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Do family farms really converge to a uniform size? The role of unobserved farm efficiency*

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We analyse the growth of family farms in Israeli cooperative villages during a period of economic turmoil. We use instrumental variables to account for the endogeneity of initial farm size, and correct for selectivity as a result of farm survival. We also include a technical efficiency index, derived from the estimation of a stochastic frontier production model, as an explanatory variable. Our aim is to check whether ignoring efficiency could have been the reason for convergence results obtained elsewhere in the literature. We found that technical efficiency is an important determinant of farm growth, and that not controlling for technical efficiency could seriously bias the results. In particular, larger farms are found to grow faster over time, while without controlling for technical efficiency the farm growth process seemed to be independent of initial farm size. The increasing polarisation of farm sizes in Israel has ramifications for the inefficiencies induced by the historical quota system, for the political power of the farm sector and for the social stability of farm communities.

Key words: farm growth, farm size, farm survival, instrumental variables, sample selection, technical efficiency.

1. Introduction and background

Farm sectors in both developed and developing 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 because of 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 US data have emphasised the importance of government's involvement in agriculture on farm growth and other aspects of structural changes in agriculture (e.g., Huffman and

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Evenson 2001; Ahearn *et al*. 2005; Goodwin *et al*. 2007; Key and Roberts 2007).

A key concept in the firm size literature is the so-called Gibrat's law (Sutton 1997). Gibrat's law states that firm growth is independent of initial firm size. Gibrat's law has been used as an assumption or as an equilibrium result in industrial organisation theories (Lucas 1978). In an industry that is not in a long-run equilibrium, whether Gibrat's law holds is an empirical question (Evans 1987; Hall 1987). Several attempts have been made to examine the validity of Gibrat's law for the agricultural sector, most of which rejected Gibrat's law on empirical grounds. For example, Sumner and Leiby (1987) found that dairy farm growth in the Southern US is negatively related to farm size, indicating that farm sizes tend to converge to a uniform size over time. Shapiro *et al*. (1987) had similar conclusions for Canadian farms.

Several empirical studies identified that the farm size distribution tends to be bimodal. 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 specialise 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 pattern of tendency of farm sizes towards a bimodal distribution 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 because of 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 because 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, then a significant relation between farm size and farm growth may be observed even if a true relation does not exist.

The purpose of this article 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 because of farm survival. In principle, one would also like to control for allocative efficiency and scale efficiency, but unfortunately data on costs and prices necessary for their estimation are not available. We apply this empirical framework to panel data on Israeli family farms, and find that the inclusion of technical efficiency as an explanatory variable not only increases the explanatory power of the model but also makes the relation between farm growth and initial farm size statistically significant. The Israeli context and the data are described in Section 2, and the empirical approach is presented Section 3. Section 4 includes the empirical results, and Section 5 provides concluding comments.

2. Background on the Israeli farm sector

The data used in this research are on Israeli family farms between 1971 and 1995. The later part of this period was characterised by extreme turbulence in the farm sector. During the 1970s the farm sector was relatively stable because of 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. Towards the end of the 1970s and into the 1980s, the government gradually reduced the planning and support of agriculture, and the Israeli economy as a whole became unstable because of 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 led 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. The increased availability of foreign workers since the early 1990s also contributed to the structural changes in agriculture and especially to farm growth (Kislev 2003). It allowed farms that were initially limited by labour availability to expand faster.

It should be noted that despite the reduced involvement of the government in the planning of agriculture, rigid quota systems for land, water and foreign labour still exist, and even some production quotas (milk, eggs). While there is some extent of tradability or exchangeability in these quotas, there is no doubt that their existence makes it more difficult for farmers to adjust to market incentives. In this context, studying the changes in the size distribution of farms may lead to interesting policy implications.

The farm sector in Israel comprises of three distinct types of farms: collective farms (*Kibbutzim*), family farms in cooperative villages (*Moshavim*) and private farms (family farms and business farms). We focus here on family farms in cooperative villages on which we have much better data. About one-third of all cultivated land in Israel is in cooperative villages, they included more than a half of the self employed in agriculture in 1995, and were responsible for almost a half of total agricultural production in Israel. A family farm in a cooperative village is a physical unit that is easy to identify and track over time. Although these farms are organised in cooperatives, each one is an independent production unit and farm operators make their own production decisions (subject to a set of institutional constraints that have largely eroded over the years). More details about the institutional structure of cooperative villages and the transformation they experienced can be found in Kislev (1992), Schwartz (1999), Sofer and Applebaum (2006), and Kimhi (2009).

3. Data

The data are extracted from the two Censuses of Agriculture, 1971 and 1981, and a 1995 representative farm survey, all conducted by the Central Bureau of Statistics in Israel. The 1971 Census dataset includes 21,929 family farms, while the 1981 Census dataset includes 27,047. The increase in the number of farms is in part because of establishment of new cooperative villages between 1971 and 1981, and in part because of a more inclusive definition of a farm in 1981, with the latter responsible for about three quarters of the increase.[1] A farm record could be matched across the Census datasets if the farm remained in the hands of the same extended family (either the same operator or one of his siblings or children). We were able to match 15,382 farm records in this way.

The 1995 farm survey covered about 10 per cent of the farms in cooperative villages. Of the roughly 3000 observations in the sample, about two-thirds were from cooperative villages. About two-thirds of those 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 exogenously random selection mechanism, so that the resulting sample is assumed to be representative of the 1995 family farm population in cooperative villages. 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 control for this type of selectivity or even asses its magnitude. A total of 1040 farms could be identified and matched across all three periods. This is a representative sample of cooperative-village family farms that remained in the hands of the same extended family from 1971 to 1995. Noting that in 1995, family farms were almost 85 per cent of the agricultural enterprises in Israel, and that about three quarters of those are in cooperative villages, our sample represents a sizeable portion of Israeli agriculture.

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

[1] A farm in a Moshav is a well-defined production unit, but there are households in the Moshav who only have access a small plot of land. These were not included in the 1971 census but were included in the 1981 census. The 1995 survey included only full-size farms, as in 1971.

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 < 4 per cent of farms in our sample became inactive between 1971 and 1981, while another 16 per cent 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) found that the overall exit rate among Israeli farmers had a similar pattern during those periods. Nevertheless, their aggregate data revealed overall exit rates that are much higher than in our limited panel. Their analysis led to the conclusion 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. This is especially suitable for crop farms. Weiss (1999), on the contrary, 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- or livestock-based measures. It should be noted that the value of output that we use is not measured directly but computed using norms: for each type of crop or livestock, the plot size or the number of livestock is multiplied by an average 'norm' of output per unit of land or livestock. These norms were derived from detailed production surveys by the Central Bureau of Statistics, and they vary 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. Recall that farm size is a continuous measure, based on the normative value of farm output (see above), evaluated at 1995 prices, and expressed in units of NIS 1000 (the size categories in the figure were determined arbitrarily). Between 1971 and 1981, the average family farm grew at about 7 per cent annually, whereas the annual rate of growth between 1981 and 1995 was about 5.5 per cent. 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

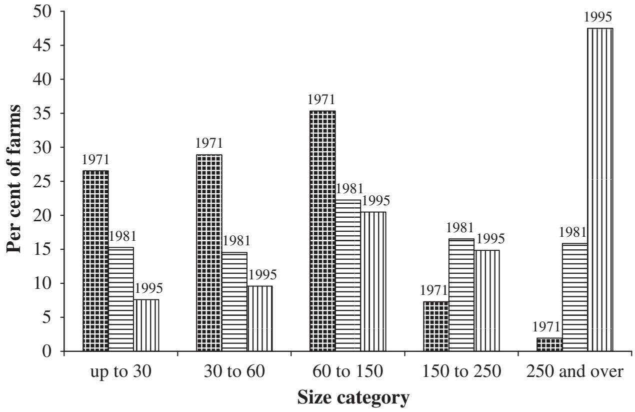

Figure 1 Farm size distributions in Israel: 1971–1995.

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 3 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 definition of farm exit (farms becoming inactive) is a crucial ingredient for 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.

4. Empirical framework

The literature on firm growth was stimulated by the observed empirical regularity that the firm's growth rate declines with its size, a violation of Gibrat's law. The modelling approach has gone through an evolutionary process. Early models were based on stochastic growth processes, whereas 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 led 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 because of 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 instru-

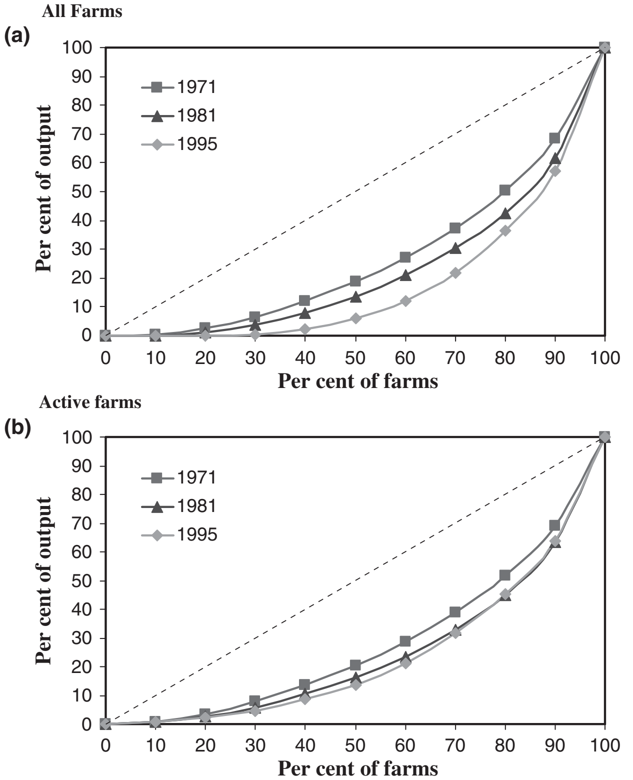

Figure 2 Lorenz curves of farm size distributions.

ment 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 study.

Earlier studies of 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, emphasising its dependence on geographical conditions and institutional factors. Ahituv and Kimhi (2006) emphasised the interdependence between farm size and off-farm labour 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.

It should be noted that while all the above applications used longitudinal micro-data on individual farm sizes, aggregate data can also be used to study farm size transitions using the Markov framework. Earlier applications of this framework (Garcia *et al*. 1987; Zepeda 1995) were limited by the dimensionality problem. Later applications (Karantinins 2002; Tonini and Jongeneel 2009) overcame this problem, but required prior information. Hence, for short panels with individual farm data the framework presented below is more appropriate.

We specify a log-linear regression of farm growth (G) on initial size (Y) and its square and a set of additional explanatory variables ($\mathbf{X}$), where $G_t = \ln Y_t - \ln Y_{t-1}$ (this relates to individual farms; farm subscripts are omitted for simplicity):

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

The log-linear specification is common practice in the firm growth literature. Evans (1987) tested it against alternative specifications and found it most satisfactory. 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 because of farm survival that is not independent of farm growth, we introduce a latent survival equation:

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

where observed survival is defined as:

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

and $\mathbf{Z}$ includes explanatory variables that are assumed to affect farm survival (see Appendix 1). We also explicitly specify that growth is observed only for farms that survived:

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

Assuming that u_t and v_t are jointly distributed as bivariate normal, we estimate the model Equations (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:

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

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 normally distributed random term 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 $\mu = 0$ 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 $\mathrm{TE}_i = E(\exp(-\mu_i) \mid \varepsilon_i)$, where $E(\cdot|\cdot)$ is the conditional expectation operator. The exact derivation can be found in Kumbhakar and Lovell (2000). We use only 1981 data for the estimation of technical efficiency because in this way we can compute the technical efficiency index for all observations included in the above model. This technical efficiency index is added to the two sets of explanatory variables **X** and **Z** in Equations (1) and (2), respectively, to give:

$$G_t = \alpha_1 \ln Y_{t-1} + \alpha_2 (\ln Y_{t-1})^2 + \mathbf{X}_{t-1}\boldsymbol{\beta} + \varphi \mathrm{TE}_{t-1} + u_t \qquad (1)'$$

$$D_t = \gamma_1 \ln Y_{t-1} + \gamma_2 (\ln Y_{t-1})^2 + \mathbf{Z}_{t-1}\boldsymbol{\delta} + \tau \mathrm{TE}_{t-1} + v_t \qquad (2)'$$

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. **Z** includes all variables in **X** as well as several farm characteristics (landholdings, capital stock, farm labour input, and specialisation).

As explanatory variables in **W**, we use landholdings, capital stock, the farm labour input of the farm operator, and the farm labour input of other family members, all in logs. We also include the level of farm specialisation. Capital stock is measured in fixed prices, and excludes the value of land which is not available because the market for farmland is too thin. Labour 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. Specialisation is measured by ΣS_i^2, where S_i is the share of branch i in total farm output (branches include field crops, fruits, vegetables, dairy, poultry and other livestock). This measure tends to zero when the number of different active branches tends to infinity, and is equal to one when the farm is specialising in a single branch.

Table 1 compares the means of the explanatory variables in **X** and **Z** 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 to the trend in the country as a whole. Farm size has increased

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	%
Europe/America	27.5	21.2	14.6	%
Asia/Africa	63.3	59.5	50.2	%
Household size	5.86	5.65	5.34	People
Farm size	81	152	264	NIS 1000 (1995)
Landholdings	5.7	3.6	3.0	ha
Capital stock	168	454	233	NIS 1000 (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	%
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			

Farm size is expressed in units of NIS 1000 in 1995 prices.
*Village location and establishment year are naturally constant over time.

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 per cent between 1981 and 1995. The level of farm specialisation increased over the years, especially between 1981 and 1995.

5. Results

Table 2 summarises the regression results. The first column shows the estimated ordinary least square (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 for relatively small farms up to a certain size threshold (that can be easily derived from the coefficient estimates), but it is increasing with farm size for relatively large farms, above the size threshold. It should be noted that the joint hypothesis in which neither initial farm size nor its square affect farm growth was strongly rejected (R^2 dropped to 7 per cent when the two variables were excluded). The coefficients imply that the size threshold beyond

Table 2 1981–1995 Farm growth results

Variable	Without TE		With TE	
	OLS	Heckman	OLS	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)
European/ 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 $\mathrm{cov}(u,v) = 0$		0.2342		0.1536
Number of cases	753	833	752	833

Coefficients of regional and establishment year dummies not shown.
t statistics in parentheses.
*Coefficient significant at 5%; **Coefficient significant at 1%.

which farm growth starts increasing with farm size is larger than practically all of the farms in the sample. Hence we conclude that the farm size distribution becomes more concentrated around its mean over the years among active farms. 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 in which the residuals are uncorrelated. However, the coefficients of initial farm size and its square are much smaller in absolute value than the OLS coefficients, and are not statistically significant. This implies that according to this model, farm growth is not systematically related to farm size.

Table 3 shows the coefficients of the stochastic frontier model Equation (5), estimated using 1981 data. Indeed, it would have been preferred to estimate this model using panel data, but using panel data would have led to the loss of a considerable number of observations. The coefficients of land, capital and labour 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-abundant

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

Notes: *t* statistics in parentheses.
*Coefficient significant at 5%; **Coefficient significant at 1%.

farms. This is because land-abundant farms tend to specialise in field crops, one of the least profitable farm branches, whereas farms with less land tend to specialise in livestock and in vegetables and flowers under cover, more profitable branches. The labour input of the head of household is more productive than the labour input of other family members. The coefficient of specialisation is negative, implying that more specialised farms are less efficient, after controlling for other covariates. This is consistent with the findings of Morrison Paul and Nehring (2005) for the US. 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 kernel density estimate of the distribution of the estimated technical efficiency index is shown in Figure 3.

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 much interesting question is whether controlling for technical efficiency changes the conclusions regarding the reduced cross-sectional variability of farm size over time. 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 relatively small farms, up to a certain size threshold, and increasing with initial farm size for relatively large farms, beyond the threshold. However, the size threshold beyond which farm growth increases with farm size is very different. In the OLS growth regression, the threshold is at the 90th percentile of initial farm size, meaning that for 90 per cent of the farms, farm growth declines with farm size. This is not very different from the result we obtained before introducing technical efficiency. After correcting for endogeneity of initial farm size and selectivity because of farm survival, we find that the threshold is at the 13th percentile of initial farm size, meaning that farm growth decreases with farm size for only 13 per cent of farms, whereas for 87 per cent of the farms, farm growth increases with farm size.

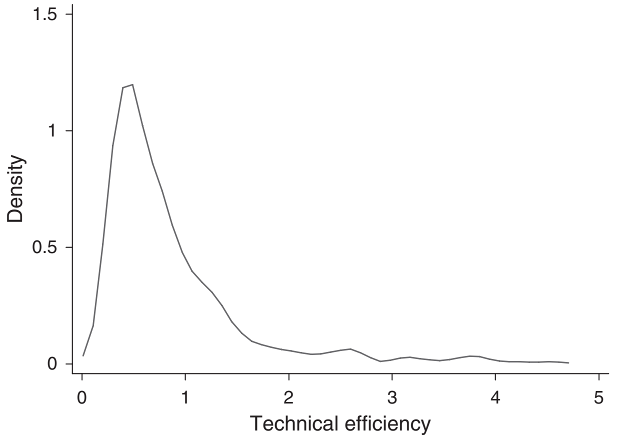

Figure 3 Kernel density estimate of the distribution of the estimated technical efficiency index.

Recall, 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 not systematically related to farm size. Here, these coefficients are highly significant, implying a systematic positive relation between farm growth and farm size except for the smallest farms. Hence, controlling for technical efficiency, when studying the evolution of the farm size distribution over time, has led to a substantially different conclusion in our case study of Israeli family farms.

The robustness of these results to the specification of the stochastic frontier model is verified by varying the distributional assumptions on μ (half normal, exponential, and truncated normal) as well as by specifying the production function as the more flexible translog instead of Cobb-Douglas (Equation (5)). Appendix 2 shows that the main results are qualitatively unchanged.

6. Concluding comments

We have analysed the growth of family farms in Israeli cooperative villages between 1981 and 1995, using panel 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 because of 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 endogeneity 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 that larger farms tend to grow faster over time after including technical efficiency. Without controlling for technical efficiency, the farm growth process seemed to be independent of initial farm size.

We conclude that previous studies of farm growth could have suffered from serious omitted variable bias because of the exclusion of farm efficiency. This conclusion comes with a caveat, that the current analysis could also suffer from an omitted variable bias because of the exclusion of allocative and scale efficiencies. Further research, using adequate data that is unfortunately not available for the current study, should examine this issue. 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, while 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 in order to make conclusions about whether the variability of farm sizes increases or decreases over time. Our results theoretically implied a concentration of farm sizes among the smallest farms, and faster farm growth among larger farms, as has been found in several previous studies, but further examination revealed that the vast majority of farms are within the size range of faster growth.

We found that growth is faster in larger farm households, which could indicate that family labour is still important, perhaps for the supervision of hired workers, even when farms grow and become more commercialised. 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, then 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 specialisation and heterogeneity have led to the collapse of cooperation. Still, given that the three major agricultural inputs, namely land, water and foreign labour, 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 polarisation 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 2009). 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 realise 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 labour.

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Appendices

Appendix 1 1981–1995 Farm survival results

Variable	Without TE	With TE
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)
European/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 labour	0.0285 (0.85)	0.0323 (0.95)
Family farm labour	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. *t* statistics in parentheses. *Coefficient significant at 5%; **Coefficient significant at 1%.

Appendix 2 Robustness checks to the specification of technical efficiency

Model/Specification	Coefficient of farm size	Coefficient of farm size squared
OLS		
Half normal distribution	−0.7093 (−3.43)**	0.0579 (2.57)**
Exponential distribution	−0.5685 (−2.65)**	0.0395 (1.71)
Truncated normal distribution	−0.5754 (−2.68)**	0.0406 (1.76)
Translog specification	−0.7766 (−3.73)**	0.0584 (2.56)*
Heckman		
Half normal distribution	−0.5336 (−3.18)**	0.0738 (3.29)**
Exponential distribution	−0.4779 (−2.86)**	0.0655 (2.93)**
Truncated normal distribution	−0.4828 (−2.89)**	0.0662 (2.96)**
Translog specification	−0.4694 (−2.75)**	0.0633 (2.78)**

Notes: The first three specifications are for different distributional assumptions on technical efficiency.
The translog specification is for the production function.
t statistics in parentheses.
*Coefficient significant at 5%; **Coefficient significant at 1%.