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Does Farm Size Really Converge? The Role of Unobserved Farm Efficiency

by

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Abstract

We analyze the growth of family farms in Israeli cooperative villages between 1981 and 1995, using longitudinal data. We use instrumental variables to account for the endogeneity of initial farm size, and correct for selectivity due to farm survival. We also include a technical efficiency index, derived from the estimation of a stochastic frontier production model, as an explanatory variable. We find that technical efficiency is an important determinant of farm growth, and that not controlling for technical efficiency could seriously bias the results. The size distribution of Israeli family farms is found to be mostly diverging, while without technical efficiency farm growth seemed to be predominantly random.

JEL classifications: Q12, L25, C34.

Keywords: Farm Size; Farm Growth; Farm Survival; Instrumental Variables; Sample Selection; Technical Efficiency.

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Abstract

We analyze the growth of family farms in Israeli cooperative villages between 1981 and 1995, using longitudinal data. We use instrumental variables to account for the endogeneity of initial farm size, and correct for selectivity due to farm survival. We also include a technical efficiency index, derived from the estimation of a stochastic frontier production model, as an explanatory variable. We find that technical efficiency is an important determinant of farm growth, and that not controlling for technical efficiency could seriously bias the results. The size distribution of Israeli family farms is found to be mostly diverging, while without technical efficiency farm growth seemed to be predominantly random.

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Introduction and background

Farm sectors in developed economies are continuously undergoing structural changes. One of the key structural features that are changing is the size distribution of farms. The interest in the size distribution of farms has increased in recent years due to the increased recognition of the multifunctional role of family farms in shaping rural landscapes, rural economies and rural societies. An increase in the size of the average farm over time, accompanied by a decline in the number of farms, has been documented in many countries, but the farm growth process seems to be far from uniform. For the United States, Hoppe et al. (2007) reported that while the number of large farms (at least $250K in sales) grew steadily from 1982 to 2002, the number of small farms (under $10K in sales) declined from 1982 to 1992 but increased from 1992 to 2002. Several studies using U.S. data have emphasized the importance of government's involvement in agriculture on farm growth and other aspects of structural changes in agriculture (e.g., Huffman and Evenson, 2001; Ahearn et al., 2005; Goodwin et al., 2007; Key and Roberts, 2007). Sumner and Leiby (1987) found that dairy farm growth in the Southern U.S. is negatively related to farm size, indicating convergence of farm sizes over time. Shapiro et al. (1987) had similar conclusions for Canadian farms.

Several empirical studies identified a convergence towards a bimodal size distribution. Garcia et al. (1987) found, using Markov analysis, that medium-size Illinois cash grain farms exhibit the fastest growth rate. Weiss (1999) found that intermediate-size Austrian farms either grow fast and specialize in farming or grow slowly and supplement their income with non-agricultural earnings. Rizov and Mathijs (2003), Juvančič (2005), and Dolev and Kimhi (2007) found a similar bimodal convergence of farm size in Hungary, Slovenia and Israel, respectively,

However, McErlean et al. (2004) concluded that farm growth is independent of initial farm size in Northern Ireland, while Kostov et al. (2005) showed, using quantile regressions, that the smallest dairy farms have lower growth rates, while growth is proportionate to size throughout the remaining parts of the size distribution. Bremmer et al. (2002) have also failed to find significant size effects on growth of Dutch farms.

Theoretically, the heterogeneity of observed farm growth patterns across countries and over time can be attributed to the evolutionary nature of the farm growth process due to limited resource mobility (Chavas 2001) and/or imperfect information that leads to a learning process (Pakes and Ericson, 1998). Empirically, most of the studies in this literature may be subject to an omitted variable bias due to unobserved farm efficiency. If larger farms are less efficient, as suggested by the development economics literature (e.g., Carter, 1984; Feder, 1985), or more efficient, as suggested by Morrison Paul et al. (2004), but farm efficiency is not observed, a significant relation between farm size and farm growth may be observed even if a true relation does not exist.

The purpose of this paper is to examine the farm size-growth relation after controlling for farm efficiency. We employ a stochastic frontier framework in order to estimate a technical efficiency index for each farm, and use this index as an explanatory variable in a farm growth regression, controlling for endogeneity of initial farm size and for selectivity due to farm survival. We apply this empirical framework to panel data on Israeli family farms, and find that the observed relation between farm size and farm growth becomes more robust after accounting for technical efficiency. The Israeli context and the data are described in the next section, and after that we present the empirical approach. The following section includes the empirical results, and the final section offers concluding comments.

Data

We use panel data on Israeli family farms for the years 1971, 1981 and 1995. The later part of this period was characterized by extreme turbulence in the farm sector. During the 1970s the farm sector was relatively stable due to the generous farm support policies that also involved almost unlimited availability of cheap credit. Farm growth was facilitated mainly by the gradual opening of export markets for fresh produce, in particular fruits and flowers. Towards the end of the decade and into the early 1980s the government gradually reduced its involvement in the planning and support of agriculture, and the Israeli economy as a whole became unstable due to the acceleration of inflation. The 1985 anti-inflationary policy resulted in a sharp rise in the real rate of interest, and caught the farm sector in deep short-term debt that could not be serviced (Kislev et al., 1991). This has lead to the collapse of the cooperative system that governed the vast majority of farm activity in the country. Exit from agriculture and other structural changes accelerated as a result of the crisis. As farm income continued to decline, remaining farmers had to increase the scale of their operation in order to make a living, and/or diversify to other income-generating activities. Another factor that contributed to the structural change in agriculture and especially to farm growth was the increased availability of foreign workers since the early 1990s (Kislev 2003). This allowed farms that were initially limited by the availability of labor to expand faster.

The data are extracted from the two recent Censuses of Agriculture, 1971 and 1981, and a 1995 representative farm survey, all conducted by the Central Bureau of Statistics in Israel. We focus on family farms in cooperative villages (*Moshavim* in Hebrew), because these are the farms for which we could link the records across time periods and generate a longitudinal file. About a third of all cultivated land in Israel is in cooperative villages, and they include more than a half of the self employed in agriculture. A family farm in cooperative villages is a physical unit that is easy to identify and track over time. The 1971 Census data set includes 21,929 family farms, while the 1981 Census data set includes 27,047. The increase in the number of farms is in part due to establishment of new cooperative villages between 1971 and 1981, and in part due to a more inclusive definition of a farm in 1981, with the latter responsible for about three quarters of the increase. A farm record could be matched across the Census data sets if the farm remained in the hands of the same extended family. We were able to match 15,382 farm records in this way.

The 1995 farm survey covered about 10% of the farms in cooperative villages. Of the roughly 3,000 observations, about half were successfully matched to the 1981 Census records. It should be noted that matching was not successful in certain villages because farm identification numbers were changed in those villages between 1981 and 1995. We consider this as an exogenous selection mechanism. Obviously, another reason for lack of matching was a transfer of ownership, which is not exogenous, but because of the backward matching process we cannot identify this type of selectivity and hence cannot do much about this. A total of 1,040 farms could be identified and matched across all three periods. These are farms (but not all farms) that remained in the hands of the same extended family from 1971 to 1995.

The description of the data and the matching process makes it clear that it is impossible to track entry and exit of farms using these data. We employ a rather narrow definition of exit that we are able to identify, namely farms that stopped producing between two consecutive data periods, conditional on remaining in the hands of the same extended family. Thus, we are not able to account for farm exit that is accompanied by the sale of the farm outside the family. It should be noted that selling a farm in Israeli cooperative villages involves selling the whole farm unit including the family residence. This limits the attractiveness of this type of farm exit and enables us to identify farm families that stopped operating their farms for all practical purposes but keep the farm for residential purposes. The data show that less than 4% of farms in our sample became inactive between 1971 and 1981, while another 16% became inactive between 1981 and 1995. This is consistent with the relative stability of Israeli agriculture during the 1970s and the turbulent subsequent periods, described above. Ahituv and Kimhi (2006) have shown that the overall exit rate among Israeli farmers had a similar pattern during those periods. Nevertheless, they reported overall exit rates that are much higher than in our limited panel, and concluded that entry and exit were responsible for most of the observed changes in farm size between 1981 and 1995.

We measure farm size by the real value of output. There is more than one way to measure farm size (Lund 1983, 2005). However, Yee and Ahearn (2005) have shown that alternative size concepts do not affect the farm growth results in a significant way. We have therefore chosen the simplest measure that was available for all three periods. Most researchers use the size of operated land as a measure of farm size. Weiss (1999), on the other hand, used the number of livestock as a measure for farm size in Austria. For Israeli cooperative-village family farms, which tend to be diversified despite their relatively small size, and engage in both crop and livestock enterprises, a measure of output is more appropriate than either land-based or livestock-based measures. It should be noted that the value of output that we use is computed normatively, whereas for each type of crop or livestock, the plot size or the number of livestock is multiplied by an average coefficient of output that varies only by geographic location. In this sense this measure of size mostly reflects the volume of inputs used on the farm and the choice of output portfolio rather than actual output. In particular, it does not reflect individual farm productivity or price heterogeneity.

Figure 1 demonstrates the considerable shift to the right of the farm size distribution. Between 1971 and 1981, the average family farm grew at about 7% annually, while the annual rate of growth between 1981 and 1995 was about 5.5%. These rates of growth are higher than the rate of increase in the quantity index of output in Israel as a whole reported by Kislev and Vaksin (2003). This could reflect a faster farm growth in cooperative villages relative to other sub-sectors, and/or selectivity due to survival that is biased towards larger farms. Ahituv and Kimhi (2006) divided this quantity index by the number of self-employed farmers and obtained somewhat lower growth rates for 1971-1981 but much higher growth rates for 1981-1995. This reflects the higher rate of exit from farming in the latter period.

Figure 2 shows Lorenz curves for farm size in the three years, for all farms and for active farms. Comparing the two parts of the figure, it can be seen that the increase in farm size inequality between 1981 and 1995 almost disappears after limiting the sample to active farms. This implies that even our minimal definition of farm exit (farms becoming inactive) is a crucial ingredient of the analysis of the Israeli data, and confirms the important role of farm survival in the analysis of farm growth at the micro level in general.

Empirical framework

The literature on firm growth was stimulated by the observed empirical regularity that the firm's growth rate declines with its size. The modeling approach has gone through an evolutionary process. Early models were based on stochastic growth processes, while later models offered frameworks in which growth depends on firm decisions as well (Sutton 1997). Some of these models focused on economies of scale in production and/or marketing. Jovanovic (1982) suggested that heterogeneous firms learn gradually about their ability, and then decide to grow or exit the industry. These theoretical developments have lead to a series of empirical applications. Evans (1987) estimated farm growth as a function of initial size and its square as explanatory variables. He also corrected for selectivity due to firm survival. Hall (1987) extended this model to account for endogeneity of initial firm size, and also used a third-degree polynomial of initial size to explain firm growth. Weiss (1999) applied this approach to a three-period panel of Austrian family farms, using the first period of data to instrument second-period farm size, which in turn was used to explain farm growth between the second and third periods. Given the nature of our data, this is the empirical model we adopt in this paper.

Earlier studies of the farm size evolution in Israel did not explicitly consider the role of farm survival. Kahanovitz et al. (1999) offered a rather descriptive analysis of farm growth, emphasizing its dependence on geographical conditions and institutional factors. Ahituv and Kimhi (2006) emphasized the interdependence between farm size and off-farm labor participation, but did not explicitly consider the dynamic aspects of farm growth. Kimhi and Rekah (2006) estimated a dynamic model of farm size for the years 1992-2001, but used village-level data, which obviously did not allow for the treatment of farm survival.

We specify a log-linear regression of farm growth (G) on initial size (Y) and its square and a set of additional explanatory variables (**X**), where $G_t = \ln Y_t - \ln Y_{t-1}$:

(1) $$G_t = \alpha_1 \ln Y_{t-1} + \alpha_2 (\ln Y_{t-1})^2 + \mathbf{X}_{t-1}\boldsymbol{\beta} + u_t$$

Potential endogeneity of Y_{t-1} is evident from the definition of G_t. Hence, we use time t-2 explanatory variables as instruments for Y_{t-1}. This implies that we can only estimate equation (1) for t=1995, where 1971 variables are used as instruments for 1981 farm size.

In order to correct for selectivity due to farm survival that is not independent of farm growth, we introduce a latent survival equation:

(2) $D_t = \gamma_1 \ln Y_{t-1} + \gamma_2 (\ln Y_{t-1})^2 + \mathbf{Z}_{t-1}\boldsymbol{\delta} + v_t$

where observed survival is defined as:

(3) $$d_t = \begin{cases} 1 & D_t > 0 \\ 0 & \text{otherwise} \end{cases}$$

We also explicitly specify that growth is observed only for farms that survived:

(4) $$G_t^o = \begin{cases} G_t & d_t = 1 \\ 0 & \text{otherwise} \end{cases}$$

Assuming that u_t and v_t are jointly distributed as bivariate normal, we estimate the model (1)-(4) using the maximum likelihood approach of Heckman (1979).

In order to estimate technical efficiency, we use the stochastic frontier production function estimation procedure (Coelli et al., 1998). Suppose that firm *i* (out of *n*) has a Cobb-Douglas stochastic production function. After a logarithmic transformation, it can be specified as:

(5) $\ln(Q_i)=\beta_0+ \Sigma_{j=1\ldots J}[\beta_j \ln(W_{ij})]+\varepsilon_i$

where Q is farm output and W are inputs. In addition, the stochastic term ε_i can be specified as a sum of two elements: $\varepsilon_i = \delta_i - \mu_i$, where δ is a "conventional" random term distributed normally with mean zero and standard deviation σ_δ, and μ is a half normal positive random variable with standard deviation σ_μ, considered as an unobserved technical inefficiency index. The extreme case of μ=0 represents maximum efficiency. Assuming that μ and δ are uncorrelated with each other and with the explanatory variables in W, the coefficients β can be estimated using maximum likelihood methods (Kumbhakar and Lovell, 2000). Then, the technical efficiency index can be computed as $E(\exp(-\mu_i \mid \varepsilon_i)$, where $E(\cdot|\cdot)$ is the conditional expectation operator. The exact derivation can be found in Kumbhakar and Lovell (2000). This technical efficiency index is added to the two sets of explanatory variables **X** and **Z** in (1) and (2), respectively.

As explanatory variables in **X**, we use demographic characteristics of the farm household, and village location and year of establishment. Demographic variables include age and age squared, and a set of country-of-birth dummies, all reported for the farm operator. Also included is household size. We have also tried to include education of the head of household, but it resulted in many missing values and did not seem to affect the results significantly. Village location is represented by a set of regional dummies, and village establishment year is also grouped categorically.

As explanatory variables in **W**, we use landholdings, capital stock, the farm labor input of the farm operator, the farm labor input of other family members, all in logs. We also include the level of farm specialization. Capital stock is measured in fixed prices, and excludes the value of land. Labor input is measured as an index ranging from 0 to 110 for each person, with 110 indicating that the person is working full-time on the farm. Specialization is measured by ΣS_i^2, where S_i is the share of crop i in total output. This measure tends to zero when the number of different crops tends to infinity, and is equal to one when the farm is cultivating a single crop. **Z** includes all variables in **X** as well as several farm characteristics (landholdings, capital stock, farm labor input, and specialization).

Table 1 compares the means of these explanatory variables across the three periods. The process of ageing of farm operators is evident, but since the increase in average age is lower than the number of years between surveys, it indicates a gradual replacement of older operators by their younger successors. This generational shift is also reflected in the increased fraction of Israeli-born farm operators. There is a gradual decrease in household size, parallel of the trend in the country as a whole. Farm size has increased dramatically, as discussed above. The size of landholdings went down, especially between 1971 and 1981, and capital stock more than doubled between 1971 and 1981 (see Ahituv and Kimhi 2002), but declined by almost 50% between 1981 and 1995. The level of farm specialization increased over the years, especially between 1981 and 1995.

Results

Table 2 summarizes the regression results. The first column shows the estimated OLS coefficients of the 1981-1995 farm growth equation (1), using only farms that were active in both 1981 and 1995. The coefficients of initial size and its squared value are negative and positive, respectively. This implies that the rate of farm growth is declining with farm size up to a certain size threshold, and beyond that threshold the growth rate starts increasing with farm size. The coefficients imply that the declining segment covers the entire 1981 farm size distribution, hence we have convergence of farm size among active farms over time. Other statistically significant effects are obtained for the ethnic origin dummies (negative) and for household size (positive).

The second column shows the coefficients of the farm growth equation (1) estimated using the Heckman selection model specified in equations (1)-(4). The coefficients of the farm survival equation (2) appear in appendix 1. The Wald test for the correlation between the error terms in the survival equation and the growth equation cannot reject the hypothesis that the residuals are uncorrelated. However, the coefficients of initial farm size and its square are much smaller in absolute value that the OLS coefficients, and are not statistically significant. This implies that according to this model, farm growth is predominantly random.

Table 3 shows the coefficients of the stochastic frontier model (5), estimated using 1981 data. The coefficients of land, capital and labor sum up to about 0.8, implying decreasing returns to scale. Capital seems to be the most important input of production. Landholdings have a negative coefficient, meaning that output is smaller on land-intensive farms. This is because land-intensive farms tend to specialize in field crop, one of the least profitable farm outputs, while farms with less land tend to specialize in livestock and in vegetables and flowers under cover. The labor input of the head of household is more productive than the labor input of other family members. The coefficient of specialization is negative, implying that specialized farms are less efficient. This is consistent with the findings of Morrison Paul and Nehring (2005) for the U.S. The hypothesis that efficiency is uniform across farms is strongly rejected. Using these results, a technical efficiency index is computed for each farm, as described above.

The last two columns in table 2 repeat the analysis after including the computed technical efficiency index among the explanatory variables. In both cases the coefficient of technical efficiency is positive and highly significant, implying that more efficient farms grow faster. This is not surprising, but the more interesting question is whether controlling for technical efficiency changes the convergence results. The coefficients of farm size and its squared value have the same pattern as before, implying that farm growth is decreasing with initial farm size for small farms and increasing with initial farm size for large farms. However, the range of sizes that are considered small and large for this purpose is very different. In the OLS growth regression, the turning point is at the 90th percentile of initial farm size, meaning that 90% of farms are within the converging range and only 10% are within the diverging range. This is not very different from the result we obtained before introducing technical efficiency. After correcting for endogeneity of initial farm size and selectivity due to farm survival, we find that the turning point is at the 13th percentile of initial farm size, meaning that only 13% of farms are within the converging range, while 87% are within the diverging range. Recall that before introducing technical efficiency, the coefficients of initial farm size and its squared value were not statistically significant, leading to the conclusion that farm growth is predominantly random. Here, these coefficients are highly significant, implying a systematically diverging growth pattern. Hence, controlling for technical efficiency, when studying farm size convergence, has led to a substantially different conclusion in our case study of Israeli family farms.

Concluding comments

We have analyzed the growth of family farms in Israeli cooperative villages between 1981 and 1995, using longitudinal data. We followed the empirical approach of Weiss (1999) by focusing on the potentially nonlinear effect of initial farm size on its subsequent growth, by using instrumental variable techniques to account for the endogeneity of initial farm size, and by correcting for selectivity due to non-random survival of farms throughout the period of analysis. Our results support the earlier findings that farm growth is non-linear in initial farm size, and that both endegeneity and sample selection are important in this kind of analysis. In addition, we introduced an estimate of farm technical efficiency into the analysis, and found that it is an important determinant of farm growth. More importantly, the farm growth pattern changes considerably, after including technical efficiency among the explanatory variables. In our case study, we found evidence of farm size divergence over time after including technical efficiency. Without controlling for technical efficiency, the farm growth process seemed to be predominantly random.

We conclude that previous studies of farm growth could have suffered from serious omitted variable bias due to the exclusion of farm technical efficiency. Another important implication of our study is that the estimated farm growth pattern is quite sensitive to model specification. This is because our OLS and sample selection models produced qualitatively different results, despite the fact that the hypothesis of no sample selectivity could not be rejected. We also learned that it is not sufficient to look at the signs of the coefficients of initial farm size and its squared value and make conclusions about farm size convergence and divergence. Our results theoretically implied a convergence of small farm and divergence of large farms, as has been found in several previous studies, but further examination revealed that the vast majority of farms are within the diverging segment of the farm size distribution.

We found that growth is faster in larger farm households, which could indicate that family labor is still important, perhaps for the supervision of hired workers, even when farms grow and become more commercialized. However, this result should be evaluated with caution, since household size is not necessarily exogenous to farm size: succeeding children and their families may be more likely to join more profitable family farms that are also growing faster. A more complete analysis of farm growth would involve these succession considerations. This is left for future research, and will necessarily require longer panels of data.

If family farms in Israel continue to diverge in size as implied by our empirical results, this could have important implications for the farm sector as a whole. First and foremost, the fact that Israeli farmers are still subject to historical non-tradable quotas of land and water implies that increased diversity of farm sizes could increase the inefficiencies associated with the quota system. Second, we have seen in the past that increased specialization and heterogeneity have led to the collapse of cooperation. Still, given that the three major agricultural inputs, namely land, water and foreign labor, are controlled and regulated by the government, the political process implies that the farm sector must gather forces in order to advocate and affect policy in order to reach common goals such as keeping the rights to the land, controlling the price of water in times of shortage, and ensuring a stable supply of foreign workers. Increased polarization in the farm size distribution may result in different interests of farms in different size categories, and this could hamper their ability to play effectively in the political field. Finally, there may also be social ramifications for farm communities, which are already struggling to redefine their identity given the increased proportions of non-farm families in these communities (Kimhi and Rekah, 2008). All this does not lead to the conclusion that policy makers should find ways to keep family farms more equal in size. Rather, policy makers should realize that the social costs of the outdated quota system are increasing over time and find ways to enhance the performance of the almost missing markets of land, water, and foreign labor.

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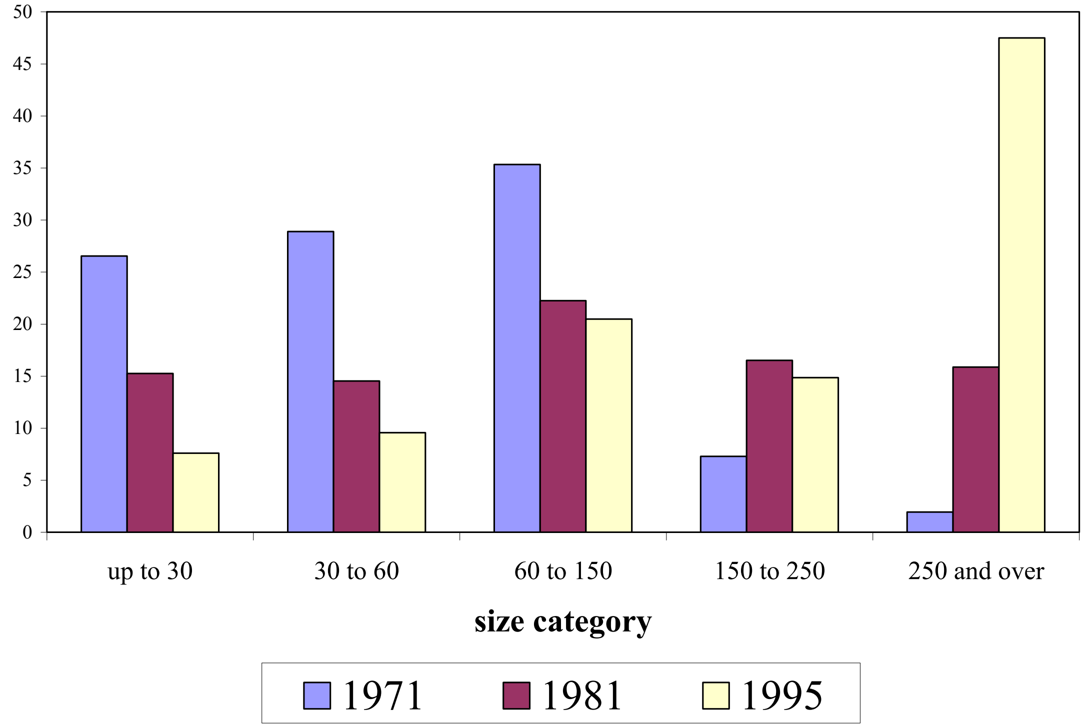
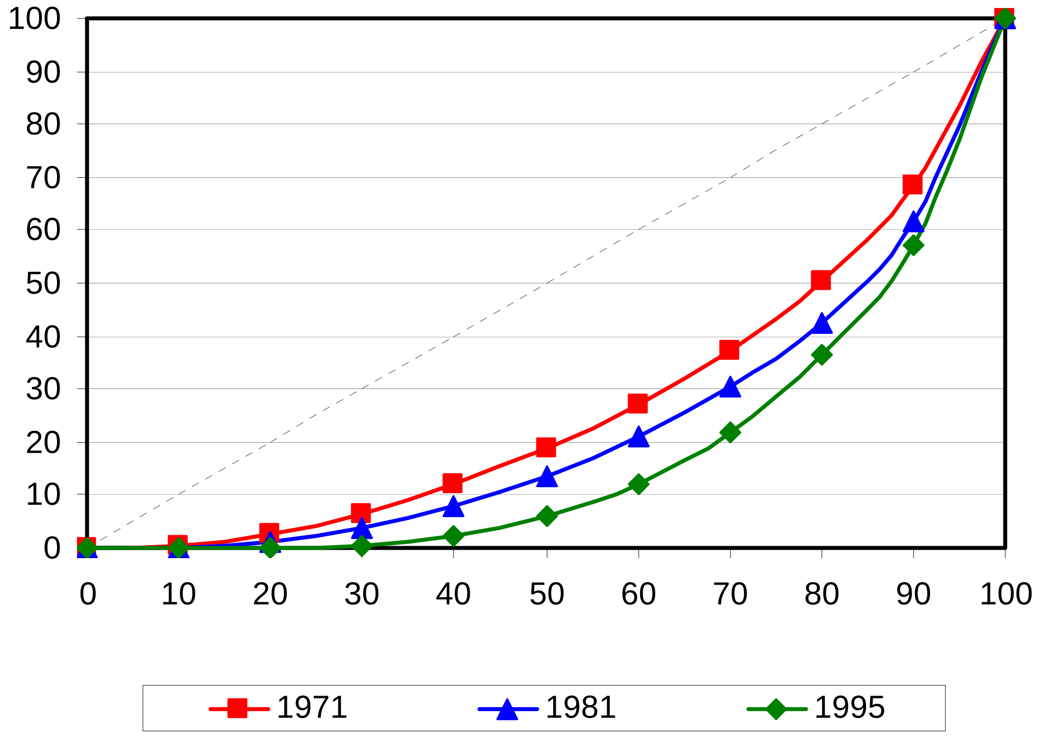


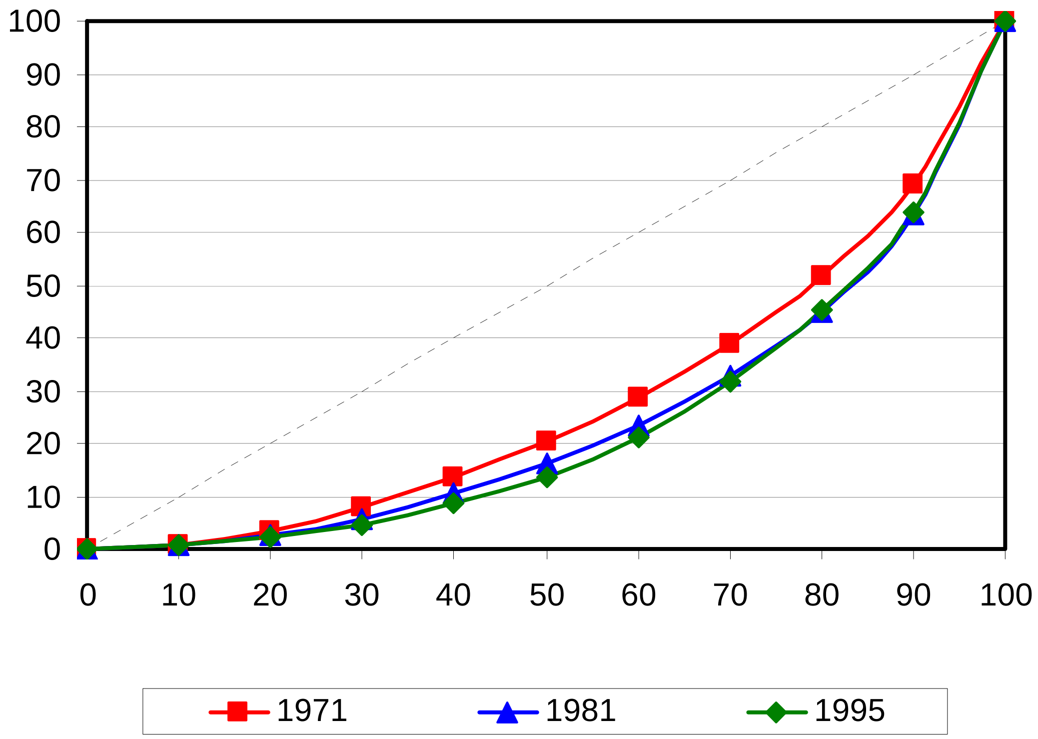
Note: Farm size is a continuous measure, based on the normative value of farm output (see text), evaluated at 1995 prices, and expressed in units of NIS 1,000. The size categories in the figure were determined arbitrarily.

Figure 1. Farm Size Distributions in Israel: 1971-1995



A.

A. All farms



B.

B. Active farms

Figure 2. Lorenz Curves of Farm Size Distributions

Table 1. Means of Explanatory Variables

Variable	1971	1981	1995	Units
Age	44	50	56	years
Country of birth				
Israel	9.3	19.3	35.2	percent
Europe/America	27.5	21.2	14.6	percent
Asia/Africa	63.3	59.5	50.2	percent
Household size	5.86	5.65	5.34	people
Farm size	81	152	264	NIS 1,000 (1995)
Landholdings	5.7	3.6	3.0	hectares
Capital stock	168	454	233	NIS 1,000 (1995)
Operator's farm labor	73	61	58	index (full time =110)
Family farm labor	76	48	55	index (full time =110)
Specialization	69	74	83	percent
Region*				
Golan and Upper Galilee		7.1		
Northern valleys		10.1		
Haifa and Akko		7.3		
Central plains		34.7		
Southern plains		18.6		
Jerusalem		6.1		
South		16.1		
Establishment year*				
Up to 1947		17.5		
1948-1956		72.5		
1957 and up		10.0		

* Village location and establishment year are naturally constant over time

Table 2. 1981-1995 Farm Growth Results

Variable	Without T.E. OLS	Without T.E. HECKMAN	With T.E. OLS	With T.E. HECKMAN
Farm size	-1.0549 (-5.26)**	-0.2353 (-1.27)	-0.7093 (-3.43)**	-0.5336 (-3.18)**
Farm size squared	0.0722 (3.16)**	0.0210 (0.85)	0.0579 (2.57)**	0.0738 (3.29)**
Technical efficiency			0.6995 (5.42)**	1.1059 (12.58)**
Age	-0.0126 (-0.60)	-0.0140 (-0.52)	-0.0251 (-1.22)	-0.0362 (-1.50)
Age squared	0.0001 (0.67)	0.0001 (0.49)	0.0003 (1.35)	0.0004 (1.54)
Europen/American origin	-0.3146 (-2.60)**	-0.3166 (-2.22)*	-0.2943 (-2.48)*	-0.2743 (-2.15)*
Asian/African origin	-0.2866 (-2.46)*	-0.2630 (-1.82)	-0.2056 (-1.78)	-0.1595 (-1.23)
Household size	0.0509 (3.07)**	0.0376 (1.98)*	0.0442 (2.71)**	0.0469 (2.76)**
Intercept	3.9960 (5.97)**	1.3522 (1.82)	2.4579 (3.44)**	1.0653 (1.61)
R^2	22.77%		25.78%	
p-value for χ^2 test	0.0000	0.0000	0.0000	0.0000
p-value for cov(u,v)=0		0.2342		0.1536
Number of cases	753	833	752	833

Notes: coefficients of regional and establishment year dummies not shown.
* coefficient significant at 5%; ** coefficient significant at 1%.

Table 3. 1981 Stochastic Frontier Estimation Results

Variable	Coefficient
Landholdings	-0.0755 (-2.91)**
Capital stock	0.7795 (27.81)**
Operator's farm labor	0.0552 (4.06)**
Family farm labor	0.0292 (2.70)**
Specialization	-0.0036 (-2.81)**
Intercept	1.0476 (5.21)**
σ_δ	0.4661
σ_μ	0.9238
Number of cases	963

* coefficient significant at 5%; ** coefficient significant at 1%.

Appendix 1: 1981-1995 Farm Survival Results

Variable	Without T.E.	With T.E.
Farm size	0.0338 (0.17)	0.0692 (0.36)
Farm size squared	0.0086 (0.29)	0.0020 (0.07)
Technical efficiency		-0.1792 (-1.50)
Age	-0.0177 (-0.48)	-0.0125 (-0.34)
Age squared	0.0001 (0.28)	0.0005 (0.14)
Europen/American origin	0.2165 (0.97)	0.2176 (0.97)
Asian/African origin	-0.1893 (-0.90)	-0.1993 (-0.95)
Household size	0.0317 (1.35)	0.0292 (1.24)
Landholdings	0.1445 (1.07)	0.1192 (0.88)
Positive landholdings	0.6007 (1.19)	0.4994 (0.98)
Operator's farm labor	0.0285 (0.85)	0.0323 (0.95)
Family farm labor	0.0057 (0.21)	0.0009 (0.03)
Capital stock	0.3865 (4.99)**	0.4052 (5.22)**
Specialization	0.0004 (0.15)	0.0012 (0.43)
Intercept	-1.8654 (-1.71)	-1.9103 (-1.76)

Notes: coefficients of regional and establishment year dummies not shown.
* coefficient significant at 5%; ** coefficient significant at 1%.

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