Journal of Economic Behavior and Organization 220 (2024) 325–342

Contents lists available at ScienceDirect

Journal of Economic Behavior and Organization

journal homepage: www.elsevier.com/locate/jebo

Property rights, labor reallocation, and gender inequality in rural China

Xinjie Shi [a,b,c,f], Bingyu Huangfu [a], Songqing Jin [d], Xuwen Gao [e,f,g,*]

[a] China Academy for Rural Development, School of Public Affairs, Zhejiang University, Hangzhou, China
[b] Research Center for Common Prosperity, Future Regional Development Laboratory, Innovation Center of Yangtze River Delta, Zhejiang University, Jiashan, China
[c] Center for Common Prosperity of Zhejiang University & Huzhou City, Huzhou, China
[d] Department of Agricultural, Food and Resource Economics, Michigan State University, East Lansing, MI, USA
[e] School of Public Affairs, Zhejiang University, Hangzhou, China
[f] Institute for Common Prosperity and Development, Zhejiang University, Hangzhou, China
[g] Center of Social Welfare and Governance, Zhejiang University, Hangzhou, China

ARTICLE INFO

JEL Classification:
Q15
D13
J16
J71

Keywords:
Property right
Labor reallocation
Gender inequality
Rural China

ABSTRACT

This study examines the gender-differentiated effects of improved land property rights on labor reallocation in China, using quasi-exogenous variation in the timing of the implementation of the Rural Land Contracting Law, which allows farmers to lease out their land. We find that while men and women tend to shift labor from the agricultural into non-agricultural sector following the land reform, women lag behind men. Our findings reveal a noticeable gender gap in the growth of off-farm labor participation (+16.69 % among men; + 1.95 % among women) and hours worked in off-farm sectors. The gender-differentiated effects on off-farm employment are likely caused by disadvantaged labor market conditions for women. These findings underscore the importance of improved land property rights in fostering rural structural transformation. Moreover, our results suggest that implementing land reforms without accompanying changes to address the root causes of gendered differences in off-farm employment could limit their full potential. This study has significant policy implications for the rural transformation of developing countries.

1. Introduction

In developing countries, it is common for up to two-thirds of the population to work in the agriculture sector, compared to less than 5 % of the population in developed countries.[1] However, the value added per worker in agriculture is lower than that in the non-agriculture sector, particularly in developing countries (Gollin et al., 2014). For instance, in most countries across Sub-Saharan Africa and South Asia, the agricultural value added per worker was typically less than $1000 in 2017.[2] This misallocation of workers is the primary cause of the low aggregate productivity in developing countries (Restuccia et al., 2008). Well-defined land property rights have long been recognized as an essential premise for efficient workforce allocation, as tenure security encourages farmers to engage in land rental and move into non-farm employment (Chernina et al., 2014; de Janvry et al., 2015). Despite efforts emphasizing equality

* Corresponding author.
E-mail address: xwgao@zju.edu.cn (X. Gao).

[1] See https://ourworldindata.org/employment-in-agriculture
[2] See https://ourworldindata.org/employment-in-agriculture

https://doi.org/10.1016/j.jebo.2024.02.020
Received 27 January 2023; Received in revised form 16 January 2024; Accepted 10 February 2024
Available online 23 February 2024

and empowerment of women in agriculture (Nelson et al., 2012), little is known about gender disparity in the reallocation of labor from agricultural to non-agricultural employment following property rights reforms. This study is among the first to explore this issue.

We examine the effects of the Rural Land Contracting Law (RLCL) in China—a property rights reform that provides farmers with legal rights to lease their farmland as part of existing protections for property rights security—on gender-differentiated labor reallocation. China is an appropriate choice for this study for two reasons. First, while women have limited land inheritance rights in some developing countries (Roy, 2015), men and women have equal land rights in China.[3] Codifying equality implies that gender disparity in land rights is unlikely to be the primary reason for gender differences in responses to land reforms. Other factors, such as child penalties, high mobility costs, human capital gap, and gender identity norms, may exist. Moreover, gender discrimination in the labor market may play a role, even in the absence of gender differences in land rights. Second, by allowing farmers to act as lessors, the RLCL provides a unique opportunity to estimate the gender-differentiated process in the labor movement to non-agricultural sectors.

To examine the relationship between land reform and labor reallocation from a gender perspective, this study utilizes five datasets. Following Chari et al. (2021), we collect data on when each province in China began implementing land reform after the central government's enactment of the RLCL in 2003. Subsequently, we combine these data with a panel dataset from a nationally representative household survey—the China Health and Nutrition Survey (CHNS)—which covers approximately 7200 households, including over 30,000 individuals in 15 provinces and municipalities with various socioeconomic indicators.[4] For a supplementary analysis of the mechanisms, we adopt data from (i) China's 2000 Population Census, which includes detailed information about individual occupations for a large sample; (ii) the second wave of China's Women Social Status Survey (CWSS), conducted in 2000, which provides information on gender identity norms; and (iii) the 2014 China Labor-force Dynamics Survey (CLDS).

Drawing on the variation in land reform timing across provinces, we employ a difference-in-differences (DID) method to analyze the issues of interest and obtain several main findings. First, we investigate gender-differentiated labor reallocation following the RLCL. Although property rights laws lead to a significant increase in off-farm employment status for both men and women, there is a significant gender gap of 5.45 percentage points regarding the off-farm work employment (5.94 percentage points increase for men and only a 0.49 percentage points increase for women). In addition, we find that, after the reform, hourly off-farm income for men increased by 28.6 % while it only increased by 9.3 % for women. As for the agricultural sector, we do not find gender-specific effects, as men and women worked 9.98 % fewer hours on average in the agricultural sector after the reform. These results suggest that although land reform encourages men and women to shift their labor away from the agricultural sector and into non-agricultural sectors, women lag behind men when transitioning into the off-farm sector.

In the mechanisms analysis, we find that this gender-specific response can be partially explained by the disadvantaged labor market conditions faced by women. Unskilled rural workers are more likely to engage in blue-collar occupations, which are male-dominated. After transitioning from agriculture, rural women encounter fewer opportunities in non-agricultural employment compared to men. We rule out other mechanisms, including child penalty, higher mobility costs, the human capital gap, and gender identity norms, which do not explain why women lag behind men after the land reform.

We conduct a series of robustness checks to verify whether the effects we estimate are from land reform. The validity of our identification relies on the conditional exogeneity of the timing of RLCL implementation. To examine this identification issue, we show that local reform timing is not associated with provincial socioeconomic characteristics, local gender culture, and differential opportunities for men and women at the baseline year. Changes in these characteristics after the land reform are not linked to reform status. Second, an event study further verifies that the trend for gender-specific effects mirrors the implementation in early-reforming provinces, those that reformed late, and those that had yet to reform.

The biggest challenge for our identification is that some factors correlated with the implementation of the land reform may have gender differentiated impacts. Our robust results control for provincial economic status, industry composition, and gender identity norms at baseline interacted with year fixed effects and then with year-by-gender fixed effects. Another concern is that other reforms in rural China may confound the RLCL's effects. For instance, eliminating agricultural tax (Chari et al., 2021) may occur in regions where women may benefit more than men. However, the results do not change significantly after controlling for indicators of interaction between tax reform year and the gender dummy. Similarly, we rule out the gender-specific confounding effects of China's accession to the World Trade Organization (WTO) (Erten and Leight, 2021; Khanna et al., 2021). Adopting wild cluster bootstrap do not change our main results and we also conduct a placebo test. Finally, we follow Callaway and Sant'Anna (2021) to address the "negative weight problem" in staggered DID design (de Chaisemartin and D'Haultfoeuille, 2020; Goodman-Bacon, 2021; Baker et al., 2022). The main results remain unchanged.

Our study relates to two strands of literature. First, we investigate the impact of property rights reform on factor reallocation and its effects on rural development. Insecure property rights may be associated with the misallocation of land and labor factors, and reforms aimed at secure property rights lead to land (Cheng et al., 2019; Chari et al., 2021) and labor reallocation (de Janvry et al., 2015; Giles and Mu, 2018; Gao et al., 2021). This type of reallocation has been identified in the case of the RLCL in China (Deininger and Jin, 2009; Chari et al., 2021). Our research complements these studies, as we investigate possible associations between the RLCL and gender-differentiated labor reallocation. We are among the first to investigate the gender gap in the agricultural sector's non-agricultural labor market caused by land reform.

The second strand of literature examines gender gaps in society. Despite a remarkable convergence in the economic roles of men and women in the labor market, persistent and significant gender gaps exist in labor supply, wages, and representation in top jobs

[3] However, women in rural China can lose their rights to land upon marriage, divorce, or widowhood (Hare et al., 2007; Judd, 2007).

[4] Refer to https://www.cpc.unc.edu/projects/china for more details.

(Cortés and Pan, 2020; Jarrell and Stanley, 2004; Thu et al., 2023; Cortés et al., 2023). Studies have demonstrated that these gender gaps could be due to (i) changes in labor demand and labor market opportunities that may be associated with physical conditions and bargaining (Card et al., 2016; Sorkin, 2017; Gerard et al., 2021; Biasi and Sarsons, 2022); (ii) differential child penalties faced by fathers and mothers after childbirth (Bertrand et al., 2010; Fitzenberger et al., 2013; Schönberg and Ludsteck, 2014); (iii) differential mobility costs for men and women (Paull, 2008; Halla et al., 2020; Kleven et al., 2019); (iv) differences in human capital that lessen over time as human capital investments between men and women converge (Goldin, 2014, 2021; Blau and Kahn, 2017); and (v) gender identity norms that assign a breadwinner role to husbands and shift the role of married women from paid labor supply to childcare and housework (Bertrand et al., 2015; Kleven and Landais, 2017; Zinovyeva and Tverdostup, 2021). We contribute to this line of work by providing quasi-experimental evidence of the role of these factors as mechanisms in determining gender differences in labor reallocation after the land reform in China. We identify the mechanisms that may cause gender disparities despite the land reform having no gender preferences.

The remainder of this study is organized as follows. Section 2 briefly reviews the institutional background of land reform and off-farm employment in China. Section 3 presents the data and identification strategy. Section 4 reports the main results of differing labor reallocation between men and women after the land reform implementation in rural China. Section 5 discusses the potential mechanisms driving this gender-specific pattern. Section 6 presents a series of sensitivity analysis. Section 7 concludes the study.

2. Background

Since 1949, most restructuring in China has involved rural land system reform. At the start of the reform and opening up in the late 1970s (Liu et al., 2014), rural land system restructuring was initiated in Xiaogang Village, Fengyang County, Anhui Province, and became the basis of the Household Responsibility System (HRS). The HRS attempted to divide land ownership property rights into two parts: ownership by village collectives and that by farmers with contractual management rights (private use rights) (Wang and Zhang, 2017). This plan increased farmers' enthusiasm for agricultural production, and rural development aligned more closely with the market economy in China. The pilot expansion of this reform in 1984 marked the beginning of the nationwide popularization of the HRS and confirmation of a 15-year land contract (Cheng et al., 2019).

Although farmers obtained private use rights for farmland after 1979, they risked losing land due to insecurity, as local governments had the right to reassign plots until the late 1990s. In 1998, the Land Management Law extended the land contract period to 30 years and enhanced land tenure security (Chari et al., 2021). However, after the relaxation of the *hukou* system,[5] some farmers were allowed to leave agriculture and obtain jobs in cities, leading them to abandon farmlands that could not be transferred to other farmers (Shi, 2018, 2022a, 2022b). This situation called for an official clarification of the legitimacy of farmland transfers. In response, the Chinese government introduced the RLCL in 2003. Although land transfers before 2003 occurred primarily among neighbors and were based on informal (verbal) land rental agreements, they sometimes caused disputes that could not be resolved. The 2003 land reform gave farmers the legal right to transfer land and outlined rules of leasing, transferring leases, and dispute settlements.

The RLCL differed from previous reforms in allowing farmers to lease farmland and standardizing the farmland transfer procedure, leading to two changes in farmland (Chari et al., 2021; Zhou et al., 2020). First, the RLCL decreased the probability of land reallocation led by local governments (land expropriation), which made farmers feel secure about their land rights. The share of villages involved in this type of land reallocation was as low as 4 % between 2001 and 2006 and decreased to zero between 2007 and 2008 (Chari et al., 2021). Second, the RLCL significantly promoted land transfers in rural China. As Appendix Fig. A1 illustrates, the percentage of transferred area to the total contracted area more than doubled from 2002 (1.4 %) to 2005 (3.1 %). After 2008, the farmland transfer scale expanded rapidly due to industrialization and urbanization.

Changes in land reallocation indicate labor reallocation from the agricultural sector to the off-farm sector. In 1978, 93 % of rural laborers (approximately 283 million) worked in agriculture; in 1983, the total number of migrants was approximately 2 million (Cai et al., 2009). Rural–urban migration increased in the 1980s with the gradual abolition of institutional obstacles; the number of migrants increased to 30 million in the late 1980s and grew to 75.5 million in 2000 because of rapid economic growth after 1992 (Cai et al., 2009). As shown in Appendix Fig. A2, the number of migrants increased to 104.7 million by 2002, which was 20 million more than in 2001. Further, it rose to 113.9 million in 2003, with migrants accounting for 44.4 % of urban employees. In 2022, the number of migrants reached 295.6 million, of which men accounted for 63.4 % while women accounted for only 36.6 %.

The rapid increase in the number of migrants underscores the shifts in farmers' off-farm employment following the reform (Öztürk and Yalçın, 2023). In China, migrant workers with lower levels of education often concentrate in labor-intensive industries, such as construction and manufacturing.[6] Undoubtedly, they were indispensable to the rapid development of China's industrial sector in the early 21st century. This trend aligns with the main question that this study seeks to explore: Was the labor reallocation trend the same for men and women after China's land reform?

[5] The Chinese *hukou* system is a method of population registration that distinguishes urban and rural residents. Before 1984, rural and urban residents were required to continue residing and working within their geographical locations, indicating that rural-urban migration was strictly controlled.

[6] National Bureau of Statistics of China: Monitoring Survey Report of China's Rural Migrant Workers, 2010-2022.See https://www.stats.gov.cn/.

3. Data and identification strategy

3.1. Data

This study uses datasets from the CHNS and the land reform implementation timing by province. In addition, we use supplementary data from three other surveys: the 2000 Population Census of China, the second wave of the CWSS, and the CLDS.

Following Chari et al. (2021), we collect information on the timing of RLCL implementation at the province level between 2003 and 2012. As the RLCL is implemented top-down, we rule out program placement bias by sorting out the specific implementation time of the RLCL by province. We add the implementation information of five additional provinces to those presented in Chari et al. (2021): Heilongjiang, Beijing, Henan, Guizhou, and Hubei. We remove the provinces not included in the CHNS, thus leaving 12 provinces in our sample (Appendix Table A1).

The CHNS comprises panel survey data collected by the Carolina Population Center at the University of North Carolina and the National Institute for Nutrition and Health at the Chinese Center for Disease Control and Prevention. This survey adopted a stratified, multilevel random cluster sampling method. It included information on household demographic characteristics, economic status, employment, and living conditions, particularly individuals' nutritional and health status, allowing us to examine the effect of land property rights in depth. The baseline national survey wave was conducted in 1989. We use rural data from six annual waves between 2000 and 2015 (2000, 2004, 2006, 2009, 2011, and 2015), including 17,765 individuals and 4994 households in 52 cities from 12 provinces (Beijing, Liaoning, Heilongjiang, Shanghai, Jiangsu, Shandong, Henan, Hubei, Hunan, Guangxi, Chongqing, and Guizhou). The level of observation for the main regressions using these data is individual by year.

Furthermore, we use data from three additional surveys to analyze the mechanisms and interpret our main results. First, we examine data from the 2000 Population Census conducted by the National Bureau of Statistics to examine gender discrimination in the labor market. The detailed industry and occupation information gathered at the individual level from the Population Census allows us to explore gender disparities in labor market conditions across provinces.

Second, to complement our investigation of gender identity norms, we analyze data from the second wave of CWSS, which was conducted jointly by the All-China Women's Federation and the National Statistics Bureau of China in 2000. The survey comprises nine information categories: health, education, economy, social security, politics, marriage and family, lifestyle, legal rights and cognition, and gender concept and attitude. Our analysis include 19,449 individuals.

Third, we supplement the data with migration information from the 2014 CLDS. The CLDS is a national social survey containing rich information on 23,594 individuals across 29 Chinese provinces.[7] The 2014 CLDS includes detailed migration information based on a retrospective history of respondents' locations, which is used to create an individual-level longitudinal panel between 2000 and 2014.

Table 1 presents the summary statistics for the outcome variables in these four datasets. This study focuses on the rural labor force aged 16–55 years in the survey year. The primary outcome variables of interest indicate whether individuals work in the off-farm or agricultural sector, daily work hours of off-farm and agricultural work, and housework hours. The data show that men were more likely to work in off-farm employment (10.5 percentage points more than women) and for 0.92 more hours than women. Men also hold an advantage in agricultural work. In contrast, women devoted more time to housework than men, with a 35.7 percentage points higher likelihood of involvement and an additional 3.2 housework hours per day.

Regarding earnings from off-farm work, men earn 7743 Chinese yuan while women only earn 4392 Chinese yuan. Furthermore, the share of married individuals and number of children are slightly higher for women compared to men. We also report the share of individuals working in manufacturing, construction, and service sectors using the 2000 Population Census. The proportion of men working in these sectors is significantly higher than that of women, especially in construction and manufacturing. In addition, men are more likely to migrate than women. Furthermore, there are differences in their perceptions of gender norms.

Thus, the descriptive statistics indicate that men are more likely to engage in off-farm employment, whereas women primarily do housework. In the next section we will explore whether these gender-differentiated outcomes suggest labor reallocation following the land property rights reforms.

3.2. Identification strategy

To investigate the gender-differentiated effects of the RLCL on the sectoral allocation of labor, we use the following regression:

$$Y_{it} = \beta_0 + \beta_1 Reform_{p,t-1} + \beta_2 Reform_{p,t-1} \times Female_i + \beta_3 Female_i + \delta_t + \eta_i + \varepsilon_{it} \tag{1}$$

where Y_{it} represents the labor market outcomes of individual i in year t, $Reform_{p,t-1}$ is an indicator of whether the reform has been implemented in province p and year $t-1$,[8] and $Female_i$ is a gender dummy equal to 1 for women and 0 otherwise. The primary independent variable of interest is the interaction between the reform and female indicators, representing a gender-differentiated

[7] A probability-proportional-to-size sampling (PPS) procedure based on population size and administrative units is adopted to ensure that the survey is nationally representative. Consequently, the sample size distribution across cities in the CLDS is consistent with the geographic distribution of China's population.

[8] As many reforms are implemented late in the year (October and November), we focus on the year after implementation.

Table 1
Summary statistics.

Variable names	Men	Women	Differences	
			Mean	P value
Panel A: Variables from CHNS				
Have off-farm work (=1)	0.356	0.251	0.105***	0.0000
Average daily off-farm work hours	3.102	2.185	0.917***	0.0000
Have farm work (=1)	0.299	0.278	0.020***	0.0000
Average daily farm work hours	5.335	5.005	0.330***	0.0000
Do daily housework and childcare (=1)	0.460	0.816	−0.357***	0.0000
Average daily housework and childcare hours	5.157	8.398	−3.241***	0.0000
Annual off-farm work earnings	7743.071	4392.008	3351.063***	0.0000
Annual farm work earnings	1283.012	1908.225	−625.213***	0.0000
Married (=1)	0.688	0.714	−0.026***	0.0000
Number of children	0.755	0.770	−0.015*	0.0780
Panel B: Variables from the 2000 Population Census				
Work in the manufacturing sector (=1)	0.129	0.124	0.005***	0.0000
Work in the construction sector (=1)	0.048	0.007	0.041***	0.0000
Work in the service sector (=1)	0.023	0.022	0.001**	0.0302
Panel C: Variables from 2014 CLDS				
Migration (=1)	0.005	0.004	0.001**	0.0188
Panel D: Variables from 2000 CWSS				
Men should focus on society, and women should focus on family (Agree=1)	0.502	0.451	0.052***	0.0000
Men are inherently more capable than women (Agree=1)	0.312	0.280	0.032***	0.0000
Women should marry well rather than do well in their careers (Agree=1)	0.308	0.359	−0.051***	0.0000
The biggest role of women is to produce children (Agree=1)	0.257	0.249	0.008	0.2015
With economic development, women's social status will not improve (Agree=1)	0.801	0.810	−0.009	0.1140

Note: This table shows the summary statistics for the main outcome and control variables. Panel A reports the main variables obtained from the *CHNS*. "Have off-farm work," "Have farm work," and "Do daily housework and childcare" are dummy variables that equal 1 if positive, and 0 otherwise. "Average daily off-farm work hours" and "Average daily farm work hours" indicate average daily off-farm work and farm work hours, respectively. "Average daily housework and childcare hours" includes time for housework, such as washing, cooking, housework, and taking care of children. "Annual off-farm work earnings" and "Annual farm work earnings" represent the annual income of off-farm and farm work, respectively. "Married" shows marital status, with 1 indicating married and 0 indicating single, widowed, or divorced. "Number of children" denotes the number of living children. Panel B reports variables calculated from the 2000 Population Census. The manufacturing, construction, and service industries are based on categories 13–43, 47–49, and 75–82 in the 1994 Industrial Classification and Codes for national economic activities (GB/T 4754-94). Panel C reports the "Migration" variable from the 2014 CLDS. "Migration" equals 1 if the individual migrated in the current year, and 0 otherwise. Panel D reports the social norms obtained from the 2000 CWSS. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

response to the reform. The regression controls for individual and year fixed effects which are denoted by η_i, and δ_t, respectively. Standard errors are clustered at the province level.

The validity of the empirical strategy relies on the conditional heterogeneity of the timing of the property rights reform. A potential concern arises if more developed regions adopted the reform earlier than less developed regions. The estimated reform effects can be biased if socioeconomic characteristics are associated with the trends of outcome variables. To evaluate whether and to what extent this affects our main results, we limit the sample to pre-reform years and the year of reform implementation and regress the year in which each province started implementing the reform on key provincial socioeconomic characteristics (de Janvry et al., 2015; Chari et al., 2021), local gender culture (Almond et al., 2019),[9] and differential opportunities for men and women in the baseline year. The results show that the local reform timing is not associated with any factors (Appendix Table A2).

Moreover, we preliminarily test the essential parallel trend assumption underlying the DID estimation. We regress each variable regarding economic condition, amenity, and government expenditure on the reform dummy to examine whether any socioeconomic changes are associated with the reform (Appendix Table A3). We find that all coefficients are statistically insignificant. In addition, we conduct an event study, discussed further in Section 6. We find no significant trends in the outcomes of interest before the land reform. This indicates that changes in off-farm employment in the province do not drive the timing of reform implementation.

In Section 6, we conduct a series of robustness checks to confirm the validity of our empirical strategy. We add different controls and rule out the confounding effects of other policies. In addition, we conduct a placebo test by randomly allocating provinces to each land reform wave. The placebo reform rollout is unable to pick up the significant gender effect.

4. Results

Table 2 presents the estimated results of Eq. (1). We control for individual fixed effects across all specifications to rule out the

[9] Provincial socioeconomic characteristics include GDP per capita, proportion of secondary industries, and proportion of tertiary industries in the baseline year. Proxies for differential opportunities include the share of men working in construction and wage gap. Proxies for local gender culture include the gender ratio and gender norms.

Table 2
Land reform and off-farm work.

Dependent variable	Have off-farm work (=1)		Off-farm work hours (off-farm work hours>0)		Log (off-farm work hours)	
	(1)	(2)	(3)	(4)	(5)	(6)
Land reform (β_1)	0.0595***	0.0594***	0.275***	0.263***	0.0307**	0.0284*
	(0.0171)	(0.0164)	(0.0729)	(0.0759)	(0.0127)	(0.0134)
Land reform × Female (β_2)	−0.0554***	−0.0545***	−0.240*	−0.239*	−0.0275*	−0.0273*
	(0.0135)	(0.0133)	(0.129)	(0.131)	(0.0143)	(0.0146)
Observations	19,419	19,419	4164	4164	4164	4164
Adjusted R-squared	0.511	0.511	0.274	0.273	0.259	0.258
Mean of the dependent variable	0.365	0.365	8.290	8.290	2.081	2.081
Demographics	No	Yes	No	Yes	No	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Individual FE	Yes	Yes	Yes	Yes	Yes	Yes

Note: This table presents results from regressions of indicators for having off-farm work (Columns 1 and 2), average daily off-farm work hours (conditional on off-farm work hours>0) (Columns 3 and 4), and log off-farm work hours (Columns 5 and 6). All regressions control for individual and year fixed effects. Columns 2, 4, and 6 also control for demographics, including age, and indicators for educational attainment. Robust standard errors clustered at the provincial level are reported in parentheses. Data are obtained from the *CHNS*. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

potential for endogeneity due to time constant individual-specific unobservable factors, such as ability. We further control for demographic attributes in columns 2, 4 and 6. Columns 1 and 2 show the gender-specific off-farm labor market effect through the extensive margin, where the outcome variable indicates whether individuals are off-farm employed. Both the land reform indicator and interaction term are statistically significant across all the specifications.[10] The coefficients imply that the land reform increases the probability of working in off-farm sectors by 5.94 percentage points for men in contrast to only 0.49 percentage points for women, indicating a gender gap of 5.45 percentage points. As 35.6 % of rural male workers and 25.1 % of rural female workers are off-farm employed, the point estimates represent a 16.69 % increase at the mean for men, while only a 1.95 % increase for women.

Columns 3–6 present the intensive margin of land reform on off-farm labor supply. We limit our sample to those who have a job. The dependent variables include both the levels of daily work hours (columns 3 and 4) and the logarithm of the daily work hours (columns 5 and 6). We find a pattern similar to the extensive margin: land property rights reform leads to increased off-farm labor supply, but this effect is statistically smaller for women, and this difference is statistically significant. Columns 3–4 show that the estimates of the interaction term are around −0.24, and those of the land reform indicator are around 0.27. Combining the two-parameter estimates implies that implementing the property rights law increases the daily work hours by 0.27 h for men, while only increases 0.03 h for women. These findings are consistent across columns 5 and 6. The results show that men's off-farm working hours increased by 2.84 %, while women's only increased by 0.11 %.

Having shown that land reform led to gender-differentiated effects on off-farm labor market supply, we now turn to the question of the effects of land reform on labor supply to other labor activities such as farm work, or household and childcare work. Ceteris paribus, as labor supply to off-farm labor market increases, a reduction in other labor activities is expected. In other words, as the increase in women's off-farm work hours is smaller than that of men after the land reform, do they spend more time in farm work, in household and childcare work, or in the combination of both activities? To explore these issues, we modify Eq. (1) by replacing the dependent variable with: (1) a dummy variable indicating involvement in farm work (defined as having non-zero farm work hours), (2) hours in farm work (conditional on working in farm work), (3) logarithm of hours in farm work, (4) a dummy variable indicating involvement in household and childcare work (defined as having non-zero household and childcare work hours), (5) hours in household and childcare work (conditional on working in household and childcare work), and (6) the logarithm of hours in household and childcare work.

The regression results are presented in Table 3, with each column corresponding to one of the six dependent variables in the same order as introduced above. Both the coefficient estimates of land reform and the interaction term are statistically insignificant in Column 1, suggesting that land reform does not affect the probability that farmers work in the agricultural sector. However, columns 2 and 3 indicate that land reform causes a 10 % reduction in average hours spent on farm work without gender differences. The reduction in farm work hours is statistically significant at the 1 % level. Therefore, land reform encourages men and women to shift their labor away from the agricultural sector and into non-agricultural sectors.

Columns 4–6 in Table 3 show that land reform increases both the likelihood (Column 4) and the amount of time women dedicate to housework and childcare (Columns 5 and 6). Women increase the time spent on housework and childcare by 21.8 % per day, while men are inclined to decrease the time allocated to housework and childcare by 30.7 %. These gender-differentiated changes in housework and childcare induced by the land reform can be reconciled with the finding on gender heterogeneity in employment transformation toward the off-farm sector for rural individuals.

Having shown the gender-specific labor supply effects of the land reform, we now shift our attention to the understanding of the

[10] β_3 (coefficient for the female dummy) is dropped due to collinearity, as we control for individual fixed effects.

Table 3
Effects of land reform on farm work and housework.

Dependent variable	Have farm work (=1) (1)	Farm work hours (2)	Log (farm work hours) (3)	Do housework and childcare (=1) (4)	Daily housework and childcare hours (5)	Log (daily housework and childcare hours) (6)
Land reform (β_1)	−0.0183	−0.521***	−0.0998***	0.0119	−3.613***	−0.307***
	(0.0331)	(0.0667)	(0.0200)	(0.0118)	(0.693)	(0.0765)
Land reform × Female (β_2)	−0.0111	0.132	0.0162	0.0262***	4.809***	0.525***
	(0.0165)	(0.109)	(0.0232)	(0.00779)	(0.468)	(0.115)
Observations	26,431	9395	9395	23,813	15,872	15,872
Adjusted R-squared	0.300	0.207	0.245	0.462	0.441	0.470
Mean of the dependent variable	0.355	5.357	1.535	0.703	7.560	1.284
Demographics	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Individual FE	Yes	Yes	Yes	Yes	Yes	Yes

Note: This table presents results from regressions of an indicator for engaging in farm work (Column 1), average daily farm work hours (conditional on farm work hours>0) (Column 2), log farm work hours (Column 3), an indicator for doing daily housework and childcare (Column 4), average daily housework and childcare hours conditional on hours>0 (Column 5), and log housework and childcare hours conditional on hours>0 (Column 6). All regressions control for individual and year fixed effects. Demographics include age and indicators of educational attainment. Robust standard errors clustered at the provincial level are reported in parentheses. Data are obtained from the *CHNS*. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

income effect of the reform. More specifically, we inquire whether the gender-specific labor supply changes are accompanied by an increasing earnings gap between men and women. Table 4 provides evidence supporting this inquiry. We distinguish between non-farm income[11] and farm income, the latter of which refers specifically to the total annual income received from collective and household farming for an individual.[12] As labor income contains zero values, we follow Chari et al. (2021) by adopting the inverse hyperbolic sine (IHS) function. It is an alternative to log transformations when dealing with variables that encompass zero values.[13] In line with the results on labor supply, land reform also leads to a gender-specific increase in off-farm income. Specifically, land reform causes men's hourly off-farm income to increase by 28.6 %, while only by 9.3 % for women. Meanwhile, we show that land reform causes a significant decrease in agriculture income, both for men and women, which is also consistent with the documented losses in agriculture labor supply.

Overall, these results show that land reform causes gender inequality in the labor market. Although men and women tend to shift their labor away from the agricultural sector into non-agricultural sectors following the land reform, women lag behind men. It is likely to be accompanied by gender-specific variations in the time spent on housework and childcare.

5. Mechanisms: why are women left behind?

We find that men shifted their labor supply to off-farm labor supply more than women due to the reform. A natural next question is why. This section explores the mechanisms underlying the link between land reform and gender-differentiated labor reallocation responses. Our results could be due to five mechanisms: labor market opportunities, child penalties, mobility costs for women, human capital gender gap, and gender identity norms. We examine each of these channels to determine their role and find evidence supporting for the first mechanism: rural workers are predominantly engaged in blue-collar jobs that offer more employment opportunities and higher wages to men.

5.1. Labor market conditions

Are fewer Labor Market Opportunities a Possible Mechanism? One potential factor contributing to the gender-differentiated effects of RLCL on labor reallocation could be the occupational differences between men and women (Blau and Kahn, 2017). As noted in numerous studies, women's participation in the labor market is highly associated with the structural transformation of the industry, given that men and women might have differential endowments of 'brawn' and 'brain' (Olivetti and Petrongolo, 2014). Some studies have also indicated that men are more concentrated in blue-collar occupations,[14] typically involving heavy physical work in manufacturing, construction, mining, or maintenance sectors, while women are more likely to be in administrative support and service occupations (Blau and Kahn, 2017; Goldin, 2014; Olivetti and Petrongolo, 2014, 2016).

[11] We use individual hourly salary, which is a more precise measure than monthly and yearly income.
[12] This does not include individual agricultural labor income (rather than farm profits) from working on one's land.
[13] The IHS function is similar to a log transformation but with zero values explicitly defined. Hence, we use it for continuous results whose distribution includes zeros.
[14] Men have advantages in these energy-intensive activities (Pitt et al., 2012).

Table 4
Effects of land reform on non-farm and farm income.

Dependent variables	IHS Non-farm income		IHS farm income	
	(1)	(2)	(3)	(4)
Land reform (β_1)	0.292***	0.286***	−0.375*	−0.388**
	(0.0666)	(0.0652)	(0.178)	(0.171)
Land reform × Female (β_2)	−0.194**	−0.193**	0.212	0.210
	(0.0732)	(0.0707)	(0.204)	(0.205)
Observations	13,684	13,684	19,705	19,705
Adjusted R-squared	0.391	0.391	0.379	0.379
Mean of the dependent variable	1.175	1.175	3.120	3.120
Demographics	No	Yes	No	Yes
Year FE	Yes	Yes	Yes	Yes
Individual FE	Yes	Yes	Yes	Yes

Note: The dependent variables are the IHS function of average hourly non-farm income (Columns 1 and 2) and the IHS function of average annual farm income (Columns 3 and 4). All regressions control for individual and year fixed effects. Columns 2 and 4 control for demographics, including age, and indicators for educational attainment. Robust standard errors clustered at the provincial level are reported in parentheses. Data are obtained from the *CHNS*. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

Table 5
Potential mechanisms.

Dependent variable	Have off-farm work (=1)				
	(1)	(2)	(3)	(4)	(5)
Land reform	0.0698***	0.0647***	0.0619***	0.0591***	0.131***
	(0.0188)	(0.0167)	(0.0156)	(0.0178)	(0.0141)
Land reform × Female	−0.0332***	−0.0308***	−0.0679***	−0.0542***	−0.103**
	(0.0105)	(0.00813)	(0.0211)	(0.0131)	(0.0366)
Land reform × Female × High initial wage gap	−0.0388*				
	(0.0202)				
Land reform × Female × High initial share of men working in construction and manufacturing		−0.0480**			
		(0.0209)			
Land reform × Female × Married with children			0.0212		
			(0.0205)		
Land reform × Female × Above middle school				−0.0666	
				(0.136)	
Land reform × Female × Gender norms (an indicator for group I with lowest values)					0.0610
					(0.0384)
Land reform × Female × Gender norms (an indicator for group II with middle values)					0.0565
					(0.0375)
Observations	19,419	19,419	19,419	19,419	19,419
Adjusted R-squared	0.512	0.511	0.511	0.511	0.512
Mean of the dependent variable	0.365	0.365	0.365	0.365	0.365
Demographics	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Individual FE	Yes	Yes	Yes	Yes	Yes

Note: The dependent variable is an indicator of off-farm employment. All regressions control for year fixed effects and individual fixed effects. All regressions control for demographics, including age, and indicators for educational attainment. "High initial wage gap" and "High initial share of men working in construction and manufacturing" are dummies representing outside options. "High initial wage gap" is a baseline variable obtained from the 1991, 1993, and 1997 wave of the *CHNS*. "High initial share of men working in construction and manufacturing" is obtained from the 2000 Population Census. Based on the initial province-level wage gap and share of men working in construction and manufacturing, we calculate dummies of whether these variables are above average and interact them with land reform and the female dummy. "Married with children" represents the marital and family status, which equals 1 for married people with children, and 0 otherwise. "Above middle school" represents education attainment, which equals 1 when the person's educational level is above middle school, and 0 otherwise. Gender identity norms are constructed using principal component analysis for five questions by province: "Men should focus on society, and women should focus on family" (Agree=1), "Men are inherently more capable than women" (Agree=1), "Women should marry well rather than do well in their careers" (Agree=1), "The biggest role of women is to produce children" (Agree=1), "The status of women will not improve with economic development" (Agree=1). To avoid endogeneity, we use the proxy for gender identity norms at the provincial level in the baseline year of 2000. We treat them as dummy variables for terciles, with the highest value group referring to more apparent gender discrimination. All interaction terms between the above mechanism dummy variables and land reform term are included in the regression. However, due to space constraints, we do not display them individually in the study. Robust standard errors clustered at the provincial level are reported in parentheses. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

To test the channel of outside options, we include all the interactions of the reform dummy and female dummy with baseline characteristics of the city that proxy for differential opportunities for men and women, such as the province-level wage gap and the share of men working in construction and manufacturing. Columns 1–2 of Table 5 present the results. We find that the triple interaction terms are all significantly negative. In other words, women face fewer opportunities for non-agricultural employment following the land reform in regions initially characterized by large gender wage gaps and higher male dominance in manufacturing and construction sectors. This indicates that gender-specific labor market opportunities are, at least partially, responsible for the gender-differentiated labor reallocation following the reform.

5.2. Other mechanisms

Is Child Penalty a Possible Mechanism? Parenthood consistently exerts unequal impacts on men and women (Goldin, 2014; Shi and Shen, 2023). Numerous studies indicate that men's and women's earnings evolve similarly before parenthood but diverge sharply after the arrival of children (Kleven et al., 2019). Women consistently perform a higher share of childcare (Maurer-Fazio et al., 2011). Consequently, women are likely to reduce their labor supply on both extensive and intensive margins after childbirth (Paull, 2008; Bertrand et al., 2010; Fitzenberger et al., 2013; Schönberg amd Ludsteck, 2014; Halla et al., 2020).

To further explore whether differential childcare roles explain why the reform affected men and women in different ways, we perform a heterogeneity analysis by parental status. In Column 3 of Table 5, we introduce all interactions of the parental dummy with the reform and female dummies to examine whether the effect differs by parental status. More specifically, we aim to determine whether women with children are less likely to increase off-farm work. However, we find that this is not the case. The triple interaction term is also insignificant, suggesting that childbearing is not likely a contributor to relatively lower off-farm employment for women. These results help to rule out the child penalty channel.

Is the lower Education Levels for Women a Possible Mechanism? Another related channel to consider is the disparity in education levels between men and women. On average, we observe that rural women are less educated than rural men (Appendix Table A4), which aligns with findings from related research on the gender human capital gap (Brown and Park, 2002). While one might anticipate that rural women's continued engagement in agricultural and housework post-reform could be due to their relatively lower education levels, our analysis suggests otherwise.

In Column 4 of Table 5, we allow the gender-specific effect to vary by education levels to investigate the potential role of the gender human capital gap. We introduce all interaction terms between education level,[15] year, and gender into our analysis. The results indicate that the triple interaction term is insignificant, suggesting that educational attainment is not a significant factor for both men and women. This finding helps to rule out the potential influence of the gender human capital gap and aligns with Molina (2021), who demonstrates that schooling and cognitive ability are less complementary in blue-collar jobs, which are typically favored by men.

Are Gender Norms toward Female Employment a Possible Mechanism? China, deeply influenced by Confucianism, has a long history of son preference (Murphy et al., 2011) and a social norm where the husband is considered the primary breadwinner of the household (Ye and Zhao, 2018). If women are anticipated to shoulder the primary responsibility for child and family caretaking, their participation in the labor market post-reform might be less likely. Given the temporal flexibility is often a top priority for women (Blau and Kahn, 2017), it would be rational for women to avoid entering the off-farm sector.

To investigate this hypothesis, we use the CWSS data to create a proxy for gender identity norms which are a set of dummy variables for terciles.[16] Specifically, we introduce all interactions between gender identity norms dummies, land reform, and gender dummy, examining how the triple interaction term affects off-farm employment. Column 5 in Table 5 reveal that gender identity norms do not seem to play a significant role in the gender-specific effects of RLCL on the number of annual off-farm days.[17]

Is higher Mobility Cost for Women a Possible Mechanism? The mobility factor, especially when children are present, could be a possible explanation for the gender disparity in participation in non-farm employment post-reform. Unfortunately, we are not able to directly tackle this mechanism due to the lack of data on whether the off-employment is local or migratory. However, we can explore this issue indirectly using the CLDS data. The CLDS 2014 survey includes a retrospective history of locations for respondents (Gao et al., 2022), allowing us to create an individual-level longitudinal panel. Using the longitude data, we re-estimate Eq. (1), and the results are reported in Table 6.

We do not observe significant evidence of an increase in the probability of migration after the reform. The absence of such evidence could be partially attributed to the institutional barriers to migration in China (Chari et al., 2021). Importantly, these results suggest that, following the reform, most individuals tend to secure off-farm employment near their homes rather than traveling far to work, which might be influenced by other pull factors like *hukou* reform and the development of industrialization in cities (Gao and Zhang, 2023). Moreover, there are no observed gender-specific differentials in this pattern. Hence, whether the higher mobility cost for women is a mechanism becomes less relevant in the context of this study.

[15] Education level is defined as an indicator for high school and above education.

[16] We use principal component analysis for five questions by province: "Men should focus on society, and women should focus on family" (Agree=1), "Men are inherently more capable than women" (Agree=1), "Women should marry well rather than do well in their careers" (Agree=1), "The biggest role of women is to produce children" (Agree=1), and "The status of women will not improve with economic development" (Agree=1). To avoid endogeneity, we use the proxy for gender identity norms at the provincial level in the baseline year 2000.

[17] This could be because grandparents often care for children in rural China (Chang et al., 2011). Mothers in these circumstances are likely to be less constrained by gender identity norms.

Table 6
Land reform and migration.

	Dependent variable: Migrate (=1)	
	(1)	(2)
Land reform (β_1)	0.000686	0.000684
	(0.000910)	(0.000915)
Land reform × Female (β_2)	−0.000817	−0.000779
	(0.000564)	(0.000574)
Observations	150,805	150,805
Adjusted R-squared	0.0590	0.0584
Mean of the dependent variable	0.004	0.004
Cohort FE	No	Yes
Year FE	Yes	Yes
Individual FE	Yes	Yes

Note: The dependent variable indicates whether the individual migrated in the currents year. All regressions control for year and individual fixed effects. Column 2 also controls for cohort fixed effects. Robust standard errors clustered at the provincial level are reported in parentheses. Data are obtained from the 2014 CLDS. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

In sum, the disadvantaged labor market condition for women provides the most credible explanation for the gender-differentiated effects on off-farm employment. Given the gender gaps in labor market conditions, women will likely lag behind when joining the non-farm sector. We do not find direct evidence to support other mechanisms.

6. Sensitivity analyses

We perform several robustness checks to examine the sensitivity of the gender-specific effects of the RLCL on labor reallocation.

6.1. Additional controls and specifications

We assess whether our empirical results are robust by controlling for potential confounders. One potential confounder is industry composition. Olivetti and Petrongolo (2014) reveal that international differences in industry structure explain approximately one-third of the overall cross-country variation in wage and hour gender gaps. In addition, differential time trends in gender identity norms across provinces could confound the results. To address these concerns, we assess the robustness of our results to local economic controls, industry composition, and gender identity norms at the baseline by interacting them with year fixed effects and then with year-by-gender fixed effects in Table A5. The parameters of land reform are more precisely estimated, when controlling for other factors, the effects of the reforms seem to be larger,[18] but the gender gap in the effects continue to be very large across specifications. Thus, our empirical baseline patterns are not influenced by gender-specific time trends.

6.2. Other reforms

Another concern with DID estimates is the possibility that, alongside the land reform, other reforms occurring during the same period could have compounding effects, introducing bias to the true impact of land reform. This concern arises primarily from the implementation of other reforms following the implementation of the RLCL.

A significant change for rural households during this period is the elimination of the agricultural tax. The central government initiated a rural tax-and-fee reform in 2000, intending to replace all agricultural taxes and fees with a uniform agricultural tax. This reform was followed by the gradual elimination of the agricultural tax in 2004, with a targeted goal to eliminate it entirely within five years. However, nearly all provinces had already abolished the agricultural tax by the end of 2005 (Wang and Shen, 2014; Chen, 2017). We address this concern in two ways. First, Wang and Shen (2014) report that the abolition of the agricultural tax does not lead to a substantial increase in farmers' net income, indicating that agricultural tax reform might not significantly impact labor and land reallocation in rural China. Second, the effect of rural tax reform could be gender-specific. For instance, the elimination of agricultural tax may first occur in developed areas that are favorable to females. To further mitigate this concern, we control for all the interactions between an indicator of whether a province conducted the tax reform in a specific year and with a female dummy. As shown in Appendix Table A6, the land reform continues to exhibit a significant and positive effect on off-farm employment, with a relatively minor effect for women. As the coefficients of interest remain identical to those in Table 2, the results suggest that the agricultural tax reform does not contaminate the gender-differentiated RLCL effect on off-farm employment.

The second change during this period is China's accession to the WTO. The determinants of structural change encompass push and

[18] There could be labor market shocks that interact with gender due to the differential distribution of industries across cities. For example, technological shocks related to automation were common in the 2000s (Keller and Utar, 2022), which could mitigate the effect of land reform. Moreover, it could have a lesser impact on women if it affects more negatively jobs in which men are disproportionately working.

pull factors (Erten and Leight, 2021). Although our study primarily focuses on push factors, trade liberalization, acting as a significant pull factor, could potentially bias our results. Specifically, trade liberalization may lead to an overestimation of the labor reallocation effect. To address this concern, we follow Erten and Leight (2021) and Khanna et al. (2021) and control for the Normal Trade Relations Gap (NTR gap). Before 2002, the maintenance of China's Most Favored Nation (MFN) status in the US required regular Congressional approval, leading considerable uncertainty with annual renewals. However, with China's accession to the WTO, the US permanently granted China a NTR status. Following the approach of Khanna et al. (2021), we construct a city-level exposure measure—the average gap between NTR and non-NTR rates across products—weighted by the industry export share in each city during the baseline year. Once again, we control for all interactions of the city NTR gap, a post-WTO dummy, and a female dummy to control for the gender-specific effects of the WTO accession. As shown in Appendix Table A6, the results indicate that our findings remain economically large and statistically significant.

6.3. Event study

To assess the parallel trend assumption, we employ an event study analysis using the following equation:

$$Y_{it} = \sum_{k=-3}^{7} \beta_k Reform_{p,t,k} + \sum_{k=-3}^{7} \alpha_k Reform_{p,t,k} \times Female_i + Female_i + \delta_t + \eta_i + \varepsilon_{it} \tag{2}$$

where $Reform_{p,t,k}$ indicates the time period relative to the reform year in province p.[19] For instance, $Reform_{p,t,-2}$ represents two years before the implementation of the reform in province p, and $Reform_{p,t,2}$ represents two years after implementation (Chari et al., 2021). We introduce interactions between $Reform_{p,t,k}$ and the gender dummy, as specified in Eq. (1) to examine whether the results differ by gender at each period. Eq. (2) allows us to test whether reform implementation is exogenous to our primary outcomes of interest.

Fig. 1 illustrates the estimated α_k values. No significant trends are observed in the number of off-farm days by gender before the reform. However, the interaction term for policy implementation and the gender dummy consistently remains below zero for six years post-implementation. These results further support (i) the assumption that the gender difference in off-farm employment would have remained the same over time if RCLC had not been implemented and (ii) the existence of gender-differentiated effects of RCLC on off-farm employment, as men are more likely than women to be employed off-farm.

6.4. Additional technical tests

6.4.1. Placebo tests

To test whether our main results are affected by potential omitted variables, we conduct a placebo test by randomly allocating provinces to each land reform wave. As this allocation is a random process, the false treatment variable is expected to have no impact on the outcome of interest. In this scenario, we do not anticipate gender-differentiated effects, which would indicate a misspecification of the DID estimation (Li et al., 2016). Using the exact specification as in Eq. (1), Panel A of Appendix Table A7 demonstrates that this false reform timing has no effect on off-farm employment for either men or women. This result suggests that unobservables do not drive the gender-differentiated effects of land reform on labor reallocation.

6.4.2. Wild cluster bootstrap

Considering the limited number of provinces in our sample (12), we employ the wild cluster bootstrap method (Roodman et al., 2019) to address the challenge of a small number of clusters (Deininger et al., 2021). Panel B of Appendix Table A7 presents the results, with figures in square brackets denoting p-values from wild cluster bootstrap (Roodman et al., 2019) with Rademacher weights and 1000 replications. The results align with the baseline results presented in Table 2, affirming the consistency and robustness of our results.

6.4.3. Issues of staggered did

Recent developments in the econometrics literature have drawn attention to the "negative weight problem" in staggered DID design (de Chaisemartin and D'Haultfoeuille, 2020; Goodman-Bacon, 2021; Baker et al., 2022). If treatment effects evolve, some units may receive negative weights when their parameters are aggregated to form average treatment effects, leading to potential biases in estimates. To investigate this possibility, we compute a weighted average of the policy-implementing wave-specific point estimates with weights equal to each wave's share of treated observations (Callaway and Sant'Anna, 2021). The point estimates in Panel C of Appendix Table A7 are highly consistent with our baseline results.

7. Conclusions

Despite the rapidly increasing evidence of the effects of property rights reform on the reallocation of factors, such as land and labor, few studies have examined gender differences. This study exploits a quasi-natural experiment in China to explore the gender-differentiated effects of land property rights reforms on labor reallocation. The results indicate that the RLCL, which allows farmers

[19] As many reforms are implemented late in the year (October and November), we focus on the year after implementation.

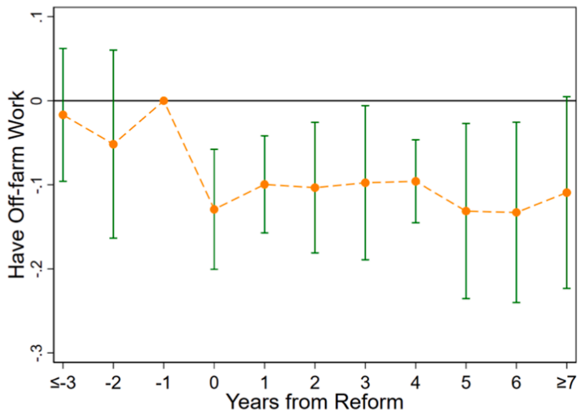

Fig. 1. Leads and Lags of Reform Effects on Off-farm Days
Note: The graphs plot the coefficients and associated 95 % confidence intervals from estimating the leads and lags of land reform on off-farm employment. All the effects are relative to the year before the reform ($k=-1$). Data are obtained from the *CHNS*.

to lease out their land, increases off-farm labor participation among men and women, suggesting a relocation of labor from the primary (farm) sector to the secondary and tertiary sectors. However, women appear to lag behind men when transitioning into the off-farm sector, as we identify a noticeable gender gap in the off-farm employment increase. This situation is mainly driven by the disadvantages in the labor market for women. Furthermore, women encounter fewer employment opportunities in the non-agricultural sector than men. The reform results in a sharp decline in agricultural job numbers without a simultaneous emergence of other employment opportunities considered suitable for women. We do not find evidence of the effects of other channels, such as child penalties, high mobility costs, gender human capital gap, and gender identity norms.

Our study is among the first to highlight the role of gender in labor reallocation among sectors in response to improved land property rights. In developing countries, where land remains the critical source of livelihood, securing land property rights may lead to an efficient allocation of labor, thereby displacing farmers out of agriculture. However, neglecting the differential effects between men and women may increase gender inequality during the labor-shifting process. This study sheds light on the reasons women lag behind men in the non-agricultural sector following the land reform. A deeper analysis of whether and how gender-differentiated responses would affect household welfare and the implications for agricultural transformation in a gender-specific framework is a worthwhile avenue for future research.

Declaration of competing interest

The authors declare no conflict of interest.

Data availability

Data will be made available on request.

Funding

This paper was supported by Zhejiang Province Philosophy and Social Science Planning Zhijiang Youth Special Project: Research on the Mechanism and Path of Increasing Income for Low Income Rural Population under the Background of Common Prosperity (Grant No.24ZJQN042YB), National Natural Sciences Foundation of China (Grant No. 72003170), Humanities and Social Sciences Program of the Ministry of Education of China (Grant No. 22JJD790077), National Social Science Fund of China (Grant No. 21&ZD091), the Major Program of National Fund of Philosophy and Social Science of China (Grant No. 23ZDA035), the Fundamental Research Funds for the Central Universities in China (Grant No. S20230210; S20220191), Philosophy and Social Science Research Special Project of the Ministry of Education of China (Grant No. 23JD20105), the National Natural Science Foundation of China Grant 71773110, Michigan State University's AgBioResearch (Grant MICL02608), and Zhejiang Province Philosophy and Social Science Planning Project (Grant 24YJRC01ZD-3YB).

Appendix: Additional Figures and Tables

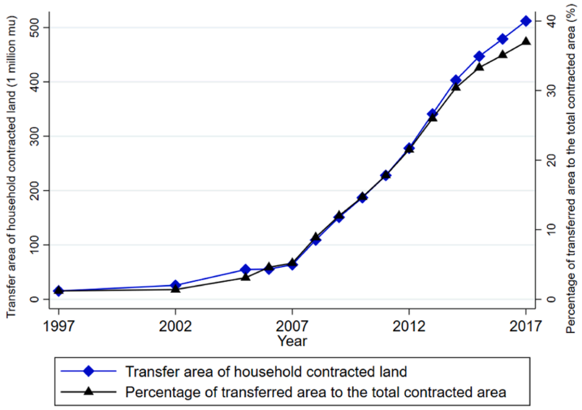

Fig. A1. Land Transfer in China (1997–2017)
Note: This figure shows rural land transfer in China between 1997 and 2017. The y-axis on the left represents the total transfer area of household contracted land each year, and the y-axis on the right represents the percentage of transferred area to the total contracted area each year. The data are from Zhou et al. (2020).

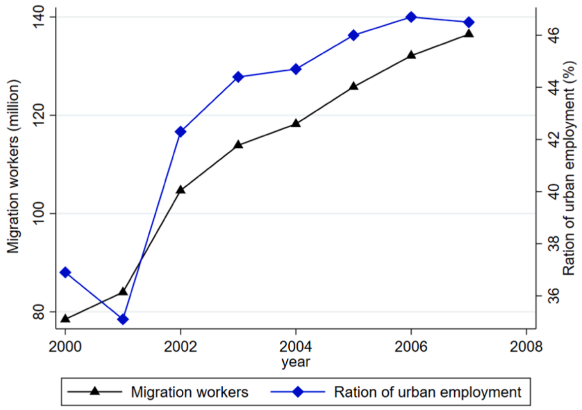

Fig. A2. Migrant Workers and Urban Employment in China (2000–2007)
Note: This figure shows migrant workers and urban employment in China from 2000 to 2007. The y-axis on the left represents the migrant workers each year, and the y-axis on the right represents the ratio of migrant workers to total urban employment. The data are from Cai et al. (2009).

Table A1
Rural land contracting law announcement by province.

Province	Year	Document Name	Issue Date	Effective Date
Shanghai	2003	Hu Fu Fa (2003) No. 29	04/25/2003	04/25/2003
Jiangsu	2004	Jiangsu Province Government Order (2003) No. 21	12/18/2003	02/01/2004
Heilongjiang	2004	Hei Zheng Ban Fa (2004) No.17	05/13/2004	05/13/2004
Beijing	2004	Jing Fa (2004) No.17	08/26/2004	08/26/2004
Henan	2004	Rules for the Transfers of Rural Land Contract and Management Rights in Henan Province	09/17/2004	09/17/2004
Hunan	2004	Hunan Province People's Congress Standing Committee (2004) No. 35	07/30/2004	10/01/2004
Shandong	2004	Shandong Province People's Congress Standing Committee (2004) No. 37	07/30/2004	10/01/2004
Liaoning	2005	Liaoning Province People's Congress Standing Committee (2005) No. 28	01/28/2005	04/01/2005
Guizhou	2005	Notice of the Guizhou Provincial Agriculture Committee on Issuing the Special Inspection Plan for the Implementation of the Province's Rural Land Contracting Laws and Policies	04/26/2005	04/26/2005
Guangxi	2006	Gui Zheng Ban Fa (2006) No. 141	11/14/2006	11/14/2006

(continued on next page)

Table A1 (*continued*)

Province	Year	Document Name	Issue Date	Effective Date
Chongqing	2007	Chongqing Municipality People's Congress Standing Committee (2007) No. 6	04/02/2007	07/01/2007
Hubei	2012	Hubei Province People's Congress Standing Committee (2012) No. 138	07/27/2012	10/01/2012

Note: Only the 12 provinces with complete information from the *CHNS* are included.

Table A2
Determinants of provincial reform timing.

	(1)	(2)	(3)
Share of Men Working in Construction	−0.825	−5.535	−50.79
	(5.479)	(6.945)	(56.03)
Wage Gap	−0.000346	−0.000247	−0.00129
	(0.000593)	(0.000825)	(0.00198)
Gender Ratio		19.43	−10.96
		(17.30)	(115.1)
Gender Norm		−1.171	−0.230
		(12.37)	(32.81)
GDP per capita			−3.894
			(4.142)
Proportion of Secondary Industry			1.567
			(1.657)
Proportion of Tertiary Industry			1.880
			(2.093)
Joint F-test	0.885	0.497	0.731

Note: The dependent variable is provincial reform year. Each observation is a provincial year. Share of Men Working in Construction is obtained from the 2000 Population Census. Wage Gap is calculated based on the wage gap between men and women based on the 2000 *CHNS.* Gender Ratio is calculated based on the rural population in the 2000 Population Census. Gender Norm is an indicator for agreeing that "Men should focus on society, and women should focus on family" in the 2000 *China's Women Social Status Survey (CWSS).* Proportion of Secondary Industry and Proportion of Tertiary Industry represent the proportion of the secondary and tertiary industries in GDP, respectively, which are obtained from the *2000 China Statistical Yearbook.* *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

Table A3
Economic condition, amenity, and government expenditure before and after land reform.

	(1)	(2)
Panel A: Economic Condition		
GDP	−1.535	2.992
	(1.803)	(1.827)
Manufacturing Employment	1347	934.9
	(1306)	(916.7)
Construction Employment	−459.5	−1866
	(449.9)	(2028)
Accommodation and Catering Employment	−5.083	15.14
	(13.02)	(10.55)
Rural Electricity Consumption	−0.0478	0.0323
	(0.0412)	(0.0369)
Wage	−758.8	119.7
	(2946)	(1912)
Panel B: Amenity		
No. of Health Institutions	−1.105	−0.172
	(3.392)	(3.450)
No. of Senior High Schools	0.0161	0.00513
	(0.0251)	(0.0166)
No. of Primary Schools	0.0650	−0.464
	(0.837)	(0.777)
Panel C: Government Expenditure		
Government Budget Expenditure	−0.169	0.228
	(0.221)	(0.152)
Province FE	Yes	Yes
Year FE	Yes	No
City-tier × Year FE	No	Yes

Note: Row names correspond to dependent variables. The independent variable is an indicator of whether the reform was implemented in a province. Robust standard errors clustered at the provincial level are reported in parentheses. Data are obtained from the *2000–2013 China City Statistical Yearbook*. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

Table A4
Gender human capital gap.

	Dependent variable: Years of schooling			
	(1)	(2)	(3)	(4)
Female	−1.253***	−1.198***	−1.210***	−1.210***
	(0.187)	(0.185)	(0.185)	(0.185)
Observations	21,624	21,624	21,624	21,624
Adjusted R-squared	0.393	0.488	0.492	0.492
Mean of the dependent variable	5.554	5.554	5.554	5.554
Demographic	No	Yes	No	Yes
Household FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Cohort FE	No	No	Yes	Yes

Note: The dependent variable is years of schooling. Demographics include age. Robust standard errors clustered at the provincial level are reported in parentheses. Data are obtained from the *CHNS*. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

Table A5
Additional controls.

	Dependent variable: Have off-farm work (=1)	
	(1)	(2)
Land reform	0.0872***	0.0962***
	(0.0129)	(0.0128)
Land reform × Female	−0.0589***	−0.0798***
	(0.0137)	(0.00508)
Observations	19,419	19,419
Adjusted R-squared	0.517	0.518
Mean of the dependent variable	0.365	0.365
Demographics	Yes	Yes
Additional Controls ×Year	Yes	Yes
Additional Controls ×Year × Female	No	Yes
Year FE	Yes	Yes
Individual FE	Yes	Yes

Note: The dependent variable is an indicator of off-farm employment. All regressions control for year and individual fixed effects. We also control for demographics, including age and indicators for educational attainment, in all columns. We add a series of province-level initial control variables related to economic development, industry composition, gender norms, including GDP per capita, share of the GDP of the secondary industry and the tertiary industry, gender ratio, and terciles of gender norms variable. We control for all the indicators with year (Column 1), and all the interactions with year and female dummy (Column 2). Robust standard errors clustered at the provincial level are reported in parentheses. Data are obtained from the *CHNS*. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

Table A6
Control for other reforms.

	Dependent variable: Have off-farm work (=1)	
	(1)	(2)
Land reform	0.0546***	0.0451**
	(0.0157)	(0.0166)
Land reform × Female	−0.0550***	−0.0314*
	(0.0139)	(0.0146)
Observations	19,074	19,074
Adjusted R-squared	0.510	0.511
Mean of the dependent variable	0.365	0.365
Agricultural tax reform	Yes	Yes
NTR gap × post-WTO	Yes	Yes
Agricultural tax reform × Female	No	Yes
NTR gap × post-WTO × Female	No	Yes
Demographics	Yes	Yes
Year FE	Yes	Yes
Individual FE	Yes	Yes

Note: The dependent variable is an indicator of off-farm employment. We control for year fixed effects and individual fixed effects for all regressions, and we also control demographics in all regressions, including age and indicators for educational attainment. In Column 1, we control for agricultural tax reform, proxied by an indicator of whether the province conducted a tax reform in a specific year, as well as accession to WTO, proxied by the interaction term between the NTR gap and post-WTO dummy. We allow these two reforms to differ by gender in Column 2 and include the indicator for the agricultural tax reform and the interaction between it and the female dummy, as well as variables for the city NTR gap, post-WTO dummy, and all interactions between them and the female dummy. Robust standard errors clustered at the provincial level are reported in parentheses. Data are obtained from the CHNS. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

Table A7
Additional technical tests.

	Dependent variable: Have off-farm work (=1)	
	(1)	(2)
Panel A: Placebo reform implementation sequence		
Land reform (β_1)	0.0277	0.0279
	(0.0212)	(0.0211)
Land reform × Female (β_2)	−0.0209	−0.0207
	(0.0124)	(0.0124)
Observations	19,419	19,419
Adjusted R-squared	0.510	0.510
Panel B: Wild-cluster Bootstrap		
Land reform (β_1)	0.060*	0.059*
	(0.0172)	(0.0163)
	[0.0660]	[0.0750]
Land reform × Female (β_2)	−0.055***	−0.054***
	(0.0134)	(0.0132)
	[0.0000]	[0.0000]
Observations	19,419	19,419
Adjusted R-squared	0.510	0.511
Panel C: Re-weighted estimates		
Land reform × Female (β_2)	−0.0572***	−0.0558***
	(0.0120)	(0.0116)
Observations	19,419	19,419
Mean of the dependent variable	0.365	0.365
Demographics	No	Yes
Year FE	Yes	Yes
Individual FE	Yes	Yes

Note: In Panel A, we randomly allocated provinces to each wave of land reform. In Panel B, Figures in square brackets are p-values from wild cluster bootstrap (Roodman et al., 2019) with Rademacher weights and 1000 replications. Wild-Bootstrap standard errors are reported in parentheses. Panel C reports the results obtained by estimating the difference-in-differences model wave by wave (provinces treated last are used as the control group) and reweighting the estimated treatment effects by the sample shares in each wave of RLCL implementation (following Callaway and Sant'Anna, 2021). All regressions control for year and individual fixed effects. Column 2 controls for demographics, including age and indicators for educational attainment. Data are obtained from the *CHNS*. *** significant at 1 %; ** significant at 5 %; * significant at 10 %.

References

Almond, D., Li, H., Zhang, S., 2019. Land reform and sex selection in China. J. Polit. Econ. 127 (2), 560–585. https://doi.org/10.1086/701030.

Baker, A.C., Larcker, D.F., Wang, C.C., 2022. How much should we trust staggered difference-in-differences estimates? J. Financ. Econ. 144 (2), 370–395. https://doi.org/10.2139/ssrn.3794018.

Bertrand, M., Goldin, C., Katz, L.F., 2010. Dynamics of the gender gap for young professionals in the financial and corporate sectors. Am. Econ. J.: Appl. Econ. 2 (3), 228–255. https://doi.org/10.1257/app.2.3.228.

Bertrand, M., Kamenica, E., Pan, J., 2015. Gender identity and relative income within households. Q. J. Econ. 130 (2), 571–614. https://doi.org/10.1093/qje/qjv001.

Biasi, B., Sarsons, H., 2022. Flexible wages, bargaining, and the gender gap. Q. J. Econ. 137 (1), 215–266. https://doi.org/10.1093/qje/qjab026.

Blau, F.D., Kahn, L.M., 2017. The gender wage gap: extent, trends, and explanations. J. Econ. Lit. 55 (3), 789–865. https://doi.org/10.1257/jel.20160995.

Brown, P.H., Park, A., 2002. Education and poverty in rural China. Econ. Educ. Rev. 21 (6), 523–541. https://doi.org/10.1016/S0272-7757(01)00040-1.

Cai, F., Du, Y., Wang, M., 2009. Migration and labor mobility in China. hum. dev. res. paper (hdrp) ser. 9 (2009) mpra paper no. 19187. institute of population and labor Economics, Chinese Academy of Social Sciences.

Callaway, B., Sant'Anna, P.H., 2021. Difference-in-differences with multiple periods. J. Econom. 225 (2), 200–230. https://doi.org/10.1016/j.jeconom.2020.12.001.

Card, D., Cardoso, A.R., Kline, P., 2016. Bargaining, sorting, and the gender wage gap: quantifying the impact of firms on the relative pay of women. Q. J. Econ. 131 (2), 633–686. https://doi.org/10.1093/qje/qjv038.

Chang, H.Q., Dong, X.Y., MacPhail, F., 2011. Labor migration and time use patterns of the left-behind children and elderly in rural China. World Dev. 39 (12), 2199–2210. https://doi.org/10.1016/j.worlddev.2011.05.021.

Chari, A., Liu, E.M., Wang, S.Y., Wang, Y., 2021. Property rights, land misallocation, and agricultural efficiency in China. Rev. Econ. Stud. 88 (4), 1831–1862. https://doi.org/10.1093/restud/rdaa072.

Chen, S.X., 2017. The effect of a fiscal squeeze on tax enforcement: evidence from a natural experiment in China. J. Public Econ. 147, 62–76. https://doi.org/10.1016/j.jpubeco.2017.01.001.

Cheng, W., Xu, Y., Zhou, N., He, Z., Zhang, L., 2019. How did land titling affect China's rural land rental market? Size, composition, and efficiency. Land Use Policy 82, 609–619. https://doi.org/10.1016/j.landusepol.2018.12.037.

Chernina, E., Castañeda Dower, P., Markevich, A., 2014. Property rights, land liquidity, and internal migration. J. Dev. Econ. 110, 191–215. https://doi.org/10.1016/j.jdeveco.2013.03.010.

Cortés, P., Pan, J., 2020. Children and the Remaining Gender Gaps in the Labor Market. National Bureau of Economic Research, Cambridge, MA. NBER Working Paper Series (No. 27980). http://www.nber.org/papers/w27980.
Cortés, P., Pan, J., Pilossoph, L., Reuben, E., Zafar, B., 2023. Gender differences in job search and the earnings gap: evidence from the field and lab. Q. J. Econ. 138 (4), 2069–2126. http://doi.org/10.3386/w28820.
de Chaisemartin, C., d'Haultfoeuille, X., 2020. Two-way fixed effects estimators with heterogeneous treatment effects. Am. Econ. Rev. 110 (9), 2964–2996. https://doi.org/10.1257/aer.20181169.
de Janvry, A., Emerick, K., Gonzalez-Navarro, M., Sadoulet, E., 2015. Delinking land rights from land use: certification and migration in Mexico. Am. Econ. Rev. 105 (10), 3125–3149. https://doi.org/10.1257/aer.20130853.
Deininger, K., Jin, S., 2009. Securing property rights in transition: lessons from implementation of China's rural land contracting law. J. Econ. Behav. Organ. 70 (1–2), 22–38. https://doi.org/10.1016/j.jebo.2009.01.001.
Deininger, K., Xia, F., Kilic, T., Moylan, H., 2021. Investment impacts of gendered land rights in customary tenure systems: substantive and methodological insights from Malawi. World Dev. 147, 105654 https://doi.org/10.1016/j.worlddev.2021.105654.
Erten, B., Leight, J., 2021. Exporting out of agriculture: the impact of WTO accession on structural transformation in China. Rev. Econ. Stat. 103 (2), 364–380. https://doi.org/10.1162/rest_a_00852.
Fitzenberger, B., Sommerfeld, K., Steffes, S., 2013. Causal effects on employment after first birth—a dynamic treatment approach. Labour. Econ. 25, 49–62. https://doi.org/10.2139/ssrn.2368230.
Gao, X., Shi, X., Fang, S., 2021. Property rights and misallocation: evidence from land certification in China. World Dev. 147, 105632 https://doi.org/10.1016/j.worlddev.2021.105632.
Gao, X., Song, R., Timmins, C., 2022. The Fertility Consequences of Air Pollution in China. National Bureau of Economic Research, Cambridge, MA. NBER Working Paper Series (No. 30165). http://www.nber.org/papers/w30165.
Gao, X., Zhang, R., 2023. A study on startup policy toolkit of off-farm workers returning hometown in China. Agric. Rural Stud. 1 (3), 0014. https://doi.org/10.59978/ar01030014.
Gerard, F., Lagos, L., Severnini, E., Card, D., 2021. Assortative matching or exclusionary hiring? The impact of employment and pay policies on racial wage differences in Brazil. Am. Econ. Rev. 111 (10), 3418–3457. https://doi.org/10.1257/aer.20181596.
Giles, J., Mu, R., 2018. Village political economy, land tenure insecurity, and the rural to urban migration decision: evidence from China. Am. J. Agric. Econ. 100 (2), 521–544. https://doi.org/10.2139/ssrn.2529339.
Goldin, C., 2014. A grand gender convergence: its last chapter. Am. Econ. Rev. 104 (4), 1091–1119. https://doi.org/10.1257/aer.104.4.1091.
Goldin, C., 2021. Career & Family: Women's Century-Long Journey Toward Equity. Princeton University Press, Princeton NJ.
Gollin, D., Lagakos, D., Waugh, M.E., 2014. The agricultural productivity gap. Q. J. Econ. 129 (2), 939–993. https://doi.org/10.1093/qje/qjt056.
Goodman-Bacon, A., 2021. Difference-in-differences with variation in treatment timing. J. Econom. 225 (2), 254–277. https://doi.org/10.1016/j.jeconom.2021.03.014.
Halla, M., Schmieder, J., Weber, A., 2020. Job displacement, family dynamics, and spousal labor supply. Am. Econ. J.: Appl. Econ. 12 (4), 253–287. https://doi.org/10.1257/app.20180671.
Hare, D., Yang, L., Englander, D., 2007. Land management in rural China and its gender implications. Fem. Econ. 13, 35–61. https://doi.org/10.1080/13545700701445298.
Jarrell, S.B., Stanley, T.D., 2004. Declining bias and gender wage discrimination? A meta-regression analysis. J. Hum. Res. 39 (3), 828–838. https://doi.org/10.2307/3558999.
Judd, E.R., 2007. No change for thirty years: the renewed question of women's land rights in rural China. Dev. Change 38 (4), 689–710. https://doi.org/10.1111/j.1467-7660.2007.00429.x.
Keller, W., Utar, H., 2022. Globalization, gender, and the family. Rev. Econ. Stud. 89 (6), 3381–3409. https://doi.org/10.1093/restud/rdac012.
Khanna, G., Liang, W., Mobarak, A.M., Song, R., 2021. The Productivity Consequences of Pollution-Induced Migration in China. National Bureau of Economic Research, Cambridge, MA. NBER Working Paper Series (No. 28401). http://www.nber.org/papers/w28401.
Kleven, H., Landais, C., 2017. Gender inequality and economic development: fertility, education, and norms. Economica 84 (334), 180–209. https://doi.org/10.1111/ecca.12230.
Kleven, H., Landais, C., Søgaard, J.E., 2019. Children and gender inequality: evidence from Denmark. Am. Econ. J.: Appl. Econ. 11 (4), 181–209. https://doi.org/10.1257/app.20180010.
Li, P., Lu, Y., Wang, J., 2016. Does flattening government improve economic performance? Evidence from China. J. Dev. Econ. 123, 18–37. https://doi.org/10.1016/j.jdeveco.2016.07.002.
Liu, Y., Fang, F., Li, Y., 2014. Key issues of land use in China and implications for policy making. Land Use Policy 40, 6–12. https://doi.org/10.1016/j.landusepol.2013.03.013.
Maurer-Fazio, M., Connelly, R., Lan, C., Tang, L., 2011. Childcare, eldercare, and labor force participation of married women in urban China, 1982–2000. J. Hum. Resour. 46 (2), 261–294. https://doi.org/10.1353/jhr.2011.0011.
Molina, T., 2021. Pollution, ability, and gender-specific investment responses to shocks. J. Eur. Econ. Assoc. 19 (1), 580–619. https://doi.org/10.1093/jeea/jvaa005.
Murphy, R., Tao, R., Lu, X., 2011. Son preference in rural China: patrilineal families and socioeconomic change. Popul. Dev. Rev. 37 (4), 665–690. https://doi.org/10.1111/j.1728-4457.2011.00452.x.
Nelson, S., Sisto, I., Crowley, E., Villarreal, M., 2012. Women in agriculture: closing the gender gap for development. In: Jägerskog, A., Jønch Clausen, T. (Eds.), Feeding a Thirsty World. Stockholm International Water Institute, Stockholm, pp. 25–30. SIWI Report Number 31.
Olivetti, C., Petrongolo, B., 2014. Gender gaps across countries and skills: demand, supply and the industry structure. Rev. Econ. Dyn. 17 (4), 842–859. https://doi.org/10.1016/j.red.2014.03.001.
Olivetti, C., Petrongolo, B., 2016. The Evolution of Gender Gaps in Industrialized Countries. National Bureau of Economic Research, Cambridge, MA. NBER Working Paper Series (No. 21887). http://www.nber.org/papers/w21887.
Öztürk, M., Yalçın, E.C., 2023. Determining factors effecting the population growth rate of villages in Turkey: an econometric perspective. Agric. Rural Stud. 1 (3), 0013. https://doi.org/10.59978/ar01030013.
Paull, G., 2008. Children and women's hours of work. Econ. J. 118 (526), F8–F27. https://doi.org/10.1111/j.1468-0297.2007.02114.x.
Pitt, M.M., Rosenzweig, M.R., Hassan, M.N., 2012. Human capital investment and the gender division of labor in a brawn-based economy. Am. Econ. Rev. 102 (7), 3531–3560. https://doi.org/10.1257/2Faer.102.7.3531.
Restuccia, D., Yang, D.T., Zhu, X., 2008. Agriculture and aggregate productivity: a quantitative cross-country analysis. J. Monet. Econ. 55 (2), 234–250. https://doi.org/10.1016/j.jmoneco.2007.11.006.
Roodman, D., Nielsen, M.Ø., MacKinnon, J.G., Webb, M.D., 2019. Fast and wild: bootstrap inference in Stata using boot test. Stata J. 19 (1), 4–60. https://doi.org/10.1177/1536867x19830877.
Roy, S., 2015. Empowering women? Inheritance rights, female education and dowry payments in India. J. Dev. Econ. 114, 233–251. https://doi.org/10.1016/j.jdeveco.2014.12.010.
Schönberg, U., Ludsteck, J., 2014. Expansions in maternity leave coverage and mothers' labor market outcomes after childbirth. J. Labor. Econ. 32 (3), 469–505. https://doi.org/10.1086/675078.
Shi, X., 2018. Heterogeneous effects of rural-urban migration on agricultural productivity: evidence from China. China Agric. Econ. Rev. 10 (3), 482–497. https://doi.org/10.1108/CAER-10-2017-0193.
Shi, X., 2022a. Moving out but not for the better: health consequences of interprovincial rural-urban migration in China. Health Econ. 31 (4), 555–573. https://doi.org/10.1002/hec.4468.

Shi, X., 2022b. Land reallocation responses to China's New Rural Pension Scheme: Evidence from a regression discontinuity design. Journal of Rural Studies 92, 35–49. https://doi.org/10.1016/j.jrurstud.2022.03.018.

Shi, X., Shen, Y., 2023. Mental health penalties of having a child: findings from the China family panel studies. NPJ Mental Health Res. 2 (1), 7. https://doi.org/10.1038/s44184-023-00026-x.

Sorkin, I., 2017. The role of firms in gender earnings inequality: evidence from the United States. Am. Econ. Rev. 107 (5), 384–387. https://doi.org/10.1257/aer.p20171015.

Thu, Y.M., Htoo, S.S., Htwe, N.N., Gomersall, K., 2023. Rural women's participation and their decision making behavior in livestock management and household activities in central dry zone area of Myanmar. Agric. Rural Stud. 1 (1), 0004. https://doi.org/10.59978/ar01010004.

Wang, Q., Zhang, X., 2017. Three rights separation: china's proposed rural rights reform and four types of local trials. Land Use Policy 63, 111–121. https://doi.org/10.1016/j.landusepol.2017.01.027.

Wang, X., Shen, Y., 2014. The effect of China's agricultural tax abolition on rural families' incomes and production. China Econ. Rev. 29, 185–199. https://doi.org/10.1016/j.chieco.2014.04.010.

Ye, B., Zhao, Y., 2018. Women hold up half the sky? Gender identity and the wife's labor market performance in China. China Econ. Rev. 47, 116–141. https://doi.org/10.1016/j.chieco.2017.08.002.

Zhou, Y., Li, X., Liu, Y., 2020. Rural land system reforms in China: history, issues, measures, and prospects. Land Use Policy 91, 104330. https://doi.org/10.1016/j.landusepol.2019.104330.

Zinovyeva, N., Tverdostup, M., 2021. Gender identity, coworking spouses, and relative income within households. Am. Econ. J.: Appl. Econ. 13 (4), 258–284. https://doi.org/10.1257/app.20180542.