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Productivity and the Growth of Japanese Agriculture in the 1930s: A Panel Data Analysis Using a Survey of the Farm Household Economy

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Productivity and the Growth of Japanese Agriculture in the 1930s:

A Panel Data Analysis Using a Survey of the Farm Household Economy[*]

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Abstract

In the early 1930s, the Showa Depression, which commenced in 1930 following the onset of the Great Depression in 1929, had substantial effects on urban and rural economies in Japan, and agricultural production stagnated in the 1930s. Many studies have analyzed Japanese agriculture by using production functions. However, there is variance among them. Additionally, many studies were based on the assumption of constant returns to scale (CRS) because of data limitations. Utilizing a detailed micro-level database, we re-examined agricultural production in this period. The results show that the values of the production elasticity of factors scored near the lowest of those shown by previous studies and that CRS is not supported. The results also show that the trend of the change in total factor productivity is in line with that of previous studies.

Key words

Farm household survey, Panel data, Production function, Total factor productivity, Prewar Japan

JEL codes

D22, N55, O13, Q12

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1. Introduction

In the period between the end of World War I and the beginning of World War II (hereafter, the interwar period), Japan experienced the Showa Depression, which commenced in 1930 following the onset of the Great Depression in 1929, and it had a large impact not only on the urban economy but also on the farm household economy. In 1931, when cheaper rice produced in colonies had begun to be imported and the weather conditions had led to a good harvest in the previous year, rice prices and farm income per labor hour were at their lowest levels (Fujie and Senda 2011, Sakane 2010). After the onset of the Depression, bad weather conditions caused poor harvests. In 1934, cold weather brought on a serious famine in Tohoku and Hokuriku region, and many people suffered from starvation, especially in mountainous areas (Hiraga 2003). However, with some fluctuation, Japan's farm household economy gradually recovered until Japan went on a war footing.

In this period, the structure of Japanese agriculture changed substantially. The existing landlord–tenant farmer system had been destabilized, and tenancy disputes arose frequently. The government and rural communities attempted to settle this instability using collective actions in the villages (Kojima 2008, Sakane 2011). Kurihara (1948) has argued that as part of this process, landed-tenant farmers with medium land size emerged. The size of land they managed was about two *cho*,[1] which was larger than the average managed land size at the time (just under one *cho*). Associated with these changes, the number of large (previously the landowners) and small farmers relatively decreased.[2] These changes, referred to as the *Chuno Hyojunka Ron* [Convergence to medium-scale farmers], have been widely discussed by Japanese economic and agricultural historians (Fujie and Senda 2011).

Previous studies analyzed agriculture and sources of its growth in the period before World War II

[1] One *cho* is approximately 9,917 m^2.

[2] Kurihara (1948) defined medium-scale farmers as farmers with an operational land size between one and two *cho*. Small- and large-scale farmers were defined as farmers with operational land sizes smaller than one *cho* and larger than two *cho*, respectively.

(hereafter, the prewar period). In particular, many production function analyses, such as Kamiya (1941) and Akino and Hayami (1974), were conducted to identify the extent of the contribution of production factors. Hayami (1973) argued that agricultural production had stagnated mainly because of decreased technological innovation in this period. Hayashi and Prescott (2008), who used a two-sector growth model, argued that one of the main reasons the Japanese economy had decreased growth in the prewar period is the patriarchic family system, which affected rural agricultural households. However, many of these studies assumed constant returns to scale (hereafter, CRS) and are based on macro-level data.

In this paper, to make some new contributions to this body of literature, we conduct a production function analysis with the standard formulation, utilizing the results of the survey conducted by the Ministry of Agriculture and Forestry (hereafter, the MAF) as a micro-level panel database. We also apply a dynamic panel model to estimate the production function to treat the various shocks that occurred in the 1930s, and we calculate the total-factor productivity (TFP) and its change from the results of the production function. Estimating the production function and calculating TFP with micro-level data has an advantage in that it enables us to mitigate problems such as multicollinearity and to treat heterogeneity among samples, and the obtained results can serve as a basis for further studies, such analyses of labor allocation.

The composition of this paper is as follows. First, we briefly review the literature on agricultural production function analysis in the pre-/interwar period of Japan in Section 2. We explain the characteristics of the dataset in Section 3 and the models and methods in Section 4. The results are given in Section 5, and Section 6 concludes the paper.

2. Literature on Agricultural Production Function Analysis in Pre-/Interwar Japan

The many previous studies about Japanese agriculture in this period include analyses of the agricultural production function. In the primary study, Kamiya (1941) estimated the

Cobb–Douglas-type agricultural production function.[3] In Kamiya's study, land was introduced as an independent explanatory variable since land is one of the most important factors in agricultural production. Using production cost survey data on rice farmers from 1937 to 1939, Ohkawa (1945) showed that the aggregated value of production elasticity of factors was close to one, that differences in productivity among large scale farmers and small scale farmers were relatively small, and that the value of the production elasticity of land was 0.4–0.5, which was consistent with the land rent rate.

Hayami (1973) and Akino and Hayami (1974) estimated the prefectural-level cross-section production function and sought to determine the factors of agricultural growth. They concluded that education and research/diffusion activities for agriculture contributed to some part of growth beyond that explained by conventional inputs such as land, labor, capital, and other input goods, and that agricultural growth stagnated because of low investment in research/diffusion activities in the interwar period. Their results also showed that the value of the elasticity of labor was around 0.4, which was higher than those of other inputs (around 0.15–0.3).

Shintani (1983) used data from the farm survey conducted by the MAF in 1934 to 1936 and showed that the values of the elasticity of labor, land, and capital were around 0.45, 0.4, and 0.15, respectively. The results in Shintani (1983) differ from those of Hayami (1973) and Akino and Hayami (1974), and Shintani maintained that the low values of the elasticity of land in their work was due to the data utilized to obtain their results (Shintani 1983). Minami (1981) also estimated the agricultural production function in the interwar period and showed that the value of the elasticity of land gradually decreased from around 0.8 in the early 1920s to around 0.6 in the 1930s.

As we have seen, many studies have been conducted on the agricultural production function in the pre-/interwar period. With respect to the values of the elasticity of inputs, the results differ even among those studies limited to the 1930s: land ranges from 0.2 to 0.6, labor from 0.1 to 0.5, and capital from 0.1 to 0.2. However, almost all of these previous studies assumed that the aggregated

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[3] According to Heady and Dillon (1961), Kamiya (1941) conducted the first estimation of a Cobb–Douglas-type production function of farms, not only in Japan but across the world.

value of the production elasticities of land, labor, and capital equaled one.[4] In other words, they assumed CRS. One of the reasons for this assumption is multicollinearity. Akino (1972), which was the base study of Hayami (1973) and Akino and Hayami (1974), stated that the imposition of CRS was finally adopted to cope with the multicollinearity, although the alternative assumption of not imposing CRS had been considered.

Data availability could be another reason for the assumption of CRS. For example, Hayami (1973) and Akino and Hayami (1974) used prefectural-level datasets, and therefore their sample sizes are relatively small. Shintani (1983) used a dataset extracted from the same survey results as those used in this paper; this survey is more accurate than the other one. At that time, however, it was not possible to use these survey results for panel data analysis.

After the studies mentioned above, production function analysis of Japanese agriculture shifted its target to the postwar period in terms of economic development. Methodologically, some studies adopted more flexible function forms, such as trans-log, while others did general equilibrium growth accounting. However, their target was the postwar period.[5] Thus, there seems have been no substantial progress in the production function analysis of pre-/interwar-period Japanese agriculture since Shintani (1983).

3. The MAF Survey Data[6]

In Japan, data on the farm household economy have been collected since the 1890s. Surveys of the farm household economy in modern Japan has their origin in the surveys conducted by Mankichi Saito,

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[4] This is in the case of the value-added term. In case of the output term, "other input goods," which includes, for example, fertilizer, is used in addition to these inputs.

[5] For the latter, Yamaguchi (1982) conducted production function analysis of both pre- and postwar period Japanese agriculture.

[6] This section is based on our previous paper (Kusadokoro, Maru, and Takashima 2012), where we used the same dataset for another research topic.

an engineer at the Ministry of Agriculture and Commerce (the precursor to the MAF) at an agricultural experiment station during the Meiji and Taisho periods. These surveys were subsequently taken over by the *Teikoku Nokai* [Imperial Agricultural Association] and then the MAF (these are known as the *Teikoku Nokai* survey and the MAF survey, respectively). In addition to the Saito surveys, a series of other surveys collected individual records using single-entry bookkeeping methods designed for the farm household economy, and the bookkeeping design and sampling method were revised several times during the pre-/interwar period. Following Inaba (1953), we categorize these surveys into the following stages after considering their procedures and contents.[7]

- The Mankichi Saito survey, 1909–1920
- The *Teikoku Nokai* survey, 1913–1915
- The MAF survey (first period), 1921–1923
- The MAF survey (second period), 1924–1930
- The MAF survey (third period), 1931–1941
- The MAF survey (fourth period), 1942–1948

In this analysis, we use individual data extracted from the third period of the MAF survey (hereafter, the third-period MAF survey). Our data covers 16 prefectures (Akita, Fukushima, Ibaraki, Tokyo, Niigata, Toyama, Yamanashi, Nagano, Shizuoka, Aichi, Osaka, Shimane, Hiroshima, Tokushima, Fukuoka, and Miyazaki) over 11 years. The first eight prefectures are in the east of Japan, while the other eight are in the west. In principle, the survey included six or nine households in each prefecture every year. The number of farm households that had completed bookkeeping as collected by the MAF averaged 86 percent throughout the survey period (Ozeki 2009).

Here, we note that there was a sort of arbitrariness in the selection of the farmers surveyed since the surveys in the pre-/interwar period did not employ random sampling methods. Selected farmers tended to manage larger amounts of land since these farmers were generally well educated and were

[7] See Ozeki (2009, p. 124) and the table in the appendix in Inaba (1953).

able to understand the bookkeeping system. In the first year of the third-period MAF survey, the selection criteria for the surveyed farmers were changed considerably from the previous survey, mainly because of this bias toward farmers with larger farms (Inaba 1953, Ozeki and Sato 2008). However, there may still be a sampling bias in terms of farm size. Hereafter, we briefly check the basic characteristics of the dataset.

Table 1 details the top 12 patterns in the panel data sequence of the sample households. The most frequent pattern was households surveyed in only one year, 1941. These patterns suggest frequent changes in the surveyed households, even though the MAF survey was in principle a continuous program of research. Only five households in our sample were surveyed every year in the survey period. There is therefore the possibility of an estimation bias caused by attrition in the panel data. Accordingly, we must interpret the results obtained in this analysis with care.

As described, the third-period MAF survey attempted to collect information from farmers with medium-sized farms. However, the survey requirement that the farmers should do the bookkeeping themselves has left a degree of upward bias in the survey because a certain level of education was necessary. Panel A in Table 2 compares the distributions of operational land size of the farmers sampled in the third-period MAF survey and the *Noji Tokei Hyo* (hereafter, the *Noji* data), the national-level statistics. We can see that more than two-thirds of the farmers sampled in the third-period MAF survey have one to three *cho* of operational land. The comparable share is only about 28 percent in the *Noji* data. Furthermore, the share of small-scale farmers (those with less than 0.5 *cho* of agricultural land) in the sample data is also obviously smaller than that in the *Noji* data. Indeed, according to the estimation in Umemura et al. (1966), the average operational land size in 1930s Japan was only about 0.94 *cho*, but it is about 1.31 *cho* in our sample data.

Panel B in Table 2 compares the distribution of land ownership by categorizing farmers into three groups on the basis of their land ownership: farmers who owned not less than 80 percent of their operational land (landed farmers), those who borrowed not less than 80 percent of their operational land (tenant farmers), and others (landed-tenant farmers). Both the sample data and the *Noji* data display similar patterns, although the sample data have a slightly lower proportion of landed-tenant farmers and a higher proportion of tenant farmers. Thus, although the third-period MAF survey retains some upward bias toward large-scale farmers, the survey appropriately collects data from tenant farmers. This is the main advantage of the survey over other representative surveys of Japan's farm household economy, such as the agricultural management survey conducted by *Teikoku Nokai* (Senda and Kusadokoro 2009).

4. Model Specification

4.1 Production Function

In this study, we estimate production functions using the dataset from the third-period MAF survey to understand agricultural production and its change in the 1930s. As mentioned above, previous studies used different datasets and different variables. This may have led to the varying results among previous studies. In Hayami (1973) and Akino and Hayami (1974), a prefectural-level dataset was used, and the variables in their analyses were selected according to the standard of the output term. Shintani (1983) used the dataset from the third-period MAF survey, which is the source of our data. However, Shintani (1983) limited the samples to landed farmers for 1934 to 1936, and the variables were selected according to the standard of the value-added term. In this study, we select variables according to the standard of the output term to determine the effects of other miscellaneous inputs and adopt a Cobb–Douglas-type functional form to match previous studies. The descriptive statistics of the variables used in this study are shown in Table 3. All the price values are deflated by the agricultural products price index.

First, some static models (the pooled OLS model, the random effect model, and the fixed effect model) are estimated to compare the results of previous studies. Then, we estimate a Blundell and Bond (2000)-type dynamic panel data model to consider the effect of shocks caused by the Showa depression and other events.

Here, we explain the Blundell and Bond (2000)-type production function estimation procedure. The basic specification of the production function is

$$Y_{it} = \mathrm{A}_{it} L_{it}^{\beta_L} H_{it}^{\beta_H} C_{it}^{\beta_C} G_{it}^{\beta_G}, \quad i = 1, \dots, N; t = 1, \dots, T, \tag{1}$$

where Y_{it} is the output value of farm i in year t, L_{it} is the size of the managed land, H_{it} is the hours of labor, C_{it} is the value of capital stock, and G_{it} is the value of the other miscellaneous inputs. Here, taking the logarithm of both sides of equation (1), we obtain

$$y_{it} = \beta_l l_{it} + \beta_h h_{it} + \beta_c c_{it} + \beta_g g_{it} + u_{it} \ , \qquad (2)$$

where y_{it}, l_{it}, h_{it}, c_{it}, and g_{it} are the logs of Y_{it}, L_{it}, H_{it}, C_{it}, and G_{it}, respectively. u_{it}, the log of A_{it}, is the error term, and is specified as follows:

$$u_{it} = \gamma_t + (\eta_i + v_{it} + m_{it}) \ ,$$

$$v_{it} = \rho v_{i,t-1} + e_{it}, \ |\rho| < 1 \ ,$$

$$e_{it}, m_{it} \sim MA(0) \ ,$$

where γ_t is a year-specific intercept, η_i is an unobserved farm-specific effect, v_{it} is a possibly autoregressive shock, e_{it} is a productivity shock, and m_{it} is a serially uncorrelated measurement error.

To estimate the parameters of the restricted model (2), either a dynamic (common factor) representation of (2),

$$\begin{aligned} y_{it} = {} & \beta_l l_{it} - \rho\beta_l l_{i,t-1} + \beta_h h_{it} - \rho\beta_h h_{i,t-1} + \beta_c c_{it} - \rho\beta_c c_{i,t-1} \\ & + \beta_g g_{it} - \rho\beta_g g_{i,t-1} + \rho y_{it-1} + (\gamma_t - \rho\gamma_{t-1}) \\ & + \{\eta_i(1-\rho) + e_{it} + m_{it} - \rho m_{it-1}\} \ , \end{aligned}$$

is adopted, or an unrestricted model,

$$\begin{aligned} y_{it} = {} & \pi_1 l_{it} + \pi_2 l_{i,t-1} + \pi_3 h_{it} + \pi_4 h_{i,t-1} + \pi_5 c_{it} + \pi_6 c_{i,t-1} + \pi_7 g_{it} \\ & + \pi_8 g_{i,t-1} + \pi_9 y_{it-1} + \gamma_t^* + (\eta_i^* + w_{it}) \ , \end{aligned}$$

subject to four non-linear common factor restrictions: $\pi_2 = -\pi_1\pi_9$, $\pi_4 = -\pi_3\pi_9$, $\pi_6 = -\pi_5\pi_9$, and $\pi_8 = -\pi_7\pi_9$, where $\gamma_t^* = \gamma_t - \rho\gamma_{t-1}$ and $\eta_i^* = \eta_i(1-\rho)$. Given consistent estimates of the unrestricted parameter vector $\pi = (\pi_1, \ \pi_2, \ \pi_3, \ \pi_4, \ \pi_5, \ \pi_6, \ \pi_7, \ \pi_8, \pi_9)$ and its $var(\pi)$, these restrictions can be imposed and tested by the minimum distance method to obtain the parameter vector in the restricted model. If there are no measurement errors, w_{it} becomes e_{it}; otherwise, $e_{it} + m_{it} - \rho m_{it-1}$.

4.2 Calculation of Total Factor Productivity

After determining the technological parameters of the restricted model, following Mary (2013), farm-level productivity is calculated to obtain the aggregated TFP and its change rate. The farm-level productivity p_{it} is calculated as follows:

$$\hat{p}_{it} = \exp(y_{it} - \hat{\beta}_l l_{it} - \hat{\beta}_h h_{it} - \hat{\beta}_c c_{it} - \hat{\beta}_g g_{it}).$$

Then, the aggregated TFP for each year is calculated by using the output share of each farm household as weights. Here, we note that this $\hat{p}_{it}$ contains error terms since this is a calculated value for each sample. However, this aggregated TFP, the average value of $\hat{p}_{it}$, can mitigate the effect of the error terms to some extent.[8]

5. Results

5.1 Static Models

First, static models (the pooled OLS model, the random effect model, and the fixed effect model) are estimated for comparison with the previous studies. Table 4 shows the results obtained. The coefficients of land fall within around 0.12 to 0.40, labor 0.22 to 0.25, capital 0.05 to 0.10, and miscellaneous inputs 0.23 to 0.28. All coefficients in all models are significant at the one percent level. The aggregated value of the coefficients of these inputs is around 0.96 in the pooled OLS model, 0.91 in the random effect model, and 0.69 in the fixed effect model. The low score in the fixed effect model is mainly due to the low score of the coefficient of land. Tests among the models support the adoption of the random effect model, and CRS is rejected at the five percent level in all models.

Here, the coefficient of land in the fixed effect model is substantially low. Hoch (1962), who used a dataset of Minnesotan farms, obtained results that support decreasing returns to scale in the production function by estimating a panel fixed effect model. Hoch (1962) interpreted this to be the result of limitations in production caused by farms' management ability. Following Hoch (1962), the

[8] In addition, by using $\hat{p}_{it}$, we can calculate the average TFP for each household. This is an advantage of using micro-level data.

low value of the coefficient of land in this paper can be interpreted as being the result of drawing a distinction between farms' management ability and the coefficient of land. Indeed, there are relatively strong positive correlations between land size and the fixed effect, and thus, the large land size of some farms can be interpreted as the result of high farm management ability. However, the fixed effect may mis-absorb some portion of the value of land elasticity in the fixed effect model, where variance in land size is small. We must keep in mind that it might lead to the low value bias of land elasticity in the fixed effect model and that this might affect the result of the Hausman test.

Next, we estimate the static models, imposing CRS, to determine how elasticity values differ from the estimation results without imposing CRS. In this case, the dependent variable is output per unit of land and independent variables are labor per unit of land, capital per unit of land and miscellaneous goods per unit of land. The elasticity value of land is calculated after the estimation. Table 5 shows the results. The coefficients of land fall within around 0.28 to 0.42, labor 0.25 to 0.29, capital 0.06 to 0.18, and other miscellaneous inputs 0.25 to 0.28. All coefficients are significant at the one percent level in all models, and tests among the models support the adoption of the fixed effect model.

The value of miscellaneous input goods is stable in both the imposing-CRS models and not-imposing-CRS models. In contrast, the value of capital varies in the imposing-CRS models. In the imposing-CRS models, the elasticity values are unevenly increased or boosted compared to the estimation results without assuming CRS, especially in the fixed effect model. These uneven changes might be caused by forcing the CRS assumption.

Here, we summarize the results of the static models. Without the CRS assumption, the random effect model is adopted and CRS is rejected. The coefficient of land is low in the fixed effect model. On the other hand, with the assumption of CRS, elasticity values are boosted unevenly, and each elasticity value shows a relatively large difference among models. From these results, we can say that imposing CRS may cause bias. However, the fixed effect here also may cause bias. Therefore, both specifications might have biases. This might be caused by the application of static models even though there were changes in the production circumstances given the Depression and other factors. Thus, we proceed to the estimation of the Blundell and Bond (2000)-type dynamic model.

5.2 Dynamic Model

The Blundell and Bond (2000)-type dynamic model is estimated to overcome the weakness of the static models pointed out above.[9] Table 6 shows the results obtained. Here, for example, *t*-2 is the specification with two period-lagged instruments in the first-differenced equation and one period-lagged difference instruments in the level equation. In the unrestricted model expression, lagged land and lagged labor show different signs in the *t*-2 specification. In contrast, only the sign of lagged land becomes positive in the *t*-3 specification model with no significance. The Arellano–Bond test rejects first-order but not second-order serial correlation in both specifications. Additionally, Hansen's over-identification test shows a sufficiently large *p*-value. As a whole, there seems to be no serious specification problem, especially in the *t*-3 specification.

The coefficients of the explanatory variables in the restricted model are recovered by the minimum distance method. Concerning the common factor restrictions, the Comfac test does not reject the null hypothesis in either the *t*-2 or the *t*-3 specification. In the *t*-2 specification, however, the value of capital becomes almost zero with no significance, and the value of land also shows no significance. There are no negative signs, no values become almost zero, and only one variable shows insignificance in the *t*-3 specification. Considering that only the value of lagged land is not significantly positive and that the signs of all the other lagged variables are negative in the unrestricted model expression, we can conclude that the common factor restriction in the *t*-3 specification is valid.

In the *t*-3 specification, the values of land, labor, capital, and other miscellaneous input goods are around 0.26, 0.15, 0.06, and 0.24, respectively. The value of land recovers and gets close to the level of the random effect model. However, the simply aggregated value of coefficients remains just above 0.7, and the Wald test rejects the validity of CRS. This indicates that previous studies might have biases.

The coefficient of land in the *t*-3 specification in this paper exceeds the results of a series of

[9] Before estimation, Fisher-type tests are implemented for all variables. The results reject the existence of a unit root.

Hayami's works and is close to that of Akino (1972), with a specification similar to other studies (Table 7). As for the proportion of each input, the results of this paper rank in the middle of those of previous studies, indicating that no extreme labor or land-responsive technology was adopted in this period of Japanese agriculture, and are somewhat similar to that of Akino (1972) and Minami (1981). The differences in each coefficient value in previous studies might be caused by both the assumption of CRS and differences in the datasets, variables, and specifications employed among the studies. For example, the coefficients of land in Hayami (1973) and Akino and Hayami (1974) are very low compared to other studies. However, Akino (1972), whose results were used in Hayami (1973) and Akino and Hayami (1974), obtained somewhat similar results as ours from testing the same specification.[10] In addition, there is substantial variation in the coefficient of labor, which might be caused by the differences in variables. Some studies used the number of laborers, while others used the number of working days; this paper uses working hours.

5.3 Total Factor Productivity

Next, we calculate TFP and its change by using the results of the Blundell and Bond (2000)-type dynamic production function. Specifically, the coefficients of the restricted model with the *t*-3 specification are used for calculation. The results are shown in Table 8 and Figures 1 to 4.

The results in Table 8 show that the percent change of the three-year average value from 1932 to 1940 was around 1.00. The change in the first half of the period is 1.05 and that in the second half is 0.94. This means that there was a relatively higher recovery right after the Showa Depression and that the speed of growth declined before World War II, with somewhat high variation (also see Figures 1 and 2).[11]

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[10] Of course, the main objective of Hayami's work was to clarify the effects of research/diffusion and education. Therefore, it must be noted that those results cannot be compared directly with ours in terms of the purpose of the analysis.

[11] This variation might be caused by calculating TFP for each year. In contrast, the results of Hayashi and Prescott (2008) seemed to have more variation in the first half of this period. Their results might more sensitively

Figures 3 and 4 show a comparison of the five-year average TFP and its change with the findings of Hayami (1973) and Hayashi and Prescott (2008). The negative value for 1936 in Figure 4 is due to low TFP values in 1934 and 1938 caused by a bad harvest. On the whole, it can be said that the change rate of the five-year average TFP is in a decreasing trend. This trend might be caused by the period of the Second Sino–Japanese War prior to World War II. Hayami (1973) insisted that productivity growth in the interwar period was lower than that in other periods because of the exhaustion of production techniques that had been stocked by veteran farmers. More precisely, Hayami (1973) pointed out that there was a recovery of production in the first half of the 1930s and that agricultural production again began to be depressed by militarism in the second half. Calculation of the average change rate of the five-year average total productivity from the results in Hayami (1973) in this period generates values around 0.23 for 1931 to 1941, around 0.27 for 1931 to 1936, and around 0.18 for 1936 to 1941.[12]

Although our results cannot be directly compared with these studies because of differences in calculations, our results show a relatively similar trend and are between the results of Hayami (1973), which shows an average change rate of the five-year average TFP of 1.181, and that of Hayashi and Prescott (2008), which shows 0.414. The reason our results and those of previous studies with a macro dataset show similar trends is as follows: TFP in previous studies contains macro-level factors that are not included in the model specifications, such as technological change and changes in weather and economic circumstances. In contrast, TFP in our results contains micro-level factors such as technological change and changes in weather and economic circumstances in addition to the macro-level factors. If the changes in the micro-level factors do not differ substantially from the changes in the macro-level factors, TFP in our results may not differ from TFP in the previous studies.

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detect the effect of famine in 1934 than ours because of the differences in specifications, the estimation period, and the dataset. Nevertheless, our results shows a similar trend to that of Hayashi and Prescott (2008).

[12] It must be noted that these values of Hayami (1973) are calculated from the five-year average total productivity and that our value in Table 8 cannot be compared with that of Hayami (1973) since the period is different.

Since our results are obtained from panel data, we also can calculate the average TFP for each household. Figure 5 shows the relationship between average land size and this average TFP classified by farm type. The figure shows that productivity trends do not have large differences among farm types. Additionally, the TFP curve in the "Total" sub-panel shows a somewhat log-form relationship and reaches its peak at right after two *cho*. In terms of production technology, around two *cho* can be considered the most efficient size of land in terms of operation. By comprehensively interpreting from the point of view of agricultural production technology, this result and decreasing returns to scale provide supporting evidence for the convergence to medium-scale farmers in this period.

6. Conclusion

This paper re-examined agricultural production function in the 1930s, the period of recovery from the shock of the Showa Depression, by utilizing the dataset extracted from the third-period MAF survey. The results obtained are summarized as follows.

First, static models were estimated. In the pooled OLS and random effect models, the coefficients of the explanatory variables are similar to those of previous studies, and the aggregated values of these models score above 0.9 and do not largely contradict previous studies. In the fixed effect model, however, the value of elasticity of land is absorbed to some extent and the aggregated value decreases to less than 0.7. Furthermore, the test for the null hypothesis of CRS, which previous studies supported and assumed, is rejected, and estimations that impose CRS display unbalanced results. These results might be caused by applying static models even though there were changes in the production circumstances accompanying the Depression.

Next, a Blundell and Bond (2000)-type dynamic model was estimated. Although the simple aggregated value of the coefficients recovered to more than 0.7, the test for the null hypothesis of CRS was rejected again. The elasticity values in the dynamic model show moderate scores, indicating no land/labor-intensive/responsive technology in agricultural production in the 1930s.

By using the coefficients of dynamic model, TFP and its change were calculated. The results obtained in this paper indicate that the stagnation of productivity growth in this period was somewhat close to Hayashi and Prescott (2008), who adopted a two-sector growth model, and was not so different from the results of Hayami (1973) and Hayashi and Prescott (2008) in terms of the trend.

The results that CRS is rejected and that the most technologically efficient size of operational land becomes around two *cho* comprise supporting evidence for the convergence to medium-scale farmers. However, it should be noted that this is an analysis of the production side, and therefore, the issue cannot be fully determined without further analysis of the cost side. This remains for future research.

By applying dynamic panel data analysis and using a more detailed and informative micro-level dataset, this paper has shown, in some part, different results from those of previous studies, although the results are generally consistent with them. However, our results should be considered tentative since the database is still under construction and only 16 prefectures are currently used. As there are plans to improve the database with more individual data, there seems to be a great deal of room to improve our analysis, not only in terms of the estimation model but also the data.

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Tables

Table 1. Panel structure of the sampled farm households (top 12 patterns)

(Obs.: 1,079)

Frequency	Cumulative frequency (%)	1931	1932	1933	1934	1935	1936	1937	1938	1939	1940	1941
16	7.1	-	-	-	-	-	-	-	-	-	-	1
11	12.1	1	1	-	-	-	-	-	-	-	-	-
10	16.5	1	1	1	1	-	-	-	-	-	-	-
7	19.6	-	-	-	-	-	-	-	-	1	1	1
7	22.8	1	-	-	-	-	-	-	-	-	-	-
6	25.4	-	-	-	-	-	-	-	-	1	-	1
5	27.7	1	1	1	1	1	1	-	-	-	-	-
5	29.9	1	1	1	-	-	-	-	-	-	-	-
5	32.1	1	1	1	1	1	1	1	-	1	-	1
5	34.4	1	1	1	1	1	1	1	1	1	-	1
5	36.6	1	1	1	1	1	1	1	1	1	1	1
142	100.0	(other patterns)										
224												

Source: Sample data extracted from the Third-Period MAF Survey

Table 2. Distribution of operational land size and land ownership

Panel A: Distribution of operational land size

	Operational land size (x: *cho*)[3]						
	$x < 0.5$	$0.5 <= x < 1$	$1 <= x < 2$	$2 <= x < 3$	$3 <= x < 5$	$5 < = x$	Total
Sample data from the third-period MAF survey (No. of obs.)[1]	36	200	408	80	13	0	737
Ratio in the sampled data (%)	4.9	27.1	55.4	10.9	1.8	0.0	100
Ratio of the *Noji* data (%)[2]	34.4	35.4	23	5.5	1.5	0.2	100

Panel B: Distribution of land ownership

	Land ownership[4]			
	Tenant	Landed Tenant	Landed	Total
Sample data from the third-period MAF survey (No. of obs.)[1]	242	267	228	737
Ratio in the sampled data (%)	32.8	36.2	30.9	100
Ratio of the *Noji* data (%)[2]	26.3	43.4	30.3	100

Sources: Sample data from the Third-Period MAF Survey and the *Noji Tokei Hyo*

Notes:

1. The figures are the total number of observations in each category in the sampled data between 1931 and 1937.
2. Average values of the *Noji Tokei Hyo* data between 1931 and 1937.
3. One *cho* is approximately 9,917 m^2.
4. A landed farmer is a farmer owning land representing not less than 80 percent of the operational land; a tenant farmer is a farmer renting land not less than 80 percent of the operational land; all others are classified as landed-tenant farmers.

Table 3. Descriptive statistics of variables

(Obs.: 1079)

variables	Description (unit)	Mean	S.D.
Y_{it}	Value of farm output (1,000 Yen)	1.167	0.563
L_{it}	Size of cultivated land (*cho*)	1.311	0.576
H_{it}	Labor hours (1,000 hours)	5.271	2.191
C_{it}	Value of capital (1,000 Yen)	1.043	0.853
G_{it}	Value of miscellaneous input goods (1,000 Yen)	0.229	0.216

Source: The Third-Period MAF Survey.

Note: One *cho* is approximately 9,917 m^2.

Table 4. Production function estimates (pooled OLS, random effect, and fixed effect)

(Obs.: 1079)

	OLS			Random Effect			Fixed Effect		
l_t	0.401	(0.020)	***	0.299	(0.027)	***	0.118	(0.042)	***
h_t	0.221	(0.022)	***	0.250	(0.023)	***	0.239	(0.027)	***
c_t	0.051	(0.013)	***	0.098	(0.019)	***	0.097	(0.032)	***
g_t	0.282	(0.012)	***	0.261	(0.014)	***	0.234	(0.017)	***
Adj. R-squared	0.777								
Breusch–Pagan test				774.85					
(p-value)				0.000					
F-test							6.68		
(p-value)							0.000		
Hausman test						20.48			
(p-value)						0.116			
Elasticity of scale	0.955			0.908			0.688		
CRS (p-value)	0.012			0.001			0.000		

Notes:

1. Year dummies are included in both models. Standard errors are in parentheses.
2. ***, **, and * indicate significance at one, five, and 10 percent, respectively.
3. CRS is a test for constant returns to scale.

Table 5. Production function estimates (pooled OLS, random effect, and fixed effect) imposing CRS

(Obs.: 1079)

	OLS			Random Effect			Fixed Effect		
l_t	0.417			0.344			0.276		
h_t	0.246	(0.020)	***	0.278	(0.022)	***	0.289	(0.026)	***
c_t	0.058	(0.013)	***	0.115	(0.018)	***	0.184	(0.029)	***
g_t	0.279	(0.012)	***	0.263	(0.014)	***	0.251	(0.017)	***
Adj. R-squared	0.577								
Breusch–Pagan test				790.36					
(p-value)				0.000					
F-test							6.33		
(p-value)							0.000		
Hausman test							45.78		
(p-value)							0.000		

Notes:

1. Year dummies are included in both models. Standard errors are in parentheses.

2. ***, **, and * indicate significance at one, five, and 10 percent, respectively.

Table 6. Production function estimates (System GMM)

(Obs.: 729)

	t-2			t-3		
Unrestricted model						
l_t	0.104	(0.117)		0.216	(0.135)	
l_{t-1}	0.006	(0.065)		0.107	(0.148)	
h_t	0.280	(0.123)	**	0.137	(0.114)	
h_{t-1}	0.083	(0.108)		-0.023	(0.118)	
c_t	0.059	(0.118)		0.126	(0.128)	
c_{t-1}	-0.139	(0.120)		-0.187	(0.129)	
g_t	0.224	(0.065)	***	0.264	(0.070)	***
g_{t-1}	-0.020	(0.052)		-0.124	(0.067)	*
y_{t-1}	0.237	(0.080)	***	0.333	(0.124)	***
Arellano–Bond test (1) (*p*-value)	0.000			0.000		
Arellano–Bond test (2) (*p*-value)	0.752			0.160		
Hansen OID test (*p*-value)	0.348			0.357		
Dif. Hansen test (*p*-value)	0.560			0.465		
Restricted model						
β_l	0.101	(0.111)		0.261	(0.103)	**
β_h	0.362	(0.094)	***	0.154	(0.093)	*
β_c	0.006	(0.078)		0.058	(0.088)	
β_g	0.229	(0.061)	***	0.242	(0.056)	***
ρ	0.364	(0.057)	***	0.476	(0.078)	***
Comfac (*p*-value)	0.205			0.660		
CRS (*p*-value)	0.009			0.003		

Notes:

1. Year dummies are included in both models. Standard errors are in parentheses.
2. ***, **, and * indicate significance at one, five, and 10 percent, respectively.
3. Comfac is a minimum distance test for common factor restrictions. We use “md_ar1.ado” for this test.
4. CRS is a test for constant returns to scale.

Table 7. Comparison with previous studies

	sys GMM (t-3)			Akino and Hayami (1974), Hayami (1973)	Akino (1972)	Shintani (1983)	Minami (1981)	
		Convert to CRS (l, h, c, g)	Convert to CRS (l, h, c)	1930–1935	1930–1935	1934–1936	1931–1935	1936–1940
β_l	0.26	0.36	0.55	0.15	0.38	0.4	0.63	0.57
β_h	0.15	0.22	0.33	0.4	0.31	0.45	0.21	0.29
β_c	0.06	0.08	0.12	0.15	0.06	0.15	0.15	0.13
β_g	0.24	0.34		0.3	0.26			

Notes:

1. Values in “Convert to CRS” are the ratio of the elasticity value of each input to the total.
2. Akino and Hayami (1974), Hayami (1973), and Akino (1972) used prefectural-level data. The specification of Akino (1972) is R305-11.
3. Shintani (1983) used the same database as ours; however, the use of the data was limited. Shintani did not impose CRS at the time of estimation.
4. Minami (1981) used prefectural data.

Table 8. TFP and its change

	TFP (1931 = 100)	Change in TFP (%)	Three-year Avg. TFP (1933 = 100)	Change in three-year Avg. TFP (%)	Five-year Avg. TFP (1933 = 100)	Change in five-year Avg. TFP (%)
1931	100.000					
1932	104.984	4.984	100.000			
1933	110.950	5.683	101.083	1.083	100.000	
1934	103.420	-6.787	101.928	0.837	101.935	1.935
1935	107.655	4.095	101.690	-0.234	103.177	1.218
1936	110.198	2.362	104.256	2.524	101.361	-1.760
1937	111.529	1.208	102.271	-1.905	104.260	2.860
1938	101.381	-9.099	104.961	2.631	105.719	1.399
1939	118.699	17.082	106.168	1.149	105.295	-0.400
1940	115.340	-2.830	108.252	1.964		
1941	107.968	-6.391				
Period Avg. (overall)				0.996		0.864
Period Avg. (1932–1936)				1.048		
Period Avg. (1936–1940)				0.945		

Figures

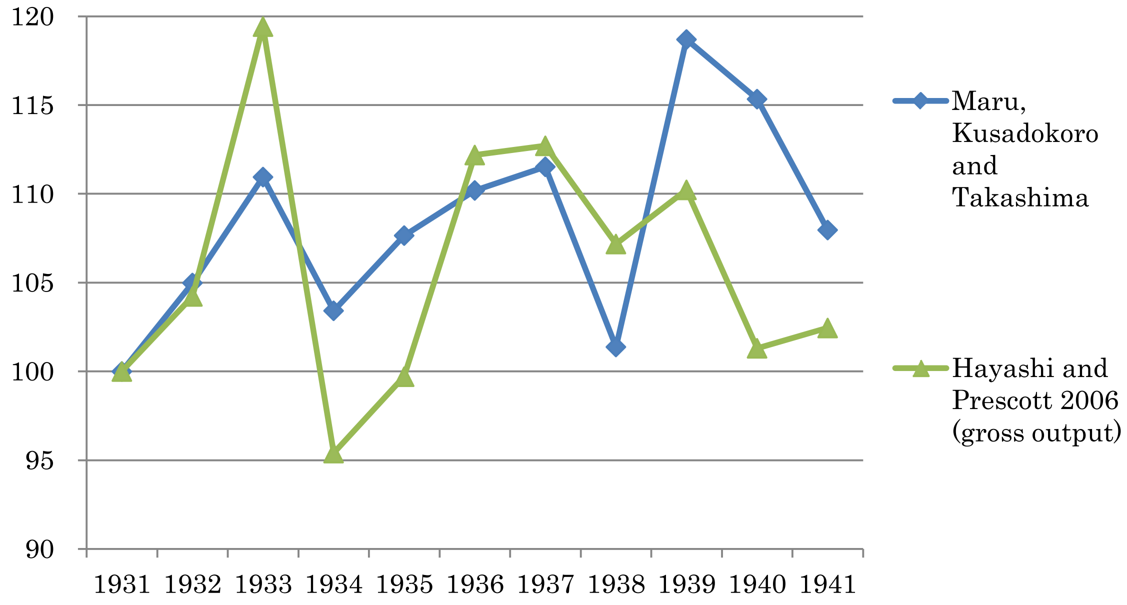

Figure 1. TFP (1931 = 100)

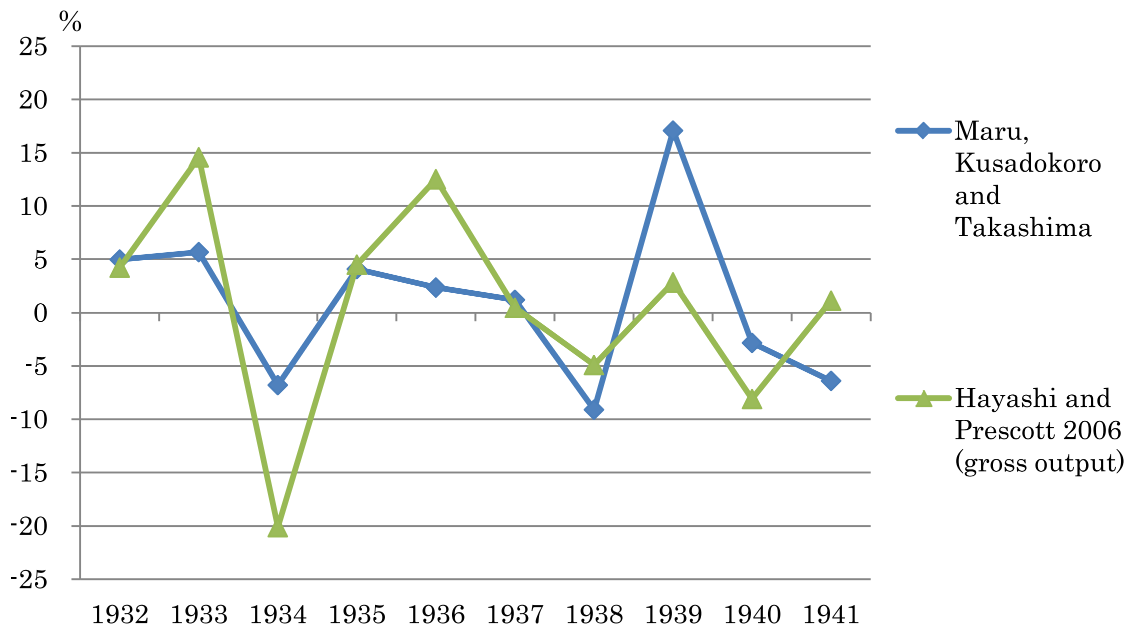

Figure 2. Percent change of TFP

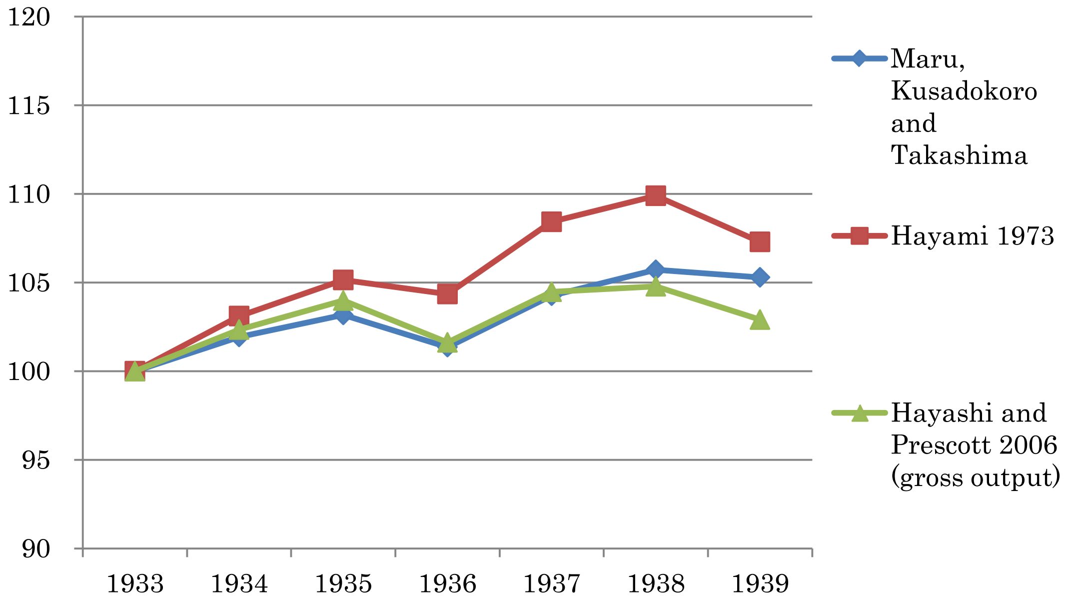

Figure 3. Five-year average TFP (1933 = 100)

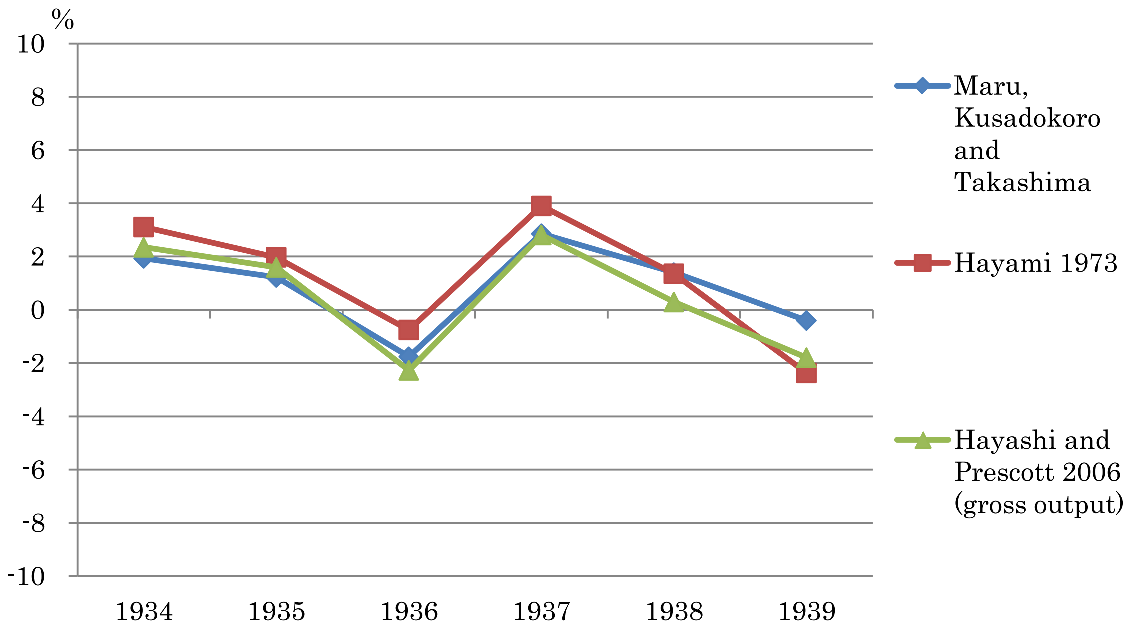

Figure 4. Percent change of the five-year average TFP

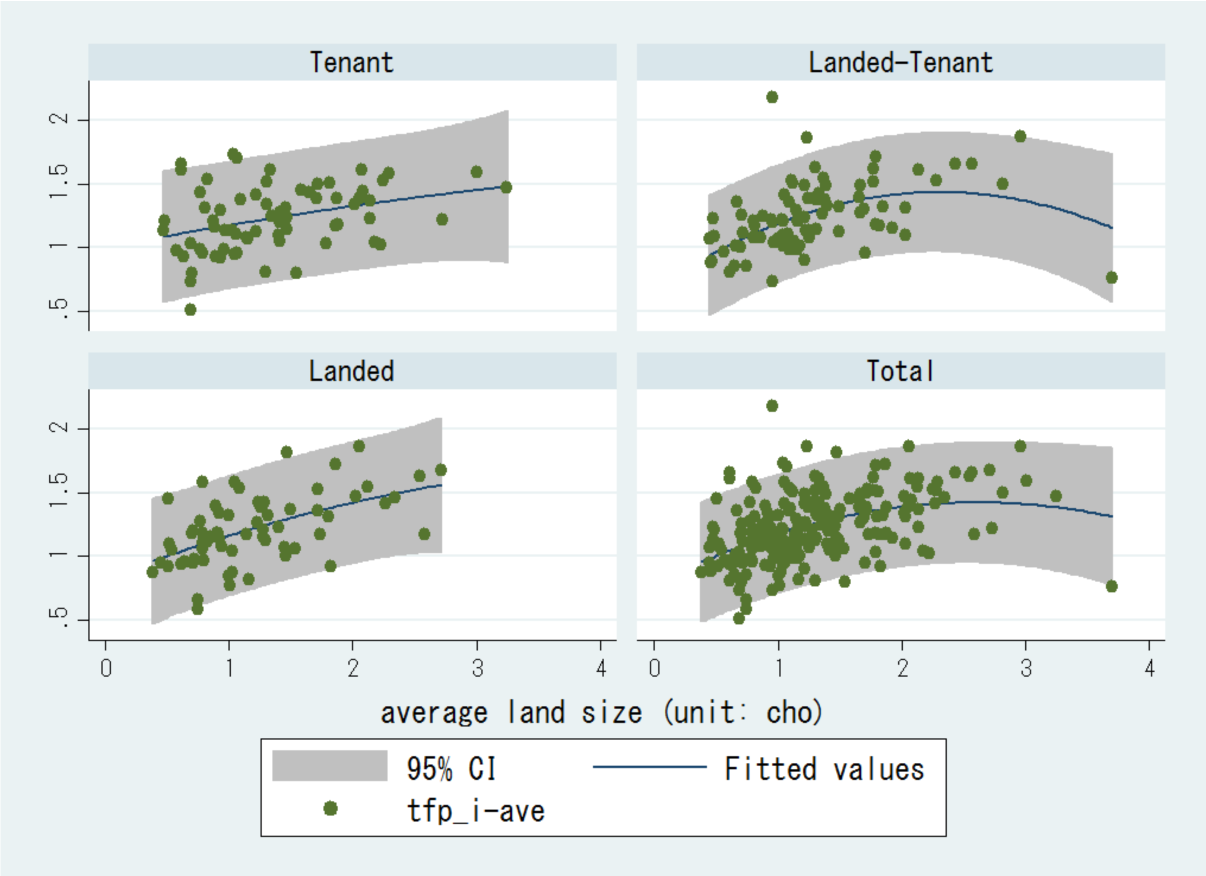

Figure 5. Average land size and average TFP of each household (tfp_i-ave) by farm type