Production Risk and Technical Inefficiency in Russian Agriculture

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Summary

This paper aims to contribute to a better understanding of possible causes of considerable production variability that characterised Russian agriculture during the last decade. The study investigates production risk and technical inefficiency as two sources that influence production variability. Using panel data from 1996 to 2001, an empirical analysis of 443 large agricultural enterprises from three regions in central, southern and Volga Russia is conducted. A production function specification accounting for the effect of inputs on both risk and technical inefficiency is found to describe production technologies of Russian farms more appropriately than the traditional stochastic frontier formulation.

Key words: Production risk, Technical efficiency, Panel data, Russian agriculture

JEL Classification: D81, Q12

1. Introduction

During transition period, the development of Russian agricultural production has exhibited a remarkably incoherent character. In general, production declined over a considerable period while at the same time a serious output variation was observable[1]. Numerous studies have been conducted in order to identify the possible causes of agricultural production decline in post-Soviet Russia (Serova, 2000, Macours and Swinnen, 2000, Liefert, 2002, Bezlepkina and Oude Lansink, 2003). Among others, deterioration in terms of trade, the elimination of producer and consumer subsidies, a weak institutional environment and undeveloped factor markets were revealed as determinants of production slowdown. Additionally, Liefert (2002) specifies weather as an important factor of production volatility in Russia; he states: "weather is volatile in Russia, such that crop harvests can vary very substantially between years". However, so far the literature has paid little attention to high production variability and its effects on the production development in Russian agriculture.

In recent years (from 1999 to 2002) Russian agricultural production has exhibited substantial growth, followed by a deceleration in 2003. Many factors have positively contributed to such a development. Brooks and Gardner (2004) refer to improved macroeconomic stability, increased demand for domestic food after the financial crisis in 1998, and an increased interest in agricultural investments from the side of more entrepreneurial producers. These factors must have induced an increase in farm productivity through their impact on technical change and technical efficiency. On the other hand, the literature (Gaidar, 2002) denotes favourable weather conditions, particularly in 1999 and 2000, as a determinant of the recent growth in agricultural production in Russia. However, weather has volatile effects on agricultural production and cannot have a long-term positive effect on farm productivity. Therefore, in these circumstances it is essential to distinguish between factors stabilizing farm productivity in the long term and volatile weather effects that can positively contribute to production growth only for short periods.

In this context the objective of this study is to analyze: whether Russian farms increased production by enhancing their productivity, and technical efficiency in particular, in the later 1990s and the beginning of the 2000s as well as to investigate the extent to which the recent growth can be attributed to some reduction in production risk due to favourable weather conditions in this period.

There have been a number of empirical investigations concerning the development of agricultural productivity in Russia. Osborne and Trueblood (2002a) found that multifactor productivity of corporate farms in Russia declined by 1.7 percent per year in the period from 1993 to 1998. A similar result for the same period was reported by Voigt and Uvarovsky (2001). On the other hand, Lerman et al. (2003) estimated that multifactor productivity rose by 7.4 percent from 1992 to 1997. This contradiction of results of different studies could be traced back to differences in production conditions in the individual years of the considered periods, especially the last years in these studies. One of the best years in the 1990s was 1997 – "a good weather and harvest year" (Liefert, 2002), when grain production reached 88.6 million tons (Goskomstat, 2002a). However, 1998 was a drought

[1] Appendix A demonstrates the development of grain production, sown areas and yields from 1980 to 2003. A substantial decline of grain production in the last decade was attributed to a gradual drop in sown grain areas, but grain yields rose, on average, compared to the pre-reform period. On the other hand, high yield variability is remarkable in both periods: in the 1980s as well as during transition (Appendix B).

year in most grain producing regions of Russia, with grain harvest amounting to only 47.9 million tons (Goskomstat, 2002a).

Sotnikov (1998) and Sedik et al. (1999) were the first to study technical efficiency in Russian agriculture during the reform era. In both studies, the authors estimate the magnitudes of technical efficiency on the oblast level. The estimates employ the stochastic frontier approach. In addition, Sedik at al. (1999) carry out data envelopment analysis. The studies provide analogous results and show that technical efficiency declined from 1991 to 1995. A study by Osborne and Trueblood (2002b) analysed Russian crop production between 1995 and 1998, and showed that the trends revealed in the earlier studies have slowed down but have not been reversed. In contrast, the estimates of technical efficiency in 75 Russian regions obtained by Voigt (2002) do not suggest serious changes in technical efficiency at the national level from 1993 to 2000. In addition, he found that the development of technical efficiency in different regions does not have any common trend.

Recently, several studies estimated technical efficiency using farm level data. Bezlepkina and Oude Lansink (2003) study technical efficiency of dairy farms in the Moscow region and consider their development with regard to capital structure and subsidising programs from 1996 to 2000. The study results show that even though technical efficiency decreases considerably in the year of financial crisis, 1998, in general it has a positive trend in the analysed period. These results are consistent with the findings by Stange and Lissitsa (2004) who compare technical efficiency of farms in the same region with regard to their specialization, size and form of organization in the years 1993 and 2000. The results of both studies suggest an increase in technical efficiency of the considered farms in recent years. However, the farms of the Moscow region located near the city can hardly be considered representative of Russian agriculture. In this context, further investigation is necessary to assess the current stage of technical efficiency development in Russian agriculture.

The studies of technical efficiency of Russian agricultural producers differ with respect to estimation techniques and subject of investigation. Additionally, many particularities can be found with regard to the objectives and background of the individual studies. However, neither of the studies analysing the development of agricultural production in Russia took proper account of the presence of risk and the farmers' responses to it, whereas it is common knowledge that economic units make their decisions under uncertainty.

The presence of risk influences not only production output but also producers' behaviour, primarily with regard to input use. If risk mitigation plays a principal role in decision-making, then a farm's technical efficiency score may alter significantly. Therefore, technical efficiency assessed considering a producer's response to uncertainty is not the same in a setting where no effect of risk on input-use decisions is taken into account. Thus, when uncertainty is pervasive, the theoretical framework for studying technical efficiency is to be extended with respect to risk and producers' responses to risk. In this study, production risk is assumed to be an important factor in Russian agriculture and to influence production decisions of Russian farmers. Hence, the present study aims to estimate the magnitudes of both technical inefficiency and production risk faced by agricultural producers in Russia and therefore to explain the pattern of Russian agricultural production development in the last decade.

Two approaches are employed in the study: the Just and Pope model (1978), and a Kumbhakar extension of this model to introduce technical efficiency (Kumbhakar, 2002). The Just and Pope model allows to distinguish between the effects of input use on production output and production risk (1978). Technical efficiency explained by a complementary function presents an additional source of production variability (Kumbhakar, 2002). Both models are extended to consider systemic production risk and are estimated using panel data (from 1996 to 2001) of 443 large agricultural enterprises from three regions in central, southern and Volga Russia. Based on the estimation results, two hypotheses with respect to the study objectives will be discussed:

- Agricultural production in Russia is subject to considerable production risk; besides systemic weather effects, output variance depends on the intensity of input use.

- Technical inefficiency (TI) enhances output variability in Russian agriculture; TI is farm specific and can be explained by input use.

The paper is organized as follows. Section 2 outlines the methodology applied to distinguish and assess two sources of production variability: production risk and technical inefficiency. Section 3 presents the specification of the models used in the study. Data and estimation results with regard to the study objectives are discussed in Section 4. Conclusions are drawn in the final section.

2. Theoretical framework

This study employs stochastic frontier analysis (SFA)[2], which requires a parametric representation of the production technology. In addition, it considers output variability by a two-part error term. The distributional assumptions for both parts of the error term have to be imposed. This approach was pioneered by Aigner, Lovell and Schmidt (1977). The general notation of the model is the following:

$$y_i = f(\mathbf{x}_i;\boldsymbol{\alpha})e^{v_i} TE_i , \tag{1}$$

where y_i is the output of producer i ($i \in \mathrm{I}$), $\mathbf{x}_i$ is a vector of inputs used by producer i, $\boldsymbol{\alpha}$ represents a vector of technology parameters, $f(\mathbf{x}_i;\boldsymbol{\alpha})$ is the production frontier, and TE_i is the output-oriented technical efficiency of producer i. In addition, v_i represents a producer-specific random component.

Technical efficiency is defined as the ratio of observed output to maximum feasible output in a state of nature depicted by $\exp\{v_i\}$:

$$TE_i = \frac{y_i}{f(\mathbf{x}_i;\boldsymbol{\alpha})e^{v_i}} . \tag{2}$$

However, the conventional specification of a stochastic production function has a feature which may seriously restrict its potential to depict production technology appropriately. An important disadvantage of the traditional multiplicative stochastic specification of production technology is the implicit assumption that if any input has a positive effect on output, then a positive effect of this input on output variability is also imposed. Just and Pope (1978) showed that the effects of input on output should not be tied to the effects of input on output variability *a priori*. Instead, they proposed a more general stochastic specification compared to the usual econometric production function approach. Accordingly, the adequate production function specification has to include two general functions: one which specifies the effects of the input on the mean of output and another which specifies the effect of input on the variance of the output:

$$y_i = f(\mathbf{x}_i;\boldsymbol{\alpha}) + g(\mathbf{x}_i;\boldsymbol{\beta})v_i , \tag{3}$$

where, $f(\mathbf{x}_i;\ \boldsymbol{\alpha})$ is the mean production function and $g(\mathbf{x}_i;\ \boldsymbol{\beta})$ is the variance production function. Furthermore, $\boldsymbol{\alpha}$ is a vector of the mean production function parameters, $\boldsymbol{\beta}$ is a vector of the variance production function parameters and v_i is a stochastic term assumed to be $i.i.d.N(0,\ 1)$[3]. Thus, $E(y) = f(\mathbf{x})$, *and* $V(y) = g^2(\mathbf{x})$. Accordingly, the effect of input changes has been separated into two effects - the effect on mean and the effect on variance. Since the variance of y is specified as a function of the production inputs $g(\mathbf{x}_i;\boldsymbol{\beta})$, the Just-Pope production function exhibits heteroscedasticity.

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[2] Empirical studies on efficiency usually utilize either Data Envelopment Analysis (DEA) or SFA. DEA is a non-parametric approach and employs linear programming to construct a piecewise-linear, best-practice frontier for each economic unit (Färe et al., 1985). No functional form for the frontier is imposed on the data. However, most non-parametric programming approaches to estimate production frontier consider production to be deterministic and do not regard the possibility of noisy data by assumption. Deviations from the frontier are considered as inefficiency, though some authors have dealt with this issue by introducing elements of statistical analysis in the DEA models (Simar and Wilson, 1995; Gstach, 1996). Since stochastic specification of the production frontier model permits taking into account random shocks that affect production but lie outside producer control, SFA is a more appropriate approach for an environment characterized by considerable random effects.

[3] i.e., independent and identically distributed standard normal random variable.

The marginal production risk, defined as

$$\frac{\partial\,\mathrm{var}(y)}{\partial x_j} = 2g(\mathbf{x};\boldsymbol{\beta})g_j(\mathbf{x};\boldsymbol{\beta}) \tag{4}$$

can be positive as well as negative, or zero, subject to the signs of $g(\mathbf{x}_i;\boldsymbol{\beta})$, and $g_j(\mathbf{x}_i;\boldsymbol{\beta})$, where the latter is the partial derivative of g with respect to input j.

Generally, there are 3 possibilities for integrating technical efficiency into the Just-Pope production function:

(i) in additive form (Battesse at al., 1997). In this case it is attached to the variance production function, together with the random term representing production uncertainty:

$$y_i = f(x_{ij};\alpha_j) + g(x_{ij};\beta_j)(v_i - u_i)\,; \tag{5}$$

(ii) in multiplicative form. Here, technical efficiency is attached to the mean production function (Kumbhakar, 2002):

$$y_i = f(x_{ij};\alpha_j)(1-u_i) + g(x_{ij};\beta_j)v_i\,, \tag{6}$$

In this case an additional assumption: $exp\{-u\}=1-u$ has to be introduced.

(iii) in the more flexible form suggested by Kumbhakar (2002), where an additional function q(x) for explaining technical inefficiency is introduced:

$$y_i = f(x_{ij};\alpha_j) + g(x_{ij};\beta_j)v_i - q(x_{ij};\gamma_j)u_i\,. \tag{7}$$

Equations (5) and (6) are special cases of (7). Depending on the choice of the $q(\mathbf{x})$ function, the model in (7) can be reduced to (5) when $q(\mathbf{x})=g(\mathbf{x})$ or to (6) when $q(\mathbf{x})=f(\mathbf{x})$.

3. Model specification

In this study, two model specifications are considered: the Just and Pope model (JP-model), and a Kumbkakars' extension of the model by considering technical efficiency as provided by (7). These specifications are extended by introducing variables that account for a systemic part of production risk (SPR) and by applying them to panel data. In the following, the subscripts $i = 1,..,N$ and $t = 1,...,T$ denote the producer and the time period, respectively. Defining $\mathbf{x} = [x_1,...x_J]$ the production function can be written as

$$y_{it} = f(\mathbf{x}_{it};\boldsymbol{\alpha}) + \exp(\mathbf{D}\beta_t)g(\mathbf{x}_{it};\boldsymbol{\beta})v_{it} \qquad \textit{(Just and Pope with SPR)} \tag{8}$$

$$y_{it} = f(\mathbf{x}_{it};\boldsymbol{\alpha}) + \exp(\mathbf{D}\beta_t)g(\mathbf{x}_{it};\boldsymbol{\beta})v_{it} - q(\mathbf{x}_{it};\boldsymbol{\gamma})u_{it} \qquad \textit{(Kumbhakar with SPR)} \tag{9}$$

where, v_{it} is assumed to be *i.i.d.* $N(0,1)$, , and u_{it} is *i.i.d.* $N^{+}(0,\sigma_u^2)$. The function $g(\mathbf{x}_{it},\boldsymbol{\beta})v_{it}$ represents the idiosyncratic component of production risk faced by selected farms. Systemic production risk is captured by a vector **D**, which consists of dummy variables for the individual years (Greene, 2003). Thus, β_t can be viewed as a proxy for the systemic component of risk, which expresses a spatial effect of annual weather conditions on production variance for the entire sample.

In the case of the model specification with TI, the mean production function and production variance function are defined at the frontier, i.e., $u_{it}=0$. Thus, for both approaches

$$E(y|_{u=0}) = f(\mathbf{x}),\ \ V(y|_{u=0}) = \{\exp(\mathbf{D}\beta_t)\, g(\mathbf{x})\}^{2}. \tag{10}$$

A single-step maximum likelihood (ML) procedure was employed to estimate the parameters of the specified models. Taking into consideration the distributional assumptions on v und u, the likelihood function of TN observations is formulated as the product of the probability density functions $f(\boldsymbol{\varepsilon}_{it})$ of TN single observations and the Jacobian $|J|$ of the undertaken transformation (ε from y):

$$L = \prod_{i=1}^{N} f(\boldsymbol{\varepsilon}_i) * |J|, \text{ with } \boldsymbol{\varepsilon}_i = [\varepsilon_{i1},...,\varepsilon_{iT}] \text{ and } f(\boldsymbol{\varepsilon}_i) = \prod_{t=1}^{T} f(\varepsilon_{it}) \tag{11}$$

where $\varepsilon_{it} = \dfrac{y_{it} - f(\mathbf{x}_{it})}{\exp(\mathbf{D}\beta_t)g(\mathbf{x}_{it})} = [v_{it} - h(\mathbf{x}_{it})u_{it}]$ with $h(x_{it}) = \dfrac{q(\mathbf{x}_{it})}{\exp(\mathbf{D}\beta_t)g(\mathbf{x}_{it})}$.

The probability density function of ε_{it} is

$$f(\varepsilon_{it}) = \sqrt{\frac{2}{\pi}} \frac{1}{\sigma_{it}} \Phi\left\{ -\frac{\varepsilon_{it}\sigma_u h(\mathbf{x}_{it})}{\sigma_{it}} \right\} \exp\left\{ -\frac{1}{2} \frac{\varepsilon_{it}^{\ 2}}{\sigma_{it}^{\ 2}} \right\} \tag{12}$$

with $\sigma_{it}^{\ 2} = 1 + h^2(\mathbf{x}_{it})\sigma_u^2$ and $\Phi(\cdot)$ being the distribution function of the standard normal random variable (Kumbhakar, 2002). The Jacobian in our case is a *TN*x*TN* diagonal matrix with the elements $\dfrac{1}{\exp(\mathbf{D}\beta_t)g(\mathbf{x}_{it})}$.

Then, the log-likelihood function to be estimated is

$$\ln(L) = const - \sum_{t=1}^{T} \left(\frac{1}{2}\sum_{i=1}^{N} \ln \sigma_{it}^{\ 2} + \sum_{i=1}^{N} \ln\left[\Phi\left\{ -\frac{\varepsilon_{it}\sigma_u h(\mathbf{x}_{it})}{\sigma_{it}} \right\} \right] - \frac{1}{2}\sum_{i=1}^{N} \frac{\varepsilon_{it}^{\ 2}}{\sigma_{it}^{\ 2}} - \mathbf{D}\beta_t - \sum_{i=1}^{N} \ln g(\mathbf{x}_{it}) \right) \tag{13}$$

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The maximization of the log-likelihood function in (13) provides the ML estimates of the parameters in *f*(**x**), *g*(**x**) and *q*(**x**), as well as of σ_u (Greene, 2003). They can be used to calculate the technical inefficiency measures of individual producers in a particular year by employing the conditional distribution of u_{it}, given ε_{it}, which were derived by Jondrow et al., (1982):

$$E\left[u \middle| (\varepsilon - u)\right] = \sigma_0 \left\{ \mu_0 / \sigma_0 + \left\{ \phi(\mu_0 / \sigma_0) / \Phi(\mu_0 / \sigma_0) \right\} \right\} \tag{14}$$

where $\mu_0 / \sigma_0 = -\{\varepsilon_{it} \cdot \sigma_u \cdot h(\mathbf{x}_{it})\} / \sigma_{it}$ and $\sigma_0^2 = \{\sigma_u^2 h^2(\mathbf{x}_{it})\} / \sigma_{it}^{\ 2}$.

4. Estimation and Empirical Results

4.1 Data and Estimation

The model is estimated using balanced panel data of 443 large agricultural enterprises from three Russian regions. 70 farms are located in Oroel (central Russia), 180 farms in Krasnodar (southern Russia) and 193 in Samara (Volga Russia). The data set was provided by Goskomstat - the Russian State Committee of Statistics. The data set covers the period from 1996 to 2001. To enable a more accurate assessment of the dependence of production on weather conditions, attention is focused on crop production. All enterprises included in the sample are large-scale farms with a crop area of more than 200 ha intensively growing grain for commercial use[4]. On average, the structure of sowing area in the selected farms is as follows: 58.6 percent grain and legumes, 8.8 percent sunflower seed, 2.4 percent sugar beet, 0.3 percent potato and field-grown vegetables; the remainder refers to other crops[5]. The sample represents between 22 and 45 percent of the total crop area in the individual regions. In the view of experts, Krasnodar and Samara are regions with a higher exposure to natural hazards. Samara and Oroel belong to a small group of Russian regions that recently have been very active in introducing Western production technologies (Schüle and Zimmermann, 2002).

Production output is measured as annual farm revenues from crop production plus the value of unsold grain (*Y*)[6]. The mean output function is a function of the area of sown land (*Land*), labor

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[4] Selecting farms was done by applying the following criteria: sown area is to be more than 200 ha, crop revenue is to be more than 50 percent of the farm revenue and the ratio of the grain sowing area to whole farm sowing area is to be higher than 0.4. Additionally, farms with very strong specialization on potato or vegetable production, as well as seed breeding farms, were excluded from the sample.

[5] The statistical data used in the study does not provide a stronger differentiation of the sowing areas with regard to individual crops.

[6] The value of unsold grain was reckoned as a difference between a farm's annual grain production and grain sales multiplied by grain prices in the year 2001.

defined as the annual average number of employees involved in crop production (*Labor*), the value of depreciation, machinery maintenance and fuel costs in crop production[7] as a proxy for capital (*Capital*), materials costs (*Materials*)[8] and time (t and t^2) as an indicator of technical change[9]. Technical inefficiency (TI) is a function of the same variables. To enable a more precise assessment of various inputs effects on production risk, some components of materials costs, such as seed costs *(Seed)*, fertilizer costs (*Fertilizer*) and other production costs (*Suppl*)[10], were considered individually in the production risk function. The variables *Land*, *Labor* and *Capital* specify this function as well[11].

All monetary data were measured in 1,000 rubles and adjusted to the year 2001 by the regional price indices for agricultural inputs and output as they are provided by Goskomstat (Goskomstat, 2002a and b, 2003a and b). However, for fertilizer and capital, these indices were not obtainable. Two options exist for adjusting these data: using the input-specific price indices defined on the country level, or using the regional price index for aggregated agricultural input (which is defined for a wide range of inputs). The first option did not reflect regional price development, while the second did not account for changes in price relations. Since the study aims to explore the effects of different input groups on production output and its variability, the former option was expected to cause a smaller bias than the latter. Therefore, in the case of fertilizer costs, their values were adjusted by employing the fertilizer price index at the country level. To deflate depreciation, the country-level price index for machinery in crop production was applied. Maintenance and fuel costs were adjusted by the regional index for aggregated agricultural input.

Additionally, no distinction between seed produced on the farm and purchased seed was possible. However, since most of the farms use self-produced seed, the regional agricultural output price index was employed.[12]

The functional forms which are used for the mean production function are the Cobb-Douglas and translog specifications. Production variance and technical inefficiency (TI) are assumed to follow the Cobb-Douglas form. The LR-test suggested a rejection of the Cobb-Douglas form in favour of translog. However, since the translog specification did not provide theoretically consistent parameter estimates, constraints were introduced to fulfill monotonicity and the necessary conditions of quasi-concavity at the approximation point[13].

Both models (8) and (9) were estimated for the 443 farms from all 3 regions and for the farms from individual regions, i.e., 180 farms from Krasnodar, 70 from Oroel and 193 from Samara. The obtained parameter estimates are presented in Table 1.

The parameters α_j, β_j and γ_j are elasticities of the factor j in the mean, output variance and TI function, respectively. Positive values of the coefficients β_j in the production risk function mean that the corresponding factor increases production variability, whereas negative values signal that the factor is a risk-decreasing one. Negative signs of the coefficients γ_j indicate that a factor reduces technical inefficiency, otherwise a factor is TI increasing[14].

4.2 Estimation Results

Table 1 presents the coefficient estimates for all 4 samples. Most parameter estimates are significant, with the exception of some cross-product variables, and the coefficients of the risk and

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[7] Since we did not have data on machinery maintenance and fuel costs separately, we decided to integrate both parts of these costs into *Capital* variable. They have both defined the intensity of machinery application in Russian farms during the transformation period.

[8] Whereas other costs are calculated as the difference of total production costs of crop production and costs of labor, seed, fertilizer, equipment and machine maintenance, and fuel.

[9] Because only cost data was available for most of the production factors, no breakdown between factor quality and quantity was possible.

[10] Usually costs of plant protection. Therefore, they could be considered as a proxy for pesticides and herbicides use.

[11] To reduce computing time, all variables were normalized by their geometrical means.

[12] Certainly for the farms purchasing high quality seed, this leads to some distortions.

Furthermore, it is assumed that all farms in the regions have the same price risk, although in our opinion this may be rather restrictive.

[13] For details, see Morey (1986).

[14] γ_j =0 means that a factor is neutral with regard to technical inefficiency.

inefficiency function. Additionally, LR-tests indicate the *Kumbhakar with SPR* model appears to be more appropriate than the *JP with SPR* model in all samples at an acceptable level of significance[15]. Thus, the following discussion of estimation results concentrates on this approach.

The results of model estimations for the individual regions show that the parameter estimates are quite different in the three regions. Therefore, a James/Welch-test (Sievers, 1989) was conducted to prove the significance of differences between the production function parameter estimates in the considered samples (H0: $\alpha_{jKrasnodar} = \alpha_{jOroel} = \alpha_{jSamara}$). The test showed that the three considered regions have significantly[16] different parameter estimates with respect to land elasticity. Consequently, estimates of a common production function for all three regions appear to be incorrect. For that reason, the results of the separate estimations for the individual regions are considered more accurate and are discussed in the following.

The estimates show that the elasticities of scale are larger than those in Oroel and Samara, 1.22 and 1.17, respectively; for Krasnodar, constant returns to scale are evident. High total elasticity of production in Oroel is explained primarily by higher land elasticity in this region compared to Krasnodar and Oroel, 0.46 against 0.30 and 0.34, respectively. As Oroel belongs to the Russian regions with the most productive soils, this finding complies with the empirical evidence. The greatest production elasticity is observable in all samples with respect to materials. Their scores vary between 0.47 and 0.53. Regarding capital, the highest proportional contribution to the production is apparent in Samara – 0.20 compared to 0.13 in Oroel and 0.09 in Krasnodar. The estimates of labor elasticity are comparable in all three regions, but relatively low compared to the other partial production elasticities, although its estimate for the farms in Krasnodar is not significant. This indicates that labor appears to be relatively less productive. This result is in line with the conclusions of Osborne and Trueblood (2002b) as well as Liefert and Swinnen (2002) that labor is an excessive production factor in Russian agriculture.

The results suggest a positive impact of technical change in the considered period for two of the three regions – Samara and Oroel (Figure 1). Only in Krasnodar is the impact of technical change not distinct - the parameter estimates for technical change are not significantly different from zero in this region.

With regard to the bias of technical change, there is evidence for land-saving and capital-using technological change in the data. Different studies (Osborne and Trueblood, 2002b, Bezlepkina and Landsink, 2002) have found that farms in Russia overuse land and have explained this by low land shadow prices. In this context, land saving can be seen as a positive development in Russian agriculture, as this might imply that the farms tend to put excessive land out of operation. On the other hand, as the land market in Russia is developing rather slowly, this land cannot be effectively transferred to other economic agents. So the area of unused crop land grows in Russia[17].

[15] Additionally, both models were compared with traditional specifications of the production function. However, the hypothesis that the coefficients of the risk function, as well as the function explaining technical efficiency are not significantly different from 0, was rejected with 1 % significance level for all samples.

[16] The null hypothesis was rejected at the 1% level of significance with respect to land.

[17] According to the Russian Statistical Committee, the area of sown land amounted to 79.6 Mio. ha in 2003, which is approximately 68 percent of its 1991 level (Goskomstat, 2004).

Variable		All 3 regions JP with SPR	All 3 regions Kumbhakar with SPR	Krasnodar JP with SPR	Krasnodar Kumbhakar with SPR	Oroel JP with SPR	Oroel Kumbhakar with SPR	Samara JP with SPR	Samara Kumbhakar with SPR
Mean Production Function	αKrasnodar	0.91	1.25***	0.81***	1.08	---	---	---	---
	αOroel	0.96	1.11**	---	---	0.88	1.20*	---	---
	αSamara	0.71***	0.89***	---	---	---	---	0.93	1.06***
	αLand	0.27***	0.12***	0.36***	0.30***	0.44***	0.46***	0.49***	0.34***
	αLabor	0.22***	0.16***	0.16**	0.14	0.11***	0.11***	0.14***	0.10***
	αCapital	0.18***	0.22***	0.12***	0.09***	0.19***	0.13*	0.14***	0.20***
	αMaterials	0.48***	0.60***	0.50***	0.47***	0.42***	0.52***	0.40***	0.53***
	αt	-0.02	-0.03	0.10**	0.05	0.00	-0.11*	-0.15***	-0.13***
	αtt	0.01*	0.01***	-0.01**	0.00	0.00	0.01	0.03***	0.03***
	αLand&t	-0.01	-0.00***	-0.02	-0.03	-0.08**	-0.08***	-0.05**	-0.04*
	αLabor&t	-0.02**	-0.02***	0.01	0.03**	0.02**	0.02***	-0.03***	-0.02***
	αCapital&t	0.02***	0.02***	0.02**	0.02**	0.04**	0.05***	0.03***	0.02***
	αMaterials&t	0.01	0.01***	0.00	0.00	0.04***	0.02	0.04***	0.03***
	αLand^2	0.09***	0.10***	0.10	0.11*	0.20***	0.18***	-0.11*	-0.02
	αLabor^2	0.11***	0.11***	0.08*	0.11***	0.11***	0.11***	0.10***	0.08***
	αCapital^2	0.05***	0.05***	0.04***	0.04***	0.04**	0.05***	0.05***	0.05***
	αMaterials^2	0.08***	0.10***	0.07**	0.09***	0.17***	0.16***	0.12***	0.15***
	αLand&Labor	-0.09***	-0.04***	0.12	0.05	0.07	0.04	-0.12*	-0.08
	αLand&Capital	-0.02	-0.03***	-0.04	-0.04	-0.06	-0.07	0.13***	0.10**
	αLand&Materials	-0.04	-0.08***	-0.04	-0.10**	-0.30***	-0.27***	-0.03	-0.13**
	αLabor&Capital	-0.04***	-0.04***	-0.07**	-0.08***	-0.02	-0.03	-0.03	-0.04*
	αLabor&Materials	-0.05**	-0.09***	-0.20***	-0.14***	-0.11**	-0.09*	0.05*	0.03
	αCap&Materials	-0.09***	-0.08***	-0.01	0.00	-0.09***	-0.09***	-0.15***	-0.12***
Production Risk Function	βLand	0.43***	0.49***	0.25***	0.13	0.11	0.00	0.80***	1.03
	βLabor	0.14***	0.24***	0.15***	0.03	0.42***	0.39***	-0.07*	-0.13***
	βCapital	0.15***	0.11***	0.18***	0.22***	0.25***	0.34***	0.04	-0.04***
	βSeed	0.17***	0.13***	0.23***	0.07	0.01	-0.02	0.21***	0.15***
	βFertiliz	0.03***	0.04***	0.07***	0.46***	0.04***	0.04***	0.03***	0.03***
	βSuppl[c]	0.02***	0.00	0.00	0.00	0.04***	0.05***	0.02**	-0.01
	βKrasnodar96	0.75***	0.67***	0.51***	0.37***				
	βKrasnodar97	0.73***	0.62***	0.45***	0.32***				
	βKrasnodar98	0.65***	0.59***	0.43***	0.32***				
	βKrasnodar99	0.72***	0.65***	0.47***	0.40***				
	βKrasnodar00	0.61***	0.54***	0.39***	0.30***				
	βKrasnodar01	0.57***	0.50***	0.35***	0.25***				
	βOroel96	0.42***	0.38***			0.43***	0.33***		
	βOroel97	0.43***	0.41***			0.44***	0.40***		
	βOroel98	0.47***	0.44***			0.48***	0.47***		
	βOroel99	0.39***	0.33***			0.36***	0.36***		
	βOroel00	0.46***	0.43***			0.32***	0.30***		
	βOroel01	0.46***	0.43***			0.36***	0.33***		
	βSamara96	0.41***	0.38***					0.55***	0.50***
	βSamara97	0.44***	0.42***					0.62***	0.60***
	βSamara98	0.31***	0.27***					0.35***	0.29***
	βSamara99	0.33***	0.30***					0.42***	0.36***
	βSamara00	0.30***	0.27***					0.44***	0.40***
	βSamara01	0.31***	0.26***					0.42***	0.36***
TI Function	γLand	---	-0.40***	---	-0.01	---	0.00	---	-0.35***
	γLabor	---	-0.38***	---	0.31	---	0.21	---	0.13
	γCapital	---	0.41***	---	0.04	---	-0.12	---	0.36***
	γMaterials	---	1.24***	---	0.38***	---	1.44***	---	1.15***
	γt	---	0.04***	---	-0.07*	---	-0.45*	---	0.11***
	$\sigma_u{}'^{\,b}$	---	0.19***	---	0.42***	---	0.32**	---	0.15***
Value of test statistics (2logLR)		76.31***		81.75***		11.16*		49.73***	

a) *, ** and *** denote significantly different from 0 at the 10%, 5%, and 1% level;

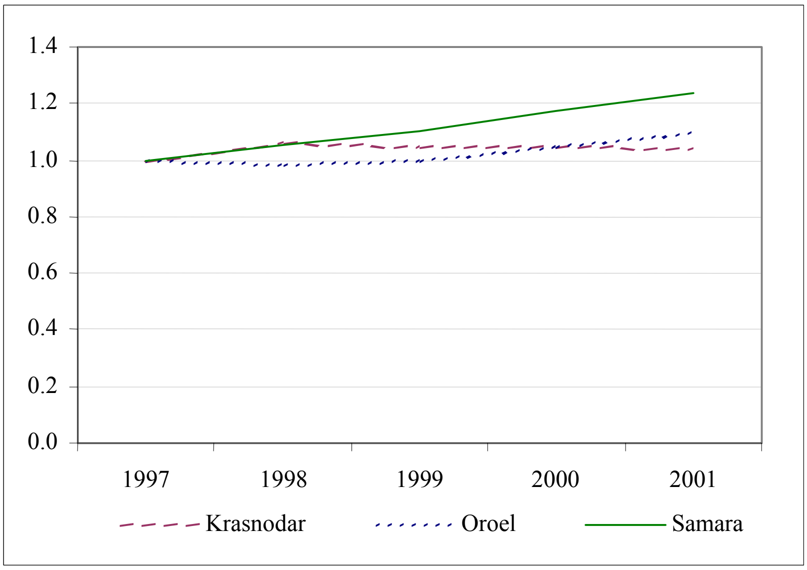
b) $\sigma_u' = \gamma 0 * \sigma_u$;

c) Suppl - other costs.

Source: authors' calculations.

Table 1. Parameter estimates.

In Samara, the capital-using impact of technical change is accompanied by labor-saving change. This indicates that the farms in this region are likely to release labor that becomes redundant. A different process seems to occur on the farms from the Krasnodar and Oroel regions, where the capital using impact of technical change does not cause labor saving. Under market conditions, the opposite would have been expected. Hayami and Ruttan (1985) argue that technical change in agriculture is governed by relative factor prices; those technologies, which allow the greatest reduction of the scarcest (most costly) input, are adopted. However, labor may not necessarily be considered as a scarce factor in Russian agriculture. During Soviet times, the authorities were concerned primarily with full employment (Macours and Swinnen, 2002). At that time agriculture served as a buffer for labor which was excessive in other sectors of the economy. During the transition process, most farms have pursued a protective employment policy, retaining surplus labor in agriculture[18]. This argument suggests that labor input is not governed by relative prices; hidden unemployment remains a widespread phenomenon in rural areas. Krasnodar and Oroel belong to the most important agrarian regions in Russia; there, agriculture contributed 0.32 and 0.36 percent, respectively, to the regional GDP in 2001 (Goskomstat, 2003b). On the other hand, Samara is considered as an industrial region (industry contributes approximately 90 percent to the regional GDP there). Respectively, in 2001 the agricultural sector employed 23.3, 21.4 and 7.6 percent (Goskomstat, 2002c) of the total labor force in these regions. Therefore, the results of the study seem to be in line with empirical evidence, that the relatively limited employment opportunities outside agriculture in Krasnodar and Oroel arguably cause difficulties in releasing the redundant labor in farms; this should be easier in Samara.



Source: authors' calculations.

Figure 1. Annual technological change in the selected farms.

In the following, two hypotheses are discussed:

(1) Agricultural production in Russia is subject to considerable production risk; besides systemic weather effects, output variance depends on the intensity of input use.

(2) Technical inefficiency enhances the output variability in Russian agriculture; TI is farm specific and can be explained by input use.

(1) As mentioned before, to prove the hypothesis on production risk, the traditional specification of the stochastic production frontier was tested against the alternatives: *JP with SPR* and *Kumbhakar with SPR* models by means of the likelihood-ratio test. For all samples the null hypothesis (H0: $\exp(\mathbf{D}\beta_t)g(\mathbf{x})= 0$) was rejected at the 1 % level of significance. In this regard the data are consistent with the hypothesis that production risk presents an important source of variability of agricultural production in all selected regions.

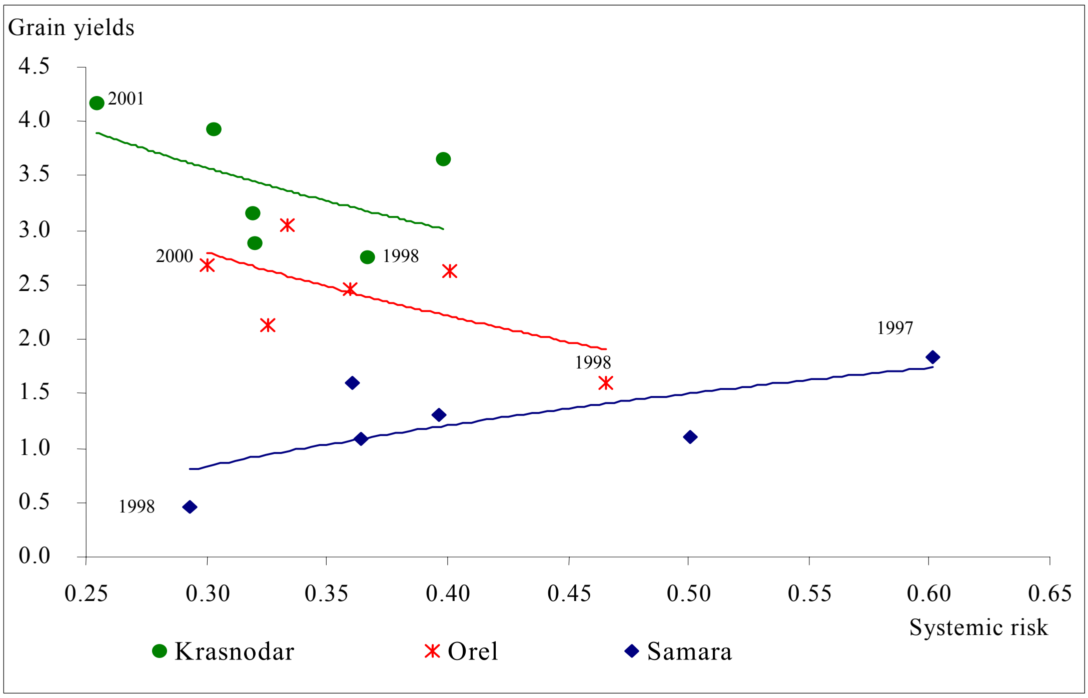
Principally, parameter estimates of the risk function are significantly different from zero. According to the estimates, production variability is quite high and many of the input coefficients in

[18] Statistical data shows that the number of employees in Russian agriculture sank from 9.7 in 1991 to 7.2 in 2001, i.e., by 26 percent. However, as sown area dropped by 32 percent in the same period, an increase in labor intensity (man-to-land ratio) was present in 2003 compared to the pre-reform period (Goskomstat, 2004).

the risk function have comparatively large values, first of all in the case of labor and capital in Oroel, and capital and fertilizer in Krasnodar.

The parameters of the systemic component of production risk are highly significant for all three regions. In addition, the estimated values of this component of production risk are relatively large, which implies that a considerable portion of the output variation can be attributed to systemic risk in the selected regions.

Two different patterns can be distinguished among the considered regions. First, systemic risk, and accordingly, yield variability increases in those years when weather conditions are unfavourable for crop production. Second, the systemic risk component is low in adverse weather years, and conversely, enhances output variability in good weather years. In Krasnodar and Oroel, a negative effect of the assessed values of systemic risk has been found. Figure 2 demonstrates grain yields development in the selected farms with respect to systemic risk[19]. The farms in these regions had their highest yields in the years 2000 and 2001. Accordingly, for these years, systemic risk estimates show that it reached its lowest values – 0.30 and 0.25 in Krasnodar, and 0.30 and 0.33 in Oroel, respectively (Table 1). In Krasnodar and Oroel, weather conditions tend to be relatively favourable for crop production; consequently, expected grain yields are rather high. So, adverse weather conditions in individual years appear to be caught by higher systemic risk values. For Samara, the opposite holds true. This region belongs to the so-called “risky crop growing” regions in Russia (Sheltikov et al., 2001), where the prevailing climatic and natural production conditions are rather harsh. This is reflected in quite low expected yields in these regions. High output variability is caused there by favourable, rather than adverse weather conditions. Therefore, in Samara, another pattern appears to be present – high parameter estimates of the systemic risk component point to favourable weather conditions in the individual years and correlate positively with the grain yields, as can be seen from Figure 10. Accordingly, the lowest value of systemic risk is assessed for 1998, when a severe drought caused widespread crop failure in Volga Russia. The highest value of systemic risk is assessed for 1997, which was a good crop year.



* - trends are assessed by applying the logarithmic functional form.
Source: authors’ calculations.

Figure 2. Systemic risk (estimated) and grain yields in the selected farms*.

19 To evaluate the estimates of the systemic risk component, only grain yields were taken into consideration in this paper. However, it is necessary to keep in mind that there are also other crops which contribute to the farm output.

One important advantage of the JP-approach is the possibility of distinguishing between an input effect on mean output and its impact on output variability, i.e., risk. In this study, serious differences among the parameter estimates in the mean and risk function could be found in the case of labor and capital in Samara. The coefficients on these variables are negative and significant for the Samara region farms. This indicates that production risk is reduced with increased use of this input. These estimates can be attributed to the fact that farms in Samara have been very active in the last decade in introducing minimum-tillage technology[20], which allows to save soil moisture and hence reduce yield variability caused by drought. With regard to labor, to the best knowledge of the authors, its risk-reducing effect has not been discussed in the literature so far. In the framework of this study, labor might be regarded as a proxy of farm size. Should this be the case, large farms turn out to manage production risk more effectively than their small counterparts.

However, according to the model estimates for two other regions, most of the considered factors have a risk-increasing effect on agricultural output there. Additionally, the estimation results do not confirm the view that pesticides are a factor of stabilizing rather than increasing agricultural production (Quiggin and Chambers, 2003) – indeed, the parameter estimates are mostly insignificant with regard to this factor. This contradicts, however, to the estimates of the model with a Cobb-Douglass specification of the mean output function where a highly significant risk-reducing effect of plant protection was found[21].

In general, the analysis demonstrates that there is only a weak response of the farms to production risk: most production factors enhance farms' production volatility. Therefore, the model estimates can serve to draw an empirically-relevant conclusion: current factor endowment of Russian farms is not adjusted to production conditions and should be adjusted to them in the future.

(2) The likelihood-ratio tests leads to a rejection of hypothesis H0: no inefficiency, i.e., $q(\mathbf{x})$ = *const*., and $\sigma_u^2 = 0$, at the 1 % significance level with respect to the model estimations for farms in Krasnodar and Samara. This suggests that the specification of the model including technical inefficiency is more appropriate, i.e., technical inefficiency enhances the variability of agricultural production in these regions. The hypothesis H0 cannot be rejected for the farms for Oroel only at the 10 % significance level, however[22].

The variance of output defined in the model with TI ($\boldsymbol{\sigma}^2 = \{\exp(\mathbf{D}\beta_t)g(\mathbf{x})\}^2 + q(\mathbf{x})^2\sigma_u^2$) is explained mostly by variance due to production risk. For most farms in all regions $q(\mathbf{x})^2\sigma_u^2 < \{\exp(\mathbf{D}\beta_t)g(\mathbf{x})\}^2$, i.e., according to model estimates, production risk contributes considerably to the volatility of agricultural production (Figures 3 to 6).

The model estimates show that only materials have an effect significantly different from zero on the technical efficiency of the farms in all considered regions. Even though, according to the parameter estimates of the mean production function, this factor is *ceteris paribus* most productive in all three regions, variable inputs seem to be used inefficiently in most farms – the parameters estimates have positive values, i.e., the farms are inefficient with regard to this factor. This finding is in line with the results of Osborne and Truebood (2002a) as well as Bezlepkina and Oude Lansink (2003) who argue that Russian farms use too much variable input.

Unlike the other two regions, in Samara, most of the parameter estimates in the TI function are significant. According to the model estimates, technical efficiency decreases with increased use of capital in this region. This result apparently indicates the overuse of capital and hence is in line with findings of Osborne and Truebood (2002a) who argue that technical and allocative inefficiency in Russian agriculture can be explained by machinery–intensive farming practices inherited from the Soviet era. Additionally, the use of old machinery can induce efficiency losses because of higher maintenance costs. On the other hand, unjustified high capital use can result from investments in new

[20] According to information from the regional government, in 2004 minimum tillage was introduced on 560,000 ha, i.e., 27 percent of the total sown area, in the Samara region (Samara government official site, 2005).

[21] The estimation results of the model with the Cobb-Douglas specification of the mean production function are available from the authors upon request.

[22] Lower significance of the estimates is possibly caused by a lower estimation efficiency due to relatively small number of observations from Oroel – 420 observations against 1,158 and 1,080 from Samara and Krasnodar, respectively.

machinery as well. In this regard, two options are available for Russian farms at the moment: investing in either highly productive Western, or relatively old-fashioned domestic equipment. This implies that investing farms introduce capital-intensive and labor-saving technologies. At the same time, due to institutional restrictions, there is hardly a labor release on farms[23]. This induces inefficiencies with respect to both capital and labor use in this region. Land has a positive effect on technical efficiency in Samara - large farms appear to be more efficient in this region.

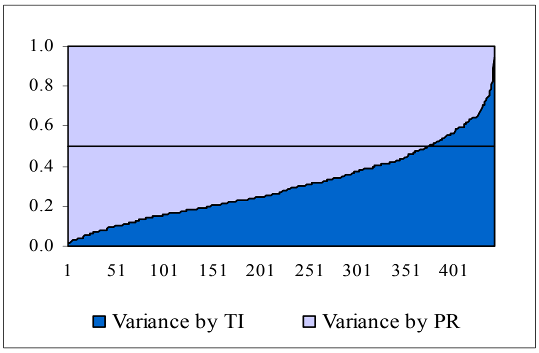


Figure 3. Sample 1 (all 443 farms)

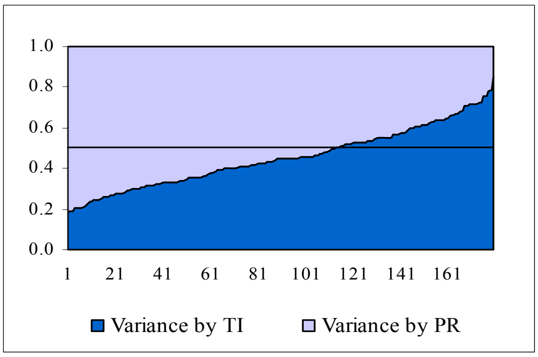


Figure 4. Sample 2 (183 farms from Krasnodar)

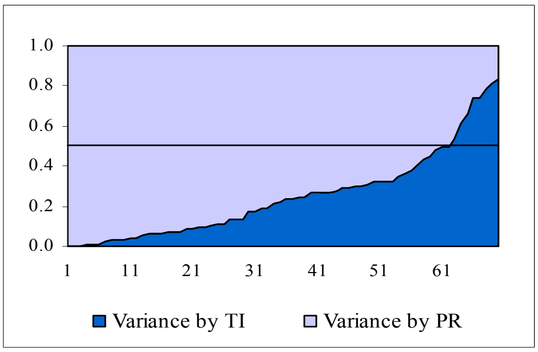


Figure 5. Sample 3 (70 farms from Oroel)

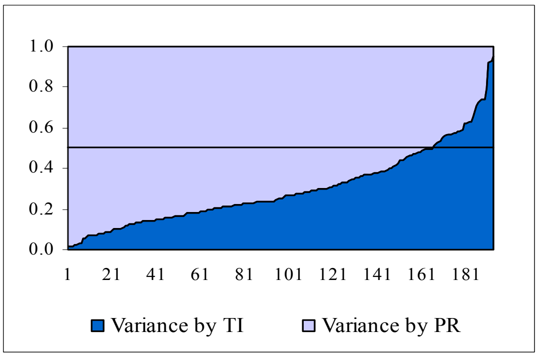


Figure 6. Sample 4 (193 farms from Samara)

Source: authors' calculations.

Figures 3 – 6. Ratio of the variance induced by TI and by production risk (PR) in the selected farms (average for the farms over 1996-2001).

The development of farms' technical efficiency over the considered period is presented in Figures 7 to10. The estimates suggest that technical efficiency increased over the considered period in both regions, Oreol and Krasnodar, while in Samara no improvement of farms' technical efficiency took place from 1996 to 2001.

In order to understand the development of technical efficiency in the transition process, both the institutional and technological aspects have to be taken into account. The institutional aspects find their expression in the "U-curve effect" (Havrylyshyn et al., 1998) suggesting that adapting to market coordination causes high transaction costs, as agents have to learn how to act in the new environment. Apparently, this initially induces an efficiency decline. Learning and forming the new incentives lead to a reduction of transaction costs during transition, as well as improved performance. In this context, increasing efficiency can be regarded as an indicator of the adjustment of economic agents to the requirements of the new environment.

On the other hand, due to the diffusion of innovations, a technological effect can induce an outward shift of the production frontier. However, for the enterprises which fail to adopt innovative

[23] The data used in this study shows that the annual average number of workers in crop production declined in the Samara region farms, from 81 to 70 per farm, on average (i.e., by 14 percent) in the 1996 – 2001 period. However, given the relatively low real wages in Russian agriculture, this was apparently not sufficient to prevent inefficiencies with respect to capital use.

techniques, the distance to the best domestic practice increases, implying an increase in technical inefficiency due to the technological effect.

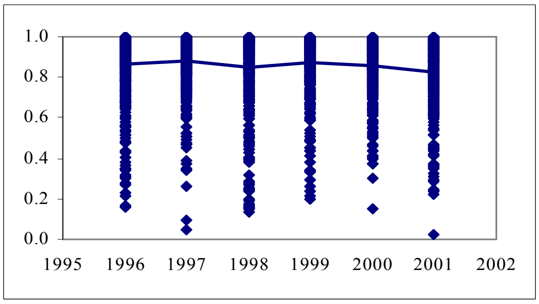


Figure 7: all 443 farms

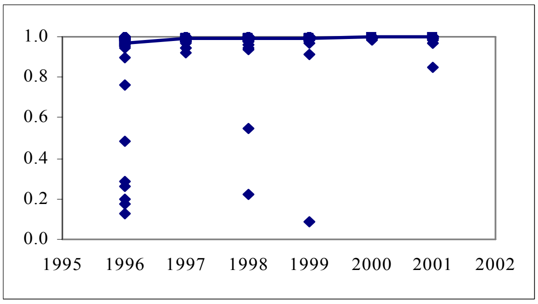


Figure 8: 183 farms from Krasnodar

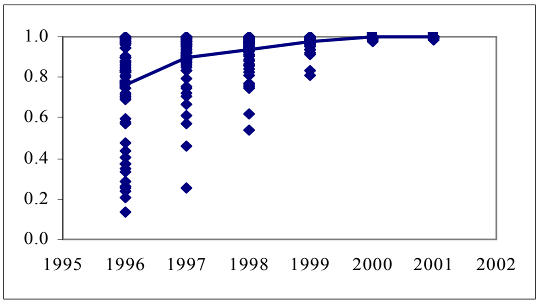


Figure 9: 70 farms from Oroel

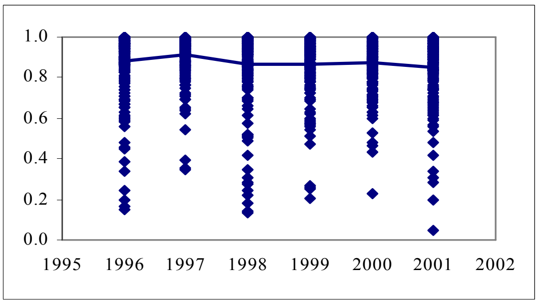


Figure 10: 193 farms from Samara

Source: authors' calculations.

Figure 7 – 10. Technical inefficiency of selected farms (1996-2001) (1.0 = 100 percent efficiency).

In accordance with the "U-curve effect", the results of technical efficiency estimates for the farms in Krasnodar and Oroel are in line with the findings of the earlier studies on technical efficiency of Russian farms, which reported a decline of technical efficiency in the 1991-1995 period (Sotnikov, 1998; Sedik et al., 1999) as well as a slowdown in this decline in the subsequent 1995-1998 period (Osborne and Trueblood, 2002b). Arguably, the results of this study present evidence suggesting the beginning of the second, i.e., upward, part of the "U-curve" in these regions.

At the same time, in Krasnodar and Oreol, only a marginal impact of technological change on the production frontier was identified. This implies that the adoption of new technologies plays only a minor role in the adjustment process in these regions. It can be expected that in this situation the "U-curve effect" dominates, with farms becoming more efficient, though under a constant technology.

In contrast, the technological effect has arguably been prevailing in the case of Samara, where farms have been very active in introducing new technologies that have enabled them to shift the production frontier outwards[24]. The "U-curve effect" might have been also present in Samara, but the development of inefficiency scores suggests that the technological effect has been dominant in this region. Thus, the study results demonstrate that the regional differences in the development of technical efficiency revealed by Voigt (2002) seem to be ongoing.

[24] One consequence of this interpretation is that the rank of the farms in Samara with regard to technical efficiency scores in the individual years is subject to significant changes. This is exactly what can be found in the results: many farms which were very efficient at the beginning of the considered period achieved only a medium level of efficiency in subsequent years. Many farms which were rather inefficient in 1996 moved into the efficient group. Moreover, once they have reached this group, their position becomes relatively stable.

5. Conclusions

This study has focused on the estimation of the magnitudes of technical inefficiency and production risk faced by agricultural producers in Russia. The study used the Just and Pope model (1978) to estimate a production function considering production risk, and its extension, by incorporating technical inefficiency as specified by Kumbhakar (2002) in the framework of cross-sectional data. The models were extended by introducing a term to account for a systemic part of production risk and by applying it to panel data.

By analyzing panel data of 443 farms from different parts of Russia, results were obtained which suggest that technical inefficiency enhances the variability of agricultural production in Russia. Moreover, according to the model estimates, production risk considerably contributes to the volatility of Russian agricultural production. For most farms, output variability is explained mainly by production risk. This indicates that when investigating agricultural production development in Russia, more attention should be paid to the presence of production risk and related farmer behavior. In particular, concerning the studies on technical efficiency of Russian farms, neglecting the effect of risk on production output and farmers' response to risk may cause incorrect estimations of technical efficiency.

The estimates indicate that there are significant differences in production technologies in the three investigated regions. This holds not only for the production elasticities but also for the impact of technological change. While in Oreol and Krasnodar, a shift of the production frontier has hardly been observed, Samara has experienced more dynamic development which has apparently significantly enhanced its production possibilities.

According to the study results, in Oreol and Krasnodar, farm efficiency increased significantly between 1996 and 2003. However, the farms in Samara have been less successful in adjusting to the best regional practice in the considered period. In this regard, the study results show that the development of technical inefficiency can be related to the effect of technological change. Under constant technology it is reasonable to assume that firms learn from past experience and, thus, are on a path towards the best production practice. On the other hand, the diffusion of innovation shifts the production possibility set outward. Thus, different farms can define the production frontier in subsequent years. Should the technological effect prevail, for the enterprises which fail to adopt innovative techniques, the distance to the best domestic practice increases, implying a decline in technical efficiency. This seems to be the case for the farms in Samara.

Additionally, the analysis demonstrates that there is only a weak response of the farms to production risk: most production factors enhance farms' production volatility. This implies that current factor endowment of Russian farms is not adjusted to production conditions. At the same time, as production risk plays an important role in the development of agricultural production at this stage, farms have to search for options to improve their responses to production risks, primarily with respect to the introduction of modern production technologies and practices that can reduce output volatility and enable a more flexible factor use subject to the state of nature. Finally, further research is needed to analyse the farmers' response to production risk. If agricultural producers do not exhibit risk-adjusting behaviour, the reasons for this have to be analysed. In this regard, it would be necessary to model farmers' risk preferences and estimate its impact on input use explicitly.

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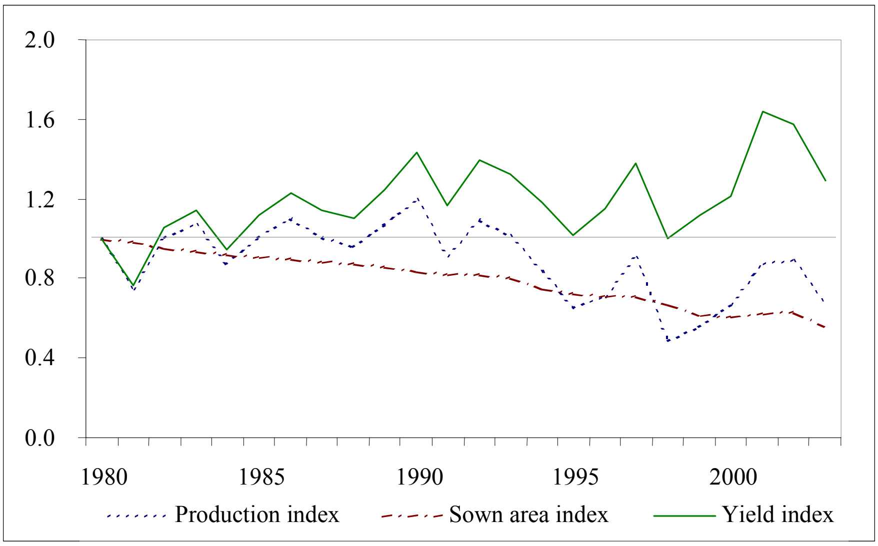
Appendix A. Grain production, sown area and yield development in Russia, 1980-2003 (1980 = 1.0).



Source: own calculations based on official statistics
(Goskomstat, 1980-1986, Goskomstat, 2002b)

Appendix B: Yield and yield variability of main Russian crops in 1980-1991 and 1992-2003 (per ha of sown area).

Product	1980 - 1991		1992 - 2003	
	Average yield, 0.1t/ha	Coefficient of variation, %	Average yield, 0.1t/ha	Coefficient of variation, %
Cereals, total	14.4	15.0	16.5	15.8
Sugar Beet	190.3	21.9	183.1	11.5
Sunflower Seed	11.1	16.8	8.8	12.7

Source : authors' calculations based on official statistics
(Goskomstat, 1980-1986, Goskomstat, 2004)

Appendix C: Descriptive statistics of the variables.

Variables and units of measurement	Output 1000 RUB of 2001	Land Hectars of sown area	Labor Workers in crop production	Capital 1000 RUB of 2001	Materials 1000 RUB of 2001	Seed 1000 RUB of 2001	Fertilizer 1000 RUB of 2001	Other costs 1000 RUB of 2001
All farms								
Min	105	206	2	24	58	3	0	0
Max	156479	28153	769	29167	64460	14774	16606	55798
Mean	17470	5228	135	3086	9500	1718	1250	6532
Std.Dev.	19517	3459	114	3444	9017	1459	1889	6850
Farms from Krasnodar								
Min	274	206	8	44	124	32	0	0
Max	156479	20255	769	29167	64460	11719	16606	55798
Mean	30066	6289	217	4086	15158	2318	2488	10352
Std.Dev.	23186	3170	122	3992	10173	1490	2337	8050
Mean 1996	27633	6638	227	3119	14698	2042	2192	10464
Mean 2001	35801	6192	203	6367	16151	2402	2897	10852
Farms from Oroel								
Min	105	250	2	31	106	3	0	0
Max	75159	14031	490	20856	36347	7596	7700	29204
Mean	10812	3464	84	2268	6244	1143	794	4307
Std.Dev.	12315	2394	79	2786	6342	1121	1073	4900
Mean 1996	10033	3615	90	1855	6587	1101	487	4999
Mean 2001	14137	3304	78	3575	5595	1084	1126	3386
Farms from Samara								
Min	108	250	2	24	58	32	0	0
Max	57538	28153	396	19310	41802	14774	4459	32003
Mean	7908	4878	76	2450	5404	1368	260	3776
Std.Dev.	7722	3706	54	2825	5013	1335	443	3951
Mean 1996	7938	5428	81	2179	7323	1643	186	5494
Mean 2001	7730	4310	70	3294	3848	1065	459	2325

Source: authors' calculations.