TECHNICAL, ECONOMIC, AND ALLOCATIVE EFFICIENCY IN PEASANT FARMING: EVIDENCE FROM THE DOMINICAN REPUBLIC

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I. INTRODUCTION

THE agricultural sector has always been an important component of the Dominican Republic's economy. During the 1980s, agriculture accounted for as much as 20 per cent of the gross domestic product (GDP), 45 per cent of overall employment, and over 50 per cent of total foreign exchange earnings. By contrast, during this period the manufacturing sector employed 20 per cent of the labor force while contributing 15 per cent of GDP (U.S. Department of Agriculture 1987; World Bank 1990).

A significant feature of Dominican Republic agriculture is the dual structure of the land tenure system. This dual structure consists of large latifundios that produce traditional export crops (i.e., sugarcane and tobacco), and small peasant farms that typically produce for subsistence and for local markets (i.e., beans, yucca, manioc, corn, and rice). Data from the 1981 agricultural census reveals that 82 per cent of the farms had less than five hectares, while the largest 1 per cent had more than two hundred hectares and occupied 36 per cent of the total arable land (Greene and Roe 1990). Most of the remaining land was owned by the Dominican Agrarian Institute, which is the government agency charged with land redistribution. This sharp concentration in landownership still remains, despite the implementation of an agrarian reform process that began around 1962 and since 1976 has given the right to farm to approximately 17 per cent of a total of 400,000 landless families (Kaiser 1988).

A more recent feature of agricultural development in the Dominican Republic

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has been the expansion of nontraditional crops.[1] Income from these crops increased by 87 per cent over the period 1976–84, and over the last decade, this type of farm output has become an increasingly important component of agricultural exports (Battese 1992; CEDOPEX 1985). The private agribusiness support of peasant farms and recent policy efforts suggest that small operations can indeed benefit from participating in nontraditional crop production and marketing. Williams and Karen (1985) concluded that the introduction of nontraditional crops in some areas of the Dominican Republic resulted in a 38 per cent increase in employment, and that the per capita income among small-scale farmers in these areas was three times higher than the national average in 1983.

It is important to emphasize that despite the potential benefits stemming from the expansion of the nontraditional crop sector, the overall productivity of Dominican agriculture remains low and the presence of the latifundio sector continues to be significant. The poor performance of agriculture is most clearly evidenced by much lower standards of living in rural areas compared to urban areas; thus, the largest concentration of absolute poverty, illiteracy, and infant mortality is found in the countryside (Quezada 1981).

There is considerable agreement with the notion that an effective economic development strategy depends critically on promoting productivity and output growth in the agricultural sector, particularly among small-scale producers.[2] Empirical evidence suggests that small farms are desirable not only because they provide a source of reducing unemployment, but also because they provide a more equitable distribution of income as well as an effective demand structure for other sectors of the economy (Bravo-Ureta and Evenson 1994; Dorner 1975). Consequently, many researchers and policymakers have focused their attention on the impact that the adoption of new technologies can have on increasing farm productivity and income (Hayami and Ruttan 1985; Kuznets 1966; Schultz 1964; Seligson 1982). However, during the last decade, major technological gains stemming from the green revolution seem to have been largely exhausted across the developing world. This suggests that attention to productivity gains arising from a more efficient use of existing technology is justified (Bravo-Ureta and Pinheiro 1993; Squires and Tabor 1991).

The presence of shortfalls in efficiency means that output can be increased without requiring additional conventional inputs and without the need for new technology. If this is the case, then empirical measures of efficiency are necessary in order to determine the magnitude of the gains that could be obtained by improving per-

[1] Nontraditional crops consist primarily of export crops, excluding sugar, coffee, and tobacco, as well as other cash crops used in the domestic markets, such as rice, corn, pigeon beans, tomatoes, citrus, yucca, and Chinese vegetables (Oberg et al. 1985).

[2] For an earlier analysis of the role of agriculture in economic development, see Johnston and Mellor (1961). For a more recent discussion, see Hayami and Ruttan (1985).

formance in agricultural production with a given technology. An important policy implication stemming from significant levels of inefficiency is that it might be more cost effective to achieve short-run increases in farm output, and thus income, by concentrating on improving efficiency rather than on the introduction of new technologies (Belbase and Grabowski 1985; Shapiro and Müller 1977).

The general objective of this paper is to assess the possibilities for productivity gains by improving the efficiency of small-scale agriculture in the Dajabon region of the Dominican Republic. This objective is pursued first by estimating a stochastic production frontier which provides the basis for measuring farm-level technical (*TE*), economic (*EE*), and allocative (*AE*) efficiency. Then, a second step analysis (Bravo-Ureta and Pinheiro 1993; Lingard, Castillo, and Jayasuriya 1983) is performed where separate two-limit tobit equations for *TE*, *EE*, and *AE* are estimated as a function of various attributes of the farms/farmers in the sample. This study has policy implications because it not only provides empirical measures of different efficiency indices, but also identifies some key variables that are correlated with these indices. In this fashion, we go beyond much of the published literature concerning efficiency because most research in this area of productivity analysis focuses exclusively on the measurement of technical efficiency (Bravo-Ureta and Pinheiro 1993).

The following is organized into four sections. In Section II, we discuss the methodological framework used in the analysis. Section III presents a brief description of the data and of the empirical frontier production model used as the basis for measuring farm-level efficiency. In Section IV, we present the empirical measures of efficiency along with the results of the two-limit tobit equations. The paper ends with a section containing some concluding remarks.

II. METHODOLOGICAL FRAMEWORK

Almost forty years ago, Michael Farrell (1957) introduced a methodology to measure economic, technical, and allocative efficiency. In this methodology, *EE* is equal to the product of *TE* and *AE*. According to Farrell, *TE* is associated with the ability to produce on the frontier isoquant, while *AE* refers to the ability to produce at a given level of output using the cost-minimizing input ratios (see Figure 1). Alternatively, technical inefficiency is related to deviations from the frontier isoquant, and allocative inefficiency reflects deviations from the minimum cost input ratios. Thus, *EE* is defined as the capacity of a firm to produce a predetermined quantity of output at minimum cost for a given level of technology (Farrell 1957; Kopp and Diewert 1982).

Over the last three decades, Farrell's methodology has been applied widely, while undergoing many refinements and improvements. The model used in this paper is based on an extension advanced by Kopp and Diewert (1982) and further

Fig. 1. Graphical Representation of Observed, Technically, and Economically Efficient Cost Measures
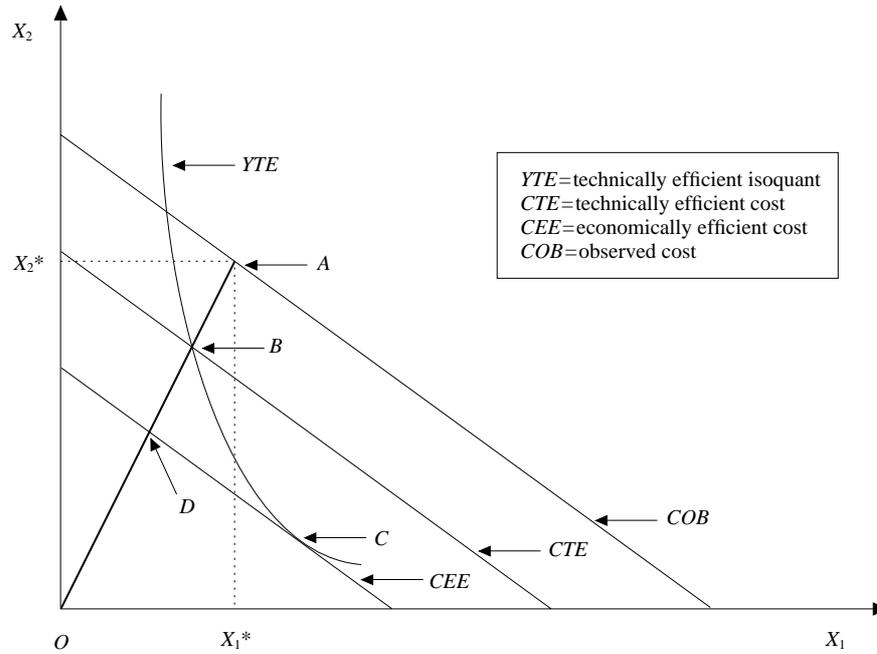
Notes: 1. According to equation (2) k_2, ($i=2$) in this figure, is equal to OX_1^*/OX_2^*.
2. According to equations (5), (6), and (7) TE, EE, and AE are equal to: $TE=OB/OA=CTE/COB$, $AE=OD/OB=CEE/CTE$, and $EE=TE\cdot AE=OD/OA=CEE/COB$.

modified by Bravo-Ureta and Rieger (1990). To begin with, assume that a deterministic production frontier is given by the equation

$$Y_j = g(X_{ij}; \beta), \qquad (1)$$

where Y_j is the output of the jth farm, X_{ij} is the ith input used by farm j, and β is a vector of unknown parameters. To simplify the exposition, the subscript j is dropped in what follows. From equation (1), it is possible to derive the technically efficient input quantities (X_{it}) for any given level of output $\bar{Y}$, by solving simultaneously the following equations:

$$\bar{Y} = g(X_i; \beta), \qquad (2)$$
$$X_1/X_i = k_i,$$

where k_i is the ratio of the observed level of inputs X_1 and X_i ($i > 1$) at output $\bar{Y}$.

Next, assume that the production frontier in equation (1) is self-dual (e.g., Cobb-Douglas) and that the corresponding cost frontier can be expressed as

$$C = h(P, Y; \alpha), \tag{3}$$

where C is the minimum cost to produce output Y, P is a vector of input prices, and α is a vector of parameters. Applying Shephard's lemma, the system of minimum cost input demand equations can be obtained by differentiating the cost frontier with respect to each input price. This demand equation for the ith input (X_{di}) is equal to

$$\partial C/\partial P_i = X_{di} = f(P, Y; \phi), \tag{4}$$

where ϕ is a vector of parameters. From the input demand equations we can obtain the economically efficient input quantities, X_{ie}, by substituting the firm's input prices P and output quantity $\bar{Y}$ into equation (4).

Thus far we have solved for the input bundles X_i, X_{it}, and X_{ie}. It is now possible to calculate the cost of the actual or observed (*COB*) input bundle as $\sum_i X_i\ P_i$, while the cost of the technically (*CTE*) and economically efficient (*CEE*) input combinations associated with the firm's observed output are given by $\sum_i X_{it}\ P_i$ and $\sum_i X_{ie}\ P_i$, respectively. These cost measures are the basis for calculating *TE* and *EE* as follows:

$$TE = \sum_i X_{it}\ P_i\ /\ \sum_i X_i\ P_i = CTE\ /\ COB, \tag{5}$$

and

$$EE = \sum_i X_{ie}\ P_i\ /\ \sum_i X_i\ P_i = CEE\ /\ COB. \tag{6}$$

As already mentioned, in the Farrell (1957) methodology, *EE* is equal to the product of *TE* and *AE*; hence, equations (5) and (6) are used to calculate *AE* as:

$$AE = EE\ /\ TE = \sum_i X_{ie}\ P_i\ /\ \sum_i X_{it}\ P_i = CEE\ /\ CTE. \tag{7}$$

A graphical illustration of the various cost and efficiency measures is presented in Figure 1. For simplicity, output Y is expressed in terms of two inputs, X_1 and X_2. Let point A be the combination of observed levels of inputs X_1 and X_2 (i.e., X_1^* and X_2^*) yielding observed output Y_o for a given firm, and isoquant *YTE* be the technically efficient or frontier isoquant associated with output Y_o. The isocost lines *COB*, *CTE*, and *CEE* reflect observed, technically efficient, and economically efficient cost levels while points B and C are, respectively, the *TE* and *EE* combinations of inputs X_1 and X_2 associated with output Y_o. The ratio k_i from equation (2) is equal to OX_1^*/OX_2^*. This ratio insures that the technically efficient input quantities are on the same ray out of the origin as the observed quantities, which is consistent with Farrell's definition of *TE*.

The Kopp and Diewert (1982) approach is based on a deterministic frontier

which imposes a limiting assumption that the entire deviation from the frontier is due to inefficiency. Schmidt (1985–86), among others, has argued that efficiency measures obtained from deterministic models are affected by statistical noise. For this reason, Bravo-Ureta and Rieger (1991) used a stochastic production frontier in order to remove the random element from the efficiency component before deriving the various efficiency indices.

The stochastic production frontier can be written as

$$Y = f(X_i; \beta) + \varepsilon, \tag{8}$$

where Y, X_i, and β are as defined earlier. The essential idea behind the stochastic frontier model is that ε is a "composed" error term (Aigner, Lovell, and Schmidt 1977; Meeusen and Van den Broeck 1977). This term can be written as

$$\varepsilon = v - u, \tag{9}$$

where v is a two-sided ($-\infty < v < \infty$) normally distributed random error (v $N[0, \sigma_v^2]$) that captures the stochastic effects outside the farmer's control (e.g., weather, natural disasters, and luck), measurement errors, and other statistical noise. The term u is a one-sided ($u \geq 0$) efficiency component that captures the technical inefficiency of the farmer. In other words, u measures the shortfall in output Y from its maximum value given by the stochastic frontier $f(X_i; \beta) + v$. This one-sided term can follow such distributions as half-normal, exponential, and gamma (Aigner, Lovell, and Schmidt 1977; Greene 1980; Meeusen and Van den Broeck 1977). In this paper, it is assumed that u follows a half-normal distribution (u $N[0, \sigma_u^2]$) as typically done in the applied stochastic frontier literature. The two components v and u are also assumed to be independent of each other.

The maximum likelihood estimation of equation (8) yields consistent estimators for β, λ, and σ^2, where β is a vector of unknown parameters, $\lambda = \sigma_u / \sigma_v$, and $\sigma^2 = \sigma_u^2 + \sigma_v^2$. Jondrow et al. (1982) have shown that inferences about the technical inefficiency of individual farmers can be made by considering the conditional distribution of u given the fitted values of ε and the respective parameters. In other words, given the distribution assumed for v and u, and assuming that these two components are independent from each other, the conditional mean of u given ε is defined by

$$E(u_j \,|\, \varepsilon_j) = \sigma_* \left[\frac{f^*(\varepsilon_j \lambda / \sigma)}{1 - F^*(\varepsilon_j \lambda / \sigma)} - \frac{\varepsilon_j \lambda}{\sigma} \right], \tag{10}$$

where $\sigma_*^2 = \sigma_u^2 \sigma_v^2 / \sigma^2$, f^* is the standard normal density function, and F^* is the distribution function, both functions being evaluated at $\lambda \varepsilon / \sigma$.

Consequently, by replacing ε, σ_*, and λ by their estimates in equations (8) and (10), we derive the estimates for v and u. Subtracting v from both sides of equation (8) yields the stochastic production frontier

$$Y^* = f(X_i; \beta) - u = Y - v, \qquad\qquad\qquad (11)$$

where Y^* is defined as the farm's observed output adjusted for the statistical noise contained in v (Bravo-Ureta and Rieger 1991). Equation (11) is used to compute X_{it} as well as to derive the cost frontier. As described at the beginning of this section, the cost frontier is then used to obtain the minimum cost factor demand equations, which, in turn, become the basis for calculating the economically efficient input levels X_{ie}.

III.   THE DATA AND THE EMPIRICAL PRODUCTION FRONTIER MODEL

The data used in the econometric analysis reported in this paper comes from a sample of small farms located in the Dajabon region. The Dajabon region, situated in the northwest corner of the Dominican Republic, was chosen for this study because it is one of the country's poorest areas, where agriculture plays a key role in the local economy. Nontraditional agriculture has become increasingly prominent in this area, since it supplies local markets, as well as markets across the Haitian border and overseas. Most importantly, nontraditional crops are the primary source of cash income for most of the small farmers in the region. Dajabon is further characterized by increasing population pressure on the land, low levels of investment in agriculture, and low average per capita income. During the last decade, however, the Dominican Agrarian Institute has focused on increasing the output of nontraditional crops in this region, due to the growing importance of markets for these products (Pinheiro 1992).

During the spring of 1988, a survey was conducted on a random sample of small farms located in Dajabon. After discarding a few incomplete records, a sample of sixty farms remained for analysis. Farm size in the sample ranges from 8 to 185 *tareas* with a mean of 43.13 *tareas*.[3] Of the sixty farms, forty are private farms and the other twenty are farm units that were created in the 1970s as a result of the agrarian reform process. Farmers that are not beneficiaries of the Agrarian Reform Program either rent or own their land. In contrast, the agrarian reform units are managed by the individual farm families, but they do not own the land.

At the time of the study, 2,498 of the total 2,588 *tareas* making up the sixty farms were under cultivation. The production of corn and yucca took up 42 per cent of the cultivated land. Rice, manioc, *gandules* (pigeon beans), and red beans were the main crops occupying the remainder of the arable land. The production of bananas, potatoes, Chinese eggplant, bitter melon, and pineapple, although common, were less significant.

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[3] 1 acre = 6.15 *tareas*; 1 *tarea* = 0.162 acres; 1 *tarea* = 0.0648 hectares.

TABLE I

ORDINARY LEAST SQUARES (OLS) AND MAXIMUM LIKELIHOOD (ML) PARAMETER ESTIMATES
BASED ON A SAMPLE OF SIXTY SMALL FARMERS FROM DAJABON

Variable	Mean (Std. Dev.)	OLS Estimates (Std. Error)	ML Estimates (Asymp. Std. Error)
Intercept	—	4.867***	5.148***
		(0.516)	(0.641)
Land (X_1)	47.80	0.359***	0.357***
	(53.56)	(0.119)	(0.117)
Labor (X_2)	729.90	0.159**	0.156**
	(801.70)	(0.079)	(0.079)
Fertilizer (X_3)	7.70	0.059***	0.057***
	(19.90)	(0.018)	(0.017)
Tools (X_4)	30.60	0.037***	0.037***
	(34.70)	(0.013)	(0.013)
Seeds and draft	887.50	0.164**	0.169**
power (X_5)	(1,625.70)	(0.071)	(0.069)
Function coefficient	—	0.778	0.777
F-statistic model	—	29.38***	—
F-statistic CRTS[a]	—	6.22**	—
Adjusted R^2	—	0.71	—
λ	—	—	0.975
σ^2	—	—	0.517**
			(0.221)
Log likelihood	—	—	34.36

[a] CRTS = constant returns to size.
*** Significant at the 0.01 level.
** Significant at the 0.05 level.

The model chosen to perform the efficiency analysis can be expressed in general form as

$$Y = f(X_1, X_2, X_3, X_4, X_5), \tag{12}$$

where Y is output and the X's are inputs. A more detailed definition of these variables is given below, while descriptive statistics are presented in Table I.

The output variable in equation (12), Y, is the farm value of all crops produced. The variable X_1 includes all cultivated land, X_2 includes family and hired labor measured in worker-days, X_3 represents fertilizer measured in 100 pound units (cwt), X_4 corresponds to total expenditures on small farm tools for the year, and X_5 is the value of seed and draft power used in the production process. The explanatory variables included in this model have been commonly used in estimating agricultural production frontiers for developing countries (Belbase and Grabowski 1985; Kalirajan 1981, 1984; Kalirajan and Flinn 1983; Kalirajan and Shand 1985; Phillips and Marble 1986; Rawlins 1985; Taylor and Shonkwiler 1986; Taylor, Drummond, and Gomes 1986).

The Cobb-Douglas functional form is used to specify the stochastic production frontier, which is the basis for deriving the cost frontier and the related efficiency measures.[4] Despite its well-known limitations, the Cobb-Douglas is chosen because the methodology employed requires that the production function be self-dual. It is also worth stating that this functional form has been widely used in farm efficiency analyses for both developing and developed countries.[5] The specific model estimated is given by

$$\ln Y = \ln A + \sum_{i=1}^{5} \beta_i \ln X_i + \varepsilon, \tag{13}$$

where A and β_i are parameters to be estimated ($i = 1, \ldots, 5$), ε is the composed error term, and Y and the X_i variables are as defined earlier.

IV. EMPIRICAL RESULTS

Based on the model discussed in the previous section, Table I presents ordinary least square (OLS) and maximum likelihood (ML) estimates of the production function parameters. The OLS function provides estimates of the "average" production function, while the ML model yields estimates of the stochastic production frontier. The similarities of the slope parameters across equations confirm that the frontier function represents a neutral upward shift of the average production function. These results are consistent with the findings of Bravo-Ureta and Evenson (1994) and Bravo-Ureta and Rieger (1990). Moreover, all parameter estimates are statistically significant at the 1 per cent level for the two models with the exception of the parameters for labor (X_2), and seeds and draft power (X_5), both of which are significant at the 5 per cent level.

The function coefficient, which measures the proportional change in output when all inputs included in the model are changed in the same proportion, is given in Table I for the two equations. The function coefficient for both the OLS and ML estimates is approximately 0.78, which indicates that returns to size are decreasing. Restricted least squares regression was used to formally test the null hypothesis of constant returns to size. The computed F statistic is 6.22, which exceeds the critical F value of 4.02 at the 5 per cent level of significance. Consequently, the null hypothesis of constant returns to size was rejected.

[4] The use of a single equation model is justified by assuming that farmers maximize expected profits (Bravo-Ureta and Rieger 1990; Caves and Barton 1990; Kopp and Smith 1980; Zellner, Kmenta, and Drèze 1966).

[5] Support for this statement can be found in the reviews of the empirical literature recently completed by Battese (1992), and by Bravo-Ureta and Pinheiro (1993). Moreover, recent work suggests that the choice of functional form might not have a significant impact on measured efficiency levels (Ahmad and Bravo-Ureta 1996; Good et al. 1993).

TABLE II

FREQUENCY DISTRIBUTION OF TECHNICAL, ALLOCATIVE, AND ECONOMIC EFFICIENCY

Efficiency (%)	Technical Efficiency		Allocative Efficiency		Economic Efficiency	
	No.[a]	%[b]	No.[a]	%[b]	No.[a]	%[b]
> 85	0	0	0	0	0	0
> 80 ≤ 85	9	15	3	5	0	0
> 75 ≤ 80	11	18	4	7	0	0
> 70 ≤ 75	11	19	4	7	0	0
> 65 ≤ 70	13	22	2	3	0	0
> 60 ≤ 65	6	10	2	3	3	5
> 55 ≤ 60	4	6	3	5	2	3
> 50 ≤ 55	4	6	8	13	4	7
> 45 ≤ 50	1	2	4	7	3	5
> 40 ≤ 45	1	2	3	5	5	8
> 35 ≤ 40	0	0	6	10	8	13
> 30 ≤ 35	0	0	4	7	6	10
> 25 ≤ 30	0	0	6	10	5	8
> 20 ≤ 25	0	0	2	3	10	17
> 15 ≤ 20	0	0	4	6	4	7
> 10 ≤ 15	0	0	4	6	6	10
≤ 10	0	0	1	2	4	7
Mean (%)	70.0		44.0		31.0	
Minimum (%)	42.0		9.5		5.3	
Maximum (%)	85.0		84.0		62.0	

[a] The number of farms.
[b] The percentage (rounded) of total farms.

The ratio of the standard error of u (σ_u) to the standard error of v (σ_v), known as lambda (λ), is 0.97. Based on λ, we can derive gamma (γ) which measures the effect of technical inefficiency in the variation of observed output ($\gamma = \lambda^2 / [1 + \lambda^2] = \sigma_u^2 / \sigma_\varepsilon^2$). The estimated value of γ is 0.49, which means that 49 per cent of the total variation in farm output is due to technical inefficiency.

The cost frontier dual to the production frontier shown in Table I is

$$\ln C = 0.005 + 0.459 \ln P_1 + 0.200 \ln P_2 + 0.074 \ln P_3 + 0.048 \ln P_4 + 0.218 \ln P_5 + 1.286 \ln Y^*, \tag{14}$$

where C is the cost of crop production per farm measured in Dominican pesos (RD\$); P_1 is the average rent per *tarea* of land in the Dajabon region which was determined to be RD\$25.00; P_2 is the average daily wage rate which is estimated at RD\$6.00; P_3 is the average price of a 100 pound unit of fertilizer estimated at RD\$20.00; P_4 is the average price of farm tools set at RD\$1.10 for each RD\$1.00, implying an operating capital cost of 10 per cent per annum; P_5 represents the price of seeds and draft power set at RD\$1.05, reflecting the cost of capital for six

TABLE III

EMPIRICAL ESTIMATES OF EFFICIENCY FROM STOCHASTIC PRODUCTION FRONTIERS

Author	Country	Product	TE (%)	AE (%)	EE (%)
This Study	Dominican Republic	Crops	70	44	31
Bagi (1982)	India	Rice	93	—	—
Bravo-Ureta and Evenson (1994)	Paraguay	Cotton	58	70	40
		Cassava	59	88	52
Huang and Bagi (1984)	India	Whole farm	89	—	—
Hussain (1989)	Pakistan	Crops	69	43	29
Kalirajan (1981)	India	Rice	67	—	—
Kalirajan (1984)	Philippines	Rice	63	—	—
Kalirajan and Flinn (1983)	Philippines	Rice	80	—	—
Kalirajan and Flinn (1983)	Philippines	Rice	50	—	—
Kalirajan and Shand (1986)	Malaysia	Rice	67	—	—
Phillips and Marble (1986)	Guatemala	Maize	75	—	—
Rawlins (1985)	Jamaica	Crops	73	—	—
Taylor and Shonkwiler (1986)	Brazil	Whole farm	71	—	—

months; and Y^* is the total farm output measured in Dominican pesos and adjusted for any statistical noise as previously specified in equation (11).

The results derived from the econometric estimation of equation (13) indicate that technical efficiency (*TE*) indices range from 42 to 85 per cent for the farms in the sample, with an average of 70 per cent (Table II). This means that if the average farmer in the sample was to achieve the *TE* level of its most efficient counterpart, then the average farmer could realize an 18 per cent cost savings (i.e., $1 − [70/85]$). A similar calculation for the most technically inefficient farmer reveals cost savings of 50 per cent (i.e., $1 − [42/85]$). The mean allocative efficiency of the sample is 44 per cent, with a low of 9.5 and a high of 84. The combined effect of technical and allocative factors shows that the average economic efficiency level for this sample is only 31 per cent, with a low of 5.3 per cent and a high of 62 per cent. These figures indicate that if the average farmer in the sample were to reach the *EE* level of its most efficient counterpart, then the average farmer could experience a cost savings of 50 per cent (i.e., $1 − [31/62]$). The same computation for the most economically inefficient farmer suggests a gain in economic efficiency of 92 per cent (i.e., $1 − [5.3/62]$). In sum, it is evident from these results that *EE* could be improved substantially, and that allocative inefficiency constitutes a more serious problem than technical inefficiency.

To provide a basis of comparison for the efficiency measures just discussed, Table III presents average efficiency indices reported in other studies that have

estimated stochastic production frontiers using farm data from developing countries. As the data shows, the 70 per cent mean *TE* found in this study is in line with the findings reported by others. Table III also shows the few estimates of *AE* and *EE* that have been reported in the literature. The 44 per cent average *AE* found in this paper is very close to the 43 per cent figure reported by Hussain (1989) for a sample of wheat and maize farmers in Pakistan. By contrast, Bravo-Ureta and Evenson (1994), in their analysis of Paraguayan peasant farms, reported higher estimates of *AE* (70 and 88 per cent). In addition, the average *EE* level of 29 per cent reported by Hussain (1989) is very close to the 31 per cent average *EE* found in this paper. Again, the *EE* estimates reported by Bravo-Ureta and Evenson (1994) are higher than those found here.

As already established, *EE* is composed of *AE* and *TE*. Therefore, economic inefficiency arises from a combination of the technical and allocative components. For policy purposes, it is useful to identify the sources of these inefficiencies, which can be done by investigating the relationship between farm/farmer characteristics and the computed *TE* and *AE* indices separately. The association between *EE* and these same characteristics can also be ascertained directly. To delve deeper into this matter, and following what is known in the literature as "second step" estimation (Bravo-Ureta and Pinheiro 1993), the following models are estimated:

$$EFFIC = f(CONT, REF, SIZE, SCHOOL, AGE, PEOP), \qquad (15)$$

where *EFFIC* is, alternatively, the natural logarithm of farm-level *TE*, *AE*, or *EE*. All variables in equation (15), with the exception of *PEOP*, are dummy variables and are defined as follows: *CONT* is equal to one for those farmers producing any crop under contract with an agribusiness firm and zero otherwise; *REF* is equal to one if the individual producer is an agrarian reform beneficiary and zero otherwise; *SIZE* is equal to one for medium-size farms, which are those that have between 50 and 100 *tareas*, and zero otherwise; *SCHOOL* is equal to one if the producer has four or more years of schooling and zero otherwise; and *AGE* is equal to one for younger farmers, who are those less than twenty-five years of age, and zero otherwise. The last explanatory variable, *PEOP*, is the natural logarithm of the number of people in the household including the household head. The variables included in equation (15) are those usually incorporated in analyses of this type (Belbase and Grabowski 1985; Bravo-Ureta and Evenson 1994; Huang and Bagi 1984; Hussain 1989; Kalirajan and Flinn 1983; Squires and Tabor 1991).

The models for *TE*, *AE*, and *EE* in equation (15) are estimated separately using the two-limit tobit procedure, given that the efficiency indices are bounded between 0 and 100 (Greene 1991; Hossain 1988). The models were also estimated using the seemingly unrelated regression (SUR) procedure. However, to obtain statistical efficiency gains using the SUR procedure it is necessary to have different regressors in the various equations (Judge et al. 1988). Unfortunately, there is no

TABLE IV

Two-Limit Tobit Equations for Technical (*TE*), Allocative (*AE*), and Economic (*EE*)
Efficiency for a Sample of Sixty Small Farmers from Dajabon

Variable	Mean (Std. Dev.)	*TE* Parameter (Std. Error)	*AE* Parameter (Std. Error)	*EE* Parameter (Std. Error)
Intercept	—	4.145***	3.288**	2.828***
		(0.042)	(0.131)	(0.137)
CONT	0.566	0.047	0.566***	0.613***
	(0.50)	(0.033)	(0.103)	(0.108)
REF	0.333	0.040	0.230***	0.270**
	(0.48)	(0.036)	(0.112)	(0.117)
SIZE	0.200	0.024	0.516***	0.541***
	(0.40)	(0.042)	(0.131)	(0.137)
SCHOOL	0.700	0.072**	−0.114	−0.043
	(0.46)	(0.036)	(0.114)	(0.111)
AGE	0.017	0.206*	0.695*	0.901**
	(0.13)	(0.128)	(0.405)	(0.422)
PEOP	1.184	0.007	−0.043*	−0.036
	(2.23)	(0.008)	(0.024)	(0.025)
Log-likelihood	—	40.2	−28.7	−31.2

*** Significant at the 0.01 level.
 ** Significant at the 0.05 level.
  * Significant at the 0.10 level.

theoretical basis for deciding which variables should be included in the different
efficiency equations. Nevertheless, arbitrary choices were made in this regard and
several experiments were performed to generate SUR estimates. This preliminary
analysis revealed that the estimates obtained from the two-limit tobit models were
similar to those obtained from the SUR experiments. Hence, the discussion that
follows focuses on the results obtained from the two-limit tobit models.

According to the results of the second step regressions, presented in Table IV,
CONT has a positive and highly significant impact on both *EE* and *AE*, while the
effect on *TE* is also positive but not significant. This finding is consistent with
Glover's argument (1984) that contract farming can be very valuable for small-
scale operations, because it facilitates access to markets and increases income and
employment for growers. In addition contract production with an agribusiness firm
provides farmers with a secure market for their crops as well as some technical
assistance (Oberg et al. 1985). Furthermore, contract farming may improve
allocative or price efficiency, and thereby economic efficiency, by reducing risk.

REF was found to have a positive and statistically significant connection with
both *EE* and *AE*. The association with *TE* is also positive, but nonsignificant. A
positive coefficient for the *REF* variable suggests that the reform process is directly
related to productivity, probably because beneficiaries have better access to exten-

sion personnel, nonformal education, and technical information. These results are also in line with the notion that local organizations that typically emerge in an agrarian reform process can lead to improvements in resource allocation (Dewalt 1979; Meyer 1989; Seligson 1982).

The positive and statistically robust relationship between *SIZE* and *EE* and *AE* supports the notion that medium-size farms (50 to 100 *tareas*) have an efficiency advantage over the other farms in the sample (less than 50 or more than 100 *tareas*). The link between efficiency and farm size has been the subject of much discussion in the literature (Berry and Cline 1979). However, only a few studies using frontier function methodology have investigated this issue in developing country agriculture, but most have found no statistically significant correlation between size and technical efficiency (Bravo-Ureta and Evenson 1994; Huang and Bagi 1984; Kalirajan 1991; Ray 1985; Squires and Tabor 1991). By contrast, in a non-frontier analysis, Kaiser (1988) found that large farms in the Dominican Republic had higher *EE* than small farms.

Formal education, commonly measured in years of schooling, is the farmer attribute that seems to have received most of the attention in the frontier function efficiency literature. The results of these studies, which have focused almost exclusively on *TE*, reveal that the association between schooling and individual farm *TE* is quite mixed. Various studies have found a positive connection (Belbase and Grabowski 1985; Kalirajan and Shand 1986; Phillips and Marble 1986), while several others have reported no statistically significant relationship between these two variables (Bravo-Ureta and Evenson 1994; Kalirajan 1984, 1991; Kalirajan and Shand 1985; Phillips and Marble 1986).[6] The results presented in Table IV reveal a positive and statistically significant correlation between *SCHOOL* and *TE*, while the coefficient for this regressor in the *EE* and *AE* equations are negative and non-significant. Therefore, these results indicate that farmers with four or more years of schooling exhibited higher levels of *TE*, but, surprisingly, this association does not carry to the *EE* equation.

The results concerning *AGE* suggest that farmers under twenty-five years of age have higher levels of *TE*, *EE*, and *AE*. Of all the variables considered in the second step analysis, *AGE* is the only one that has uniformly the same sign and is statistically significant in all three efficiency equations. These results are consistent with the findings of Kalirajan and Flinn (1983), Kalirajan and Shand (1985), Belbase and Grabowski (1985), and Bravo-Ureta and Evenson (1994). According to Hussain (1989), older farmers are less likely to have contacts with extension agents and are less willing to adopt new practices and modern inputs. Furthermore, younger farmers are likely to have some formal education, and therefore might be

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6 For additional evidence see Bravo-Ureta and Pinheiro (1993), Kalaitzandonakes and Dunn (1995), and Kalirajan and Shand (1985).

more successful in gathering information and understanding new practices, which in turn will improve their economic efficiency through higher levels of *TE* and/or *AE*.

The last variable included in the tobit equations is the number of people in the household, *PEOP*, which was found to have a negative and significant correlation with *AE*. This result suggests that larger households might utilize family labor beyond the point where the marginal value product of labor is equal to the wage rate. By comparison, the coefficients for *PEOP* in the *TE* and the *EE* equations are positive and negative, respectively, but statistically nonsignificant in both cases.

An interesting implication emerging from the second step analysis is that improvements in allocative efficiency offer a higher potential, compared to technical efficiency, for enhancing economic efficiency. From a policy perspective, contract production, farm size, and agrarian reform status are the variables found to be most promising for improvements in economic efficiency, primarily through gains in allocative efficiency.

With regard to agrarian reform, in this age of market-oriented policies, it is somewhat difficult to recommend the enactment of land redistribution schemes to foster gains in economic efficiency. However, it is reasonable to argue that the greater *AE* reached by agrarian reform beneficiaries, relative to those peasant farmers in the sample that were not beneficiaries, might be due to technical and managerial assistance, as well as to better access to information, which sometimes are components of agrarian reform processes. If so, this result would be consistent with several other studies that have found a positive connection between farm-level efficiency and the availability of extension services and access to information (Bravo-Ureta and Evenson 1994; Kalirajan 1981, 1984, 1991; Kalirajan and Flinn 1983; Kalirajan and Shand 1986; Ray 1985; Shapiro and Müller 1977). In fact, the positive association between extension and efficiency appears to be the most robust finding in the farm efficiency literature focusing on developing country agriculture (Bravo-Ureta and Pinheiro 1993). Consequently, the analysis is consistent with the notion that public investment geared to improving the provision of managerial support and the dissemination of information to peasant farmers via extension programs or other forms of nonformal education are likely to lead to higher levels of efficiency.[7]

V.   CONCLUDING REMARKS

This paper has presented measures of technical, allocative, and economic efficiency for a sample of sixty peasant farmers in the Dajabon region of the Domini-

[7] A rationale for the public funding of extension services can be found in Birkhaeuser, Evenson, and Feder (1991).

can Republic. Maximum likelihood techniques were used to estimate a Cobb-Douglas production frontier, which was then used to derive its corresponding dual cost frontier. These two frontiers are the basis for deriving farm-level efficiency measures.

The analysis reveals average levels of technical, allocative, and economic efficiency equal to 70 per cent, 44 per cent, and 31 per cent, respectively. These results suggest that substantial gains in output and/or decreases in cost can be attained given existing technology. The results also point to the importance of examining not only *TE*, but also *AE* and *EE* when measuring productivity. We would like to point out that despite the role that higher efficiency levels can have on output, productivity gains stemming from technological innovations remain of critical importance in agriculture. Hence, research efforts directed toward the generation of new technology should not be neglected.

In a second step analysis, relationships between *TE*, *AE*, and *EE*, and various attributes of the farm and farmer were examined. The second step analysis relied on two-limit tobit regression techniques to estimate three separate equations, where *TE*, *AE*, and *EE* were expressed as functions of six farm/farmer characteristics: contract farming, agrarian reform status, farm size, schooling, producer's age, and household size. The results show that younger, more educated farmers exhibited higher levels of *TE*. In addition, it was found that contract farming, medium-size farms, and being an agrarian reform beneficiary have a statistically positive association with *EE* and *AE*. By contrast, the number of people in the household has a negative association with *AE*. Younger farmers are found to have uniformly higher levels of *TE*, *AE*, and *EE* than their older counterparts.

An important conclusion stemming from the analysis of our sample of Dominican Republic peasant farmers is that *AE* appears to be more significant than *TE* as a source of gains in *EE*. From a policy point of view, contract production, farm size, and agrarian reform status are the variables found to be most promising for action. The analysis suggests that policymakers should foster the development of medium-size farms, while promoting contract arrangements between peasant farmers and agribusinesses. Concerning agrarian reform, we argue that the statistically significant positive relationship between *AE* and being an agrarian reform beneficiary is likely a result of better access to managerial assistance and to information. If this is the situation, then our analysis supports the argument for public sector involvement in the provision of information to peasant farmers as a means to improve efficiency levels, and thus household incomes.

As is the case with any empirical study, the findings reported in this paper should be interpreted with caution. The model used in the analysis does not incorporate several factors that might influence efficiency such as risk, market imperfections, cash constraints, and social status. Finally, it is worth emphasizing that this study is one of the few that has measured and attempted to explain the variation in *TE*, *AE*,

and *EE*. Most studies concentrate on the measurement of *TE* and a few have gone on to explain its variation. One notable exception is the recent work by Ferrantino, Ferrier, and Linvill (1995). It seems clear from their work, and from the analysis presented in this paper, that additional methodological and empirical efforts are needed to improve our understanding of the determinants of the various efficiency indicators.

REFERENCES

Ahmad, Munir, and Boris E. Bravo-Ureta. Forthcoming. "Dairy Farm Technical Efficiency Measures Using Panel Data and Alternative Model Specifications." *Journal of Productivity Analysis*.

Aigner, Dennis J.; C. A. Knot Lovell; and Peter Schmidt. 1977. "Formulation and Estimation of Stochastic Frontier Production Function Models." *Journal of Econometrics* 6, no. 1: 21–37.

Bagi, Faqir Singh. 1982. "Economic Efficiency of Sharecropping: Reply and Some Future Results." *Malayan Economic Review* 27, no. 2: 86–95.

Battese, George E. 1992. "Frontier Production Functions and Technical Efficiency: A Survey of Empirical Applications in Agricultural Economics." *Agricultural Economics* 7, nos. 3/4: 185–208.

Belbase, Krishna, and Richard Grabowski. 1985. "Technical Efficiency in Nepalese Agriculture." *Journal of Development Areas* 19, no. 4: 515–25.

Berry, R. Albert, and William R. Cline. 1979. *Agrarian Structure and Productivity in Developing Countries*. Baltimore, Md.: Johns Hopkins University Press.

Birkhaeuser, Dean; Robert E. Evenson; and Gershon Feder. 1991. "The Economic Impact of Agricultural Extension: A Review." *Economic Development and Cultural Change* 39, no. 3: 607–50.

Bravo-Ureta, Boris E., and Robert E. Evenson. 1994. "Efficiency in Agricultural Production: The Case of Peasant Farmers in Eastern Paraguay." *Agricultural Economics* 10, no. 1: 27–37.

Bravo-Ureta, Boris E., and António E. Pinheiro. 1993. "Efficiency Analysis of Developing Country Agriculture: A Review of the Frontier Function Literature." *Agricultural and Resource Economics Review* 22, no. 1: 88–101.

Bravo-Ureta, Boris E., and Laszlo Rieger. 1990. "Alternative Production Frontier Methodologies and Dairy Farm Efficiency." *Journal of Agricultural Economics* 41, no. 2: 215–26.

————. 1991. "Dairy Farm Efficiency Measurement Using Stochastic Frontiers and Neoclassical Duality." *American Journal of Agriculture Economics* 73, no. 2: 421–28.

Caves, Richard E., and David R. Barton. 1990. *Efficiency in U.S. Manufacturing Industries*. Cambridge, Mass.: MIT Press.

CEDOPEX (Dominican Export Promotion Center). 1985. *Boletines estadísticos anuales, 1977–85*. Santo Domingo, Dominican Republic.

Dewalt, Billie R. 1979. *Modernization in a Mexican Ejido: A Study in Economic Adaptation.* New York: Cambridge University Press.

Dorner, Peter. 1975. *Land Reform and Economic Development*. Baltimore, Md.: Penguin Books.

Farrell, Michael. 1957. "The Measurement of Productivity Efficiency." *Journal of the Royal Statistics Society*, series A, 120, part 3: 253–90.

Ferrantino, Michael J.; Gary D. Ferrier; and Carl B. Linvill. 1995. "Organizational Form and Efficiency: Evidence from Indian Sugar Manufacturing." *Journal of Comparative Economics* 21, no. 1: 29–53.

Glover, David J. 1984. "Contract Farming and Smallholder Outgrower Schemes in Less-developed Countries." *World Development* 12, nos. 11/12: 1143–57.

Good, David H.; M. Ishaq Nadiri; Lars-Hendrik Roller; and Robin C. Sickles. 1993. "Efficiency and Productivity Growth Comparisons of European and U.S. Air Carriers: A First Look at the Data." *Journal of Productivity Analysis* 4, nos. 1/2: 115–25.

Greene, William H. 1980. "Maximum Likelihood Estimation of Econometric Frontier Functions." *Journal of Econometrics* 13, no. 1: 27–56.

———. 1991. *LIMDEP: User's Manual and Reference Guide*. New York: Econometric Software, Inc.

Greene, Duty D., and Terry L. Roe. 1990. "República Dominicana." In *Economía política de las intervenciones de precios agrícolas en America Latina*, ed. A. Krueger, M. Schiff, and A. Valdés. Santiago: Alfabeta Impresores.

Hayami, Yujiro, and Vernon W. Ruttan. 1985. *Agricultural Development: An International Perspective*. Baltimore, Md.: Johns Hopkins University Press.

Hossain, Mahabub. 1988. *Nature and Impact of the Green Revolution in Bangladesh.* Research Report 67. Washington, D.C.: International Food Policy Research Institute.

Huang, Cliff J., and Faqir S. Bagi. 1984. "Technical Efficiency on Individual Farms in Northwest India." *Southern Economic Journal* 51, no. 1: 108–15.

Hussain, Syed S. 1989. "Analysis of Economic Efficiency in Northern Pakistan: Estimation, Causes and Policy Implications." Ph.D. diss., University of Illinois.

Johnston, Bruce F., and John W. Mellor. 1961. "The Role of Agriculture in Economic Development." *American Economic Review* 51, no. 4: 566–93.

Jondrow, James; C. A. Knot Lovell; Ivans. Materov; and Peter Schmidt. 1982. "On the Estimation of Technical Efficiency in the Stochastic Frontier Production Function Model." *Journal of Econometrics* 19, nos. 2/3: 233–38.

Judge, George G.; R. Carter Hill; William E. Griffiths; Helmut Lutkepohl; and Tsoung-Chao Lee. 1988. *Introduction to the Theory and Practice of Economics*. New York: John Wiley & Sons.

Kaiser, Harry M. 1988. "Relative Efficiencies of Size and Implications for Land Redistribution Programs in the Dominican Republic." *Applied Agricultural Research* 3, no. 3: 144–52.

Kalaitzandonakes, Nicholas G., and Elizabeth G. Dunn. 1955. "Technical Efficiency, Managerial Ability and Farmer Education in Guatemalan Corn Production: A Latent Variable Analysis." *Agricultural and Resource Economics Review* 24, no.1: 36–46.

Kalirajan, K. 1981. "An Econometric Analysis of Yield Variability in Paddy Production." *Canadian Journal of Agricultural Economics* 29, no. 3: 283–94.

———. 1984. "Farm-Specific Technical Efficiencies and Development Policies." *Journal of Economics Studies* 11, no. 3: 3–13.

———. 1991. "The Importance of Efficient Use in the Adoption of Technology: A Micro Panel Data Analysis." *Journal of Productivity Analysis* 2, no. 2: 113–26.

Kalirajan, K., and J. C. Flinn. 1983. "The Measurement of Farm-Specific Technical Efficiency." *Pakistan Journal of Applied Economics* 2, no. 2: 167–80.

Kalirajan, K., and R. T. Shand. 1985. "Types of Education and Agricultural Productivity: A

Quantitative Analysis of Tamil Nadu Rice Farming." *Journal of Development Studies* 21, no. 2: 232–43.

———. 1986. "Estimating Location-Specific and Firm-Specific Technical Efficiency: An Analysis of Malaysian Agriculture." *Journal of Economic Development* 11, no. 2: 147–60.

Kopp, Raymond J., and W. Erwin Diewert. 1982. "The Decomposition of Frontier Cost Functions Deviations into Measures of Technical and Allocative Efficiency." *Journal of Econometrics* 19, nos. 2/3: 319–31.

Kopp, Raymond J., and V. Kerry Smith. 1980. "Frontier Production Function Estimates for Steam Electric Generation: A Comparative Analysis." *Southern Economic Journal* 46, no. 4: 163–73.

Kuznets, Simon. 1966. *Modern Economic Growth: Rate, Structure, and Spread*. New Haven, Conn.: Yale University Press.

Lingard, John; L. Castillo; and S. Jayasuriya. 1983. "Comparative Efficiency of Rice Farms in Central Luzon, the Philippines." *Journal of Agricultural Economics* 34, no. 2: 37–76.

Meeusen, Wim, and Julien van den Broeck. 1977. "Efficiency Estimation from Cobb-Douglas Production Function with Composed Error." *International Economic Review* 18, no. 2: 435–44.

Meyer, Carrie A. 1989. "Agrarian Reform in the Dominican Republic: An Associative Solution to the Collective/Individual Dilemma." *World Development* 17, no. 8: 1255–67.

Oberg, Keith, et al. 1985. *Linking Small Farmer Organizations to Non-Traditional Agribusiness in the Dominican Republic*. Washington, D.C.: United States Agency for International Development.

Phillips, Joseph M., and Robert P. Marble. 1986. "Farmer Education and Efficiency: A Frontier Production Function Approach." *Economics of Education Review* 5, no. 3: 257–64.

Pinheiro, António E. 1992. "An Econometric Analysis of Farm-level Efficiency of Small Farms in the Dominican Republic." Master's thesis, University of Connecticut.

Quezada, Norberto A. 1981. "Endogenous Agricultural Price and Trade Policy in the Dominican Republic." Ph.D. diss., Purdue University.

Rawlins, Glenville. 1985. "Measuring the Impact of I.R.D.P.II upon the Technical Efficiency Level of Jamaican Peasant Farmers." *Social and Economic Studies* 34, no. 2: 71–96.

Ray, Subhash C. 1985. "Measurement and Test of Efficiency of Farms in Linear Programming Models: A Study of West Bengal Farms." *Oxford Bulletin of Economics and Statistics* 47, no. 4: 371–86.

Schmidt, Peter. 1985–86. "Frontier Production Functions." *Econometric Reviews* 4, no. 2: 289–328.

Schultz, Theodore W. 1964. *Transforming Traditional Agriculture*. New Haven, Conn.: Yale University Press.

Seligson, Mitchell A. 1982. *Peasant Participation in Costa Rica's Agrarian Reform: A View from Below*. Ithaca, N.Y.: Rural Development Committee, Cornell University Center for International Studies.

Shapiro, Kenneth H., and Jürgen Müller. 1977. "Sources of Technical Efficiency: The Roles of Modernization and Information." *Economic Development and Cultural Change* 25, no. 2: 293–310.

Squires, Dale, and Steven Tabor. 1991. "Technical Efficiency and Future Production Gains in Indonesian Agriculture." *Developing Economies* 24, no. 3: 258–70.

Taylor, Timothy G., and J. Scott Shonkwiler. 1986. "Alternative Stochastic Specifications of the Frontier Production Function in the Analysis of Agricultural Credit Programs and Technical Efficiency." *Journal of Development Economics* 21, no. 1: 149–60.

Taylor, Timothy G.; H. Evan Drummond; and Aloisio T. Gomes. 1986. "Agricultural Credit Programs and Production Efficiency: An Analysis of Traditional Farming in Southeastern Minas Gerais, Brazil." *American Journal of Agricultural Economics* 68, no. 1: 110–19.

U. S. Department of Agriculture. 1987. "Annual Agricultural Situation Report—Dominican Republic." Washington, D.C.: Foreign Agricultural Service.

Williams, Simon, and Ruth Karen. 1985. *Agribusiness and the Small-Scale Farmer: A Dynamic Partnership Worldwide for Development.* Boulder, Colo.: Westview Press.

World Bank. 1990. *World Tables, 1989–90.* Baltimore, Md.: Johns Hopkins University Press.

Zellner, A.; J. Kmenta; and J. Drèze. 1966. "Specification and Estimation of Cobb-Douglas Production Function Models." *Econometrica* 34, no. 4: 784–95.