Productivity, Geography, and the Export Decision of Chilean Farms

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Abstract: This article analyzes the export participation of Chilean farms and the relative importance of farm-specific and geographic characteristics in this decision. An export behavior model is estimated using data on 8,284 Chilean farms and a two-stage conditional maximum likelihood procedure. Farm efficiency has a relatively stronger effect than the combined effect of geographic characteristics in increasing the probability of export participation. Farms with skilled (managerial) labor and in regions with higher human capital also have a relatively higher probability of producing for the export market. However, for geographic characteristics to positively affect export participation, farms must achieve a minimum level of efficiency.

Key Words: Agricultural Trade, Chile, Export Participation, Geography, Productivity.

JEL Codes: F11, O13, O18

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Why do some farms decide to produce exportables while others produce domestic-market oriented commodities? In the context of manufacturing industries, firms' decision to produce for foreign markets, popularly termed the export decision, has been extensively addressed beginning with the contribution of Bernard and Jensen (1995). The emerging theoretical and empirical literature on factors that underlie a firm's decision to export, continue to export or exit a foreign market have improved our understanding of exporting firms' characteristics (Aw and Hwang, 1995; Aitken, Hanson and Harrison, 1997; Bernard and Jensen, 1997, 1999, 2004a; Roberts and Tybout, 1997; Helpman, Melitz and Yeaple, 2004). The accumulated evidence indicates high-productivity firms self-select into export markets, and exporters survive longer and pay higher wages relative to nonexporters in high- and low-income economies (Wagner, 2005).

The export behavior modeled in the context of manufacturing sector does not apply directly to the case of primary agriculture, where farms often do not export and marketing firms make the export decision. However, farms decide on producing goods where the export intensity, i.e., share of exports in domestic production is either larger or smaller (Bernard and Jensen, 1995; Pavncik, 2002). Why is the decision to participate in exportable production important in the agricultural sector? Because, agriculture is one of the highly protected segments of developed and developing economies, and attempts to bring about successful liberalization of the agricultural sector have often been countered with structural-adjustment concerns (Aksoy and Beghin, 2005). Most studies of agricultural trade liberalization claim long-run benefits to reform, but cite significant structural adjustment and the short- to medium-term harm to farm and rural communities. These studies do not necessarily model farm-level decision making, which is a significant factor in the structural adjustment process from a protected regime to a market-based economy. Understanding farms' export participation would aid in creating successful exporters, and making liberalized and open-market policies politically feasible. Successful exporters bring about stable income growth to the specific industry and the broader economy.

The objective of this article is to analyze the export participation of Chilean farms and to identify the relative importance of farm-specific and geographic characteristics in this decision. Indeed, Chile is an excellent example of the export-led growth theory with relatively open markets including the agricultural sector (World Bank, 2001). We set up an export behavior model along the lines of Aitken, Hanson and Harrison (1997). Data on 8,284 Chilean farms are assembled for 1997 and an export participation rule similar to that of Pavcnik (2002) is utilized. The farm-specific characteristics in our analysis include efficiency (productivity), size and ownership structure. In farms' export participation, we consider the role of geographic characteristics, measured by infrastructure, natural advantages, human capital, and institutional quality at the regional level (Krugman, 1991; Limao and Venables, 2001). Since the export decision literature has addressed simultaneous determination of export decision and firm characteristics, we test and correct for possible endogeneity of regressors using a two-stage conditional maximum likelihood procedure.

An Export Behavior Model

Our approach to modeling export participation of Chilean farms is similar to that of Aitken, Hanson and Harrison (1997). Suppose firms choose to produce for the domestic market or the foreign market or both. Assume that firms incur additional costs to produce for different markets, which are termed as distribution costs. Furthermore, distribution costs in the domestic market are different from that of the foreign market. Total cost for a firm, indexed by *j*, is:

$h^j(q_d^j + q_f^j) + m_d^j(q_d^j) + m_f^j(q_f^j)$, where *d* and *f* index domestic and foreign market, *h()* and *m_i()*, $i = d, f$, are the production and distribution cost functions, respectively. Separability in production and distribution costs is assumed, and *h()* and *m_i()* are increasing and convex in their respective arguments. The production decision of the *j*-th firm is given by the solution to:

(1) $$\max_{q_d, q_f}\{P_d q_d^j + P_f q_f^j - h^j(q_d^j + q_f^j) - m_d^j(q_d^j) - m_f^j(q_f^j)\} \quad s.t. \quad q_d, q_f \geq 0$$

where P_d and P_f are prices (not necessarily specific to the firm) received in domestic and foreign market, respectively.[1] The optimal output choice may be zero in either market. All firms produce positive quantities for the domestic market but, in practice, some firms produce zero exports. As in Aitken, Hanson and Harrison (1997), we only consider the possibility of a corner solution for the variable q_f^j. Let $q_f^{j^*}$ be the latent variable such that,

(2) $$\begin{cases} q_f^{j^*} = q_f^j & \text{if} \quad q_f^j > 0 \\ q_f^{j^*} = 0 & \text{otherwise.} \end{cases}$$

Given our interest in the firm export decision, we choose to focus on the estimation of the probability that a firm exports. Let the dummy variable y_j take value *1* if $q_f^j > 0$ and 0 otherwise. The estimation of the discrete choice model based on the latent variable in equation (2) allows us to obtain consistent estimates of the underlying solution to $q_f^{j^*}$. It follows from equation (2) and the definition of y_j that the probability that the *j*-th firm exports is given by:

(3) $$\Pr(y_j = 1) = \Pr(\alpha + \beta X_j + \delta T_j^k + \gamma Z_j^k + \varepsilon_j > 0),$$

where ε_j is normally distributed, which permits the estimation of equation (5) as a binary probit

[1] Similar to Aitken, Hanson and Harrison (1997), we have a static export-behavior model based on distribution-cost differences. Roberts and Tybout (1997), Bernard and Jensen (2004), and Helpman, Melitz and Yeaple (2004) find evidence of sunk costs in firms' export decision. Such dynamic considerations are beyond the scope of this article since we have cross-sectional data.

model. In equation (3), X_j is a vector of firm-characteristics including size and productivity arising from the production and distribution cost functions, and β is the associated parameter vector of interest; T_j^k is the terms of trade between exportables and domestic production (P_f / P_d) with its corresponding parameter δ; Z_j^k is the *k*-th regional or geographic characteristics within which the firm operates originating either from the output prices or cost functions or both, and the parameter vector γ measures their relative importance to the probability of export production. Examples of Z_j^k include infrastructure, natural advantages, and human capital.

Chilean Data

Data from the Chilean Agricultural Planning Office (ODEPA) show that fruits and fruit-derivatives (e.g., wine) accounted for about 80 percent of Chilean agricultural exports during 1990-2000 (ODEPA, 2001). Moreover, 15 fruits accounted for 93 percent of all fruit exports. The second largest export group is vegetables and flowers accounting for another 11 percent of exports, while traditional agricultural commodities like grains and animal products (e.g., beef) constituted a small share (< 6 percent) of exports. Therefore, we identify a set of twelve crops that are considered traditional and not exportable. Consequently, farms market participation or orientation can be defined according to what they produce:

- Exporter ($q_f^{j^*} > 0$): Farms producing for the export market, if they only produce some or all of the (15) exportable fruits
- Traditional ($q_d^{j^*} > 0$, $q_f^{j^*} = 0$): Farms producing for the domestic market, if they only produce some or all of the (12) non-exportable traditional crops

Note that the above categorization is not unlike that of Pavcnik (2002), who uses plant-level data from the Chilean manufacturing sector (see also Alvarez and Lopez, 2004). The above criterion is not a classification of exportables and traditionals (which differ by country), but one to identify farms' willingness and ability to participate in exportable production.

Farm-level data are obtained from the VI Chilean Census of Agriculture (1997) conducted by the Chilean National Institute of Statistics (CNIS), which is the only agricultural census since 1976. Data include location of farm (county and region), number of employees, area and production of individual crops and animal products, total land area and demographic information on farm households including age (experience) and size. The database has over 300,000 farms. However, we face the problem of farms producing some of exportables (15) and traditional (12) crops as well as products not included in either list.[2] Therefore, we select a set of farms that only produce at least one of the fifteen fruits or one of the twelve traditional crops, and that are not involved in any other crop or animal production.[3] The sampled farms use land to only produce either selected traditional crops or exportable fruits, but do not have land allocated to produce other crops or fruits.[4] Thus, farms that have land for producing vegetables, seeds, flowers, annual or permanent pastures and forages (dairy or cattle farms) are eliminated. The resulting sample of 13,478 farms could still be producers of other products. So, we select farms that have an area for "other uses" less than or equal to the 25 percent of total land area. The application of the above selection criteria yielded a sample of 8,284 farms.[5]

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[2] It is not possible to group farms by specific crops or fruits (e.g., grapes) since they produce more than one product.

[3] A key question addressed in this study is whether or not higher productivity causes participation in exportable production. A measure of productivity should reflect the true, overall farm efficiency of using inputs to generate outputs. Situations where a farm could be inefficient producing an exportable fruit, but highly efficient producing a non-exportable fruit should therefore be avoided.

[4] In the multinomial logit model, with a sample that also included farmers producing exportable and traditional products, the independent of irrelevant alternative assumption was rejected.

[5] Setting the cutoff to zero or 50 percent produced a sample of 4,430 or 13,478 farms but, did not affect the quality of our results in the next section.

The primary farm-specific characteristic in the vector X_j in equation (5) is the farm-level productivity or efficiency index (Coelli, Rao and Battese, 1998). Using data envelopment analysis (DEA) farms are ordered according to their technical efficiency, which is defined as the distance to the production frontier. A linear programming problem is solved for each farm, whose solution θ corresponds to technical efficiency; whenever $\theta = 1$, a farm is technically efficient.[6] Other variables in the vector X_j in equation (3) include farm size represented by total land (traditional crops and exportables) and employment (family and hired labor); farm-owner's experience represented by age; and a dummy for the presence of a manager or operator hired by the farm-owner (manager). A variable representing terms of trade between exportables and traditionals is computed for each farm based on prices from ODEPA.

To represent Z_j^k, i.e., geographic characteristics of the k-th region, data are obtained from two main sources: *The Regional Competitiveness Report, 2001* (Informe de Competitividad Regional, 2001), published by the Chilean Ministry of Economy (CME), and *The Chile-Environmental Statistics 1998-2002* (Chile – Estadísticas del Medio Ambiente 1998-2002) of CNIS.[7] The normalized indexes from these two sources measure geographic characteristics in the following categories. Infrastructure is represented by the non-farm capital index, which includes industrial (mining and manufacturing) capital, roads, potable water and sewer coverage. Natural advantage is represented by the soil type of a region.[8] The human capital index is a weighted combination of average schooling coverage, performance of schools, performance in

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[6] The efficiency scores under the constant and variable returns to scale assumptions are of similar magnitude. Reported results in the next section are quantitatively and qualitatively similar to those obtained under constant returns to scale. We represent efficiency scores using $1/\theta$, where inefficient farms have scores less than 1.

[7] In our estimation, data from CME reports for only 1997 are used.

[8] Natural advantages specific to the region can be measured by factors such as type of soil, temperature, precipitation, and topography. We chose soil type since there exists an established scientific classification of soils based on potential uses or use capacities.

college entry tests, workforce's years of schooling, health facilities and health indicators of workers. The government quality index measures the performance of a local government in creating a favorable environment for businesses and its inhabitants. The above representations of geographic characteristics are common to most studies of export behavior and consistent with the economic geography literature (Krugman, 1991; Bernard and Jensen, 1995; Roberts and Tybout, 1997; Aitken, Hanson and Harrison, 1997; Limao and Venables, 2001).

Estimation Procedure

Based on equation (5), initial versions of the binary probit model are specified to address the problem of endogenous regressors: efficiency, labor and land at the farm level. Due to space limitations, we do not present the estimated coefficients of the initial specifications.

Prior theoretical and empirical analyses support the link between higher efficiency or productivity and export participation but, two competing hypotheses explain the directionality in this linkage. The first is the self-selection hypothesis, which states that only higher productivity firms will become exporters: Bernard and Jensen (1995, 1999) in the case of the United States; Clerides, Lach and Tybout (1998) for Colombia, Mexico and Morocco; Alvarez and Lopez (2004) in Chile; and Girma, Greenaway and Kneller (2004) in UK. In the case of a farm, there are extra costs associated with the production of exportable products, such as growing high-quality varieties sought by foreign markets, investing to preserve post-harvest quality costs, and related costs. These higher costs can only be afforded by high-productivity farms (self-selection), and therefore, the decision to export is impacted by farms' efficiency:

(4) $$y_{1j} = \alpha + \delta_1 y_{2j} + \beta_1 X_{1j} + \delta T_j^k + \gamma Z_j^k + u_{1j} \qquad j = 1, \ldots, n,$$

where efficiency, indexed by y_2, and a set of exogenous variables, X_1, T^k and Z^k, explain the export decision, y_1. Note that y_1 can only take the observable sign of a latent variable $q_f^{j^*}$.

The second hypothesis, learning-by-exporting, suggests that firms improve their productivity by participating in the exportable market (Clerides, Lach and Tybout, 1998; Aw, Chang and Roberts, 2000). In the case of Chilean agriculture, an export-oriented producer is exposed to demanding buyers/exporting firms, who require high-quality products to compete in international markets. Hence, farms learn from their export participation, which leads to higher productivity relative to those only producing for the domestic market:

$$y_{2j} = \gamma_2 y_{1j} + \beta_2 X_{2j} + u_{2j} \qquad j = 1, \ldots, n, \tag{5}$$

where X_2 is a set of explanatory variables for y_2. Similarly, decisions on labor and land allocation are likely determined jointly with the decision to produce exportables.

From equations (4) and (5) it is apparent that $E(u_1 u_2) \neq 0$, and so, the application of standard binary probit methods to equation (4) will yield inconsistent parameter estimates. To test and correct for regressors' endogeneity, we utilize the two-stage conditional maximum likelihood (2SCML) procedure developed by Rivers and Vuong (1988). The 2SCML estimator is consistent and asymptotically efficient for probit models with continuous exogenous variables. Rivers and Vuong (1988) proposed a reduced form for y_2 and assumed that the residuals in the reduced form of y_2 have a normal distribution. In the 2SCML procedure, the endogenous variable y_2 is regressed on selected instruments and all the explanatory variables of the system. Then, in the probit regression, the binary variable is regressed on y_2, explanatory variables and the residuals from the regression of y_2. To test y_2's exogeneity, Rivers and Vuong (1988) proposed a likelihood ratio (LR) test, which is given by $LR = -2(\ln \hat{L}_R - \ln \hat{L}_U)$, where $\hat{L}_U$ and

$\hat{L}_R$ are the log-likelihood values of the probit with and without the residuals from y_2 regression as explanatory variables, respectively. The LR statistic has a chi-squared distribution with degrees of freedom equal to the number of endogenous variables in the probit equation.

Based on the above, we test the assumption of endogeneity for three farm-specific variables: efficiency, labor and land.[9] The LR test rejects the null hypothesis that the farm efficiency index is exogenous (LR statistic, 61.80). Similarly, the labor input is found to be endogenous (LR statistic, 38.54) but, we cannot reject the null that the land area of a farm is an exogenous variable.

Two-Stage Probit Results

In table 1, we present four versions of the two-stage binary probit model for the export decision. All four specifications include the residuals from the instrumental regressions of efficiency and labor. Model (1) is a basic specification with farm-specific characteristics and the county-level terms of trade. Model (2) includes the physical (geographic) characteristics: non-farm capital and soil type, and model (3) includes people and institutional characteristics: human capital and government quality. Model (4) includes farm-specific characteristics, county-level terms of trade and all geographic indexes. Results in last three rows of table 1 show that the log likelihood value improves in the presence of either or both sets of geographic indexes, confirming their significant role in the export decision. The pseudo R^2 indicates that the regression lines of all four models well fit the observed data (82% to 83%). Furthermore, based on LR tests, model 4 fits the data best among the four alternative specifications in table 1.

The coefficient on efficiency score is positive and significant in model 4, which is

[9] The lack of additional instruments preclude us from testing geographic characteristics' endogeneity. The Rivers and Vuong (1997) procedure is not applicable to the manager dummy.

consistent with studies on export decision in the manufacturing sector (e.g., Bernard and Jensen, 1999; Helpman, Melitz and Yeaple, 2004). Good firms, i.e., firms with higher productivity, tend to export or in our case, participate in the exportable sector. The significance of farm-owner's age shows that more experienced producers have a higher probability of participating in exportable production in all 4 models. Based on model 4, the coefficient on the manager dummy suggests that the presence of a manager in a farm has a significantly positive effect on the decision to participate in exportable production. The results on age and manager variables suggest that the presence of skilled labor on a farm increases its probability of export participation (Bernard and Jensen, 1997). The other farm-specific variables mostly had relatively lower effects on export participation.

The coefficient on the county-level terms of trade is significantly positive in all models. Implicit in this index are the uniform tariffs of either 31.5 or 11 percent for most traditionals, i.e., $\frac{P_E}{P_T^W(1+t)}$, where P_E represents weighted exportables' price, while the denominator corresponds to tariff-adjusted (t), world price of traditionals. Therefore, a reduction in tariffs on traditionals' imports will increase a farm's probability of export participation.

With regard to geographic characteristics, the result that higher soil quality has a significantly negative effect on the probability of export participation is likely due to the Chilean soil classification system. The Chilean soil quality index documents sidehills and uneven topography, where a significant share of exportable fruits is produced, as low-quality soils. The advantage of sidehills is that they provide better exposure to sunlight relative to even topography. Thus, the negative relationship between the probability of export participation and "low-quality soils" of sidehills does not come as a surprise (Suelos. El Principio de la Vid, 2001). As noted earlier, the reason to include the non-farm capital index in the export behavior model is the hope that it will mimic infrastructure. However, the index is dominated by the presence or absence of mining and manufacturing capital with little impacts from roads and other public infrastructure components. Therefore, the higher the non-farm capital endowment, the lower is the probability of agricultural export participation.[10]

The human capital index, which is a source of productivity spillovers to the agricultural sector, has a positive and statistically significant effect on the participation decision (model 4). The quality of regional government has a statistically significant positive effect on the probability of export participation. Thus, regions with better local governmental performance appear to provide favorable conditions for export production.

Productivity versus Geographic Effects

In table 2 we present four sets of marginal effects: case 1 evaluates the marginal effects when all variables are evaluated at their respective means, the most commonly reported results in the discrete choice literature (Aitken, Hanson and Harrison, 1997); case 2 evaluates the marginal effects holding the human capital index at 0.9 and the rest of the variables at their respective means; case 3 holds the efficiency score at 0.9 and others at respective means; and case 4 evaluates marginal effects holding the efficiency score and human capital index at 0.9 and the other variables at their respective means. Case 1 shows that efficiency has the largest marginal effect (2.029) followed by that of the manager dummy (0.480), government quality (0.279), and human capital (0.206). In cases 2 and 3, holding the human capital index or efficiency score at 0.9 yields a ranking of marginal effects similar to that in case 1. When both efficiency and human capital effects are combined, one would anticipate that the probability of export participation will likely increase more than those in case 2 and 3. However, the marginal effects

[10] High correlation of non-farm capital with other available indexes in the CME report, and the long and narrow landscape of Chile make it harder to further decompose the effects of components of some these indexes.

in case 4 are lower than those of case 1, 2 or 3. These results arise from the nonlinearity inherent in discrete choice models.

Figure 1(a) and 1(b) demonstrate that the effect of changes in farm-specific productivity is stronger than the effect of the locational productivity on export participation. For example, if the human capital index increases from 0.75 to 0.90, the productivity of farms is affected in a way that lowers the threshold required for export participation as shown in figure 1(a). On the other hand, if the efficiency score increases from 0.75 to 0.90, then the effect on predicted probabilities of the human capital index is greater than the effect produced by a corresponding change in itself (not presented due to the space constraint). Alternatively, if a farm increases its productivity, its probability of export participation will be higher than when the overall productivity of its surroundings increases. Moreover, to participate in exportable production, a farm has to satisfy a minimum efficiency threshold. In other words, even when the geographic characteristics favor exports, if a farm does not meet the efficiency threshold, its probability of participating in exportable production will not increase.

Summary and Conclusions

In this article, we analyze Chilean farms' export participation and the relative importance of farm-specific and geographic characteristics in this decision. An export behavior model is estimated using data on 8,284 Chilean farms for 1997. Farm-specific characteristics (efficiency, size and ownership structure), an indicator for terms of trade between exportables and traditionals, and geographic characteristics (e.g., education, institutional quality) are considered in farms' export behavior. We test and correct for possible endogeneity of regressors (e.g., efficiency, labor) using a two-stage conditional maximum likelihood procedure.

Results suggest that trade liberalization, i.e., lowering tariffs on traditional commodities, improves the terms of trade of exportables, which significantly affects Chilean farms' decision to produce exportables. Farm-specific efficiency effects appear relatively stronger than the combined effects of geographic characteristics on export participation. When a highly efficient farm locates in a region with better geographic characteristics, its likelihood of producing for export markets is higher. On the other hand, opposite results are obtained when low-efficiency farms are located in regions with higher human capital or government quality. The latter is due to an efficiency threshold for farms to participate in exportable production. The effects of other farm and geographic characteristics such as farm-owner's age, land and soil type appear to be lower than that of farm efficiency, regional human capital and local government quality.

The results on productivity highlight the role that government policies can have in transforming domestic-market oriented farms into export producers. Policies should focus on improving farm-specific productivity up to and beyond the threshold when geographic characteristics become important in the export participation. Unless farms achieve a minimum efficiency level, investments in regional productivity or infrastructure appear to have relatively lower effects on export participation. Exporting is positively associated with profits and income, and to create successful exporters, a farm's efficiency is relatively more important than the characteristics of the region within which it operates.

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Table 1. Two-Stage Binary Probit Specifications for Export Participation

	Model 1	Model 2	Model 3	Model 4
Intercept	-17.727*	-13.417*	-18.941*	-15.190*
	(0.594)	(0.920)	(0.720)	(1.057)
Residual_Efficiency	-12.323*	-8.317*	-12.876*	-8.292*
	(0.586)	(0.818)	(0.782)	(1.098)
Residual_Labor	0.059*	0.036	0.361*	0.553*
	(0.028)	(0.031)	(0.052)	(0.095)
Farm Specific Characteristics				
Efficiency	20.728*	16.849*	22.417*	18.677*
	(0.760)	(0.954)	(0.990)	(1.207)
Age	0.021*	0.015*	0.027*	0.029*
	(0.003)	(0.003)	(0.003)	(0.004)
Manager	0.412	0.534*	2.630*	4.414*
	(0.237)	(0.250)	(0.380)	(0.730)
Labor	-0.060*	-0.025	-0.359*	-0.541*
	(0.028)	(0.031)	(0.052)	(0.095)
Land	-0.018*	-0.016*	-0.006*	0.005
	(0.002)	(0.002)	(0.002)	(0.004)
County-Level Terms of Trade	0.014*	0.011*	0.011*	0.010*
	(0.001)	(0.001)	(0.002)	(0.002)
Geographic Characteristics				
Non-Farm Capital		0.034		-2.664*
		(0.177)		(0.613)
Soil Type		-0.104*		-0.074*
		(0.018)		(0.019)
Human Capital			-0.288	1.899*
			(0.290)	(0.666)
Government Quality			1.793*	2.572*
			(0.308)	(0.448)
Number of observations	8284	8284	8284	8284
LR chi-squared statistic[1]	7945.65*	8035.44*	8035.54*	8070.84*
Log-likelihood value	-866.157	-821.256	-821.209	-803.563
Pseudo R^2	0.8210	0.8303	0.8303	0.8339

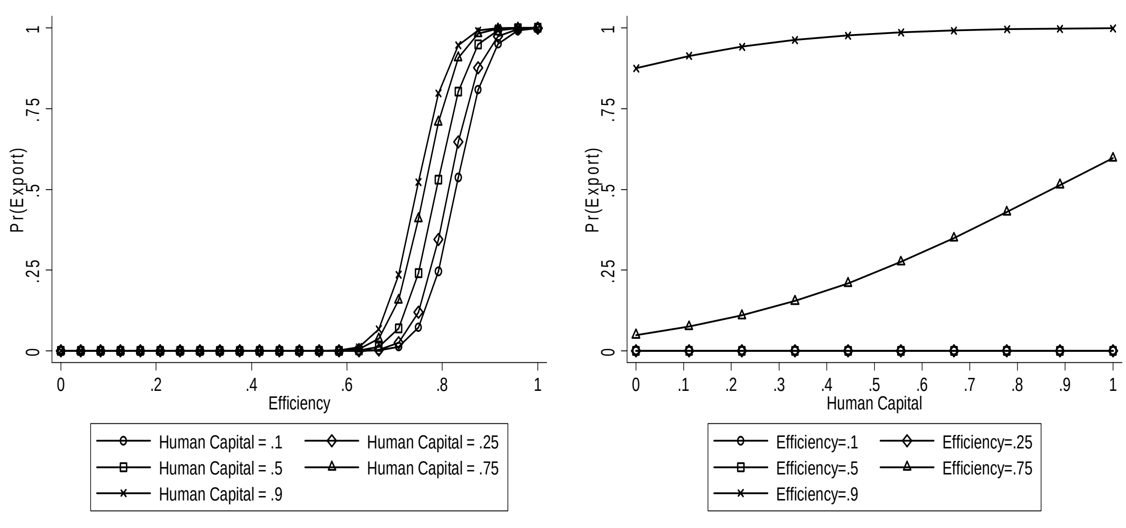
Numbers in parentheses are standard errors; * denote statistical significance at the 5% level.
[1]Test statistic has 8, 10, 10, and 12 degrees of freedom for each model, respectively.

Table 2. Marginal Effects From Two-Stage Probit Model (4)

	Case 1	Case 2	Case 3	Case 4
Efficiency	2.029	5.831	1.123	0.126
Labor	-0.059	-0.169	-0.033	-0.004
Land	0.001	0.002	0.000	0.000
Age	0.003	0.009	0.002	0.000
Manager	0.480	1.378	0.265	0.030
County-Level Terms of Trade	0.001	0.003	0.001	0.000
Infrastructure	-0.289	-0.832	-0.160	-0.018
Soil Type	-0.008	-0.023	-0.004	-0.001
Human Capital	0.206	0.593	0.114	0.013
Government Quality	0.279	0.803	0.155	0.017

Case 1: All variables at mean; Case 2: Human Capital at 0.9, Rest at respective mean, Case 3: Efficiency at 0.9, Rest at respective mean, Case 4: Efficiency at 0.9, Human Capital at 0.9, Rest at respective mean.

Figure 1. Predicted Probabilities of Alternative Efficiency and Human Capital Indexes



(a) (b)