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IFPRI Discussion Paper 01525

April 2016

Structural Transformation and Intertemporal Evolution of Real Wages, Machine Use, and Farm Size–Productivity Relationships in Vietnam

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ABSTRACT

This paper explores the evolution of real agricultural wages, machinery use, and the relationship between farm size and productivity in Vietnam during its dramatic structural transformation over the course of the 1990s and 2000s. Using six rounds of nationally representative household survey data, we find strong evidence that the inverse relationship between rice productivity and planting area attenuated significantly over this period and that the attenuation was most pronounced in areas with higher real wages. This pattern is also associated with sharp increases in machinery use, indicating a scale-biased substitution effect between machinery and labor. The results suggest that rural-factor market failures are receding in importance, making land concentration less of a cause of concern for aggregate food production.

Keywords: farm size–productivity relationship; structural transformation; Vietnam

ACKNOWLEDGMENTS

This paper was undertaken as a part of, and partially funded by, Bill & Melinda Gates Foundation (BMGF) and the CGIAR Research Program on Policies, Institutions, and Markets (PIM) led by IFPRI. This paper has not gone through IFPRI's standard peer-review process. The opinions expressed here belong to the author, and do not necessarily reflect those of BMGF, PIM, IFPRI, or CGIAR.

1. INTRODUCTION

The structural transformation of low-income agrarian economies is both cause and consequence of the steady amelioration of rural-factor market imperfections that have led to intersectoral and interhousehold heterogeneity in shadow prices of land, labor, and capital (Timmer 1988, 2007). Urbanization, the rise of modern industrial and service sectors, and a declining share of agriculture in gross domestic product and employment go hand in hand with economic growth and diminution of market failures that obstruct factor price equalization among production units. In low-income agrarian economies, household-specific market failures appear commonplace (de Janvry, Fafchamps, and Sadoulet 1989). When multiple markets fail, factor (shadow) prices vary across units, leading to heterogeneous input application rates and partial productivity measures, the most common of which is the oft-observed inverse relationship between farm size and crop yields (Feder 1985). The existence of any relationship between farm productivity and farm size—negative or positive—has attracted much scholarly and policy maker attention, as perhaps indicative of key market failures that motivate agricultural and rural development policy interventions. Examples of such interventions include progressive land reform and smallholder agricultural credit subsidies that might generate both efficiency and equity gains. If economic transformation were to bring improved factor market participation and performance, this might shift the farm size–productivity relationship and the economic rationale for concerted intervention to address rural market failures.

In the past twenty-five years, Vietnam has undergone some of the most rapid transformation of any low-income agrarian nation. During that time, the country experienced rapid real gross domestic product growth of 4–8 percent annually, mainly driven by the development of nonfarm sectors and the activation of markets for credit, labor, land, machinery, and other factors of production throughout the country (McCaig and Pavcnik 2013). As the nonfarm sectors provided more job opportunities, people moved out of farming, resulting in increased agricultural real wages and shrinkage in the wage gap between agricultural and nonagricultural sectors, as predicted by dual economy models (Lewis 1954; Gollin 2014).

Such changes likely have important implications for agriculture. It has long been observed that, in developing-country agriculture, smaller farms are typically more productive per unit area cultivated than larger ones (Chayanov 1926/1986; Sen 1962; Berry and Cline 1979; Carter 1984; Barrett 1996; Benjamin and Brandt 2002; Barrett, Bellemare, and Hou 2010; Carletto, Savastano, and Zezza 2013). The evidence of such an inverse farm size–productivity relationship has often justified land policies supporting small landholders and deterring farm size expansion, as well as agricultural credit policies to promote smallholder access to commercial inputs.

However, as a low-income agrarian economy undergoes rapid structural transformation, do factor markets for agricultural labor and machinery become more active, driving up real wages and attenuating the inverse relationship? Conceptually, Otsuka (2013) suggested that increasing real wages would reduce demand for agricultural labor, promote the use of machinery as a substitute for labor, and decrease the disadvantage of larger farms—perhaps even flipping the inverse relationship to a direct relationship—due to scale economies in machine use. Similarly, using nationally representative household data from India, Foster and Rosenzweig (2011) estimated an increase in optimal farm size due to the substitution of machinery for labor.

This paper explores the evolution of real agricultural wages, machinery use, and the relationship between farm size and productivity in Vietnam during its dramatic structural transformation over the course of the 1990s and 2000s. This inverse relationship has been attributed to imperfections in multiple factor markets, especially for land, labor, and credit (Sen 1966; Feder 1985). Even without significant changes in land policy, if the inverse relation had been partially driven by imperfect credit or labor markets, improved factor market functioning through structural transformation, as manifested in increased real wage rates and more active machinery rental markets, may have lessened or may even have reversed the long-standing inverse farm size–productivity relationship. Because the inverse relationship has long served as a powerful metaphor for pervasive rural-factor market failures that motivate government interventions, any such evolution has powerful implications for policy.

This paper uses six rounds of nationally representative household survey data from Vietnam, from 1992 through 2008, to examine whether and the extent to which increasing real wages and increased machinery rentals are associated with change in the inverse relationship between rice productivity and cultivated area.[1] To our knowledge, this study is the first to look at intertemporal change in the inverse farm size–productivity relationship over such a long period. There is strong evidence that the inverse relationship between rice productivity and planting area attenuated significantly over this period. Those effects are most pronounced in areas with higher real wages, which are also associated with sharp increases in machinery use, indicating a scale-biased substitution effect between machinery and labor. The empirical results of this study suggest that, in Vietnam, larger farmers' disadvantage in land productivity, compared to smaller farmers, has lessened sharply, which is probably attributable to improved labor and machine rental markets.

[1] Vietnam Living Standards Surveys was changed to Vietnam Household Living Standards Surveys in the rounds in 2000s. However, the survey instruments and sampling framework were virtually unchanged.

2. THEORETICAL FRAMEWORK

The empirical analysis that is the primary contribution of this study is framed with a simple theoretical model. Assume the land market does not function, so farmers have fixed landholdings. In this case, labor and machinery are the only two variable inputs. The production function of farm i is

$$Q_i = Q(L_i, M_i | h_i), \tag{1}$$

where Q_i is output, h_i is landholding, L_i is labor input, and M_i is machine use. Assuming constant returns to scale, equation (1) is rewritten as

$$y_i \equiv Q_i/h_i = f(l_i, m_i), \tag{2}$$

where y_i is household-specific yield (output per hectare) and $l_i = L_i/h_i$ and $m_i = M_i/h_i$ represent input application rates per hectare. Assuming that farmer i maximizes profitability and the price of output serves as numeraire, the factor demand functions can be derived based on the first-order conditions.

$$l_i = l(\theta_i, k_i) \tag{3}$$

$$m_i = m(\theta_i, k_i), \tag{4}$$

where θ_i is shadow price of labor and k_i is price of machine use. If $\theta_i = \theta$ and $k_i = k\ \forall i$, then factor prices are exogenous to individual households, and markets will allocate factors so as to equalize productivity across the fixed land endowments. But if a second market (besides land) functions imperfectly, then the (shadow) price of one or both factors will vary across production units, leading to a relationship between farm size and productivity (Feder 1985).

Following Feder (1985), Benjamin and Brandt (2002), and Barrett, Sherlund, and Adesina (2008), let θ_i increase in h_i to capture search and or supervision costs; following Foster and Rosenzweig (2011) and Feder (1985), let k_i decrease in h_i to reflect high-capacity (that is, lower unit price) machinery's greater suitability to larger plots, borrowing costs that decrease in collateralizable landholdings, or nontrivial fixed search or contracting costs of machinery rental. Plugging (3) and (4) into (2) gives

$$y_i = f(l_i, m_i) = f\big(l(\theta_i, k_i), m(\theta_i, k_i)\big). \tag{5}$$

Assuming machine and labor are pure substitutes, so that $\frac{\partial l_i}{\partial k_i}, \frac{\partial m_i}{\partial \theta_i} > 0$, from (3) and (4):

$$\frac{\partial l_i}{\partial h_i} = \frac{\partial l_i}{\partial \theta_i}\frac{\partial \theta_i}{\partial h_i} + \frac{\partial l_i}{\partial k_i}\frac{\partial k_i}{\partial h_i} < 0 \tag{6}$$

$$\frac{\partial m_i}{\partial h_i} = \frac{\partial m_i}{\partial \theta_i}\frac{\partial \theta_i}{\partial h_i} + \frac{\partial m_i}{\partial k_i}\frac{\partial k_i}{\partial h_i} > 0. \tag{7}$$

Combining (2), (6), and (7) results in

$$\frac{\partial y_i}{\partial h_i} = \frac{\partial f}{\partial l_i}\frac{\partial l_i}{\partial h_i} + \frac{\partial f}{\partial m_i}\frac{\partial m_i}{\partial h_i} = \frac{\partial f}{\partial l_i}\frac{\partial l_i}{\partial \theta_i}\frac{\partial \theta_i}{\partial h_i} + \frac{\partial f}{\partial l_i}\frac{\partial l_i}{\partial k_i}\frac{\partial k_i}{\partial h_i} + \frac{\partial f}{\partial m_i}\frac{\partial m_i}{\partial \theta_i}\frac{\partial \theta_i}{\partial h_i} + \frac{\partial f}{\partial m_i}\frac{\partial m_i}{\partial k_i}\frac{\partial k_i}{\partial h_i}\ . \tag{8}$$

Further assume that marginal productivity responds more to own-price than to cross-price effects—that is, $\left|\frac{\partial f}{\partial l_i}\frac{\partial l_i}{\partial \theta_i}\right| > \frac{\partial f}{\partial m_i}\frac{\partial m_i}{\partial \theta_i}$ and $\frac{\partial f}{\partial l_i}\frac{\partial l_i}{\partial k_i} < \left|\frac{\partial f}{\partial m_i}\frac{\partial m_i}{\partial k_i}\right|$—which is strictly true for a Cobb-Douglas production function. Notice that a nonconstant farm size–production relationship arises from incomplete or imperfect markets; if both labor and machinery rental markets function perfectly, then $\frac{\partial \theta_i}{\partial h_i} = \frac{\partial k_i}{\partial h_i} = 0$ implies that $\frac{\partial y_i}{\partial h_i} = 0$, even with the assumed absence of a land market. But if either the labor or machinery rental market functions imperfectly, then at least one term in (8) will be nonzero. If labor (machinery rental) market imperfections dominate, then $\frac{\partial y_i}{\partial h_i} < (>)\ 0$. The inverse relationship holds when labor market imperfections favor small farms; it is reversed when capital market imperfections favor larger farms.

If the process of structural transformation reduces market frictions and drives $\frac{\partial \theta_i}{\partial h_i}$ or $\frac{\partial k_i}{\partial h_i}$ toward zero, then the relationship between crop productivity and farm size will attenuate. The empirical regularity in low-income agriculture is that labor market imperfections dominate, leading to an inverse farm size–productivity relationship. Therefore, as structural transformation proceeds, driving up real wages, there should be attenuation of the commonplace downward-sloping relationship between crop yields and farm size. The next section explores that prediction using data from Vietnam.

3. EMPIRICAL SPECIFICATION

Based on the model of the preceding section, subindex *t* represents period and *c* represents commune. The shadow price of labor is specified as a function of the prevailing local (real) wage rate, w_{ct}, adjusted for the household-specific landholding that influences labor search and supervision costs:

$$ln\,\theta_{ict} = \alpha_i + \alpha_1\, ln\, w_{ct} + \alpha_2\, ln\, w_{ct} \times\, ln\, h_{ict} + \alpha_3 z_{ict} + \alpha_4 D_t + \varepsilon_{ict}\,, \tag{9}$$

where α_i is a household fixed effect that also captures time-invariant, location-specific effects; z_{ict} is a vector of household-specific, time-varying characteristics; D_t is a time dummy; and ε_{ict} is an independent and identically distributed (iid), mean zero, normally distributed random error term. The main coefficient of interest, α_2, is expected to be positive. The price of machine use is specified as

$$ln\, k_{ict} = \gamma_i + \gamma_{1t} + \gamma_2\, ln\, h_{ict} + u_{ict}, \tag{10}$$

where γ_i is a household fixed effect that also captures the time-invariant, location-specific effects; γ_{1t} captures period-specific fixed effects (including interest rates, which are assumed uniform across communes); γ_2 captures the shadow price effect of landholdings, which is expected to be negative; and u_{ict} is an iid, mean zero, normally distributed random error term.

Equations (3), (4), (9), and (10) combine to create the log linear factor demand equations:

$$ln\, l_{ict} = \beta_i + \beta_1\, ln\, w_{ct} + \beta_2\, ln\, w_{ct} \times ln\, h_{ict} + \beta_3\, ln\, h_{ict} + \beta_4 z_{ict} + \beta_5 D_t + e_{ict}, \tag{11}$$

$$ln\, m_{ict} = \delta_i + \delta_1\, ln\, w_{ct} + \delta_2\, ln\, w_{ct} \times ln\, h_{ict} + \delta_3\, ln\, h_{ict} + \delta_4 z_{ict} + \delta_5 D_t + \epsilon_{ict}, \tag{12}$$

where β_i and δ_i capture household fixed effects and D_t is a period fixed effect. In the labor demand equation (11), β_1 and β_3 should be negative and β_2 positive. In the machine demand equation (12), δ_1 and δ_3 are expected to be positive and δ_2 negative.

After plugging the factor demand functions into the crop yield equation:

$$ln\, y_{ict} = \xi_i + \xi_1\, ln\, w_{ct} + \xi_2\, ln\, w_{ct} \times ln\, h_{ict} + \xi_3\, ln\, h_{ict} + \xi_4 z_{ict} + \xi_5 D_t + v_{ict}, \tag{13}$$

where ξ_i is the household fixed effect, ξ_1 is expected to be negative, ξ_2 is expected to be positive, and ξ_3 can be either positive or negative depending on whether the capital or labor market imperfection dominates. Data permitting, one could estimate the system of equations (11)–(13) to test the model predictions about coefficients of interest. By testing the hypothesis that the parameters ξ_2 and ξ_3 change over the course of structural transformation in the economy, one can explore whether any initial relationship between farm size and productivity is indeed attenuated by increased factor market activity, as theory would predict.

4. DATA AND DESCRIPTIVE EVIDENCE

This study uses data from the 1992 and 1998 Vietnam Living Standards Surveys (VLSS) and from the 2002, 2004, 2006, and 2008 Vietnam Household Living Standards Surveys (VHLSS) (General Statistics Office of Vietnam). VLSS and VHLSS use similar survey instruments and are nationally representative. The survey instruments include a household questionnaire and a commune questionnaire. VLSS used the 1990 census in its sampling framework. VLSS 1998 intended to interview all households from VLSS 1992. VHLSS sampled based on the 2000 census. Each round in the VHLSS survey, except for 2002, re-interviewed some households sampled in the previous round and added some newly sampled households. (The 2002 round was based off a new sample according to the most recent census.) Any two or more rounds of VHLSS can form a household panel, though the sample size decreases as more rounds are included in the panel.

The data from household and commune surveys were merged to construct three rural household panel data sets: VLSS 1992/1998, VHLSS 2002/2004, and VHLSS 2006/2008. The VLSS 1992/1998 panel contains 3,034 households from 103 rural communes; the VLSS 2002/2004 panel, 2,303 households from 794 rural communes; and the VLSS 2006/2008 panel, 2,346 households from 956 rural communes.

The commune survey provides information on local agricultural wage by gender and by task (land preparation, planting, tendering, and harvesting). We generate median wage by gender, combining across all tasks for each commune for each panel round. To calculate the real wage, we deflate local nominal wage rates by the national consumer price index (CPI), which captures intertemporal inflation, and by the regional CPI, which captures spatial price variation.[2] Figure 4.1 plots the median female and male real agricultural wages from 1992 to 2008. The real wage increased significantly from 1992 to 1998. It leveled off from 1998 to 2004, probably reflecting the lagged effects of the Asian financial crisis of 1997/1998. From 2004 to 2008, the real wage again picked up rapidly, at a rate even faster than seen during 1992–1998. Table 4.1 reports the median male real wage by regions from 1992 to 2008, using VLSS and VHLSS commune survey data.[3] Although the real wage was consistently lower in the northern regions than in the southern regions, regional wage differences narrowed considerably by 2008, indicating an increasingly spatially integrated national labor market. This pattern is consistent with the stylized facts of structural transformation (Timmer 2007).

[2] The regional CPIs are provided in the dataset.

[3] It is not possible to generate a population-weighted national average because the wage data come from the commune survey, and commune-level weights are not available for these data sets.

Figure 4.1 Median daily real male and female agricultural wages, 1992–2008

Source: Authors' computations based on commune surveys in VLSS 1992 and 1998 and VHLSS 2002, 2004, 2006, and 2008.

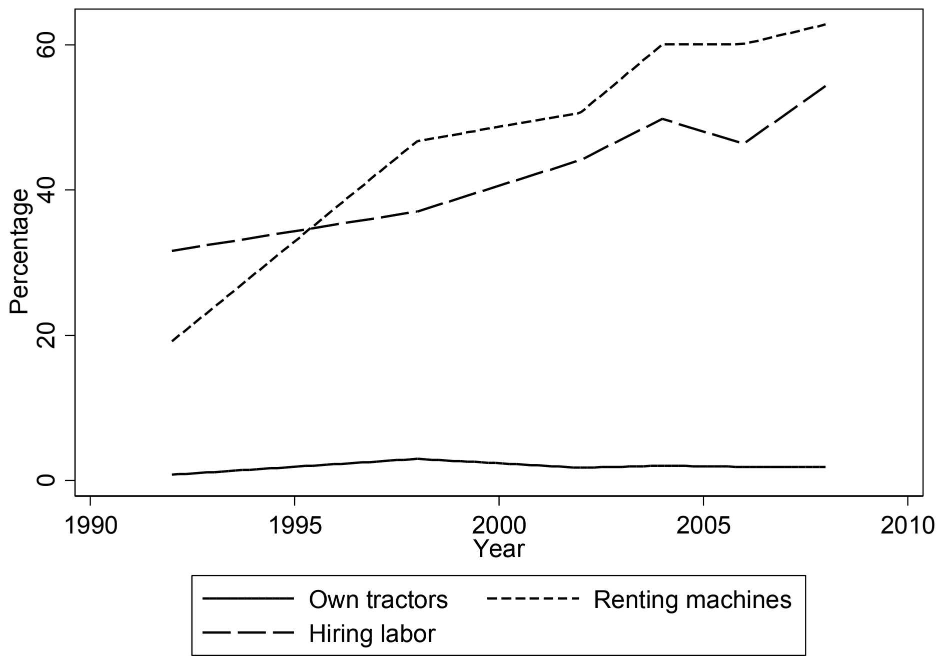
Table 4.1 Real median daily wage of male agricultural labor (000 Vietnamese dong)

Region	1992	1998	2002	2004	2006	2008
Red River delta	7.49	14.41	13.59	15.28	19.90	28.06
Northeast	5.16	11.17	11.00	13.00	15.78	22.22
Northwest	6.96	9.05	9.28	9.38	14.48	18.66
North central coast	7.67	12.12	12.96	13.22	19.95	23.58
South central coast	7.34	15.56	14.39	16.17	17.60	23.51
Central highlands	9.21	13.40	13.38	13.89	18.46	26.53
Southeast	11.44	15.70	17.26	17.14	21.60	25.63
Mekong River delta	15.01	18.69	17.51	19.04	22.41	25.53

Source: Authors' computations based on commune surveys in VLSS 1992 and 1998 and VHLSS 2002, 2004, 2006, and 2008.

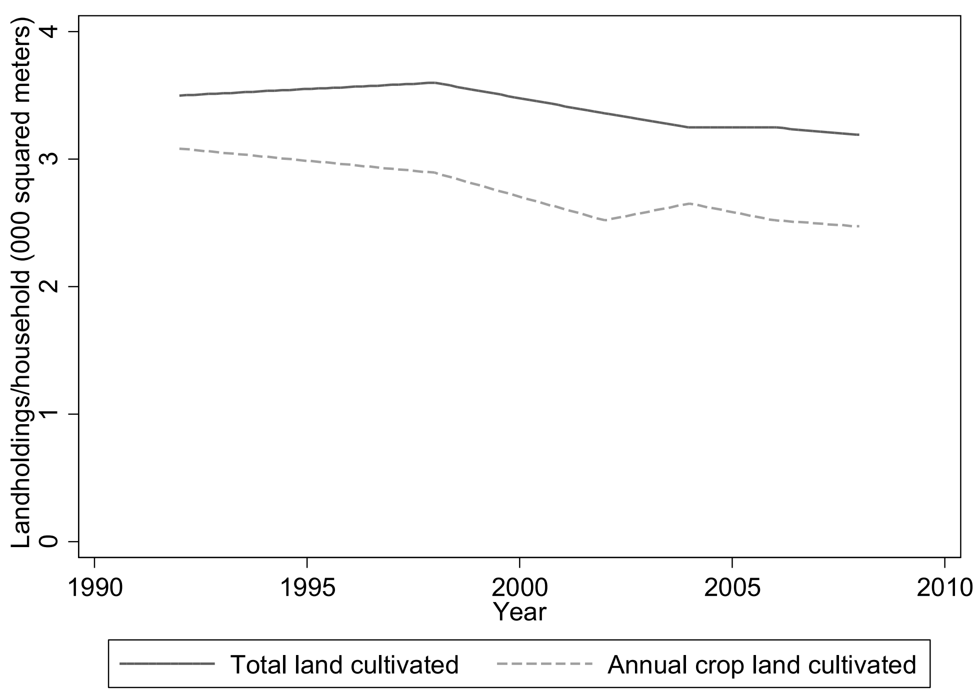
The household survey data provide information on household demographics, rice output by type of rice, rice planting area, landholdings, machinery use, labor hiring, and production cost by category. Figure 4.2 plots the proportion of households that rented machinery, owned tractors, or hired labor from 1992 to 2008. There is not much change in tractor ownership; the rate remains almost zero, which is not surprising given the rather low median landholdings in Vietnam (Figure 4.3). The percentage of cultivating households that rented machines more than tripled, however, from 19 percent in 1992 to 63 percent in 2008. This observation mirrors recent findings from China (Yang et al. 2013). The percentage of households that hired labor also increased sharply, from 32 percent in 1992 to 55 percent in 2008; by 2008, most cultivating households hired both labor and machinery services. Figure 4.3 plots the median area of land cultivated per household from 1992 to 2008. A median farm household cultivated 0.35 ha and 0.32 ha in 1992 and 2008, respectively. The cultivated area of annual crops dropped from 0.31 ha in 1992 to 0.25 ha in 2008 for the median farm household.

Figure 4.2 Trend of machine ownership, machine renting, and labor hiring, 1992–2008



Source: Authors' computations based on household surveys in VLSS 1992 and 1998 and VHLSS 2002, 2004, 2006, and 2008.

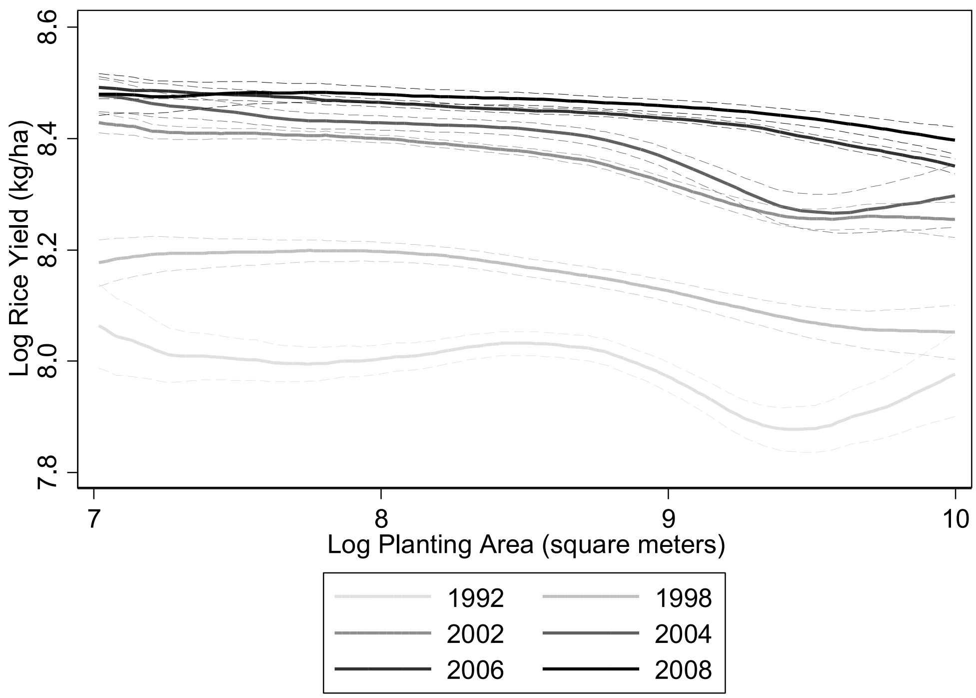
Figure 4.3 Trend of total land cultivated per household and total annual cropland cultivated per household, 1992–2008



Source: Authors' computations based on household surveys in VLSS 1992 and 1998 and VHLSS 2002, 2004, 2006, and 2008.

Figure 4.4 plots the locally weighted polynomial regression result of log rice yield against log planting area for 1992, 1998, 2002, 2004, 2006, and 2008.[4] The figure shows point estimates with solid lines and the 95 percent confidence intervals with dotted lines of the same color as the corresponding point estimates. Although the curve is downward sloping throughout, pointing to the presence of an inverse farm size–productivity relationship, it flattens considerably in 2006 and 2008, suggesting that productivity advantage for small farmers had decreased. The sharp upward displacement of the yield curves between each successive survey round, especially the first three, demonstrates the rapid growth in rice productivity.

Figure 4.4 Rice yield versus planting area, 1992–2008



Source: Authors' computations based on household surveys in VLSS 1992 and 1998 and VHLSS 2002, 2004, 2006, and 2008.

[4] The procedure used is the "lpoly" in Stata 13 SE with default optimal bandwidth.

5. ESTIMATION RESULTS

Ideally, we would estimate equations (11)–(13) as a system for improved efficiency. Due to the lack of total labor input data, however, we cannot estimate the labor demand function, equation (11). For machine use, we only know the total amount of money spent on machine rental and fuel but not the days of machine use, as required for estimation of equation (12).[5] We therefore estimate only the yield equation (13). For comparison, we use the 1992/1998 panel, the 2002/2004 panel, and the 2006/2008 panel separately. We also demean the variables $\ln w_{mt}$ and $\ln h_{mit}$ around the sample means, so the coefficients ξ_1 and ξ_3 can be interpreted as the mean partial effects (or partial effects evaluated at the same mean). Columns (1)–(3) of Table 5.1 report regression results on rice yield aggregated over all rice varieties, for the three panels.[6]

Table 5.1 Regression results on land productivity of rice (all varieties included)

Variable	(1) 1992/1998	(2) 2002/2004	(3) 2006/2008	(4) (2)-(1)	(5) (3)-(1)	(6) (3)-(2)
Log total area of rice (all varieties)	**–0.150*** **	**–0.119*** **	**–0.0607*** **	**0.0307**	**0.0894*** **	**0.0587* **
	(0.0296)	(0.0309)	(0.0152)	(0.0427)	(0.0332)	(0.0344)
Log total area of rice x log male real agricultural wage	**0.00554**	**–0.00384**	**0.0758*** **	**–0.00938**	**0.0702* **	**0.0796* **
	(0.0355)	(0.0365)	(0.0212)	(0.0508)	(0.0412)	(0.0421)
Log male real agricultural wage (Vietnamese dong in 1992)	–0.0235	0.0218	0.0223	0.0453	0.0458	0.000566
	(0.0793)	(0.0225)	(0.0248)	(0.0821)	(0.0828)	(0.0334)
Male household head	0.00126	0.0430	0.00506	0.0417	0.00381	–0.0379
	(0.0328)	(0.0400)	(0.00925)	(0.0516)	(0.0339)	(0.0410)
Age of household head	0.000920	0.000606	0.000243	–0.000313	–0.000676	–0.000363
	(0.00125)	(0.00128)	(0.000301)	(0.00179)	(0.00128)	(0.00132)
Highest education of household members	0.0239**	–0.00155	–0.000449	–0.0255**	–0.0243**	0.00110
	(0.0106)	(0.00586)	(0.00398)	(0.0121)	(0.0113)	(0.00708)
Number of male members	0.0172	–0.0221	–0.00258	–0.0393	–0.0198	0.0195
	(0.0175)	(0.0169)	(0.00455)	(0.0243)	(0.0180)	(0.0175)
Household size	0.00903	0.0194*	0.00489	0.0103	–0.00414	–0.0145
	(0.00902)	(0.0111)	(0.00600)	(0.0143)	(0.0108)	(0.0126)
Year dummy	0.334***	0.0499***	0.0178*	–0.284***	–0.316***	–0.0321**
	(0.0611)	(0.0106)	(0.00921)	(0.0618)	(0.0616)	(0.0141)
Household fixed effects	Yes	Yes	Yes			
Observations	4,744	2,777	3,780	7,521	8,524	6,557
R-squared	0.226	0.069	0.045	0.197	0.197	0.058

Source: Authors' estimation using VLSS 1992 and 1998 and VHLSS 2002, 2004, 2006, and 2008.
Notes: The variables "log total area of rice" and "log male real agricultural wage" are centered around their sample means. Standard errors in parentheses: * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

[5] As explained earlier, large farms are likely to face a lower price of machine use. It is thus inappropriate to use spending on machine rental to proxy for time of machine use due to heterogeneous pricing. In addition, if spending on machine rental were used as proxy for the actual machine use, we would underestimate machine use of the farm households that use their own machines. Dropping machine-owning households would also create a sample selection bias that is difficult to address without a credible instrument that affects machine purchase but does not affect machine rental. This study thus only looks at the extensive margin by generating a dummy variable indicating machine use. The model specification and estimation results are in the appendix.

[6] The rice varieties include winter-spring ordinary rice, summer-autumn ordinary rice, tenth-month or autumn-winter rice, ordinary rice planted in terraced field, year-round ordinary rice, year-round glutinous rice, and year-round specialty rice.

There are two main findings. First, the coefficient estimate on planting area is statistically significantly negative in all panels, suggesting the existence of an inverse farm size–productivity relationship, which is consistent with most of the published literature. However, with a value of –0.061, the estimated coefficient in the 2006/2008 panel is lower than that in the 1992/1998 panel (–0.150) and in the 2002/04 panel (–0.119). The result is consistent with the model prediction that the inverse farm size–productivity relationship attenuates as the labor and machine rental markets become more active and spatially integrated.

Second, the coefficient estimate on the interaction term of planting area and real wage is statistically significantly positive in the 2006/2008 panel, though it is insignificant in the 1992/1998 and 2002/2004 panels. This suggests that the inverse relationship attenuates most quickly in areas where the real wage rate is higher, creating greater incentives to substitute machinery for labor. The statistically insignificant estimate from 1992/1998 and 2002/2004 is consistent with the existence of imperfect labor markets in the earlier years in the survey, before wages began to reach a level at which substituting capital for labor began to appear potentially profitable to farmers.

To test for the significance of changes in farm size–productivity relationship (that is, to test the differences in key coefficient estimates between the panels), we pool the three panels, interact all the control variables with panel dummies, and run the same regression with the pooled dataset. The results are reported in columns (4)–(6) of Table 5.1. Although the coefficient estimate of the planting area is not significantly different between the 1992/1998 panel and the 2002/2004 panel, indicating that attenuation of the inverse relationship was modest at best, the effect was statistically significantly lower for the 2006/2008 panel than for either of the earlier two panels. Similarly, the coefficient estimate on the interaction term of planting area and real wage for the 2006/2008 panel was statistically significantly higher than for either the 1992/1998 or 2002/2004 panels. Together, these results reinforce the story that by the latter rounds of the survey, increasingly active rural-factor markets significantly reduced the inverse farm size–productivity relationship, and these effects were most pronounced in areas with higher real wages. The point estimates for the 2006/2008 panel (column (3)) suggest that an 80 percent increase in the real wage for male workers offsets the effect of doubling farm size.

In Table 5.1, the dependent variable is land productivity of rice aggregated over all rice varieties. These results may be biased if the choice of rice varieties is correlated with farm size and if the productivity differs across rice varieties. We thus run the same regressions for spring ordinary rice and autumn ordinary rice separately as a robustness check. The 2002 VHLSS does not distinguish between these rice varieties; therefore, the 2002/2004 panel is left out of these analyses. The results, reported in Tables 5.2 and 5.3, are similar to those reported in Table 5.1, showing a significantly decreasing inverse relationship for both spring and autumn rice over 1992/1998 and 2006/2008. The interaction term is significantly positive for both spring and autumn rice from the 2006/2008 panel, in line with the result from Table 5.1. For the 1992/1998 panel, the interaction is insignificant for spring rice but significantly positive for autumn rice.

Table 5.2 Regression results on land productivity of spring ordinary rice

Variable	(1) 1992/98	(2) 2006/08	(3) (2)-(1)
Log total area of spring ordinary rice	**–0.129***	**–0.0550***	**0.0743***
	(0.0256)	(0.0131)	(0.0287)
Log total area of spring ordinary rice x log male real agricultural wage	**–0.00845**	**0.0552***	**0.0636***
	(0.0326)	(0.0209)	(0.0386)
Log male real agricultural wage (Vietnamese dong in 1992)	0.0679	–0.00214	–0.0700
	(0.0665)	(0.0235)	(0.0702)
Male household head	–0.00877	0.0116	0.0204
	(0.0448)	(0.0110)	(0.0459)
Age of household head	–0.000110	0.0000715	0.000181
	(0.00120)	(0.000365)	(0.00125)
Highest education of household members	0.0119*	–0.0000726	–0.0119
	(0.00638)	(0.00570)	(0.00853)
Number of male members	0.00510	0.00493	–0.000172
	(0.0207)	(0.00491)	(0.0212)
Household size	0.00431	–0.00259	–0.00691
	(0.0115)	(0.00627)	(0.0131)
Year dummy	0.268***	–0.0118	–0.279***
	(0.0559)	(0.00875)	(0.0564)
Household fixed effects	Yes	Yes	
Observations	3,952	3,035	6,987
R-squared	0.249	0.024	0.222

Source: Authors' estimation using VLSS 1992 and 1998 and VHLSS 2002, 2004, 2006, and 2008.
Notes: The variables "log total area of spring ordinary rice" and "log male real agricultural wage" are centered around their sample means. Standard errors in parentheses: * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Table 5.3 Regression results on land productivity of autumn ordinary rice

Variable	(1) 1992/98	(2) 2006/08	(3) (2)-(1)
Log total area of autumn ordinary rice	**–0.216*** **	**–0.0893*** **	**0.126*** **
	(0.0406)	(0.0212)	(0.0456)
Log total area of autumn ordinary rice x log male real agricultural wage	**0.181** **	**0.115** **	**–0.0661**
	(0.0751)	(0.0542)	(0.0923)
Log male real agricultural wage (Vietnamese dong in 1992)	–0.200	0.0125	0.213
	(0.218)	(0.0473)	(0.223)
Male household head	–0.0321	0.0111	0.0432
	(0.0567)	(0.0161)	(0.0587)
Age of household head	–0.000106	0.000804	0.000911
	(0.00227)	(0.000508)	(0.00232)
Highest education of household members	0.0236*	–0.00417	–0.0277*
	(0.0125)	(0.00677)	(0.0141)
Number of male members	0.0341	0.00703	–0.0271
	(0.0242)	(0.00728)	(0.0251)
Household size	0.0206	0.0100	–0.0106
	(0.0142)	(0.00842)	(0.0165)
Year dummy	0.358***	0.0563***	–0.302***
	(0.104)	(0.0166)	(0.105)
Household fixed effects	Yes	Yes	
Observations	3,520	2,413	5,933
R-squared	0.150	0.076	0.139

Source: Authors' estimation using VLSS 1992 and 1998 and VHLSS 2002, 2004, 2006, and 2008.

Notes: The variables "log total area of autumn ordinary rice" and "log male real agricultural wage" are centered around their sample means. Standard errors in parentheses: * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

6. CONCLUSION

This study uses three household panel data sets from the 1990s and the 2000s from the same survey to explore the changing relationship between land productivity and rice planting areas in Vietnam. The findings show that the inverse relationship attenuates considerably over the course of those two decades. This change is associated with rising real wages and increasingly active machine rental and agricultural labor markets in rural Vietnam. In addition, the inverse relation attenuated most in areas with higher agricultural real wages by the mid-2000s, as real wages reached levels sufficiently high enough to induce some substitution of machinery for labor. As a result, the long-standing productivity advantage assumed to exist among smaller farmers appears to have diminished or disappeared altogether by the latter part of the period. Indeed, as real wages keep increasing, the inverse relationship may be reversed, leading to increased land concentration among farmers increasingly likely to employ machinery, without adverse effects on aggregate food production or prices.

APPENDIX: ESTIMATION OF MACHINE USE EQUATION

For the machine use equation, we generate a dummy variable as a dependent variable from the cost of machine rental and machine ownership. The variable takes value 1 if the household owned any tractors or spent on machine rental; otherwise, it takes 0. The estimated equation is:

$$d_{mit} = \delta_{0,mi} + \delta_1 \ln w_{mt} + \delta_2 \ln w_{mt} \times \ln h_{mit} + \delta_3 \ln h_{mit} + \delta_4 z_{mit} + \delta_5 D_t + \epsilon_{mit}, \qquad \text{(A1)}$$

where d_{mit} is the dummy variable indicating machine use for household i in commune m and year t. We estimate equation (A1) using the 1992/98, the 2002/04, and the 2006/08 panels separately.

The results are presented in Table A.1. For all three panels, larger farmers are more likely to use machines, consistent with theoretical prediction. Interestingly, machine use is not responsive to real agriculture wage in the 1990s panel or in the early 2000s panel. In contrast, it is significantly higher when real wage is higher in the 2006/08 panel, suggesting the efficiency improvement of rural-factor markets.

Table A.1 Regression results on machine use equation

Variable	(1) 1992/98	(2) 2002/04	(3) 2006/08
Log total land cultivated	0.0787***	0.0919***	0.0649***
	(0.0233)	(0.0209)	(0.0192)
	0.0174	0.0602	–0.0214
Log total land cultivated x log male real agricultural wage	(0.0364)	(0.0475)	(0.0279)
	–0.0545	0.0778	0.102**
Log male real agricultural wage (Vietnamese dong in 1992)	(0.0931)	(0.0519)	(0.0410)
Male household head	–0.0596	–0.0364	–0.00887
	(0.0441)	(0.0665)	(0.0191)
Age of household head	–0.0000624	0.000839	–0.000588
	(0.000908)	(0.00189)	(0.000517)
	0.00143	0.0104	–0.000792
Highest education of household members	(0.00664)	(0.00805)	(0.00600)
Number of male members	0.0321	–0.00308	0.00489
	(0.0197)	(0.0277)	(0.00825)
Household size	–0.0288**	0.00789	–0.00357
	(0.0124)	(0.0178)	(0.0106)
Year dummy	0.296***	0.0599***	–0.00142
	(0.0535)	(0.0158)	(0.0145)
Constant	0.311***	0.430***	0.693***
	(0.0750)	(0.138)	(0.0690)
Observations	5,241	3,378	4,431
R-squared	0.193	0.037	0.015

Source: Authors' estimation using VLSS 1992/98, VLSS2002/04, and VHLSS 2006/08.
Notes: "Log total land cultivated" and "log male real agricultural wage" are centered around their sample means. Standard errors in parentheses: * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

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