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# **Long-Run Trends in the Farm Size Distribution in Israel: The Role of Part-Time Farming**

Ayal Kimhi and Nitzan Tzur  
Department of Agricultural Economics and Management  
Faculty of Agriculture of the Hebrew University  
P.O. Box 12, Rehovot 76100, Israel



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## Abstract

This article proposes a nonparametric analysis in which the change in the distribution of farm size between two periods is decomposed into several components, and the contributions of subgroups of farms to this change are analyzed. Using data on Israeli family farms, we analyze the changes in the farm size distribution in two separate time periods that are characterized by very different market conditions, focusing on the different contributions of full-time farms and part-time farms to the overall distributional changes. We find that between 1971 and 1981, a period characterized by stability and prosperity, the farm size distribution has shifted to the right with relatively minor changes in higher moments of the distribution. On the other hand, between 1981 and 1995, a largely unfavorable period to Israeli farmers, the change in the distribution was much more complex. While the overall change in the size distribution of farms was smaller in magnitude than in the earlier period, higher moments of the distribution were not less important than the increase in the mean. Between 1971 and 1981 the contributions of full-time farms and part-time farms to the change in the size distribution are quite similar. Between 1981 and 1995, however, full time farms contributed mostly to the growth in the average farm size, while average farm size among part-time farms actually decreased, and their contribution to the variance of farm size was quantitatively larger. We conclude that the contribution of part-time farming to the increase in farm size inequality is not straightforward. Rather, it depends on the economic environment.

## Introduction

A well-known stylized fact in agricultural economics is that the number of farms in developed economies declines over time while the size of the average farm increases. These trends have been documented and analyzed for the U.S. (e.g., Huffman and Evenson 2001; Ahearn *et al.* 2005; Key and Roberts 2007), Canada (Shapiro *et al.* 1987), Britain (Upton and Haworth 1987), Austria (Weiss 1999), The Netherlands (Bremmer *et al.* 2002), Hungary (Rizov and Mathijs 2003; Bakucs and Fertő 2009), Slovenia (Juvančič 2005), and Israel (Ahituv and Kimhi 2006; Dolev and Kimhi 2010), among other countries. The existing literature has used various regression specifications to estimate the determinants of average farm size. Some of the applications allowed farm growth to depend on initial farm size, thereby allowing for differential growth rates for farms of different sizes. The results show trends of increased concentration of farm sizes in several cases and trends of increased dispersion of farm sizes in other cases, while in some other cases no significant effect of farm size on farm growth was found.

The limitation of this line of literature is the reliance on a parametric regression model that allows for a limited class of distributional changes. Two alternatives have been proposed in the literature. Chavas and Magand (1988) and Zepeda (1995) used a Markov analysis to estimate transition probabilities between size classes. Alternatively, Kostov *et al.* (2005) and Bakucs and Fertő (2009) estimated the farm growth equation by quantile regression, thereby allowing different growth rates in different parts of the size distribution. These methods allow for more flexible changes in farm growth rates across the farm size distribution. Still, they do not capture the entire change in the farm size distribution over time.

The purpose of this article is to propose a method for examining the changes over time of the entire farm size distribution, and to identify determinants of these changes. Wolf and Sumner (2001) looked at the changes in the farm size distribution using kernel density estimates, but did not go further than a visual inspection of the density plots. We take this approach a step further. Our proposed method analyzes the changes in the size

distribution of farms by decomposing the change in the density function into changes in subgroup shares and changes in subgroup densities, after dividing the farm population into subgroups according to some key characteristics. The changes in subgroup densities are decomposed further, as suggested by Jenkins and van Kerm (2005), into changes in the location (mean), spread (variance), and higher moments of the distribution. This allows the identification of types of farms that contribute to the changes in the farm size distributions in specific ways. This approach is nonparametric in nature, and is superior to regression-based parametric approaches, such as the one proposed by Miljkovic (2005), who used a regression framework to analyze the determinants of an index of farm size inequality. Several semiparametric alternatives have been proposed in the literature. For example, Melly (2005) uses a quantile regression in order to decompose inequality into the share of covariates, the share of coefficients and the share of residuals. This allows for a richer set of covariates than the Jenkins and van Kerm (2005) procedure, but it relies on a parametric assumption about the dependence of conditional quantiles on the covariates.

We choose to divide the sample into two subgroups: full-time farms and part-time farms. A full-time (part-time) farm is a farm whose operator does not work (works) off the farm. Separating the sample into full-time and part-time farms enables to examine the interaction between farm type and the change in the farm size distribution. Previous research has shown that off-farm work is one of the most important determinants of farm growth (Ahituv and Kimhi 2006; Upton and Haworth 1987; Weiss 1999).

We use cross-sectional data on Israeli family farms for three different time periods. The first two are derived from the two recent Censuses of Agriculture, 1971 and 1981, which include the entire population of farm households. The third data source is the 1995 farm survey, covering about 10% of the population. All three data collection efforts were conducted by the Central Bureau of Statistics in Israel. We focus on family farms in cooperative villages (*Moshavim*), because for these we have the most detailed information. Using data from three periods allows us to analyze the changes in the farm size distribution in two sub-periods: 1971-81 and 1981-95. This is particularly important in the case of Israel, since the 1970s were a relatively stable and favorable period for Israeli farmers, while the latter period was characterized by much turmoil, including high inflation, a debt crisis, and hired labor shortages due to security issues. Therefore, we expect quite different trends in the farm size distribution in these two sub-periods. Figure 1 confirms this expectation. The top panel presents the kernel density plots of farm size distributions in the three time periods. While the change from 1971 to 1981 seems to be mostly an increase in the average farm size, the change from 1981 to 1995 involves both an increase in the average farm size and an increase in the variance of farm size. The two other panels in figure 1 present the kernel density plots of full-time farms and part-time farms, respectively. It can be seen that the increase in mean farm size between 1981 and 1995 is entirely due to full-time farms, while the increase in farm size inequality between those years is mostly due to part-time farms. Since the fraction of part-time farms went down from 44% in 1971 to 37% in 1981 and 28% in 1995, the changes in the farm size distribution could be rooted in the intensive margin and/or in the extensive margin. In the empirical section of this paper, we will further decompose the distributional changes and assess their relation to the full-time/part-time dichotomy.

In the next section we present the density decomposition methodology. After that we provide a more detailed description of the data we use and the measurement of farm size. Then we present the decomposition results. The final section summarizes the findings.

## Methodology

Suppose that the farm population can be divided into  $K$  different subgroups indexed  $1 \dots K$ .<sup>1</sup> The density function of the farm size distribution can be written as:

$$(1) \quad f(x) = \sum_{k=1}^K v^k \cdot f^k(x),$$

where  $f(x)$  is the density function of farm size ( $x$ ) over the entire farm population,  $v^k$  is the population share of subgroup  $k$ , and  $f^k(x)$  is the density function of farm size within subgroup  $k$ . In addition, the change in the density function between time period 0 and time period 1 can be written as:

$$(2) \quad \Delta f(x) = \sum_{k=1}^K w^k \Delta f^k(x) + \sum_{k=1}^K z^k(x) \Delta v^k = c_D(x) + c_S(x),$$

where  $c_D(x)$  is the contribution of the changes in subgroup densities,  $c_S(x)$  is the contribution of the changes in the subgroup shares, and the weights  $w^k$  and  $z^k(x)$  are defined as:

$$(3) \quad w^k = \pi \cdot v_0^k + (1 - \pi) \cdot v_1^k$$

$$(4) \quad z^k(x) = (1 - \pi) \cdot f_0^k(x) + \pi \cdot f_1^k(x)$$

where  $0 \leq \pi \leq 1$  can be chosen arbitrarily. In our application we use  $\pi = 0.5$ .

Following Jenkins and van Kerm (2005), we now move to further decompose the change in subgroup densities  $c_D(x)$  into three components: sliding, stretching and squashing. Sliding reflects a horizontal shift of the entire density function. Stretching reflects an increase in the spread of the density without changing the mean. Squashing reflects all other changes in the density function, holding the mean and the spread constant. We begin by assuming the existence of a subgroup-specific function ( $g_k$ ) that describes end-period farm size ( $x_1$ ) as a function of beginning-period farm size ( $x_0$ ):  $x_1 = g_k(x_0)$ . Using the inverse of  $g_k$ , we can express the end-period density as:

$$(5) \quad f_1^k(x) = \left| \frac{d(g_k^{-1}(x))}{dx} \right| f_0^k(g_k^{-1}(x)).$$

By using specific functional forms for  $g_k$ , we can construct specific approximations of the changes in the farm size density. For example, suppose that we choose a linear function:

$$(6) \quad x_1 = \alpha_k + \beta_k x_0.$$

Under the linearly assumption, our approximation for the farm size density is:

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<sup>1</sup> This section draws heavily on Jenkins and van Kerm (2005).

$$(7) \quad \zeta^k(x) = \left| \frac{1}{\beta_k} \right| f_0^k \left( \frac{x - \alpha_k}{\beta_k} \right).$$

Now suppose that we impose the constraint  $\beta_k = 1$ . The linear transformation  $g_k$  now reflects an additive increase of a constant number of units,  $\alpha_k$ , in the size of all farms in subgroup  $k$ . In terms of the density function, this is reflected in a horizontal shift of the entire function, which is denoted as sliding. Calibrating to the increase in average farm size, we obtain  $\alpha_k = E(f_1^k) - E(f_0^k)$ . Using these parameters, (7) is now denoted  $\zeta_1^k(x; \mu_1^k, \sigma_0^k)$ , where the subscript "0" of the standard deviation means that we keep the spread of the initial period, and the subscript "1" of the mean of the distribution means that the approximated distribution has the same mean as the actual distribution in the final period.

We now move to an alternative parameterization of (6):  $\beta_k = s$ ,  $\alpha_k = (1-s)E(f_0^k)$ . It is easy to verify that this transformation does not change the mean of farm size, but increases the standard deviation by a factor of  $s$ . Hence, the calibration to the final-period standard deviation requires setting  $s = \sqrt{\text{Var}(f_1^k) / \text{Var}(f_0^k)}$ . Using these parameters, (7) is now denoted  $\zeta_1^k(x; \mu_0^k, \sigma_1^k)$ , where the subscript "0" of the mean of the distribution means that we keep the mean of the initial period, and the subscript "1" of the standard deviation means that the approximated distribution has the same standard deviation as the actual distribution in the final period.

We can also merge these two transformations into a single transformation that allows changes in both mean and standard deviation. Calibration to final-period mean and standard deviation requires setting  $\beta_k = s = \sqrt{\text{Var}(f_1^k) / \text{Var}(f_0^k)}$  and  $\alpha_k = E(f_1^k) - E(f_0^k)$ . The resulting approximated density based on (7) is denoted as  $\zeta_1^k(x; \mu_1^k, \sigma_1^k)$ . We are now in the position to decompose the change in the subgroup density function of farm size into the three components: sliding, stretching and squashing. Note that both sliding and stretching can be obtained in two ways. Sliding, for example, is the change in the mean, but it can be conditioned on the standard deviation of either the initial period or the final period. Similarly, stretching is the change in the standard deviation, but it can be conditioned on the mean of the initial period or the final period. We solve this problem by weighting each of these possibilities in a way that leaves squashing as a residual. The resulting decomposition is:

$$(8) \quad \begin{aligned} \Delta f^k(x) &= \eta(\zeta_1^k(x; \mu_1^k, \sigma_0^k) - f_0^k(x)) + (1-\eta)(\zeta_1^k(x; \mu_1^k, \sigma_1^k) - \zeta_1^k(x; \mu_0^k, \sigma_1^k)) \\ &\quad \text{[Subgroup mean effect (sliding)]} \\ &+ \eta(\zeta_1^k(x; \mu_1^k, \sigma_1^k) - \zeta(x; \mu_1^k, \sigma_0^k)) + (1-\eta)(\zeta_1^k(x; \mu_0^k, \sigma_1^k) - f_0^k(x)) \\ &\quad \text{[Subgroup variance effect (stretching)]} \\ &+ f_1^k(x) - \zeta_1^k(x; \mu_1^k, \sigma_1^k) \\ &\quad \text{[Subgroup residual effect (squashing)]} \end{aligned}$$

The weight  $\eta$  is set at 0.5 in the empirical analysis. Once computed, (8) can be plugged into (2) to obtain the overall decomposition.

## Data

The 1971 Israeli Census of Agriculture data set includes 19,147 observations on family farms in cooperative villages, while the 1981 Census data set includes 18,614. The 1995 representative farm survey covered 2,049 farms, representing a population of 15,546 farms. This latter survey focused on active farms, and hence only farms with annual value added of more than NIS3,000 were included. Therefore, we trimmed the 1971 and 1981 samples accordingly, with thresholds that reflect the changes in the consumer price index. The resulting number of farms in 1971 and 1981, are, respectively, 19,005 and 18,499.

We measure farm size by the real value of output. This is the simplest measure that was available for all three periods. The value of output is computed "normatively", in a way that is similar to the computation of Standard Gross Margin by the European Commission. Specifically, for each type of crop or livestock, the plot size or the number of livestock is multiplied by an average coefficient of output, derived from specific field surveys, that varies only by geographic location. In this sense this normative measure of output 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. Hence, it can legitimately be considered a measure of farm size. This is particularly important because most family farms in Israel are diversified, and therefore simpler measures of size such as operated land or number of livestock are not adequate. We would have preferred to use value added rather than output to measure farm size (Lund 1983), but unfortunately value added was not computed in the 1971 census. We did repeat the 1981-1995 decomposition using value added instead of value of output, and the results were quite similar.

## Decomposition results

In this section we apply the decomposition methodology described above to the case of changes in the farm size distribution in Israel. Figure 2 shows the decomposition of the changes in the farm size distribution, for the two sub-periods, 1971-81 and 1981-95. The top panel shows the total change in the distribution. The total change is a simple vertical subtraction of the initial-period density function from the end-period density function. For both sub-periods, the top panel indicates that the farm size distribution has shifted to the right: relatively small farm sizes show mostly negative values while relatively large farm sizes show mostly positive values. This is just a replication of what we saw in figure 1.<sup>2</sup>

The remaining panels show the relative importance of the different components of the distributional changes, in each sub-period. The first observation is that the component of the share of each subgroup in the farm population is negligible. This implies that farm size transitions are driven by factors other than farms changing from full-time to part-time or the other way around. Secondly, we can see that between 1971 and 1981 the sliding component is very similar in shape to the overall change, indicating that the remaining components are relatively not important as a set. Specifically, we can see that the stretching component and the squashing component have considerably lower magnitudes compared to the sliding component, and they also effectively cancel each other in most ranges of the farm size distribution.

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<sup>2</sup> Note that the vertical scales of the 1971-81 and 1981-95 graphs are not identical, and hence the changes in 1981-95 are smaller in magnitude than the changes in 1971-81.

The situation is different in the case of the farm size distribution change between 1981 and 1995. Here, the magnitudes of the sliding, stretching and squashing components are not very different from each other. While the sliding component still indicates that farms got larger on average, the stretching and squashing components indicate that a non-negligible number of farms actually got smaller. This is similar to the conclusions of Dolev and Kimhi (2010). Hence, the phenomenon of the "disappearing middle" of the farm size distribution was much more important during the 1980s and beginning of the 1990s than during the 1970s.

Figures 3 and 4 separate the components of the decomposition into the contributions of full-time farms and part-time farms. In figure 3 we can see that the dominant sliding components of the distributional change between 1971 and 1981 are not very different for full-time and part-time farms, although for part-time farms the changes due to sliding seem to be spread relatively more evenly along the range of farm sizes. The same is true for the stretching components. However, in figure 4 we can see that the different components of the distributional changes between 1981 and 1995 are quite different among full-time and part-time farms. In particular, the top panel shows that while among full-time farms it is quite clear that the entire size distribution has shifted to the right, we observe a "disappearing middle" among part-time farms, i.e., the size distribution of part-time farms became flatter between 1981 and 1995, confirming our earlier conclusion from figure 1.

## Conclusion

Analyses of changes in farm size distributions should be based on minimal distributional assumptions. This article proposes a nonparametric analysis in which the change in the distribution between two periods is decomposed into sliding, stretching and squashing components, as well as a subgroup component if the farm population is broken down to several subgroups. We apply this method to the case of Israeli family farms, and analyze the changes in the farm size distribution in two separate time periods that are characterized by very different market conditions. In particular, we focus on the different contributions of full-time farms and part-time farms to the overall distributional changes.

Our analysis shows that between 1971 and 1981, a period characterized by stability and prosperity of the Israeli farm sector, the change in the farm size distribution is almost entirely attributed to the sliding component, i.e., the whole distribution has shifted to the right with relatively minor changes in higher moments of the distribution. In addition, the difference between the contributions of full-time farms and part-time farms to the change in the size distribution is quite small.

The change in the distribution between 1981 and 1995 was much different. This period was unfavorable to Israeli farmers, with sharp changes in the economic and institutional environment. It is not surprising that the change in the distribution was much more complex than in the earlier period. In particular, while the overall change in the size distribution of farms was smaller in magnitude, higher moments of the distribution were not less important than the sliding component. In addition, full-time and part-time farms contributed quite differently to the change in the farm size distribution, with full time farms contributing mostly to the growth in the average farm size, while average farm size among part-time farms actually decreased, and their contribution to the variance of farm size was quantitatively larger.

The method proposed in this article was proved to be quite useful, but it is still limited in its ability to examine multiple determinants of the distributional change. Of course, one can separate the sample into multiple subgroups that reflect more than one

determinant, but this still falls short of a complete multivariate analysis. This issue is left for further research.

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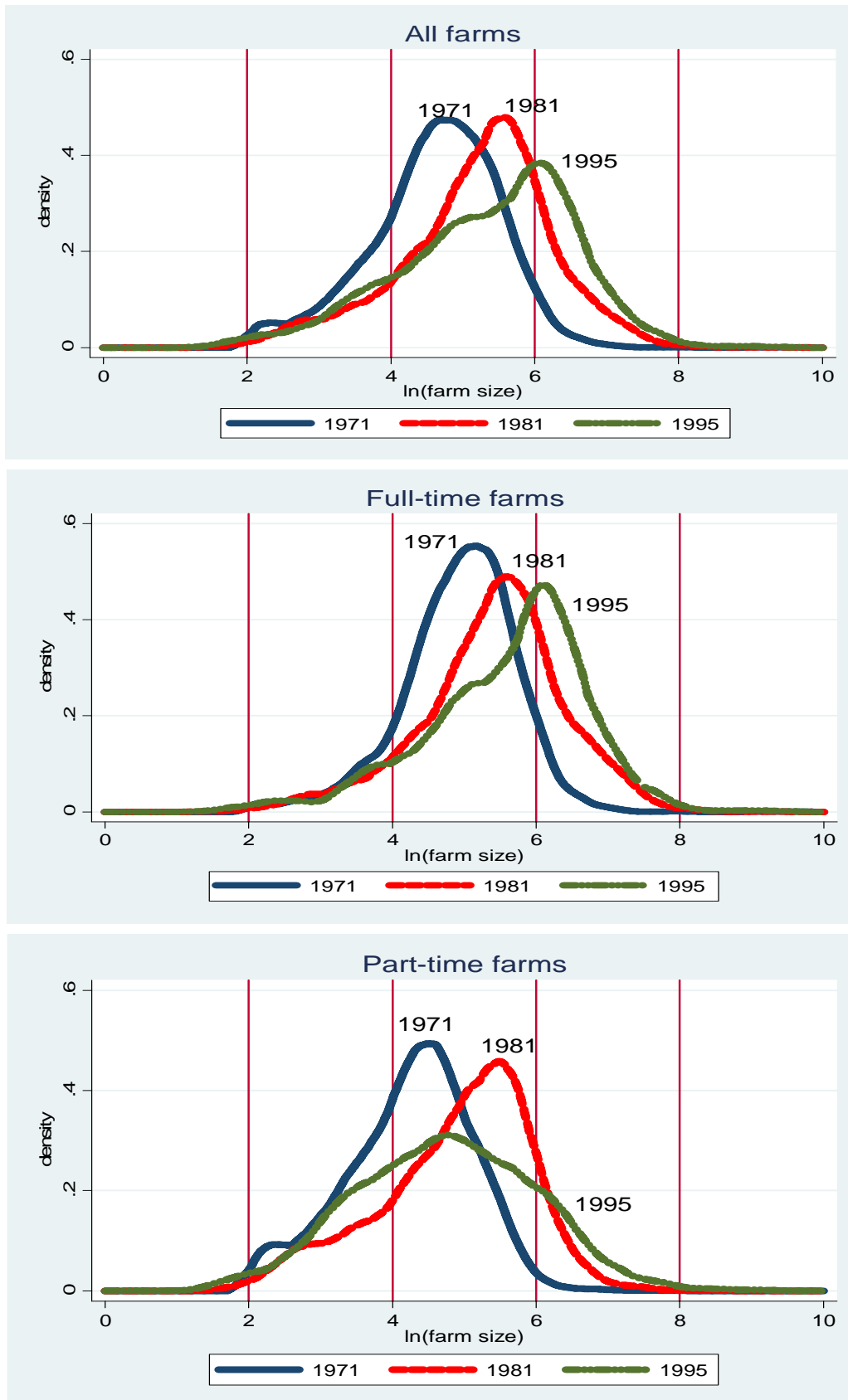


Figure 1. Changes in the farm size distribution by full-time/part-time status

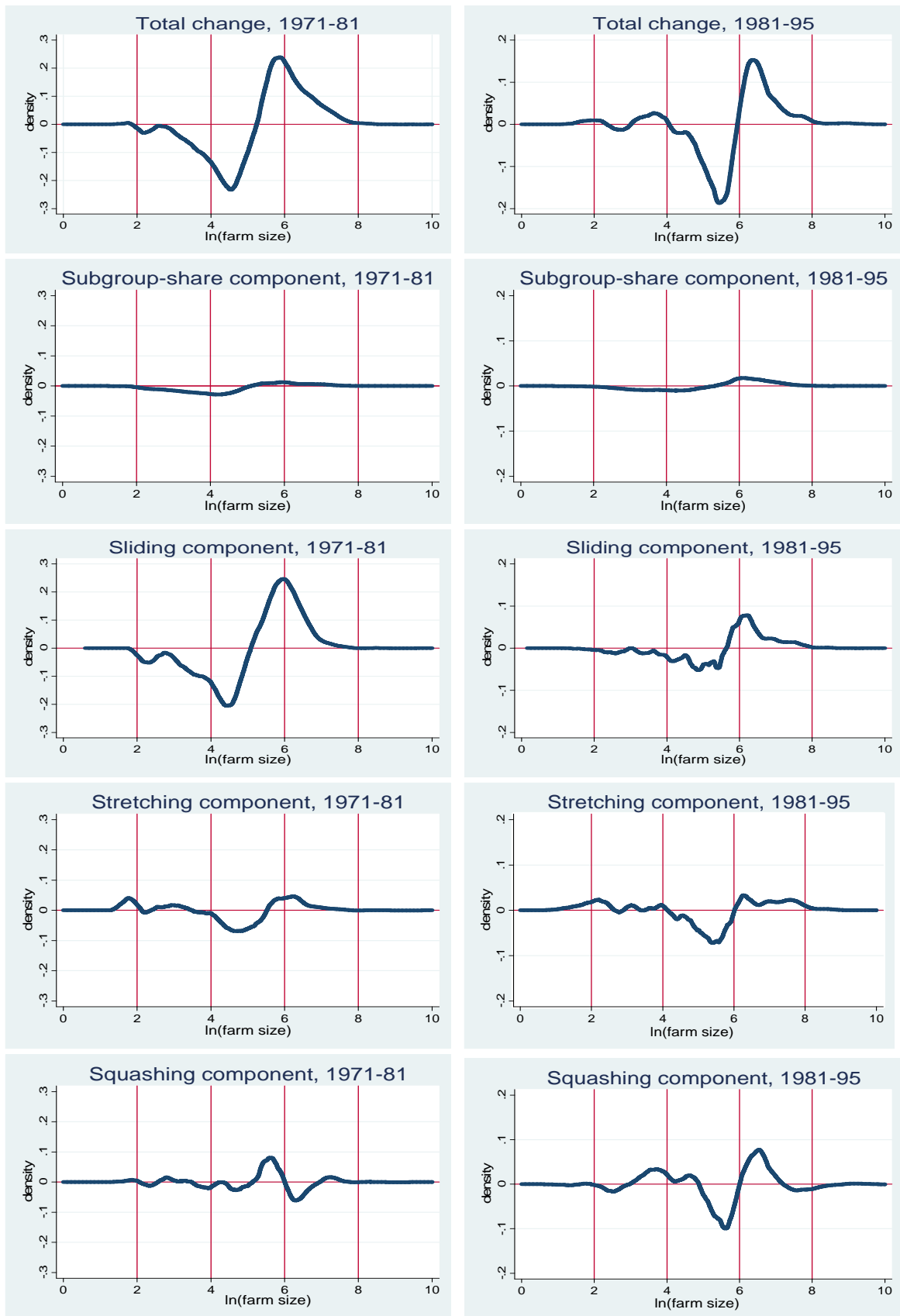


Figure 2. Decomposition of changes in farm size distribution, 1971-81 and 1981-95

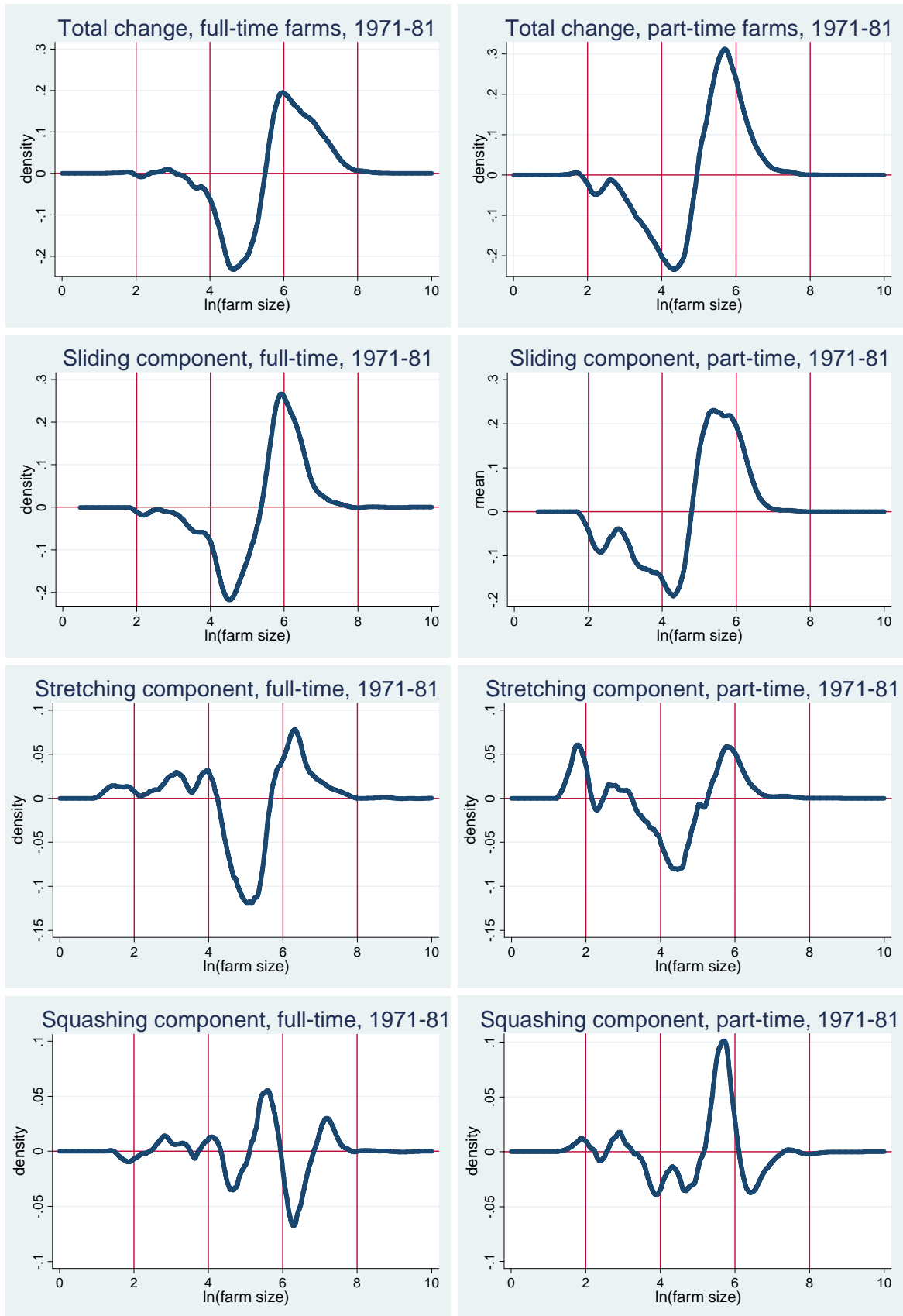


Figure 3. Decomposition of changes in farm size distribution, 1971-81

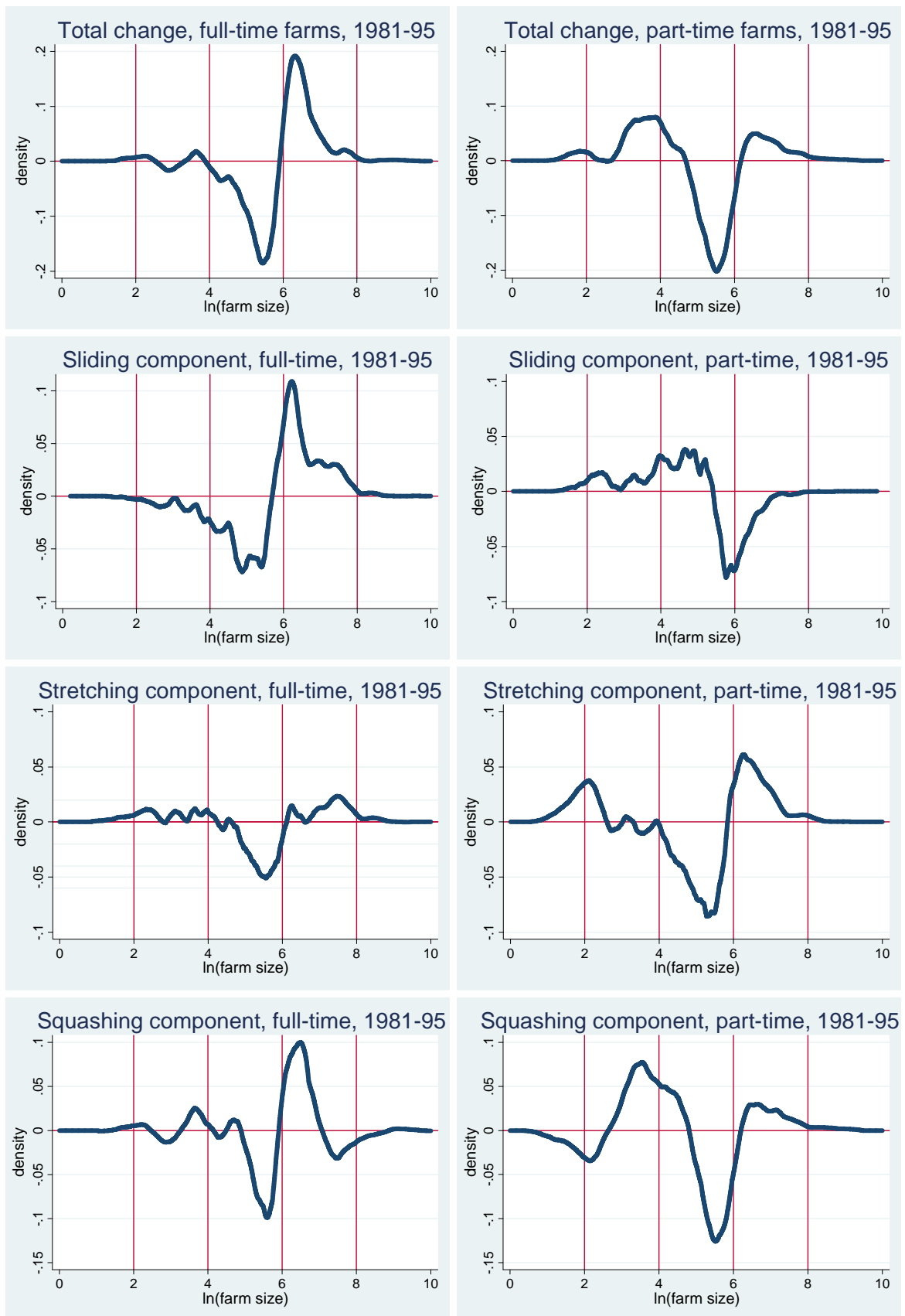


Figure 4. Decomposition of changes in farm size distribution, 1981-95