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# An Analysis of the Impact of Alternative EU Dairy Policies on the Size Distribution of Dutch Dairy Farms: an Information Based Approach to the Non-Stationary Markov Chain Model

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**An analysis of the impact of alternative EU dairy policies on the size distribution of Dutch dairy farms:  
an information based approach to the non-stationary Markov chain model**

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**Abstract**

This paper analyses the impact of the dairy quota scheme on the size distribution of the Dutch dairy industry. A non-stationary Markov model approach is used, where the transition probabilities are explained by a set of exogenous (policy) variables. Using an information theoretical approach, a model is estimated for The Netherlands and used to simulate the impacts of alternative EU dairy policies. Several results emerged: a) There is an autonomous over time decline in farm numbers (implying increase in farm size). b) The dairy quota regime positively influences 'small' and 'medium' farm sizes; c) Abolition of the dairy quota will negatively affect the total number of active farms and favours further increase of farm scale. d) Targeting support according to needs increases the number of active dairy farms as compared with the status quo.

**Keywords:** Farm size structure, dairy, milk quota, policy, maximum entropy

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# **An analysis of the impact of alternative EU dairy policies on the size distribution of Dutch dairy farms: an information based approach to the non-stationary Markov chain model**

## **1. Introduction**

Farm numbers have been declining drastically over the past decades, whereas farm size increases. Farm size and structure have long been issues considered by agricultural policy both in Europe (Keane and Lucey, 1997) and the US (Sumner, 1985). Given the main policy aim of supporting farmer's incomes and the close relationship between agricultural income distribution and farm size, this concern for distributional issues is no surprise. Given this correlation it is somewhat surprising that the agricultural policies have often been lacking effective policy instruments for influencing the firm size distribution. Benefits of the commodity programs, dominating traditional agricultural policies, are known to be roughly distributed in accordance with output and thus particularly benefit large farms which were less in the need for income support. The shift in agricultural policies from traditional price support to direct payments, initiated in the EU with the MacSharry reform of 1992, increases the available policy tools to make farm support more specific with respect to farm size and distribution. In both the 1990 and 1996 US farm bill debates, for example, several proposals were developed and sometimes implemented that were aimed at directing the bulk of farm program payments towards mid-sized farms, as measured in terms of gross sales (Wolf, 2001, 78). However, the potentials to pursue a targeted social agricultural policy are by far not realized (Podbury, 2000).

The aim of this paper is to analyse the farm size distribution of the Dutch dairy industry, with a particular focus on how the common agricultural policy (CAP) affected this distribution. More in particular, this research should provide a framework to analyse the implications of changes in the EU dairy policy (for example a substantial change in the dairy quota system or even its abolition) on the dairy farm size distribution.

A tool often used to describe changes in firm size distributions over time is the Markov process. This approach has the advantage that it relies on aggregated data of finite size categories -- the so-called Markov states-- at given discrete time intervals. Therewith it avoids the requirement longitudinal time-ordered micro data describing movement of individuals between different states, data which are only sparsely available. The main result of this analysis is the transition probability matrix, which describes the probability of a variable in a certain Markov state (for example a firm size class) to enter another Markov state. See Lee (1977), Zepeda (1995a,b) and Karantininis (2001) for a list of both general and agriculture related Markov studies<sup>1</sup>.

Based on Gibrat's Law, which states that firm growth is independent of firm size, firm size was initially often modelled as a purely stochastic Markov process, where the transition probability matrix (TPM) is assumed to be constant over time (Karantininis, 2001, 1). Applications following this "stationary Markov model" approach are Adelman (1958), Padberg (1962), Lee et al. (1977), and Oustapassidis (1986). However, this approach neglects the impact of the 'environment' on an industry's firm structure as well as the behavioral response of the entrepreneurs to these factors.

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<sup>1</sup> Reviewing a number of studies I have the feeling that Markov approaches tend to underpredict the number of small farms being active in the equilibrium (see also Zepeda, 1995b, 850).

Non-stationary Markov chain analysis explicitly allows variables characterising the industry's environment, among which policy variables aimed at influencing the sector, to explain the non-stationary transition probabilities. Examples of this approach are Lucas (1978) and Jovanovic (1987).

Most studies using a non-stationary transition probability matrix make very strong parametric distributional assumptions and other restrictions. Traditional estimation techniques like OLS fail, or require strong restrictions, because the estimated parameters must satisfy probability assumptions (non-negative probabilities, adding up). MacRae (1977) suggested a Logit transformation, which automatically satisfied the probabilistic constraints (see Zepeda 1995a,b for applications). However, there often is still a degrees of freedom problem which restricts the researcher to the choice of a limited number of explanatory variables. Even if sufficient degrees of freedom are available there can be problems with the convergence of the estimation algorithms (see Geurts, 1995).

In this paper we use the generalised maximum entropy (GME) formalism, which is based Shanon's (1948) information theory and Jaynes (1957a,b) as an estimation procedure. We employ the generalised cross entropy (GCE) formalism by Golan et al. (1996), and rely on recent Markov model applications using this approach by Golan and Vogel (2000), Courchane et al., (2000), and Karantininis (2001). The GCE formalism is used to recover coefficients of the effects of exogenous variables on individual transition probabilities when a specific (linear) functional form of the relationship is imposed. This method allows the use of an extensive set of explanatory variables. The impact of each variable on the individual probabilities and size categories is evaluated in the form of impact elasticities. Prior information on the TPM is introduced using the GCE formalism.

In the next section (Section 2) the general model structure and the selected explanatory variables are discussed. In Section 3 the GCE estimator for the non-stationary Markov model is presented. It also includes a description of the way prior information is used. Section 4 discusses the data, describes trends in the Dutch dairy farm size distribution, and provides the estimation results. Section 5 presents the simulation results indicating the impacts on the farm size distribution of four alternative dairy policies: status quo, Agenda 2000, quota abolition, and support rebalancing. Finally, Section 6 closes with some concluding and qualifying remarks.

## 2 The Markov model

Assume the firm size in the dairy industry is divided into  $J$  size categories and denote by  $n_{jt}$  the number of firms in the  $j$ -th size category ( $j=1, \dots, J$ ). Then a Markov chain process can be expressed as

$$n_{jt} = \sum_{i=1}^J p_{ij} n_{it-1}; \quad j = 1, \dots, J \quad (1)$$

where  $p_{ij}$  is the probability of transition from size  $n_i$  at time  $t-1$  to size  $n_j$  at time  $t$ , and  $i$  and  $J$  similar to  $j$  and  $J$ . The total number of farms existing at time  $t$ ,  $N_t$ , is equal to  $\sum_{i=1}^J n_{it}$ . In matrix notation equation (1) can be written as

$$\mathbf{n}(t) = \mathbf{P}' \mathbf{n}(t-1) \quad (2)$$

where  $\mathbf{n}(t) = (n_{1t}, \dots, n_{Jt})'$  is a  $K \times 1$  column vector and  $\mathbf{P} = (\mathbf{p}_1 \ \mathbf{p}_2 \ \dots \ \mathbf{p}_K)$  is the transition probability matrix (TPM) with each vector  $\mathbf{p}'_i = (p_{1i}, p_{2i}, \dots, p_{Ki})$ . The probability matrix is a stochastic matrix satisfying:

$$p_{ij} \geq 0 \text{ and } \sum_{j=1}^K p_{ij} = 1. \quad (3)$$

Besides the evolution of the size distribution an important and related issue is the modelling of entry and exit from the industry. The number of assumed potential entrants to the industry is known to have an important effect on both (short-run) projections and equilibrium solutions, even though it will not affect the estimated proportions of active firms falling in each size category (Stanton and Kettunen, 1967). By defining 'no production' as an additional category (say corresponding with state  $i=0$ ) it allows the modelling of entry and exit in the industry as well as the change in the size distribution of the 'active' or producing firms. In a fully competitive environment the number of 'firms in the 'no production' category is indeterminate, but might be expected to be large relative to the total number of 'active' farms (Stanton and Kettunen, 1967, 639). However, with respect to the dairy industry, in particular under the milk quota system, entry conditions seem a limiting factor. Therefore, the total number of dairy farms at the initial date (1972), will be used as an indicator of the total number of firms implying that the number of firms in state  $i = 0$  at that date is zero.

The probability matrix  $\mathbf{P}$  is unlikely to be constant but will rather be dependent on the economic situation both inside and outside the dairy industry. For that purpose it is assumed that  $p_{ij}$  from (1) is a function of a set of explanatory variables, or

$$p_{ij}(t) = f_{ij}(\mathbf{z}(t-1), \beta_{ij}) \quad (4)$$

with  $f_{ij}(\cdot)$  denoting a general function of a vector of  $N$  exogenous variables  $\mathbf{z}(t-1) = (z_{1,t-1}, \dots, z_{N,t-1})$  and  $\beta_{ik}$  a vector of parameters. This corresponds to  $\mathbf{P}(t)$  now being a time dependent or non-stationary transition probability matrix.

Examining the literature yields a number of variables likely to affect transition probabilities are relative prices, technological change, economies of size, farm debt, sunk costs, policy variables, demographic variables, indicators related to off-farm employment, etc. (see Goddard et al, 1993 and Zepeda, 1995b for an overview)<sup>2</sup>. Since in this paper the specific aim is to analyse farm size distribution with respect to key policy the selected explanatory variables (ignoring a constant) are limited to:

1. the level of aggregate milk output dairy quota;
2. a dummy trend variable indicating policy regime (see Section 4 for details);
3. the actual farm gate price of milk (based on actual fat and protein content);
4. a technology shifter (based on estimated autonomous milk yield development).

Although dairy cows play a non-negligible role in EU beef production, no explicit variable (like for example the beef price) is taken into account in this case. The Dutch dairy sector is highly specialized in dairying and therefore beef is considered as a by product.

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<sup>2</sup> Sometimes a distinction between the set of variables affecting entry/exit category and the set of variables affecting other categories (examples are Chavas and Magand, 1988, and Zepeda, 1995)

### 3. The information approach to recovering Markov transition probabilities

#### 3.1 The GCE estimator for the non-stationary Markov Model with prior information

The statistical model to be estimated consists of (2) and (4), to each of which a vector of disturbances is added ( $\mathbf{u}(t)$  and  $\mathbf{e}(t)$ ).

$$\mathbf{x}(t) = \mathbf{P}' \mathbf{x}(t-1) + \mathbf{u}(t) \quad (2a)$$

$$p_{ij}(t) = f_{ij}(\mathbf{z}(t-1), \beta_{ij}) + e_{ij}(t) \quad (4a)$$

and  $\mathbf{x}(t)$  the vector of proportions, obtained after normalization of the farm numbers in each size class  $n_{it}$  by the total number of farms in the first period,  $N_0$ .

A stationary TPM estimator using GCE is developed by Lee and Judge (1996), and Golan, et al., (1996). Assume  $\mathbf{u}(t)$  is a vector of disturbances with zero mean bounded within a specified support vector  $\mathbf{v}$ . Each element of the  $\mathbf{u}_t$  is parameterised as  $u_{it} = \sum_m^M v_m w_{itm}$ , where  $\mathbf{w}$  is an  $M$ -dimensional vector of weights (in the form of probabilities) for each  $u_{it}$ ,  $\mathbf{v}$  is an  $M$ -dimensional vector of supports. With  $\mathbf{x}(t)$  being a vector of proportions, the support vector can be set to ( $n_{it} \in [0,1]$ ) or to  $\mathbf{v} = [-1/K\sqrt{T}, \dots, 0, \dots, 1/K\sqrt{T}]'$  (e.g. Golan and Vogel, 2000 and Golan et al, 1996, 96-100).

By using GCE, any prior information about  $\mathbf{P}$  can be incorporated in the form of a matrix of priors  $\mathbf{Q}$ . Some research has indicated that farms typically do not decrease in size without going out of business, whereas other studies argue that might scale up or down in size, but with no more than one size category per transition (Zepeda, 1995b, 842). The latter assumption, which seem to be rather plausible when growth is considered as a continuous process, would imply that in general

$$x_{it} = p_{i-1,i,t} x_{i-1,t-1} + p_{i,i,t} x_{i,t-1} + p_{i+1,i,t} x_{i+1,t-1} \quad (5)$$

with all other elements in the  $i$ -th row of the probability matrix expected to be equal to zero. Rather than imposing this as a restriction which should be satisfied, like was done in Zepeda (1995b), here this information is used as prior information, which seems likely, but may be overruled by the data. Since the number of dairy farms is consistently diminishing over time, Geurts (1995) assumes that the probabilities of re-entry are equal to zero, or

$$p_{0j} = 0 \quad \text{for all } j = 1, \dots, K \quad (6)$$

with the zero subscript denoting the entry-exit category<sup>3</sup>. Another prior restriction could be to limit the number of non-movers to be not lower than a certain fraction  $c$ . The prior information can be directly included in the  $\mathbf{Q}$  prior-matrix of the GCE estimator (see below).

Prior information about the disturbance  $\mathbf{u}_T$ , call it  $\mathbf{w}_{itm}^0$ , can be incorporated as well. Since no clearly directed a-priori information was available, they are assumed to be uniformly symmetric about zero.

The objective of the GCE estimator is to minimize the joint entropy distance between the data and the priors. Let  $H(\cdot)$  be the measure of cross entropy, then the GCE is:

$$\min_{\mathbf{p}, \mathbf{w}} \left\{ H(\mathbf{P}, \mathbf{W}, \mathbf{Q}, \mathbf{W}^0) = \sum_i \sum_j p_{ij} \ln(p_{ij} / q_{ij}) + \sum_i \sum_t \sum_m w_{itm} \ln(w_{itm} / w_{itm}^0) \right\} \quad (7)$$

<sup>3</sup> Another prior restriction, not further considered here, could be to limit the number of non-movers to be not lower than a certain fraction.

subject to three sets of constraints: (a) The  $K \times T$  data consistency constraints (Equations (2)); (b) The normalization constraints for both the transition probabilities ( $K$  constraints) and the error weights ( $K \times T$  constraints):  $\sum_j^K p_{ij} = 1$ ,  $\sum_m^M w_{itm} = 1$  (proper distributions); and (c) the  $K^2$  non-negativity constraints for  $\mathbf{P}$  and the  $K \times T \times M$  constraints for  $\mathbf{w}$ :  $\mathbf{P} \geq 0$ , and  $\mathbf{w} \geq 0$ .  $H(\cdot)$  can be interpreted as a dual-loss function, which gives equal weights to prediction and precision. The solution to the above system of equations is derived Golan, et. al., (1996, Chapter 6).

### 3.2 A Markov model with a linear explanation function

In this section it is assumed that the transition probabilities can be explained by a linear function of exogenous variables. Allowing for non-stationarity of the TPM (i.e. substituting 4a into 2a), the Markov process can now be expressed as:

$$\mathbf{x}(t) = \mathbf{P}'(t) \mathbf{x}(t-1) = [\boldsymbol{\beta} \mathbf{z}(t) + \mathbf{e}(t)]' \mathbf{x}(t-1) + \mathbf{u}(t); \quad t=1, \dots, T \quad (8)$$

MacRae (1977) points out that in most estimation methods, each row of transition probabilities must be formulated to depend on the exact same set of exogenous variables. Furthermore, Lee, et al. (1977) and MacRae, (1977) develop the statistical properties of the disturbance terms  $\mathbf{e}$  and  $\mathbf{u}$  in (4). The GCE formalism allows the disturbances  $\mathbf{e}$  and  $\mathbf{u}$  to be recovered separately.

Let each  $\beta_{ijn}$  and each  $e_{ijt}$  be parameterized over a discrete finite support space:  $\beta_{ijn} = \sum_s^S d_{ijns} \theta_s$ , and  $e_{ijt} = \sum_h^H g_{ijth} \varphi_h$ , where  $\boldsymbol{\varphi}$ ,  $\boldsymbol{\theta}$  are support vectors of size  $S$  and  $H$  respectively, and  $\mathbf{d}$  and  $\mathbf{g}$  are the corresponding probabilities to be recovered. The Markov process in (8) now becomes:

$$x_{jt} = \sum_i^K x_{i,t-1} \left[ \sum_{\eta}^{N_{ij}} \left( \sum_s d_{ijns} \theta_s \right) z_{n,t-1} + \sum_h^H g_{ijth} \varphi_h \right] + \sum_m^M v_m w_{jtm}; \quad j=1, \dots, K, \quad t=1, \dots, T \quad (9)$$

where  $N_{ij}$  is the number of covariates in the  $(ij)$ th cell. Applying GCE  $\boldsymbol{\beta}$ ,  $\mathbf{e}$ , and  $\mathbf{u}$  can be determined through the recovered values of  $\mathbf{d}$ ,  $\mathbf{g}$ , and  $\mathbf{w}$  respectively.

The objective function corresponding to this model, which has to be minimized to obtain the GCE estimates, is

$$H(\mathbf{D}, \mathbf{G}, \mathbf{W}; \mathbf{D}^0, \mathbf{G}^0, \mathbf{W}^0) = \sum_i \sum_j \sum_n \sum_s d_{ijns} \ln(d_{ijns} / d_{ijns}^0) + \sum_i \sum_j \sum_t \sum_h g_{ijth} \ln(g_{ijth} / g_{ijth}^0) + \sum_i \sum_t \sum_m w_{itm} \ln(w_{itm} / w_{itm}^0) \quad (10)$$

where zero-superscripts denote prior values of matrices or parameters. As can be seen the probabilities are no longer direct elements in the objective function. As a consequence stochastic prior information on the probability structure  $\mathbf{P}$  can no longer directly be included, as was the case in (7).

The linear model (9) does not automatically satisfy the standard normalisation and non-negativity constraints on transition probabilities (see equation 3). Therefore a number of additional restrictions are imposed. Firstly, non-negativity constraints are imposed on the parameters, i.e.  $\mathbf{d} \geq 0$ ,  $\mathbf{g} \geq 0$ , and  $\mathbf{w} \geq 0$ . Secondly, the parameter and error weights are restricted  $\sum_s^S d_{ijns} = 1$ ,  $\sum_h^H g_{ijth} = 1$ , and  $\sum_m^M w_{itm} = 1$  to guarantee them to reflect proper distributions. However, this still not guarantees



that the probabilities of the transition matrix will satisfy the normal regularity conditions. In order to guarantee non-negativity of the probabilities and guaranteeing the sum of the probabilities being equal to 1 therefore

$$\sum_n^{N_{ij}} \left( \sum_s d_{ijns} \theta_s \right) z_{n,t-1} \geq 0, \quad i,j = 1, \dots, K; t = 1, \dots, T \quad (11a)$$

and

$$\sum_j^K \sum_n^{N_{ij}} \left( \sum_s d_{ijns} \theta_s \right) z_{n,t-1} = 1 \quad i = 1, \dots, K; t = 1, \dots, T \quad (11b)$$

should hold.

The prior information regarding the structure of the transition probability matrix (see equations 5 and 6) can also be formulated in terms of additional constraints. To exclude firms to move more than 1 stage upwards or downwards in the farm size distribution the following constraints satisfy

$$\sum_n^{N_{ij}} \left( \sum_s d_{ijns} \theta_s \right) z_{n,t-1} = 0, \quad i,j = 1, \dots, K \wedge j \geq i + 2; t = 1, \dots, T \quad (12a)$$

$$\sum_n^{N_{ij}} \left( \sum_s d_{ijns} \theta_s \right) z_{n,t-1} = 0, \quad i,j = 1, \dots, K; \wedge j \leq i - 2; t = 1, \dots, T \quad (12b)$$

The restriction excluding re-entry is<sup>4</sup>

$$\sum_n^{N_{ij}} \left( \sum_s d_{ijns} \theta_s \right) z_{n,t-1} = 0, \quad i=0, j = 1, \dots, K; t = 1, \dots, T \quad (13)$$

Since the error vector  $\mathbf{e}$  is not necessarily zero for each observation, the prior restrictions on upward and downward movement and re-entry do not imply any probabilities to be strictly restricted to zero.

The restrictions specified so far guarantee that the non-stationary Markov chain model satisfies all the regularity properties on the transition probabilities and the prior ideas, with respect to the sample information. When the estimated model is used to simulate out of the sample environment, say for  $\mathbf{z}$  values out of the range implied by  $\mathbf{z}(t)$ , there is no guarantee that the probabilities will satisfy the regularity properties<sup>5</sup>. The imposed restrictions only guaranteed the regularity conditions to locally (only for the considered sample) hold. A way to make it more likely that the model will be well-behaved when it is used in out of sample simulations, is to create a  $R \times K$  vector of 'extreme' explanatory variables, say  $\mathbf{z}^0$ , and add the following additional restrictions to the optimisation problem:

$$\sum_n^{N_{ij}} \left( \sum_s d_{ijns} \theta_s \right) z_{n,t-1}^0 \geq 0, \quad i,j = 1, \dots, K; t = 1, \dots, T \quad (14a)$$

and

<sup>4</sup> Note that for  $j \geq 2$  the re-entry restrictions are already implicit in (12a).

<sup>5</sup> This might be in particular a problem when the model is used to explain a highly disaggregated farm size distribution pattern.

$$\sum_j^K \sum_n^{N_{ij}} \left( \sum_s d_{ijns} \theta_s \right) z_{n,t-1}^0 = 1 \quad i = 1, \dots, K; t = 1, \dots, T \quad (14b)$$

The vector  $\mathbf{z}^0$  can be interpreted as a vector spanning up the discretion space of the exogenous policy (and non-policy) variables. If the  $\mathbf{z}$  values used for the out of sample simulation remain within this space, the model will not become subject to theoretical inconsistencies<sup>6</sup>.

### 3.3 Impact measures, diagnostics and inference

The direct effects of  $\mathbf{X}$  and  $\mathbf{Z}$  on the TPM are captured through the marginal effects  $\partial \tilde{p}_{ijt} / \partial \bar{x}_j$  and  $\partial \tilde{p}_{ijt} / \partial \bar{z}_n$ , with  $\tilde{p}_{ijt}$  the estimated transition probability (at time  $t$ ), and  $\bar{x}_j$  and  $\bar{z}_n$  denoting the mean values (over  $t=1, \dots, T$ ). These marginal effects evaluate the impact of each farm class  $j$  or explanatory variable  $n$  on the non-stationary TPM. The first one can best be interpreted in terms of the reallocation of the firm size distribution over time. The second one gives the marginal effect of a change in the explanatory variables on the TPM. A more convenient form is the "probability elasticity" (Zepeda, 1995b):

$$E_{ijtn}^p = \frac{\partial \tilde{p}_{ijt}}{\partial z_{tn}} \frac{\bar{z}_{tn}}{\tilde{p}_{ijt}} = \frac{\partial \sum_n \beta_{ijtn} z_{tn}}{\partial z_{tn}} \frac{\bar{z}_{tn}}{\tilde{p}_{ijt}} = \frac{\beta_{ijtn} z_{tn}}{\sum_n \beta_{ijtn} z_{tn}} \quad (15)$$

Where  $E_{ijtn}^p$  measures the percentage change of the  $n^{th}$  (exogenous) variable on the transition probability between states  $i$  and  $j$  at time  $t$ . The "impact elasticity" measures the (indirect) effect of the  $n$ -th exogenous variable on the number of farms in the  $j$ -th category (evaluated at sample average):

$$E_{jn}^{x(t)} = \frac{\partial x_{j,t}}{\partial z_{t-1,n}} \frac{\bar{z}_{t-1,n}}{\bar{x}_{j,t}} = \sum_i \frac{\partial p_{ij}}{\partial z_{t-1,n}} \bar{x}_{i,t-1} \frac{\bar{z}_{t-1,n}}{\bar{x}_{j,t}} = \left( \sum_i \beta_{ijtn} \bar{x}_{i,t-1} \right) \frac{\bar{z}_{t-1,n}}{\bar{x}_{j,t}} \quad (16)$$

Theil's inequality measure will be used as a goodness of fit indicator. This statistic, calculated for each size class, can be interpreted as a goodness of fit indicator.

## 4. Data and estimation results

The data represent the Dutch dairy farms size distribution from 1972-1999 and comprise 7 size classes. A graphical illustration of the evolution of the dairy farm size distribution in the Netherlands is given in Figure 1. The smaller size classes show a strong decline over time. The two largest size classes (70-99 and 100-...), in the following labelled as the 'large' farms, show an increase over the pre-quota period, a decline in the first 5 years after the introduction of the quota, and more or less stabilise thereafter. Class 50-69 shows a similar pattern, but is still going to slightly decrease from 1989 on onward. The mid-size class (30-49) shows a cyclical behaviour, with, however, a clear downward trend. In the following the size classes 30-49 and 50-69 are labelled as the category medium-sized farms. The 'small farms', consisting of size classes (1-29),

<sup>6</sup> Alternatively, one can assume a multinomial Logit transformation, which satisfies both the normalisation and the non-negativity constraints globally (MacRae, 1977; Golan et al., 1997).

show a sharp decline up till 1984, which is continued after the introduction of the milk quota, but at a lower rate of decline. Because the milk quota regime introduces a change in trends rather than a shift-effect, a dummy trend<sup>7</sup> variable was used capture the dairy quota effect. Inspection of Figure 1 suggests that the introduction of the milk quota system slowed down farm size adjustment in the Dutch dairy sector. However, the sectoral adjustment process did not come to a standstill but continued also after 1984. Over the considered period the total number of active farms declines by 72,235 farms or about 70%.

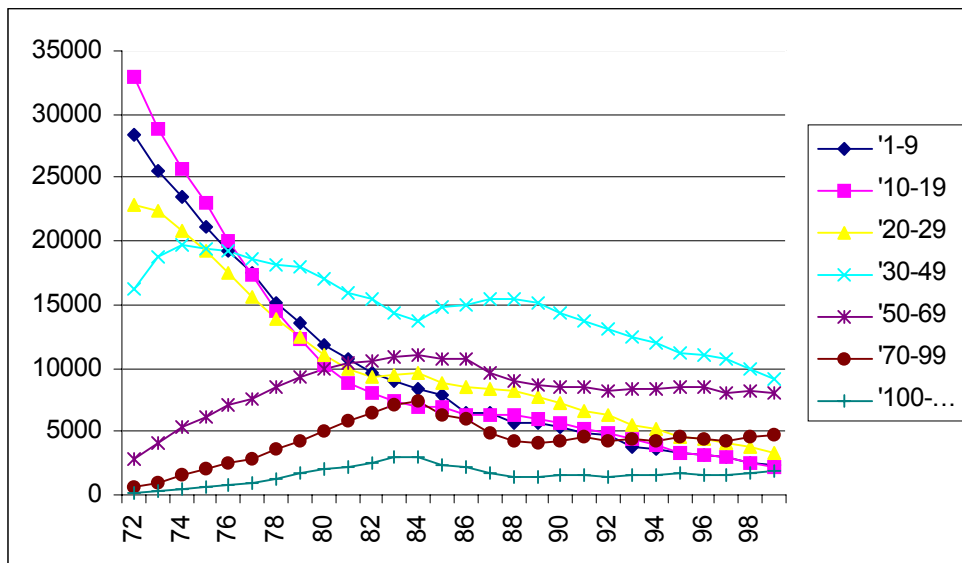


Figure 1 Dairy farm size evolution (absolute number of farms)<sup>8</sup>

In the estimated model the farm size distribution is aggregated into three classes (small, medium, and large) of active farms (see discussion above) and one class of inactive farms<sup>9</sup>. The GCE optimisation problem was programmed and solved in GAMS (see Brooke et al, 1992 for a further description of this algebraic modelling package). The estimated TPM is given in Table 1 (average evaluated 1997-1999)<sup>10</sup>. As the diagonal elements of the matrix show farms are very likely to stay in their current farm size class from one year to the next. However, in particular within the medium and large size classes the situation is far from stable, but large proportions are moving in and out.

<sup>7</sup> In contrast with Geurts (1995) who uses a shift-factor. The dummy trend variable is zero in the case of no quota, 1 in 1984, 2 in 1985 and so on.

<sup>8</sup> The figure is similar to the figure which one would get when the proportions (expressed in terms of the total number of active and inactive farms of the initial period 1972/73) would have been calculated. So Figure 1 gives the pattern of proportions the model has to explain.

<sup>9</sup> Aggregation (which implies a loss of information) is done as a first step in the analysis, in the final research paper the aim is to provide a more disaggregated analysis. Preliminary efforts show that it is technically possible to estimate such a disaggregated model (8 size classes (including inactive) with 7 explanatory variables).

<sup>10</sup> Note that since the TPM is not evaluated at the sample average some residuals in  $\mathbf{e}(t)$  might be non-zero causing some slightly negative probability values or slight deviations from the adding up condition.

*Table 1 Transition probability matrix*

	inactive	small	medium	large
inactive	1.00	-0.00	0.00	0.00
small	0.06	0.73	0.21	0.00
medium	0.00	0.29	0.53	0.18
large	0.00	-0.00	0.55	0.45

The impact of the exogenous variables on the number of farms in each class size are given by the impact-elasticities (Table 2)<sup>11</sup>. As the Table shows the impact of the technology shift is in favour of an increasing farm scale. In contrast, the milk quota dummy in particular favours the smaller farms.

*Table 2 Impact-elasticities of exogenous variables*

	inactive	small	medium	large
milk output	-0.038	0.076	0.005	-0.13
dummy trend	-0.031	0.056	0.012	-0.108
milk price	-0.025	0.036	0.025	-0.098
trend	-0.03	0.053	0.013	-0.104

An increase in the total milk output has a negative impact on the number of inactive farms, viz. profitability helps farms in surviving. The introduction of the milk quota regime leads to a decline in the number of inactive farms and large farms *ceteris paribus*. In contrast the quota regime favours the number of farms in the 'small' and 'medium' size classes. An increase in the milk price shows a similar impact than the quota introduction. This implies that a milk price decline leads to an increase of the number of inactive farms as well as the number of 'large' farms. It is strange that the overall impact of the growth trends works in favour of the number of 'small' farms and negatively influences the number of 'large' farms. This observation is in contrast with a lot of evidence and needs further study.

The value of the Theil inequality coefficient for classes 'inactive', 'small', 'medium' and 'large' is 0.0008, 0.0015, 0.008, and 0.029 respectively. Since the values are close to zero, the model does an acceptable job in explaining the evolution of farm size distribution in the past for all classes. Roughly seventy percent of the estimated parameters were significant.

## **5. Simulated Structural Changes in the Dutch dairy industry**

The model is used to investigate how the Dutch dairy farm size distribution may change in the future, depending on alternative policy scenarios. For the period 2000-2015 the following EU dairy policy scenarios are considered:

- 1) Status quo (no policy change) (SQ). Agenda 2000 is not implemented. Milk prices and milk quota remain at their year 2000 values, only trend variable changes.

<sup>11</sup> The estimated impact of the explanatory variables on the probability matrix are not presented here to save space, but are available upon request and will be included when final results are obtained.

- 2) Agenda 2000 (AG). Starting in 2004/05, the milk price is reduced by 15% in 3 steps as compared with 2000/01 and kept stable from 2007/08 and onwards. In conjunction milk quota are increased 1.5% and also kept constant from 2007/08 and onwards. A decoupled milk premium is introduced, which increases from 8.3 Euro/t in 2004/05 to 25.0 Euro/t in 2006/07<sup>12</sup>.
- 3) Abolition of milk quotas (QA). Milk quotas are abolished in 2004/05. The milk price is assumed to immediately drop by 30% and to further decline by 1% per year afterwards<sup>13</sup>. Milk output is assumed to increase with 1% per year<sup>14</sup>.
- 4) Rebalancing of support (RS). Similar to Agenda 2000, but with an adjusted compensatory payments regime. The introduced milk premium is differentiated between farm size classes (and no longer considered to be decoupled)<sup>15</sup>. Small farms and medium farms get an effective milk price of 120% and 110% of the actual milk price respectively. Large farms get no price premium support.

The simulation results are compared to the reference scenario Status quo (indicated as scenario SQ) and summarized in Figure 2 and Table 3. Figure 2 gives the evolution of the total number of active dairy farms under the alternative scenarios. As the Figure shows, the decline in the total number of farms will continue in the future, irrespective of the policies pursued. In SQ the total farm number declines from 31719 in 2000/01 (see Table 3) to 24871 in 2014/15 (-21%). This implies an annual reduction in total dairy farm numbers of 2%, which is not unrealistic, although lower than in the last decade. The quota abolition scenario QA leads to a total farm number decline of 26% in 2014/15, which is 5 percentage points more than SQ. In contrast, for the same year the support rebalancing scenario SR has a total farm number which is more than 6% above the SQ. The total difference between SR and SQ is 2800 farms (which is roughly 10 percent of the total farm population in 2014/15).

*Table 3 Summary of results policy scenarios*

	#-SQ 00/01	#-SQ 14/15	dev-AG	dev-QA	dev-SR
small	7897	12415	-394	-812	894
medium	17172	9516	-202	-484	448
big	6650	2940	18	14	176
total	31719	24871	-578	-1282	1518

<sup>12</sup> Since the milk premium is considered to be decoupled its impact is not taken into account in the simulation.

<sup>13</sup> The estimated milk price decline with quota abolition is based on Kleinhanss *et al* (2001) and Benjamin *et al* (1999) who calculate or use milk price reductions varying from -18% to -53%.

<sup>14</sup> Estimate based on assumed average milk shadow price of 70% of actual milk price, a milk supply elasticity of 0.5, an autonomous yield increase of about 1% per year, and a total herd size expansion of 0.5% per year.

<sup>15</sup> In this sense the support rebalancing (SR) scenario has similar characteristics as a negative income tax system.

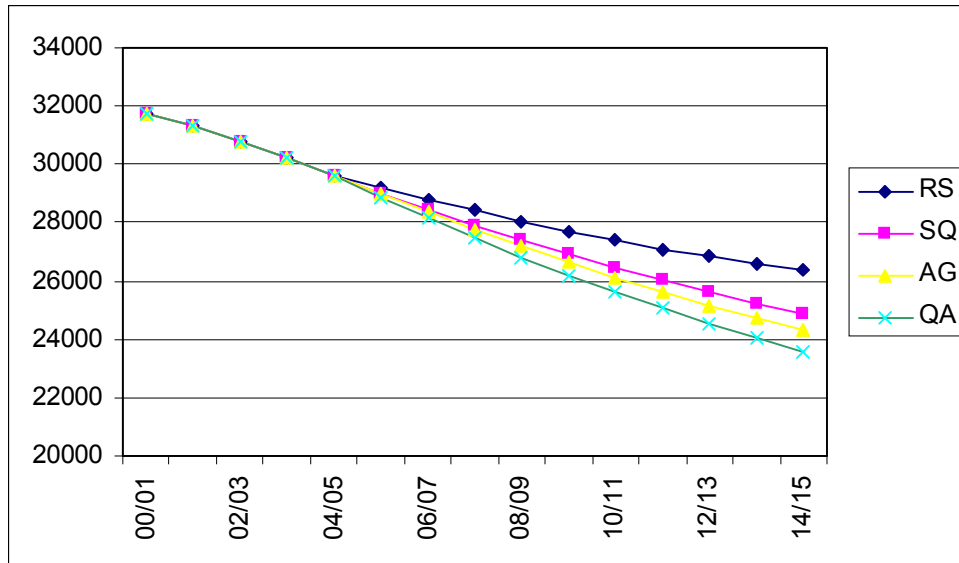


Figure 2 The evolution of total dairy farm numbers

Table 3 shows the impacts of the distribution of dairy farms. As comparing the first two columns show the total number of 'small' farms increases from 25% in 2000/01 to about 50% in 2014/15. This is unlikely when thinking from the current conditions, where there is relatively little part-time farming. At the same time the shares of medium sized farms declines from 54% to 38% and of large sized farms from 21% to 12%. As already noted in the end of the previous section this is still a puzzling outcome which needs further studying. As compared with SQ the abolition of the milk quota regime leads to lower number of 'small' and 'medium' farms and an increase of 'large' farms. This seems plausible, although in order for the milk balance to be hold, the average herd number in the various classes has to change. From the disaggregated data we see this is happening. The support rebalancing scenario has a positive impact on all (active) size classes, but in particular favours 'small' farms. The increase in the number of farms in all categories, with still the quota in place, seems unlikely, even with a significant increase in part-time farming.

## 6 Concluding remarks

Since this paper reflects work in progress these results have a preliminary character. Although the predicted adjustment in the aggregate number of dairy farms was not in contrast with expectations, as noted before, some puzzling findings about the trend pattern in the farm size distribution were found. The general result that keeping the milk quota system in place slows down the adjustment in the farm size evolution and therewith has a positive impact on the total number of active farms is however a plausible result which is also confirmed by Geurts (1995). Abolition of quota will increase the dynamics and is likely to improve the changes for the 'large' farms. Since 'large' farms due to their scale have a lower cost price of milk, this favours efficiency in production. If the main aim is to pursue a social policy supporting small and medium sized family farms, our first results suggest that the targeting of support makes it possible to do a much better job than is realised with the current policies.

Future research will focus on disaggregation of the analysis, a further refinement in the

explanatory variables (for example the inclusion of a revenue variable, which will allow to study the coupledness of direct payments). With respect to the estimation procedure a non-zero covariance structure of the error terms across size categories will be taken into account (along lines suggested in Golan, 1996, 186). Also a multinomial logit specification, which automatically satisfies the theoretical requirements, will be analysed, although our first results with this specification indicate that the multinomial logit is more sensitive with respect to convergence of the optimal solution as the linear specification chosen here. Moreover, we will elaborate on the use of prior information. Although for example the milk output condition (or quota constraint) is automatically satisfied by the sample data, it might still be possible to use production balance conditions to impose further constraints on parameters<sup>16</sup>. Finally, the entry/exit decision will be re-examined.

## References

- Chavas, J-P and G. Magand (1988) "A Dynamic Analysis of the Size Distribution of Firms: The Case of the US Dairy Industry". *Agribusiness* 4(4), 315-329.
- Benjamin, C. A. Gohin and H. Guyomard (1999) "The future of European dairy policy". *Canadian Journal of Agricultural Economics* 47, 91-101.
- Brooke, A., D. Kendrick and A. Meeraus (1992) GAMS User's guide; release 2.25. San Francisco: The Scientific Press.
- Courchane, M., A. Golan, and D. Nickerson (2000) "Estimation and evaluation of loan discrimination – An information approach." *Journal of Housing Research*, 11
- Geurts, J.A.M.M., (1995) *Technical & structural changes in Dutch dairy farms* (Msc-thesis). Wageningen: Wageningen University.
- Goddard, E., A. Weersink, K. Chen (1993) "Economics of Structural Change in Agriculture". *Canadian Journal of Agricultural Economics* 41, 475-489.
- Golan, A., and S.J. Vogel, (2000) "Estimation of non-stationary social accounting matrix coefficients with supply side information." *Economic Systems Research*, 12.
- Golan, A., G. Judge, and D. Miller (1996) *Maximum entropy econometrics*. Chichester: John Wiley and Sons.
- Hennessy, Th. (2001) "Modelling Farmer Response to Policy Reform: An Irish Example". Rural Economy Research Centre, Teagasc, Dublin, mimeo.
- Jaynes, E.T., (1957a) "Information theory and statistical mechanics". *Physics Review* 106, 620-630
- Jaynes, E.T., (1957b) "Information theory and statistical mechanics II." *Physics Review* 108, 171-190
- Jovanovic, B., (1982) "Selection and the evolution of industry." *Econometrica*, 50(3), 649–670
- Karantininis, K. (2001) *Information Based Estimators for the Non-stationary Transition Probability Matrix: An Application to the Danish Pork Industry*. Copenhagen: Department of Economics and Natural Resources, Royal Agricultural University of Copenhagen.
- Keane, M. and D. Lucey (1997) "The CAP and the Farmer" in: C. Ritson and D.R. Harvey eds. *The Common Agricultural Policy* (2<sup>nd</sup> edition) Wallingford: CAB International, 227-239.

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<sup>16</sup> For example, taking into account (15) and given fixed total milk output, it can be shown that  $\sum_{j=1}^J E_{j,n}^{x(t)} s_j^q = 0$ ,

where  $s_j^q$  represents the share of the  $j$ -th size category in total milk output, which imposes a restriction on the parameters.

- Kleinhanss, W., D. Manegold, M. Bertelsmeier, E. Deeken, E. Giffhorn, P. Jägersberg, F. Offerman, B. Osterburg and P. Salamon (2001) "Mögliche Auswirkungen eines Ausstiegs aus der Milchquotenregelung für die deutsche Landwirtschaft". Arbeitsbericht 5/2001 (Arbeitsgruppe "Modellgestützte Politikfolgenabschätzung" der FAL, Braunschweig).
- Lee, T.C. and G. Judge (1996) "Entropy and Cross-Entropy Procedures for Estimating Transition Probabilities from Aggregate Data." In Berry, D.A., C.M. Chaloner, and J.K. Geweke, eds, *Bayesian Analysis in Statistics and Econometrics*, Chichester: John Wiley and Sons, 153–164.
- Lucas, R.E.Jr., (1978) "On the size distribution of firms." *Bell Journal of Economics* 9: 508-203
- MacRae, E.C., (1977) "Estimation of time-varying Markov processes with aggregate data." *Econometrica* 45, 183–198.
- Mittelhammer, R.C., G.G. Judge and D.J. Miller (2000) *Econometric Foundations*. Cambridge: Cambridge University Press.
- Oustapassidis, K. (1986) "Greek Agricultural Cooperatives: radical change of trend recognition?" *Oxford Agrarian Studies* XV, I, 216-39.
- Padberg, D.I. (1962) "The use of Markov processes in measuring changes in market structure." *American Journal of Agricultural Economics*, 189-199.
- Podbury (2000) "US and EU agricultural support; Who does benefit?" *Current issues* 2000(2), ABARE (project 1633).
- Sumner, D. A. (1985) "Farm Programs and Structural Issues". In: B. Gardner, ed., *U.S. Agricultural Policy: The 1985 Farm Legislation*. AEI Symposia series, No. 85c, Washington DC.
- Shannon, C.E., (1948) "A mathematical theory of communication." *Bell System Technical Journal*, 379-423.
- Stanton, B.F. and L. Kettunen (1967) "Potential Entrants and Projections in Markov Process Analysis". *Journal of Farm Economics*, 49, 633-642
- Wolf, C.A. and D.A. Sumner (2001) "Are Farm Size Distributions Bimodal? Evidence from Kernel Density Estimates of Dairy Farm Size Distributions". *American Journal of Agricultural Economics* 83(1): 77-88
- Zepeda, L., (1995a) "Asymmetry and non-stationarity in the farm size distribution of Wisconsin milk producers: An aggregate analysis." *American Journal of Agricultural Economics*, 77: 837–852.
- Zepeda, L., (1995b) "Technical change and the structure of production: A non-stationary Markov analysis." *European Review of Agricultural Economics*, 22: 41–60.