.021)	0.710 (0.019)	0.090 (0.013)	0.201 (0.016)

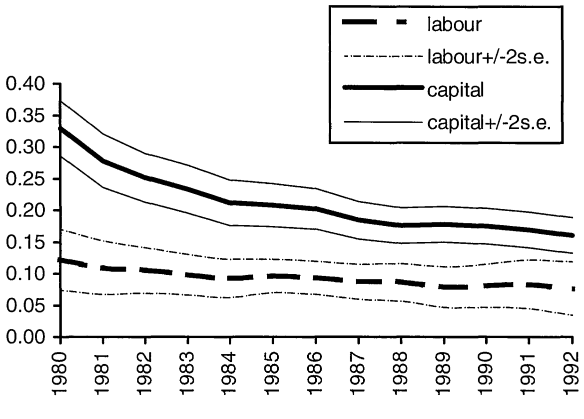

Fig. 1. Labour and capital utilisation elasticities (at the sample mean).

Hence, the trend is towards reducing excess capacity, which makes CU move from about 0.55 to around 76 in 1992 (Fig. 2).

Before the introduction of supply control, the real price of milk favoured the growth of imports over domestic supply, which explains the positive differential between rental and shadow prices of quasi-fixed factors. On the other hand, farmers did not initially perceive the introduction of production quota in 1984 as individual constraint. At least unofficially, Unalat and the ministry encouraged farmers to exceed their quotas as Italy was a deficit country, and the exact level of production had not been established. As a result, farmers increased their production and formed expectations of a more favourable determination of quota allotments (e.g. based on higher historical outputs) in the future. Failure to apply the system meant that Italy ran up a huge fine, which, in the logic of a supply-control measure, should have been paid by farmers. In fact, in the face of farmer protest, the fine was passed on to Italian taxpayers and the whole agricultural sector. Our data set shows this peculiar national story: while the adjustment of land is negligible (the average farm acreage remained about constant), milk output per cow (an average of 47 tonnes in 1980) and herd size (from an average of 137–176 units) increased at the significative rates of 2.2 and 1.8% a year, respectively, during the 1980–1992 period. Clearly, these adjustments contrast both with economic models seeking to explain the optimal adjustment of milk production under a delivery quota (Rasmussen and Nielsen, 1985) and with the experiences of other European countries (Burrell, 1989).

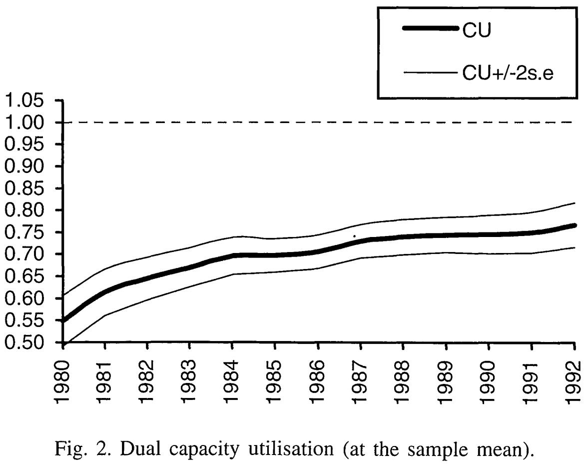

Fig. 2. Dual capacity utilisation (at the sample mean).

4.4. Technical efficiency

Before commenting on efficiency estimates we report on the test for temporal change in technical efficiency. The null hypothesis, H_0: $\mu_{1f}=\mu_{2f}=0$ ($f=1,2,\ldots,41$), is rejected soundly, the LR statistic being $\chi^2_{(82)}=303.35$. This result can be compared to other studies. Ahmad and Bravo-Ureta (1996) find the time-varying fixed effects superior in terms of statistical and plausibility criteria to alternative specifications. Kumbhakar et al. (1997) reach simi-

Table 7
Mean efficiency levels by farm size and selected years

Classification	1980	1982	1984	1986	1988	1990	1992	1980/1992	Minimum
<50 ha	0.718	0.647	0.614	0.615	0.649	0.714	0.638	0.657	0.477
50–100 ha	0.684	0.655	0.645	0.654	0.684	0.726	0.609	0.670	0.493
>100 ha	0.735	0.665	0.628	0.622	0.645	0.691	0.595	0.656	0.469
Mean	0.709	0.654	0.628	0.631	0.661	0.713	0.618	0.662	0.469
Minimum	0.532	0.516	0.494	0.498	0.530	0.586	0.469	0.469	

lar results comparing competing models proposed in earlier research. A specification in which technical efficiency is a quadratic function of time and varies across firms produces the most reasonable levels and temporal patterns.[14]

Regarding technical efficiency (Table 7), one can observe smooth changes over time and only small differences across types of farms. There seems to be no conclusive evidence that bigger farms are more technically efficient than small farms. This is not in line with the findings in a number of other studies (Kumbhakar et al., 1989; Bravo-Ureta and Rieger, 1991), and could be partly due to the fact that size is not properly captured by area.

Mean technical efficiency is predicted as being 0.66, with a minimum of 0.47. On average, then, the same level of output could have been produced at about 34% lower cost if farms had according to best practice. These measures tend to be somewhat lower than those derived from a variety of primal/dual stochastic frontiers (Bravo-Ureta, 1986; Bravo-Ureta and Rieger, 1991; Kumbhakar and Hjalmarsson, 1993; Kumbhakar and Heshmati, 1995; Maietta, 1998). But they are comparable to others using the distribution-free approach (Maietta, 2000; Hallam and Machado, 1996; Ahmad and Bravo-Ureta, 1996). This result possibly reflects the fact that individual dummies may pick up other latent features along with technical efficiency. Recent applied literature seeks to simultaneously control for these explanatory variables, which typically include farm size, rented/tenanted land, soil quality, geographical characteristics and farmer age and education[15] (Battese and Broca, 1997; Battese and Coelli, 1995; Maietta, 1998). A common finding is that age has a negative effect on technical efficiency. In our panel, the mean age is more than 55, with one fifth of farmers above 66 and a maximum age of 80.

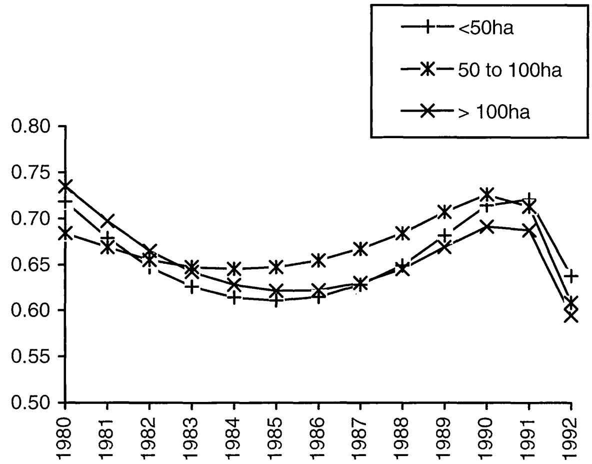

Fig. 3. Mean efficiency levels by farm size.

From Fig. 3, one notices some improvement after the introduction of the production quota followed by deterioration at the turn of the decade. Interestingly, this has been particularly true for medium and large farms. If this ranking reversal is based on valid comparisons, it contrasts with the opinion that most efficient farms are likely to persist in their pre-eminence. A tentative explanation is that large farms exploited their market power and exceeded their own allotments as long as quotas were granted to dairies and unused

[14] We also tested for the normality of residuals based on the normalized least squares residuals of Eq. (9) (Davidson and MacKinnon, 1993). The null of symmetry could not be rejected but there was no evidence of zero excess kurtosis at the 1% level of significance.

[15] More recently hypotheses have been advanced regarding the process by which financial exposure may exert an influence on efficiency (Nasr et al., 1998).

Table 8
Frequency distribution of efficiency levels by period

Efficiency levels	1980/1992 (%)	1980/1984 (%)	1985/1992 (%)
Below 50%	1.7	1.0	2.1
50.1–60%	27.9	32.7	25.0
60.1–70%	44.3	38.0	48.2
70.1–80%	16.7	19.0	15.2
80.1–90%	5.3	4.9	5.5
90.1–100%	4.1	4.4	4.0
Cumulate	100	100	100
Standard deviation	0.10	0.11	0.10

volumes could be reallocated to other producers. When the scope for this type of compensation was somewhat reduced, small farms in the panel turned out to be more flexible in adjusting their productive capacity.

From the frequency distribution in Table 8, it is apparent that, overall, only less than 10% of the observations have a predicted technical efficiency of 80% or more. Farms are mostly concentrated (about 44.3%) in the efficiency class 60.1–70%, which is also the class that has changed the most between periods, increasing from 38 (1980/1984) to 48% (1985/1992). The percentage of more efficient farms is about constant whereas that of farms having predicted technical efficiency of 50% or less doubled between pre- and post-quota periods. In a sense, the median class shifted towards slightly lower efficiency levels.

5. Concluding comments

In this study we have estimated a short-term specification of the SGM cost function that allows for quasi-fixed inputs and variable returns to scale so that the role of temporary equilibrium and economies of scale can be investigated. A two-stage procedure is used to estimate first the restricted cost function parameters and then farm-level technical efficiency. A balanced panel of Italian dairy farms observed over the years 1980–1992 serves as the case study. Production technology is analysed through a set of price elasticities of both variable inputs and shadows prices. Input technical efficiency is based on the fixed effect model, so that no distributional assumptions are required to separate out the first-stage estimated residuals. Individual scores, which vary according to a second-degree polynomial of time, reflect input over-utilisation as compared to the most efficient farm in the sample. The main results of this exercise can be summarised as follows.

In the short term, variable inputs are found to be inelastic, substitutes for one another and much more responsive to output than price changes. There exist scale economies and excess capacities in all farm groups; both quasi-fixed inputs are under-utilised, although the tendency is towards reducing disequilibrium over time.

Farms are characterised by a relatively high rate of cost reduction: 3.5% per year at the panel mean. Technological bias is towards the use of other inputs and economising in hired labour and purchased feeds.

Mean technical efficiency is 66%. This result might be interpreted as a measure of a disappointing technical performance after the introduction of milk quotas, but to certain extent it also reflects the approach chosen, as suggested by previous studies. There is little evidence that larger farms tend to be more efficient, which of course may depend on the hectare-based definition of size. Estimates show that small farms are only slightly less efficient as well as less heterogeneous than medium/large farms. There is, however, some evidence of smooth variation of efficiency over time, with a notable break at the turn of the decade when the rating reverses and small farms take the lead. Another finding is the narrow spread of efficiency scores: around 88% of farms are concentrated in the efficiency class 50–80% with a stable distribution over time.

Given the illustrative nature of our study, we believe the model has displayed some potential and has given pertinent answers within the chosen framework. Of course, further work remains to be done. First, the panel is not entirely representative. Hence, the behavioural insights and policy implications outlined above are not straightforwardly extendable to the whole Italian dairy industry. Second, generally farms are multi-output firms, with varying degrees of specialisation. So, using aggregate output as measure of economic performance hinders the possibility of appreciating the effects of production quota on decision processes on dairy farms. A multi-output multi-input specification of the restricted SGM cost function would be a more promising and appropriate framework of analysis. Moreover, given that the organisation of the dairy sector in the EU has changed significantly over time, it would be interesting to have an extended period of investigation and see whether the recently introduced transferability of quotas has brought about significant changes in the estimated parameter values and technical efficiencies. Last but not least, individual efficiency estimates have limited utility for policy and management purposes if empirical studies do not investigate the possible sources of inefficiency.

Responding to these questions will provide a better understanding of the role of production quota in the Italian dairy sector and represents an interesting agenda for future research.

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Appendix A. Parameter estimates of the SGM cost function

Parameter	Estimate	Standard error
s_{FF}	−38.75	17.62
s_{FI}	20.20	16.31
s_{II}	−29.94	16.01
b_{FF}	266.0	17.97
$b_{FF}D_M$	−28.11	4.323
$b_{FF}D_L$	−36.15	5.822
b_{II}	43.55	13.81
$b_{II}D_M$	21.34	3.180
$b_{II}D_L$	44.93	4.307
b_{HH}	60.35	9.473
$b_{HH}D_M$	26.01	2.942
$b_{HH}D_L$	28.23	3.855
b_F	−9.674	3.539
b_I	5.180	2.662
b_H	23.57	2.121
b_{Ft}	16.05	13.81
b_{It}	38.41	10.69
b_{Ht}	3.235	6.841
b_t	−14.46	5.858
b_{YY}	−88.83	12.86
b_{tt}	−33.09	8.394
d_{FL}	−35.94	12.57
d_{FK}	−86.49	16.92
d_{IL}	−5.986	9.695
d_{IK}	−7.848	13.04
d_{HL}	−30.64	6.211
d_{HK}	−10.56	8.398
c_{LY}	3.002	5.605
c_{KY}	25.38	7.494
c_{Lt}	0.4651	5.401
c_{Kt}	9.315	6.739
c_{LL}	8.470	4.367
c_{LK}	9.748	3.131
c_{KK}	11.22	6.742

The log-likelihood is −7344.88. Positive semidefiniteness of the Hessian of quasi-fixed inputs imposed with the semiflexible technique. Standard errors computed from the quadratic form of analytic first derivatives (delta method). Glossary of parameter subscripts: F, feeds; I, other inputs; H, hired labour; Y, output; t, trend; L, family labour; K, capital; D_M, classification dummy (1 if 50–100 ha, 0 otherwise); D_L, classification dummy (1 if more than 100 ha, 0 otherwise).

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