Economies of Scale in Agriculture: A Reexamination of the Evidence

Yoav Kislev and Willis Peterson

Prevented from conducting controlled experiments, the data economists work with are far from ideal. Measurement errors abound, important variables are unobserved, not all simultaneity can be accounted for, and we often use state, regional, or national aggregates to estimate structural parameters relating to individual firms or households. As a result, every measurement and estimated parameter can be questioned. Confidence is gained when quantitative results can be corroborated by observations on independent sets of data and, particularly, by analyses of different aspects of the same problem. With these observations in mind, we reexamine in this paper the evidence for economies of scale in U.S. agriculture.

Average farm size has grown markedly in the U.S. Between 1929 and 1987, output per farm increased more than six times, and land per farm expanded almost threefold (table 1). This long-term growth of farm size has led people to believe that large farms are more efficient than small ones. As a result there appears to be widespread concern that family farms are an endangered species, eventually to be replaced by large corporate enterprises. This implies a belief that there are economies of scale in U.S. agriculture. This belief is critically examined in the paper.

For most of us, "economies of scale in agriculture" mean (1) that the production function for the typical firm in the industry is characterized by increasing returns to scale, and (2) that small farms are less efficient than larger ones. We start with the first meaning of the scale economies and comment on issues associated with size distribution at the end of the section.

We consider four dimensions of the problem: sources of scale economies, measurement of the phenomenon, the implied consequences and ex-

Table 10.1 Per-farm output, inputs, and prices

		1929	1949	1959	1969	1974	1978	1982	1987
Gross output	($1000)	11.2	27.2	37.2	52.1	78.6	81.3	73.1	71.5
Value added	($1000)	8.2	20.4	24.7	31.2	47.6	45.9	39.7	36.7
Land	(acres)	157	216	303	389	440	449	440	461
Family labor	(labor-years)	1.2	1.2	1.3	1.1	1.0	1.0	1.0	0.9
Hired labor	(labor-years)	0.4	0.3	0.5	0.4	0.4	0.5	0.7	0.6
Part-time farming	(percent)	12	23	30	40	36	42	43	35
Machinery	($1000)	5.1	11.5	26.1	33.2	39.9	56.3	63.4	77.2
Land prices	($ per acre)	319	253	365	524	650	879	935	472
Land rent	($ per acre)	—	—	54	81	146	125	101	58
Diesel fuel	($ per gallon)	—	0.71	0.65	0.55	0.84	0.80	1.31	0.71
Custom rate	($ per acre)	16.62	13.91	13.56	11.52	18.63	15.76	16.12	12.43
Manufacturing wages	($ per week)	165	257	344	393	407	427	388	406
Wage rental	(index)	30	57	78	104	67	83	74	100

Note: All dollar values in constant 1987 prices, deflated by the Consumer Price Index. See appendix for description of data and sources.

planation of farm size growth. We argue that the scale economies hypothesis can be questioned on both conceptual and statistical grounds, and that it is not, in general, consistent with historical evidence. But we should perhaps make it clear at the outset that we do not think that size never matters. We do not believe, for example, that a dairy farm with one cow produces at the same average cost as a two hundred-cow operation. What we do propose is that agriculture is not typically characterized by increasing returns at the prevailing size distribution. The paper details this assertion.

Sources

It is a simple truism that if the world is exactly multiplied by two, production will also exactly double (Friedman 1962). But we are not interested in the returns to multiplication of every aspect of production. Rather, changes in scale are changes in the physical dimensions of production assets and in factor flows, while the environment, entrepreneurship, and often also management stay the same. Then scale effects are mixed. For example, it takes less than twice as much material to build a fence around a 640-acre farm as it does for a 320-acre operation. On the other hand, increasing management difficulties reduce efficiency as scale expands. Scale economies in an industry reflect the sum total of these effects.

Economies of scale in agriculture have been attributed to indivisibility of assets. Hayami and Ruttan assert that "the scale economies usually stem from the lumpiness or indivisibility of fixed capital" (p. 146). But it is hard to find long-run indivisibilities on the farm. Returns to scale is a long-run concept, and in the long run the size distributions of machinery, land,

structures, irrigation systems, herds, and flocks are continuous, not "lumpy." Tractors and their implements come in a variety of sizes, from the 8-horsepower hand-driven garden types to the 350-horsepower four-wheel drive behemoths. Other machines also come in a variety of sizes. In the few cases where large machines are the most efficient, such as combine harvesters and cotton pickers, rental markets develop.

The human agent may be the most "lumpy" input. But in this case the option of part-time farming on the one hand, and a large operation with hired labor on the other, makes it possible to scale the size of the farm and the labor of the family to the available managerial ability. We consider part-time farming below. At this point we come to the conclusion that among the conventional inputs of land, labor, and capital, long-run indivisibilities should not be given serious consideration as a source of economies of scale.

In spite of the apparent lack of indivisibilities, suppose for the sake of argument that the typical farm's long-run average total cost curve is U-shaped. As the industry adjusts toward the least cost point, scale economies should be gradually eliminated. At the minimum point, returns to scale are constant. In other words, returns to scale are a temporary phenomenon (Griliches 1963, p. 232). Taking the sum of the coefficients from a Cobb-Douglas production function as a measure of scale economies, the value should converge to one. Yet, as will be shown in the following section, this has not occurred.

Another possibility is that economies of scale are created by reduced uncertainty due to government intervention in agriculture (Madden and Partenheimer 1972, p. 103). However, government has been intervening since the 1930s. Adjustment should have been completed by now.

In a dynamic version of the U-shaped cost curve it is hypothesized that technological progress results in the production of new and larger machines that are more efficient than the older and smaller ones. Perhaps this is what Bieri, de Janvry, and Schmitz (1972) had in mind when they wrote that "the rate of increase in farm size in the future will be determined, to a large extent, by the rate at which the machine companies manufacture larger and larger equipment" (p. 802). But there is an identification problem here. Although the dynamic hypothesis is consistent with the observation that farms and machines have grown in size, and that measured economies of scale have not declined, it is not consistent with the growth and continued popularity of part-time farming. Nor is this hypothesis consistent with the observed cessation of the growth in farm size which occurred from the mid-1970s to the early 1980s without a corresponding halt to the progression of new technology, or the ability to produce still larger machines. Below we argue that the growth in optimum farm size induced changes in size of machines, rather than the other way around.

We turn now to size distribution. An "asset" that is clearly fixed on the

farm, particularly for the family unit, is entrepreneurial ability and management. This fixity raises important econometric issues which are discussed below. Here we note that even if management is fixed for the individual farm, it varies markedly among farms. With varying managerial ability, farms will differ in size in an industry in equilibrium, even if they all faced identical input prices and operated on a constant returns production function. Some of them may be part-time farms. Here again, as with machines, adjustment should have eliminated the measured scale effects, if farms were not of optimal size (ability adjusted) in the past.

Differences in managerial ability are, however, not the only generators of variation in farm size. Another factor is differences in the prices of inputs, labor in particular. People with low opportunity cost of labor—old farmers, spouses without permanent employment, regular workers who draw nonpecuniary returns from farming and rural life—may operate relatively small farms, many of them part-time operations, which do not reach optimal size when labor is priced at full cost. In this sense economies of scale might be said to exist in agriculture even when the industry is in equilibrium.

Can we attribute the phenomenon of part-time farming with the dimensions it reaches in modern agriculture to the low-cost hypothesis? Between 1949 and 1969, a period when farm size grew rapidly, the incidence of part-time farming, as measured by the share of operators working one hundred days or more off their farms, nearly doubled. In 1987 this figure was nearly three times larger than in 1929 (table 1). Hence the low-opportunity cost hypothesis implies that over the last several decades the share of people willingly accepting relatively low returns to labor grew significantly. Moreover, with technological advancement, the share of purchased inputs increased and terms of trade worsened, squeezing out inefficient farms—while part-timing increased. Considering this evidence, it is hard not to come to the conclusion that even if economies of scale exist in agriculture, they are of minor magnitude and significance.

Another type of observed economy of scale, this time transitory, arises when some farmers lag behind the others as the economic environment changes and optimal farm size is growing. Then operators who are late to recognize the changing circumstances produce on suboptimal farms, and at genuinely high costs. Latecomers may affect measured returns to scale in the industry, but being transitory this phenomenon is not an indication of economies of scale, a long-run static concept.

Measurement

Several methods have been employed to test for the existence of scale economies, including the survival technique, synthetic firm studies, and the sum of the coefficients of Cobb-Douglas production functions.

According to the survival technique (Stigler 1958; Saving 1961), growth of firms in an industry is an indication that large firms are more efficient than small ones; that is, economies of scale exist. However, the application of this method to agriculture is problematic. An implicit assumption of the method is that economic conditions have not changed. This is not true for American agriculture. Also, in this sector the labor input per farm has not grown (table 1). Thus the large increase in size cannot be the textbook case of increasing returns to scale.

If both increasing returns and rising cost of labor prevail, farms may grow in size while labor input is reduced (relatively or absolutely). But then an identification problem arises: which was the factor causing growth? We suggest below that changing relative cost of labor, not scale economies, induced farm growth. Moreover, by the logic of the survival technique, the existence of part-time farming, and particularly the fact that it increased during the 1950s and 1960s, suggests the absence of economies of scale in agriculture.

Synthetic firm studies are reviewed by Madden and Partenheimer. The difficulty with these studies is that problems of management and coordination and differences in skills and ability of farmers cannot be appropriately accounted for. These problems are considered the major sources of diseconomies of scale. Unless they can be taken account of, the measurement of scale economies by synthetic cost studies is meaningless.

Summing the coefficients of Cobb-Douglas production functions is probably the most common method of measuring scale economies in agriculture, although not the only method (see Kumbhakar, Biswas, and Bailey 1981; Hornbaker, Dixon, and Sonka 1989; Moschini 1990). To update the results of previous studies we fitted an aggregate agricultural production function to U.S. data for the census years 1978, 1982, and 1987. To ensure comparability with the earlier studies we utilized as much as possible the original Griliches (1964) specification and the Cobb-Douglas form. With five inputs (fifteen variables) and only forty-eight observations, the outcome of the more general translog form cannot be relied upon. The results are presented in table 2.

A summary of the sum of coefficients for the census years 1949 to 1987 obtained from previous studies and from table 2 is shown in table 3. The Davis study (1979) also followed the Griliches specification.

With the exception of 1964, the sum of coefficients averaged in the neighborhood of 1.3 from 1949 through 1974, with no discernible trend. The 1978-82-87 sums are almost identical to the earlier estimates. Thus there is no evidence to suggest that the sum of coefficients is converging to one. The remarkable thing about these results is that the sum of coefficients is relatively constant and larger than one during the entire 1949–87 period. This includes the time when farm size grew rapidly—1949–1974,

Table 10.2 Production functions

INPUTS	1978	1982	1987
Land and buildings	0.103 (2.63)	0.110 (2.25)	0.128 (3.73)
Labor	0.267 (5.17)	0.270 (4.52)	0.219 (3.70)
Machinery	0.231 (3.75)	0.272 (3.29)	0.150 (2.79)
Fertilizer and chemicals	0.274 (9.24)	0.206 (7.00)	0.272 (9.06)
Other	0.427 (14.5)	0.427 (10.9)	0.521 (13.7)
R^2	0.977	0.962	0.971
Sum of coefficients	1.30	1.29	1.29

Note: Figures in parentheses are t-ratios. See appendix for description of variables.

Table 10.3 Summary of findings

Source	Year	Sum of coefficients	Farm size (acres)	Wage-rental ratio (1987 = 100)
Griliches (1964)	1949-54-59	1.28	216-242-303	56-70-78
Davis (1979)	1964	1.18	352	89
	1969	1.28	389	104
	1974	1.28	440	67
Table 2	1978	1.29	449	83
	1982	1.29	440	74
	1987	1.30	461	100

when farm size was relatively stable, 1974–1982, and when farm size grew modestly, 1982–1987. If the sum of the coefficients is an indicator of scale economies, one would expect it to be larger when farm size grew rapidly than when it did not. Whatever the sum of coefficients is picking up, it does not appear to be measuring economies of scale.

The findings of scale economies reported in tables 2 and 3 as well as similar results from previous studies can be questioned on the grounds that the inability to measure management causes the sum of coefficients to be biased upward. Unlike the labor input that can in principle be measured correctly, management is unobservable. In general, one may expect management to be complementary to most other inputs—an operator with greater ability can manage a larger operation. Thus management is correlated with size, and economies of scale will tend to be overestimated. Other unobserved fixed factors have a similar effect on the estimates. The following analysis of this issue is due to Mundlak (1968); it was later expanded by Mundlak and Hoch (1965), and by Hoch (1962). It is recapitulated here in some detail because of its importance.

Let the firm (farm) production function be

$$Y = F(X, M), \qquad (1)$$

where Y = output,

X = a vector of observable inputs, and

M = firm-specific factors, management, or other fixed factors.

In general M is unobserved and not measured. The estimated production function is then,

$$Y = f(X). \qquad (2)$$

With the specification in equation (2), firms are on different functions—each according to its M value.

Better environmental conditions, more productive soil, superior location, as well as better management—are all reflected in higher (often unobserved) firm-specific factors. Consequently, complementarity prevails between the firm-specific factors and other inputs. As a result, total inputs (observed plus unobserved) will be underestimated on large farms relative to small farms, biasing regression coefficients, and giving the appearance that large farms are more efficient than small ones.

Given panel data and assuming that the farm-specific effects are constant over the period of observation, the bias can be eliminated if, instead of ordinary least squares (OLS) estimates of the pooled function, a covariance analysis is employed; that is, if firm dummy variables are included in the analysis. In covariance analysis the regression is estimated not from the original observations, but from deviations from the firm means. If, as assumed, M is constant, the deviations are all zeros and all firms are on the same function. The estimated coefficients are then unbiased "within" firm estimates.

Hoch (1962) summarized six previous studies that had reported both OLS and analysis of covariance of farm level production functions. In all six cases the sum of the coefficients in the covariance analysis was less than 85 percent of the sum calculated from the OLS estimates, and in all cases the sum was smaller than one. In his own analysis of dairy farms, reported in the same paper, Hoch found that covariance analysis reduced the sum of coefficients to levels smaller than one for ten samples of farms producing milk for market. He found, however, increasing returns to scale (and higher sums of coefficients) for two samples of dairies producing for manufacturing. To our knowledge this is the only reported evidence of covariance analysis increasing the sum of the coefficients, and the only case in which increasing returns were found with covariance analysis.

When farm-level data are not used, agricultural production functions are usually estimated for the average farm in a state, a region, or a country. Random aggregation within regions cancels out farm-specific effects and eliminates the associated specification biases. However, regions are also

characterized by regional-specific effects, such as soil, climate, or economic circumstances, and in samples of individual farms coming from different regions, the regional effects have to be accounted for. Again, one way to do it is to include regional dummy variables in the regressions. Failure to include such dummies, failure to account for the region-specific effects, results in specification biases. These specification biases are augmented by regional aggregation. Hence the bias in the estimated sums of the coefficients in Cobb-Douglas functions is likely to be more serious when the observations are regional or country averages than when they are farm-level data (Kislev 1966).

Since soil, climate, and other growing conditions differ among the states, one would expect the sum of coefficients to be biased upward. For an additional test of this hypothesis we conducted a covariance analysis for the intercountry productivity estimates of Hayami and Ruttan (1985). They found increasing returns to scale for the developed countries in their sample and constant returns for the developing countries. We recalculated their regression for the developed countries (Q21 in their table 6-2) but with country dummies. The sum of the coefficients was 1.320 without the dummies and 1.077 with the dummies. The latter sum was not significantly different from 1 (the standard error of the sum was 0.119). In another study, Capalbo (1988) reports decreasing returns to scale ($\Sigma b_i = 0.8$) from an aggregate agricultural production function estimated by OLS from time series data. However, the data employed in this study were aggregate output and inputs rather than per-farm observations. Consequently, nothing can be said about scale economies from these results. Scale economies relate to the size of the firm, not to the size of the industry.

Before leaving this section, a caveat should be added about covariance analysis. In general, random errors in the measurement of the observed variables cause underestimation of the regression coefficients. Griliches (1985) shows that covariance analysis may exacerbate the effect of measurement errors. By giving up "between" variability we run into the danger of increasing the "noise to signal" ratio. For our discussion of economies of scale, this possibility is particularly worrying—empirical observations are never error free. The message here is, again, that one cannot rely on a single source of evidence. Continuing with the survey of evidence, we turn now to the examination of the consequences of scale economies.

Consequences

Returns to scale, if they actually exist in agriculture, affect significantly both the economics of agriculture and our understanding of the farm sector. As long as the sources of the phenomenon are not clearly defined, the term *economies of scale*, just like *technical change*, is, in a sense, a name for our

ignorance. To see how ignorant, consider the following. Between 1929 and 1987 gross output per farm in the U.S. increased 6.4-fold (table 1). If inputs were all accounted for and perfectly measured, and constant returns to scale prevailed, inputs also would have increased 6.4 times and output per unit of input would have remained constant. If, in fact, the agricultural production function is homogeneous of degree 1.3, then the input growth necessary to achieve a 6.4-fold growth in output would have been 4.2 instead of 6.4 ($4.2^{1.3} = 6.4$). In other words, over one-third of the increase in per-farm output would be left unexplained [$(6.4 - 4.2)/6.4 = 0.35$]. As a degree of homogeneity, 1.3 is a very large number.

Economies of scale also affect the distribution of the returns to the factors of production. If, again, the production function is homogeneous of degree 1.3, and inputs are paid the value of their marginal products (VMPs), the sum total of payment to all factors will be $1.30 for each dollar of output (Euler's theorem). If, however, land and management constrain the size of the farm, and all other factors are purchased freely on the market (including own labor which can also be regarded as "purchased" in a free market), then prices will equal VMPs for the purchased inputs— land and management becoming the residual claimants. By the estimates of table 2, in which management is not included, the production elasticity of land (its factor share) is 0.12. It follows that the purchased inputs receive $0.88 \times \$1.30 = \$1.14$ for each dollar of output. This leaves land with a negative return. Of course, such an outcome is implausible. Land rents and land prices have not been zero (table 1).

Dynamic consideration can, however, be used to explain positive land values in the face of negative residual returns. If scale economies are viewed as a temporary disequilibrium, then as farms grow to a size sufficiently large to capture all scale economies so that constant returns prevail, the residual return to land will at this point become positive. Given a sufficiently low rate of interest, the discounted present value of these positive future returns also could be positive. However, this argument does not explain the positive land rents which prevailed throughout the period. Nor does it explain the apparent inability of land buyers to revise their expectations after decades of allegedly negative residual returns to land. A realization that increasing returns were not temporary should have led to a decrease in real land prices during the 1970s, but just the opposite occurred. Changes in inflationary expectations, along with changes in real rates of interest, appear to have had a larger impact on real land prices during the 1970s and 1980s than changes in expectations regarding the attainment of a constant returns-to-scale equilibrium. A likely reason is that land buyers and sellers did not perceive a disequilibrium. Nor is there any evidence in the land market literature of the existence of such a disequilibrium.

Farm Growth

The major consequence attributed to economies of scale is growth of farm size. And indeed American farms grew significantly up to the mid 1970s (table 1). However, economies of scale are not the only force driving farm growth. In an earlier paper we presented an alternative explanation, namely, that capital intensity and farm size are determined by the price ratio between labor and machinery (Kislev and Peterson 1982). For farming, the price of labor is the alternative earning opportunities outside agriculture (table 1). The cost of machinery is harder to measure. Theoretically it is the rental price. In agriculture, a rental market for machinery services exists in the form of custom work for hire. We, therefore, measured the cost of machinery services as the real value of custom rates in combine harvesting of wheat.

From 1929 to 1974 real custom rates declined. The increase in quality of machines and the decrease in the real cost of energy likely were important factors contributing to this decline. At the same time, the opportunity cost of farm labor increased causing the wage to rental ratio to increase even more (table 1). These relative price changes provided farmers with an incentive to increase the amount of machinery on farms, enabling the family farm to cultivate a larger amount of land and in order to maintain parity income growth with the nonfarm sector. Similarly, operators of livestock enterprises were able to enlarge their herds and flocks due to a higher capital intensity. In our paper virtually all of the growth in farm size over the period 1930–70 is explained by the model without reference to "economies of scale" or "technological change." The increase in real nonfarm earnings provided an incentive for many farmers to leave agriculture for alternative occupations. The "pull" effect of higher nonfarm wages appears to have dominated the "push" effect of lower machine costs (Peterson and Kislev 1982). In turn, off-farm migration increased the land available to each remaining farmer. Seen from this perspective, as urban wages rise, farmers are induced to either move to town, or in the case of those that remain full-time farmers, to increase the amount of resources at their command to keep income in parity with alternative occupations.

During the 1970s history extended our experiment in a new direction and provided new evidence when the wage to rental ratio sharply declined (table 1). As is shown in table 3, there is a strong correlation between changes in the wage-rental ratio and changes in farm size. From 1949 to 1969 the wage-rental ratio steadily increased, along with farm size. Coinciding with the energy crisis in the mid 1970s, the wage-rental ratio exhibited a sharp decrease. At the same time, growth in farm size abruptly

stopped. Average farm size remained relatively constant over the next ten years. During the mid 1980s the wage-rental ratio resumed its upward trend. The latest figures show a resumption of the growth in farm size during this time.

In general, the wage-rental ratio provides a better explanation for changes in farm size, and a better predictor of these changes, than the sum of coefficients, which exhibited no decline toward unity when farm size growth ceased.

Because farms are operated by a constant amount of quality adjusted labor, much of the increased machinery input has taken the form of larger machines, although new machines such as tomato and fruit harvesters and cotton pickers, also have been developed. Technologically, large machines could have been built a long time ago, and indeed, they were; take earth moving equipment, for example. In fact, the first four-wheel-drive tractors were assembled by farmers in the 1950s, so widespread and simple was the technology. Only afterwards did the machinery companies produce these large tractors. Here, as in other cases, the companies reacted to economic changes once they were reflected in demand. The decision of manufacturers to build larger machines is not an exogenous factor determining farm size.

One might argue that the U.S. income tax law, which taxed the earnings from capital at lower rates than earnings from labor through investment credits and accelerated depreciation, promoted the growth of large capital-intensive farms. This may be true, although Batte and Sonka (1985) report that increasing marginal income tax rates reduced the advantages of large-scale farms. To the extent that the income tax law reduced the price of capital relative to that of labor, it would, according to our farm size model, have contributed to the increase in farm size. But unless the tax law continued to lower the price of capital, it would have resulted in a once and for all change in equilibrium farm size, not a continued increase.

Farm mechanization can be viewed as a process of induced innovation. Changing factor prices for agriculture, including the price of energy, induced the machine companies to produce more efficient and larger machines. These changes also induced agricultural research to produce appropriate biological and chemical technologies. Changes in factor prices induced the suppliers of factors and technology to innovate. An alternative view, that innovations occurred on the farm, which in turn increased the demand for machinery even at higher real prices, cannot be supported by any evidence that we know of (Kislev and Peterson 1981).

Concluding Remarks

We have conducted a multidimensional examination of the hypothesis that agricultural production is characterized by increasing returns to scale and

find that it is not supported by the evidence. We focused on general trends: determinants of average farm size and its growth. Our conclusion does not deny the possibility that economies of scale exist in certain types of farms or periods, but that such economies do not explain overall growth in unit size in agriculture. Of course, at a point in time farm size will exhibit substantial variation around the mean because of differences in managerial ability among farmers (Sumner and Leiby 1987). The better the manager, the larger the output that minimizes average total cost.

We found again that the evidence supports the hypothesis that farm size is endogenously determined by the wage rental ratio as it affects machinery on farms and farm size. Since labor per farm has been relatively constant, deepening capital intensity has taken the form of larger machinery. As always, some farmers are earlier adopters of larger machines; others follow with a lag. Statistical estimates and synthetic analysis may then indicate returns to scale as the latecomers close the gap. But short-run disequilibria should not be confused with scale economies, which is a static concept to be viewed in a long-run context after adjustment has run its course.

Discarding economies of scale as an explanation of farm growth eliminates a lacuna in our understanding of the economics of the farm sector. At the same time, the discussion highlights an old-new puzzle: what are the economic, technological, and institutional factors that support the family unit as the dominant form of organization and maintain it through periods of large economic and technical change? The solution to this puzzle will do much to improve our understanding of the structure of agriculture.

Appendix

Data Sources and Definitions for Table 1

Unless otherwise noted, all references are published by the Government Printing Office in Washington.

Gross output: cash receipts from farming plus value of home consumption from *Agricultural Statistics* (respective years). In all cases, number of farms are from *Agricultural Statistics.*

Value added: gross output less feed, seed, livestock, fertilizer, and miscellaneous expenses from *Agricultural Statistics* (respective years).

Land: acres per farm from the Agricultural Census (respective years).

Family labor: number of family laborers per farm as reported in *Agricultural Statistics* (1957, 1967, 1984) adjusted for quality (income by years of schooling completed for rural males, ages 25–64, as reported by the 1980 *Census of Population*, was multiplied by the proportion of rural farm males in each schooling category and summed). The resulting weighted

averages for the census years 1940–80 were then utilized to construct a labor quality index, 1980 = 100. The labor quality indexes for the other four census years are as follows: 1940, 80; 1950, 82; 1960, 86; 1970, 93. The 1940 quality index is used to adjust the 1929 labor figure, while the 1970 index was used to adjust the 1969 and 1974 figures. The 1980 index of 100 was applied to the 1978 and 1987 figures.

Hired labor: data source and quality adjustment procedure are the same as for family labor.

Part-time farming: proportion of farm operators working 100 days or more off the farm as reported by the *Census of Agriculture* (1929, 1949, 1959, 1974, 1978, 1987).

Machinery: stock of machines on farms obtained by summing the value of shipments over the preceding fifteen years, each year deflated by the CPI. Data sources: *Statistical Abstract* (1928) and *Agricultural Statistics* (respective years).

Land price: value of land per acre, excluding buildings, as reported by U.S. Department of Agriculture, Economic Research Service, "Farm Real Estate Historical Series Data," *Statistical Bulletins 520* (1973) and *738* (1985), and *Agricultural Statistics* (1988).

Land rent: cash rent per acre of cropland, 1969–82, in the following eight midwestern states: Michigan, Wisconsin, Minnesota, Ohio, Indiana, Illinois, Iowa, and Missouri. Data source: USDA, ERS and ESCS, *Farm Real Estate Market Developments* (1972, 1977, 1981, 1985, 1989). Rental data for these states are not available prior to 1967. The 1959 figure was estimated by the following procedure: Burt (1986) reports a rental figure of $17 per acre for Illinois in 1959. His 1969 figure for Illinois is $30 per acre. The corresponding figure of 1969 for the eight midwestern states is $25 per acre, or 83 percent of the Illinois rent. We therefore estimated the 1959 rent for the eight states as 0.83 × $17 = $14 in current year prices, or $54 in constant 1987 prices. Figures for 1987 are from USDA, ERS, Agricultural Land Values and Markets: *Outlook and Situation Report* (June 1988).

Diesel fuel: data source: *Agricultural Statistics* (1953, 1963, 1973, 1977, 1981, 1988). Figure for 1949 is price of distillate.

Custom rate: charge per acre of custom harvesting of wheat in Kansas. The 1929 figure is from Reynoldson et. al. (1928). The 1949 rate is from Friesen, et al. (1953). The remaining custom rates are from an annual publication, *Kansas Custom Rates* (1970, 1975, 1980, 1989), published by the Kansas Crop and Livestock Reporting Service.

Manufacturing wages: gross weekly earnings in manufacturing as reported in the *Economic Report of the President* (1969, 1990).

Manufacturing wages divided by custom rate, expressed as an index, 1987 = 100.

Description of Production Function Variables

All variables are measured on a per-farm basis.

Output: sales plus value of home consumption plus inventory change plus government program payments. The sales figures are from the *Census of Agriculture* (respective years). Figures other than sales are from USDA, ERS, *Economic Indicators of the Farm Sector* (1988). All values are in constant 1977 prices. Each state's output was deflated by a weighted average USDA price index for livestock and crops with weights equal to the proportion of livestock and crop output in each state.

Land and buildings: value from *Census of Agriculture* (respective years) deflated state by state by the land and buildings price index, with each state equal to 100 in 1977.

Labor: operator plus hired labor from the *Census of Agriculture.* Operator labor was adjusted for proportion of farmers working two hundred or more days off the farm. Hired labor was converted to days by dividing by the farm wage rate from *Agricultural Statistics.* In the 1980 *Census of Population,* differences among states in the educational level of farm people are negligible. Hence, adjusting labor for quality differences had no effect on the coefficients.

Machinery: annual depreciation from "Farm Income Data: A Historical Perspective," USDA, ERS, *Statistical Bulletin* 740 (May 1986), deflated by custom rate price index for combine harvesting of wheat in Kansas.

Fertilizer and chemicals: value from *Census of Agriculture* (respective years), deflated by the USDA fertilizer price index.

Other: all other production expenses from *Census of Agriculture* (respective years), deflated by the USDA farm and motor supplies price index. All price indexes, except custom rates, are from *Agricultural Statistics.*

References

Batte, Marvin, T., and Steven T. Sonka. 1985. "Before and After Tax Size Economies: An Example for Cash Grain Production in Illinois." *Amer. J. Agr. Econ.* 67:600–609.

Bieri, Jurg, Alain de Janvry, and Andrew Schmitz. 1972. "Agricultural Technology and the Distribution of Welfare Gains." *Amer. J. Agr. Econ.* 54:801–8.

Burt, Oscar. 1986. "Capitalization Formula for Farmland Prices." *Amer. J. Agr. Econ.* 68 (Feb.): 26.

Capalbo, Susan M. 1988. "A Comparison of Econometric Models of U.S. Agricultural Productivity and Aggregate Technology." *Agricultural Productivity, Measurement and Explanation.* Ed. Capalbo and John M. Antle. Washington, D.C.: Resources for the Future.

Davis, Jeffrey. 1979. "Stability of the Research Production Coefficient for U.S. Agriculture." Ph.D. diss., University of Minnesota.

Friedman, Milton. 1962. *Price Theory*. Chicago: Aldine.

Friesen, H. J., et al. 1953. "1952–53 Custom Rates for Farm Operators in Central Kansas." Kansas Agricultural Economic Report no. 59.

Griliches, Zvi. 1963. "The Source of Measured Productivity Growth: United States Agriculture, 1940–60." *J. Pol. Econ.* 71:331–46.

———. 1964. "Research Expenditures, Education, and the Aggregate Agricultural Production Function." *Amer. Econ. Rev.* 54:961–74.

———. 1985. "Data Problems in Econometrics." *Handbook of Econometrics*. Ed. Griliches and M. Intriligator. Amsterdam: North Holland.

Hayami, Yujiro, and Vernon W. Ruttan. 1985. *Agricultural Development: An International Perspective*. 2d ed. Baltimore: Johns Hopkins University Press.

Hoch, Irving. 1962. "Estimation of Production Function Parameters Combining Time Series and Cross-Section Data." *Econometrica* 30:34–53.

Hornbaker, R. H., B. L. Dixon, and S. T. Sonka. 1989. "Estimating Production Activity Costs for Multioutput Firms with a Random Coefficient Regression Model." *Amer. J. Agr. Econ.* 71:167–77.

Kislev, Yoav. 1966. "Overestimates of Returns to Scale in Agriculture—A Case of Synchronized Aggregation." *J. Farm Econ.* 48:967–83.

Kislev, Yoav, and Willis Peterson. 1981. "Induced Innovations and Farm Mechanization." *Amer. J. Agr. Econ.* 63:562–65.

———. 1982. "Prices, Technology and Farm Size." *J. Pol. Econ.* 90:578–95.

Kumbhakar, S. C., B. Biswas, and D. Bailey. 1981. "A Study of Economic Efficiency of Utah Farmers: A System Approach." *Rev. Econ. and Statist.* 71:595–604.

Madden, Patrick J., and Earl J. Partenheimer. 1972. "Evidence of Economies and Diseconomies of Farm Size." *Size, Structure, and Future of Farms*. Ed. Gordon A. Ball and Earl O. Heady. Ames: Iowa State University Press.

Moschini, G. 1990. "Nonparametric and Semiparametric Estimation: An Analysis of Multiproduct Returns to Scale." *Amer. J. Agr. Econ.* 72:589–96.

Mundlak, Yair. 1968. "Empirical Production Function Free of Management Bias." *J. Farm Econ.* 43:44–56.

Mundlak, Yair, and Irving Hoch. 1965. "Consequences of Alternative Specifications in Estimation of Cobb-Douglas Production Functions." *Econometrica* 33:814–28.

Peterson, Willis, and Yoav Kislev. 1986. "The Cotton Harvester in Retrospect: Labor Displacement or Replacement?" *J. Econ. Hist.* 46:199–216.

Reynoldson, L. A., et al. 1928. "The Combined Harvester—Thresher in the Great Plains." USDA Technical Bulletin no. 70 (Feb.).

Ruttan, Vernon W. 1988. "Scale, Size, Technology and Structure: A Personal Perspective." *Determinants of Farm Size and Structure*. Ed. Lindon J. Robinson. Department of Agricultural Economics, Michigan State University.

Saving, Thomas R. 1961. "Estimation of Optimum Size of Plant by the Survivor Technique." *Quar. J. Econ.* 75:569–607.

Stigler, George, J. 1958. "The Economies of Scale." *J. Law and Econ.* 1:54–71.

Sumner, D. A., and J. D. Leiby. 1987. "An Econometric Analysis of the Effects of Human Capital, Size and Growth Among Dairy Farms." *Amer. J. Agr. Econ.* 69:465–70.