RESEARCH

TECHNOLOGICAL CHANGE AND TECHNICAL EFFICIENCY FOR DAIRY FARMS IN THREE COUNTRIES OF SOUTH AMERICA

Boris E. Bravo-Ureta[1], Víctor H. Moreira[2]*, Amilcar A. Arzubi[3], Ernesto D. Schilder[4], Jorge Álvarez[5], and Carlos Molina[5]

ABSTRACT

The progressive liberalization of agricultural markets, along with the threat that imported products can pose to local producers, reveals the importance of productivity growth as a mechanism to improve competitiveness. Technical efficiency measurement is the most studied component of productivity because it can help to generate valuable information for policy formulation and farm level decisions focused on the improvement of farm performance. This study uses unbalanced panel data sets for dairy farms from Argentina, Chile and Uruguay, to estimate stochastic production frontier models. These frontiers are then used to estimate economies of size, technological change and technical efficiency. All estimations are based on the Battese and Coelli (1992) model, which is widely used in empirical productivity studies. The models for all three countries exhibit increasing returns to scale, which suggests that the dairy farms in the samples are operating at a sub-optimal size. The average annual rate of technological change for Argentina was 0.9%, for Chile 2.6% and for Uruguay 6.9%, while average technical efficiency was 87.0%, 84.9% and 81.1%, respectively.

Key words: stochastic frontiers, dairy farms, Argentina, Chile, Uruguay.

INTRODUCTION

Beginning with the Uruguay Round of the World Trade Organization (WTO), which opened in 1997, the multilateral liberalization of agricultural markets has been an important goal that has been strengthened by the reduction of protective tariffs in many countries (Hanrahan and Schnepf, 2005). The same authors point out that another element that helps to project a scenario of increasing liberalization of agricultural markets is the leading role that this sector played in the failed Doha Round in Qatar, which began in 2001. The opening up of trade has brought the increased competition of imported products in many markets, and it can be expected that new agricultural products will be incorporated into this process of liberalization. In this new scenario, the dairy sector plays a very important role in all the discussions about the liberalization of markets, fundamentally because of the high level of protection in many countries, and also because it is expected that the milk sector will experiment significant expansion globally, especially in Asia, North Africa, the Middle East, Central America and the Russian Federation (FAO, 2003; Blayney and Gehlhar, 2005; European Commission, 2005).

The progressive liberalization of agricultural markets, and the threat that the competition of imported products represents for local producers reveals the importance of improving productivity as a key factor to counteract pressures from countries that present greater competitiveness (Pinstrup-Andersen, 2002; Ruttan, 2002). An example that illustrates this process is the case of New Zealand, where the unilateral dismantling

[1] University of Connecticut, Department of Agricultural and Resource Economics, Storrs, 06269, Connecticut, USA. E-mail: boris.bravoureta@uconn.edu
[2] Universidad Austral de Chile, Instituto de Economía Agraria, Campus Isla Teja, Valdivia, Chile. E-mail: vmoreira@uach.cl
*Corresponding author.
[3] Universidad Nacional de Lomas de Zamora, Facultad de Ciencias Agrarias, Ruta Provincial 4, kilómetro 2, (1836) Llavallol, Buenos Aires, Argentina.
[4] Instituto Nacional de Tecnología Agropecuaria (INTA), Dirección Nacional, Rivadavia 1439 (1033), Buenos Aires, Argentina.
[5] Universidad de la República, Facultad de Agronomía, Av. 18 de Julio 1968, Montevideo, Uruguay.
Received: 11 October 2007. Accepted: 06 March 2008.

of agricultural subsidies has been accompanied by significant increases in farm efficiency and in total factor productivity at the national level (Sandrey and Scobie, 1994; Evans *et al.*, 1996; Paul *et al.*, 2000; ABARE, 2006). Today, New Zealand is the major exporter of milk products in the world (Blayney and Gehlhar, 2005).

To achieve significant increases in productivity and competitiveness, policies are required that encourage the adoption of new technologies, as well as measures that promote the efficient use of existing technologies. It is also necessary that policy makers, producers and agricultural extensionists have access to empirical studies that allow them to elucidate the effects of different factors that influence increases in productivity (Russell and Young, 1983; Kalirajan, 1984). Despite the importance of this topic, there are only a few specialized studies that consider the case of the Southern Cone countries of South America. Consequently, the objective of this paper was to analyze the level of technical efficiency (TE) at the micro-economic level (farm), considering samples of dairy farms in Argentina, Chile and Uruguay. In addition, technological change (TC) and economies of size (EOS) were analyzed for the three samples.

MATERIALS AND METHODS

Stochastic production frontier

To achieve the proposed objectives, stochastic production frontiers (SPF) were estimated, using unbalanced panel data. Frontier models can be classified into two basic types: parametric and non-parametric. Parametric frontiers require the specification of a particular functional form and can be classified as deterministic and stochastic. The deterministic model assumes that any deviation from the frontier is due to inefficiency, while the stochastic incorporates statistical noise. In this respect, in the case of deterministic frontiers, any measurement error and any other source of stochastic variation in the dependent variable is attributed to inefficiency, making the estimations of TE sensitive to extreme values (Greene, 1993). On the other had, the SPF resolves the problem of extreme values incorporating a compound error with a two-sided symmetrical term and a one-sided component. The latter reflects inefficiency, while the two-sided error captures random effects outside the control of the production unit.

The production frontier used here follows the structure of Battese and Coelli (1992), which has become very popular in recent years. In 1995 these authors published an extension of the original model, which is typically used when there are data that can be used to explain the variation in TE (Battese and Coelli, 1995). However, in the present study, such data were not available.

In accordance with Battese and Coelli (1992), the SPF can be represented as:

$$Y_{it} = exp(x_{it}\beta + V_{it} - U_{it}) \qquad [1]$$

where Y_{it} is the output of the i-th farm in the t-th time period; x_{it} is a vector $(1 \: x \: K)$ of inputs and other explanatory variables for the i-th farm in the t-th time period; β is a vector $(K \: x \: 1)$ of the unknown parameters to be estimated; V_{it} is the random error, which is supposed to have a normal distribution with mean zero and constant variance (σ_v^2), that is $V_i \sim iid \: N(0, \: \sigma_v^2)$; and U_{it} is the non-observable and non-negative random error, that captures technical inefficiency for the i-th farm.

Following Battese and Coelli (1992), U_{it} can be defined as:

$$U_{it} = \{exp[-\eta(t - T)]\} \: U_i \qquad [2]$$

where η is an unknown scalar to be estimated, t is the time period analyzed and T is the total number of periods. TE increases, remains constant or decreases with time when $\eta > 0$, $\eta = 0$ or $\eta < 0$, respectively. The U_{it} term can have different specifications and the most popular are the non-negative distribution of a truncated normal with an average μ and a constant variance ($U_i \sim iid/N(\mu, \: \sigma_U^2)/$) and the half-normal distribution ($U_i \sim iid/N(0, \: \sigma_U^2)/$). Coelli *et al.* (1998) suggests that the choice of a more general distribution, such as the truncated normal, is generally preferable. However, this is an empirical question and consequently, in this paper, the truncated normal distribution was tested against the half-normal.

TE for the i-th farm is given by:

$$TE = exp(-U_i) \qquad [3]$$

where U is specified in equations [1] and [2]. The TE for each farm is calculated by using the conditional expectation of $-U_i$ ($exp(-U_i)$), given the compound error term (V-U) (Jondrow *et al.*, 1982; Battese and Coelli, 1988). All these calculations were done using the software FRONTIER 4.1, which provides maximum likelihood estimates for the parameters of the stochastic frontier model (Coelli, 1996). FRONTIER 4.1 is a program that is often used in estimations of stochastic frontiers and its distribution is free (Sena, 1999), which is not the case with other alternative programs (e.g., LIMDEP, STATA).

Considering the specification indicated above, the null hypothesis that technical inefficiency is not present in the model was tested. This is equivalent to saying that $\gamma = 0$, taking into account that this parameter corresponds to the ratio between the variance of the one-sided error (σ_U^2) and the total variance ($\sigma_v^2 + \sigma_U^2$), that is $\gamma = \sigma_U^2/(\sigma_v^2 + \sigma_U^2)$ and consequently it ranges between 0 and 1 (Battese and Corra, 1977). The null hypotheses that technical inefficiency is time invariant (H_0: $\eta = 0$) and that it follows a half-normal distribution (H_0: $\mu = 0$) were also tested.

Data and empirical model

Descriptive statistics for the data used in this analysis are presented in Table 1. The data from Argentina came from a sample of dairy farms located in the Abasto Sur basin, in Buenos Aires Province. The data was collected by researchers from the Universidad

Table 1. Descriptive statistics of dairy farm data for Argentina, Chile and Uruguay.

Country/Variable	General average	Average per period				
Argentina		**1997-1998**	**1999-2000**	**2001-2002**		
Milk, L yr^{-1}	1 028 372	1 064 015	980 191	1 019 305		
SD [a]	523 977	521 096	575 399	501 568		
Cows, number yr^{-1}	160	151	159	174		
SD	72	63	76	80		
Labor, equivalent-worker	4.6	5.0	4.3	4.2		
SD	1.8	2.0	1.5	1.6		
Costs in feed, US$ yr^{-1} [b]	110 387	145 721	108 919	64 006		
SD	67 724	69 786	63 874	31 406		
Veterinary costs, US$ yr^{-1}	11 634	14 028	12 900	7 389		
SD	7 364	7 938	7 824	3 593		
Number of farms [c]	82	35	21	26		
Chile		**1996-1997**	**1998-1999**	**1999-2000**	**2000-2001**	**2001-2002**
Milk, L yr^{-1}	55 010	33 016	45 000	56 810	84 285	95 469
SD	47 929	20 677	39 759	44 098	57 600	69 135
Cows, number yr^{-1}	25	21	24	27	29	31
SD	15	10	14	20	13	17
Labor, equivalent-worker	3.3	3.9	3.6	3.2	2.2	2.5
SD	2.3	1.4	3.2	1.9	0.4	1.2
Costs in feed, US$ yr^{-1} [b]	3 214	2 063	1 837	2 919	3 806	9 388
SD	3 896	1 174	1 772	3 167	2 362	7 214
Veterinary costs, US$ yr^{-1}	316	287	187	337	510	544
SD	299	180	206	224	416	471
Number of farms	92	20	33	18	10	11
Uruguay		**1999-2000**	**2000-2001**	**2001-2002**	**2002-2003**	
Milk, L yr^{-1}	848 321	875 259	739 653	914 308	892 313	
SD	655 832	537 802	615 381	699 623	833 278	
Cows, number $year^{-1}$	213	206	188	233	235	
SD	153	131	134	169	192	
Labor, equivalent-worker	5.1	5.3	4.8	5.3	5.3	
SD	2.6	2.3	2.6	2.6	3.1	
Costs in feed, US$ yr^{-1} [b]	63 588	88 892	57 634	55 849	42 771	
SD	56 189	63 690	52 515	51 409	41 932	
Veterinary costs, US$ yr^{-1}	15 542	19 606	13 801	16 058	10 944	
SD	11 447	11 282	10 928	10 905	11 641	
Number of farms	147	42	43	37	25	

[a] SD: standard deviation; [b] US$, the reference year is July 2004-June, 2005 or 2004-2005; [c] Corresponds to total farms.

de Lomas de Zamora in Buenos Aires, incorporating three agricultural periods (1997-1998, 1999-2000 and 2001-2002), and included 46 dairy farms with a total of 82 observations. The data from the Chilean sample came from 48 farms of small milk producers belonging to the Farm Management Center of Paillaco, Valdivia, located in the south of the country in the communities of Paillaco, Los Lagos and Futrono, and covers the agricultural years of 1996-1997, 1998-1999, 1999-2000, 2000-2001 and 2001-2002, with a total of 92 observations. The Uruguayan data, collected by researchers from the Universidad de la República, in Montevideo, cover four agricultural years (1999-2000, 2000-2001, 2001-2002 and 2002-2003) and include 70 dairy farms and a total of 147 observations. It is evident that all three data sets are unbalanced panels. All variables expressed in monetary values were available in nominal dollars for each period and country. The nominal dollars were adjusted according to the purchasing power of each country (Argentina, Chile or Uruguay) in comparison to the United States, using the Consumer Price Index (CPI) of each country. This was done by multiplying the nominal dollars by the CPI of the United States, and subsequently dividing it by the specific CPI of each country, for each one of the periods considered. All the adjusted variables are then converted to real dollars using the period of July 2004 - June 2005 as the base.

Three separate SPF models were estimated, one for each country, using a translog (TL) specification. This model can be represented as:

$$y_{it} = \alpha_0 + \sum_{k=1}^{4} \beta_k x_{kit} + \frac{1}{2}\sum_{k=1}^{4}\sum_{l=1}^{4} \beta_{kl} x_{kit} \times x_{lit} + \sum_{k=1}^{4} \delta_k x_{kit} \times t + \lambda_l t + \frac{1}{2}\lambda_{11} t^2 + V_{it} - U_{it}, \quad [4]$$

where the sub-indexes k represent the k-th explanatory variable, i reflects a specific farm, t is a tendency variable to capture TC, and the sub-index j, which indicates the country, is omitted to simplify the exposition.

The dependent variable (y) is the natural logarithm for annual output per farm, measured in liters. The explanatory variables, also expressed in natural logarithms, are the following: average number of cows per dairy farm (*CO*); labor, measured in equivalent workers (*LB*); purchased feed (*FD*) including concentrate feed, hay, minerals and all the costs associated with the production of hay, ensilage and grass; and veterinary input costs (*VE*). The definition of t as a tendency variable that captures TC is modified with each data set. In this manner, the value of t for Argentina is 1 = 1997-1998, 3 = 1999-2000, and 5 = 2001-2002; for Chile 1 = 1996-1997, 3 = 1998-1999, 4 = 1999-2000, 5 = 2000-2001, and 6 = 2001-2002; and for Uruguay 1 = 1999-2000, 2 = 2000-2001, 3 = 2001-2002 and 4 = 2002-2003. The random terms V_{it} and U_{it} are the same as were already defined in Equations [1] and [2], and the Greek letters represent the parameters to be estimated. As is common practice in the TL models, all variables are expressed as deviations from the geometric mean, which makes it possible to interpret the linear coefficients directly as partial elasticities of production.

RESULTS AND DISCUSSION

Different specifications of the stochastic frontier were estimated for the three countries, with the objective of determining the most consistent model for the available data. The preliminary analysis revealed that the TL functional form is superior to the Cobb-Douglas, which is consistent with the production frontier literature (Coelli *et al.*, 2005; Bravo-Ureta *et al.*, 2007). The term that captures TE follows a half-normal distribution, and TE is statistically significant and constant over time. Consequently, the analysis developed from here on is limited to the results from the model with these predominant characteristics. More details regarding the procedures used for model selection can be found in Moreira (2006).

In the model for Argentina the linear parameters were highly significant, with the exception of the coefficient for the time trend t. The parameter for t^2 (time squared) was positive and significant, while the only significant parameters for the interaction terms were those for t and *VE* (veterinary inputs) (Table 2). The coefficient for *CO* (cows) was the most important among the partial elasticities, which implies that a percentage change in the number of cows would have a greater influence on milk production than a similar change in any other input. These results are consistent with many other studies, including Kumbhakar *et al.* (1991), Arias and Álvarez (1993), Heshmati and Kumbhakar (1994), Ahmad and Bravo-Ureta (1996), Jaforullah and Devlin (1996), Cuesta (2000), Lawson *et al.* (2004a; 2004b), and Moreira *et al.* (2006).

The model for Chile also presented highly significant first order parameters, with the exception of *LB* (labor) and t (Table 2). Most of the comments made for Argentina are also valid for Chile, except for the parameter for t^2,

Table 2. Parameter estimates for translog production frontiers of dairy farms in Argentina, Chile and Uruguay[a]

	Argentina		Chile		Uruguay	
Const.	0.012	*0.034*[b]	0.173**	*0.081*	0.268***	*0.048*
CO	0.774***	*0.055*	0.488***	*0.096*	0.554***	*0.084*
LB	0.193***	*0.047*	0.195*	*0.150*	0.103	*0.093*
FD	0.115***	*0.042*	0.289***	*0.055*	0.232***	*0.059*
VE	0.149***	*0.045*	0.128***	*0.042*	0.179***	*0.052*
T	0.009	*0.012*	0.026	*0.026*	0.069***	*0.022*
CO^2	0.116	*0.146*	0.031	*0.177*	-0.307	*0.226*
LB^2	0.144	*0.113*	0.176	*0.171*	0.017	*0.294*
FD^2	0.134**	*0.058*	0.015**	*0.007*	0.074	*0.085*
VE^2	0.003	*0.007*	0.005	*0.005*	-0.118**	*0.057*
t^2	0.037***	*0.008*	-0.057***	*0.017*	0.005	*0.021*
CO×LB	-0.334*	*0.201*	-0.362	*0.342*	0.364	*0.422*
CO×FD	-0.061	*0.088*	-0.031	*0.126*	0.076	*0.202*
CO×VE	-0.024	*0.123*	0.017	*0.100*	0.230	*0.218*
CO×t	-0.057	*0.039*	-0.014	*0.077*	0.103	*0.104*
LB×FD	-0.030	*0.095*	-0.056	*0.161*	-0.222	*0.253*
LB×VE	0.043	*0.138*	0.089	*0.145*	0.012	*0.231*
LB×t	0.033	*0.040*	-0.059	*0.092*	-0.157	*0.108*
LB×VE	-0.025	*0.082*	0.038	*0.057*	-0.103	*0.136*
FD×t	0.008	*0.026*	0.012	*0.040*	0.067	*0.054*
VE×t	0.073***	*0.023*	0.029	*0.039*	-0.054	*0.064*
FC	1.231		1.100		1.068	
LLF	67.193		0.288		54.194	
σ^2	0.036***	*0.011*	0.092***	*0.028*	0.090***	*0.028*
γ	0.860***	*0.064*	0.512***	*0.201*	0.844***	*0.073*

*10% of the level of significance. ** 5% of the level of significance. *** 1% of the level of significance.
[a] Argentina: 46 farms and 82 observations; Chile: 48 farms and 92 observations; Uruguay: 70 farms and 147 observations; [b] Standard error in italics.
CO: number of cows; LB: labor; FD: feed cost; VE: veterinary input costs; t: time; FC: function coefficient; LLF: log-likelihood; σ^2: sigma squared; γ: gamma parameter.

which was negative and significant and the parameters for the interaction between *t* and the other inputs that were not significant. The results in Table 2 also show that the first order parameters for the Uruguayan model were highly significant, with the exception of LB. The t^2 parameter was not significant and, as in the Chilean case, none of the parameters for the interaction between *t* and the other inputs were significant.

The function coefficient is the indicator commonly used to measure economies of size (EOS) in primal models, such as the SPF being used here. For the model presented in Equation [4] the function coefficient is equal to:

$$\mathrm{EOS} = \sum_{k=1}^{4} \beta_k + \sum_{k=1}^{4} \beta_{kl} \times x_{lit} + \sum_{k=1}^{4} \delta_k \times t \qquad [5]$$

Given that the data is normalized by the geometric mean, the function coefficient calculated at this mean is equal to the sum of the linear parameters ($\mathrm{EOS} = \sum_{k=1}^{4} \beta_k = \beta_{CO} + \beta_{LB} + \beta_{FD} + \beta_{VE}$). Consequently, the samples for the three countries reveal the presence of EOS, the largest being for Argentina (1.231), followed by Chile (1.100) and then by Uruguay (1.068). To obtain a broader vision of the behavior of EOS, Table 3 shows the function coefficient for four groups of farms of different size. The groups were defined according to the number of cows per quartile for each country. For Chile and Argentina, EOS decreased monotonically as the average number of cows rose. The data suggests that the average cost curve for Argentina is L-shaped, while that of Chile is U-shaped. In the case of Uruguay, EOS increases monotonically with the number of cows and this insinuates an average cost curve with a negative slope.

Another important dimension of the productive structure, which can be analyzed, based on the econometric estimations of a production frontier with panel data, is TC. For the TL model in Equation [4] the expression for TC is equal to:

$$TC = \sum_{k=1}^{4} \delta_k x_{kit} + \lambda_1 + \lambda_{11} t \qquad [6]$$

The annual rate of TC, at the geometric mean, is equal to the linear coefficient of the tendency variable (λ_1), which in this case is 0.9, 2.6 and 6.9% for Argentina, Chile and Uruguay, respectively. To examine the variation over time, Table 4 provides the annual rate of TC for each country according to the available information. It can be observed that for Argentina there is a very low and negative annual rate between 1997-1998 and 1998-1999 (-1.1%), which then rose between 1999-2000 and 2000-2001 (2.1%). In the case of Chile, there was a clear deterioration between 1996-1997 and 1997-1998 (5.5%) and 2000-2001 and 2001-2002 (-4.4%). In the case of Uruguay, an important TC rate was observed during the period under analysis, which experienced a reduction between 1999-2000 and 2000-2001 (11.9%) and 2000-2001 and 2001-2002 (8.0%), and then an increase in 2001-2002 and 2002-2003 (13.6%).

Table 5 shows descriptive statistics for the TE measures for each country. The average TE for Argentina was 87.0% with a minimum of 69.1% and a maximum of 97.9%. The average in the case of Chile was 84.9% with extremes of 64.4 and 94.8%. The statistics for Uruguay show an average of 81.1% with a variation between 49.3 and 97.1%. Several studies of dairy

Table 3. Economies of size (EOS) of Argentinean, Chilean and Uruguayan dairy farms.

	Argentina	Chile	Uruguay
		Nº of cows	
First quartile	1.436	1.383	0.733
Second quartile	1.228	1.123	1.040
Third quartile	1.185	1.040	1.174
Fourth quartile	1.014	0.876	1.363
Geometric mean	**1.231**	**1.100**	**1.068**

Table 4. Technological change (TC) for Argentinean, Chilean and Uruguayan dairy farms.

	Argentina[a]	Chile[b]	Uruguay[c]
		%	
1996-1997 to 1997-1998	na[d]	5.5	na
1997-1998 to 1998-1999	-1.1	5.5[e]	na
1998-1999 to 1999-2000	-1.1[e]	3.1	na
1999-2000 to 2000-2001	2.1	-0.1	11.9
2000-2001 to 2001-2002	2.1[e]	-4.4	8.0
2001-2002 to 2002-2003	na	na	13.6
Geometric mean	**0.9**	**2.6**	**6.9**

[a] Base year for Argentina is 1997-1998; [b] Base year for Chile is 1996-1997; [c] Base year for Uruguay is 1999-2000; [d] na: not available; [e] The results of the previously period were used because of lack of information.

Table 5. Technical efficiency (TE) of Argentinean, Chilean and Uruguayan dairy farms.

	Average	Standard deviation	Minimum	Maximum
		%		
Argentina	87.0	7.6	69.1	97.9
Chile	84.9	6.7	64.4	94.8
Uruguay	81.1	10.9	49.3	97.1

farms that use the stochastic frontier model have TEs close to the results of this research, as was reported by Moreira (2006).

It is important to note that the different indicators of productivity reported in this paper were calculated individually for each country and their respective production frontiers, so they are not directly comparable. In order to compare these indicators it is necessary to formally analyze whether the samples have access to the same level of technology. This idea is the basis of the meta-frontier production function, an approach recently developed by Battese *et al.* (2004) and refined by O'Donnell *et al.* (2008).

CONCLUSIONS

This study used unbalanced panel data for dairy farms from Argentina, Chile and Uruguay to estimate stochastic production frontiers. These frontiers were used to evaluate economies of size (EOS), rates of technological change (TC) and technical efficiency (TE). In the preliminary analysis it was shown that the translog functional form (TL) is superior to the Cobb-Douglas. The term that captures TE follows a half-normal distribution and is statistically significant and constant over time, presenting mean values of 87.0, 84.9 and 81.1% for Argentina, Chile and Uruguay, respectively. This result makes evident that the dairy farmers included in the sample of the three countries could increase their milk production by 13.0, 15.1 and 18.9%, respectively, without increasing the use of inputs. Average TC was 0.9% for Argentina, 2.6% for Chile and 6.9% for Uruguay. Finally, increasing EOS were found, which implies that the farms in the sample operate at a sub-optimal size.

RESUMEN

Cambio tecnológico y eficiencia técnica en predios lecheros de tres países de Sudamérica. Boris E. Bravo-Ureta[1], Víctor H. Moreira[2]*, Amilcar A. Arzubi[3], Ernesto D. Schilder[4], Jorge Álvarez[5], y Carlos Molina[5]. La liberalización progresiva de los mercados agrícolas, junto a la amenaza que implica para productores nacionales la competencia de productos importados, dejan en claro la relevancia del incremento en la productividad como un elemento para mejorar la competitividad. La medición de la eficiencia técnica es uno de los componentes de la productividad más estudiados, debido a que proporciona información valiosa al momento de formular políticas y tomar decisiones destinadas a mejorar la administración predial. Este trabajo utiliza datos de panel desbalanceados de predios lecheros provenientes de Argentina, Chile y Uruguay, para estimar fronteras estocásticas de producción. Luego estas fronteras se usaron para analizar economías de tamaño, tasas de cambio tecnológico y eficiencia técnica. En todas las estimaciones se empleó el modelo de Battese y Coelli (1992), ampliamente usado en la literatura de productividad. Los modelos para los tres países exhiben economías de escala crecientes, lo que implica que los predios en las muestras operan a un tamaño sub-óptimo. La tasa promedio anual de cambio tecnológico fue 0,9% para Argentina, 2,6% para Chile y 6,9% para Uruguay mientras que la eficiencia técnica media fue igual a 87,0%, 84,9% y 81,1%, respectivamente.

Palabras clave: fronteras estocásticas, predios lecheros, Argentina, Chile, Uruguay.

LITERATURE CITED

ABARE. 2006. Agricultural economies of Australia & New Zealand, past, present, future. Australian Bureau of Agricultural and Resource Economics. New Zealand Ministry of Agriculture and Forestry. Available at http://www.maf.govt.nz/mafnet/rural-nz/statistics-and-forecasts/iaae-australia-nz-agriculture/ (Accessed March 2007).

Ahmad, M., and B.E. Bravo-Ureta. 1996. Technical efficiency measures for dairy farms using panel data: A comparison of alternative model specifications. J. Prod. Anal. 7:399-415.

Arias, C., y A. Álvarez. 1993. Estimación de eficiencia técnica en explotaciones lecheras con datos de panel. Invest. Agrar. Econ. 8:101-109.

Battese, G.E., and T.J. Coelli. 1988. Prediction of firm-level technical efficiencies with a generalized frontier production function and panel data. J. Econometrics 38:387-399.

Battese, G.E., and T.J. Coelli. 1992. Frontier production functions, technical efficiency and panel data: With application to paddy farmers in India. J. Prod. Anal. 3:153-169.

Battese, G.E., and T.J. Coelli. 1995. A model for technical inefficiency effects in a stochastic frontier production function for panel data. Empir. Econ. 20:325-332.

Battese, G.E., and G.S. Corra. 1977. Estimation of a production frontier model: With application to the pastoral zone of Eastern Australia. Aust. J. Agric. Econ. 21:169-179.

Battese, G.E., D.S.P. Rao, and C.J. O'Donnell. 2004. A metafrontier production function for estimation of technical efficiencies and technology gaps for firms operating under different technologies. J. Prod. Anal. 21:91-103.

Blayney, D.P., and M.J. Gehlhar. 2005. U.S. Dairy at a new crossroads in a global setting. Amber Waves 3:1-6.

Bravo-Ureta, B.E., D. Solís, V.H. Moreira, J.F. Maripani, A. Thiam, and T.E. Rivas. 2007. Technical efficiency in farming: A meta-regression analysis. J. Prod. Anal. 27:57-72.

Coelli, T.J. 1996. A guide to Frontier version 4.1: A computer program for stochastic frontier production and cost function estimation. Australia: CEPA Working Paper 96/7, Department of Econometrics, University of New England, Armidale, New South Wales, Australia.

Coelli, T.J., D.S.P. Rao, and G.E. Battese. 1998. An introduction to efficiency and productivity analysis. 275 p. Springer, New York, USA.

Coelli, T., D.S.P. Rao, C. O'Donnell, and G. Battese. 2005. An introduction to efficiency and productivity analysis. 349 p. 2nd ed. Springer, New York, USA.

Cuesta, R.A. 2000. A production model with firm-specific temporal variation in technical inefficiency: With application to Spanish dairy farms. J. Prod. Anal. 13:139-158.

European Commission. 2005. Prospects for agricultural markets and income: 2005-2012. European Commission Directorate-General for Agriculture. Available at http://ec.europa.eu/agriculture/publi/caprep/prospects2005/index_en.htm (Accessed March 2007).

Evans, L., A. Grimes, B. Wilkinson, and D. Teece. 1996. Economic reform in New Zealand 1984-95: The pursuit of efficiency. J. Econ. Lit. 34:1856-1902.

FAO. 2003. World agriculture: Towards 2015/2030 an FAO perspective. FAO Corporate document repository. Available at http://www.fao.org/docrep/005/y4252e/y4252e00.htm (Accessed April 2007).

Greene, W.H. 1993. The econometric approach to efficiency analysis. p. 67-119. *In* Fried, H.O., C.A.K. Lovell, and S.S. Schmidt (eds.) The measurement of productive efficiency. Oxford University Press, New York, USA.

Hanrahan, C., and R. Schnepf. 2005. WTO Doha round: Agricultural negotiating proposals. The Library of Congress. Available at http://fpc.state.gov/documents/organization/57217.pdf (Accessed December 2006).

Heshmati, A., and S.C. Kumbhakar. 1994. Farm heterogeneity and technical efficiency: Some results from Swedish dairy farms. J. Prod. Anal. 5:45-61.

Jaforullah, M., and N.J. Devlin. 1996. Technical efficiency in the New Zealand dairy industry: A frontier production function approach. N.Z. Econ. Pap. 30:1-17.

Jondrow, J., C.A.K. Lovell, I.S. Materov, and P. Schmidt. 1982. On the estimation of technical inefficiency in the stochastic frontier production function model. J. Econometrics 19:233-238.

Kalirajan, K.P. 1984. Farm-specific technical efficiencies and development policies. J. Econ. Stud. 11:3-13.

Kumbhakar, S.C., S. Ghosh, and J.T. McGuckin. 1991. A generalized production frontier approach for estimating determinants of inefficiency in U.S. Dairy farms. J. Bus. Econ. Stat. 9:279-286.

Lawson, L.G., J.F. Agger, M. Lund, and T.J. Coelli. 2004a. Lameness, metabolic and digestive disorders, and technical efficiency in Danish dairy herds: A stochastic frontier production function approach. Livest. Prod. Sci. 91:157-172.

Lawson, L.G., J. Bruun, T.J. Coelli, J.F. Agger, and M. Lund. 2004b. Relationships of efficiency to reproductive disorders in Danish milk production: A stochastic frontier analysis. J. Dairy Sci. 87:212-224.

Moreira, V.H. 2006. Dairy farm technical efficiency and milk total factor productivity growth in three South American countries. 162 p. Ph.D. Dissertation. University of Connecticut, Storrs, Connecticut, USA.

Moreira, V.H., B.E. Bravo-Ureta, B.L. Carrillo, and J.A. Vásquez. 2006. Technical efficiency measures for small dairy farms in Southern Chile: A stochastic frontier analysis with unbalanced panel data. Arch. Med. Vet. 38:25-32.

O'Donnell, C.J., D.S.P. Rao, and G.E. Battese. 2008. Metafrontier frameworks for the study of firm-level efficiencies and technology ratios. Empir. Econ. 34:231-255.

Paul, C.J.M., W.E. Johnston, and G.A.G. Frengley. 2000. Efficiency in New Zealand sheep and beef farming: The impacts of regulatory reform. Rev. Econ. Stat. 82:325-337.

Pinstrup-Andersen, P. 2002. Food and agricultural policy for a globalizing world: Preparing for the future. Am. J. Agric. Econ. 84:1201-1214.

Russell, N.P., and T. Young. 1983. Frontier production functions and the measurement of technical efficiency. J. Agric. Econ. 34:139-150.

Ruttan, V.W. 2002. Productivity growth in world agriculture: Sources and constraints. J. Econ. Perspect. 16:161-184.

Sandrey, R.A., and G.M. Scobie. 1994. Changing international competitiveness and trade: Recent experience in New Zealand agriculture. Am. J. Agr. Econ. 76:1041-1046.

Sena, V. 1999. Stochastic frontier estimation: A review of the software options. J. Appl. Econ. 14:579-586.