Technical Efficiency of Grain Production in Ukraine

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Abstract

A stochastic production frontier model with technical inefficiency effects is estimated on a representative sample of Ukrainian state and collective grain-producing farms. Technical efficiency declined from 1989 to 1992. More experienced managers were found to be more productive, with the effect of experience diminishing with age. We found that on-farm provision of production infrastructure was associated with higher efficiency, a result supporting the hypothesis that disorganization causes output to decline in transition economies.

Key words: Technical efficiency, stochastic production frontier, economies in transition, grain production, disorganization hypothesis

TECHNICAL EFFICIENCY OF GRAIN PRODUCTION IN UKRAINE

In Soviet Ukraine, the farming system was an integral part of the centrally planned national economy and farm performance was judged not by the financial results of production, but by how well the centrally planned production targets were met. Artificially low food prices and subsidies left producers with little motivation for improving efficiency and competitiveness (Csaki and Lerman, 1997a). One of the goals of the reforms, begun in the early 1990s, has been to transform Ukrainian agriculture into a productive system motivated by private incentives in order to improve overall productive efficiency.

The reforms have progressed slowly, however. The restructuring and reorganization of the former state and collective farming system have been purely formal to date; most collective and state farms reorganized into the new legal form of collective farm enterprise with no internal reorganization (Csaki and Lerman, 1997b). Although 35,000 private farms appeared by 1997, the private sector still accounts for only 15 percent of the country's agricultural land. Collective forms of organization continue to dominate Ukrainian agriculture as they control most of the land and remain the major employer in rural areas.

The reforms, which included limited price liberalization and the introduction of private property, resulted in a sharp decline in production and consumption. By 1996, Ukrainian gross domestic product (GDP) dropped to 43 percent of its 1990 level, and gross agricultural product dropped to 59 percent of the 1990 level (Csaki and Lerman, 1997b).[1] Agricultural productivity, measured as output per worker, crop yields, milk per cow, and animal slaughter weights, for example, clearly deteriorated. And, even though the area sown in crops has decreased since 1990 by only 5 percent, the production of main cash crops, cereals, and sugar beets dropped by 30 to 40 percent between 1990 and 1995 (Csaki and Lerman, 1997b).

In general, an output decline can be attributed to movements along a path on or beneath the production surface (input use decline), movement away from the production surface (technical efficiency decline), and/or shifts in the production surface (technical change). Because there is no indication that the technology changed over the first reform years, we do not believe there was a systematic shift in the production surface for this reason, except, perhaps, for year-to-year movements due to varying weather. In this study, we hypothesize that, in addition to the input use decline, inefficiency increased in collective agricultural production, and we seek to quantify the changes in and the determinants of technical inefficiency in grain production over the period 1989 to 1992.

Grain is one of Ukraine's most important agricultural products. Ukraine produced an average 47.4 million tons of cereals per year in 1986 to 1990. The area under grain averaged 14,541 thousand hectares on average over the years 1989 to 1992, about 44 to 45 percent of the total Ukrainian area sown. Grains are the primary crops for both livestock production and human consumption. During the period 1989 to 1993, wheat accounted for 49 percent of the total area cultivated under cereals; barley, the most important feed grain, accounted for 19 percent, followed by maize, 15 percent (World Bank, 1994). Although state and collective farms were quite diversified agricultural enterprises, most of them produced grains. Currently, over 90 percent of Ukrainian grain is still produced by the former state and collective farms (Valdes et al., 1997).

Blanchard and Kremer (1997) argue that the decline in output and in productivity at the beginning of the transition in former Soviet republics can be attributed to the disorganization that occurred as old production links, such as the state input distribution system, deteriorated significantly, while new production links had not yet emerged. In particular, contractual agreements enforceable by the coercive power of a central planner in Soviet times became much less obligatory with the transition. Essentially, the mechanisms of contract enforcement that exist in the West, such as law and reputation, could not be created overnight. Despite the intuitive appeal of this theory, there has been minimal empirical support, and, to our knowledge, no study has investigated the effect of disorganization on transition agriculture. This paper provides empirical support to the hypothesis that disorganization (in addition to decreased use of inputs) contributed to a loss of agricultural output in the early 1990s.

We use stochastic production frontier modeling, the framework that has contributed to other related agricultural policy analyses (see Battese, 1992; Bravo-Ureta and Pinheiro, 1993; Coelli, 1995b). Knowing the effects of farm-specific variables—such as characteristics of the farm manager and farm management system, the experience of managers, and the distance from supply and distribution points—allows for better understanding of farm production efficiency. Because aggregation of the data across farms might mar the impact of these factors, farm-level data are preferred for inefficiency analysis.

To date, little research exists on the efficiency of agriculture in the formerly planned economies. Data have been limited, especially, at the farm level. Many earlier studies have used country-level data (Koopman, 1989; Carter and Zhang, 1994) and regional within-country aggregates (Boyd, 1987, 1988; Brada and King, 1993; Hofler and Payne, 1993, 1995; Sedik et al., 1999; Sotnikov, 1998). The existing research also has been limited by aggregation of outputs and inputs that uses "synthetic" Soviet-time prices, the prices that rarely reflected the relative scarcity of resources in the economy. Brock (1994, 1997) analyzed farm-level data to study production efficiency using a "whole farm" production function. A few other studies have examined farm-level technical efficiency of a single crop, but aggregated inputs through artificial monetary valuations (e.g., Skold and Popov, 1990, 1992; Johnson et al, 1994; Bayarsaihan et al., 1998). Our research avoids these aggregation problems by using farm-level survey data on grain production, all reported in physical units.

In this study of Ukrainian grain-producing farms, we estimate a frontier production function, examine the changes in technical efficiency at the earliest stages of the economic reforms, and evaluate the relationship between technical efficiency and selected farm characteristics. With limited price data available, we also attempt to evaluate allocative efficiency.

The combination of the three strengths—the time period analyzed, the quality of data, and the model employed—makes our study different from previous research. Our study covers the beginning of the reforms, utilizes farm-level data reported in physical units, and employs an inefficiency model that allows for simultaneous estimation of the parameters of both the frontier production function and the inefficiency effects. To our knowledge, no study has conducted similar analysis for any of the countries of the former Soviet Union.

The paper is organized as follows:

- data and farm structure are described together with productivity indicators,
- a stochastic frontier production model is presented,
- a discussion of results of estimation,
- a discussion of allocative efficiency, and
- a summary of our findings with concluding remarks is provided.

Data

The data come from a random survey of state and collective farms in Ukraine during the period 1989 to 1992. Because little internal restructuring has occurred since this period, the clear advantage of the detailed input and output data reported in physical units outweighs the possible disadvantage of using eight-year-old data.

The data were collected in 1992 retrospectively for 1989 to 1991. The survey was designed as a random sample of state and collective farms across agro-climatic zones and was stratified by farm size (Carriquiry, 1993). The Ukrainian Institute for Agrarian Economics (UIAE) supervised the administration of the survey. The data were based on farm-kept written records that are the source for standard statistical questionnaires filled out at the end of each year. Out of the original 80 farms surveyed, data for 41 farms from two administrative regions, the *Kyivska oblast* and *Cherkaska oblast* of the mixed soil-climatic zone, were complete and were used for the analysis. The mixed soil-climatic zone has average annual precipitation of 450 to 600 mm and has, for the most part, highly favorable black soils. This zone takes up about one-third of the total Ukrainian agricultural land. Comparison of sample means with those of census data confirms that the sample is representative for the mixed soil-climatic zone of Ukraine.[2]

Descriptive Statistics

Table 1 shows changes over the period 1989 to 1992 typical for Ukraine's agricultural sector as large collective and state farms started to downsize. The average decline in the land holdings of the farms over the period was 7.6 percent, because of obligatory transfer of land to state reserves. The reserves serve as a source of land for subsidiary household plots of state and collective farm members and for independent private farms (Csaki and Lerman, 1997b). On average, farm employment decreased by 13 percent, and the number of pensioners per worker increased by 17 percent over the period. The decrease in the working population on the farms can be attributed to young people leaving farms and going into cities and retired workers remaining on the farms (World Bank, 1994; UIAE, 1992). The UIAE (1992) reports specifically that the rural population of the two administrative regions represented in the data declined by more than 22 percent from 1970 to 1990.

During Soviet times, the farms not only produced agricultural output, but also provided most of the social and municipal services for the communities where they were located. In addition, the farms often did their own construction of production facilities, maintenance, and processing. In 1992, about 13 percent of farm employees were social, maintenance, repair, construction, or processing workers (Table 1). On average, the share of farm workers devoted to activities other than agricultural production declined by 8 percent. In contrast, the share of nonagricultural production expenditures increased from less than 5 percent in 1989 to almost 20 percent of total farm expenditures in 1992. The complex system of subsidies, bonuses, and other price distortions, as well as the problems of maintaining adequate financial accounting under conditions of high inflation[3] make clear the need to use physical input measures in preference to monetary valuations.

Partial Productivity Indicators

Table 2 shows the grain production and input use based on the survey results. On average, total grain production declined by 35 percent over the four years, although the area under grains declined by only 12 percent. A part of the decline in yield can be attributed to poor weather in 1991 and moderately inferior weather in 1992 (prolonged drought during the summer combined with high temperatures) (IMF, 1993), although input shortages aggravated the situation. The application of inputs changed dramatically, as application of chemicals per hectare decreased 20 percent, organic fertilizer per hectare dropped 16 percent, and labor use per hectare increased more than 90 percent on average over the four years. The decline in the application of chemicals must be attributed to sharp increases in the prices of agricultural inputs relative to prices of agricultural output. The decrease for organic fertilizer came from the downsizing of livestock operations (Csaki and Lerman, 1997b). In this situation, farms substituted the inputs that were readily available (labor and land) for more expensive and relatively scarce inputs (chemicals and fertilizer).

The reported diesel fuel used per hectare, a proxy for machinery services, did not change over time—a result that seems surprising because fuel is the agricultural input that experienced the highest price increase (Csaki and Lerman, 1997b). The lack of change may reflect deliberate overreporting of fuel use on farm accounts to ensure enough allotment from the state in the future. Although Brock (1994) did not find deliberate overreporting in farm annual accounts, it is not clear whether Brock refers to both input utilization and output or to the output only. Alternatively, the unchanged fuel use may reflect the adequate provision of inputs due to the preferred status of grain production. Because we have no other measures of capital expenditures and no means to check for or to correct the possible error in the measurement of reported fuel input, we make no adjustment to the values of this input.

Method

We estimate a frontier production model with technical inefficiency effects as proposed by Battese and Coelli (1995). To disentangle the effect of the decline in input use from that of efficiency change, we employ a model with time-varying technical efficiency, in which the movements away from the production surface over time are represented by a linear trend. Among others, Cornwell et al. (1990), Kumbhakar (1990), and Lee and Schmidt (1993) have considered other possible time-varying technical efficiency models for panel data. The distinctive advantage of Battese and Coelli's model is that in addition to modeling the time-varying inefficiency, it allows estimation of the effect of farm-specific variables on inefficiency.

The issue of explaining inefficiency has been of interest in many studies in agriculture (e.g., Bravo-Ureta and Pinheiro, 1993; Coelli and Battese, 1996).[4] Battese and Coelli's (1995) model avoids the inconsistency of a two-step approach to modeling the production frontier and inefficiency by modeling and estimating the frontier production function and the inefficiency effects model simultaneously. The ability to separate the factors affecting the frontier from the factors affecting the inefficiency comes at a price of explicit assumptions about the distribution of the error terms and inefficiencies.

In our study, the inefficiency explanatory variables are chosen to track the changes over time and to explain the variation across farms by the variation in farm organization, in managerial ability, in access to markets, and in availability of resources.

A Cobb-Douglas stochastic frontier production function is assumed to be the appropriate model for the analysis of the state and collective farm data for the two administrative regions.[5] The model estimated is defined by

$$Y_{it} = \beta_0 + \beta_{91} d_{91} + \sum_{j=1}^{5} \beta_j x_{ijt} + V_{it} - U_{it}\,, \tag{1}$$

where the subscript i indicates the observation for the i-th farm in the survey (i = 1,2,…,41), and the subscript t indicates the observation for the t-th year (t = 1,2,3,4). Y represents the logarithm of the total grain production (in metric tons) on the given farm in the given year; β_j, (j = 0,1,…,5, 91) represent the unknown parameters associated with the explanatory variables in the production function; d_{91} is a dummy variable, which has value 1 if t = 3, and value 0 otherwise; and the x_is (i = 1,2,…,5) represent the logarithms of the total amounts of land under grain production (in hectares), labor in grain production (in 1,000 hours), organic fertilizer applied for grain production (in 100 tons), chemicals applied for grain production (in tons), and diesel fuel used in grain production (in 1,000 liters), respectively.

The V_{it}s are assumed to be iid N(0, σ_V^2) random errors, independently distributed of the U_{it}s. The U_{it}s are nonnegative random variables, associated with technical inefficiency of production, which are assumed to be independently distributed such that U_{it} is obtained by truncation (at zero) of the normal distribution with variance σ_u^2, and mean, μ_{it}, where the mean is defined by

$$\begin{aligned}\mu_{it} = \delta_0 &+ \delta_1 \log(nonagw_{it} / totw_{it}) \\ &+ \delta_2 \log(agw_{it} / totland_{it}) \\ &+ \delta_3 \log(dis_i) \\ &+ \delta_4 t \\ &+ \delta_5 age_i + \delta_6 age_i^2 \quad ,\end{aligned} \tag{2}$$

where δ is a (7 x 1) vector of unknown parameters to be estimated. The variable $nonagw_{it} / totw_{it}$ is the ratio of the number of workers on the farm that are not involved in agricultural production to the total number of workers on the farm; $agw_{it} / totland_{it}$ is the number of agricultural workers on the farm per hectare of total farm land; dis_i is the distance from a given farm to a nearest city in kilometers; t is the year of observation (t = 1,2,3,4); and age_i is the age of the given farm manager in years.

The binary variable d_{91} is included into the stochastic frontier specification to account for poor weather conditions in 1991.

The *nonagw/totw* ratio reflects farm organization and measures the extent of the farm's self-reliance in the maintenance of its productive and social infrastructure. The higher this ratio, the less the farm depends on the state-provided infrastructure in its operation. A positive effect of this farm organization measure on technical efficiency would be consistent with the disorganization hypothesis. The higher the proportion of non-agricultural workers in the total number of workers, the less the farm relies on outside provision of production infrastructure. The less the farm relies on the outside provision, the less is the chance that a supplier will not fulfill a contract. The fewer disturbances from broken contracts, the smoother is the production process, and the fewer losses the farm incurs.

The *age* variable is included to check whether the younger and, presumably, more reform-oriented managers or the older, more experienced ones achieve higher levels of technical efficiency. Schultz (1975) argued that education and experience enhance a person's ability to deal with economic disequilibria. The ability is defined as being able to "perceive, to interpret correctly, and to undertake action that will appropriately reallocate their resources" (Schultz, 1975, p.827) in response to a changed economic environment. The random disturbances affecting production, such as weather changes, droughts, and diseases, alter an "equilibrium" of normal production. In this situation, better manager's ability to regain the "equilibrium" may lead to higher technical efficiency. We included the age^2 variable to account for a possible non-linear effect of age on the efficiency.

The distance to the nearest city was included to measure the effect, if any, of the access to markets. The *agw/totland* ratio is included in the model to control for the relative labor abundance of the farm. We think of this variable as a reflection of past farm performance (the more successful the farm was, the more people remain on this farm). The time variable captures the changes in inefficiency over time. The descriptive statistics of the variables used in estimation are presented in Table 3.

The parameters of the model, i.e. the β's, the δ's, and the variance parameters $\sigma^2 = \sigma_u^2 + \sigma_v^2$ and $\gamma = \sigma_u^2 / (\sigma_u^2 + \sigma_v^2)$, are simultaneously estimated using the method of maximum likelihood. We used program FRONTIER 4.1 developed by Coelli (1996) that computes the parameter estimates by maximizing a nonlinear function of the unknown parameters in the model subject to the constraints.

Results

Table 4 reports maximum likelihood estimation results. Several generalized likelihood-ratio tests regarding the stochastic frontier coefficients, inefficiency model, and variance parameters, are summarized in Table 5.[6]

Production Frontier

The generalized likelihood-ratio test failed to reject the composite hypothesis that second-order variables in the Translog model are zero. The test statistic $\lambda \overset{Approx.}{\sim} \chi^2_{15}$ took a value of 1.02, which is clearly smaller than the critical value 25.00. Consequently, we assumed that a Cobb-Douglas function is an adequate representation of the stochastic frontier function.

The elasticities of frontier output with respect to inputs were estimated to be 0.34 for land, 0.12 for labor, 0.07 for organic fertilizer, 0.22 for chemicals, and 0.27 for fuel. These results are not strictly comparable with most of the known studies on agricultural efficiency in Eastern Europe and the former Soviet Union, because the inefficiency models used in previous studies are not for grain, but rather for an aggregate farm output (either for the aggregate “crops,” or for the “whole farm” production function). In comparison, our land elasticity is much smaller than the value 0.71 found by Johnson et al. (1994) in a study of grain production efficiency; that study used a different model and relied on farm-level data in rubles for several inputs for the years 1986 to 1991. Wyzan (1981) found the land elasticity to be 0.62 when estimating a grain production function for the whole USSR with republic-level aggregated data for 1960 to 1976. At the same time, our land elasticity is close to the ones obtained for the Stavropol region of Russia, 1986 to 1988, by Skold and Popov (1990, 1992); and for Mongolia, 1986-1989, by Bayarsaihan et al. (1998); their estimates were in the range of 0.21 to 0.34.

The estimated labor elasticity of 0.12 falls in the range from 0.040 to 0.223 reported in the grain studies mentioned previously. In addition, labor elasticities in this range were reported in many of the efficiency studies that aggregated output [see, e.g., Hofler and Payne (1995) for Yugoslavia, 1961 to 1979; Brock (1994) for Russia, 1991; Carter and Zhang (1994) for nine centrally planned economies, 1965 to 1989]. The other input data (organic fertilizer, chemicals, and fuel) were not available in the previous grain efficiency studies.

Because the generalized likelihood-ratio tests of the hypotheses are preferred to the asymptotic t-tests in maximum likelihood estimation, the hypotheses that the coefficients of the corresponding input variables are zero were tested for each input variable separately. The tests were rejected, as reported in Table 5, which means that the impact of each of the input variables on the frontier production function is statistically significant. The returns-to-scale parameter was found to be 1.02, implying constant returns-to-scale for grain production on the state and collective farms. This result is consistent with earlier studies for Ukraine (Johnson et al., 1994) and for Russia (Skold and Popov, 1990, 1992).

Technical Inefficiency

The major interest of our study is the inefficiency model. Figure 1 provides frequency distributions of the efficiency estimates. The average technical efficiency in the sample was estimated as 0.82, 0.76, 0.68, and 0.60 for the four years of data (1989 to 92), respectively. Estimates in this range are found in earlier studies of inefficiency in agricultural production (Battese, 1992; Bravo-Ureta and Pinheiro, 1993). The null hypothesis that inefficiency effects are absent from the model is strongly rejected at the 5 percent level of significance, as is the joint null hypothesis that the explanatory variables in the model for the technical inefficiency effects have zero coefficients. The null hypotheses that individual effects of the explanatory variables in the model for the technical inefficiency effects are each zero were tested as well. The results presented in Table 5 show that all five null hypotheses were rejected.

The estimate for the variance parameter, γ, is estimated to be close to one. If this parameter is zero, then σ_u^2 in (1)-(2) is zero, and the model reduces to a traditional production function with the explanatory variables (*nonagw/totw*, *age*, *dis*, dis^2, *agw/totland*, and, *t*) all included in the production function. This would mean that inefficiency effects are not stochastic. The last null hypothesis, H_0: γ =0, which specifies that the explanatory variables in the model for the technical inefficiency effects are not stochastic, is rejected by the data. [7]

A comparison of technical efficiency of individual farms over the four years shows some consistency among the farms in terms of ranking by efficiency. Rankings by technical efficiency show a correlation between the years, with the correlation coefficients between one year and the next ranging from .64 to .72. In particular, the farms with lower technical efficiency ranks also had lower technical efficiency in subsequent years.

Inefficiency Effects Model

As expected, the estimated coefficient of year is positive, which means that technical efficiency declined over time, a result that is consistent with earlier findings obtained with a different model and a different Ukrainian farm data set (Johnson et al., 1994).

We found a positive effect of the self-reliance in provision of farm infrastructure on technical efficiency. This result is consistent with the disorganization hypothesis. The early stages of the transition were associated with less contract discipline that in turn worsened the ability of farms to achieve the maximum possible output. In this situation, vertical integration, which in the case of collective agriculture took the form of provision of some farming infrastructure on the farm, became a tool to alleviate the dependence of the farms on outside provision of inputs.

The number of agricultural workers per hectare was found to have positive effect on technical efficiency (i.e., a negative effect on inefficiency), which suggests that abundance of labor resources for production is important for achieving effective utilization of inputs. The model employed by Johnson et al. (1994) did not estimate the impact of farm-specific variables on inefficiency simultaneously with estimation of the stochastic frontier. But, they found a similar relationship by comparing the means of the number of agricultural workers per hectare for the fifty least and fifty most efficient farms (the number of farms in the sample was 3,798).

The effect of distance to the nearest city was found to be negative, i.e., *ceteris paribus*, farms located farther from the cities were more technically efficient (i.e., less inefficient). One interpretation is that the advantage in location may have allowed the farms located more distant from cities to do better in competition with cities for workers.

More energetic workers from rural farms located closer to cities could commute to city jobs, lowering the average skill/effort level of the available labor on these farms.[8] In addition, the farms located closer to cities had easier access to the less productive (in agricultural tasks) city workers and students recruited for harvest time. In this way, relative efficiency would be related to the distance from the city through its effect on the productive quality of farm labor, even if workers did not leave the rural area permanently. Unfortunately, the lack of additional data on commuters and temporary urban labor prevents further investigation of this argument, and the explanation offered remains only a conjecture.

The estimates of δ_5 and δ_6 imply a positive effect of the age of the manager on technical efficiency at the sample mean, i.e., other things equal, the older the manager, the less technical inefficiency the farm displayed. This is consistent with the results of Skold and Popov (1992), who found a similar, though weak, relationship between a manager's experience and technical efficiency in grain production for a sample of 136 Russian farms observed over the years 1986 to 1987. A similar positive impact of a manager's experience on technical efficiency has been found in studies on Third World agriculture (see the survey by Bravo-Ureta and Pinheiro, 1993).

The findings related to age can be explained by the manager's ability to deal with disequilibria. The process of reforms in formerly planned economies entails adjustment and reallocation of resources as the economy moves from one equilibrium state (the planned system) to another, more market-oriented, system. Under the old system, the production environment had remained stable over many years and farm managers knew from their experience how the system worked and what were the objectives and incentives of economic agents involved in production (i.e., the farm workers, local administration, government procurement agencies, and party officials). This knowledge allowed farm managers to successfully achieve their goals, such as maintaining appropriate social status and local power through fulfillment of state production plans. The production possibilities frontier was well known from experience, and the managers gained knowledge on how to organize, motivate, and monitor employees.

The start of the reforms meant major changes in the known economic environment: prices began to reflect the scarcity of economic resources, and financial results now played an increasingly important role in the valuation of farm performance. The altered possibilities increased demand for the managers' ability to make efficient choices to achieve a new equilibrium in the economic system with the changed rules, objectives, and constraints. The estimation results suggest, then, that those more experienced managers were better able to achieve technical efficiency.

The positive effect of age on technical efficiency, as expected, does not hold on all sample points. The marginal effect of the age on the mean μ_{it} is a decreasing function of age, and achieves zero around the age of 53. This result is consistent with the phenomenon that aging brings not only accumulated experience, but also health limitations. The concave effect of age on productivity is a common finding of human capital studies (Huffman), yet we found no technical efficiency study that considered the second-order age or experience term.

Allocative Efficiency

Allocative efficiency refers to the ability to choose the optimal input proportions given relative prices and output level. Ideally, when farm-level price data are available, efficiency can be estimated directly via estimation of dual (cost or profit) models. The absence of farm-level price data rules out this approach in this study. Therefore, we follow Bravo-Ureta and Rieger (1991) who suggest combining aggregate price data with a cost frontier derived analytically from the estimated production frontier under the assumption that the farms minimize the variable costs of expected production.

Following Koop and Diewert (1982), allocative efficiency is measured as the ratio of two costs: the variable costs when the cost-minimizing quantities of variable inputs $\hat{x}_{it}^{MC}$ are used, relative to the variable costs when the technically efficient quantities of variable inputs $\hat{x}_{it}^{TE}$ are used. The cost-minimizing quantities of variable inputs $\hat{x}_{it}^{MC}$ are estimated by analytically solving the cost-minimization problem under the assumptions that the frontier production function is as estimated, and that the target output $\widehat{Y}_{it}$ is the output observed adjusted for the statistical noise,

$$\begin{aligned}\widehat{Y}_{it} &\equiv Y_{it} - \widehat{V}_{it} \\ &= \widehat{\beta}_0 + \widehat{\beta_{91}} d_{91} + \sum_{j=1}^{5} \widehat{\beta}_j x_{ijt} - \widehat{U}_{it}\end{aligned}.$$

Here "hats" over the betas, the V_{it}, and the U_{it} refer to the estimates obtained using the method of maximum likelihood discussed previously. The technically efficient quantities of variable inputs $\hat{x}_{it}^{TE}$ are estimated by equiproportionately scaling back the observed variable input quantities so that the output $\widehat{Y}_{it}$ can still be produced.

In Ukraine, an analysis of allocative efficiency faces important methodological issues because of distorted input markets and high inflation. First of all, there are no land or organic fertilizer markets in Ukraine; consequently, no land or fertilizer prices are available. We circumvent this obstacle by treating the observed quantities of these inputs as quasi-fixed, but the resulting allocative efficiency measures are inevitably "partial" in the sense that they reflect optimization behavior with respect to some, but not all, inputs. A second, related issue is that even when the markets exist, because of rationing, bonuses, and compensatory payments, the collective and state farms do not always pay the input prices as reported in the official statistics. Given the existing input distribution system, farms may not respond to the official relative prices if they do not reflect the relative scarcity of resources as *faced by the farms*. In particular, our data show little change in fuel use over time—a finding that suggests great rigidity in allocation of this input, even in 1992. Consequently, we treat fuel as a quasi-fixed input in the allocative efficiency analysis. Thus, our analysis is based on only two variable inputs, labor and chemicals.

Finally, given that prices began to rise rapidly in 1990, there is a question of which period prices are relevant for decision-making. We have assumed that the 1989 prices are suitable for the 1989 year, because these prices had remained stable over many pre-reform years. But for 1992, we do not have any information on exactly when the allocation decisions were made, a time that is crucial given high inflation. The mid-calendar-year price estimates were used for 1992 because they correspond roughly to the beginning of an agricultural production year.

The price data come from various World Bank publications (World Bank, 1993, 1994). Between 1989 and 1992, the nominal price of plant protection chemicals grew 37.2 times, whereas that of labor grew only 17.7 times. We found a slight increase in allocative efficiency from the mean of 0.808 in 1989 to the mean of 0.836 in 1992. Given the data limitations, however, we have reservations about these findings. Correlation analysis suggests that the two types of efficiency are positively related in a given year: the coefficient of correlation between estimated farm-specific 1989 technical and 1989 allocative efficiencies is 0.20 (p-value of 0.21), and that for 1992 is 0.08 (p-value of 0.60). Also, the farms that were relatively allocatively efficient in 1989 were likely to choose the right input mix in 1992 as well (estimated correlation coefficient 0.33 with a p-value of 0.04). These results suggest that, similar to the technical efficiency, allocative efficiency is farm-specific.

Discussion and Concluding Remarks

Grain production and input use data in physical units, together with overall farm operations information, were used to estimate a stochastic frontier model in which inefficiency effects were modeled as a function of farm-specific variables and time. The magnitudes of the production function and efficiency estimates are consistent with other findings obtained for the formerly planned economies with different models and different data sets. The results illustrate that the process of transformation of Ukrainian agriculture begun in 1990 has been costly in terms of technical efficiency: efficiency declined over the years 1989 to 1992. The relative abundance of labor and the distance to a nearest city were both found to have a positive effect on technical efficiency. Our results indicate that the traditional production function model is likely to be inadequate for the farm-level analysis of grain production.

Investments of the farms’ labor resources in infrastructure also improved technical efficiency. Unfortunately, the data used do not discriminate among different types of farms’ non-agricultural production activities. These activities could include investment in other areas of production (e.g., production facilities construction, processing, and marketing) or improvement in farm living conditions (e.g., catering, child care provision, and road maintenance). Improved farm living conditions would likely increase the quality of available labor resources directly by improving human capital (health) and reducing shirking and absenteeism due to health and child care problems.

The average quality of labor resources also may be affected indirectly through preventing productive workers from quitting. The potential loss of farm-provided social benefits is considered one of the main reasons for a farm employee's decision to remain on the collective farm as opposed to starting his or her own private farm (Lerman et al., 1994). Hence, investment into farm social infrastructure could be a valuable tool used by farm managers to restrain workers from leaving the farms. Losing workers is undesirable because those who leave, on average, possess above-average skills, both general and agriculture-specific.[9]

Independently of whether the increase in the share of farm non-agricultural production activities means more production or more social workers, the increase in the share means also that the farm-provided jobs not only enhance farm production and/or general infrastructure but also lead to additional income for the rural population. The importance of this additional income is underscored by the shrinking scale of main production as outlined previously, and consequently, lower farm revenues and wage bill. Moreover, the state, the major buyer of the farm output, has been consistently late with payments for produce. Wage arrears have begun to cause distress for the agricultural enterprises. The World Bank surveys showed that more than half of the farms were unable to meet their payrolls on time at least once in 1993. The additional non-agricultural production jobs would have provided greater income security for farm families, thus reducing further the possibility of families' leaving collective farms. In sum, the farm's non-agricultural production activities may make collective farms a more attractive place for living, and in this way may keep the average skill level of workers from declining.[10]

Our study complements that of Blanchard and Kremer (1997) who studied the effect of disorganization caused by economic transition on decline in industrial output. In particular, using country-level data for production of 159 goods, they found strong empirical evidence that output had fallen furthest for the goods with the most complex production process. That is, the more complex the production process, the more suppliers the production depends on, and hence the more important is the role of contracts (implicit or explicit). Also, the more suppliers the production depends on, the higher are the chances that the contracts will be broken. With bad contract discipline in the transition period, the feasible way to protect a firm from the losses due to broken contracts is to integrate vertically.

In our case, the units of observations, the farms, do not differ in production processes. However, the share of non-agricultural workers in the total number of farm workers varies positively with reduced reliance on outside provision of production inputs. Our findings may be interpreted as empirical support for the positive effect of vertical integration as a loss-prevention tool in the transition period.

The results illustrate that the introduction of reforms has not immediately reversed the decline in efficiency in Ukraine's agriculture. Moreover, the more efficient farms were found to be less market-oriented, a result associated with maintenance of farm infrastructure. Further research is needed on how farm organization may affect agricultural production efficiency in Ukraine and other countries undergoing economic transition. The results highlight the importance of analysis of production at the farm level because production efficiency varies across farms and this finding should be taken into account for both research and policy considerations.

What policy implications can be drawn from our findings? One of the major current agricultural policy issues debated in Ukraine is how to approach restructuring of the collective sector and, in particular, whether or not to break up the large collective farms. Our findings on virtually constant returns-to-scale suggest that from the purely technical point of view, breaking up the farms may not have a major effect. Rather, our results suggest shifting the attention to another area—farming infrastructure. We found support for the great relative importance of farming infrastructure in the transition period. If we were to advise a farm manager, we would recommend vertical integration to prevent inefficiencies resulting from broken contracts. If we were to advise a policy maker, we would recommend actions on development of input supply markets and improvement of overall contractual discipline.

Table 1. General farm-level indicators [a]

Indicator	Units	1989	1990	1991	1992	Avg. change per farm 1989 to 1992
Agricultural experience of manager	years				25.6	
					(7.4)	
Agricultural land	hectares	2403	2350	2236	2196	-7.6%
		(971)	(948)	(857)	(830)	(9.6%)
Total farm workers	number	384	376	359	335	-13%
		(123)	(121)	(118)	(120)	(10%)
Agricultural workers	number	327	317	305	290	-11%
		(97)	(97)	(97)	(98)	(12%)
Ratio of non-agricultural to total	number	0.145	0.154	0.146	0.129	-8%
farm workers		(0.053)	(0.054)	(0.053)	(0.050)	(26%)
Agricultural workers per hectare of	number	0.143	0.142	0.142	0.137	-3%
agricultural land		(0.033)	(0.032)	(0.029)	(0.030)	(15%)
Share of agricultural land under grains	number	0.482	0.484	0.485	0.458	-5%
		(0.053)	(0.059)	(0.053)	(0.056)	(10%)
Pensioners	number	369	370	372	373	1%
		(130)	(132)	(133)	(130)	(10%)
Ratio of pensioners to total farm workers	number	0.97	0.99	1.05	1.14	17%
		(0.22)	(0.24)	(0.24)	(0.30)	(17%)
Ratio of non-agricultural to total	number	0.047	0.078	0.098	0.196	600%
farm expenditures		(0.045)	(0.087)	(0.104)	(0.155)	(815%)

[a]All the indicators reported are average per farm, the numbers in parentheses are the standard deviations.

Source: UIAE Survey of Ukrainian farms.

Table 2. Partial productivity and input use indicators [a]

Indicator	Units	1989	1990	1991	1992	Avg. change per farm 1989 to 1992
Yield	tons per hectare	4.27	4.00	2.88	3.18	-25.4%
		(0.79)	(0.73)	(0.60)	(0.74)	(8.5%)
Production	tons	4936	4562	3205	3183	-35%
		(2690)	(2558)	(1796)	(1778)	(10%)
Area planted	hectares	1173	1149	1104	1020	-12%
		(550)	(542)	(510)	(469)	(10%)
Labor per hectare planted	hours	24	27	28	35	93%
		(12)	(13)	(15)	(17)	(184%)
Fertilizer per hectare planted	kilograms	7603	7624	6598	6385	-16%
		(5952)	(5950)	(5264)	(5556)	(23%)
Chemicals per hectare planted	kilograms	6.5	6.3	5.7	5.2	-19%
		(1.5)	(1.6)	(1.4)	(1.4)	(15%)
Fuel per hectare planted	liters	83	83	82	82	-0.9%
		(14)	(14)	(13)	(12)	(5.7%)

[a] All the indicators reported are average per farm, the numbers in parentheses are the standard deviations.

Source: UIAE Survey of Ukrainian farms.

Table 3. Summary statistics for variables in the stochastic frontier production function[a]

Variable	Units	Sample Mean	Sample Standard Deviation	Minimum	Maximum
Production	tons	3972	2361	1219	18574
Land	hectares	1112	517	268	2850
Labor	1,000 hours	32	29	6	219
Fertilizer	100 tons	79	78	14	596
Chemicals	tons	6.6	3.7	1.6	21.4
Fuel	1,000 liters	93	51	24	285
Ratio of non-agricultural to total workers (*nonagw / totw*)	number	0.143	0.053	0.041	0.317
Agricultural workers per agricultural land (*agw / totland*)	number per hectare	0.141	0.031	0.081	0.245
Distance to city (*dis*)	km	35	16	10	85
Manager's age (*age*)	years	46.9	8.2	30	65

[a] 41 farms, 4 years, 164 observations in total.

Table 4. Maximum-likelihood estimates for parameters of the stochastic frontier production model for farms of Kyivs'ka and Cherkas'ka oblasti

Variable	**Parameter**	**Estimate**	**Standard Error of Estimator** [a]
Stochastic Frontier			
Constant	β_0	3.99	0.32
Year 1991 Dummy	β_{91}	-0.205	0.027
ln (land)	β_1	0.338	0.089
ln (labor)	β_2	0.119	0.018
ln (fertilizer)	β_3	0.071	0.030
ln (chemicals)	β_4	0.220	0.032
ln (fuel)	β_5	0.271	0.067
Inefficiency Model			
Constant	δ_0	0.24	0.36
ln (Non-ag./Total Workers)	δ_1	-0.085	0.039
ln (Ag. Workers/Total Land)	δ_2	-1.00	0.15
ln (Distance to City)	δ_3	-0.211	0.070
Year	δ_4	0.093	0.012
Age	δ_5	-0.033	0.015
Age Squared	δ_6	0.00031	0.00016
Variance Parameters			
	σ^2	0.0175	0.0023
	γ	1.0000	0.0029
ln (Likelihood)		114.96	

[a] The standard errors for the estimators are obtained by the computer program Frontier 4.1; they are correct to two significant digits.

Table 5. Generalized-likelihood ratio tests of hypotheses involving parameters of the stochastic frontier inefficiency model[a]

Null Hypothesis	Meaning of Hypothesis	ln (H_0)	λ	D.F.	Critical Value	Decision
Stochastic Frontier						
H_0: $\beta_1 = 0$	Var. *land* does not affect stochastic frontier	105.24	19.44	1	3.84	Reject H_0
H_0: $\beta_2 = 0$	Var. *labor* does not affect stochastic frontier	100.39	29.13	1	3.84	Reject H_0
H_0: $\beta_3 = 0$	Var. *fertilizer* does not affect stochastic frontier	109.75	10.42	1	3.84	Reject H_0
H_0: $\beta_4 = 0$	Var. *chemicals* does not affect stochastic frontier	96.48	36.96	1	3.84	Reject H_0
H_0: $\beta_5 = 0$	Var. *fuel* does not affect stochastic frontier	110.62	8.69	1	3.84	Reject H_0
Inefficiency Model						
H_0: $\gamma = \delta_0 = ... = \delta_6 = 0$	Inefficiency effects are absent from the model	54.83	120.26	8	14.85	Reject H_0
H_0: $\delta_1 = ... = \delta_6 = 0$	Inefficiency effects are not a linear function of the Explanatory variables	57.29	115.34	6	12.59	Reject H_0
H_0: $\delta_1 = 0$	*ln (nonagw/totw)* does not affect inefficiency linearly	69.04	91.84	1	3.84	Reject H_0
H_0: $\delta_2 = 0$	*ln (agw/totland)* does not affect inefficiency linearly	68.25	93.43	1	3.84	Reject H_0
H_0: $\delta_3 = 0$	*ln (dis)* does not affect inefficiency linearly	107.94	14.04	1	3.84	Reject H_0
H_0: $\delta_4 = 0$	Inefficiency does not change linearly with time	86.98	55.96	1	3.84	Reject H_0
H_0: $\delta_5 = \delta_6 = 0$	*Age* does not affect inefficiency	75.78	78.36	2	5.99	Reject H_0
H_0: $\delta_6 = 0$	*Age* does not affect inefficiency quadratically	112.73	4.46	1	3.84	Reject H_0
H_0: $\gamma = 0$	Inefficiency effects are not stochastic	109.63	10.65	2	5.14	Reject H_0

[a] The critical values correspond to 5 percent level of significance.

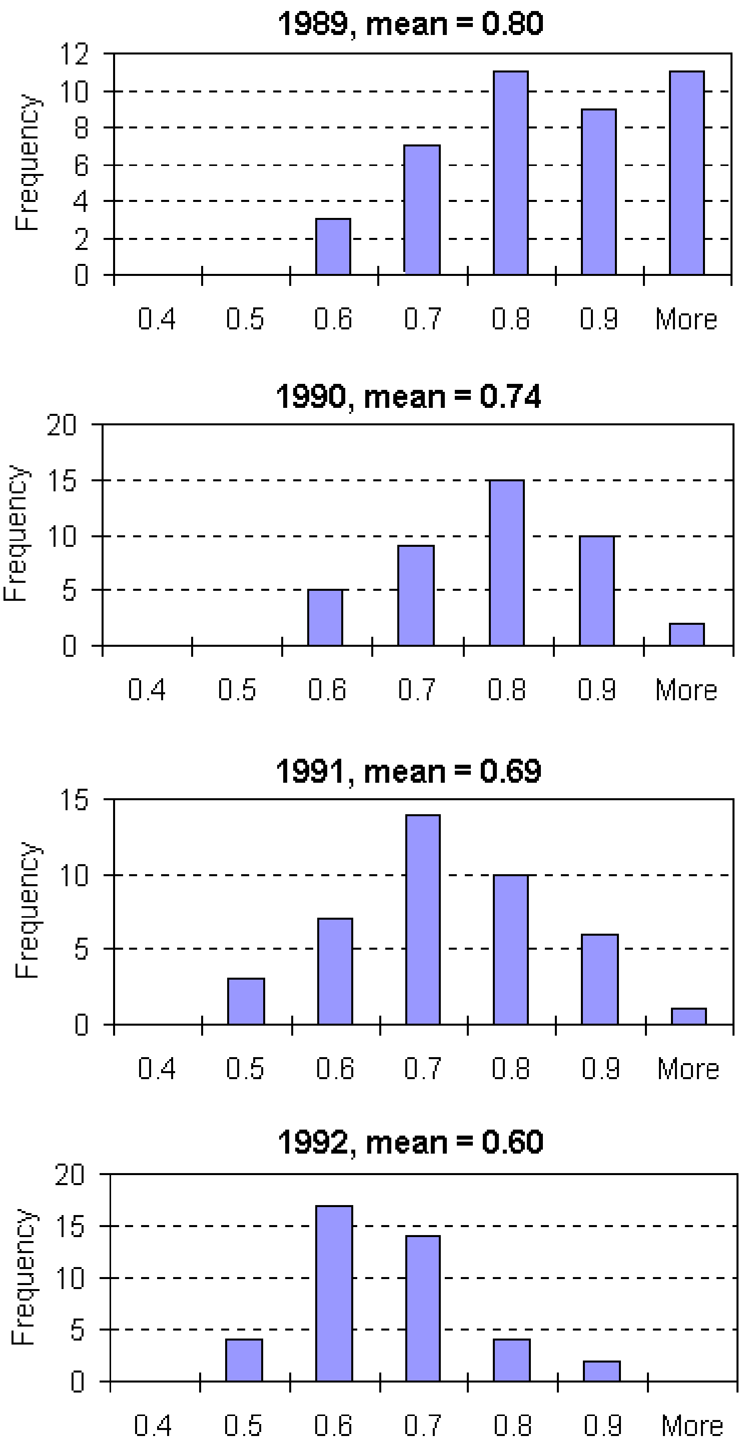

Figure 1. Frequency distribution of technical efficiencies by years

Endnotes

1. The GDP and gross agricultural product figures ought to be taken with a caution as both the price regimes and producer incentives differed substantially between 1990 and 1996. We thank an anonymous reviewer for drawing our attention to this point. Nevertheless, the figures reflect a large decline in production in both the agricultural sector and in the economy as a whole.

2. We tested the null hypotheses specifying that the means of our sample are equal to the corresponding means of the Mixed zone collective and state farm census for 9 variables: farm land, total number of workers, and the number of agricultural workers, each for the years 1989, 1990, and 1991. Using a standard t-test, we failed to reject any of the 9 null hypotheses. Results of the tests are available from the authors.

3. Prices doubled from 1990 to 1991 and increased almost 20-fold from 1991 to 1992 (IMF, 1993).

4. Earlier studies adopted a logically inconsistent two-step approach, in which at the first step, stochastic frontier production function was estimated and the inefficiencies were predicted under the assumption that the inefficiency effects are identically distributed. Yet at the second step, the assumption that the means of the inefficiencies are the functions of some farm attributes was introduced, and the inefficiency predictions were regressed on the candidate explanatory variables.

5. A translog specification was tested and rejected by the data. See Results Section on the particulars of the test.

6. The generalized likelihood-ratio statistic is computed as $\lambda = -2\log\left[L(H_0)/L(H_1)\right]$, where $L(H_0)$ and $L(H_1)$ are the likelihood functions evaluated at the restricted and unrestricted maximum-likelihood estimator for the parameters of the model. If the null hypothesis, H_0 , is true, and does not involve $\gamma = 0$, then the statistic has approximately chi-squared distribution with parameter equal to the number of restrictions imposed by H_0. If a null hypothesis includes $\gamma = 0$, then, since by its definition γ has to be non-negative, the statistic has asymptotically a mixed chi-squared distribution (Coelli, 1995a). Koddle and Palm (1986) provide critical values for the statistics in such tests.

7. In this case, the parameter δ_0 is not identified, and consequently, the number of degrees of freedom for the test statistic is two.

8. An estimated 160,000 workers commuted to the city of Kyiv from nearby rural communities at the beginning of the 1990s (Bohdan, 1992).

9. According to official statistics, the rural share of the total working age population has remained stable, around 28.4 percent, over the past several years. In rural areas, the major alternative to the former state or collective farm employment would be private farming. Csaki and Lerman (1997b) found that private farmers are, on average, better educated than collectivist farm employees, suggesting that the departure of farm workers to private farming would lower the average level of education of collectivist farm workers. As education and experience became increasingly important in successful adjustment to the rapidly changing economic environment of Ukraine, the departure of the more highly educated workers from the collectivist farms may have contributed to the lower productivity on those farms.

10. Following this argument, another explanatory variable of inefficiency, the number of agricultural workers per hectare, captures a part of the effect of the share of non-agricultural workers in the total number of farm workers on inefficiency. Further separation of the effects of these two explanatory variables on technical inefficiency would require specification of a labor mobility model. The data available are not able to support this analysis.

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