Modeling Soviet Agriculture for Assessing Commanding Economy Policies

By Padma Desai, Columbia University
and Balbir Sihag, University of Lowell

February 16, 1994
Discussion Paper No. 689

Abstract

A fixed effect production function model is adopted and estimated in this paper for evaluating the agricultural policies of the Soviet planners in Russia (formerly RSFSR) and Kazakhstan from 1953 to 1980. In particular, the functional form allows the variance of output to increase or decrease as one input is increased.

Our estimates suggest the following assessment of the Soviet planners' agricultural policies during the command economy days:

First, the extension of farming in Kazakhstan with generally inferior soils and climate resulted in lower crop levels (with identical input applications). Second, the massive inflow of capital resources and sharp rise in fertilizer use in agriculture during the Brezhnev years (1964-1982) did not contribute significantly to crop levels in the two republics. Finally, the extension of the cultivated area into marginal lands generally, and of fertilizer use in wrong mix and form contributed to the instability of agricultural output.

In conclusion, we emphasize the need for specifying and estimating appropriate models for evaluating Soviet agricultural policies especially because farm-level data are rapidly becoming available in the republics of the former Soviet Union.

February 16, 1994
Preliminary
Comments welcome

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Padma Desai
Columbia University

and

Balbir Sihag
University of Lowell

The increasing availability of firm and farm level data in Russia and other republics of the former Soviet Union is contributing to scholarly research in several directions. In particular, the use of the production function as a microeconomic tool of empirical analysis with the aim of drawing policy conclusions in specific activities is much in evidence.

Thus, Danilin, Materov, Rosefielde and Lovell (1984) estimate technical efficiency in Soviet cotton refining from a 1974 sample of 151 enterprises and conclude that its production efficiency is high in relation to the estimated stochastic frontier, and exhibits little inter-enterprise dispersion. Johnson et al. (1994) employ a similar stochastic frontier analysis to data for 11,400 farms for six years (from 1986 to 1991) and provide measures of technical efficiency and its inter-farm variation. Both studies draw on current estimation procedures and extend them to increasingly complex functional specifications. At the same time, they speculate on the impact of Soviet planning procedures and managerial incentive systems on the performance, as indicated by their estimates, of the relevant activity.

These pioneering studies herald the liberation of planned economy empirical research from its early efforts constrained by the paucity of data, and their availability by industrial branches and sectors of economy. The use of large, microeconomic data set for econometric estimation with continuing refinements in the functional form and estimation procedures makes the estimates credible. When these are employed for explaining the role of the planners in influencing the performance of this or that activity, the exercise comes closer to being authentic. As these economies move to free markets, the initial studies can also provide a meaningful anchor for contrasting the performance of the original production units (in the panel data set) with their fortunes at a future date.

While such data are becoming increasingly available for post-Soviet agriculture, the critical issue consists in linking the proposed analysis with the available, state-of-the-art scientific work. The important contributions of Anderson, Buccola, Griffiths, Just, McCarl and Pope provide the necessary foundations from two perspectives.

First, the production function must be defined such that it captures the peculiarity of agricultural activity. In the standard production function with the multiplicative error term, the marginal risk defined as the partial derivative of the variance of output with respect to a given input is positive for all inputs.[1] This implies that the reduction at the margin of, say, pesticide use will reduce crop variability whereas actually it might raise such variability. Therefore, the functional form must permit the variance of output to increase or decrease as one input is increased. Following Just and Pope (1978, 1979), and Griffiths and Anderson (1982), it can be stated as follows:

$$Q = f(X) + h(X)\epsilon = \prod_{i=1}^{k} X_i^{\alpha_i} + \epsilon \prod_{i=1}^{k} X_i^{\beta_i} \qquad (1)$$

where Q is the level of output, X_i is the level of the ith input and ϵ is a stochastic error term with zero mean and variance of σ^2.

1 For a rigorous proof, see Just and Pope (1979, p. 277).

This more general formulation permits greater analytical flexibility with respect to the effect of an input on mean farm output and its variance.[2]

Second, Griffiths and Anderson (1982) suggest that time series data may be needed to estimate β_i's in equation (1) since in a cross-section series, much of the variation in output "...can be attributed to weather variation and if the increased use of some inputs is to have mitigating effect on output variation, observations over time are likely to be needed to capture this effect." (p. 529). On the other hand, cross-section data are necessary, in their view, to estimate α_i's in equation (1) because input levels in a farm may not vary significantly over time to allow estimation of α_i's. Therefore, they recommend the pooling of cross-section and time-series data to estimate equation (1). However, appropriate statistical tests must be undertaken to justify such pooling.

Additionally, the model in (1), which is nonlinear and includes the stochastic function $h(X)\epsilon$ can be specified in various ways.

Therefore, in Section I, we state our model and discuss the statistical tests necessary for accepting the parametric estimates. The sequence of steps for estimating the model is stated in Section

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[2] However, the residual $h(X_t, \beta)\epsilon_t$ is generally heteroscedastic and therefore creates problems about hypothesis testing and the statistical properties of the estimates of α_i and β_i. Despite the presence of heteroscedasticity, the ordinary least squares (OLS) and nonlinear least squares (NLS) estimates of α_i are unbiased and consistent; however, they are asymptotically inefficient. (Buccola and McCarl, 1986, p. 732.)

II. In Section III, we present some estimates for the two republics of RSFSR (currently Russia) and Kazakhstan based on our (fixed effect) specification. In conclusion, we interprete our results and link them to the decisions implemented by Soviet planners in the command economy days. In particular, our estimates suggest that while increased fertilizer use raised mean agricultural output modestly, it increased its variance.[3] In conclusion, we emphasize that the policies of planners in Soviet agriculture cannot be assessed properly without incorporating the special features of farming in the model. More so as detailed, farm-level information becomes available.

I. The Model and the Required Statistical Tests

The fixed effect model (alternatively known as the covarinace model) adopted by us is stated as follows:

$$Q_{it} = (\lambda + \sum_{i=2}^{N} \theta_i Z_{it} + \sum_{t=2}^{T} \delta_t W_{it}) \prod_{k=1}^{K} X_{kit}^{\alpha_k} + \epsilon_{it} \prod_{k=1}^{K} X_{kit}^{\beta_k} \quad (2)$$

where

Z_{it} = 1 for the i^{th} cross-sectional unit

= 0 otherwise (i=2,3...N)

W_{it} = 1 for the t^{th} time period

= 0 otherwise (t=2,3...T)

θ_i measures the unit-specific effect

and δ_t measures the time-specific effect

$E(\epsilon_{it}) = 0$, $Var(\epsilon_{it}) = \sigma^2$ and $Cov(\epsilon_{it}, \epsilon_{jt}) = 0$ for i=j

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[3] In a pathbreaking study of Indian foodgrain production, Mehra (1981) showed that the improved seed/fertilizer-based technologies (beginning from the mid-1960s) raised grain yields but increased their instability compared to the earlier years.

Note that in this fixed effect model, the dummy variables Z_{it} and W_{it} are included to allow for, say, differences in soil quality in the two republics and of weather over time.[4]

What are the implications of this procedure? Suppose we adopt a dummy of 1 for soil quality in the RSFSR and 0 in Kazakhstan.[5] The soil quality of the RSFSR agricultural belt is a mixture of the chernozem (black), forest-steppe and steppe soils.[6] On average, the soils in Kazakhstan are inferior with a combination of steppe and semi-desert soils. The adoption of the dummies here implies that

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4 It must be emphasized that soil quality is mentioned here as an illustrative feature distinguishing one republic from the other. Another distinguishing element can be the extent of irrigation facilities. The relative breakdown of farms in collective (kolkhoz) and state (sovkhoz) types with differing incentives to farm workers is yet another aspect.
Again, these features can be measured in the unit-specific effect for a set of farms when they are pooled together according to a distinguishing characteristic such as a soil zone.

5 Instead of the 0-1 dummies, one can devise soil indexes for the republics (or for a group of farms in a given soil zone) based on a detailed mapping of soil types in their agricultural belts. Such an index will reflect soil differences in a republic (or the selected farms) in relation to average soil. It is also possible to construct an irrigation index reflecting the availability of irrigation water in a republic (or in a set of farms). Similarly, time series of weather indexes can be constructed for a republic (or farms in a climate zone) on the basis of weather data used by Desai (1986).

6 For a discussion of the variety of soils in the Soviet agricultural belt, see Desai (1986, Chapter 3).

the soil differences in the two republics and their impact on output are nonrandom.[7]

In any case, it is necessary to test if there are systematic differences between the two republics with respect to, say, soil quality and other features, or weather, or both. For this purpose, it is necessary to apply appropriate tests.

Statistical Tests:

Suppose we want to establish in our fixed effect model that there are no systematic differences of soil quality and weather (from year to year) between the two republics. For this purpose, we must test the null hypothesis that $\theta_i = 0$, $\delta_t = 0$. Additionally, we must test the hypothesis that the coefficients α_k are identical for each republic in order to justify the pooling of the time-series and cross-section data. The detailed steps of applying these tests for the fixed effect model are stated in the Appendix.

II. Estimation Methodology

The procedure for estimating the fixed effect model is essentially similar to those for estimating the random effect model

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[7] By contrast, in the random effect model (discussed extensively by Just and Pope, and Anderson and Griffiths) the likely magnitude of the impact of republic-specific feature such as soil quality is included in the error term and is random.
Among the statistical differences of the two models, note that the fixed effect model adopted by us sacrifices many degrees of freedom whereas the random effect specification saves degrees of freedom and the estimators are, therefore, likely to be efficient. On the other hand, its major limitation is that if the omitted variables, such as soil quality and weather, are related to the explicitly specified inputs on the right hand side of the equation, the estimated coefficients can be biased and inconsistent implying a misspecification.

developed by Just and Pope (1979) and further refined by Griffiths and Anderson (1982). We state the sequence of steps for our fixed effect model because we have employed the first two steps for deriving the results reported in Tables 1 and 2.

Step 1. Run a nonlinear regression on the first part of equation (2) (i.e. without $\epsilon_{it}\prod_{k=1}^{K} X_{kit}^{\beta_k}$ and compute the residuals

$$\hat{u}_{it} = \{Q_{it} - (\hat{\lambda}+\Sigma\hat{\theta}_i Z_{it}+\Sigma\hat{\delta}_t W_{it})\prod_{k=1}^{K} X_{kit}^{\hat{\alpha}_k}\}$$

Step 2. Regress the residuals $\hat{u}_{it}$ on the inputs as follows:

$$\ln(|\hat{u}_{it}|) = \beta_o + \Sigma\beta_k \ln(X_{kit}) + \epsilon_{it}$$

Step 3. Compute $\hat{h}_{it} = \prod_{k=1}^{K} X_{kit}^{\hat{\beta}_k}$ and transform equation (2)

by dividing both sides by $\hat{h}_{it}$. Run again a nonlinear regression as in step 1 to obtain $\hat{\lambda}$, $\hat{\theta}_i$, $\hat{\delta}_t$ and $\hat{\alpha}_k$'s in equation (2).

Step 4. Iterate the above steps two or three times to obtain efficient estimators of α_k's. In this connection, the discussion in Buccola and McCarl (1986) regarding the small sample properties of an estimator and suitable correction of β_o in step 2 is relevant.

III. Data and Estimates

Production function data for the two republics from 1953 to 1980 are put together from official statistical handbooks. The output refers to ruble value of agricultural output (in billion 1970 rubles); the inputs include land (in million sown hectares), capital stock (in constant 1973 prices), labor (in million manhours) and fertilizers (in million metric tons). We found it difficult to update the series in view of the nonavailability of republic handbooks for recent years.

In Tables 1 and 2, we present the estimates of α_k's and β_k's for the two republics of RSFSR and Kazakhstan resulting from an estimation of equation (2) which is specified as Cobb-Douglas constant returns to scale with pooled cross-section and time-series data. Note that only steps 1 and 2 of the estimation procedure have been

Table 1. Estimates of α_k's

Row	Republic	λ	θ_2	α_1 (Land)	α_2 (Capital)	α_3 (Labor)	α_4 (Fertilizer)	SER	LLF	DW
(1)	Kazakhstan	0.3498 (1.4900)	--	0.6414 (3.6770)	0.2384 (1.5264)	0.1339 (0.7151)	-0.0137	0.5311	-19.8527	1.9159
(2)	RSFSR	1.0516 (1.6972)	--	0.6017 (3.2047)	0.0530 (0.6714)	0.2485 (1.1860)	0.0968	2.3834	-61.8906	1.4064
(3)	Pooled	2.5980 (1.9748)	0.0000	0.4548 (2.6982)	-0.1399 (2.2270)	0.4524 (2.4328)	0.2327	2.2388	-122.519	0.9193
(4)	Pooled	0.6639 (2.3304)	0.3727 (2.5833)	0.6045 (4.7347)	0.0554 (1.0071)	0.2447 (1.7175)	0.0954	1.6777	-105.817	1.4820

Table 2. Estimates of β_k's

Row	Republic	β_0	β_1 (Land)	β_2 (Capital)	β_3 (Labor)	β_4 (Fertilizer)	R^2	SER	DW
(1')	Kazakhstan	8.5924 (1.7684)	0.1152 (0.1016)	-3.7832 (2.7662)	0.9990 (0.7407)	2.1316 (2.5609)	0.1450	0.8748	2.7463
(2')	RSFSR	5.9134 (0.1874)	0.1405 (0.0178)	-0.4143 (0.2783)	-3.7382 (1.0483)	0.9157 (0.7520)	0.1452	1.0107	1.8762
(4')	Pooled	-1.7107 (0.4063)	1.6971 (1.6372)	-1.7034 (2.3199)	-1.0791 (0.9931)	1.3282 (2.3716)	0.4102	0.9274	1.8931

Notes to Tables 1 and 2

1. The estimates in rows (1) and (2) of Table 1 are derived by fitting equation (2'), and in rows (3) and (4) by fitting equation (2") with pooled data for the two republics of Kazakhstan and RSFSR. However, as stated in the text, θ_2 is constrained to be zero in row (3) and is estimated in row (4). The estimates of Table 2 are derived by fitting the equation in step 2 on page 7. The equations in Table 1 are nonlinear whereas in Table 2, they are specified in double-log formulation. Furthermore, the dependent variables in the equations of rows (1') and (2') are the (natural) log of the absolute values of the error terms estimated from the equations in rows (1) and (2), and the explanatory variables are also in (natural) log. Finally, the estimates in row (4') of Table 2 are derived by pooling the data of the two republics. Here the dependent variable is the (natural) log of the absolute values of the error terms of the equation in row (4).

2. Values in parentheses are t values of the estimated parameters. SER is the standard error of the regression, LLF is the log of likelihood function, DW is the Durbin-Watson statistic and R^2 is (correlation coefficient)^2.

carried out. In other words, we have not corrected for heteroscedasticity to achieve efficient estimates. Also, the statistical test with regard to the equality of α_k's (k=1,2,3,4) across the republics in justification of pooling the data for the two republics has not been performed. In Table 1, $\Sigma\alpha_k=1$ and in Table 2, β_k's, as already stated, are not constrained to be non-negative. Finally, the dummy variables W_{it} representing time-related effects such as year-to-year variations in weather are not included in the estimation. However, the dummy variable Z_{2t} representing republic-specific effect such as different soil quality is incorporated. This implies that the equation for estimating the parameters for Kazakhstan (specified as republic 1) and RSFSR (specified as republic 2) in rows (1) and (2) of Table 1 is:

$$Q_{it} = \lambda_i \prod_{k=1}^{K} X_{kit}^{\alpha_{ki}} + \epsilon_{it} \prod_{k=1}^{K} X_{kit}^{\beta_{ki}} \qquad (2')$$

where i=1,2 and k=1,2,3,4 for the four inputs; whereas the specification for estimating the parameters for the pooled data of the two republics in rows (3) and (4) is:

$$Q_{it} = (\lambda + \theta_2 Z_{2t}) \prod_{k=1}^{K} X_{kit}^{\alpha_k} + \epsilon_{it} \prod_{k=1}^{K} X_{kit}^{\beta_k} \qquad (2'')$$

where i=1,2; k=1,2,3,4; and Z_{2t} is the dummy of 1 for RSFSR and 0 for Kazakhstan. However, note that in row (3) of the

Table, $\theta_2=0$ implying that λ is identical for the two republics whereas in row (4), θ_2 is estimated which implies that if λ is the constant term for Kazakhstan, then the constant term for RSFSR is $(\lambda+\theta_2)$.

In row (2) of Table 1, the coefficients α_k have positive signs but in row (1), the coefficient of fertilizer is negative. The estimating equation for both is (2'). When the data for the two republics are pooled in row (3) and θ_2 is constrained to be zero implying that the values of α_k's and the constant term λ are equal for both the republics, the estimates of the coefficients α_k change significantly and the coefficient with respect to capital becomes negative. On the other hand, in row (4), θ_2 is allowed to vary. As a result, the standard error of the regression falls, all the coefficients have expected signs, and most of the coefficients and, in particular, θ_2, are statistically significant. In view of the statistically significant value of θ_2, we reject the null hypothesis that $\theta_2 = 0$.

What do the results of row (4) in Table 1 imply? First, the coefficient of $(\lambda+\theta_2)$ for RSFSR estimated at 1.0366 (0.6639 + 0.3727) exceeds the corresponding term in Kazakhstan with a value of 0.6639. This implies that, for identical input levels, crop levels in RSFSR would be higher than in Kazakhstan. This may result from say the superior soils in RSFSR. Or perhaps in the RSFSR, relatively more farms are organized along collective lines implying an edge in farm incentives. The dummy →

with an estimated parameter of 0.3727 incorporates the combined impact of several such features. Clearly, it is important to separate these by employing properly-specified indexes each incorporating these aspects for the republics. Second, the ⟶ estimates suggest very small output elasticities with respect to capital and fertilizers. If these estimates are correct, a further application of capital and fertilizers does not contribute significantly to augmenting crop levels. (Note that the impact of these inputs on crop variability is a separate issue and is analyzed in terms of the estimates of β_k presented in Table 2).

The estimates of β_k of Table 2 are derived on the basis of step 2 on page 7 . The estimated constant term here is β_0. The results in row (4') with the pooled data provide a better fit than in rows (1') and (2'). Most of the coefficients are statistically significant. Also, the signs of the coefficients are in line with our expectations: the variance in output increases with increases in the area under cultivation and fertilizers,[8] and the variance decreases with increases in capital and labor.

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8 The presumption here is that increased use of fertilizers without matching applications of new seed varieties, pesticides and water can raise crop variability. For evidence in support of this argument with respect to Soviet graingrowing, see Desai (1987, chapter 6).

Our estimates throw light on the agricultural policies pursued by the Soviet planners in RSFSR and Kazakhstan for almost three decades ending in 1980.

First, the extension of farming in Kazakhstan with generally inferior soils and climate resulted in lower crop levels (with identical input applications).

Second, the massive inflow of capital resources and sharp rise in fertilizer use in agriculture during the Brezhnev years (1964-1982) did not contribute significantly to increased crop levels in the two republics. Investments were channeled into agriculture, without consideration of costs or returns, over wide-ranging activities such as land drainage and reclamation, rural road-building and provision of social infrastructure, increasing the supplies of machines without regard to quality or spare parts, and setting up agroindustrial complexes. By 1982, outlays in agriculture had reached 27 percent of total investment in the economy. At the same time, fertilizer use per kilogram of hectare under grain had jumped sixfold from 8.9 kilograms in 1964 at the start of the Brezhnev leadership to 54 kilograms in 1982.

Finally, the extension of the cultivated area into the marginal lands generally, and fertilizer use in wrong mix and form contributed to instability of agricultural output.

Even at the republic level, our results can be improved by including suitable soil features and weather indexes. However, our preliminary estimates mark a significant step in specifying and estimating an appropriate model for evaluating Soviet agricultural

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Details are in Desai (1987 p. 243, and 1992) policies. These refinements are equally important when farm-level data are used for assessing agricultural policies in the former Soviet Union.

Appendix

METHOD OF HYPOTHESIS TESTING

Testing the Assumptions of the Fixed Effect Model

Testing the Null Hypothesis $\theta_i=0$

We estimate equation (2) with and without the constraint. For example, the null hypothesis $\theta_2=0$ means that the second republic (RSFSR) is not different from the first (Kazakhstan) and a rejection of the null hypothesis implies that RSFSR is diffrent from Kazakhstan.

We apply the likelihood ratio test to test the null hypothesis. That is

$$LR = -2[L(\tilde{\alpha},\tilde{\sigma}^2) - L(\hat{\alpha},\hat{\sigma}^2)] \sim \chi^2_1$$

where $L(\tilde{\alpha},\tilde{\sigma}^2)$ is the log of likelihood function with the constraint (i.e. $\theta_2=0$) and $L(\hat{\alpha},\hat{\sigma}^2)$ is the log of likelihood function without the constraint. Similarly, the likelihood ratio test can be used to test the null hypothesis $\delta_t=0$ and the equality (between the republics) of the coefficients α_k.

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