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Determinants of off-farm work and temporary migration in China

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Abstract Existing research inadequately explains the factors that drive temporary internal migration in China. Using data for 2005 drawn from 1,903 households in 43 rural villages, we calculate binomial and multinomial logit (BL, MNL) models of probabilities that an adult belongs to one of three categories of worker—on-farm, off-farm, or temporary migrant—as a function of individual and household characteristics. We control for village xed effects, paying close attention to male/female differences. Nearly all coef cients—even for village dummies—vary signi cantly by sex. For two variables—age and schooling—the relationships are non-linear. There is an optimal age and amount of schooling that maximizes the probability that a worker will be employed away from the family farm. For schooling, this is low, suggesting that educated workers are underemployed. This might indicate that schooling beyond primary grades is poor quality, or at least inappropriate for the job market.

Keywords China Education Rural labor force Migration Hukou Floating population

Introduction

Internal migration drives urbanization; therefore, an understanding of what drives migration in China is necessary if we are to understand the urbanization process

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and, ultimately, the environmental changes that accompany it. It is the purpose of this paper to contribute to this first stage of understanding, namely the migration decision. It is important to emphasize from the outset that we are analyzing temporary migration rather than permanent settlement, but similar forces might be expected to drive both types of migration.

China's household registration system (*hukou*) was modeled after the Soviet *propiska* (internal passport) system and continues to constrain internal migration. In recent years, the central government has devolved responsibility for *hukou* policies to local governments, but it is still difficult if not impossible for peasants to qualify for permanent residency rights and associated social benefits, such as free access to urban jobs, and to government services such as health care, pensions, and public schooling. Nearly all rural migrants are non-*hukou*, that is, legally they are "temporary" migrants, even though they may have lived and worked in the given destination for years (for details, see Chan 2009).

There is a large and growing literature on the determinants of migration in China, but we still lack a clear picture of what is driving temporary migration from rural areas. To some extent, this reflects spotty data, but it is also a product of poor research design. More than a decade ago, Yang and Guo (1999: 930) complained that researchers pay scant attention to community factors and even less "to the role of gender in temporary migration and to possible differences in the determinants of temporary migration between men and women." Sadly, this is still true today. What has changed in the last decade or so is researchers' increasing use of the sophisticated multinomial logit (MNL) model instead of the previously popular binary logit (BL) model. Zhao (1997, 1999) was the first to apply the MNL model to Chinese migration data. Knight and Song (2003) followed, then Xia and Simmons (2004), Liu (2008), Chen and Hamori (2009), Démurger et al. (2009), Knight et al. (2010), and Wu (2010). Shi et al. (2007) estimate a multinomial *probit* model that resembles the MNL model, but is computationally more difficult. In reviewing this work, what struck us most was the lack of attention to gender differences. Nine of these ten studies include a gender dummy (male = 1, female = 0, or the reverse[1]), but none allow interaction between gender and other explanatory variables. Attention to community factors is often better. Wu (2010), for example, controls for fixed effects of all 33 villages covered in his survey data. Chen and Hamori (2009), in contrast, completely ignore community-specific effects in their sample, which was drawn from 288 villages spread over nine provinces. Chen and Hamori do include a dummy for region (residents of four provinces = 1, residents of any of the other five provinces = 0), but rural residents spread over an enormous region are not members of a community in any meaningful sense.

We "go with the flow" of this research and use available data to estimate a multinomial choice model. This promises more accurate estimates of the effect of explanatory variables on the probability of migrating. Equally important, it allows us to measure the determinants of working locally but off-farm, a way to increase the incomes of rural families without moving them to urban areas. Following Zhao

[1] The exception is Xia and Simmons (2004), who in lieu of a single gender dummy, include three dummy variables: Single male, single female, and married male, but no interaction terms.

(1999) and Liu (2008), we also estimate a binary choice model: A person either chooses to migrate ($P = 1$) or chooses not to ($P = 0$). We do this in part because BL coefficients are easier to interpret, and in part to discover what a complex MNL model adds to the simpler BL model.

We move beyond the existing literature in part by analysing a large and unique set of data. Most importantly, however, we control for fixed effects of villages and pay close attention to differences between men and women in the determinants of where they work: on the family farm, off-farm in the township, or farther away as a non-*hokou* migrants. We allow for fixed effects with a gender dummy, but we also test for differences by gender in the coefficients of *all* explanatory variables, allowing each coefficient to vary by gender when appropriate.

Description of the data

The Research Centre for Rural Economy (RCRE) at China's State Council has carried out surveys of rural households for more than 20 years. As of 2005, the RCRE surveyed 24,000 households in 31 provinces, autonomous regions, and cities (Gu 2005). The RCRE household survey focuses on land use and characteristics of households, but in recent years has added questions on migration, schooling, and other information related to individual household members.

We use RCRE survey data for the year 2005, which recently became available for 43 villages in three provinces: Shandong, Zhejiang, and Jilin. Shandong and Zhejiang are thriving provinces located in the eastern coastal region of the country. Landlocked Jilin is less prosperous; it is located in the northeastern part of the country, bordering North Korea in the southeast and Russia in the east (see Fig. 1). According to the National Bureau of Statistics (2006), Zhejiang had the third largest net income per capita (6,660 yuan) of China's 31 provinces. Shandong had the eighth highest income (3,931 yuan), while Jilin ranked eleventh (3,264 yuan). Jilin was the most agrarian of the three provinces, with 73% of its labor force engaged in agriculture compared to 54% in Shandong and only 34% in Zhejiang.

The initial data set contained information for 2,020 households: 520 from Shandong, 501 from Zhejiang, and 999 from Jilin. We deleted a small number of households because data were missing for one or more adult members. Other households were removed because they contained not a single adult between the ages of 16 and 59, which are prime working years for rural Chinese. The full data set after "cleaning" consists of 1,903 households: 485 from Shandong province, 458 from Zhejiang province, and 962 from Jilin province. These households contain 5,588 persons aged 16–59, all of whom legally reside in one of 43 rural villages, although 996 of them migrate at least part of the year to jobs outside their township of residence. One-third of the migrants—332 to be precise—are women. As Table 1 indicates, men have a greater propensity than women to work away from home. The total difference of 21% points is spread almost equally between working off-farm and migrating. Women, with 27% working off-farm, are 10% points behind men and, with a 12% migration rate, are 11% points behind the more mobile men.

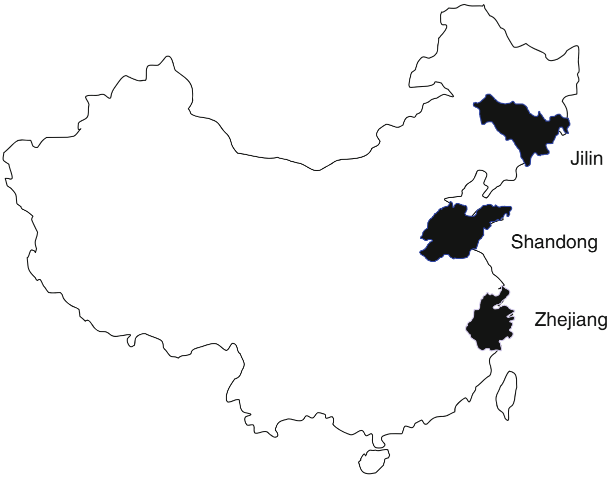

Fig. 1 The three survey provinces

Table 1 Median age and years of schooling, by work location and gender

	Location (%)		Median age		Median schooling	
	Male	Female	Male	Female	Male	Female
On-farm	40	61	38	42	8	6
Off-farm	37	27	43	41	8	6
Migrant	23	12	30	25	9	9
Total	100	100	39	40	8	7

Compiled by the authors from the full sample of 1,903 households, comprising 5,588 persons aged 16–59 years. The sample is drawn from the Ministry of Agriculture's 2005 Rural Household Survey of 43 rural villages in three provinces: Shandong (13 villages), Zhejiang (10 villages), and Jilin (20 villages)

Some personal characteristics of the workers are also presented in Table 1. The typical migrant man is 30 years old, 13 years younger than his off-farm, non-migrant counterpart, and 8 years younger than men who work only at home. Migrant women are even younger (25 years at the median). A typical migrant man has as much schooling as a migrant woman (9 years, which represents completion of compulsory education), but this is an advantage of only 1 year over a non-migrant man. In contrast, the median schooling of migrant women exceeds that of non-migrants by 3 years.

A model of multinomial choice

We t a multinomial logit (MNL) model to our data in order to estimate the probability that a rural resident aged 16–59 years is in one of three categories of workers: On-farm, off-farm in the township, or temporary migrant. The dependent variable is coded 0 for persons aged 16–59 years who work exclusively at home on the family farm, 1 for those employed off-farm in the township of residence, and 2 for temporary migrants. We designate the probability of each of these three events as P_0, P_1, and P_2, respectively. Any logit model ensures that each estimated P lies within the bounds of zero and unity. Negative probabilities and probabilities greater than one are impossible by design. The multinomial logit (MNL) model also ensures that the relevant probabilities (in our case, P_0, P_1, and P_2) sum to unity.

In the MNL model, one possibility—on-farm work—is denoted as the base or reference position. The logarithm of the odds (relative to the base) of each remaining response is assumed to follow a linear model:

$$\ln(P_1/P_0) = z_1 = \sum \beta_{1i} x_i \tag{1a}$$

$$\ln(P_2/P_0) = z_2 = \sum \beta_{2i} x_i \tag{2a}$$

where β_{1i} and β_{2i} are coefficients of the ith explanatory variable, x_i. The two equations are estimated simultaneously. There is no need to estimate a third equation, as the missing comparison between P_1 and P_2 can be obtained from the fact that $\ln(P_1/P_2) = \ln(P_1/P_0) - \ln(P_2/P_0)$.

Additional equations can be added to accommodate four, five, or more responses. When there are two responses, the model reduces to a binary logit (BL) model, leaving only one equation. In our notation, the equation for migration would be $\ln\{(P_2/(P_0 + P_1)\}$ equivalent to $\ln(P_2/(1 - P_2))$. Logarithms of odds are known as "logits", and hence the name "logit regression". Odds ratios have no upper limit, but they do have a lower bound of zero. Logits can have any value, positive or negative. A curve that is linear for the logit of P—easily estimated using linear regression techniques—is non-linear in P, taking a familiar S-shaped logistic curve that approaches, yet never reaches values of zero and unity. The slope of the logistic curve is steepest (marginal effects are greatest) at the point of inflection, where the odds are equal and $P = 1/2$. Non-linearities can be accommodated by adding squared terms to the list of explanatory variables; in this case, the curve for P will be U-shaped or inverse U-shaped with the tails of the U (inverted U) approaching, but never reaching the upper (lower) bound of unity (zero).

Maximum likelihood estimation (MLE) is used in lieu of ordinary least squares (OLS) for both binary and multinomial logit regression. MLE is an iterative procedure that produces results with excellent large sample properties. The technique is straightforward and intuitive, but unusual in that none of the observed values of P lie on the logistic curve. Moreover, the logit of any observed P is either negative infinity or positive infinity, neither of which is an actual number and therefore does not lie on the logit curve, either.

The MNL model may also be written in terms of probabilities (P's) rather than odds ratios. Exponentiating Eqs. 1a and 2a above yields $P_1 = P_0 * \exp(z_1)$ and $P_2 = P_0 * \exp(z_2)$. Considering that $P_0 + P_1 + P_2 = 1$, we know that the base probability $(P_0) = 1/\{1 + \exp(z_1) + \exp(z_2)\}$; the other two probabilities are

$$P_1 = \exp(z_1)/\{1 + \exp(z_1) + \exp(z_2)\} \tag{1b}$$

$$P_2 = \exp(z_2)/\{1 + \exp(z_1) + \exp(z_2)\} \tag{2b}$$

This way of writing the MNL model elucidates that choices are determined simultaneously, with the determinants of one affecting the determinants of the other. It is moreover helpful to note that the regression coefficients (β_{1i} and β_{2i}) measure effects relative to the base (working on-farm), since all coefficients of the base equation (the β_{0i}) are equal to zero by definition.

The explanatory variables

We "explain" the probability of a worker migrating or working off-farm as a function of four characteristics of individuals, three characteristics of households, and the fixed effects of the village of his or her official residence. In addition to the fixed effect of gender, captured by two dummy variables, we allow for interaction effects between gender and *all* of the other explanatory variables, estimating separate coefficients for men and women whenever these interaction effects are statistically significant.

The three variables (in addition to gender) that measure personal characteristics of individuals comprise AGE, SCHOOL, and HEAD (see Table 2). Researchers often find that age either has a positive effect or is insignificant for off-farm work in China, while it has a negative effect on migration (e.g., Zhao 1999; Shi et al. 2007; Knight et al. 2010). This is expected because older people might prefer to work close to home since they have fewer years in which to recover the fixed costs of migration and because those costs—as perceived by the individual—increase with age. Schooling is generally expected to promote job mobility and migration, but this is not always reflected in Chinese data. Zhao (1999), Shi et al. (2007), and Démurger et al. (2009), for example, found weak effects for formal education on migration, but strong effects for shifting from rural farm to local non-farm work.

Table 2 Variables and descriptive statistics (mean values for men and for women)

Variable		Full sample: Males	Full sample: Females	Reduced sample: Males	Reduced sample: Females
Age	Sample limited to working ages, 16–59	37.49	38.11	37.29	38.05
School	Schooling completed, 0–18 years	7.82	6.74	7.81	6.73
Head	Equals 1 if head of household	0.57	0.01	0.56	0.01
Land	Arable land—mu (1/15 h.) per hh member	3.31	3.12	3.18	2.98
Child	Equals 1 if household has child <5 years	0.12	0.14	0.13	0.15
Dependency	Ratio of old + young to adults aged 16–59	0.24	0.29	0.25	0.29
	Sample size (individuals)	2855	2733	2577	2519
	Number of villages	43	43	38	38

For reasons explained in the text, five villages with extreme values were deleted, resulting in a smaller but more representative sample of villages

A quadratic term is added for AGE and for SCHOOL to test for nonlinearities and to allow for effects to be initially positive, then negative or vice versa.

HEAD is a dummy variable that takes the value of unity if a person is the head of his or her rural household. Other things equal, this responsibility would make migration more difficult (Stark and Taylor 1991). This variable is especially appropriate for China since workers seldom migrate with their families because of the discriminatory *hukou* system. Nonetheless, this variable is rarely taken into account in Chinese migration studies. Knight and Song (2003) included HEAD in their regressions and found it to have a negative impact on migration, but not on off-farm work in the village. More than half the men in our sample are the heads of their household. Few women head a household; they number only 40, five of whom are migrants[2] (see Table 2 once again).

Three variables refer to characteristics of each worker's household: LAND, CHILD, and DEPENDENCY. LAND refers to the total amount of land cultivated per household member at the beginning of the year 2005. This comprises the amount of land allocated to the household under the Household Responsibility System *plus* any land rented temporarily from other households *minus* any land temporarily rented to other farmers. The average amount of land under cultivation is about one-fifth of a hectare, and the maximum amount in our sample is only seven hectares (105 *mu*). A negative coefficient is expected for this variable, both for off-farm work and for migration (Zhao 1999; Liu 2008).

CHILD is a dummy variable that equals 1 if an individual's official residence is in a household with a child younger than 5 years of age. Only 12% of the men and 14% of the women in our sample are members of households with children this young. Any coefficient is possible for CHILD (Zhao 1999; Yang and Guo 1999; Shi et al. 2007). On the one hand, having responsibility for young children encourages generation of income from off-farm jobs or migration. On the other hand, women in particular are needed to care for young children, unless grandparents can be entrusted with their care.

DEPENDENCY refers to the number of persons in a household who are 60 years of age and older *plus* children aged 0–15 years divided by the number of household members of working age (16–59). Residency depends solely on *hukou* status, so temporary migrants are counted as part of the household. The expected sign for the coefficient of DEPENDENCY is indeterminate, since more dependents implies a need for higher incomes, but may also imply the need for more time to care for them, thus less opportunity to work off-farm or migrate (Shi et al. 2007).

Our database contains information on household income and productive assets, but neither variable was significant in any of the regression equations and consequently, both were dropped. Démurger et al. (2009) report a significant positive effect of household wealth on local off-farm employment, but not on migration. Other researchers report significant effects of household income or wealth on the probability of migration, varying from negative (Liu 2008) to positive

[2] The careful reader might calculate that this leaves more than 200 households without a head. The heads are lacking only because we excluded all individuals 60 years of age and older from our sample. Many of these excluded individuals head a household.

(Chen and Hamori 2009) to an “inverted U” peaking just above the poverty level of income (Du et al. 2005). We were unable to discern any linear or non-linear relationship in our sample between household income or assets and the probability of working off-farm or migrating.

Finally, and most important, are village characteristics. There are numerous reasons to expect villages to have an independent effect on migration, unrelated to the characteristics of residents or households. Villages differ in wage levels, the availability of off-farm jobs, access to paved highways and railways, communication (radio, TV, telephone), and other amenities, all of which might be expected to impact positively on off-farm work and negatively on migration. The existence of village-based networks of migrants, in contrast, can facilitate migration by providing information on employment and living conditions in migrant destinations.

We have no information on the rural villages in our sample, but we do know to which of the 43 villages each household belongs. This allows us to add dummy variables to control for fixed effects that are unobserved, but vary from village to village. To avoid perfect multicollinearity and a singular matrix, the fixed effect of one village is set to zero; the fixed effects of the other 42 villages are measured relative to the fixed effect of this arbitrarily selected village. There are substantial differences between villages in the propensity of their residents to migrate and to work off-farm. The share of migrants in the villages of our sample ranges from 2.3 to 47% of the working-age population, and the share engaged in local, off-farm work varies even more, from zero to 91%.

Empirical results for the MNL model

The maximum likelihood estimates of the parameters for our MNL model of off-farm work and temporary migration are depicted in Table 3. The estimation of the model as a whole is highly significant, and most of the coefficients are statistically different from zero at the 1% level. The likelihood ratio test is a test of the joint significance of all coefficients, except for those of the four gender dummies. Regressions (not reported) to test for female/male differences in the coefficients for individual/household variables reveal, with one exception, that the differences are statistically significant. The exception is DEPENDENCY, where differences between the male and the female coefficients are small and statistically insignificant.

The MNL regression results for the full sample are very satisfactory, although there is one problem. The range of coefficients on the village dummies is extremely large, especially for women (both off-farm and temporary migration), but also for men (only for off-farm). An examination of the data reveals that this is due to five outlier villages: Two in Jilin, two in Shandong, and one in Zhejiang (see Table 4). The first village has no off-farm workers and no female migrants. The next two villages report no female off-farm workers, and the fourth reports no female migrants. In the fifth village, half the female workers are employed off-farm and the other half as migrants. This might reflect coding errors, unrepresentative samples, or the true condition of these villages. Regardless of the reason, the data for these five villages differ sharply from those of the other 38 villages, so we removed them from

Table 3 Full sample: MNL results (robust t statistics in brackets)

Variable	Off-farm work		Temporary migration	
	Male	Female	Male	Female
Gender dummy	−8.493***	−6.821***	−6.070***	−3.136***
	[−11.46]	[−9.08]	[−8.83]	[−4.07]
Age	0.368***	0.281***	0.323***	0.141***
	[10.96]	[8.81]	[9.62]	[3.22]
Age^2/100	−0.446***	−0.379***	−0.430***	−0.293***
	[−10.73]	[−9.07]	[−9.78]	[−4.48]
School	0.338***	0.273***	0.354***	0.140
	[4.47]	[4.53]	[4.26]	[1.61]
School^2/100	−2.362***	−2.404***	−2.271***	−0.577
	[−4.88]	[−5.38]	[−4.32]	[1.06]
Head	0.105		−0.980***	
	[0.59]		[−4.80]	
Land	−0.100***	−0.093***	−0.117***	−0.045
	[−3.65]	[−2.93]	[−4.35]	[−1.21]
Child		−0.392**		−0.655***
		[−2.48]		[−3.08]
Dependency	−0.045	−0.045	−0.523***	−0.523***
	[−0.47]	[−0.47]	[−3.78]	[−3.78]
Village fixed effects (#)	43	43	43	43
Average effect	0.130	−0.540	−0.088	−0.298
Maximum	4.129	28.873	2.669	29.122
Minimum	−27.466	−37.030	−1.989	−25.122
Observations				5588
Log-likelihood (gender dummies)				−5558.45
Log-likelihood (all variables)				−4178.19
Likelihood ratio test				2760.52***

The symbols *, ** and *** denote statistical significance at the 10, 5, and 1% levels in two-tailed tests. The fixed effects of one village are set at zero to avoid perfect multicollinearity with the two gender dummies. There is a single coefficient (male = female) for DEPENDENCY, because coefficients left free to vary by gender were nearly identical in repeated regressions

Bold values refer to the two coefficients constrained to be equal for men and women, and to results for the complete set of 4 MNL equations

the sample. In the reduced sample, the lower bound of the range for the share of migrants in each village increases slightly from 2.3 to 2.7%, and the lower bound for the share in off-farm work increases from zero to 3.6%. The upper bounds remain unchanged at 47 and 91%, respectively.

Table 5 presents the results for the reduced sample. Surprisingly, removal of the 492 observations from the five outlier villages only affected the coefficients for the individual/household variables slightly. The ranges for the fixed village effects are now more reasonable and similar for both men and women. Not all of the female/

Table 4 Outlier villages in the full sample

	Off-farm work		Temporary migration	
	Male	Female	Male	Female
Fixed effects of				
Village 1 (Jilin)	−27.47	−37.03	−1.30	−25.74
Village 2 (Jilin)	−2.25	−27.24	−0.06	1.41
Village 3 (Shandong)	−2.08	−28.87	−1.65	−1.19
Village 4 (Shandong)	1.96	0.90	−1.15	−19.43
Village 5 (Zhejiang)	−0.01	28.87	−0.13	29.12
Destination of workers (%)				
Village 1 (Jilin)	0.00	0.00	8.20	0.00
Village 2 (Jilin)	2.63	0.00	34.21	37.10
Village 3 (Shandong)	61.22	0.00	10.20	4.88
Village 4 (Shandong)	72.73	29.55	4.55	0.00
Village 5 (Zhejiang)	29.17	50.00	22.92	50.00

Sample data and fixed effect estimates from the MNL model of Table 2

Bold values refer to the outliers (extreme values) for village dummies

male differences in coefficients were statistically significant. In fact, only 8 of the 74 coefficients differ significantly at the 5% level, but that is twice the number that might be expected by chance. In any case, it is best to accept or reject the entire set of gender*village interaction variables as a block. The likelihood ratio test for the joint significance of all female/male differences is 169.1, which is larger than 105.2, the critical .01 value of chi-square with 74 degrees of freedom. Fixed village effects thus differ significantly between males and females, even though the effects on average are nearly the same. As for estimated coefficients of the village dummies, at the 5% level, 36 are significant for men and 27 are significant for women.

In their study of temporary migration from 32 rural villages in Hubei province, Yang and Guo (1999) raise the interesting possibility that men might be more responsive than women to the effects of community level factors. The proof they offer in support of the hypothesis is unfortunately not compelling, because they test for statistical significance rather than quantitative importance.[3] What evidence is there in our own regression results that village effects are stronger for men than for women? None, we would argue. The village fixed effects have a smaller range for men than for women (4.6 vs. 5.1) with equal dispersion (standard deviation = 1.12). It is true that more of the village dummies are significant for men than for women,

[3] Yang and Guo (1999) estimate separate BL regressions for men and for women. With only individual/household variables, the pseudo R^2 is 0.0266 for the men's regressions and 0.1601 for the women's. Adding four village variables (distance to a city, per capita income, population density, population growth) raises the pseudo R^2 to 0.0856 for men and 0.1997 for women. The increase is greater for men than for women, but this suggests only that the four coefficients in a joint test are significant at a higher level in the men's regression than in the women's. One cannot conclude from this that village fixed effects are "larger" for men than for women, any more than one can conclude from looking only at t statistics that a coefficient is quantitatively important. Small coefficients, after all, can have large t statistics.

Table 5 Reduced sample: MNL results (robust t statistics in brackets)

Variable	Off-farm work		Temporary migration	
	Male	Female	Male	Female
Gender dummy	−8.706***	−6.680***	−6.261***	−3.044***
	[−11.44]	[−8.831]	[−8.70]	[−3.83]
Age	0.381***	0.272***	0.334***	0.144***
	[11.06]	[8.43]	[9.49]	[3.10]
Age^2/100	−0.463***	−0.366***	−0.444***	−0.303***
	[−10.80]	[−8.68]	[−9.63]	[−4.34]
School	0.339***	0.281***	0.357***	0.133
	[4.39]	[4.63]	[4.08]	[1.50]
School^2/100	−2.383***	−2.482***	−2.310***	−0.580
	[−4.88]	[−5.47]	[−4.14]	[−1.04]
Head	0.146		−0.868***	
	[0.72]		[−4.15]	
Land	−0.102***	−0.095***	−0.126***	−0.052
	[−3.59]	[−2.98]	[−4.35]	[−1.25]
Child		−0.405**		−0.681***
		[−2.55]		[−3.18]
Dependency	−0.062	−0.062	−0.482***	−0.482***
	[−0.65]	[−0.65]	[−3.43]	[−3.43]
Village fixed effects (#)	38	38	38	38
Average effect	0.934	1.061	0.008	0.078
Maximum	4.160	5.235	2.617	2.743
Minimum	−2.094	−1.247	−1.986	−2.381
Observations				5096
Log-likelihood (gender dummies)				−5101.064
Log-likelihood (all variables)				−3936.62
Likelihood ratio test				2328.90***

The symbols *, ** and *** denote statistical significance at the 10, 5, and 1% levels in two-tailed tests. The fixed effects of one village are set at zero to avoid perfect multicollinearity with the two gender dummies. There is a single coefficient (male = female) for DEPENDENCY, because coefficients left free to vary by gender were nearly identical in repeated regressions

Bold values refer to the two coefficients constrained to be equal for men and women, and to results for the complete set of 4 MNL equations

but this does not mean that quantitative effects are “greater” for men. If anything, the opposite is true.

Take a closer look now at the coefficients of the variables for individual and household characteristics in Table 5. The gender dummy is smaller (more negative) for men than for women, especially for temporary migration. This does not mean that, other factors equal, men are less likely than women to migrate or to work off-farm. This is not true because almost all the remaining coefficients in the model

Fig. 2 Effect of age on the probability of off-farm work (MNL). Source Calculated from the equations of the MNL model in Table 5, for a hypothetical person with 8 years of schooling who is a member of (but does not head) a household with 3 mu per capita of arable land, no small children, and a dependency ratio of 0.25. Residence in the base village is assumed

differ between men and women. The gender dummies are only intercepts; by themselves, they have no meaning.

For both men and women, the relationship between AGE and the logit of the decision to work off-farm or to migrate is quadratic, an inverted U. All eight relevant coefficients are significantly different from zero at the 1% level. The effect of an increase in age on the off-farm odds ratio (P_1/P_0) is positive until about age 41 for men and age 37 for women, when it becomes negative. For temporary migration, the effect on the odds ratio (P_2/P_0) peaks at about age 37.5 for men and age 24 for women.[4] If this were a binomial logit (BL) model, the effect on probability would peak at the same time as the effect on the odds ratio. This is MNL, however, not BL, so interpretation of the coefficients is more difficult. For this purpose, it is better to examine results in terms of the probability Eqs. 1b and 2b above.

Figures 2 and 3 illustrate the effect of AGE on the probability of working off-farm and the probability of migrating, respectively, for a hypothetical person with 8 years of schooling who is a member of (but does not head) a household with 3 mu per capita of arable land, no small children, and a dependency ratio of 0.25. We assume further that this person resides in the base village, with fixed effects equal to zero. These four curves, like the logit curves, are inverted-U curves, but they peak 2.5–5.5 years later for off-farm work (age 46.5 for men, 39.5 for women) and 2–3 years earlier for temporary migration (age 34.5 for men, 22 for women). It is difficult to summarize these curves in words, other than to note that for women, increased years of age has a negative effect on the probability of migrating from a very early age. This is not true for off-farm employment, where age is an asset for more than half a woman's normal working life. What we cannot determine is how

[4] The turning point of each odds ratio can be calculated by setting the derivative of the logit equation with respect to AGE equal to zero and solving for AGE. For male off-farm work, for example, $d(P_1/P_0)/d_{AGE} = 0.381 - 2*0.00463AGE$, which equals zero when $AGE = 0.381/0.00926$, which is approximately 41 years.

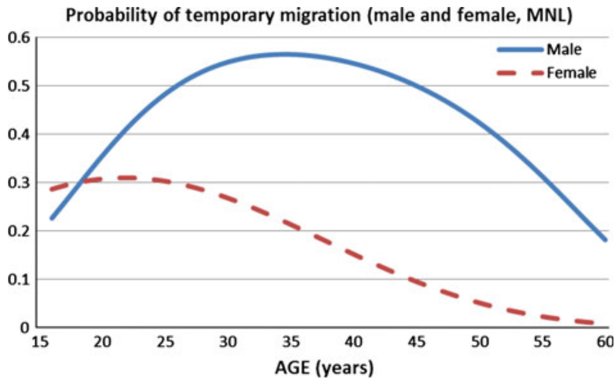

Fig. 3 Effect of age on the probability of temporary migration (MNL). Source Calculated from the equations of the MNL model in Table 5, for a hypothetical person with 8 years of schooling who is a member of (but does not head) a household with 3 mu per capita of arable land, no small children, and a dependency ratio of 0.25. Residence in the base village is assumed

much this reflects supply (the preferences of workers) and how much it reflects demand (the preferences of potential employers).

Since the coefficients of the MNL model are difficult to interpret, it is customary to calculate "marginal effects" for each coefficient. Marginal effects come in many forms (Cameron and Trivedi 2005: 122–124). In the best studies (e.g., Zhao 1999), they are clearly labeled and refer to the slope of the probability curve (or the effect of finite differences on probabilities), with all continuous variables set to (and evaluated at) their means and all dummy variables set to zero. Computer programs sometimes provide an option that computes all marginal effects at the sample means. These computations find their way into published studies without labels, warning, or explanation. It is difficult to make sense of a marginal effect computed at the mean of a dummy variable. What if average gender is 0.5? Is that a person who is half male and half female? Even worse, computer programs may treat a squared term as just another variable. Relying on these programs, some researchers who specify age as a quadratic function report marginal effects at the mean of AGE and at the mean of AGE^2 (e.g., Xia and Simmons 2004; Liu 2008; Wu 2010; Knight et al. 2010).[5] This is not correct, because AGE and AGE^2 are not independent variables. Proper calculation of the marginal effects of AGE must take into account both terms at the same time (e.g., Zhao 1999).

We find the calculation of marginal effects to be very unhelpful for understanding our MNL results. On Fig. 3, for example, consider marginal effects at the mean age of men (37) and the mean age of women (38). At these points, the effect of an additional year of age on the probability of migration is −0.4% points for a man and −1.2% points for a woman. Should we infer that age penalizes both genders, but women more than men? Yet between the ages of 22 and 34, marginal effects for

[5] Xia and Simmons (2004) rely on the variables experience and experience squared rather than age and age squared. They define experience as the number of years a person has lived following completion of his or her schooling. The other four MNL studies use the variables age and age squared.

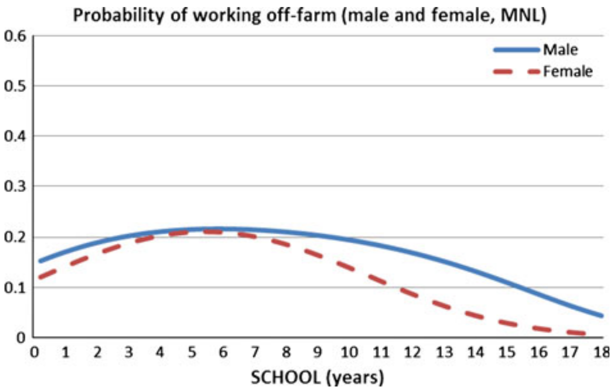

Fig. 4 Effect of schooling on the probability of off-farm work (MNL). Source Calculated from the equations of the MNL model in Table 5, for a hypothetical person, 37 years of age who is a member of (but does not head) a household with 3 mu per capita of arable land, no small children, and a dependency ratio of 0.25. Residence in the base village is assumed

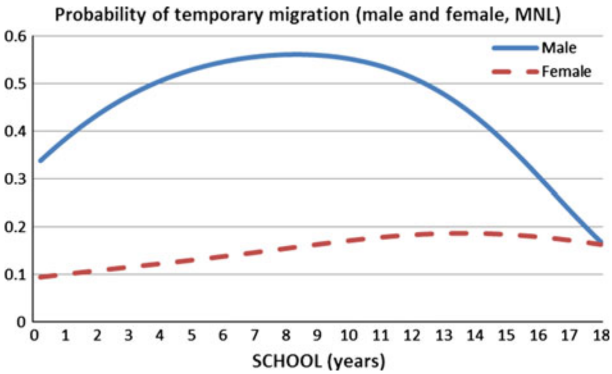

Fig. 5 Effect of schooling on the probability of temporary migration (MNL). Source Calculated from the equations of the MNL model in Table 5, for a hypothetical person, 37 years of age who is a member of (but does not head) a household with 3 mu per capita of arable land, no small children, and a dependency ratio of 0.25. Residence in the base village is assumed

men are positive, while they are negative for women. At ages younger than 22 years, the marginal effects are positive for both genders. In sum, we see no alternative to curves for a clear picture of the relationship between age and probabilities. Even single curves have limitations, for the curves change with any modification of assumed values for other independent variables.

The relationship between SCHOOL and the decision to work off-farm or to migrate is also quadratic for both men and women. Figures 4 and 5 illustrate these results for a hypothetical person who is 37 years of age, lives in the base village and is a member of (but does not head) a household with 3 mu per capita of arable land, no small children, and a dependency ratio of 0.25. All coefficients are highly significant, except for those relating to the probability of female migration, reported

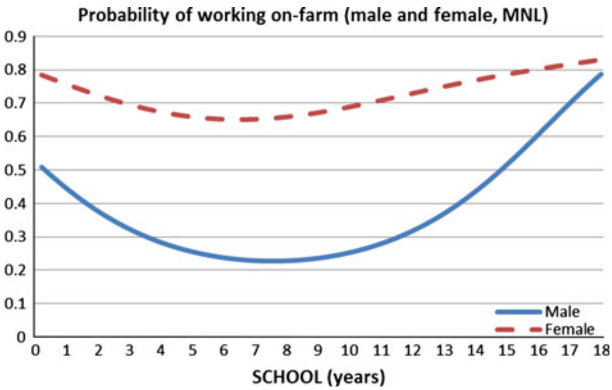

Fig. 6 Effect of schooling on the probability of on-farm work (MNL) Source Calculated from the equations of the MNL model in Table 5, for a hypothetical person, 37 years of age who is a member of (but does not head) a household with 3 mu per capita of arable land, no small children, and a dependency ratio of 0.25. Residence in the base village is assumed

in the last column of Table 5. We experimented by dropping the female SCHOOL^2 variable from the model. This produces a positive coefficient for SCHOOL in the migration equation ($t = 1.89$, significant at the 6% level), at the expense of the loss of statistical significance for this variable in the off-farm work equation. The likelihood ratio for the addition of female SCHOOL^2 to the MNL model is 27.78, significant at the 1% level, and based on this, we opted to retain the quadratic term.

The probability of working off-farm peaks in Fig. 4 at a similar number of years of schooling for men (about 6 years) and women (about 5.5 years), but the curve is steeper for women below and above these optimal years of schooling. At their maxima, the probabilities of working off-farm are similar for both men (0.22) and women (0.21); with 9 years of schooling (now compulsory in China), these probabilities fall slightly to 0.20 for men and sharply to 0.16 for women. The probability of migrating peaks in Fig. 5 at about 8 years for men and 13.5 years for women. The probability curve for men is very steep, with severe penalties for schooling that is less than or exceeds 8 years. The probability curve for women, in contrast, is rather flat, but with a positive slope over most of its range.

A full understanding for this variable is easier if we examine probabilities for the third employment option: Remaining on the farm. For each amount of schooling, this is simply the difference between unity and the sum of the other two probabilities. The resulting curves are shown in Fig. 6. Both curves are U-shaped, the inverse of the other curves. The minimum probability comes at about 6.5 years of schooling for women and at about 7.5 years for men. Up to these points, schooling decreases the probability that a person will remain on the farm. Beyond these minimum points, each additional year of education increases the probability that a person will be self-employed at home on the farm. The curve is particularly steep for men. The flatter curve for women reflects the positive effect that schooling has over a long range on the probability of migrating. In sum, especially for men,

Table 6 Binary logit (BL) models of temporary migration (robust t-statistics in brackets)

Variable	Equation 1 Male	Equation 1 Female	Equation 2	Equation 3
Male dummy	−3.537***		−1.652***	2.225***
	[−5.58]		[−3.49]	[−5.29]
Female dummy		−1.766**	−2.562***	−3.05***
		[−2.29]	[−5.27]	[−6.99]
Age	0.179***	0.053	0.074***	0.091***
	[5.84]	[1.20]	[3.07]	[3.89]
Age^2/100	−0.253***	−0.180***	−0.167***	−0.183***
	[−6.16]	[−2.65]	[−4.87]	[−5.52]
School	0.219***	0.069	0.180***	0.139***
	[2.78]	[0.80]	[3.15]	[2.60]
School^2/100	−1.341***	−0.012	−0.831**	−0.606*
	[−2.70]	[−0.02]	[−2.36]	[−1.89]
Head	−0.998***		−0.018	−0.042
	[−5.76]		[−0.13]	[−0.344]
Land	−0.088***	−0.029	−0.070***	−0.065***
	[−3.32]	[−0.75]	[−3.17]	[−5.27]
Child		−0.526**	−0.123	0.088
		[−2.53]	[−1.05]	[0.80]
Dependency	−0.490***	−0.490***	−0.387***	−0.233*
	[−3.69]	[−3.69]	[−3.00]	[−1.95]
Village fixed effects (#)	38	38	38	
Average effect	−0.523	−0.249	−0.469	
Maximum	1.411	1.971	1.516	
Minimum	−2.650	−2.360	−2.515	
Observations		5096	5096	5096
Log-likelihood (gender dummies)		−2341.8	−2341.8	−2341.8
Log-likelihood (all variables)		−1929.3	−2001.9	−2189.7
Likelihood ratio test		825.0***	679.8***	304.3***

The symbols *, ** and *** denote statistical significance at the 10, 5, and 1% levels in two-tailed tests. The fixed effects of one village are set at zero to avoid perfect multicollinearity with the two gender dummies. There is a single coefficient (male = female) for DEPENDENCY in Equation 1, because coefficients left free to vary by gender were nearly identical in repeated BL regressions. There is a single coefficient for all variables except the constant in Equations 2 and 3, because interaction effects with gender were removed

Bold values refer to the two coefficients constrained to be equal for men and women, and to results for the complete set of 4 MNL equations

both the most schooled and the least schooled are least likely to leave the family farm.

Zhang et al. (2002) collected Chinese survey data and estimated a binary choice model for off-farm employment in 1988, 1992, and 1996. They modeled schooling as a quadratic function, with findings very similar to ours. Off-farm employment in their study presumably includes temporary migration as well as local jobs. Neither schooling nor its square is significant in the 1988 cross-section, but both are significant in 1992 and especially in 1996. The authors do not mention that the shape of the function is an inverted U and note the positive effect of schooling while ignoring the negative effect of the square of schooling. They report only marginal effects (labeled dF/dx), not the estimated coefficients[6] in five regression equations (their Tables 5, 6), so we have no way of calculating the point at which the probability of off-farm labor participation peaks. Fortunately, Zhang et al. (2002, Table 7) report actual coefficients from a sixth binary choice regression that combines the three panels, allowing for interaction between independent variables and dummy variables for the last 2 years (1992 and 1996). Coefficients for schooling are not significant in the base year or in 1992, but they are significant at the 5% level in 1996. The coefficient on schooling for that year is 0.32 and the coefficient on the square of schooling is −0.02, so the function is an inverted U, with a positive but decreasing slope through 8 years of education, at which point the slope (marginal effect) becomes negative.

These findings for schooling are not consistent with a view of China as a dual economy, with unskilled surplus labor migrating from traditional agriculture to seek employment in the modern sector (Lewis 1954; Zhang 2009). Nor do the findings lend support to the suggestion by Katz and Stark (1987), taken up by Lall et al. (2006), that the effect of schooling on migration might be U-shaped: High for workers with low or high skills, but low for workers who have acquired an intermediate level of skills.[7] Indeed, what a U-shape describes is the effect of schooling on the probability that a worker remains on the family farm (see Fig. 6 again). Beyond a very modest number of years, additional schooling increases the likelihood that a worker remains on the farm. This suggests that high schools and colleges may not prepare students adequately for off-farm jobs and that underemployment of educated workers may be a problem in rural China.

The final variable relating to individual characteristics is HEAD, a dummy variable that equals unity if a person heads his or her household. For men, the coefficient of HEAD is negative and highly significant as a determinant of temporary migration, but not significantly different from zero for off-farm work (see Table 5 once again). The variable was not significant for women, possibly because few women are heads of households in China, so the female*HEAD interaction term was deleted from the model, leaving only male*HEAD interaction. Knight and Song (2003) do not allow for gender differences, but nonetheless were able to obtain similar MNL results.

The remaining three explanatory variables relate to characteristics of the rural household rather than the individual. All coefficients of LAND are negative as

[6] Démurger et al. (2009) and Wu (2010) estimate MNL rather than BL models but follow Zhang et al. (2002) in reporting only marginal effects, not the actual coefficients.

[7] "Both low- and high-skilled individuals are more likely to migrate but usually for different reasons: "surplus" low-skilled individuals have strong incentives to move to the city in search of a manual job they may not find in the rural area, while "scarce" educated workers may find that their human capital is better rewarded in cities than in rural areas" (Lall et al. 2006: 4).

expected and highly significant for both genders in the off-farm work equation. In the temporary migration equation, the coefficient is statistically significant only for men. We assume that land availability affects migration and that migration has no effect on land size. Regressions measure correlation only and the direction of causation might be the reverse of what we assume or there could be two-way causation (endogeneity). Removal of the variable has little effect on estimates of the other coefficients of the MNL model; however, so any reverse causation is not likely to be a serious problem for these data.

The reasoning behind the assumption of causation from LAND to work decisions is that less land means reduced household income, hence, greater incentive for a family worker to leave the farm to seek work (Zhao 1999). This implicitly assumes that each household has a fixed amount of land to cultivate. Until recently, this was a valid assumption for China. The Household Responsibility System that dates from the late 1970s allocates land-use rights to rural villagers on a very egalitarian basis. Migrating farmers had no incentive to rent out their land since this might send a signal to village officials that they are free to reallocate the land to others (Rozelle et al. 2002; Deininger and Jin 2007). In recent years, the central government has taken steps to increase land tenure security, beginning with the Rural Land Contract Law of 2002, which guarantees tenure for 30 years (Tao and Xu 2007). On October 19, 2008, the Communist Party issued a policy document on rural development that calls for farmers' "entitlement to subcontract, rent, exchange, transfer, and swap their land-use rights" (Xinhua News Agency 2008). There is every expectation that this will be written into law, with the freeing of land rental markets following quickly. This is likely to encourage migration since owners will be able to lease their land to others without fear of losing their rights, and they will even be able to sell their rights or use them as collateral for loans to finance migration (Lall et al. 2006).

For the year 2005, it is probably safe to assume that there is little reverse causation for the LAND variable, but this will not be true in the future if the promised reform of land tenure takes place. When that happens, it will be important to measure the amount of land over which a household has rights rather than the amount of land cultivated, for the latter might shrink when family members emigrate or work elsewhere in the township.

CHILD is a dummy variable equal to unity if there is a baby or toddler younger than 5 years of age in the household. This variable was not significant for men, so we dropped the male*CHILD interaction term from the model. The female*CHILD variable is highly significant for both off-farm work and for temporary migration. Its significance is all the more remarkable since we are holding DEPENDENCY—the ratio of the young plus the old to the number of adults aged 16–59 years—constant. DEPENDENCY is the only individual/household variable for which interaction with gender was not statistically significant. Therefore, the coefficients for this variable were estimated without regard to gender. The estimated coefficient is not significant in the off-farm equation, but it is negative and highly significant in the migration equation (see Table 5 once again). This suggests that small children constrain only female participation in the off-farm and migrant labor force. The presence of dependents in general—young and old—has a negative effect on

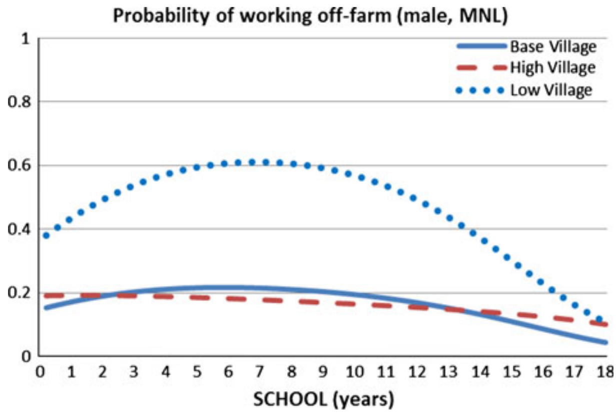

Fig. 7 Effect of schooling on the probability of off-farm work (MNL). Source Calculated from the equations of the MNL model in Table 5, for a hypothetical male, 37 years of age who is a member of (but does not head) a household with 3 mu per capita of arable land, and a dependency ratio of 0.25. This person is assumed to reside in the base village, in the village with the highest fixed migration effects or in the village with the lowest fixed migration effects

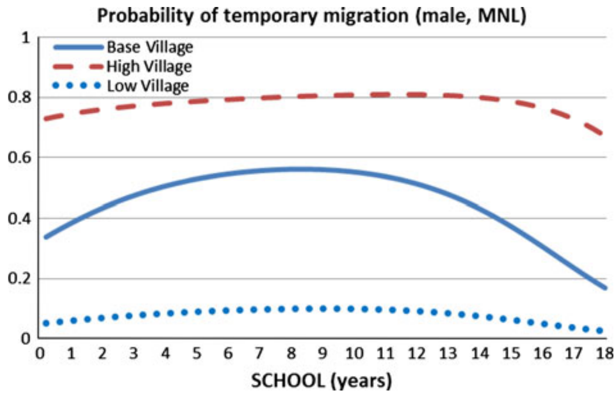

Fig. 8 Effect of schooling on the probability of temporary migration (MNL). Source Calculated from the equations of the MNL model in Table 5, for a hypothetical male, 37 years of age who is a member of (but does not head) a household with 3 mu per capita of arable land, and a dependency ratio of 0.25. This person is assumed to reside in the base village, in the village with the highest fixed migration effects or in the village with the lowest fixed migration effects

migration of men and women, but has no significant effect on the probability of off-farm work in the local community.

The coefficients of the village dummies for men range from −2.09 to +4.16 in the off-farm work equation and from −1.99 to +2.62 in the temporary migration equation. The coefficients for women have an even larger range, from −1.25 to +5.24 in the off-farm work equation and from −2.38 to +2.74 in the temporary migration equation. The village of registration is the single most important determinant of individuals’ place of work in our sample.

To illustrate the magnitude of these village fixed effects, consider a hypothetical man, aged 37 years who is a member of (but does not head) a household with 3 mu per capita of arable land, and a dependency ratio equal to 0.25. We allow his years of schooling to vary from 0 to 18. Initially, suppose that this person lives in the base village, where fixed effects are set at zero. The relationship between years of schooling and the probability of working off-farm or of migrating is graphed by the solid line in Fig. 7 and in Fig. 8, respectively. These are equivalent to the solid lines in Figs. 4 and 5.

Suppose now that we move this person to a similar household in the rural village with the highest fixed effects for migration. This village has fixed migration effects of $+2.62$ and fixed off-farm effects of $+2.07$. The result of this change of village is illustrated by the dashed lines in Figs. 7 and 8. In Fig. 8, note that the curve shifts up quite sharply and becomes nearly flat over a wide range of SCHOOL, hovering around the $P_2 = 0.8$ line. The curve in Fig. 7 flattens, but does not shift much at all, hovering just below the $P_1 = 0.2$ line. This is not obvious from the coefficients, which increase to a large positive number in each equation. The intuition is that what matters is not only the absolute size of the coefficients, but also their relative size, and 2.62 is larger than 2.07. This is a result that is difficult to visualize without the aid of a graph.

Now, move this hypothetical person to a similar household in the rural village with the lowest fixed effects for migration. This village has fixed migration effects of -1.99 and fixed off-farm effects of $+0.80$. The results of the move on probabilities are illustrated by the dotted lines in Figs. 7 and 8. The curves shift in the same direction as the coefficients change: Down for temporary migration and up for off-farm work. The temporary migration curve becomes very low and rather flat, whereas the off-farm curve is high and peaked.

For years of schooling ranging from 2 to 17 years, the gap between the low and high villages in this example exceeds 69% points. This is a huge increase in the probability of migrating that dwarfs the effect of even large changes in age or in schooling. Village effects are very, very important. The three villages, incidentally, are from three different provinces: Jilin (base), Shandong (low), and Zhejiang (high). But this is just coincidence. There is no evidence of inter-provincial differences, neither in the size of village effects nor in the size of other coefficients in the MNL model.

The BL model, gender, and village fixed effects

The MNL model is complex and its coefficients are difficult to interpret. It is reasonable to ask, then, what difference it makes to use it rather than the simpler binomial logit (BL) model. In this spirit, we used the same variables and data to estimate a BL model for the probability of temporary migration. There are only two choices: Temporary migration or no migration. The results of this exercise are reported as "Equation 1" in Table 6. Recall that the dependent variable in the MNL model is $\ln (P_2/P_0)$, whereas it is $\ln\{P_2/(P_0 + P_1)\}$ in the BL model, so the coefficients cannot be directly compared. Nonetheless, for men, all eight

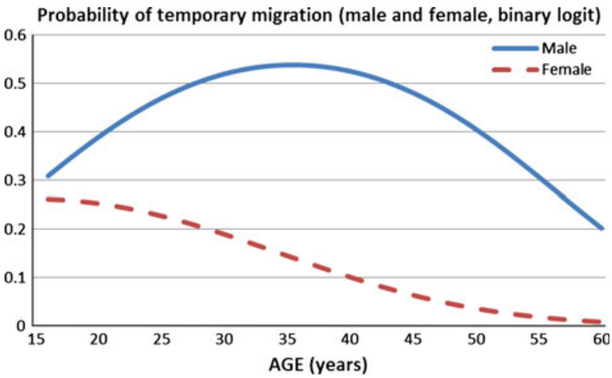

Fig. 9 Effect of age on the probability of temporary migration (BL). Source: Calculated from binary choice Equation 1 reported in Table 6, for a hypothetical person with 8 years of schooling who is a member of (but does not head) a household with 3 mu per capita of arable land, no small children, and a dependency ratio of 0.25. Base village fixed effects are assumed

coefficients of the individual/household variables have the same sign and statistical significance in the BL model as they do in the MNL model. For women, however, the BL model does not perform as well as the MNL model: The sign of each coefficient of the eight individual/household variables is the same in each model, but only two are significant at the 1% level and two at the 5% level in the BL model, compared to five at the 1% level in the MNL model. Female interaction with SCHOOL and SCHOOL^2 in the BL model is significant in a joint test, but only at the 6% level. SCHOOL alone, entered for women without the quadratic term, has a positive coefficient that is significant at the 2% level.

The only way to compare the magnitude of the BL results to the MNL results is to solve each model for P_2—the probability of migration—and graph the results for each explanatory variable of interest. This is done in Fig. 9 for AGE, using the same assumed values for other dependent values as were used to graph the MNL results in Fig. 3. Comparing the two figures reveals that the BL curve for men differs little from that of the MNL curve. The BL curve peaks a year later (at 35.5 years), but is otherwise very close to the MNL curve. The BL curve for women, however, differs noticeably from the MNL curve. The BL curve starts lower and has a negative slope throughout, whereas the MNL curve starts somewhat higher and has a positive slope until it peaks at about age 22.

Like Zhao (1999) and Liu (2008), we fail to find major differences between the BL and MNL model results. The MNL model does, however, provide visibly different—and presumably more accurate—results for the female portion of our sample. We cannot generalize this conclusion, because neither Zhao (1999) nor Liu (2008) allowed for interaction between gender and other explanatory variables.

How important is it to specify interaction between gender and other variables? We have shown that there is an important difference between men and women in the size of the estimated coefficients. What about the statistical significance of the coefficients? Does removal of interaction terms affect this as well? Equation 2 of Table 6 answers this question with results from a BL regression that estimates “unisex” coef cients for all variables except for the constant. A key result is that HEAD and CHILD lose statistical signi cance in the gender-free model. Otherwise, the results are acceptable in terms of the signi cance of explanatory variables. Many key variables are signi cant at the 1% level, and even the square of schooling manages to pass a statistical test at the 5% level.

What about village xed effects? These clearly add to the explanatory value of the model, but what impact would failure to control for these have on the statistical signi cance of other variables in the model? The consequences of removing all village dummies are shown in Equation 3 of Table 6. The most damaging result is the sharp fall in t-statistics for SCHOOL, SCHOOL^2, and DEPENDENCY. The signi cance of SCHOOL^2 and DEPENDENCY both drop to the 10% level. The t-statistics for AGE, AGE^2, and LAND increase notably, but from an already high level, so there would be no danger of false inference of statistical signi cance from failure to control for village xed effects. Equation 3 is useful, but Equation 1 provides a much better picture of the determinants of temporary migration from rural China. We conclude that it helps considerably to control for the interaction of variables with gender and for the xed effects of villages. This is at least as important as the choice of model (MNL vs. BL) in research design.

Removal of interaction terms and village dummies from the MNL model has effects that are very similar to their removal from the BL model. The results of this exercise are not shown here, but are available from the authors on request.

Conclusion

Using survey data for the year 2005 drawn from nearly two thousand households in 43 rural Chinese villages, we model off-farm work and migration as a multinomial choice that is a function of characteristics of the individual and the household, subject to “village of registration” xed effects. “Off-farm” is de ned as working away from the household farm at least part of the time, somewhere in the township. “Migration” is de ned as working away from the township of legal residence for at least part of the year. Our main contribution with this research is the close attention we pay to village effects and to gender differences in all aspects of the migration decision. A main conclusion is that there are signi cant differences between genders and between villages, independent of differences in the characteristics of individuals and households. Any model—multinomial or binomial—that ignores these differences is the worse for it.

Three variables, in addition to gender, measure personal characteristics of individuals: age, years of schooling, and a dummy variable indicating whether a person does or does not head a household. For both men and women, the relationship between age and the decision to work off-farm or to migrate is an inverted U: rst positive and then negative. For women, increased age has a negative effect on the probability of migrating from an early age, much younger than for men. This is less true for off-farm work in the local community. The relationship between schooling and the probability of off-farm work or migration is also an inverted U for both men and women, but the curves of the two genders are similar only for off-farm work. For temporary migration, the curve for men peaks at about 9 years of compulsory schooling; further schooling beyond this level has a large negative effect on migration. For women, in contrast, the curve peaks well beyond high school, but is rather flat, so education has little effect in general on female migration. Counterintuitively, we find that those with least schooling and those with most schooling are most likely to stay on the farm rather than leave to find work. This is especially true for men. Heading a household has no independent effect at all on working off-farm; it has a negative effect on the migration of men, but not of women.

The findings for schooling are especially disturbing. They suggest that schools in rural China poorly prepare workers for off-farm employment and that many of those educated beyond basic grades are underemployed. This is an area that requires further research. It would be interesting, for example, to modify our model to allow for interaction between villages and schooling, since quality of schooling most likely varies from village to village.

Three variables measure household characteristics: The amount of land per capita available to the household for cultivation, a dummy variable indicating whether an individual does or does not live in a household with a child younger than 5 years of age, and the dependency ratio of the household, measured as the ratio of the number of old plus young members to the number of working-age adults. The coefficients of the land variable are negative for off-farm work and for migration—and statistically significant as well—except in the case of female migration. The child dummy was negative and statistically significant for both off-farm work and for migration, but only in the case of women. A high ratio of dependents to working-age adults has a significantly negative effect on migration, but not on off-farm work in the local township. This variable is the only one of the individual/household coefficients that did not differ significantly by gender.

Finally, the village in which a person is registered turns out to be the single most important determinant of migration and off-farm work; it can dwarf the effect of even large differences in age or in years of schooling. Moreover, these unobserved village effects differ significantly between men and women, so it was necessary to include gender interaction terms for the village dummies. We have no information about the villages in our sample, but we do know that only a small part of the substantial differences between villages in the propensity to migrate or work off-farm can be explained by differences in the characteristics of households or of individuals sampled in those villages. Future work on the determinants of migration in China should focus not only on variables at the individual and household levels, but also on variables that vary by village, such as wage levels, availability of off-farm employment, distance from migrants' main destinations, access to railways and paved roads, communications (radio, TV, and telephone), and other amenities, in addition to village-based networks of migrants in townships and provinces of destination.

In contrast, the province in which a person is registered has no effect on the probability of migration or of off-farm work, once we control for characteristics of individuals, characteristics of households, and village fixed effects. Nor is there any evidence of signi cant inter-provincial differences in village xed effects or in other coef cients of the model.

To sum up, our main ndings are as follows: (1) Village of household registration is the single most important determinant of an individual's place of work. (2) Gender differences are very important. Nearly all coef cients—even those of the village dummies—vary signi cantly between men and women. (3) The relationship between age and the probability of working off-farm or migrating is non-linear, an inverted U, for both men and women, although the curves differ signi cantly by gender. (4) The relationship between years of schooling and the probability of leaving the family farm to work in the township or to migrate, surprisingly, is also an inverted U for both men and women. Beyond a modest number of years (less than the supposedly compulsory 9), further education *increases* the probability that a worker will remain on the family farm rather than work for salary or wages. This is evidence of underemployment of educated workers, re ecting schooling that is of poor quality, or at least inappropriate for the job market. These ndings have important implications for migration, labor and education policy in China.

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