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TECHNICAL EFFICIENCY MEASUREMENT OF DAIRY FARMS IN POLAND: AN APPLICATION OF BAYESIAN VED MODEL[1]

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Abstract: The purpose of this paper is to measure the technical efficiency of Polish dairy farms using a Bayesian Varying Efficiency Distribution (VED) model. In particular, the paper presents the design and assumptions of frontier stochastic production function for panel data. Furthermore, it specifies the microeconomic production function based on panel data, derived from the Polish FADN (Farm Accountancy Data Network). The main part of the paper presents key findings which form the basis of understanding the technological characteristics and average efficiency of Polish dairy farms. Moreover, the exogenous variables affecting the level of average farm efficiency are identified. They are the source of significant differences in levels of efficiency of dairy farmers surveyed.

Keywords: stochastic frontier models, Bayesian VED model, technical efficiency, dairy farms

INTRODUCTION

In 2010, Poland was the 12th-largest milk producer in the world and the 4th in the European Union [Statistical Yearbook of Agriculture 2012]. Although the number of dairy farms in Poland decreased from 874,000 in 2002 to 424,000 in 2010 [Raport z wyników, Powszechny Spis Rolny 2010], the production of milk increased from 11.5 billion liters in

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2004 to 12.1 billion liters in 2011 [Statistical Yearbook of Agriculture 2012]. However, the share of milk production in gross agricultural output (in current prices) decreased in 2011 compared with 2005: from 17.1% to 14.9% [Statistical Yearbook of Agriculture 2012]. These changes in the Polish dairy sector motivate our study, which analyses the technical efficiency of Polish farms, post-accession to the EU.

In previous studies of technical efficiency on different types of farms, inefficiency proved to be an inherent element in farming. The consequence of inefficiency is higher production costs, which of course negatively affect competitiveness. Therefore, the study of the causes of inefficiency has proven to be an important one [Alvarez and Arias 2004]. We can distinguish the following determinants of inefficiency among those commonly analysed: subsidies, the size of the land, the farm's economic size and its degree of specialization.

Since the milk market in the European Union is strictly regulated, the influence of Common Agricultural Policy (CAP) subsidies on the performance of dairy farms is relevant to a discussion of technical efficiency. The various CAP initiatives influence the farmer's optimal decisions through different mechanisms. Therefore, the impact of subsidies on the farms economic performance is an interesting question for policy makers who want to evaluate the effects of their decisions [Zhu et al. 2008].

Moreover, the empirical studies of technical efficiency of farms in Central and Eastern European (CEE) countries have caused disagreements about the relationship between farm size and efficiency. The commonly used measure of farm size is land area, but as this can be inappropriate for intensive livestock production, the weighting approach (e.g. European Size Unit – ESU) seems more appropriate. However, it is rarely used in efficiency studies [Gorton and Davidova 2004].

Another factor influencing the inefficiency of farms, which many authors have investigated in the studies on farming efficiency in CEE countries, is the degree of specialization in farm production.

The aim of this study is to perform a quantitative analysis of technical efficiency on Polish dairy farms and to identify the determinants of inefficiency using Bayesian VED model.

PREVIOUS STUDIES CONCERNING TECHNICAL EFFICIENCY OF LIVESTOCK AND DAIRY FARMS IN POLAND

The review of papers analysing efficiency of agriculture in Central and Eastern Europe is found in Gorton and Davidova [2004]. Papers on the efficiency of Polish farms include Munroe [2001] and Latruffe et al. [2004]. Research which in particular assesses the technical efficiency of Polish dairy farms can be found in Brümmer et al. [2002]. Analysis which presents the influence of CAP subsidies on dairy farms is found in Zhu et al.. [2008], Latruffe et al. [2012]. The relationship between technical efficiency and farm size is investigated in many papers [see review Alvarez and Arias, 2004]. The results of this relationship in Polish agriculture are presented by Van Zyl et al. [1996], Munroe [2001] and Latruffe et al. [2005]. Studies on how specialization influences technical efficiency in CEE countries are found in Brümmer [2001], Mathijs and Vranken [2001], Bojenc and Latruffe [2009].

In Polish scientific papers on efficiency analysis in agriculture, the dominant methodology of research was Data Envelopment Analysis (DEA) see, e.g., Świtłyk [1999, 2011], Rusielik [2002], Ziółkowska [2008], Rusielik and Świtłyk [2009], Kagan et al. [2010], Czyżewski and Smędzik [2010], Smędzik [2010, 2012], Bezat [2011]. The studies in which parametric approach i.e. stochastic frontiers models was used are for example: Kulawik [2008], Czekaj [2008], Czekaj et al. [2009], Rusielik and Świtłyk [2012]. This paper, on the other hand, utilizes the Bayesian approach to technical efficiency measurement to evaluate the efficiency of Polish dairy farms. The advantages of this approach are often discussed in the literature.

MODEL SPECIFICATION

There is a long history of economists quantifying inefficiency measures in production. Farrell [1957] was the first to measure productive efficiency. Presently, there are two main approaches which identify inefficiency as a deviation from a production or cost frontier: the parametric stochastic approach (Stochastic Frontier Analysis - SFA) and the nonparametric deterministic approach (DEA) [Prędki 2003]. Stochastic frontier models were simultaneously introduced by Aigner, Lovell and Schmidt [Aigner at al. 1977], as well as Meeusen and van den Broeck [1977]. In brief, SFA is based on an econometric model, which uses a conventional production function with two independent random disturbances, a symmetric around zero pure stochastic noise and a nonnegative error term representing inefficiency.

The model for a farm i (i=1,…,N) in period t (t=1,…,T) is written as follows:

$$\ln y_{it} = h(\ln x_{it}, \beta) + v_{it} + z_i \tag{1}$$

where:

y_{it} is the observed output quantity,

h is the production function,

x_{it} is the vector of the input quantities used by the farm,

β is the vector of parameters to be estimated,

v_{it} is a random error term representing random shocks ($v_{it} \sim N(0,\sigma_v^2)$).

Technical inefficiency relative to the stochastic production frontier is represented by the one-sided error component $z_i \geq 0$. Several distributions have been proposed for z_i, with the most common being the half normal, truncated normal or gamma distribution. Conventional assumption is that z_i and v_{it} are distributed independently of each other. Technical efficiency will be measured as $r_{ti} = \exp(-z_i)$, which is easily quantifiable (0,1]. A higher value for z_i equates to an increase in technical inefficiency. If z_i is zero, the farm is perfectly efficient. In many situations, the researcher is interested in making the inefficiency (a individual specific effect) depends on certain farm characteristics. It can be quite reasonable to assume that groups of similar farms, e.g. defined through their size or other factors, have similar efficiencies. Nevertheless the inefficiency distribution varies between groups.

This paper uses the Varying Efficiency Distribution model (VED) proposed by Koop, Osiewalski and Steel [Koop et al. 1997], which is more flexible than traditional frontier models. One of the advantages of this model is that it allows efficiency to vary while retaining certain individual characteristics of farms. The authors mentioned above propose that z_i represents an exponential distribution with a mean (and standard deviation) λ_i. The mean of z_i can depend on some (m-1) dummy exogenous variables s_{ij} (j = 2,...,m) explaining possible systematic differences in efficiency levels. The parameterization for the average efficiency takes the form

$$\lambda_i = \prod_{j=1}^{m} \phi_j^{-s_{ij}} \qquad (2)$$

where $\phi_j > 0$ is the unknown parameters and, by construction, $s_{i1} \equiv 1$. m = 1 is an important special case and is called the Common Efficiency Distribution (CED) model. The parameter vector ϕ indicates how the mean of the inefficiency distribution changes with the farm characteristics in s. In Bayesian analysis the parameters treated as random variables. This means that the inefficiency of farms are a priori linked through the ϕ.

The estimation of the model in equations (1) and (2) is possible using the maximum likelihood method if the parameters are assumed to be nonrandom constants, of course. However, in practice, this method is hampered by computational difficulties. Most often some non-Bayesian method used a two-step approach. Firstly, the model is estimated without the determinants of efficiency. Afterwards, at the second stage, the efficiency estimates obtained at the first stage were regressed on these farm characteristics. Therefore, this paper employs the Bayesian approach and in particular the Gibbs sampling algorithm for performing Monte Carlo integration (see [Osiewalski, Steel 1998], [Marzec, Osiewalski 2008]). In this context, it is worth mentioning that the continuous variables for s_{ij} (j > 1) can be used to explain inefficiency. But it causes some numerical difficulties because it requires a hybrid algorithm that combines Metropolis-Hastings and Gibbs sampling.

It is commonplace in frontier literature to impose regularity conditions drawn from economic theory. This is because the imposition of regularity conditions is relatively simple when employing Bayesian techniques compared to classical estimation. This paper makes use of a translog production function, but with no monotonicity conditions because it was satisfied for the entire sample data.

DATA AND RESULTS

The data used in the present study is taken from Polish Farm Accountancy Data Network[2]. The data covers farms whose main source of revenue in the analysed period came from milk production. The estimation of the Bayesian frontier model in this study involves a balanced panel data from 1,212 Polish dairy farms over the period 2004–2011.

The variables, names and symbols, used to construct output and inputs in the model, are according to European Commission document no. RI/CC 882 Rev.9 "Definitions of Variables Used in FADN Standard Results" dated November 2011. The output (Q)

[2] The authors are very grateful to dr inż. Dariusz Osuch from IERiGŻ, Warsaw, for providing access to the data.

includes production (SE131) and subsidies calculated according to the Polish methodology (SE605PL). Based on the literature and the data available, our empirical model includes the following 4 inputs: fixed capital (K), measured in Polish currency (PLN), labour (L), intermediate consumption, including materials and energy (M), measured in PLN and total utilised agricultural area (A), measured in hectares (SE025). The variable K is the sum of the value of buildings (SE450), machinery (SE455) and breeding livestock (SE460). The variable L is the total labour input, expressed in hours (SE011). Intermediate consumption is the sum of total specific costs (SE281) and total farming overheads (SE336). The construction of output and inputs is slightly different from those proposed in the literature, such as Bezat–Jarzębowska et al. [2012].

The production technology of dairy farms in this study is assumed to be specified by the translog function defined as follows:

$$\begin{aligned} h(x_{it};\beta) = {} & \beta_0 + \beta_1 \ln K_{it} + \beta_2 \ln L_{it} + \beta_3 \ln M_{it} + \beta_4 \ln A_{it} + \beta_5 \ln K_{it} \ln L_{it} \\ & + \beta_6 \ln K_{it} \ln M_{it} + \beta_7 \ln K_{it} \ln A_{it} + \beta_8 \ln L_{it} \ln M_{it} + \beta_9 \ln L_{it} \ln A_{it} \\ & + \beta_{10} \ln M_{it} \ln A_{it} + \beta_{11} \ln^2 K_{it} + \beta_{12} \ln^2 L_{it} + \beta_{13} \ln^2 M_{it} + \beta_{14} \ln^2 A_{it} + \beta_{15} t, \end{aligned} \tag{3}$$

where

t = time trend and symbols K, M, L and A represent the input set, which have been explained above.

The translog belongs to the class of so-called flexible functional forms. In contrast to a Cobb–Douglas production function, where returns to scale are global, translog functional form allows the estimated returns to scale to be different for each observation.

In the present study, there is an additional analysis, the aim of which is to identify exogenous determinants on dairy farm efficiency. The assumed factors determining inefficiency are: farm size measured by land size (classified by UAA – Utilised Agricultural Area) and the economic size (SE005). Both are expressed on a binary scale, i.e., if SE005 > 3 (on a 6-point ordinal scale), the dummy variable takes the value of 1, and zero otherwise. In addition, during the eight-year period cited above, farms earned as much as 39% of their primary income from agricultural activities other than the production of milk. Thus, the next additional determinant is the dummy variable indicating strict specialization. Indicated by an abbreviation, "specialization" is set to one when milk production is the main source of farm income in each of the eight years, or set to zero otherwise. Furthermore, as eighty-two percent (82%) of dairy farmers do not receive subsidies, this fact is reflected in this analysis: the variable "coupled subsidies" equals 1 if total subsidies on livestock (SE615) are greater than zero. These four dummy variables reflect the potential variation in farm efficiency.

Table 1, below, shows the mean values of the samples of individual variables. The average annual milk production is 187,000 PLN. Other characteristics show some features of the variables, including the skewed distribution in the population.

Table 1. Basic characteristics of sample farms: descriptive statistics

Variable	Mean	StDev	Percentile			Min	Max
			25th	50th	75th		
Output (‘000 PLN)	187	187	83	134	224	7	3486
Capital (‘000 PLN)	407	368	189	301	497	17	3750
Labour (in hours)	4 446	1 530	3 652	4 378	4 994	484	29 572
Intermediate consumption (‘000 PLN)	89	96	39	62	105	3	2176
Utilised agricultural area (in ha)	29	29	16	22	35	3	699

Source: own calculations based on data from Polish Farm Accountancy Data Network
Note: min and max denote minimum and maximum, respectively

Table 2 shows posterior mean parameter estimates of the Bayesian estimation of the stochastic frontier model.

Table 2. Posterior results of estimation the translog production frontiers (posterior means and standard deviation)

Variable	Mean	SD	Variable	Mean	SD
const	5.872	1.549	lnL lnM	0.002	0.026
lnK	-0.221	0.200	lnL lnA	-0.027	0.028
lnL	-0.913	0.297	lnM lnA	-0.042	0.019
lnM	0.786	0.217	$\ln^2K$	0.018	0.010
lnA	0.814	0.269	$\ln^2L$	0.043	0.021
lnK lnL	0.030	0.022	$\ln^2M$	0.009	0.012
lnK lnM	-0.025	0.018	ln2A	0.011	0.014
lnK lnA	0.002	0.017	T	0.025	0.001

Source: own calculations

A test comparing the translog versus Cobb-Douglas (C-D) specifications, with the restriction $\beta_5=\ldots=\beta_{15}=0$, revealed that this restriction (which has the Wald-test-statistic p-value of 1.2×10^{-5}) is rejected by the data. Therefore, the translog function seems a better representation of production technology on dairy farms.

Intermediate consumption (M) has the greatest impact on the volume of production. A 1% increase in the quantity of this factor results in an increase in production of about 0.56% (± 0.01%), ceteris paribus. A 1% increase in the utilised agricultural area (A) results in an increase in production quantity of about 0.22% (± 0.01%), ceteris paribus. The elasticity of buildings, machinery and breeding livestock (C) is 0.21%, so the impact of this factor is slightly smaller than utilized agricultural area factor. The smallest change (0.11%) in the quantity of production is effected by 1% growth in hours spent on farming.

Table 3. Posterior means and standard deviations of elasticities for a sample mean (variables on a logarithmic scale)

Variables	Average value	Mean	SD[3]
Capital (C)	305,300 PLN	0.21*	0.01
Labour (L)	4,246 h	0.11*	0.01
Intermediate consumption (M)	64,700 PLN	0.56*	0.01
Utilised agricultural area (A)	23 ha	0.22*	0.01
Returns to scale (RTS)	-	1.10*	0.01

Source: own calculations; Note: * significance at 1% levels

The coefficient on the time-trend variables in equation (1) is interpreted as a measure of pure technical change. The estimate of the parameter suggests that farms achieve an increase in production due to technical change. The growth rate in production over the past eight years has been 2.5% per year.

Another important issue is measuring economies of scale. A typical Polish producer of milk is characterized by increasing returns to scale, which is about 1.1 (±0.01). Almost all of the farms are characterized by increasing returns to scale, which does not exceed 1.3. Thus, a proportionate increase in input observably led to a more than proportionate increase in the production function. The opposite was true in only about 0.43% of the cases observed. Detailed information about the sample is presented in Table 4.

Table 4. Frequency distribution posterior means for RTS

Interval	[0.8; 0.9)	[0.9; 1.0)	[1.0; 1.1)	[1.1; 1.2)	[1.2; 1.3)
Frequency	1	41	4238	5393	23
Structure	≈0.0%	0.4%	43.7%	55.6%	0.2%

Source: own calculations

One of the main objectives of this study is to assess the technical efficiency of the dairy farms surveyed. The average efficiency of dairy farms is 0.86, which means that observed production amounts to 86% of potential output, i.e. the maximum output from the given inputs. The median efficiency score is 0.88 and the standard deviation is 0.05, reflecting also the low dispersion of efficiency scores along the sample. The efficiency level for the least efficient farms is 0.56 (±0.04). More than 35% of dairy farms should have efficiency levels in the range [0.8. 0.9], and almost 41% of farms have efficiency scores greater than 0.9.

In this study, there are four dummy variables to reflect the variation in farm efficiency. Table 5 also reports the posterior parameter estimates for the explanatory variables included in equation (2). A negative sign indicates that this variable has a negative impact on technical efficiency. Only two variables, the specialization and the land size, seem statistically significant. The Wald test indicates the null hypothesis that other dummy variables – "coupled subsidies" and "economic size" – are none-significant, i.e. not rejected (p-value equals 0.83).

[3] Identical values of standard deviation are due to rounding only.

Table 5. Sources of the technical inefficiency - posterior results

Variable (0-1)	Average value of variable	Parameter	Mean	SD
Specialization (s_{i2})	61%	$\ln(\phi_2)$	-0.23*	0.06
Coupled subsidies (s_{i3})	82%	$\ln(\phi_3)$	0.05	0.08
Land size (s_{i4})	24%	$\ln(\phi_4)$	-0.37*	0.08
Economic size (s_{i5})	68%	$\ln(\phi_5)$	-0.01	0.08

Source: own calculations; Note: * significance at 1% levels

SUMMARY AND DISCUSSION

The parameters of the translog production function are in line with the results of Brümmer et al. [2002], who also showed that the intermediate consumption factor had the highest elasticity and the labour factor, the lowest. Similar results for the Cobb – Douglas production function were obtained by Latruffe et al. [2004]. However, the parameters of the translog production function contradicted the results of Bezat–Jarzębowska et al. [2012] for the C–D production function. Furthermore, the results proved that elasticities vary over farms, revealing the Cobb–Douglas specifications to be inadequate.

The rate of technical change in the present study is higher than in Brümmer et al. [2002], who reported a technical regress (nearly 9% p.a.) in the period 1991 – 1994 for dairy farms. The negative rate of technical change was also found over the period 1996 – 2000 by Latruffe et al. [2008].

In the present study the majority of dairy farms were operating under increasing returns to scale. This result is in line with Latruffe et al. [2005] who reported that in 1996, 37% livestock farms had increasing RTS, but in 2000 this number jumped to 64%. The values obtained for RTS are confirmed by the results presented by Bezat–Jarzębowska et al. [2012] for the C-D function. However, Bezat–Jarzębowska et al. [2012] reported also decreasing RTS for a CES production function specification for two-and three factors of production.

The average technical efficiency level (0.86) in the covered period 2004 – 2011 is consistent with the results reported by Latruffe et al. [2004] who reported an average technical efficiency of 0.88 for a livestock farms panel in 2000, while in research conducted by Brümmer et al. [2002] for a sample of dairy farms in the Poznań region, average efficiency over the period 1991 – 1994 was equal to 75%. The mean total technical efficiency obtained by Latruffe at al [2005] using DEA method for a sample of livestock farms in 1996 was 0.85, which decreased to 0.71 in 2000. Because the results vary and our study was done using panel methods of estimation while previous studies used a cross-sectional data, it is not possible to clearly state if the average efficiency level increased after the accession the EU.

The results indicate that more diversified dairy farms are more technically efficient. This finding is confirmed by studies of farms in Slovenia conducted by Brümmer [2001]. However, this finding contrasts with the results reported by Mathijs and Vranken [2001] for dairy farms in Hungary, and Bojenc and Latruffe [2009] for Slovenian farms.

The analysis of the relationship between size and technical efficiency show that large farms (above 20 ha) are less technically efficient than smaller farms. These results are in line with Van Zyl et al. [1996], Munroe [2001] and Latruffe et al. [2005]. However it contradicts the conclusions drawn by Davidova et al. [2002]. Moreover, these results are not consistent with the previous study for the same sample of dairy farms by Marzec and Pisulewski [2013], but that research was conducted using a simple method.

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