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Persistence of full- and part-time farming in Southern China

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Abstract

The goal of this study is to assess the dynamics of rural households' labor market participation in the wake of China's efforts to develop rural labor markets in a manner that is conducive to its transition to a market economy. Based on a theoretical model that emphasizes the impact of duration, i.e. of the number of years households spent part-time farming or full-time farming, respectively, we investigate the shifts between these two states. We also identify socioeconomic factors that determine these shifts. The empirical study is based on discrete time hazard approaches, using micro-level panel data from Zhejiang, Hubei, and Yunnan provinces from 1995 to 2002.

Estimation results suggest relatively high chances of shifts from full-time to part-time farming and a considerable lower risk for the shift in the opposite direction. Significant negative duration dependence is found for the move from full-time to part-time farming suggesting lock-in effects. In addition, we find that labor market participation decisions are significantly related to several household, farm and village characteristics. In particular, the likelihood of shifts from full-time farming to part-time farming is positively related to the educational level of households' workforce.

JEL Classifications: C41, J22, J 24, J43, J62, Q12

Keywords: farm household, labor market participation, full-time farming, off-farm employment, hazard analysis, dynamics

1 Introduction

Households have the privilege to allocate their available time to different uses to maximize their utility (Singh, Squire and Strauss, 1986; Huffman, 1991). Their choices among different kinds of income generating work and leisure reflect individual characteristics and preferences and they may be restricted by factors like capital market constraints and the political and legal framework. A well studied example of the impact of various policy reforms on time allocation of rural households is the case of China, moving from a very restrictive system of village communes and brigades to an agricultural sector dominated by small peasant farms with a high prevalence of off-farm activities. Some selected cornerstones of this liberalization were the distribution of collectively used land to private households starting in 1978, the dismantling of commune and brigade enterprises and their transformation into township and village enterprises in the mid 1980s, weakening of public grain procurement and the household registration system¹, as well as the abolition of agricultural taxes and introduction of direct payments to grain producers in 2004.² However, agricultural land is still owned by the state and the limited transferability of land might be one of the major political constraints for households' flexibility in factor allocation. A number of studies reveal that factor proportions are sometimes determined in a way which is hardly consistent with unconstrained labor allocation. For instance, Benjamin and Brandt (1997) and Liu et al. (1998) identify an inverse relationship between farm size and labor use. Meng (2000) finds evidence that off-farm participation is influenced by restrictive land tenure arrangements and the mandatory grain procurement system. Results by Bowlus and Sicular (2003) as well as Wang et al.

¹ In order to maintain the use right on the distributed land, households were obliged to fulfill a grain quota in kind or in cash to the state. The household registration or residency permit system registers rural and urban households separately and firmly determines the access to public services, e.g. education, housing, or public welfare.

² For a more detailed description of agricultural and rural reforms see OECD (2005), People's Daily Online (2004), Fan et al. (2002), Albersen et al. (2000) as well as Huang, Lin and Rozelle (1999).

(2007) indicate non-separability between Chinese agricultural households' consumption and production decisions.

To study households' labor market participation behavior it is clear that cross-sectional approaches cannot adequately capture important dynamic aspects (Nakamura and Nakamura, 1985). One reason is, that duration effects can not be accounted for. Ample literature shows the existence of employer-specific human capital, its foundation is that experience gained within an occupation raises productivity. See Farber (1999) for a review of the relevant literature and the development of a formal model of employment-specific human capital formation in the case of wage employment. Self-employment like agricultural production or non-agricultural self-employment could additionally benefit from the establishment of occupation-related social networks. As illustrated by Fafchamps and Minten (2002) in the case of agricultural traders, larger social networks affect sales and value added significantly. Usually, the opportunity costs of taking up a new job off the farm compared to continuing the current full-time farming activity will increase over time because farm profits benefit from extending networks and accumulating human capital. A similar duration dependency holds for the transition from part-time farming to full-time farming. For instance, a household member who is working off-farm, acquires occupation specific human capital and increases his labor productivity which raises opportunity costs of leaving this job. This aspect is neglected in the above mentioned previous studies. Thus, more recently, several analyses in agricultural economics literature use longitudinal data sets to analyze farm households' labor market behavior. Some of these studies focus on the link between decisions on long-term farm development and farmers' labor market behavior. Kimhi (2000), for example, links the decisions on off-farm participation and farm exit. Weiss (1997) and Kimhi (2006) investigate the relationship between off-farm labor market participation and farm growth and farm size, respectively. Giles (2006) analyzes potential impacts of the accessibility of labor markets on the variability of Chinese rural households' incomes. He

finds that households' vulnerability to shocks in agricultural production is substantially reduced by improved access to local and migrant labor markets.

A couple of studies explicitly focus on farmers' labor market decisions at different points in time. They particularly investigate the persistence of labor market decisions. Gould and Saupe (1989) and Weiss (1997) use two-period panel data sets to account for state dependence of off-farm labor market participation decisions. Corsi and Findeis (2000) apply a dynamic model of off-farm labor market participation to distinguish between true state dependence and unobserved heterogeneity. Juvancic and Erjavec (2005) analyze asymmetries and other dynamic aspects of Slovenian farmers' labor allocation during the transition period between 1991 and 2000. Chen et al. (2004) extend a dynamic discrete choice approach to account for state dependence of Chinese farm households' labor market decisions. Zhao (2002) particularly links Chinese agricultural households' migration activities to the number of earlier migrants from the respective village. Zhang et al. (2001) focus on Chinese agricultural households' off-farm labor market participation during times of recession. Their study is the only analysis so far which looks at the determinants not only of *beginning* off-farm participation but also of *ending* this state.

A growing number of studies in labor economics literature apply duration or hazard models (Kalbfleisch and Prentice, 2002) to study households' labor market histories. This approach allows quantifying the impact of duration, i.e. of the time spent in a certain state, e.g. full-time farming, on the instantaneous probability to leave this state.³ In this way it controls for the accumulation of occupation-specific human capital and social networks as determinants of labor market participation. Considering duration avoids potential bias in other variables' coefficients which could arise through neglecting duration. Additionally, duration dependence implies that households' responsiveness to policy interventions depends on their

³ Examples of the methodology applied to similar research questions can be found in Bradley et al. (2003) as well as Blau and Riphahn (1999). Applications in the case of transition countries are presented by Orazem and Vodopivec (1997), Sorm and Terell (2000) and Appleton et al. (2002).

occupational history. This knowledge can be used to design effectively targeted policies. For instance, if a government intends to foster sectoral change by making farm households diversify their income generating activities, and if households' inclination towards such diversification decreases with duration, then the government should target policies to those households which spent least time exclusively in farming. Brosig et al. (2007) apply a discrete time hazard approach to analyze rural households' labor market histories in Zhejiang province. The authors consider movements between labor market participation of any kind (local or migrant off-farm work, employment of hired labor) and non-participation. Estimation results suggest significant negative duration dependence of the probabilities to move between these states.

The present study defines part time farming by local off-farm work and it investigates the flexibility of Chinese farm households in three provinces regarding their transition from full-time farming to part-time farming and vice versa⁴ taking into account the length of time spent in the original participation state. In this way the goal of the study which is to assess the dynamics of rural households' labor market participation will be pursued. Are there any differences between rich and poor provinces? Do households show an asymmetric behavior in starting and stopping off-farm work? After a brief sketch of the theoretical framework we present a discrete time hazard model, applied to individual household data over the period 1995-2002 from the provinces Zhejiang, Hubei, and Yunnan. The model accounts for unobserved heterogeneity to examine both the likelihood of transitions between the different labor market participation states over time and factors affecting these likelihoods. Knowledge about rural households' labor flexibility is required to develop appropriate and targeted policy measures to increase labor mobility and development.

⁴ In principle, an alternative for both transitions mentioned could be to give up farming entirely. We have decided not to consider this option in our analysis because it is only chosen in extremely rare cases (by less than 0.1 percent of the farms).

In the specific way that we apply a hazard model we go beyond the existing literature focusing on dynamic aspects, the transitions between participation states rather than taking a static view focusing on the states themselves. For three Chinese provinces we analyze the duration dependence of movements between participation states. In analyzing duration dependence of the probabilities of transitions we provide a complement to results on state dependence as e.g. in Weiss (1997).

The remainder of the work is organized as follows. The next section provides the theoretical background. This is followed by an overview over the data. Section 4 presents the methodology and the empirical results, and section 5 concludes.

2 A simple model of movements between full-time and part-time farming

The decision to move between alternative labor market participation states is illustrated by a simple model of occupation choice. It ignores some aspects of farmers' decisions, like risk (Fafchamps, 1992; Finkelstein and Chalfant, 1991; Folz 2004; Serra et al., 2004), capital market constraints (de Janvry et al., 1991; Benjamin and Phimister, 2002), as well as the role of costs associated with transactions on product markets (Key et al., 2000) and other input markets.

Consider a farm household that selects between two mutually exclusive labor market participation states: full-time farming and part-time farming. Part-time farming households allocate any amount of family labor to non-farm activities while they may or may not employ any hired labor on their farm. Full-time farming households allocate all family labor (plus potentially hired labor) to work on their farm.

The model interprets households' choices between full-time and part-time farming as the outcome of a utility maximizing process (Brosig et al. 2007, p. 168ff.). Different levels of expected utility are associated with farming types $j \in \{FF \text{ (full-time farming); } PF \text{ (part-time farming)}\}$ which can be represented by indirect utility functions $V^j(p, Y, Z)$. While p denotes a vector of consumption prices, Y indicates potential (farm- and non-farm) income,

and Z contains exogenous utility shifters corresponding to socio-demographic household and regional characteristics. We propose in this paper that the V^j may depend significantly on factors which change with the number of years that a household occupies its labor market participation state. Duration dependence of households' labor market participation decisions has been indicated by several studies (e.g. Chan and Stevens, 2001) while state dependence was emphasized e.g. by Weiss (1997) as well as Corsi and Findeis (2000). A rationale is that accumulated capital and skills, say agricultural or non-agricultural working skills, might increase the income potential of any chosen labor market state, as full-time or part-time farming, after a couple of years of experience. Thus, a longer stay in the respective state would increase the utility level attainable with the respective occupation. On the other hand, remaining in one state for years might reduce the income potential that the household has in the state not occupied, due to lost skills, learning foregone, or shrinking social networks. Finally, households' preferences for each of the alternative occupations may change over the years. Hence, by introducing a variable DUR which denotes the number of years that a household has so far remained in its present participation state, the indirect utility functions can be written as:

$$\begin{aligned}\Psi_{t^*}^{FF} &= \int V_t^{FF}(p_t, Y_t^{FF}, Z_t, DUR_t) e^{-rt} dt \\ \Psi_{t^*}^{PF} &= \int V_t^{PF}(p_t, Y_t^{PF}, Z_t, DUR_t) e^{-rt} dt\end{aligned}\tag{1}$$

Ψ^j indicates the discounted working time utility from full-time farming and part-time farming, respectively, and r denotes the rate of discount. The model predicts the movement of a household from one labor market state j to an alternative labor market state $i \in \{FF, PF\}$ when the expected working time utility associated with i , say part-time farming, exceeds the utility of the current state j , say full-time farming, net of the utility loss due to the cost associated with the transfer between the states. Associated utility losses are denoted here as transfer cost: $TC_{t^*}^{ji} = \int C_t^j e^{-rt} dt$. This disutility may include both pecuniary and non-

pecuniary cost; for example, the disutility of loosing the status of a full-time farmer.⁵ The net utility gain of a status change from j to i is hence expressed as (Breustedt and Glauben, 2007; Barkley, 1990):

$$H_{t^*}^{ji} = \Psi_{t^*}^i - \Psi_{t^*}^j - TC_{t^*}^{ji} \quad (2)$$

A change of the labor market participation state at time t^* thus occurs when the net gain of the change is positive ($H_{t^*}^{ji} > 0$). The household remains in its current state otherwise ($H_{t^*}^{ji} < 0$). That is, the farm household will quit full-time farming at time t^* , if the discounted utility of full-time farming is lower than the discounted utility of part-time farming less associated transfer cost.

An implication of the behavioral model sketched here is, that duration, i.e. the time that a household has remained in the participation state that it presently occupies, may affect the probability of an occupational change. Our study will evaluate the empirical backing of this implication.

3 Data

For the empirical analysis we use individual household and village survey data from the three Southern Chinese provinces Zhejiang, Hubei, and Yunnan over the period 1995-2002. The Ministry of Agriculture's annual survey is representative of rural households and the data were provided by the Research Center of the Rural Economy (RCRE), Beijing. The number of households in the sample is around 500 for Zhejiang and Yunnan, respectively, and 900 for Hubei. Table 1 allows comparison of rural households' social situation and of the structure of agricultural household businesses. Per capita income is near the national average (of 2622 Yuan in 2002) in Hubei while it is significantly above average in Zhejiang and below average in Yunnan. Household business activities - particularly farming - generate by far the biggest

⁵ In the case of current Chinese agricultural policy, the status of full-time farming might have not only emotional aspects. The lack of full property rights of land and the risk of loosing access to land if it is e.g. leased out to other households may add to utility of full time farming (Jacoby, Li and Rozelle, 2002).

component in household income in landlocked provinces Hubei and Yunnan while in the more diversified economy of coastal Zhejiang it represents 43 percent.

The agricultural sectors of the three provinces are dominated by small scale farms. The vast majority of the households in the sample is holding arable or orchard land. Most of these farms' (76 percent) land area is between 0.1 and 0.5 ha. Average farm sizes in the sample reflect that average land area per household in South East China is below the national average of nearly 0.6 ha (2000, excl. grassland, Stat. Yearbook 2001). The annual production value per farm is around six to seven thousand Yuan (in prices of 2000) in the three provinces. In Zhejiang, half of the produce in value terms is consumed by household members (or transferred to other households without payment) while the share of non-marketed output is above 70 percent in Hubei and Yunnan. Production is diversified. The most important products in terms of output value are fruit and vegetable as well as rice in Zhejiang and Hubei and pork and corn in Yunnan.

Table 1: Household income and farming activity by province, 1995-2002

	Zhejiang	Hubei	Yunnan
Households in sample holding arable land [%]	79	97	93
Per capita net income rural households, 2002 [1000 RMB] ⁶	4,3	2,3	1,5
Income share of household business in total income (2003) ^{7 8} [%]	43%	70%	73%
Average land endowment of household [ha]	0.27	0.36	0.52
Most common size class (arable land + orchard)	0.1-0.5 ha		
Share of farms with 0.1-0.5 ha	76%	77%	56%
Volume of annual production [1000 RMB]	5.9	6.9	6.4
within this: share of home consumption	50%	74%	71%
Main products (% in value of production)	Fruit & Veg. (25) Rice (22), Pork (22)	Fruit & Veg. (32) Rice (23) Pork (15)	Pork (21) Corn (20) Other grain (16)

Source: own computations based on RCRE data.

We now consider the prevalence of farm households' participation in local off-farm labor markets, i.e. the prevalence of part-time farming as reflected in our sample.⁹ Almost all of the farm households (92 percent in Hubei and 98 percent in Zhejiang and Yunnan, respectively) reported to have gone through phases of part-time farming during the eight years from 1995 to 2002. However, the terms of this involvement are sometimes relatively short so that periods of part-time farming and full-time farming alternate. Consequently, if individual years rather than the sample period as a whole are considered, the mean participation rates are

⁶ Statistical Yearbook 2004, Tab 10-20/ 10-22

⁷ Household business income originates to a large extent (40%) from agricultural activities.

⁸ Value of production in prices of 2000. Mean (deflated) unit values of *sales* over the 1995-2002 period were used in order to have a sufficiently large data base to compute average prices for each product.

⁹ We define a household to be part-time farming in a given year if it a) uses any agricultural land and b) has devoted any number of working days to income generating local off-farm activities (in non-agricultural family business or in wage work). Work as long-term migrants in distant locations was not considered in this definition because a very different set of patterns and determinants would likely be necessary to describe decisions for this kind of off-farm activity than it is for local off-farm work (De Brauw et al. 2002, Shi et al., 2007, Xia and Simmons, 2004).

considerably lower; they are 80, 65, and 66 percent in Zhejiang, Hubei, and Yunnan, respectively.

These frequent changes between participation states motivate to analyze the dynamics of participation behavior. Aggregate participation rates vary considerably over years and the majority of households that were participating in the labor market changed between full-time and part-time farming more than once.

Given the high frequency of changes we focus on the length of the periods in each of the labor market participation states. We analyze the duration dependence of the probability that households move between full-time and part-time farming and assess the impact of a number of household, farm, and village characteristics on this probability. Table 2 lists these variables and presents their means, conditional on the households' participation in the off-farm labor market. The selection of variables is based on findings of comparable studies and on theoretical considerations. For most of the variables the distribution differs significantly between part-time and full-time farming households (see column Sig. diff.) which hints at their relevance for our analysis.¹⁰

¹⁰ Given that the distribution of all our covariates was found to be significantly different from normal, we applied appropriate nonparametric tests (Wilcoxon and Median Two-Sample tests) for comparisons between two independent samples (SAS NPAR1WAY procedure). Detailed results are available from the authors.

Table 2: Household, farm, and village characteristics by labor market participation state*

		All holds (N=14739)	house- Part-time farms (N=10033)	Full-time farms (N=4706)	Sig. diff.
Household characteristics					
Household size [persons]	<i>HSIZE</i>	4.14 (1.45)	4.26 (1.42)	3.87 (1.45)	**
No. of females among family laborers	<i>FEM</i>	1.24 (0.71)	1.26 (0.70)	1.19 (0.74)	**
Secondary school graduates among laborers [percent]	<i>EDUS</i>	38 (35)	40 (35)	34 (34)	**
Communist party member [1/0]: % of households	<i>PARTY</i>	0.15 (0.36)	0.16 (0.37)	0.15 (0.36)	
Employing hired labor [1/0]: percent of households	<i>EMPLOY</i>	0.07 (0.26)	0.07 (0.26)	0.07 (0.26)	
Any members working as migrants [1/0]: % of households	<i>MIGRAN</i>	0.40 (0.49)	0.60 (0.49)	0.31 (0.46)	**
Farm Characteristics					
Agricultural land [ha per capita]	<i>LAND</i>	0.10 (0.10)	0.10 (0.11)	0.09 (0.09)	**
Animal production, weighted output [kg/100]	<i>LIVES</i>	1.71 (6.23)	1.60 (7.02)	1.93 (4.75)	**
Village characteristics					
Annual p.c. income [Yuan in 2000 prices]	<i>ANIPP</i>	2863 (3114)	2940 (3192)	2599 (2816)	**
Share of leased land on arable land	<i>RRAL</i>	0.08 (0.14)	0.08 (0.15)	0.06 (0.11)	**

*Standard deviations in parentheses

** Distribution differs significantly between full-time and part-time farms (see footnote 7)

Source: own computations based on RCRE rural household panel.

The first variables are supposed to capture the available labor capacity in the household and its composition which is of primary importance for labor market involvement. The total number of household members (*HSIZE*) comprises the labor force, children and persons above working age. Previous literature on labor supply indicates that male and female laborers exhibit different patterns of labor market participation (de Brauw et al., 2002; Meitzen, 1986). Thus, the variable *FEM* controls for the number of females in a family's total

labor force. The relevance of education and skills for taking up off-farm work has been emphasized among others by Chaplin et al. (2004), Cook (1999) and Zhang et al. (2002). We consequently account for education by considering a variable (*EDUS*) that indicates the share of persons who graduated at least from secondary school in the number of working age members of the household. The mean of this variable is 34 percent for full-time farming households and it is significantly higher, 40 percent, for households that do also supply off-farm work. It has been found earlier that membership in the Chinese Communist Party has linkages with employment opportunities on the one hand and with access to scarce farm inputs on the other hand. To capture this effect a dummy variable (*PARTY*) is included in the model. The variable *EMPLOY* indicates whether a household does employ any hired labor in (agricultural or non-agricultural) family business. It is hypothesized that the decisions on supplying and employing labor to/from the market are interdependent. On the one hand, families with highly labor demanding farm business may have little capacities to supply off-farm work, on the other hand, increasing specialization could mean that specially skilled or educated family members may be released from agricultural work if hired workers are employed. Around 7 percent of the households employ any hired labor.¹¹ The variable *MIGRANT* identifies households with any member staying and working in distant locations. This affects the available labor capacity of the household as well as its income via remittances.

Farm characteristics are supposed to control for differences in factor endowment and production structure that are likely to affect the family labor capacity available for off-farm engagement. The total area of arable and orchard land per capita (*LAND*) characterizes the economic size of the farm relative to household size. It is likely to influence the households'

¹¹ While for all other variables the hypothesis of equality of distribution between part-time and full-time farming households was rejected, for the variables *EMPLOY* and *PARTY* it was not. However, this does not preclude the possibility that in a multivariate context and looking at durations (or probabilities of status change) the same variables are informative as determinants of households' decisions on off-farm work.

involvement in the labor market because it affects the income capacity and labor requirements of the farm business. Labor requirements and income capacity also depend on the production structure. Particularly some forms of animal husbandry require high amounts of labor input. As a proxy to control for this we use a weighted (according to labor requirement) sum of the output quantities of the main livestock products beef, pork, poultry, milk, and eggs (*LIVES*). Although comparisons of the farm characteristics must be interpreted carefully because of their large variance it is worth noting that the mean of land area is larger in part-time farms than in full-time farms while the extent of livestock production is larger in full-time farms.

Aside from household composition and farm characteristics, the village environment is likely to affect the attractiveness and the costs of labor market participation. The transaction costs incurred with finding and doing off-farm work are likely to vary with regional structures. The number of available employment opportunities can be assumed to reduce transaction costs. Some studies use the distance to the nearest city (Huffman, 1991) or the frequency of public transport services (Chaplin et al., 2004) to account for transaction costs. While comparable information is lacking in the data set used here, we assume the average income (*ANIPP*) to be related with the economic activity and the transaction cost level. It is related to a village's general economic development and hence to (local) non-farm work opportunities. We find part-time farms more often in villages with relatively high income levels. The second village variable indicates the share of leased land among total agricultural land in a village (*RRAL*). One precondition for structural change in farming is that land can be transferred between households. This maintains that some households can intensify their farming and potentially employ hired labor while others may reduce or abandon farming activities and start or increase non-agricultural occupations.¹² Carter and Yao (2002) found an

¹² Although land transactions are still restricted in China to a regionally differing extent a land rental market has emerged in the last two decades (Kung, 2002).

impact of land transfer rights not only on the land-labor ratio but also on agricultural efficiency.

The initial analysis of the labor market participation of sample households has shown that changes between participation states occur very frequently. Moreover, the conditional distributions of some of the household, farm, and village characteristics considered differ markedly between part-time and full-time farm households.

4 Empirical approach and results

We analyze the shifts from the labor market participation state 'part-time farming' to 'full-time farming' and vice versa using methods of duration (or failure time) analysis (e.g. Kalbfleisch and Prentice, 2002). We identify how the risk of shifting from full-time to part-time farming and vice versa changes with the length of time a household spent in its respective state. Further, the additional impact of a number of covariates, such as farm, household, and regional characteristics on the probability to move between participation states is assessed. The two models were estimated separately for data sets from three provinces resulting in a total of six models. A central concept in duration analysis is the hazard function:

$$\lambda(t) = \lim_{\Delta t \rightarrow 0} \frac{\Pr(T \geq t)}{\Delta t}. \quad (4)$$

It expresses the instantaneous probability that a state, say a household's labor market participation state, ends after a duration of $T=t$, conditional upon having lasted for t periods. Or roughly spoken, $\lambda(t)$ is the rate at which spells are completed at duration t , given that they last until t . The hazard function provides a convenient definition of duration dependence. Positive duration dependence and an increasing hazard to leave a state respectively exists at t if $d\lambda(t)/dt > 0$. That is, the probability to leave a state, e.g. full-time farming, increases with the number of periods a household remains in this state. Negative duration dependence and a

decreasing hazard respectively, exists at t if $d\lambda(t)/dt < 0$. Here, the probability of leaving the present state falls with increasing length of the respective participation spell.

Our parametric specification of the hazard approach to analyze the movements between the labor market states is based on a proportional hazard model allowing for time-varying covariates. The hazard function can in this case be represented by

$$\lambda_k^{ji}(t, z_k, \beta^{ji}, \theta^{ji}) = \lambda_0^{ji}(t, \theta^{ji}) \exp[z_k(t), \beta^{ji}]. \quad (5)$$

Here $\lambda_k^{ji}(t, z_k, \beta^{ji}, \theta^{ji})$ denotes for household k the hazard of the transition from state j to state i . Further, $\lambda_0^{ji}(t, \theta^{ji})$ indicates the baseline hazard of a transition event j, i meaning the hazard under the condition of no heterogeneity among the individuals. It may depend on duration t and distributional parameters θ^{ji} . However, behavioral heterogeneity among individuals might change the individual hazard. Part of such variation can be accounted for by controlling for household k 's observed individual (time varying) characteristics $z_k(t)$. In the proportional hazard model this is accomplished by a multiplicative term, $\exp[z_k(t), \beta^{ji}]$ in the case of specification (5). The vector $z_k(t)$ comprises in our study household, farm and regional characteristics as introduced in section 3. Parameters β^{ji} measure the impact of the covariates and need to be estimated empirically. If $\exp[z_k(t), \beta^{ji}] > 1$, then the probability of the event j, i , that is the movement between the labor market participation states j and i , is above the baseline for household k . If $\exp[z_k(t), \beta^{ji}] < 1$, the opposite holds. The covariates z_k are usually scaled such that at their mean value $\exp[z_k(t), \beta^{ji}, \theta^{ji}] = 1$. Aside from the observed covariates $z_k(t)$ an unobserved spell-specific random effect is assumed to change the baseline hazard in a multiplicative way. To avoid estimation bias due to such unobserved heterogeneity (frailty) the parameter θ^{ji} is included representing the frailty variance assuming a gamma distribution (e. g. Blau and Riphahn, 1999; Meitzen, 1986; Meyer, 1990).

Spell lengths of our (annual) data are grouped according to years of duration. The number of consecutive years in which a household has reported to have remained in a participation state is interpreted as the spell length in years.¹³ The length of each period (one year) is relatively long in comparison with spell durations in the sample (less or equal 8 years) which necessitates application of estimation techniques for discrete duration data (Kalbfleisch and Prentice, 2002; 46ff.). We applied the grouped data approach suggested by Prentice and Gloeckler (1978) in Meyer's (1990) modification to account for frailty. The 'easy estimation method' proposed by Jenkins (1995) was parameterized using a complementary log-log link function so as to make it the discrete counterpart of a proportional hazard specification with a Weibull baseline hazard function. The method requires a data set which contains one observation for each year that a household remains in the respective participation state. A dichotomous variable ξ is defined that takes the value one if the participation spell is completed in that respective year and the value zero otherwise, i.e. for censored spells and for spells not yet completed. The λ -parameter of the Poisson-distribution of ξ conditional on covariates (e.g. household characteristics) and on the number of years that the participation spell lasted so far is then estimated as a binary choice model.¹⁴ In most models a high percentage of observations on individual spells had to be considered censored. We treated all spells neighboring unobserved periods on either side as "right-" censored, hinting at the possibility that the actual spell lasted longer than the recorded duration. This assures that the spell length information of these spells is 'interpreted' appropriately in the likelihood function,

¹³ Any intermissions of participation during the year are hence ignored, which implies interval censoring with respect to actual non-participation spells. However, this treatment seems appropriate because households which participate during parts of a year can be considered participants of the labor market, in the sense that they have market information and are *potential* off-farm workers throughout the year.

¹⁴ We used the STATA program `pgmhaz8` written by Stephen Jenkins (Jenkins 1997) which implements the procedure sketched here.

i.e. as greater or equal to the indicated length rather than equal. For instance, among 723 spells of part-time farming in Zhejiang this applies to 602 spells.

Tables 3a and 3b present results of duration model estimations for the movement from full-time to part-time farming and vice versa. The variable indicating movements between the participation states is regressed on a duration variable $DUR = \ln(\text{duration})$, on selected covariates describing household, farm, and village characteristics, and on the variance of unobserved heterogeneity among participation spells (*frailty*). Regarding the covariates we experimented with multiple specifications using variables that have been found influential in other labor market studies and our final specification contains the variables introduced in section 3. Given the parsimony of the parametric duration specification, degrees of freedom were not a primary concern but optimization problems were pervasive due to the highly unbalanced samples with few spell completions in comparison with the large number of observations from censored and continued spells (cf. Appleton et al., 2002). Table 3 refers to the models with the final selection of covariates.

Table 3a: Estimation results of duration model: FULL-TIME TO PART-TIME (taking up off-farm work)

	ZHEJIANG	HUBEI	YUNNAN
<i>DUR</i> [ln(years)]	-0.912** (0.176)	-0.633** (0.090)	1.212 (0.750)
<i>CONSTANT</i>	-1.071** (0.289)	-2.854** (0.252)	-3.430** (0.749)
Household characteristics			
<i>HHSIZE</i> [no.]	0.158* (0.068)	0.040 (0.034)	0.285* (0.125)
<i>FEM</i> [No.]	-0.009 (0.126)	-0.001 (0.002)	-0.104 (0.165)
<i>EDUS</i> [%]	0.005** (0.002)	0.004** (0.001)	-0.003 (0.004)
<i>PARTY</i> [1/0]	-0.061 (0.192)	-0.045 (0.131)	-0.304 (0.415)
<i>EMPLOY</i> [1/0]	-0.634* (0.333)	-0.069 (0.206)	-0.419 (0.469)
<i>MIGRANT</i> [1/0]	0.050 (0.176)	0.041 (0.102)	-0.110 (0.233)
Farm characteristics			
<i>LAND</i> [ha per capita]	-0.364 (0.893)	0.913* (0.459)	6.895* (2.874)
<i>LIVES</i> [kg]	-0.022 (0.015)	-0.003 (0.016)	-0.110* (0.049)
Village characteristics			
<i>ANIPP</i> [Yuan]	-0.010** (0.003)	0.064** (0.008)	0.043* (0.018)
<i>RRAL</i> [% of area]	0.012** (0.003)	0.001 (0.009)	0.034 (0.021)
Model diagnostics			
<i>H₀: all parms except const = 0</i>	<i>LR-statistic</i> 123.6	184.1	80.6
	<i>p-value</i> < 0.001	< 0.001	< 0.001
<i>Share of correct predictions¹</i>	60%	61%	31%
<i>State changes / Observations</i>	224/ 687	445/ 2535	238/ 1399
<i>H₀: Variance of frailty = 0</i>	<i>LR-statistic</i> 0.000	0.000	9.835
	<i>p-value</i> 0.50	0.50	<0.001

Standard errors in parentheses; * p<0.10, ** p<0.01

¹⁾ The dichotomous variable $\hat{\xi}$ 'predicted completion of participation state' was set to one if the predicted probability of changing the state ($\xi=1$) exceeded the relative frequency of status changes in the original sample and zero otherwise.

Table 3b: Estimation results of duration model PART-TIME TO FULL-TIME (stop off-farm work)

	ZHEJIANG	HUBEI	YUNNAN
<i>DUR</i> [ln(years)]	-0.847* (0.340)	-1.022** (0.094)	-1.071** (0.153)
<i>CONSTANT</i>	-0.678 (0.674)	-2.499** (0.259)	-0.354 (0.375)
Household characteristics			
<i>HHSIZE</i> [no.]	-0.443** (0.144)	-0.147** (0.046)	-0.337** (0.068)
<i>FEM</i> [No.]	0.035 (0.200)	-0.012 (0.099)	0.420** (0.118)
<i>EDUS</i> [%]	0.002 (0.003)	0.046 (0.064)	0.003 (0.003)
<i>PARTY</i> [1/0]	0.497 (0.325)	-0.500** (0.178)	-0.006 (0.288)
<i>EMPLOY</i> [1/0]	0.091 (0.532)	0.017 (0.226)	-0.009 (0.340)
<i>MIGRANT</i> [1/0]	0.324 (0.212)	0.213* (0.123)	0.119 (0.183)
Farm characteristics			
<i>LAND</i> [ha per capita]	-0.691 (1.5161)	0.593* (0.260)	-4.266** (1.283)
<i>LIVES</i> [kg]	0.002 (0.009)	-0.025 (0.036)	0.019 (0.031)
Village characteristics			
<i>ANIPP</i> [Yuan]	-0.008* (0.005)	0.065** (0.009)	-0.049** (0.011)
<i>RRAL</i> [% of area]	-0.001 (0.005)	-0.035** (0.012)	0.047* (0.018)
Model diagnostics			
<i>H₀: all parms except const = 0</i>	<i>LR-statistic</i> 118.6	200.2	150.3
	<i>p-value</i> < 0.001	< 0.001	< 0.001
<i>Share of correct predictions¹</i>	63%	64%	69%
<i>State changes / Observations</i>	121/ 2811	331/ 4613	143/ 2597
<i>H₀: Variance of frailty = 0</i>	<i>LR-statistic</i> 0.143	0.000	0.000
	<i>p-value</i> 0.353	0.50	0.50

Standard errors in parentheses; * p<0.10, ** p<0.01

¹⁾ The dichotomous variable $\hat{\xi}$ 'predicted completion of participation state' was set to one if the predicted probability of changing the state ($\xi=1$) exceeded the relative frequency of status changes in the original sample and zero otherwise.

In each of the models the null hypothesis that all coefficients except the constant are zero is clearly rejected by likelihood ratio tests. As an overall goodness-of-fit measure we present the models' power to predict completions ($\xi=1$) and non-completions ($\xi=0$) of spells. In order to map predicted hazards of completion (computed as predicted values from the fitted model) onto the dichotomous variable $\hat{\xi}$ we need to choose a critical limiting value. There is no unambiguously best choice as discussed by Greene (2003; 684f.) and we decided to use the relative frequency of completions in the respective original sample. The rates of successful predictions range between 60 and 70 percent except for the estimation of the move from full-time to part-time farming in Yunnan (31 percent). Aside from this latter estimation which was discarded for reasons discussed below, the predictive power of the models is satisfactory and no stringent differences between models or provinces are apparent. The parameter representing the variance of unobserved heterogeneity (*frailty*) was not found to be significant in any of the models (with the exception mentioned before: full-time to part-time in Yunnan) which is indicated by p-values between 0.35 and 0.5. Spell specific effects have obviously been captured sufficiently by controlling for the most influential covariates and hence the final models could be estimated without accounting for frailty.

The coefficients in the upper part of the tables show a diverse picture of the impact of duration and of the heterogeneity captured by the covariates. Five of the estimated coefficients of (the log of) duration (*DUR*) are significant at least at the 10 percent level and they are all negative. In the sixth model (full-time to part-time farming in Yunnan) the duration parameter is positive and numerically large but not significantly different from zero. The whole model suffers from a low predictive power as mentioned above. This and the significance of the frailty variance suggest that determinants which we could not include in our specification had a strong impact on the data generating process.¹⁵ Our specification

¹⁵ This result proved to be stable over a number of distributional assumptions regarding unobserved heterogeneity (gamma-, normal- distributed and mass-point frailty).

which has proved its capabilities in the case of the other five models does not yield reliable results with the data available to estimate the full-time to part-time model for Yunnan. We hence do not further interpret its results.

The interpretation of the coefficients on the *DUR*-variable is illustrated in **Figure 1** which shows the predicted hazard rates (hazards) at different durations based on the results of the parameter estimations. They indicate the probabilities that a household ends its participation status after a given duration conditional on having occupied this state up to that point. The predicted hazards are computed for a hypothetical household characterized by the mean value of each of the covariates considered in the estimations.¹⁶

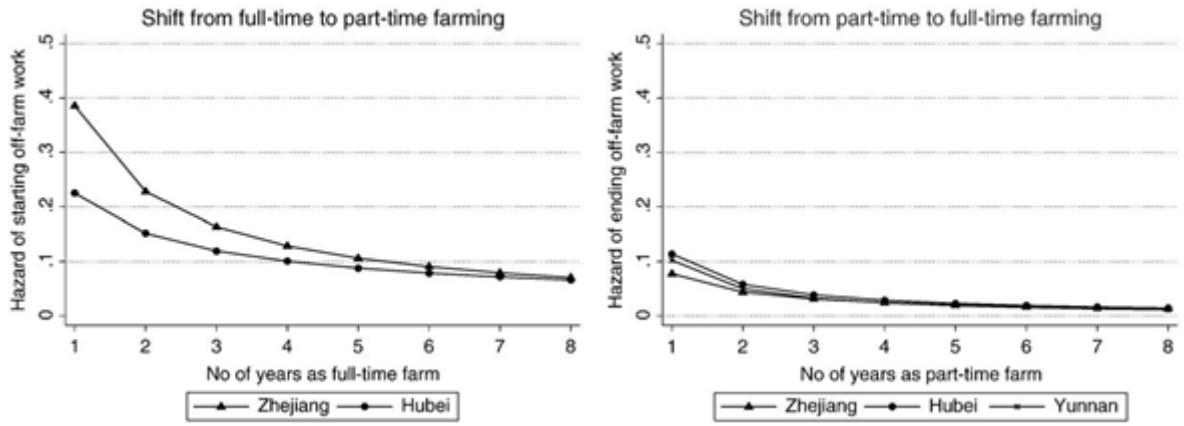


Figure 1: Predicted hazards (at means of covariates) for shifts between full- and part-time farming

All five hazard functions are decreasing which indicates negative duration dependence: Leaving a state (full-time or part-time farming) becomes less likely the longer that state has been already occupied.

This is particularly pronounced in the case of the movement from full-time to part-time farming. Among full-time farming households, which represent a comparatively small group,

¹⁶ Since the link function for parameter estimation is complementary log-log the predicted hazard rate (baseline hazard) $\hat{\lambda}_t^{ji}$ for a duration of t years is given by

$$\hat{\lambda}_t^{ji} = 1 - \exp \left(- \exp \left(\sum_{k \neq DUR} \bar{X}^k \beta^k + \ln(t) \beta^{DUR} \right) \right), \text{ with } \bar{X}^k \text{ being the mean of the } k\text{-th covariate.}$$

it is relatively likely that one of its members takes up non-agricultural work, i.e. starts supplying off-farm labor. After a single year of full-time farming this hazard is 39 percent in Zhejiang, and 23 percent in Hubei. However, it decreases to 23 and 15 percent after only one further year of non-supplying off-farm labor. The hazard then continues decreasing gradually over durations of up to eight years to 7 percent in both provinces. A possible explanation for this negative duration dependence is that some prerequisites of finding off-farm work such as specific skills and multiple social contacts decrease over the years. At the same time, specialization of the farmer and family members on farming as well as investments may enhance the technical and organizational structures over time and hence the economic efficiency of the farm. By way of a lock-in effect, off-farm work may become comparatively less attractive. It is worth noting that the decrease of the hazard is significantly faster in Zhejiang than in Hubei. Obviously, the more dynamic and developed economy of Zhejiang requires *and* rewards specialization and commitment in a path once chosen more quickly than in the more traditional economy of Hubei.

The movement from part-time to full-time farming is relatively rare in the three provinces considered. The hazards to give up off-farm work are comparatively low from the first year on (11, 10, and 8 percent in Hubei, Yunnan, and Zhejiang) and they decrease to negligible magnitudes over duration. However, predicted hazards towards long durations must be interpreted with care even if the duration dependence is statistically significant. Hazard rates for durations above five years are the weakly supported 'ends' of the hazard functions which have been fitted largely based on considerably shorter spells. The hazard of withdrawal from the off-farm labor market is lower in Zhejiang during the first years than in Hubei and Yunnan. Particularly in these latter provinces with less diversified economies off-farm involvement seems to be slightly less stable than in Zhejiang. Interpretation of the move from part-time to full-time farming is complicated by the fact that two different kinds of behavioral patterns may be underlying. It may either be that family members, who had been working off-farm before, decide to allocate their labor to farming. Another possibility is that such persons

leave the household, e.g. when establishing a new family. These two forms of return to full time farming can be assumed to have partly different determinants.

To sum up, **Figure 1** particularly indicates relatively high chances that members of full-time farming households are taking up off-farm-work, while the risk to stop this kind of involvement in the rural labor market is remarkably lower. Thus, part-time farming is a relatively stable occupational state.

Regarding the impact of covariates Table 3 shows that many of the household, farm, and village characteristics considered, significantly influence the hazard of leaving an occupied participation state. In the six models presented, between 2 and 6 (out of 10) covariates have significant coefficients. The covariates' signs indicate whether an increase in the respective variable increases or decreases the respective hazard rate. The relative change of the hazard rate associated with a one unit change of the respective covariate is given by the hazard ratio hr which, in our proportional specification, can be computed from the model parameter β as $hr = d\ln\lambda(z,\beta)/dz = \exp(\beta)$.

To give an example, the parameter value of 0.158 for household size ($HSIZE$) in the Zhejiang model '*Full-time to Part-time*' implies a hazard ratio of $\exp(0.158) = 1.171$. That means that households with one member more than the average have (for any given duration) a 17.1 percent higher hazard rate to switch from full-time to part time farming, i.e. to start supplying off-farm labor.¹⁷ For large households that do not participate in off-farm labor markets yet, it seems indeed very likely that at least one member starts to work off-farm, i.e. that spells of full-time farming are relatively short. The impact of household size on the probability of giving up off-farm work is – consequently – found to be negative. Focusing only on statistically significant coefficients we find one positive estimate for the variable indicating the number of women among laborers (FEM): in agriculturally dominated Yunnan province off-farm involvement is more likely to end in households comprising more females.

¹⁷ For dichotomous covariates (dummies) like e.g. *PARTY*, the impact of a 'one-unit change' means the hazard associated with party membership as opposed to the one for non-membership.

This is consistent with the notion that particularly females quit off-farm work when additional presence is required in the household or for child care. The educational level among full-time farming households' labor force (*EDUS*) is found to have a positive impact on the probability to start off-farm involvement in Zhejiang and Hubei. The small size of the absolute coefficient values (0.005 and 0.004) is due to measuring the share of educated laborers in percentage points. An increase of one percentage point increases the hazards by 0.53 and 0.41 percent respectively. The significance of the coefficients confirms findings of numerous earlier studies indentifying education as a key factor for income diversification of households. For the variable 'membership in Communist Party' (*PARTY*) we find a strong negative impact on the hazard to move from part-time to full-time farming (Hubei). Chances to take up and keep attractive off-farm occupations are obviously particularly high for families of party members. On the other hand this result and the insignificance of the other coefficients of this variable contrasts with earlier findings that party members are in a particularly favorable situation also for successful agricultural activities. For example, Knight and Yueh (2002) noted, access to land, water, and inputs may be relatively easy for CCP members.

Our specification includes two covariates characterizing households' labor allocation: the employment of hired labor and the work of household members as migrants in distant locations. We find that the hazard of the move from full-time to part-time farming in Zhejiang is significantly less likely if hired labor is employed (*EMPLOY*). The parameter of -0.634 implies a hazard ratio of 0.53 and hence a 47% decrease in the hazard. This contrasts with the expectation that household members may be released for off-farm occupation if hired labor is employed for farm work. The decreased chance of off-farm involvement may on the one hand be due to increased requirement for supervision and managerial work, on the other hand to size or specialization effects. The coefficient of the variable (*MIGRANT*) in the model 'part-time to full-time' in Hubei is 0.213. This indicates that households with any members working as migrants in distant locations have a 24 percent higher probability to stop local off-farm activities than households without any migrant member (Hubei). This result suggests a

substitutional relationship between local off-farm and migrant work. Local off-farm work is more likely to stop if any household members work as migrants.

With regard to the farm characteristics it turned out that the size of the households' agricultural land area per capita (*LAND*, i.e. arable plus orchard land) increases the probability to move from full-time to part-time farming in Hubei: the coefficient of 0.913 implies that households with one hectare of land per capita above average have more than twice the average hazard to start supplying off-farm labor. In Hubei, also the hazard of the movement in the opposite direction, i.e. towards full-time farming is positively affected to a significant extent by the land endowment. The farm size obviously increases flexibility to adjust in both directions the labor market participation state. However, Yunnan has a negative coefficient indicating that in this province the hazard to give up local off-farm work is lower for households with larger land area. In contrast to these results on the farms' land area and its impact on labor market participation decisions we did not identify significant impact of labor intensive livestock production measured by output quantities of major livestock products (*LIVES*).

The two village characteristics considered were found to be significantly influential for households' participation decisions: The estimates for the village income level (*ANIPP*) differ between provinces for both directions of state change: In Zhejiang, households in more wealthy villages move less flexibly between participation states. Also in Yunnan the hazard of giving up local off-farm work decreases with the village income level. In Hubei the opposite holds for both directions of move. Here relatively wealthy villages obviously provide a favorable environment for any of the considered changes of participation state. The relationship between wealth of the village and flexibility of labor market participation may be of non-linear nature. Finally, significant parameters of either sign were found for the percentage of land area leased by the farmers in a village rather than owned (*RRAL*). In villages with a high share of leased land taking up off-farm work is relatively likely (Zhejiang). This conforms to expectations that vivid markets for leased land indicate easy

land transfer and hence the possibility that some farms grow and become full-time farms while other households give up farming. The positive coefficient for the movement from part-time to full-time farming in Yunnan is also consistent with this interpretation. However, the coefficient for Hubei is negative which contradicts this interpretation: in villages with high land lease activity the probability to move from part-time to full-time farming is relatively low.

To sum up, the coefficients of most of the covariates give a heterogeneous picture regarding the determinants rural households' hazards to move between labor market participation states. The differences between provinces do not show a stringent pattern and do not suggest straightforward explanations. However, some of the results allow conclusions which are relevant for understanding household behavior and draw policy conclusions.

5 Summary and conclusions

The study contributes to the ongoing debate over the participation of agricultural households in rural labor markets in China. The choice between full-time and part-time farming is empirically analyzed. Using household data of the period 1995 - 2002 from the provinces Zhejiang, Hubei, and Yunnan we apply a discrete time hazard approach to analyze households' labor market participation histories that is the frequency of transitions between full-time and part-time farming.

Results suggest that the likelihood of moving from full-time to part-time farming is relatively high and shows significant negative duration dependence. This implies decreasing chances of taking up off-farm work with an increasing time that a household confines itself to full-time farming. The shift from part-time into full-time farming occurs relatively rarely and its probability also decreases over duration. The negative duration dependence of shifts in both directions may be a consequence of accumulated human or physical capital that makes any chosen state (full-time or part-time agriculture) more advantageous after a couple of years. At the same time, remaining in one state may also reduce the household members'

flexibility and chances to change their participation status even if conditions have changed in a way to favor such transition. Skills needed in the other state of participation may have become obsolete or forgotten due to lack of practice.

Finally, we find that several household, farm and regional characteristics are significantly related to the hazard of moving between labor market participation states. In particular we find the household size, the endowment with agricultural land, and the income level of the village to be influential in most of the models. Education of households' labor force has a significantly positive impact on the likelihood for full-time farming households to engage in off-farm work. In some cases the pervasiveness of land lease in the villages has a positive impact on the flexibility of households to choose between participation states. Given that factors like public availability of education and the ease of land transfer between households can be influenced by policy, the government can strengthen the tendency of economic diversification by setting framework conditions that encourage households to take up off-farm work.

A caveat remains that motivates further studies. It is a relatively crude measure for off-farm labor market participation to look at a dichotomous variable for household participation. Even if considering individuals rather than households is precluded by lack of data, an interesting extension could be to look at the *intensity* of participation as measured by the number of days or its contribution to income or the *type* of off-farm occupation, wage work versus self-employment.

The relatively high and increasing prevalence of farm households' labor market participation shows a rural development trend that could in medium term contribute to reducing rural poverty. Frequent changes between participation states, in particular the movement into off-farm labor market participation, are an indicator for widespread and relatively frictionless labor markets access. Low frequencies of movements between participation states are on the other hand not necessarily an indication of harmful rigidities. In

particular, if hazards of change decrease with duration this can also be an indication that households specialize and accumulate capital which is part of economic development.

From a policy perspective the setting of framework conditions that enhance households' chances to adapt flexibly to changing economic situations should be given high priority. This will enhance rural labor mobility as a key to further economic development. The analysis gives several suggestions here. A variable susceptible to policy influence, the activity of land rental markets, is highly dependent on the legal framework. By improving this framework households' labor flexibility could be increased significantly. Education is another key factor which increases farm households' probability to diversify their income generating activities. For government investment in education to be most effective it could be designed in a way to account for duration dependence of households' labor flexibility. If e.g. educational support to rural households would be targeted via scholarships to households with recent non-farm work histories, the impact on future off-farm participation would be larger than if long-term full-time farming households with very low probabilities of leaving this state would be included.

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Tables

Table 1: Household income and farming activity by province, 1995-2002

	Zhejiang	Hubei	Yunnan
Households in sample holding arable land [%]	79	97	93
Per capita net income rural households, 2002 [1000 RMB] ¹⁸	4,3	2,3	1,5
Income share of household business in total income (2003) ^{19 20} [%]	43%	70%	73%
Average land endowment of household [ha]	0.27	0.36	0.52
Most common size class (arable land + orchard)	0.1-0.5 ha		
Share of farms with 0.1-0.5 ha	76%	77%	56%
Volume of annual production [1000 RMB]	5.9	6.9	6.4
within this: share of home consumption	50%	74%	71%
Main products (% in value of production)	Fruit & Veg. (25) Rice (22), Pork (22)	Fruit & Veg. (32) Rice (23) Pork (15)	Pork (21) Corn (20) Other grain (16)

Source: own computations based on RCRE data.

¹⁸ Statistical Yearbook 2004, Tab 10-20/ 10-22

¹⁹ Household business income originates to a large extent (40%) from agricultural activities.

²⁰ Value of production in prices of 2000. Mean (deflated) unit values of *sales* over the 1995-2002 period were used in order to have a sufficiently large data base to compute average prices for each product.

Table 2: Household, farm, and village characteristics by labor market participation state*

		All holds (N=14739)	house- Part-time farms (N=10033)	Full-time farms (N=4706)	Sig. diff.
Household characteristics					
Household size [persons]	<i>HSIZE</i>	4.14 (1.45)	4.26 (1.42)	3.87 (1.45)	**
No. of females among family laborers	<i>FEM</i>	1.24 (0.71)	1.26 (0.70)	1.19 (0.74)	**
Secondary school graduates among laborers [percent]	<i>EDUS</i>	38 (35)	40 (35)	34 (34)	**
Communist party member [1/0]: % of households	<i>PARTY</i>	0.15 (0.36)	0.16 (0.37)	0.15 (0.36)	
Employing hired labor [1/0]: percent of households	<i>EMPLOY</i>	0.07 (0.26)	0.07 (0.26)	0.07 (0.26)	
Any members working as migrants [1/0]: % of households	<i>MIGRAN</i>	0.40 (0.49)	0.60 (0.49)	0.31 (0.46)	**
Farm Characteristics					
Agricultural land [ha per capita]	<i>LAND</i>	0.10 (0.10)	0.10 (0.11)	0.09 (0.09)	**
Animal production, weighted output [kg/100]	<i>LIVES</i>	1.71 (6.23)	1.60 (7.02)	1.93 (4.75)	**
Village characteristics					
Annual p.c. income [Yuan in 2000 prices]	<i>ANIPP</i>	2863 (3114)	2940 (3192)	2599 (2816)	**
Share of leased land on arable land	<i>RRAL</i>	0.08 (0.14)	0.08 (0.15)	0.06 (0.11)	**

*Standard deviations in parentheses

** Distribution differs significantly between full-time and part-time farms (see footnote 7)

Source: own computations based on RCRE rural household panel.

Table 3a: Estimation results of duration model: FULL-TIME TO PART-TIME (taking up off-farm work)

	ZHEJIANG	HUBEI	YUNNAN
<i>DUR</i> [ln(years)]	-0.912** (0.176)	-0.633** (0.090)	1.212 (0.750)
<i>CONSTANT</i>	-1.071** (0.289)	-2.854** (0.252)	-3.430** (0.749)
Household characteristics			
<i>HHSIZE</i> [no.]	0.158* (0.068)	0.040 (0.034)	0.285* (0.125)
<i>FEM</i> [No.]	-0.009 (0.126)	-0.001 (0.002)	-0.104 (0.165)
<i>EDUS</i> [%]	0.005** (0.002)	0.004** (0.001)	-0.003 (0.004)
<i>PARTY</i> [1/0]	-0.061 (0.192)	-0.045 (0.131)	-0.304 (0.415)
<i>EMPLOY</i> [1/0]	-0.634* (0.333)	-0.069 (0.206)	-0.419 (0.469)
<i>MIGRANT</i> [1/0]	0.050 (0.176)	0.041 (0.102)	-0.110 (0.233)
Farm characteristics			
<i>LAND</i> [ha per capita]	-0.364 (0.893)	0.913* (0.459)	6.895* (2.874)
<i>LIVES</i> [kg]	-0.022 (0.015)	-0.003 (0.016)	-0.110* (0.049)
Village characteristics			
<i>ANIPP</i> [Yuan]	-0.010** (0.003)	0.064** (0.008)	0.043* (0.018)
<i>RRAL</i> [% of area]	0.012** (0.003)	0.001 (0.009)	0.034 (0.021)
Model diagnostics			
<i>H₀: all parms except const = 0</i>	<i>LR-statistic</i> 123.6	184.1	80.6
	<i>p-value</i> < 0.001	< 0.001	< 0.001
<i>Share of correct predictions¹</i>	60%	61%	31%
<i>State changes / Observations</i>	224/ 687	445/ 2535	238/ 1399
<i>H₀: Variance of frailty = 0</i>	<i>LR-statistic</i> 0.000	0.000	9.835
	<i>p-value</i> 0.50	0.50	<0.001

Standard errors in parentheses; * p<0.10, ** p<0.01

¹⁾ The dichotomous variable $\hat{\xi}$ 'predicted completion of participation state' was set to one if the predicted probability of changing the state ($\xi=1$) exceeded the relative frequency of status changes in the original sample and zero otherwise.

Table 3b: Estimation results of duration model PART-TIME TO FULL-TIME (stop off-farm work)

	ZHEJIANG	HUBEI	YUNNAN
<i>DUR</i> [ln(years)]	-0.847* (0.340)	-1.022** (0.094)	-1.071** (0.153)
<i>CONSTANT</i>	-0.678 (0.674)	-2.499** (0.259)	-0.354 (0.375)
Household characteristics			
<i>HHSIZE</i> [no.]	-0.443** (0.144)	-0.147** (0.046)	-0.337** (0.068)
<i>FEM</i> [No.]	0.035 (0.200)	-0.012 (0.099)	0.420** (0.118)
<i>EDUS</i> [%]	0.002 (0.003)	0.046 (0.064)	0.003 (0.003)
<i>PARTY</i> [1/0]	0.497 (0.325)	-0.500** (0.178)	-0.006 (0.288)
<i>EMPLOY</i> [1/0]	0.091 (0.532)	0.017 (0.226)	-0.009 (0.340)
<i>MIGRANT</i> [1/0]	0.324 (0.212)	0.213* (0.123)	0.119 (0.183)
Farm characteristics			
<i>LAND</i> [ha per capita]	-0.691 (1.5161)	0.593* (0.260)	-4.266** (1.283)
<i>LIVES</i> [kg]	0.002 (0.009)	-0.025 (0.036)	0.019 (0.031)
Village characteristics			
<i>ANIPP</i> [Yuan]	-0.008* (0.005)	0.065** (0.009)	-0.049** (0.011)
<i>RRAL</i> [% of area]	-0.001 (0.005)	-0.035** (0.012)	0.047* (0.018)
Model diagnostics			
<i>H₀: all parms except const = 0</i>			
<i>LR-statistic</i>	118.6	200.2	150.3
<i>p-value</i>	< 0.001	< 0.001	< 0.001
<i>Share of correct predictions¹</i>	63%	64%	69%
<i>State changes / Observations</i>	121/ 2811	331/ 4613	143/ 2597
<i>H₀: Variance of frailty = 0</i>			
<i>LR-statistic</i>	0.143	0.000	0.000
<i>p-value</i>	0.353	0.50	0.50

Standard errors in parentheses; * p<0.10, ** p<0.01

¹⁾ The dichotomous variable $\hat{\xi}$ 'predicted completion of participation state' was set to one if the predicted probability of changing the state ($\xi=1$) exceeded the relative frequency of status changes in the original sample and zero otherwise.

Figures

Figure 1: Predicted hazards (at means of covariates) for shifts between full- and part-time farming

